Taxation, Labor Supply, and the Distribution of Income

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Contents

Preface

Motivation

This dissertation consists of five empirical chapters that explore the causes of increasing economic inequality in Germany, the design of a just tax and transfer system, and of labor supply decisions under uncertainty.

Chapter 1 adds to the vibrant debate on increasing income inequality since 2000. It quantifies the determinants of increasing income inequality in Germany from 2002 to 2011. This period saw controversial reforms of the tax and transfer system, which contrary to common belief have led to a *decrease* in income inequality.

Differences in measures of income inequality gain a deeper meaning when they are framed in the context of social welfare functions and associated weights. Value judgments about the desirable distribution of income can be derived from the hypothetical situation where one chooses a distribution in complete ignorance of one's own relative position. If individuals maximize expected utility and assign the same probability to ending up in different positions, they will choose a utilitarian social welfare function (Harsanyi, 1953).[1] While many economists probably subscribe to this consequentialist – or welfarist – view, social welfare judgments are not limited to these considerations. In stark contrast and following the deontological tradition, Nozick (1974) formulated the entitlement theory according to which redistribution should only rectify *unjust* holdings of property. While the notion of just and unjust holdings of property arguably are not well-defined, from this point of view the redistribution of labor income is rarely justified – in contrast to the redistribution, e.g., of stolen goods. An important difference between this and the consequentialist point of view is that the "starting point" matters: When the distribution of gross incomes is just, the scope for redistribution is heavily limited (see also Feldstein, 1976). In practice, at least in the US, many individuals do not desire the degree of redistribution that would follow from welfarist considerations, but favor a tax schedule where the distribution of gross incomes matters for the desirable distribution of net incomes (Weinzierl, 2014). In a positive optimal taxation exercise, Chapter 2 evaluates under what normative criterion the 2015 German tax and transfer system is optimal. In particular, it contrasts the welfarist idea that follows Harsanyi (1953) with more general ideas closer to Nozick (1974), where gross incomes matter for the desirable distribution of net incomes.

Consequentialist value judgments make redistribution of income desirable. However, when designing a tax and transfer system, the efficiency costs of redistribution have to

[1] If individuals are extremely risk averse under this *veil of ignorance*, they only consider the possibility that their position will be the one of the worst off individual. This leads to Rawls' (1971) well-known criterion that welfare of the least well off person is maximized (see Nechyba, 2016, for a derivation).

be taken into account too (Mirrlees, 1971). In particular, labor income is not exogenous and individuals adjust their labor supply to taxation. The social planner therefore faces the equity-efficiency trade-off that the 'size of the cake' depends on how it is distributed. Chapters 3, 4, and 5 contribute to our understanding of the determinants of labor supply.

Chapter 3 provides a new method of estimating the Marshall elasticity, the relevant elasticity for evaluating the impact of tax-transfer reforms (Blundell and Macurdy, 1999). Moreover, it quantifies the contributions of wage and hours risk to overall labor income risk.

Labor supply not only reacts to changes in the net-of-tax wage, but to uncertainty itself. This is the case because individuals may work more in the face of wage risk in order to increase their precautionary savings. This mechanism of *precautionary labor supply* is studied in Chapter 4. Finally, Chapter 5 tests a standard model of life-cycle labor supply and finds that precautionary labor supply is less important than predicted by the model.

Main Findings and Contributions

The analysis is organized in five chapters, each of which is devoted to a specific research question and is based on evidence from micro data – the German Socio-Economic Panel (SOEP) in case of Chapters 1, 2, and 4 and the US-American Panel Study of Income Dynamics (PSID) in case of Chapters 3 and 5.

Chapter 1 proposes a method to decompose changes in income inequality into the contributions of policy changes, wage rate changes, and population changes while considering labor supply reactions. This method is applied to decompose the increase in income inequality in Germany from 2002 to 2011. The analysis complements previous papers that used different methods to "explain" increases in German income inequality (see Biewen and Juhasz, 2012; Peichl et al., 2012). In particular, I am the first to quantify the impact of labor supply adjustments to tax-transfer reforms in this context. The simulations show that tax and transfer reforms have had an inequality reducing effect as measured by the Mean Log Deviation and the Gini coefficient. For the Gini, these effects are offset by labor supply reactions. Changes in wage rates have led to a decrease in income inequality. The result implies that the increase in income inequality is due to changes in the population that are not explicitly modeled.

The main contribution in Chapter 2 is the generalization of an optimal taxation model (Saez, 2002) to allow for non-welfarist aims of the social planner. The exercise is motivated by the puzzling fact that if a social planner maximizes a weighted sum of utilities, high transfer withdrawal rates in many countries are only optimal if social weights for the working poor are very low. In contrast, optimal taxation studies usually assume that the social planner puts more weights on low income households than on higher income households.

We allow for the objective function for each individual to depend on pre-government income and calculate the planner's weights under different ideas of justness. The German tax and transfer schedule is in line with decreasing social weights if the social planner minimizes absolute sacrifice, i.e., if the social planner minimizes a function of the tax liability for each individual.

The following three chapters examine labor supply under uncertainty. In Chapter 3 overall labor income risk is decomposed into contributions from wage and hours uncertainty. This is a departure from most of the extant literature, which focuses on wage risk. Wage and hours uncertainty are in turn decomposed into permanent and transitory shocks. Permanent wage shocks might be caused by the obsolescence of human capital or promotions due to the acquirement of new skills. Permanent hours shocks might, e.g., be caused by injuries. In contrast, transitory hours shocks might occur because of brief needs of, e.g., children. In contrast to permanent wage shocks, permanent hours shocks are virtually nonexistent. Transitory wage shocks are more important than transitory hours shocks in terms of income. However, both types of transitory shocks play an economically significant role.

In order to disentangle hours shocks from labor supply reactions to wage shocks, we use the structure of a life-cycle labor supply model and estimate the transition of wage shocks to the marginal utility of wealth. We thereby obtain a sufficient statistic for the Marshall elasticity. This method of estimating the Marshall elasticity with labor supply data alone is the second main contribution of this chapter. The estimate of the Marshall elasticity is significantly negative, i.e., individuals decrease their labor supply as a reaction to permanent wage shocks.

Chapter 4 is the first study to empirically quantify the importance of precautionary labor supply defined as the difference between hours supplied in the presence of risk and hours under perfect foresight. We find that married men in Germany choose about 2.8% of their hours of work on average to shield against wage shocks. The effect is strongest for self-employed, but also relevant for other groups. If the self-employed faced the same wage risk as the median civil servant, their hours of work would reduce by 4.5%.

Finally, Chapter 5 proposes a simple test of a standard life-cycle labor supply model with constant relative risk aversion (CRRA) utility. It uses the second-order approximation of the labor supply Euler equation to test whether married men in the US react to uncertainty to the extent that is predicted by the model. The first result is that the estimated Frisch elasticity based on the second-order approximation is similar to the one obtained from the first-order approximation commonly estimated in the literature. But while agents use precautionary labor supply to cushion against wage risk, they do so to a lesser extent than predicted by CRRA utility. Therefore the model is rejected.

1. Why Has Income Inequality in Germany Increased from 2002 to 2011? [1]

1.1. Introduction

Income inequality has increased considerably in Germany from 2002 to 2011. The Gini coefficient of equivalized net household income has increased from 28.5 to 29.5 (own calculation). From a policy perspective it is important to learn about the determinants of increasing income inequality, in order to take appropriate countermeasures, e.g., if policy reforms have had an inadvertent inequality increasing effect. The time span from 2002 to 2011 is particularly interesting regarding the interaction of inequality and tax-transfer policy as it witnessed a strong increase in inequality as well as major reforms to the tax and transfer system: the controversial Hartz IV reforms of the transfer system as well as part of the phasing in of major tax reforms started in 2001. Increasing wage dispersion is another potential explanation for the increase in income inequality. These potential factors in increasing inequality are described in detail in section 1.3.

The aim of this study is to quantify the impact of policy reforms, changes in conditional wage rates, and remaining changes to the population on income inequality. To allow for the joint analysis of these factors, the decomposition framework by Bargain (2012a,b) is extended to explicitly account for the effect of changes in conditional wage rates in the spirit of Bourguignon et al. (2008). The decomposition method combines microsimulation, a structural labor supply model, and the construction of counterfactual wage rates using Mincer-style wage regressions. The decomposition is done in an entirely disaggregated way that is not limited to a specific class of inequality indices. It allows for the graphical representation of counterfactual distributions. Marginal effects of particular factors on inequality are calculated by comparing actual and counterfactual distributions and thus can be interpreted as ceteris paribus changes unconfounded by demographic or business cycle changes. The decomposition method is explained in section 1.4.

This study contributes to the literature on the decomposition of differences between two income distributions and in particular to the literature using microsimulation techniques. Bargain and Callan (2010), Bargain (2012b), Liégeois and Dekkers (2014), and Bargain et al. (2017) simulate counterfactual net incomes by applying the tax and transfer system of a given period to the population of another period using a detailed tax and transfer calculator to obtain intermediate distributions. Creedy and Herault (2011)

[1]This chapter is based on Jessen (2016).

and Bargain (2012a) expand the microsimulation approach by simulating counterfactual labor supply decisions. Bargain et al. (2015) simulate responses of taxable income. Herault and Azpitarte (2016) allow for the simulation of a wide range of additional determinants. The study at hand combines the simulation of counterfactual labor supply with the prediction of counterfactual wages following Bourguignon et al. (2008) akin to the decomposition method introduced by Blinder (1973) and Oaxaca (1973). As pointed out by Bourguignon et al. (2008), the combination of strictly parametric techniques offers the advantage of a straight-forward economic interpretation (see also Brewer and Wren-Lewis, 2015; Herault and Azpitarte, 2016).

Apart from the methodology, the analysis conducted in this study adds to a developing literature on the causes of increases in income inequality in Germany in recent years (Arntz et al., 2007; Biewen and Juhasz, 2012; Peichl et al., 2012; Bargain et al., 2017; Biewen et al., 2016). These studies are summarized in the next section.

The results are presented in section 1.5. The decomposition shows that changes of the tax and transfer system have slightly alleviated inequality as measured through the Gini index and the Mean Log Deviation (MLD). The negative effect of policy changes on the Gini is offset by labour supply reactions. In contrast, policy changes have led to an increase in the ratio between the 90th and the 50th income percentile (Q90/50). The overall effect of changes in wage rates on inequality is found to be negative. Thus, the overall increase in income inequality was caused by changes in characteristics of the population that are not explicitly modelled, e.g. in the household structure.

1.2. Previous Studies on Germany

A few papers decompose the overall change in income inequality in Germany between two periods into different factors. Table 1.1 summarizes the methods and results of these studies. Biewen and Juhasz (2012) apply a reweighting technique (DiNardo et al., 1996) along with parametric techniques to study the rise of income inequality from 1999/2000 to 2005/2006. They find that changes in household characteristics as well as changes in the transfer system have had a minor effect. Changes in household structure, labor market returns, conditional employment outcomes and changes in the tax system have led to an increase in income inequality. Their measure of conditional labor market returns is not limited to the effect of wage changes, but, given their broad definition of employment outcomes, includes hours adjustments. Biewen et al. (2016) carry out a similar analysis for the periods 2005/2006 to 2010/2011. They find that income inequality did not increase in this period. While inequality in individual monthly labor incomes increased, changes in conditional annual labor market returns had no significant impact on the Gini, but led to a decrease in the Theil index. The impact of changes in capital returns is found to have had a negligible effect on income inequality.

Table 1.1.: Overview of Previous Studies on Germany

	Biewen and Juhasz (2012)	Biewen et al. (2016)	Peichl et al. (2012)	Bargain et al. (2017)	Arntz et al. (2007)
Periods	1999/2000 – 2005/2006	2005/2006 – 2010/2011	1991 – 2007	2008 – 2013	2003 – 2005*
Policy changes	0 (transfer); + (tax)**	- (transfer); 0 (tax)		0	0 (total population); - (transfer recipients)
Method	Taxes: estimation of tax schedule with polynomial; transfers: microsimulation			Microsimulation	Microsimulation with labor supply simulation
Household structure	+	0	+ (Mean Log Deviation)		
Method	Reweighting of hhs with types distinguished by number of adults and children and whether adults are pensioners		Subgroup decomposition by number of adults and children***		
Household characteristics	0	+			
Method	Reweighting along hh members' characteristics				
Labor supply	+	0	+ (Mean Log Deviation)		
Method	Reweighting along hh members' employment outcomes conditional on hh characteristics		Subgroups additionally defined by number of employed individuals****		
Labor market returns	+	- (Theil)			
Method	Linear prediction of log household labor income conditional on hh characteristics and employment				

Note: Effect on the Gini coefficient, if not otherwise noted. +: inequality increase; -: inequality decrease; 0: very small or insignificant effect. * Ex-ante analysis of 2005 *transfer* reforms only. ** No significant effect of taxes conditional on all other effects. *** Re-weighting along subgroups yields similar results. **** Employment effect is not disentangled from the effect of changes in the household structure. *Source:* Author's own table

Peichl et al. (2012) use subgroup decomposition and reweighting to quantify the impact of changes in household size and employment outcomes on the increase of income inequality from 1991 to 2007. They find that the decreasing average household size in Germany is associated with an increase in inequality. Bargain et al. (2017) focus on static policy effects for the period 2008 to 2013. Here, policy changes have had no effect on overall inequality and a positive effect on poverty measures. Arntz et al. (2007) conduct an ex ante study of the distributional effect of the 2005 Hartz IV reform of the transfer system described in section 1.3.1. They find no direct effect of the reform on the Gini coefficient, while some other inequality measures decreased. For people directly affected by the reform the changes in the transfer system have led to a substantial decrease in the Gini coefficient.

The present paper is the first to estimate the effect of tax and transfer reforms in Germany on inequality taking labor supply reactions into account. It is also the first to evaluate the impact of conditional hourly wage rates on the inequality of household incomes.

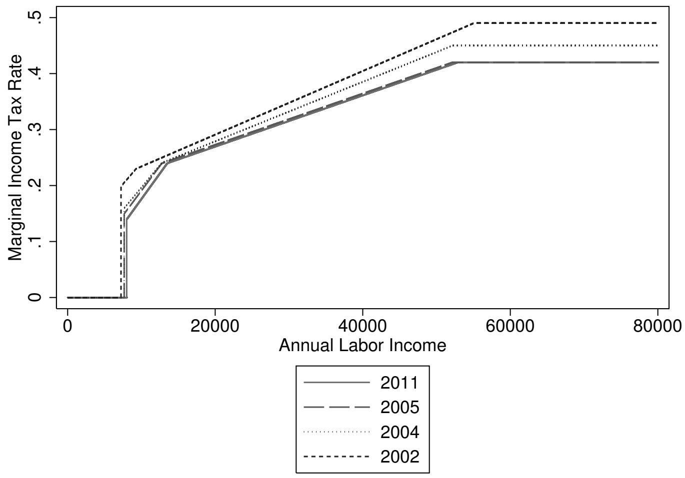
1.3. Factors in Increasing Inequality

1.3.1. Policy Changes

Figure 1.1 shows marginal income tax rates for a single household for 2002 and 2011, the two years analyzed in this study, as well as for 2004 and 2005. The figure was constructed using the STSM (Steuer-Transfer-Mikrosimulationsmodell, see Steiner et al., 2012), a tax- and transfer microsimulation model for Germany. Note that the aim of this paper is to estimate the effect of the overall change in the tax and transfer system. Therefore the 'intermediate' tax schedules of 2004 and 2005 are displayed as additional information, but are not used for the construction of counterfactual distributions. For all levels of gross income the marginal income tax rate of 2011 is lower than the one in 2002. The initial marginal tax rate was decreased gradually from 19.9 % to 14 % in 2009. The top marginal tax rate applicable for incomes exceeding 55 000 Euro (year 2002) was decreased gradually from 49 % to 42 % in 2005 and the top marginal tax rate income threshold was decreased from 55 007 (2002) to 52 151 Euro (2004). In 2007 the so-called rich people's tax of 45 % for gross incomes exceeding 250 000 Euro per year came into force (not displayed in Figure 1.1). Additionally, the size of the tax brackets was regularly adjusted slightly to account for inflation. Previously, capital income was part of the income tax base, the year 2009 saw the introduction of a capital income tax of 25 %, leading to a tax reduction for earners of capital income with a marginal income tax rate exceeding this figure.

The transfer system has been radically overhauled in the course of the Hartz IV reform. While the short-term unemployed, who have previously been employed, generally are entitled to a transfer called Unemployment Benefit (colloquially referred to as "Un-

Figure 1.1.: Marginal Income Tax Rates for a Single Household



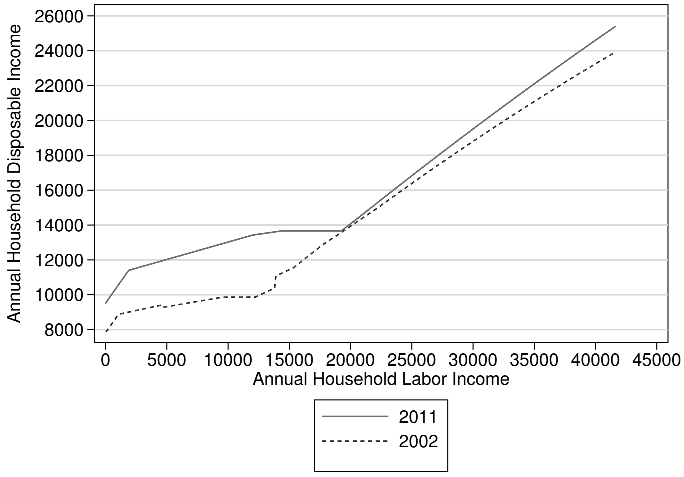
Source: Own graph based on the STSM

employment Benefit I"),[2] two kinds of means-tested transfers existed for the long-term unemployed before the reform: Unemployment Assistance, which amounted to 53% of previous labor income (57% if a child lived in the household) and Social Assistance covering the social existence minimum. In 2005, these two transfers were replaced with the so-called Unemployment Benefit II, which only ensures the social existence minimum. Individuals deemed able to participate in the labor market were subject to these changes. Former recipients of Unemployment Assistance experienced a potentially severe reduction of income due to the introduction of Unemployment Benefit II. The aim of the reform in this regard was to improve incentives for the unemployed to accept job offers. However, the level of Unemployment Benefit II is slightly higher than Social Assistance, so that former recipients of the latter were better off. Overall, the Hartz IV reform has led to an increase in government spending (Biewen and Juhasz, 2012) and an ex-post evaluation has shown that average equivalized net income of previous recipients of Unemployment Assistance was higher a year after the reform than before (Bruckmeier and Schnitzlein, 2007). As this reform of the transfer system implied lower transfers for some and higher transfers for others, the distributional effect is a priori ambiguous.

In both years 2002 and 2011, marginal employment (so-called 'mini jobs' for gross incomes of up to 325 Euro per month in 2002 and 400 in 2011) was exempted from taxes and social security contributions. However, in 2002, when gross income exceeded the

[2] The period of entitlement to this transfer was reduced from up to 36 months to 12 months and 18 months for individuals over 55 years of age. The entitlement period for the elderly was further increased in 2006 and in 2008. In 2011 the maximum entitlement period for individuals of at least 58 years of age was 24 months. Compared to the year 2002, this still means a reduction in the maximum entitlement period and could potentially have led to an increase in inequality.

Figure 1.2.: Budget Constraint of a Single Household in 2011 Euro



Source: Own graph based on the STSM

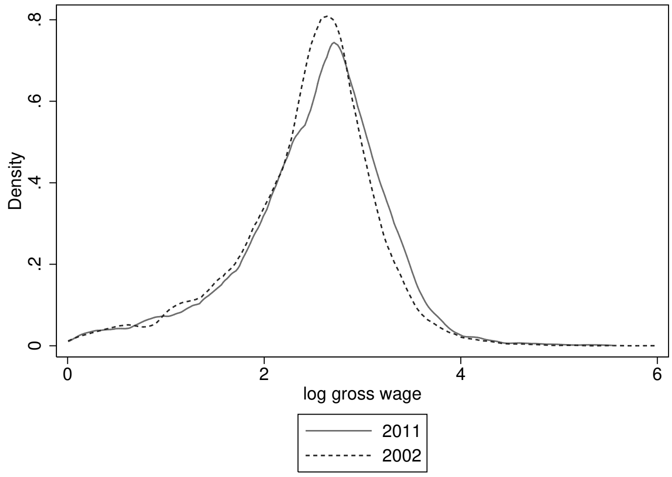
threshold for marginal employment, the average social security contribution rate paid by employees jumped from 0 to the full rate of regular jobs. This implied negative incentives for the marginally employed to work slightly more. Since 2003 average social security contributions paid by employees increase slightly with increasing gross income until they reach 20 percent at a monthly gross income of 800 Euro (year 2011). Jobs with gross incomes in this range are called 'midi jobs'.

Finally, the Citizen Relief Act (Bürgerentlastungsgesetz) – in effect since July 2010 – brought about an increase in the possible tax allowances for insurance premia. Overall, tax reforms in the analyzed time-span produced lower marginal tax rates both at the upper and at the lower end of the income distribution, so the distributional effect is unclear a priori. If the substitution effect dominates the income effect, decreased marginal tax rates lead to increases in labor supply over the entire distribution. On the other hand, increased generosity of transfers for the long-term unemployed implies lower labor supply incentives for this group. This is expected to have an inequality increasing effect.

Figure 1.2 shows the change in the budget constraint for a single without children and without wealth.[3] For low values of labor income the household is eligible for Social Assistance in 2002 and for Unemployment Benefit II in 2011. Only labor income is varied along the horizontal axis and the corresponding net income is displayed on the vertical axis. In contrast to figure 1.1, which shows marginal income tax rates, figure 1.2 additionally accounts for transfers and social security contributions. For low levels of gross labor income, the transfer received in the 2011 regime is far more generous. The

[3]For the 2002 budget constraint, gross labor incomes have been deflated to 2002 levels and – along with simulated net incomes – inflated back to 2011 levels.

Figure 1.3.: Densities of Log Hourly Wage in 2011 Euro



Source: Own calculation base on the SOEP.

lower marginal tax rates of 2011 translate into a steeper slope of the budget constraint starting at an annual gross income of about 20000 Euro.

1.3.2. Wage Dispersion

Wages in Germany have dispersed considerably since the 1990s, see, e.g., Fuchs-Schündeln et al. (2010). Several studies attest that this is partly due to polarization, which is consistent with skill-biased technological change, see, e.g., Dustmann et al. (2009). However, there is less evidence for this phenomenon in the time-span beginning in 2002. Therefore it is not to be expected that changes in conditional wage rates have led to an increase in income inequality. An alternative explanation for increasing wage dispersion is selection into employment. A recent employment boom in Germany (see, e.g., Biewen et al., 2016) is likely to have changed the composition of the work force, possibly at roughly constant *conditional* wage rates.

Figure 1.3 depicts the estimated Epanechnikov kernel density of log hourly wage densities in the two years.[4] It shows a marked increase in mass at the right of the distribution from 2002 to 2011 implying a relative increase in the number of high-paying jobs.

[4]Following Biewen and Juhasz (2012), a fixed bandwidth of 0.175 is used throughout the paper.

1.4. Empirical Strategy: Decomposition

1.4.1. Counterfactual Distributions and Decomposition

The decomposition is restricted to parametric techniques that have a straightforward economic interpretation. Let y_{bd}^{ce} be a matrix that describes socio-demographic characteristics and market incomes of the people observed in period b who receive the conditional wage rates of period d with work hours given the incentives of the tax and transfer regulations of period c and the incentives given conditional gross wages of period e. As described in subsection 1.4.4, work hours are simulated conditional on the budget constraints of individuals, which in turn are determined by the tax and transfer system and the gross hourly wage. Let x_a be the tax and transfer function that translates market income and socio-demographic characteristics into net income of each household and denote I an inequality index so that $I\left(x_a\left(y_{bd}^{ce}\right)\right)$ denotes inequality in a given observed or counterfactual situation.

The decomposition relies on the construction of counterfactual net incomes for observed households. *Household* gross income is the sum of individual labor incomes L of all household members and other pre-government household income, e.g., capital income.

Specifically, let $I\left(x_{2011}\left(y_{2011,2011}^{2011,2011}\right)\right)$ be an inequality index of the actually observed outcomes of 2011 and $I\left(x_{2002}\left(y_{2002,2002}^{2002,2002}\right)\right)$ inequality of observed outcomes in 2002. Marginal effects are given by the change in income inequality obtained by changing one factor while keeping everything else equal.

Policy effect — The *static* marginal effect of policy changes on income inequality is

$$I\left(x_{\mathbf{2011}}\left(y_{2002,2002}^{2002,2002}\right)\right)-I\left(x_{2002}\left(y_{2002,2002}^{2002,2002}\right)\right). \tag{1.1}$$

The *total* policy effect is given by

$$I\left(x_{\mathbf{2011}}\left(y_{2002,2002}^{\mathbf{2011},2002}\right)\right)-I\left(x_{2002}\left(y_{2002,2002}^{2002,2002}\right)\right), \tag{1.2}$$

i.e., the difference between actual inequality in 2002 and inequality of the counterfactual distribution where net incomes are calculated using the 2011 tax transfer system and labor supply reactions are simulated conforming to incentives in 2011 for the 2002 sample.

To obtain this marginal effect, counterfactual gross incomes need to be calculated. For the total policy effect, counterfactual individual labor income of a given household member is given by

$$\hat{L}=\left(\hat{h}|T_{2011},w_{2002},z_{2002}\right)\times w_{2002}, \tag{1.3}$$

where $\hat{h}|T_{2011}, w_{2002}, z_{2002}$ denotes predicted annual hours of work given the tax and transfer system T of 2011 while observed wage rates w and household characteristics z of the year 2002 are used. Net incomes x_{2011} are calculated applying microsimulation to gross incomes, taking into account relevant household characteristics and income sources.

Wage effect — Similarly, the *static* wage effect is given by

$$I\left(x_{2002}\left(y^{2002,2002}_{2002,\mathbf{2011}}\right)\right) - I\left(x_{2002}\left(y^{2002,2002}_{2002,2002}\right)\right). \tag{1.4}$$

The *total* wage effect is

$$I\left(x_{2002}\left(y^{2002,\mathbf{2011}}_{2002,\mathbf{2011}}\right)\right) - I\left(x_{2002}\left(y^{2002,2002}_{2002,2002}\right)\right), \tag{1.5}$$

which is the difference between 2002 income inequality and inequality of the intermediate distribution with wages as in 2011 predicted for all workers and labor supply adjusted according to these counterfactual wages.

Counterfactual individual labor incomes for this calculation are obtained from

$$\hat{L} = \left(\hat{h}|T_{2002}, \hat{w}_{2011}, z_{2002}\right) \times \hat{w}_{2011}. \tag{1.6}$$

Predicted hours of work are obtained by simulating labor supply given the counterfactual household budget constraint obtained when substituting actual hourly wages with their predicted counterparts. Counterfactual wages conditional on individual characteristics in 2002 are given by

$$\hat{w_{2011}} = c_{2002} \times \hat{\beta}_{2011} + \epsilon_{2002}, \tag{1.7}$$

where the coefficients $\hat{\beta}_{2011}$ are obtained from a wage regression using the 2011 population and c_{2002} are actual individual characteristics. ϵ_{2002} is the readjusted residual of 2002 (see subsection 1.4.2 for details on the entire procedure).

Combined effect — The combined effect of changes in conditional wage rates and the tax-transfer system is

$$I\left(x_{\mathbf{2011}}\left(y^{2002,2002}_{2002,\mathbf{2011}}\right)\right) - I\left(x_{2002}\left(y^{2002,2002}_{2002,2002}\right)\right) \tag{1.8}$$

without labor supply reactions and

$$I\left(x_{\mathbf{2011}}\left(y^{\mathbf{2011},\mathbf{2011}}_{2002,\mathbf{2011}}\right)\right) - I\left(x_{2002}\left(y^{2002,2002}_{2002,2002}\right)\right) \tag{1.9}$$

with labor supply reactions. In this case counterfactual labor incomes are given by

$$\hat{L} = \left(\hat{h}|T_{2011}, \hat{w}_{2011}, z_{2002}\right) \times \hat{w}_{2011}. \tag{1.10}$$

To give a concrete example for the procedure, the counterfactual distribution for the combined policy and wage effect including labor supply is obtained following four steps: 1) Estimate the wage equation using the 2011 sample and predict counterfactual wages for the 2002 population. 2) Use microsimulation to calculate the counterfactual budget constraint (i.e., net household incomes for different labor supply choices) for the 2002 population given the 2011 tax-transfer system and 2011 wages. 3) Estimate the structural labor supply model using the observed 2002 population, wages and tax-transfer system. 4) Use these labor supply model estimates to predict labor supply choices given the counterfactual budget constraint.

Population effect — The effect of changes in the population, i.e., everything that is not explicitly modeled, is calculated by subtracting the 2002 status quo from a counterfactual distribution of the 2011 population with counterfactual 2002 wages, tax and transfer system and labor supply:

$$I\left(x_{2002}\left(y^{2002,2002}_{\mathbf{2011},2002}\right)\right)-I\left(x_{2002}\left(y^{2002,2002}_{2002,2002}\right)\right). \tag{1.11}$$

Counterfactual labor incomes for this step are given by

$$\hat{L}=\left(\hat{h}|T_{2002},\hat{w}_{2002},z_{2011}\right)\times\hat{w}_{2002}, \tag{1.12}$$

i.e. the actual population of 2011, where 2002 wages are predicted and hours of work are simulated given the household budget constraint if the tax-transfer system and wages conform to 2002.

In section 1.5, marginal effects of wage and policy changes are reported using the year 2002 as base as in the equations above. As a robustness test, results using the year 2011 as base year are reported.[5] While the interpretation of the effects of wage and policy changes is straight-forward, the population effect represents a residual capturing all household characteristics that are not explicitly modeled, e.g. demographic changes, changes in assortative mating, changes in the distribution of capital income, changes in education choices, etc.

1.4.2. Changes in Wage Rates

The effect of conditional wages is analyzed by running a regression of log hourly wages on years of education[6], work experience and experience squared as well as years not worked

[5]Another possibility would be to calculate 'intermediate contributions', i.e., calculate the difference between two counterfactual distributions. For instance, one could first calculate the contribution of wage rate changes and in a second step calculate the contribution of tax-transfer changes *conditional on wage changes.* One could then calculate the average contribution of, e.g, wage rate changes over all – essentially arbitrary – decomposition orders. Instead, this paper focuses on marginal effects, since they have a precise and intuitive economic interpretation.

[6]An alternative estimation with categorical education variables is reported in tables A.2 and A.3.

in the last ten years to capture loss of human capital. Heckman's (1979) method is used to account for selection bias with variables capturing the number of children, family status, and the income of other household members as exclusion restriction. Seperate regressions are run for women and men and East and West Germany.

The coefficients and the constant for the years 2002 and 2011 are used to predict counterfactual wages for the respective other years' populations.[7] The entire labor incomes of employees are replaced with the counterfactual predictions.

Table 1.2.: Wage Regression 2002

	Men East	Women East	Men West	Women West
Ln(Hourly Wage)				
Years of Schooling	0.0656***	0.0647***	0.0650***	0.0719***
	(0.00529)	(0.00716)	(0.00241)	(0.00360)
Years not Worked	-0.150***	-0.112***	-0.0952***	-0.0369***
	(0.0197)	(0.0150)	(0.00898)	(0.00500)
Experience	0.0541***	0.0776***	0.0677***	0.0644***
	(0.00662)	(0.00844)	(0.00322)	(0.00443)
Experience^2/100	-0.104***	-0.164***	-0.123***	-0.128***
	(0.0165)	(0.0224)	(0.00793)	(0.0120)
Constant	1.183***	0.921***	1.444***	1.120***
	(0.101)	(0.148)	(0.0454)	(0.0776)
Mills	0.0737	0.0472	0.0227	0.0942*
lambda	(0.0713)	(0.0861)	(0.0314)	(0.0447)
N	2616	2899	7586	8253

Standard errors in parentheses
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Source: Own calculation based on the SOEP

For instance, for the wage effect with base year 2002, equation (1.4), hourly wages of the 2002 sample are replaced with predicted wages using coefficients of the 2011 wage regression. Following Bourguignon et al. (2008) and Bourguignon and Ferreira (2004), each individual's residual is multiplied by the ratio of standard deviations of residuals of the counterfactual and the observed period and added to the deterministic (predicted) part of the counterfactual wage.[8] Gross labor incomes are calculated by multiplying the counterfactual hourly wage with actual hours of work. Counterfactual wages are only

[7] 2002 wages are inflated to 2011 levels for the regressions. Counterfactual predicted wages for the 2002 sample are deflated to 2002 levels.

[8] The ratio of standard deviations of 2002 and 2011 is 1.002 implying virtually no change in within-group wage inequality.

predicted for employees. For the self-employed observed wages are used. For the status quo distribution, observed instead of predicted values are used in the analysis.

Table 1.3.: Wage Regression 2011

	Men East	Women East	Men West	Women West
Ln(Hourly Wage)				
Years of Schooling	0.0550***	0.0677***	0.0546***	0.0565***
	(0.00522)	(0.00643)	(0.00255)	(0.00330)
Years not Worked	-0.137***	-0.113***	-0.146***	-0.0484***
	(0.0159)	(0.0129)	(0.00990)	(0.00548)
Experience	0.0711***	0.0544***	0.0682***	0.0520***
	(0.00721)	(0.00743)	(0.00387)	(0.00429)
Experience^2/100	-0.156***	-0.106***	-0.131***	-0.0948***
	(0.0187)	(0.0188)	(0.00968)	(0.0111)
Constant	1.235***	1.035***	1.486***	1.330***
	(0.0902)	(0.125)	(0.0465)	(0.0740)
Mills	0.0818	-0.00145	0.0970**	0.0274
lambda	(0.0718)	(0.0717)	(0.0370)	(0.0458)
N	2419	2695	6898	7825

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Source: Own calculation based on the SOEP

The results of the wage regressions are reported in Tables 1.2 and 1.3. The signs of the coefficients are as expected, implying positive returns to schooling, positive and decreasing returns to experience, a wage penalty to human capital loss and – if significant – a positive selection term. They offer no evidence for skill-biased technological change in the observed period, instead, the returns to schooling have decreased for all groups except East German women. However, it should be kept in mind that changes in conditional wage rates reflect changes in both labor demand, e.g., because of skill-biased technological change, and labor supply. Moreover, the finding of decreasing education premia is not robust to the use of categorical education variables, see Tables A.2 and A.3 in the Appendix. But this does not change the results regarding the impact of conditional wage changes on income inequality.

1.4.3. Tax and Transfer System: Simulated Net Incomes

Counterfactual net incomes and budget constraints are calculated using the microsimulation model STSM, see Steiner et al. (2012) for additional information and Jessen et al.

(2017b) for a detailed depiction of budget constraints and marginal tax rates simulated with the STSM. The STSM covers the German tax and transfer system and accounts for deductions, allowances, social security payments and child benefits as well as interactions of the components of the tax and transfer system. When simulating counterfactual net incomes, all monetary variables in the data set are inflated or deflated respectively to the policy year. The simulated net incomes are then deflated or inflated back to the data year.

1.4.4. Behavioral Effects

Labor supply reactions to policy and wage changes are simulated via a random utility discrete choice model following van Soest (1995). For the estimation of the labor supply model the sample is restricted to household heads and partners with flexible labor supply, i.e., working age individuals excluding self-employed, civil servants, the severely disabled and people in parental leave. Households are assumed to jointly maximize utility, which depends on disposable household income and leisure of the male and female partner.

The coefficients of the utility function in turn depend on household characteristics such as the household members' age and the number of children. Weekly labor supply is discretized into six categories for women, five for men, and thus 30 for couples mimicking the observed distribution of labor supply. The net income for each labor supply category is calculated using the STSM. Gross labor income is given by the product of work hours and the (actual or counterfactual) hourly wage. Potential hourly wages of the unemployed are predicted using the selectivity corrected wage regressions described above.[9] Let Lf denote leisure of the female partner, Lm leisure of the male partner, C consumption, and ε a random disturbance. Then the utility of household i of choice alternative j is given by

$$V_{ij} = U(Lf_{ij}, Lm_{ij}, C_{ij}) + \varepsilon_{ij}. \tag{1.13}$$

The translog specification of the deterministic part of individual utility is used, allowing for interactions of the components of the utility fuction, i.e.:

$$\begin{aligned} U_{ij} &= \beta_1 ln(C_{ij}) + \beta_2 ln(C_{ij})^2 + \beta_3 ln(Lf_{ij}) + \beta_4 ln(Lf_{ij})^2 + \beta_5 ln(Lm_{ij}) \\ &+ \beta_6 ln(Lm_{ij})^2 + \beta_7 ln(C_{ij}) ln(Lf_{ij}) + \beta_8 ln(C_{ij}) ln(Lm_{ij}) + \beta_9 ln(Lf_{ij}) ln(Lm_{ij}). \end{aligned} \tag{1.14}$$

Heterogeneity between households' utility functions is incorporated through taste shifters – observed household characteristics that affect some of the coefficients of the utility function:

[9] For simulations with counterfactual wages, wages of the employed are predicted as well, see subsection 1.4.2.

$$\beta_1 = \alpha_0^C + X_1'\alpha_1^C$$
$$\beta_2 = \alpha_0^{C^2} + X_2'\alpha_1^{C^2}$$
$$\beta_3 = \alpha_0^{Lf} + X_3'\alpha_1^{Lf}$$
$$\beta_5 = \alpha_0^{Lm} + X_4'\alpha_1^{Lm}.$$
$$\beta_9 = \alpha_0^{Lf x Lm} + X_5'\alpha_1^{Lf x Lm}.$$

X_1, X_2, X_3, X_4, and X_5 contain individual and household characteristics like age, disability indicators, whether the observed person is a German citizen, and number and age of children.

The error terms ε_{ij} are assumed to be independently and identically distributed across hour categories and households according to the extreme value type I distribution. As shown in McFadden (1974), the probability that alternative k is chosen by household i is then given by:

$$P_{ik} = Pr(V_{ik} > V_{ij}, \forall j \in 1 \dots J) = \frac{exp(U_{ik})}{\sum_{j=1}^{J} exp(U_{ij})}. \tag{1.15}$$

Alternative k is chosen if it implies a higher utility than any other alternative. Changes in net income associated with specific hours points lead to changes in the choice probabilities given by equation (1.15). These allow for the calculation of labor supply effects of the hypothetical tax and transfer systems or gross wages.

Estimation results and resulting elasticities are reported in the appendix in tables B.5 and B.6. The uncompensated labor supply elasticity for women in couples is particularly large and cross-wage elasticities are negligible, in line with common previous findings in the literature summarized, e.g., in Blundell and Macurdy (1999).

Note that the model assumes constant wage rates. In practice, increases in labor supply lead to decreases in market wage rates, which, in turn, lead to decreases in labor supply. Neglecting this effect is likely to lead to an overestimation of labor supply effects. However, as will be seen, estimated labor supply reactions to policy and wage changes are small. Therefore, equilibrium effects are likely limited.

1.4.5. Data

This study is based on the Socio-Economic Panel (SOEP)[10], a yearly representative survey of German households. See Wagner et al. (2007) for further information. The concept of income in this study is annual equivalent post-government income. Like most surveys the SOEP does not capture the very top of the income distribution. Bach et al. (2009)

[10] Socio-Economic Panel (SOEP), data for years 1984-2012, version 29, SOEP, 2014, doi: 10.5684/soep.v29.

combine the SOEP with income tax return data to cover the entire distribution of market incomes until the year 2003. They find that the SOEP serves reasonably well to describe the evolution of income inequality as measured with the inequality indices used in this study, while it fails to describe the change of the top-focused entropy index GE(2). Table 1.4 shows descriptive statistics for the most important variables for the years 2002 and 2011. The means of net household income as well as personal net income equivalized according to the OECD modified equivalence scale[11] have increased from 2002 to 2011. The SOEP provides information for weekly work hours and for annual labor incomes. Annual work hours are given by 52 times weekly work hours and hourly wages are calculated by dividing annual labor income through annual work hours.[12] Counterfactual labor incomes as predicted by the labor supply model are obtained by multiplying the hourly wage by the counterfactual weekly hours of work times 52. Average hourly wages, hours of work as well as the numbers of adults and children per household have decreased slightly from 2002 to 2011.

Table 1.4.: Descriptive Statistics

	2002		2011	
	Mean	SD	Mean	SD
Net household income	36 856.27	24 821.02	37 405.98	31 402.50
Equivalent Net household income	20 933.44	12 617.48	21 712.71	16 864.70
Hourly gross wage	15.89	12.45	15.24	12.73
Weekly work hours	36.35	12.30	35.89	12.57
Years of education	11.04	3.64	11.44	3.85
Household members with age> 13	2. 18	.93	2.14	.95
Children in household	.59	.92	.50	.89
Observations	27633		24780	

Monetary variables in 2011 real Euro
Only positive wage and work hours
Source: Own calculation based on the SOEP

1.5. Decomposition Results

This section shows marginal effects of wage rate and tax and transfer changes. They have been calculated as ceteris paribus effects of changes in labor market returns and

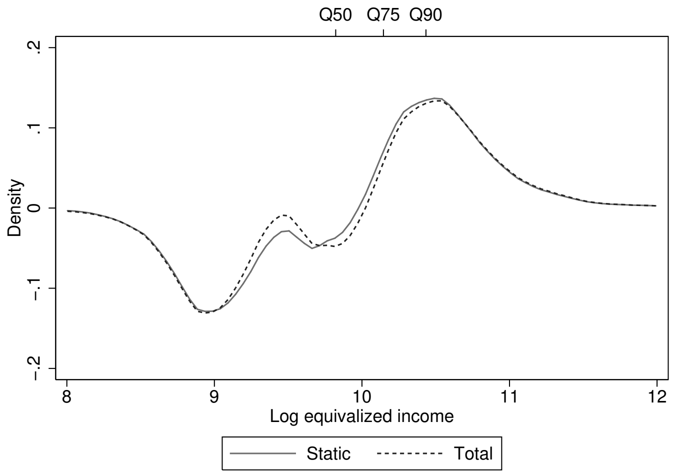
[11] I.e., net household incomes are divided by 1 plus 0.5 for each additional adult and 0.3 for each child under 14 years.

[12] Using less than 52 weeks per year in order to incorporate holidays would lead to larger hourly wages. This normalization would leave the results regarding the distribution of annual income unchanged.

the tax and transfer system, i.e., everything is kept at the 2002 level and only one factor is changed. Following Biewen and Juhasz (2012), this comes closest to the "effect" of a particular factor. A second exercise, where everything is kept at the 2011 level and only one factor is changed to the 2002 level is briefly described in subsection 1.5.4 and reported in Appendix A.

1.5.1. Policy Effect

Figure 1.4.: Policy Effect – Base Year 2002



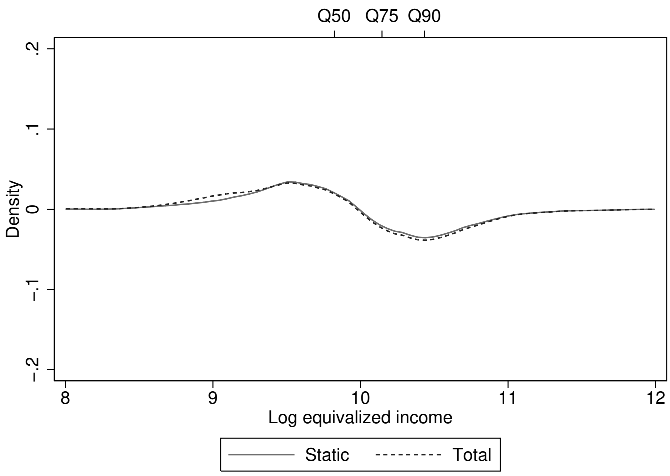
Difference between actual distribution in 2002 and counterfactual distributions applying the 2011 tax and transfer system to the 2002 population (static, $x_{\mathbf{2011}}(y^{2002,2002}_{2002,2002}) - x_{2002}(y^{2002,2002}_{2002,2002})$) and additionally simulating labor supply reactions (total, $x_{\mathbf{2011}}(y^{\mathbf{2011},2002}_{2002,2002}) - x_{2002}(y^{2002,2002}_{2002,2002})$). Source: Own calculation based on the SOEP and the STSM

The solid line in figure 1.4 shows the static policy effect. It is the difference between the actual estimated Epanechnikov kernel densities of log equivalized annual net income for the population of 2002 and the counterfactual distribution where the tax and transfer system is that of 2011 but work hours remain as in 2002. The dashed line shows the total policy effect, i.e., the counterfactual distribution where the tax and transfer system is as in 2011 and labor supply reactions to the tax and transfer changes are simulated. The 50th, 75th, and 90th percentile of the income distribution in the status quo are labeled on the upper horizontal axis.

The static effect of policy reforms yields a decrease in density at the bottom of the distribution, which is in line with the findings in Biewen and Juhasz (2012). It can be explained with former recipients of Social Assistance receiving the more generous Unemployment Benefit II. Moreover, density at the right of the distribution is increased due to policy reforms – this is the effect of tax reductions. Compared to the static counterfactual, labor supply reactions to policy reforms seem to have partly offset the static effect. The density at the bottom of the distribution is closer to the status quo. In the lower half of the distribution labor supply effects shift the distribution to the left and between the 50th and the 90th percentile, labor supply leads to a shift to the right. This reflects the change in labor supply incentives: more generous transfers have decreased incentives for low income households, while lower tax rates have increased incentives for higher income households to work.

1.5.2. Wage Effect

Figure 1.5.: Wage Effect – Base Year 2002



Difference between actual distribution in 2002 and counterfactual distributions applying the 2011 conditional wage rates to the 2002 population (static, $x_{2002}(y^{2002,2002}_{2002,\mathbf{2011}}) - x_{2002}(y^{2002,2002}_{2002,2002})$) and additionally simulating labor supply reactions (total, $x_{2002}(y^{2002,\mathbf{2011}}_{2002,\mathbf{2011}}) - x_{2002}(y^{2002,2002}_{2002,2002})$)). Source: Own calculation based on the SOEP and the STSM

Figure 1.5 shows the difference between the actual log income distribution of 2002 along with counterfactual distributions applying 2011 wage rates with (dashed line) and without (solid line) labor supply reactions to wage changes. Applying the coefficients of the 2011 wage regression to the 2002 population leads to a slight shift of the distribution to the left, the labor supply effect is negligible.

1.5.3. Summary of Effects

Table 1.5 shows the Gini along with two entropy measures, the Theil index (GE(1)) and the Mean Log Deviation (MLD, GE(0)), as well as the ratio between the 90th and the 50th income percentile (Q90/50) for the year 2002 (Status quo) and the difference in inequality between the actual 2002 distribution and the counterfactual distributions depicted in Figure 1.4 and Figure 1.5. Additionally, the joint effect of policy and wage changes as well as the effect of population changes are reported. Inequality indices for the status quo are calculated using actual observations with net incomes calculated using microsimulation. Wage effects are based on wage regressions, policy effects on microsimulation, and labor supply effects on structural labor supply simulation.

Table 1.5.: Decomposition with Base 2002

Marginal effect		Gini	Theil	MLD	Q90/50
$x_{2002}(y^{2002,2002}_{2002,2002})$	Status quo	28.5***	.144***	.143***	1.82***
		(.32)	(.0051)	(.0038)	(.0203)
$x_{\mathbf{2011}}(y^{2002,2002}_{2002,2002})$ $-x_{2002}(y^{2002,2002}_{2002,2002})$	Static tax and transfer	-.4***	0.00	-.01***	.026***
		(.11)	(.0017)	(.0022)	(.009)
$x_{\mathbf{2011}}(y^{\mathbf{2011},2002}_{2002,2002})$ $-x_{2002}(y^{2002,2002}_{2002,2002})$	Total tax and transfer	0.0	.002	-.008***	.044***
		(.13)	(.0018)	(.0025)	(.0101)
$x_{2002}(y^{2002,2002}_{2002,\mathbf{2011}})$ $-x_{2002}(y^{2002,2002}_{2002,2002})$	Static wage	-.4***	-.004***	-.003**	-.021***
		(.09)	(.0012)	(.0013)	(.0099)
$x_{2002}(y^{2002,\mathbf{2011}}_{2002,\mathbf{2011}})$ $-x_{2002}(y^{2002,2002}_{2002,2002})$	Total wage	-.3 ***	-.003**	-.002	-.024**
		(.11)	(.0015)	(.0016)	(.0104)
$x_{\mathbf{2011}}(y^{2002,2002}_{2002,\mathbf{2011}})$ $-x_{2002}(y^{2002,2002}_{2002,2002})$	Static wage and tax transfer	-1.0***	-.006***	-.016***	.010
		(.15)	(.0022)	(.0024)	(.0116)
$x_{\mathbf{2011}}(y^{\mathbf{2011},\mathbf{2011}}_{2002,\mathbf{2011}})$ $-x_{2002}(y^{2002,2002}_{2002,2002})$	Total wage and tax transfer	-.7***	-.003	-.013***	.030**
		(.16)	(.0023)	(.0027)	(.0126)
$x_{2002}(y^{2002,2002}_{\mathbf{2011},2002})$ $-x_{2002}(y^{2002,2002}_{2002,2002})$	Population	1.9***	.028***	.018***	.141***
		(.51)	(.0102)	(.0057)	(.0273)

$x_a(y^{ce}_{bd})$: household net incomes according to the tax and transfer regulations of period a of gross incomes of the population of period b with wages according to labor market prices of period d with labor supply outcomes given the incentives of the tax and transfer regulations of period c and wages of period e.
Bootstrapped standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Source: Own calculation based on the SOEP and the STSM

The Gini index, which is sensitive to changes in the middle of the distribution, was 28.5 in 2002 and a ceteris paribus change to the tax and transfer system of 2011 would have reduced it by 0.4. The effect on the MLD, which is more sensitive to changes at the lower end of the distribution, is negative as well (-0.01). In contrast, the effect of policy reforms on the Q90/50 ratio is positive, so part of its increase can be explained with policy reforms. These three effects are highly significant. The Theil index, which puts the same weight on inequality at all parts of the distribution, remains unchanged.

The increased generosity of the transfer system has reduced inequality as measured by the MLD and the Gini, but when an equal weight is put on all parts of the distribution (Theil index), this is offset by the inequality increasing effect of tax reductions for high income earners. Lower tax rates have led to an increase in the Q90/50 ratio. The total tax and transfer effect shows that labor supply reactions to policy changes have offset the inequality reducing effect of policy changes as measured through the Gini. As the Gini index is sensitive to changes in the middle of the distribution, it was substantially affected by the labor supply adjustments depicted above in figure 1.4. Labor supply reactions have also led to an additional increase in the Q90/50 ratio.

A change in wage rates to 2011 levels with and without behavioral adjustments would have led to slight decreases in all reported inequality indices. Including behavioral effects renders the wage effect on the MLD insignificant. The last two lines show how changes in wage rates and policy changes interact. These effects are negative as measured by the Gini, MLD and Theil index. The point estimates of the total effects are closer to zero than the static effects. In contrast the total effect of policy and wage effects on the Q90/50 ratio is positive (0.03) and statistically significant. Thus, part of the increase in this measure is explained through wage and policy changes.

Overall, the decomposition show that policy changes from 2002 to 2011 have reduced inequality (Gini index and MLD) and this reduction was partly offset by labor supply reactions. The Q90/50 ratio was increased through policy changes. The effect of wage rate changes on income inequality was negative. Thus, the increase in income inequality was mostly due to changes in the population. The last line in Table 1.5 confirms this. It shows that the difference in the Gini index of the actual 2002 population and the 2011 population with tax and transfer system and conditional wages as in 2002 and labor supply simulated on this basis is 1.9.

1.5.4. Robustness Tests

Appendix A reports the results of two robustness tests. First, the order of the decomposition is changed, i.e., marginal effects are reported relative to the year 2011. This exercise demonstrates what would have happened if wage rates or the tax and transfer system had not changed since 2002 apart from adjustment to inflation. While the size of the coefficients changes, the main message remains the same: Wage changes and policy changes have both led to a decrease in the Gini and the MLD and the increase in income inequality is due to population changes.

Second, wage effects are re-estimated using categorical education variables instead of years of education for the wage regressions. The results are very close to the results obtained using a continuous variable.

1.6. Conclusion

This paper suggests a decomposition of changes in inequality into contributions from policy changes, changes in conditional wage rates, and population changes while considering both static and behavioral effects. In the application of the decomposition method to the increase in income inequality in Germany from 2002 to 2011, changes in the tax and transfer system are found to have had a small inequality reducing effect as measured by the Gini and MLD and a negligible effect on the Theil index. The reduction of the Gini was offset by labor supply reactions to the policy reforms. Tax reductions have increased the ratio between the 90th and the 50th income percentile.

The effect of changes in wage rates on income inequality was significantly negative as well. Behavioral reactions to wage changes are rather limited. Regarding both wage and tax-transfer effects, the impact of labor supply adjustments on the distribution is small.

This study confirms findings in Arntz et al. (2007) and Biewen and Juhasz (2012) regarding the distributional effects of the most important reforms of the German transfer system in recent years, which, contrary to common belief, have had a slight inequality reducing effect. The policy reforms undertaken in the analyzed time span, an increase of the generosity of the transfer system and a tax reduction, have had a negative impact on the government budget. Future research should study the distributional effects of the funding of these policy measures.

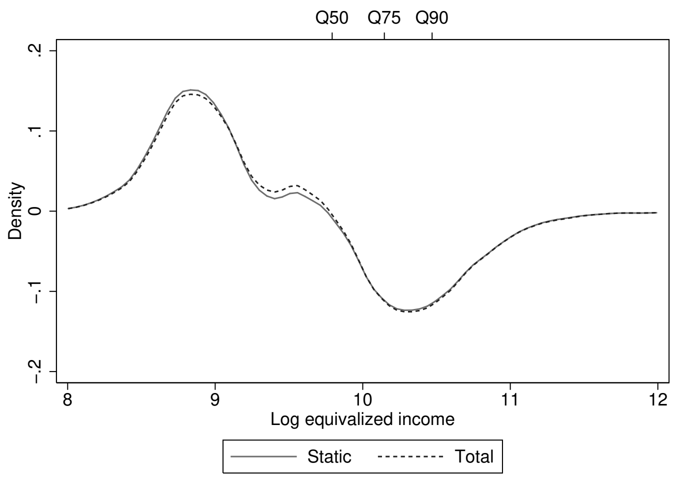
The decomposition exercise shows that most of the change in inequality cannot be explained with policy and wage rate changes, but is due to changes in the population. These include changes in household structure and the distribution of non-labor income (Biewen and Juhasz, 2012; Peichl et al., 2012; Biewen et al., 2016) as well as changes in employment patterns unrelated to changes in wage rates and the tax and transfer system. Further research into the driving forces of the population changes is warranted.

Appendix

A. Robustness

Base year 2011

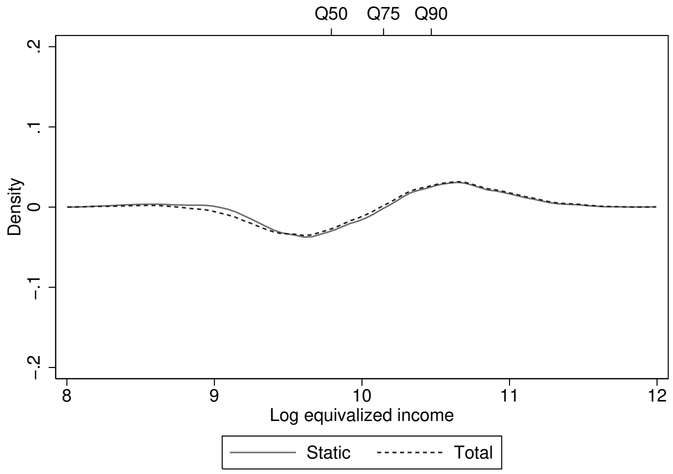
Figure A.1.: Policy Effect – Base Year 2011



Difference between actual distribution in 2011 and counterfactual distributions applying the 2002 tax and transfer system to the2011 population (static, $x_{\mathbf{2002}}(y^{2011,2011}_{2011,2011}) - x_{2011}(y^{2011,2011}_{2011,2011})$) and additionally simulating labor supply reactions (total, $x_{\mathbf{2002}}(y^{\mathbf{2002},2011}_{2011,2011}) - x_{2011}(y^{2011,2011}_{2011,2011})$). Source: Own calculation based on the SOEP and the STSM

Figure A.1 shows static and behavioral counterfactual distributions using the year 2011 as base and applying the tax and transfer system of 2002. It is the “inverse” of Figure 1.4, which displays ex-ante effects of the policy reforms, so the interaction with different populations of 2002 and 2011 does not change the direction of the distributional effects of the policy reforms. Applying the tax and transfer system of 2002 to the population of 2011 leads to an increase in density at the bottom of the distribution due to less generous transfers in 2002. The higher top marginal tax rate of 2002 leads to a decrease in density at the top of the distribution and the labor supply effects are similar to those depicted in Figure 1.4. Behavioral adjustments lead to an increase in density at log equivalized net income of 9.5 and a slight decrease in density at lower parts relative to the solid line. This can be interpreted as the effect of higher labor supply in the 2002 counterfactual compared to the 2011 distribution. This is caused by the decrease in labor supply incentives for low income households in 2011 due to the increase in the generosity of transfers.

Figure A.2.: Wage Effect – Base Year 2011



Difference between actual distribution in 2011 and counterfactual distributions applying 2002 conditional wage rates to the base year population (static, $x_{2011}(y^{2011,2011}_{2011,\boldsymbol{2002}}) - x_{2011}(y^{2011,2011}_{2011,2011})$*) and additionally simulating labor supply reactions (total,* $x_{2011}(y^{\boldsymbol{2002},2011}_{2011,\boldsymbol{2002}}) - x_{2011}(y^{2011,2011}_{2011,2011})$*). Source: Own calculation based on the SOEP and the STSM*

Applying 2002 wage rates to the 2011 population (Figure A.2) leads to an increase in density at the right and a decrease at the left of the distribution which might be attributable to the higher education premium of 2002 documented in Tables 1.2 and 1.3 in subsection 1.4.2. Labor supply effects are small, but the dynamic counterfactual distribution has a slightly higher density in the upper part and lower density at the lower part than the static counterfactual distribution indicating that some former low income households have slightly higher labor supply in the counterfactual.

A comparison of the first line of Table A.1, which displays actual and counterfactual inequality measures with the year 2011 as base, with the first line of Table 1.5 shows the overall increases in the four inequality measures. Note that the signs of the effects in the following lines of Table A.1 have the opposite meaning than those of Table 1.5. For instance, a positive tax and transfer effect with base 2011 means that applying the 2002 tax and transfer system to the 2011 sample increases inequality relative to the status quo in 2011. In other words, policy changes from 2002 to 2011 have led to a *decrease* in income inequality.

While the base year changes the magnitude of the results, the results for the tax and transfer system are qualitatively similar:[13] Changes in the tax and transfer system have led to a decrease in inequality as measured by the Gini index and the MLD, the effect on

[13]It is common that the order of the decomposition has a strong influence on the estimated effect of a single factor on the change in inequality. For instance, Bargain and Callan (2010) decompose the

Table A.1.: Decomposition with Base 2011

Marginal Effect		Gini	Theil	MLD	Q90/50
$x_{2011}(y^{2011,2011}_{2011,2011})$	Status quo	29.5***	.168***	.148***	1.93***
		(.49)	(.0136)	(.006)	(.0232)
$x_{\mathbf{2002}}(y^{2011,2011}_{2011,2011})$	Static tax and transfer	.3***	-.001	.008***	-.036***
$-x_{2011}(y^{2011,2011}_{2011,2011})$		(.11)	(.0030)	(.0018)	(.0091)
$x_{\mathbf{2002}}(y^{\mathbf{2002},2011}_{2011,2011})$	Total tax and transfer	.1	-.003	.005**	-.04***
$-x_{2011}(y^{2011,2011}_{2011,2011})$		(.12)	(.0030)	(.0022)	(.0094)
$x_{2011}(y^{2011,2011}_{2011,\mathbf{2002}})$	Static wage	1.0***	.009***	.009***	.064***
$-x_{2011}(y^{2011,2011}_{2011,2011})$		(.07)	(.0016)	(.0012)	(.0132)
$x_{2011}(y^{\mathbf{2002},2011}_{2011,\mathbf{2002}})$	Total wage	1.1***	.01***	.01***	.066***
$-x_{2011}(y^{2011,2011}_{2011,2011})$		(.1)	(.0023)	(.0014)	(.013)
$x_{\mathbf{2002}}(y^{2011,2011}_{2011,\mathbf{2002}})$	Static wage and tax transfer	1.2***	.007*	.016***	.022*
$-x_{2011}(y^{2011,2011}_{2011,2011})$		(.13)	(.0036)	(.0022)	(.0121)
$x_{\mathbf{2002}}(y^{\mathbf{2002},\mathbf{2002}}_{2011,\mathbf{2002}})$	Total wage and tax transfer	.9***	.004	.013***	.026**
$-x_{2011}(y^{2011,2011}_{2011,2011})$		(.16)	(.0041)	(.0025)	(.0131)
$x_{2011}(y^{2011,2011}_{\mathbf{2002},2011})$	Population	-1.7***	-.027***	-.018***	-.085***
$-x_{2011}(y^{2011,2011}_{2011,2011})$		(.53)	(.0120)	(.0059)	(.0313)

$x_a(y^{ce}_{bd})$: household net incomes according to the tax and transfer regulations of period a of gross incomes of the population of period b with wages according to labor market prices of period d with labor supply outcomes given the incentives of the tax and transfer regulations of period c and wages of period e.
Bootstrapped standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Source: Own calculation based on the SOEP and the STSM

the Theil is insignificant. In contrast, the effect of policy changes on the Q90/50 ratio is positive and highly significant. When taking labor supply responses into account, the effect on the Gini becomes insignificant.

As is the case for the the decomposition using the year 2002 as base, changes in wage rates from 2002 to 2011 have had a significant decreasing impact on all four indices.

The interaction of policy and wage effects are shown in the second and third line from the bottom in table A.1. Similarly to the decomposition with base 2002, the overall effects on the Gini, MLD and Theil imply that wage and policy changes together have had a negative impact on income inequality. In contrast to the decomposition with base 2002, the total effect on the Q90/50 ratio is inequality decreasing as well as it is dominated by the wage effect, which is stronger in the base 2011 decomposition. The last line shows the effect of population changes. As is the case relative to the base year 2002, population changes are the main drivers of the increase in income inequality.

Categorical education variable

To test whether the results for the marginal effects of conditional wage rate changes are sensitive to functional form assumptions regarding education, this subsection reports results obtained using categorical education variables. Tables A.2 and A.3 show wage regressions for the two years. While the point estimates of the high and medium education dummies increased slightly for men, they decreased for women. However, using categorical variables makes the estimates substantially less precise, the effect of medium education is even insignificant for East German women in 2002.

change in inequality in Ireland from 1994 to 2000. When using 1994 as base year, the effect of policy changes on the Gini coefficient is 1.4. When using 2000 as base year, the effect is only 0.7.

Table A.2.: Wage Regression 2002 with Categorical Education Variables

	(1) Men East	(2) Women East	(3) Men West	(4) Women West
Ln(Hourly Wage)				
Medium education	-0.0105	0.183*	0.142***	0.209***
	(0.0672)	(0.0854)	(0.0247)	(0.0333)
High education	0.449***	0.551***	0.595***	0.724***
	(0.0714)	(0.0943)	(0.0275)	(0.0422)
Years not Worked	-0.130***	-0.0995***	-0.0902***	-0.0313***
	(0.0198)	(0.0157)	(0.00912)	(0.00509)
Experience	0.0478***	0.0653***	0.0665***	0.0605***
	(0.00672)	(0.00861)	(0.00329)	(0.00452)
Experience^2/100	-0.0917***	-0.138***	-0.121***	-0.119***
	(0.0168)	(0.0231)	(0.00813)	(0.0123)
Constant	1.991***	1.586***	2.029***	1.766***
	(0.0977)	(0.139)	(0.0402)	(0.0665)
Mills	-0.0487	-0.0421	-0.0154	0.0241
lambda	(0.0721)	(0.0898)	(0.0323)	(0.0466)
N	2616	2899	7586	8253

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

High education: university, *medium education:* vocational training or high school, *base category:* lower

Source: Own calculation based on the SOEP

Table A.3.: Wage Regression 2011 with Categorical Education Variables

	(1) Men East	(2) Women East	(3) Men West	(4) women west
Ln(Hourly Wage)				
Medium education	0.149*	-0.0677	0.184***	0.162***
	(0.0744)	(0.0948)	(0.0313)	(0.0349)
High Education	0.628***	0.350***	0.617***	0.606***
	(0.0778)	(0.0994)	(0.0335)	(0.0411)
Years not Worked	-0.128***	-0.108***	-0.139***	-0.0373***
	(0.0159)	(0.0134)	(0.0102)	(0.00551)
Experience	0.0709***	0.0481***	0.0724***	0.0531***
	(0.00712)	(0.00753)	(0.00396)	(0.00428)
Experience^2/100	-0.155***	-0.0930***	-0.136***	-0.0928***
	(0.0185)	(0.0192)	(0.0100)	(0.0111)
Constant	1.666***	1.925***	1.824***	1.773***
	(0.0937)	(0.130)	(0.0473)	(0.0661)
Mills	0.0373	-0.0621	0.0709	-0.0447
lambda	(0.0719)	(0.0745)	(0.0399)	(0.0459)
N	2419	2695	6898	7825

Standard errors in parentheses
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
High education: university, *medium education:* vocational training or high school, *base category:* lower
Source: Own calculation based on the SOEP

Table A.4 shows the marginal static and total effects of wage changes as well as the joint effect of wage and policy changes obtained using categorical education variables. These effects are very close to those obtained using years of education. The result that changes in conditional wage rates do not explain the increase in income inequality is robust to the use of different education variables.

Table A.4.: Wage Effects with Base 2002 and Categorical Education Variables

Marginal effect		Gini	Theil	MLD	Q90/50
$x_{2002}(y^{2002,2002}_{2002,\mathbf{2011}})$	Static wage	-.3***	-.002*	-.002*	-.001
$-x_{2002}(y^{2002,2002}_{2002,2002})$		.(09)	(.0013)	(.0013)	(.0104)
$x_{2002}(y^{2002,\mathbf{2011}}_{2002,\mathbf{2011}})$	Total wage	-.2*	-.001	-.001	.003
$-x_{2002}(y^{2002,2002}_{2002,2002})$		(.10)	(.0013)	(.0016)	(.0110)
$x_{\mathbf{2011}}(y^{2002,2002}_{2002,\mathbf{2011}})$	Static wage and tax transfer	-.9***	-.005**	-.014***	.033***
$-x_{2002}(y^{2002,2002}_{2002,2002})$		(.15)	(.0023)	(.0025)	(.0130)
$x_{\mathbf{2011}}(y^{\mathbf{2011},\mathbf{2011}}_{2002,\mathbf{2011}})$	Total wage and tax transfer	-.8***	-.003	-.013***	.038***
$-x_{2002}(y^{2002,2002}_{2002,2002})$		(.16)	(.0023)	(.0026)	(.0136)

$x_a(y^{ce}_{bd})$: household net incomes according to the tax and transfer regulations of period a of gross incomes of the population of period b with wages according to labor market prices of period d with labor supply outcomes given the incentives of the tax and transfer regulations of period c and wages of period e.
Bootstrapped standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Source: Own calculation based on the SOEP and the STSM

B. Estimation Results for Labor Supply Model

Table B.5.: Estimation Results for Labor Supply Model 2002

Variables	Flexible Couples	Women with Inflexible Spouse	Men with Inflexible Spouse	Single Men	Single Women
Log Net Income	-4.541*	-4.728	7.295	1.063	-2.690
	(2.054)	(4.142)	(6.018)	(2.507)	(2.373)
Log Net Income^2	0.575***	0.597***	-0.0702	0.266***	0.285***
	(0.0582)	(0.179)	(0.242)	(0.0787)	(0.0664)
Log Net Income × East	-1.905	-1.030	-4.428	1.799	-1.504
	(1.900)	(7.409)	(7.005)	(1.330)	(2.548)
Log Net Income^2 × East	0.105	-0.0286	0.244	-0.107	0.120
	(0.107)	(0.388)	(0.361)	(0.0894)	(0.158)
Log Net Income × German Female	0.674*	1.762	-0.324		0.531
	(0.273)	(0.935)	(0.341)		(0.405)
Log Leisure Female	128.3***	118.5***			124.0***
	(6.900)	(9.573)			(10.49)
Log Net Income × Log Leisure Female	-0.447*	-1.027***			-0.403
	(0.191)	(0.250)			(0.384)
Log Leisure Female ^{2}	-14.65***	-12.37***			-14.43***
	(0.752)	(1.111)			(1.096)
Log Leisure Female × German Female	-0.331	0.0445			1.118
	(0.345)	(0.640)			(0.621)
Age Female *x* Log Leisure Female	-0.372***	-0.525***			-0.397***
	(0.0621)	(0.0813)			(0.0771)
Age ^{2} × Log Leisure Female	0.588***	0.820***			0.621***
	(0.075)	(0.094)			(0.089)
Log Leisure Female × Disability I	-0.134	-0.274			0.263
	(0.323)	(0.470)			(0.542)
Log Leisure Female × Disability II	1.036	1.086			1.196
	(0.574)	(0.780)			(0.885)
Log Leisure Female × East	-8.175***	-2.578***			0.755
	(1.717)	(0.490)			(0.561)
Log Leisure Female × Children Under 3 Years	4.892***	3.897***			5.962***
	(0.267)	(0.438)			(0.738)
Log Leisure Female × Children 7 to 16 Years	2.426***	2.011***			1.370***
	(0.158)	(0.275)			(0.315)
Log Leisure Female × Children 4 to 6 Years	2.140***	2.238***			2.838***
	(0.225)	(0.409)			(0.493)
Log Leisure Female × Children over 17 Years	0.468**	0.466			0.245
	(0.160)	(0.259)			(0.349)
Female Part Time I	-2.114***	-2.498***			-3.053***
	(0.0793)	(0.130)			(0.183)
Female Part Time II	-2.126***	-2.134***			-2.572***
	(0.0971)	(0.149)			(0.146)

Table continued on next page.

Variables	Flexible Couples	Women with Inflexible Spouse	Men with Inflexible Spouse	Single Men	Single Women
Log Net Income × German Male	-1.459**	-1.161*	-0.253	-0.193	
	(0.475)	(0.488)	(0.867)	(0.410)	
Log Leisure Male × Log Net Income	-0.492**		-1.073**	-1.068**	
	(0.172)		(0.400)	(0.362)	
Log Leisure Male	31.44***		43.65***	38.21***	
	(3.308)		(6.227)	(6.751)	
Log Leisure Male^2	-2.797***		-3.472***	-3.585***	
	(0.182)		(0.398)	(0.529)	
Log Leisure × German Male	-1.086**		-0.694	-1.281	
	(0.406)		(0.763)	(0.769)	
Log Leisure Male × Age Male	-0.231***		-0.315***	0.0216	
	(0.0469)		(0.0782)	(0.0787)	
Log Leisure Male × Age Male^2	0.326***		0.402***	0.0518	
	(0.053)		(0.087)	(0.091)	
Log Leisure Male × Disability I	0.792***		0.973*	0.312	
	(0.237)		(0.489)	(0.529)	
Log Leisure Male × Disability II	1.537***		3.493***	1.721*	
	(0.409)		(1.032)	(0.683)	
Log Leisure Male × East	-6.092***		1.041	0.660	
	(1.815)		(0.533)	(0.549)	
Male Over Time	-1.683***		-1.307***	-1.522***	
	(0.0800)		(0.164)	(0.202)	
Male Part Time	-2.803***		-2.463***	-2.428***	
	(0.112)		(0.199)	(0.195)	
Log Leisure Male × Log Leisure Female × German Male	-0.118				
	(0.103)				
Log Leisure Male × Log Leisure Female	0.137				
	(0.288)				
Log Leisure Male × Log Leisure Female quad × East	1.648***				
	(0.456)				
Observations	117395	10031	4130	3960	6854
Pseudo R^2	0.29	0.27	0.35	0.41	0.28
Log-Likelihood	-9498	-2189	-871	-754	-1474
Uncompensated own-wage elasticities					
Male	0.15		0.12	0.24	
Female	0.27	0.57			0.23
Uncompensated cross-wage elasticities					
Male	-0.03		0.00		
Female	-0.01	-0.00			

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Source: Own calculations based on the SOEP and the STSM

Table B.6.: Estimation Results for Labor Supply Model 2011.

Variables	Flexible Couples	Women with Inflexible Spouse	Men with Inflexible Spouse	Single Men	Single Women
Log Net Income	6.056**	-20.63**	-5.577	0.996	0.697
	(2.126)	(6.515)	(5.947)	(2.681)	(2.729)
Log Net Income^2	0.239***	1.101***	0.365	0.254**	0.258***
	(0.0517)	(0.269)	(0.221)	(0.0776)	(0.0704)
Log Net Income × East	-1.908	-10.53	-1.301	-0.893	-2.260
	(1.582)	(10.23)	(9.729)	(1.766)	(2.025)
Log Net Income)^2 × East	0.126	0.540	0.0750	0.0770	0.145
	(0.0928)	(0.515)	(0.502)	(0.116)	(0.125)
Log Net Income × German Female	0.355	0.662	-0.260		-1.440
	(0.486)	(1.140)	(0.388)		(1.111)
Log Leisure Female	118.1***	107.3***			121.1***
	(7.369)	(9.449)			(9.709)
Log Net Income × Log Leisure Female	-0.975***	-0.0934			-0.616
	(0.217)	(0.315)			(0.357)
Log Leisure Female ^{2}	-12.67***	-11.96***			-13.47***
	(0.777)	(1.051)			(0.992)
Log Leisure Female × German Female	-0.543	-1.041			-2.504*
	(0.411)	(0.691)			(1.021)
Age Female x Log Leisure Female	-0.209**	-0.593***			-0.324***
	(0.0717)	(0.0839)			(0.0719)
Age ^{2} Female x Log Leisure Female	0.384***	0.840***			0.499***
	(0.083)	(0.093)			(0.081)
Log Leisure Female × Disability I	0.155	0.968*			0.910*
	(0.347)	(0.444)			(0.409)
Log Leisure Female × Disability I	0.689	1.665*			1.363*
	(0.669)	(0.811)			(0.615)
Log Leisure Female × East	-12.85***	-1.519**			0.0807
	(2.214)	(0.464)			(0.459)
Log Leisure Female × Children Under 3 Years	4.763***	4.267***			5.033***
	(0.301)	(0.423)			(0.733)
Log Leisure Female × Children 7 to 16 Years	2.005***	1.948***			1.949***
	(0.189)	(0.283)			(0.279)
Log Leisure Female × Children 4 to 6 Years	2.207***	2.328***			2.288***
	(0.272)	(0.437)			(0.498)
Log Leisure Female × Children over 17 Years	0.969***	0.702**			0.189
	(0.191)	(0.269)			(0.299)
Female Part Time I1	-1.614***	-2.070***			-2.888***
	(0.0857)	(0.123)			(0.160)
Female Part Time II	-1.605***	-1.778***			-2.279***
	(0.102)	(0.140)			(0.131)

Table continued on next page.

Variables	Flexible Couples	Women with Inflexible Spouse	Men with Inflexible Spouse	Single Men	Single Women
Log Net Income × German Male	0.650	-0.570	1.872	0.436	
	(0.693)	(0.388)	(1.939)	(0.602)	
Log Leisure Male × Log Net Income	-1.397***		-0.652	-1.143**	
	(0.218)		(0.420)	(0.391)	
Log Leisure Male	57.74***		37.55***	30.19***	
	(4.205)		(6.469)	(6.835)	
Log Leisure Male^2	-4.360***		-3.589***	-2.809***	
	(0.236)		(0.393)	(0.476)	
Log Leisure × German Male	-0.158		0.472	1.418	
	(0.515)		(1.094)	(0.853)	
Log Leisure Male × Age Male	-0.334***		-0.285**	0.0455	
	(0.0632)		(0.0933)	(0.0746)	
Log Leisure Male × Age Male^2	0.431***		0.369***	-0.0668	
	(0.070)		(0.104)	(0.088)	
Log Leisure Male × Disability I	0.750**		1.489***	1.398***	
	(0.270)		(0.431)	(0.422)	
Log Leisure Male × Disability II	1.488**		2.131*	1.487*	
	(0.516)		(0.902)	(0.597)	
Log Leisure Male × East	-11.19***		0.275	1.005*	
	(2.349)		(0.590)	(0.493)	
Male Over Time	-1.647***		-1.535***	-1.805***	
	(0.0928)		(0.165)	(0.193)	
Male Part Time	-2.592***		-2.247***	-2.405***	
	(0.124)		(0.216)	(0.185)	
Log Leisure Male × Log Leisure Female × German Male	-0.0430				
	(0.127)				
Log Leisure Male × Log Leisure Female	-0.936**				
	(0.360)				
Log Leisure Male × Log Leisure Female quad × East	2.927***				
	(0.580)				
Observations	87236	9690	3749	4212	8090
Log-Likelihood	-7180	-2349	-782	-847	-1910
Pseudo R^2	0.27	0.19	0.35	0.38	0.21
Uncompensated own-wage elasticities					
Male	0.11		0.05	0.30	
Female	0.28	0.04			0.11
Uncompensated cross-wage elasticities					
Male	0.01		0.00		
Female	-0.03	-0.01			

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Source: Own calculations based on the SOEP and the STSM

2. Optimal Taxation Under Different Concepts of Justness [1]

2.1. Introduction

Optimal taxation is only relevant if it is able to capture the actual aims of the social planner. Therefore, we extend a standard optimal taxation model to reconcile it with tax transfer practices in 2015. The standard approach in the welfarist optimal taxation literature is to assume that social weights decrease with income (e.g., Saez, 2001, 2002; Blundell et al., 2009) because this pattern lies within the bounds confined by the two extreme cases of Rawlsian and Benthamite objective functions. Intuitively, the hypothesis of decreasing welfarist weights expresses the idea that the social planner values an increase of net income of the poor by one Euro more than an increase of net income of higher income groups by one Euro. Saez and Stantcheva (2016) describe it as one of their two polar cases of interest. In contrast, tax transfer systems in many countries are only optimal if the social planner had chosen weights in a non-decreasing way.[2] As we show, a major reason for this lies in high transfer withdrawal rates for the working poor.[3]

In this paper, we generalize the optimal taxation framework by Saez (2002) to divert from welfarism. In an exercise of positive optimal taxation, we calculate the social weights under different concepts of justness. First, we apply the standard welfarist concept. Second, we apply the concept of minimum sacrifice and, third, a concept based on subjectively just net incomes. This third concept utilizes a novel question in the German Socio-Economic Panel (SOEP): respondents state what net income they would consider just.[4] We term the latter concept *subjective justness*. We find that the minimum absolute sacrifice principle is in line with decreasing social weights.

Our paper is related to studies analyzing optimal taxation when the preferences of the social planner and individuals differ (Blomquist and Micheletto, 2006; Kanbur et al., 2006). Gerritsen (2016) derives the optimal tax schedule for a government that optimizes a weighted sum of subjective well-being, while individuals maximize utility

[1]This chapter is based on Jessen et al. (2017a).

[2]Appendix A reviews a number of studies with this finding.

[3]Lockwood (2016) shows that under present bias and with job search, optimal marginal tax rates are even lower than conventionally calculated. This might be especially relevant for marginal tax rates for the working poor.

[4]We use respondents who consider their current gross income as just. Thus, just net incomes can be interpreted conditional on given gross incomes.

instead of well-being. He expresses the tax-schedule in terms of sufficient statistics in the continuous framework. In contrast, we use the discrete sufficient statistics framework that allows for labor supply adjustments at the intensive and at the extensive margin following Saez (2002).

The first main contribution of our paper is a generalization of the Saez (2002) model to non-welfarist aims of the social planner. To our knowledge, we are the first to derive the general optimal taxation schedule in this framework. In a recent study, Saez and Stantcheva (2016) propose generalized marginal welfare weights that may depend on characteristics that do not enter utility.[5] In contrast, in our approach, the social planner maximizes an objective function that allows for non-welfarist concepts of justice. The approach in our paper offers the advantage that we can directly quantify the value the social planner puts on a marginal improvement in a specific justness criterion for a given group compared to other groups. Thus, we can show which criterion is in line with social weights that decrease with income.

The second main contribution is the operationalization of two specific ideas of justice: minimum sacrifice and subjective justness. Minimum sacrifice is related to the equal sacrifice principle (see Mill, 1871; Musgrave and Musgrave, 1973; Richter, 1983; Young, 1988), which stipulates that all individuals should suffer the same 'sacrifice' through taxes. The sacrifice is usually defined as the burden of taxes in terms of utility. Evidence that the equal sacrifice concept is likely to capture the preferences of the majority is only documented for the U.S: Weinzierl (2014) shows in a survey for the U.S. that around 60 percent preferred the equal sacrifice tax schedule to a welfarist optimal tax schedule. While equal sacrifice equalizes the sacrifice due to taxes across the population, minimum sacrifice minimizes the (weigthed) sum of these utility losses. The concept of minimum sacrifice is very close to the libertarian concept studied in Saez and Stantcheva (2016).[6]

The second approach, subjective just income, is novel as we use new questions from the SOEP to measure the perceived justness of gross and net incomes. These survey questions are representative for the working population in Germany. To the best of our knowledge, we are the first to use such a rich assessment of subjective preferences for just taxation in an optimal taxation framework. We analyze the social weights implied by subjective justness for subgroups of the population that might adhere to different concepts of justness: females and males, East Germans and West Germans who lived under different political systems for more than a generation, as well as supporters of different political parties. The third main contribution is the application to the German tax and transfer system, as of 2015, for which we estimate the labor supply elasticities using microsimulation and a structural labor supply model.

[5] Similar to Saez and Stantcheva (2016), we take society's preferences as given and do not analyze how they could arise through the political process.

[6] Saez and Stantcheva (2016) allow for *welfarist* weights to increase with the amount of taxes paid. Thus decreasing taxes for those with a high tax burden is a high priority for the social planner.

Our main result is that the concept of minimum sacrifice is in line with positive, declining social weights. The explanation for this finding is that the marginal sacrifice increases with the amount of taxes paid and the working poor pay only a low amount of taxes. Although the costs of redistributing a Euro to this group are relatively small, the reduction in sacrifice is small too. In contrast, the increase in utility is high in the welfarist case. A second finding is a confirmation of previous studies: the welfarist approach implies very low weights for the working poor under the 2015 German tax and transfer system. Finally, we find that the German tax and transfer system is roughly in line with the minimization of absolute deviations from subjective just net incomes and decreasing social weights.

The next section introduces our optimal taxation model for different concepts of justness, section 2.3 describes how we calculate actual and just incomes as well as how we estimate extensive and intensive labor supply elasticities for Germany. In section 2.4, we describe the resulting weights for different concepts of justness, while section 2.5 concludes.

2.2. A Model of Optimal Taxation for Different Concepts of Justness

We develop a generalization of the model in Saez (2002). As shown, the difference between the Saez (2002) model and our generalization is that in Saez (2002) the social planner maximizes the weighted sum of utility. The main advantage of our approach is that we allow for the social planner to maximize the weighted sum of 'justness functions' f_i. These functions can depend on various variables and incorporate different concepts of justness. We show that welfarism as in Saez (2002) is a special case.

2.2.1. The General Framework

We generalize the canonical model by Saez (2002), which combines the pioneering work by Mirrlees (1971) and Diamond (1980), beyond utilitarian social welfare functions. See Appendix B for a formal derivation. Net income equals consumption and is given by $c_i = y_i - T_i$, where $i = 0, ..., I$ income groups are defined through gross income y_i.[7] T_i denotes total taxes paid by the individual to finance a public good G. Each income group has the share h_i of the total population. These shares are endogenous as individuals adjust their labor supply to the tax-transfer system. The social planer chooses tax liabilities

[7]The number of income groups is assumed to be fixed. In the empirical application, we define groups $1, .., I$ as quintiles of the gross income distribution. Bargain et al. (2014) show that changing the cut-off points between groups does not affect the results substantially.

T_i to optimize a weighted sum L based on individual justness functions f_i (described in subsection 2.2.2), which may depend on c_i or on other factors that do not enter the utility function of individuals. The optimization is subject to the government budget constraint:

$$L = \sum_{i=0}^{I} \mu_i h_i f_i \quad \text{s.t.} \quad \sum_{i=0}^{I} h_i T_i = G, \tag{2.1}$$

where μ_i are the primitive social weights associated with the income level of group i.[8] Together with the Lagrange multiplier λ, they define the explicit weights $e_i \equiv \frac{\mu_i}{\lambda}$, which we focus on in this study.[9] It is important to note that our approach does not require explicit utility functions but nests the welfarist approach as a special case. Following Saez (2002), we consider the benchmark case with no income effects, where $\sum_{i=0}^{I} \partial h_j / \partial c_i = 0$. Summing the first order conditions (equation (2.14) in the appendix) over all $i = 0, \ldots, I$ we obtain the normalization of weights such that:[10]

$$\sum_{i=0}^{I} h_i e_i \frac{\partial f_i}{\partial c_i} = 1. \tag{2.2}$$

Following Saez (2002), we assume that labor supply adjustment is restricted to intensive changes to "neighbor" income groups and extensive changes out of the labor force. Thus h_i depends only on differences in after-tax income between "neighbor groups" $(c_{i+1} - c_i,\ c_i - c_{i-1})$ and differences between group i and the non-working group $(c_i - c_0)$. The intensive mobility elasticity is

$$\zeta_i = \frac{c_i - c_{i-1}}{h_i} \frac{\partial h_i}{\partial (c_i - c_{i-1})} \tag{2.3}$$

and the extensive elasticity is given by

$$\eta_i = \frac{c_i - c_0}{h_i} \frac{\partial h_i}{\partial (c_i - c_0)}. \tag{2.4}$$

[8] Positive values of μ_i imply that the social planner aims at 'improving' f_i.

[9] For welfarist applications it is common in the literature to report implicit weights, $g_i \equiv e_i \frac{\partial f_i}{\partial c_i}$, which offers the advantage to remain agnostic about utility functions. In the standard welfarist approach, implicit weights are defined as the product of the explicit weights and the marginal utility of consumption, $g_i \equiv e_i \frac{\partial u(c_{i*}, i^*)}{\partial c_i}$. We calculate *relative* social welfare weights e_i / e_0 as in Blundell et al. (2009). As will be made clear, relative explicit social welfare weights equal relative implicit weights under the welfarist approach with neither income effects nor preference heterogeneity. Thus, social weights of all approaches are comparable.

[10] In the welfarist approach, this normalization reduces to the corresponding equation in Saez (2002): $\sum_{i=0}^{I} h_i g_i = 1$.

The main result is that the optimal tax formula for group i expressed in terms of the participation elasticities η_j and the intensive elasticity ζ_i is

$$\begin{aligned}\frac{T_i - T_{i-1}}{c_i - c_{i-1}} &= \frac{1}{\zeta_i h_i}\left\{\sum_{j=i}^{I}\left[1 - e_j\frac{\partial f_j}{\partial c_j} - \eta_j\frac{T_j - T_0}{c_j - c_0}\right]h_j\right. \\ &\quad \left. -(e_i f_i - e_{i-1}f_{i-1})\zeta_i\frac{h_i}{c_i - c_{i-1}} - \sum_{j=i}^{I}\eta_j\frac{e_j f_j - e_0 f_0}{c_j - c_0}h_j\right\}.\end{aligned} \tag{2.5}$$

Multiplying equation (2.5) with $\zeta_i h_i dT$ clarifies the intuition of the optimal tax formula. Consider an increase of dT in all T_j for income groups $j = i, i+1, ...I$. The left hand side shows the negative effect on tax revenue due to individuals switching from job i to $i-1$.[11] At the optimum, this must equal the mechanical tax gains, which are valued at $\sum_{j=i}^{I}\left(1 - e_j\frac{\partial f_j}{\partial c_j}\right)$, minus tax losses due to individuals moving to group 0, $\sum_{j=i}^{I}\eta_j\frac{T_j - T_0}{c_j - c_0}h_j$, and the effect on the objective function of individuals moving into different jobs due to the tax increase, captured by the second line of the equation. The first term in the second line captures the effect of individuals moving from group i to $i-1$ and the second term captures the effect of individuals adjusting at the extensive margin.

The main difference between equation (2.17) and the mixed model in Saez (2002) is the second line, which does not appear in Saez (2002). While in the welfarist approach, changes due to behavioral responses drop out due to the envelope theorem, in our approach we consider changes in the justness function, which may change non-negligibly with a change in behavior. The second difference is that we replace the implicit weights $g_j = e_j\frac{\partial u(c_{j*}, j^*)}{\partial c_j}$ with $e_j\frac{\partial f_j}{\partial c_j}$. The optimal tax schedule in Saez (2002) depends on elasticities and weights g_j, whereas in the generalized model, they additionally depend on the justness functions f_j.

The system of equations defining the optimal tax schedule consists of I equations like (2.5) and equation (2.2). In our application, we use the 2015 German tax system, i.e. we calculate the actual tax liability T_i of each income group, and solve for $e_1, ..., e_I$. Alternatively, one could assume justness weights and calculate the optimal tax schedule that maximizes equation (2.1).

2.2.2. Operationalization of Justness Concepts

The key advantage of our approach is that the justness function can be defined very generally, thus allowing us to capture a broader set of concepts of justness than the standard approach. In principle, the function can depend on individual and aggregate

[11] Due to the assumption of no income effects and because the differences in net income between groups $i, i+1, ...I$ are unchanged, groups $i+1, i+2, ...I$ will only adjust at the extensive margin.

variables. The variables included in the justness function determine the dimensions along which the social planner considers a redistribution to be just. These variables do not need to be included in the utility function. For instance, utility is defined on after-tax income c_i and the choice of income group i in the standard welfarist approach. Our approach allows considering non-welfarist concepts of justness that rely, e.g., on before-tax income y_i.

Our approach nests the welfarist approach with quasilinear preferences.[12] This special case is given if

$$f_i = u_i = v(i) + b \times c_i, \tag{2.6}$$

where $v(i)$ denotes the disutility of work in income group i and $b \times c_i$ is the linear utility of consumption. By introducing a general justness function f_i, we may operationalize other moral judgments that depend directly on variables that do not enter the utility function as in the concept of *minimum sacrifice*. We operationalize two forms of minimum sacrifice: Minimum absolute sacrifice based on the absolute tax liability and relative minimum sacrifice based on the tax liability relative to the net income.

Sacrifice is defined as the difference in utility derived from net income and the hypothetical utility derived from gross income, i.e., if there were no taxes:

$$\text{Sacrifice} = u(y_i) - u(c_i) \tag{2.7}$$

We focus on the case of quasi-linear preferences, see equation (2.6), so the sacrifice simplifies to $y_i - c_i$. We formulate a loss function that captures the penalty to the objective function of the social planner if individuals pay taxes, i.e., if there is a positive sacrifice. This loss function is the justness function associated with minimum sacrifice.

In the case of minimum *absolute* sacrifice the loss that captures deviations of c_i from gross income y_i is determined by the parameters γ, α, and δ:[13]

$$\begin{aligned} f_i &= -(y_i - c_i)^{\gamma} \text{ if } y_i > c_i, \\ f_i &= \alpha(c_i - y_i)^{\delta} \text{ if } c_i > y_i, \\ &\gamma > 1, 0 \leq \alpha \leq 1, \delta \leq 1. \end{aligned} \tag{2.8}$$

The first line gives the penalty of paid taxes. $\gamma > 1$ implies that the penalty increases more than proportionally with the amount of taxes paid. The second line captures the gains if individuals receive transfers. If δ is smaller than one, the marginal benefits of transfers are decreasing. The parameter α scales the gains relative to sacrifices. A positive

[12]The absence of income effects, i.e. the assumption of quasi-linear preferences, is common in the optimal taxation literature following Saez (2002). In this case relative explicit welfare weights equal relative implicit welfare weights: $\frac{\partial f_j}{\partial c_j} = b$ cancels out, i.e, $\frac{g_i}{g_0} = \frac{e_i}{e_0} \frac{\partial u(c_{i*}, i^*)/\partial c_i}{\partial u(c_{0*}, 0^*)/\partial c_0} = \frac{e_i}{e_0}$.

[13]We leave for future research empirical identification of penalty functions. Note however, that this is only possible if the social weights are known.

α guarantees Pareto optimality if weights e_i are positive, as it guarantees that f_i increases with increases of c_i. With positive e_i, the social planner never chooses points on the right hand side of the Laffer curve (which are not Pareto optimal).[14] This justness function respects two properties of minimum sacrifice. First, losses from negative deviations from zero sacrifice, i.e., from positive tax liabilities, increase more than proportionally with the size of the deviation. Second, positive deviations, i.e., transfers, of the same size do not offset these losses.[15] In our empirical application, we set γ to two and δ and α to one. The latter two parameters affect mainly the unemployed, the only group that receives net transfers in our application and thus has a 'positive sacrifice'. The aim of this paper is to show which concepts of justness are in line with declining social weights under a reasonable calibration. Therefore, investigating how results change in a wide variety of calibrations is not particularly insightful. However, we have experimented with alternative values for α. Smaller values increase the social weight of the unemployed and keep the weights of the other groups relative to one another virtually unchanged.[16]

Similarly, we also consider minimum *relative* sacrifice where the function includes deviations of consumption c_i from gross income y_i relative to the level of consumption such that

$$\begin{aligned} f_i &= -\left(\frac{y_i - c_i}{c_i}\right)^{\gamma} \text{ if } y_i > c_i, \\ f_i &= \alpha\left(\frac{c_i - y_i}{c_i}\right)^{\delta} \text{ if } c_i > y_i, \\ &\gamma > 1, 0 \leq \alpha \leq 1, \delta \leq 1. \end{aligned} \tag{2.9}$$

A major advantage of our study is that we have observations of individual just levels of after-tax income for given gross incomes that are representative for the working population in Germany. Our framework allows using this information in the optimal tax formulae. We specify the justness functions similarly to the case of minimum sacrifice and set as reference point the level of just after-tax income taken from the survey. Thus the absolute formulation of the justness function is

[14] Starting from a point on the right-hand side of the Laffer curve for group i, improvements in the objective function of the social planner are possible by decreasing taxes T_i. This would increase f_i and increase tax revenues. This would, in turn, allow reducing taxes for some other group $j \neq i$. This increase in the objective function of the social planner would be a Pareto improvement as long as individual utility increases with net income.

[15] As noted in Weinzierl (2014), this is consistent with loss aversion (Kahneman and Tversky, 1979).

[16] See appendix F for variations of δ and γ.

$$
\begin{aligned}
f_i &= -(c_i^{\text{just}} - c_i)^\gamma \text{ if } c_i^{\text{just}} > c_i, \\
f_i &= \alpha(c_i - c_i^{\text{just}})^\delta \text{ if } c_i > c_i^{\text{just}}, \\
&\gamma > 1, 0 \leq \alpha \leq 1, \delta \leq 1
\end{aligned}
\tag{2.10}
$$

and the relative one is

$$
\begin{aligned}
f_i &= -\left(\frac{c_i^{\text{just}} - c_i}{c_i}\right)^\gamma \text{ if } c_i^{\text{just}} > c_i, \\
f_i &= \alpha\left(\frac{c_i - c_i^{\text{just}}}{c_i}\right)^\delta \text{ if } c_i > c_i^{\text{just}}, \\
&\gamma > 1, 0 \leq \alpha \leq 1, \delta \leq 1.
\end{aligned}
\tag{2.11}
$$

The parameters are calibrated as for minimum sacrifice. Note that the resulting *absolute* weights from an inverse optimal taxation simulation with different justness functions differ in magnitude because derivatives of the f_i functions differ. To make the comparison of weights between concepts of justness easier, we therefore calculate relative weights by dividing the obtained absolute weights e_i through the absolute weight of group 0 as in Blundell et al. (2009).

2.3. Empirical Calibration

2.3.1. The Data

We use data from the 2015 wave of the German Socio-Economic Panel (SOEP), a representative annual household panel survey. Wagner et al. (2007) provide a detailed description of the data.

As the model does not cover spousal labor supply, we restrict the analysis to working-age singles. We exclude individuals with children, heavily disabled and people who receive Unemployment Benefit I,[17] because their budget constraints and labor supply behavior differ substantially. Group 0 consists of the unemployed receiving Unemployment Benefit II.[18] We exclude the long-term unemployed with transfer non-take up, as they differ substantially from the standard case and face a different budget constraint. For the analysis we make use of a question in the SOEP, introduced in the 2015 wave,

[17]This transfer is targeted to the short-term unemployed and depends on the previous labor income.
[18]This transfer is targeted at the long-term unemployed and covers the social existence minimum.

Table 2.1.: Summary Statistics

	Mean	Std. Dev.	N
Monetary variables			
Monthly Gross Income	2626.75	1925.41	1119
Monthly Net Income	1766.18	991.86	1119
Just Net Income*	2150.85	1040.89	572
Demographics			
Sex (1=men, 2=women)	1.41	0.49	1119
Weekly Hours of Work**	41.66	9.51	990
Age	43.97	10.47	1119
East Germany Dummy	0.27	0.45	1119
Party supported in percent			
CDU/CSU (conservatives)	13.2	0.339	1119
SPD (social democrats)	8.9	0.285	1119
Bündnis 90/Die Grünen (green)	8.7	0.282	1119
DIE LINKE (left)	3.4	0.182	1119
FDP (classical liberal)	0.3	0.054	1119

*Only individuals who perceive their gross income as just
**Excluding the unemployed
Source: Own calculations based on the SOEP

that asks individuals what monthly income they would consider just. This question is discussed in more detail in the following subsection.

Table 2.1 shows summary statistics for our sample. Net incomes equal gross incomes and transfers minus income taxes and social security contributions. Only the currently employed are asked questions about what income they would consider as just.[19] Therefore, average just net income is substantially larger than average actual net income, which includes the unemployed.

2.3.2. Just and Actual Budget Constraints

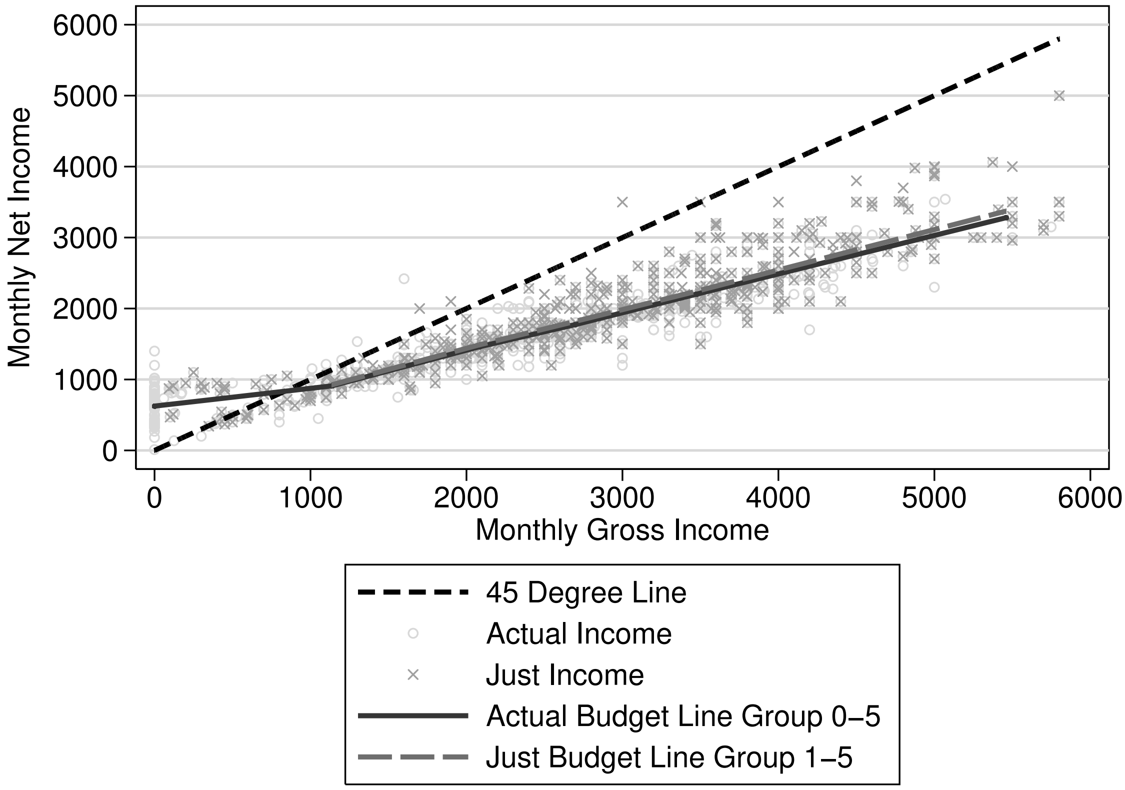
In the 2015 wave, the SOEP introduced new questions that ask what amount of income respondents would consider just in their current occupation. In particular, individuals state how high their gross income and net income would have to be in order to be just. A screenshot of this part of the questionnaire is provided in Appendix C.

Compared to other approaches to obtain information about individuals' ideas of justness, the advantage of the question is that individuals do not need to have a worked

[19]For the working poor, we add actual transfers to stated just net incomes, as these do not include transfers. Transfers include child benefits and supplements, Unemployment Benefit II, housing benefits and alimonies.

out theory of just taxation in mind to answer the question. Moreover, interviewees do not need a thorough understanding of tax schedules.

Figure 2.1.: Just Net and Gross Incomes.



Source: Own calculations based on SOEP

The German tax and transfer system is characterized by relatively generous transfers for the unemployed and high transfer withdrawal rates of up to 100 percent. Figure 2.1 shows the status quo of the German tax and transfer system and the just tax and transfer system based on our sample. The first segment of the actual budget line is almost horizontal at a net income of about 600 Euro. This represents transfer recipients. The slope of the budget line is steeper further to the right, representing individuals who do not receive transfers, but pay income taxes and social security contributions.

Gray circles represent the actual net incomes for given gross incomes. Some circles are crossed by x. This means either that an individual considers his or her actual income just or the actual income of another person. The 45 degree line marks the points where no taxes are paid. Points above this line represent actual transfer recipients or those who deem receiving transfers as just. However, most individuals perceive net incomes to be fair, where taxes have to be paid. It is likely that status quo bias explains this pattern.

Nonetheless, the answers of the respondents reflect actual perceptions of just incomes. The solid blue and the dashed red lines summarize this information. The solid blue line depicts the average actual budget constraint for six income groups that we use in the main analysis. The dashed red line shows the just budget constraint for the same groups. The just budget line is slightly above the actual budget line. The groups are defined as the unemployed and quintiles of those with positive gross labor incomes. The budget lines are based on averages for the groups. The actual budget line is relatively flat for the working poor, implying high withdrawal rates. The *just* budget line is defined only for those with positive labor income and lies slightly above the actual budget line. This reflects the preferences for paying less taxes. The distribution of net incomes for a given value of gross income is skewed toward the no tax line. Deviations in this direction can be explained with allowances. The positive skew of just net incomes is due to more people perceiving substantially higher net incomes as just than less. The incidence of crossed circles, i.e., persons who perceive their current income as just is higher below and around the average budget lines.

2.3.3. Labor Supply Elasticities

Similar to Blundell et al. (2009) and Haan and Wrohlich (2010), we use a random utility discrete choice labor supply model where each individual can choose between five work hours choices and unemployment. Each hours-person combination is associated with a gross income and net income calculated using the microsimulation model STSM. See Jessen et al. (2017b); Steiner et al. (2012) for further details on the STSM and the labor supply model.

To estimate mobility elasticities we first assign each hours-person combination in the data to an income group $i = 1, .., I$.[20] Then we predict changes in relative employment shares of income groups due to changes in relative net incomes $c_i - c_{i-1}$ and $c_i - c_0$ and calculate the mobility elasticities given by equations (2.3) and (2.4). The elasticities are reported in the tables in the next section.

[20]For instance, a person with an hourly wage of 20 Euro earns a gross income of approximately 860 Euro per month if she works 10 hours per week and about 1720 Euro if she works 20 hours. If she works 10 hours, she is assigned to group I. If she works 20 hours, she is assigned to group II. In contrast, a person with an hourly wage of 50 Euro is assigned to income group II if she works 10 hours, earning about 2150 Euro per month.

2.4. Results

2.4.1. Main Results

Table 2.2 shows average monthly individual gross incomes (column I) and corresponding average net incomes (column II) for six income groups. As is apparent from the increase in net incomes from group 0 to group 1, the marginal transfer withdrawal rate is substantial in the status quo. Column III shows average net incomes perceived as just. These average just net incomes are slightly above average actual net incomes for all groups. As only employed persons respond to the SOEP question about just net income, just net income is set marginally above the actual average transfer income of group 0.[21] Column IV shows the population share of each income group and columns V and VI display the extensive and intensive mobility elasticities, which have been estimated as described in subsection 2.3.3. For group 1, there is only one elasticity, see equations (2.3) and (2.4). The last five columns show relative explicit social weights for the different justness concepts.

Table 2.2.: Resulting Relative Weights for Different Justness Concepts

	I	II	III	IV	V	VI	VII	VIII	IX	X	XI
Group	Gross Income	Net Income	Just Net Income	Share	η	ζ	Welfarist	Minimum Sacrifice		Subjective Justness	
								Abs	Rel	Abs	Rel
0	0	625	630*	0.11	-	-	1	1	1	1	1
1	1137	910	925	0.19	0.08**	0.08**	0.239	0.0020	1.426	0.0797	0.1675
2	2082	1461	1488	0.17	0.10	0.08	0.364	0.0007	0.8488	0.0674	0.3645
3	2697	1773	1819	0.19	0.09	0.07	0.357	0.0005	0.7300	0.0390	0.3083
4	3472	2200	2242	0.17	0.07	0.06	0.392	0.0003	0.8059	0.0467	0.5722
5	5458	3279	3373	0.18	0.05	0.08	0.368	0.0002	0.9048	0.0196	0.5298

Note: German single households; own calculations based on the SOEP and the STSM.
*Just net income for this group is set as explained in the text.
**Overall elasticity of group one is 0.16.
η: extensive labour supply elasticity, ζ: intensive elasticity.

The welfarist approach (column VII) is an application of Saez (2002) as in Blundell et al. (2009). Group 0 has the highest social weight, the working poor (group 1) have the lowest weight in line with previous studies described in Appendix A. At the optimum, the welfarist weights show the costs of redistributing one Euro from individuals in group 0 to individuals in other groups. For instance, an increase in income for individuals in group 1 would reduce income in group 0 by only 0.239 Euro because individuals would

[21] We experimented with different values for this number. While changing the just net income of group 0 has a substantial impact on this group's subjective social justness weights relative to other groups, the weights of other groups relative to one another remain virtually the same.

move from group 0 to group 1, reducing the transfer burden of the state. Equivalently, the social planner values increasing the income for group 1 by one Euro 0.239 times as much as increasing the income of group 0 by one Euro. The low weights for the working poor are related to the high marginal tax rate for individuals moving from group 0 to group 1.[22] Relative weights of the upper four income groups are close to each other, in line with previous findings for Germany by Bargain et al. (2014).

Table D.1 in Appendix D shows the optimal welfarist tax schedule with weights decreasing with income. The resulting optimal tax schedule implies a substantially lower marginal transfer withdrawal rate for the working poor than in the status quo and higher net incomes for groups 1, 2, and 3. This underlines our finding that decreasing welfarist weights would imply lower transfer withdrawal rates.

Column VIII of Table 2.2 displays optimal weights for the minimum absolute sacrifice approach. These weights show how much it costs in terms of sacrifice of group 0 to reduce the sacrifice for members of a particular group as defined in equation (2.8). We focus the interpretation on the working groups as the unemployed are net recipients of transfers and thus 'pay a positive sacrifice', see section 2.2.2. The weight of this group depends strongly on the choice of parameters, especially α, but this does not change the ranking of the working groups. A comparison of the weights of tax-paying groups shows the highest weight for the working poor, 0.002,[23] and decreasing weights with income. The social planner is indifferent between imposing a slightly higher sacrifice on the working poor and imposing four times this additional sacrifice on the middle class (group 3). As the sacrifice increases quadratically with taxes paid, the marginal sacrifice for the working poor is relatively small. Consider the benchmark case with fixed incomes and the same marginal sacrifice for all groups. In this case, all weights would be the same. This is the notion of equal marginal sacrifice. In comparison, in our analysis the marginal sacrifice is lower for the working poor. Therefore, weights are higher for this group.[24] A similar reasoning applies to the other groups, which results in declining social weights. Consequently, the minimum absolute sacrifice principle is in line with the 2015 German tax and transfer system.

Column IX shows results for the minimum *relative* sacrifice principle. Again, the working poor have the highest weight of the groups with a positive tax burden. However, in contrast to the absolute sacrifice principle, weights are not decreasing with income but U-shaped. Top income earners have relatively high weights according to the *relative* sacrifice principle, because the tax paid is divided through a high consumption level. Thus a small increase in taxes would not increase the relative sacrifice of this group by much. In fact, the middle class (group 3) has the lowest weight according to this principle

[22] *Ceteris paribus*, higher elasticities and higher marginal tax rates imply a position further to the right of the Laffer curve and thus lower social weights.

[23] Again, note that the *absolute* value of this weight depends on the calibration of α, which determines the weight of group 0. Therefore, the focus is on the relative difference between working groups.

[24] As the welfarist weights indicate, the deadweight loss of increasing taxes for group 1 is very high. If it was lower, this group's minimum sacrifice weight would be even higher.

as one would have to redistribute less to members of this group to reduce their sacrifice. Thus, the German tax and transfer system does not imply decreasing social weights under the minimum relative sacrifice principle.

Columns X and XI show social weights according to the absolute and relative subjective justness principles respectively. The subjective justness principle implies penalties for the deviation of net incomes from perceived just net incomes. As discussed above, there is no information on perceived just net incomes of the unemployed, so we focus on the interpretation of the social weights of working groups. For the absolute justness principle, the working poor have the highest social weights of the working population because their average net income deviates from just net income by only 15 Euros. Social weights are decreasing except for group 4, as individuals in this group would consider a net income of only 42 Euros more than their current income just. When considering *relative* deviations from just net income, group 4 has the highest social weights of all working groups since the deviation from just income is smaller relative to the high consumption level of group 4.

Only the minimum absolute sacrifice principle is in line with decreasing social weights. For absolute subjective justness, weights are declining except for group 4. The working poor have the lowest weight of all working groups in the welfarist and the relative subjective justness approach,[25] while they have the highest weight of all working groups in the absolute subjective justness approach.

To sum up, we find that the minimum absolute sacrifice principle is in accordance with declining social weights in the status quo. Thus, the minimization of absolute sacrifice is a good description of the aims of the German society regarding the tax and transfer system.

2.4.2. Results for Subsamples

To explore whether the 2015 tax transfer schedule is designed with a particular concept of justness with focus on a specific group in mind, we split the sample into different groups. These groups differ substantially regarding the income distribution and elasticities, which might lead to different social weights. Moreover, perceived justness of taxation in these groups might differ systematically.

First, the sample is split into females and males. We find that women have a more elastic labor supply than men and lower incomes. In light of the discussions regarding the gender wage gap, subjective justness could differ systematically between women and men as well. Then we present our results for East Germans and West Germans, respectively. These two groups lived under different political systems for more than a generation. We show that West Germans have higher incomes, less unemployment, but lower extensive elasticities than East Germans. Additionally, the tax schedule might be

[25]The explanation is that the costs of decreasing the relative sacrifice for the working poor are low because of the relatively small denominator of f_1 and the fact that redistribution to this group is cost-effective.

Table 2.3.: Resulting Relative Weights for Different Justness Concepts for Women without Children

	I	II	III	IV	V	VI	VII	VIII	IX	X	XI
Group	Gross	Net	Just Net	Share	η	ζ	Welfarist	Minimum Sacrifice		Subjective Justness	
	Income	Income	Income					Abs	Rel	Abs	Rel
0	0	615	620*	0.05	-	-	1.	1	1	1	1
1	976	863	865	0.19	0.09**	0.09**	0.126	0.0043	3.8757	0.3059	0.6062
2	1903	1271	1352	0.20	0.12	0.10	0.143	0.0006	0.7603	0.0090	0.0362
3	2548	1715	1747	0.19	0.10	0.10	0.200	0.0006	1.2620	0.0311	0.2395
4	3342	2083	2122	0.23	0.07	0.10	0.174	0.0003	0.9403	0.0222	0.2522
5	4948	3122	3226	0.15	0.06	0.12	0.182	0.0002	1.5273	0.0088	0.2206

Note: German single households; own calculations based on the SOEP and the STSM.
*Just net income for this group is set as explained in the text.
**Overall elasticity of group one is 0.18.

more in line with preferences of supporters of particular political parties. To this end we exploit the rich collection of household characteristics in the SOEP, in particular, which political party, if any, individuals support.

Results for Men and Women

In Table 2.3 we report results for the subsample of women without children, which we compare, in the following, with the results for the main sample and, later, to men. As expected, gross and net incomes in all income groups are lower and labor supply elasticities are slightly higher. For the welfarist case, the working groups have smaller weights relative to the unemployed than in the main sample. As before, we find that the working poor have the lowest weight. The finding that social weights for the minimum absolute sacrifice concept are decreasing with income is robust for this subsample. The working poor have higher weights than in the main sample as they pay considerably less taxes. As before, in the relative sacrifice case, the working poor have the highest weights and top income earners have the second highest weights.

For the absolute subjective justness concept, weights are decreasing except for group 2. The working poor have a high weight because for women this group's actual income is very close to its just net income. For relative justness, the working poor have the highest weight of the working groups and the three highest income groups have similar weights. Again, group 2 is the odd one out with a very low weight.

Table 2.4 shows results for the subsample of men. Incomes are higher and elasticities are lower than for women. In the welfarist case, weights of working groups are higher than for women. This is caused by lower elasticities, which lead to men being further on the left of the Laffer curve. Nevertheless, the working poor again have the lowest weight. The finding that weights in the absolute sacrifice case decrease with income holds for men as well. The weight of the working poor is lower for men than for women because

Table 2.4.: Resulting Relative Weights for Different Justness Concepts for Men without Children

	I	II	III	IV	V	VI	VII	VIII	IX	X	XI
Group	Gross	Net	Just Net	Share	η	ζ	Welfarist	Minimum Sacrifice		Subjective Justness	
	Income	Income	Income					Abs	Rel	Abs	Rel
0	0	627	632*	0.15	-	-	1	1	1	1	1
1	1265	971	997	0.17	0.05**	0.05**	0.438	0.0015	0.7015	0.0846	0.1992
2	2228	1547	1565	0.18	0.08	0.04	0.513	0.0006	0.4911	0.1426	0.8650
3	2875	1889	1944	0.16	0.07	0.04	0.522	0.0004	0.4461	0.0477	0.4240
4	3622	2316	2381	0.17	0.06	0.04	0.551	0.0003	0.4873	0.0426	0.5698
5	6124	3561	3652	0.16	0.05	0.06	0.509	0.0002	0.4768	0.0281	0.8907

Note: German single households; own calculations based on the SOEP and the STSM.
*Just net income for this group is set as explained in the text.
**Overall elasticity of group one is 0.1.

the male group 1 pay substantially more taxes than their female counterparts. Again, in the relative minimum sacrifice case, the working poor have the highest weight and the middle class has the lowest weight of working groups. For the absolute subjective justness concept, weights are decreasing apart from group 2. For relative subjective justness, the working poor have the smallest weight.

Results for East and West Germany

Gross, net, and net just incomes are higher across all groups in West Germany (see Table 2.6) compared to East Germany (see Table 2.5). In contrast to the main sample and the previously analyzed subsamples, in the sample of East Germans the working poor are net transfer recipients and the marginal withdrawal rate when moving from group 1 to group 2 is still substantial.

The welfarist weights show highest social weights for the unemployed and lowest for the working poor (group 1 in the West, groups 1 and 2 in the East). An increase in income for individuals in group 1 would reduce income in group 0 by only 0.21 Euro in West Germany and by about 0.34 in East Germany. The relative weights of the four (three for East Germany) higher income groups are very similar and higher than the weights for the working poor.

As in our main findings, optimal weights under minimum absolute sacrifice are decreasing in both samples, though the weight of group 1 is closer to the weight of group 0 than group 2 for East Germany as group 1 are net transfer recipients and thus enjoy a 'positive tax sacrifice'. Regarding groups with a positive tax burden, the weights imply that the social planner is roughly indifferent between imposing a slightly higher sacrifice on the working poor (group 1 in West Germany, group 2 in East Germany) and imposing twice this additional sacrifice on group 2 in the case of West Germany and group 3 in the case of East Germany. This shows that the minimum absolute sacrifice principle is in line with the 2015 German tax and transfer system for East and West Germans.

Table 2.5.: Resulting Relative Weights for Different Justness Concepts for East Germany

	I	II	III	IV	V	VI	VII	VIII	IX	X	XI
Group	Gross Income	Net Income	Just Net Income	Share	η	ζ	Welfarist	Minimum Sacrifice		Subjective Justness	
								Abs	Rel	Abs	Rel
0	0	591	596*	0.18	-	-	1	1	1	1	1
1	774	837	851	0.17	0.10**	0.10**	0.339	0.9957	1.0308	0.1211	0.2408
2	1581	1192	1222	0.18	0.16	0.08	0.342	0.0011	1.0580	0.0573	0.2294
3	2200	1574	1594	0.17	0.13	0.08	0.424	0.0007	0.9845	0.1059	0.7481
4	2808	1875	1920	0.14	0.11	0.07	0.430	0.0005	0.8241	0.0482	0.4772
5	4039	2607	2625	0.16	0.09	0.08	0.428	0.0003	0.9393	0.1188	2.3145

Note: German single households; own calculations based on the SOEP and the STSM.
*Just net income for this group is set as explained in the text.
**Overall elasticity of group one is 0.2.

Table 2.6.: Resulting Relative Weights for Different Justness Concepts for West Germany

	I	II	III	IV	V	VI	VII	VIII	IX	X	XI
Group	Gross Income	Net Income	Just Net Income	Share	η	ζ	Welfarist	Minimum Sacrifice		Subjective Justness	
								Abs	Rel	Abs	Rel
0	0	653	658*	0.08	-	-	1	1	1	1	1
1	1408	1004	1030	0.21	0.07**	0.07**	0.210	0.0010	0.7161	0.0405	0.094
2	2324	1585	1616	0.16	0.09	0.08	0.309	0.0005	0.7465	0.0499	0.2905
3	2907	1898	1946	0.19	0.08	0.08	0.300	0.0004	0.6963	0.0314	0.2608
4	3699	2322	2378	0.19	0.06	0.06	0.323	0.0003	0.7449	0.0289	0.3593
5	6010	3516	3632	0.17	0.05	0.08	0.298	0.0002	0.7991	0.0129	0.3652

Note: German single households; own calculations based on the SOEP and the STSM.
*Just net income for this group is set as explained in the text.
**Overall elasticity of group one is 0.14.

Results for the minimum *relative* sacrifice principle show that the working poor have the highest weight of the groups with a positive tax burden in East Germany, but not in West Germany, where weights for the top income group are highest. The difference arises because top income earners in West Germany earn considerably more than their East German counterparts. As explained in section 2.4.1, this implies higher weights for this justness concept because the denominator of the sacrifice is higher. In both samples the middle class (group 3 in the West, group 4 in the East) has lowest weights. Thus, the German tax and transfer system does not result in decreasing social weights under the minimum relative sacrifice principle.

The last two columns report social weights under the absolute and relative subjective justness principles, respectively. When considering the absolute justness principle, the working poor in group 1 in the East have the highest social weights of the working population because their average net income deviates from just net income by only 14 Euros. While the weights jump in the East German sample, for West Germans social subjective weights decrease starting from group 2. The *relative* deviations from just net income imply increasing weights in West Germany. Taking into account that the number of observations is smaller for East Germany, this pattern could be prevalent for this group as well.

Results for Supporters of Political Parties

We show results for subjective justness for three sets of political party supporters. This is interesting because subjective just incomes might differ substantially between supporters of different parties. This allows us to analyze if the tax transfer schedule is in line with the preferences of a specific coalition. Unfortunately, the number of observations is too low to allow a party-specific analysis, as most respondents do not identify themselves as supporters of a particular party. We investigate three groups. First, supporters of the current grand coalition of the conservative Christian Democratic Union of Germany (CDU) and Christian Social Union in Bavaria (CSU) and the Social-Democratic Party (SPD). At any point of time since World War II at least one of these parties has been in power in West Germany. Additionally, we look at two passionately debated possible future coalitions: (1) a left-wing coalition including the SPD, the Green party and the socialist Left party; and (2) a coalition including the CDU/CSU, the Greens, and the classical liberal Free Democratic Party (FDP).

Table 2.7 shows results for supporters of the CDU/CSU and SPD governing coalition, in power in spring 2017. The expectation for this group is that party supporters are relatively content with the status quo. Compared to the main sample, incomes are higher in all groups. As expected, just incomes are close to actual incomes. Strikingly, the pattern for the absolute justness weights is the same as in the main sample. Weights are decreasing, except for group 4. The pattern for relative justness is very similar to the main sample as well: The highest income earning groups have the highest weights.

Table 2.8 shows results for supporters of center left parties. One would expect that high income supporters of these parties are content with paying relatively high taxes and that lower income earners would prefer more redistribution. The income distribution of this subsample is similar to that of supporters of the grand coalition. For both subjective justness concepts, the highest income group has the highest weight because this group would consider paying only 15 Euros less taxes as just. In contrast, in the main sample, the difference between actual and just net income for group five is about 100 Euros. However, in the left-wing sample, group 4 would perceive paying about 90 Euros less taxes as just and consequently has relatively low social weights.

The working poor have low weights as well even though they would consider paying only 15 Euros less taxes as fair. This is because the dead weight loss of redistribution to the working poor is low while this figure is high for higher income groups as indicated by the low welfarist weight for the working poor (not reported for this subsample).

Table 2.7.: Resulting Relative Weights for Subjective Justness Concepts for SPD/CDU/CSU supporters

	I	II	III	IV	V	VI	VII	VIII
Group	Gross Income	Net Income	Just Net Income	Share	η	ζ	Subjective Justness Abs	Rel
0	0	689	694*	0.09	-	-	1	1
1	1298	924	929	0.19	0.07**	0.07**	0.1201	0.2164
2	2317	1641	1660	0.19	0.10	0.08	0.0730	0.4121
3	2946	1910	1944	0.16	0.09	0.08	0.0373	0.2834
4	3641	2288	2314	0.21	0.07	0.06	0.0538	0.5911
5	6272	3553	3604	0.15	0.05	0.08	0.0255	0.6723

Note: German single households; own calculations based on the SOEP and the STSM.
*Just net income for this group is set as explained in the text.
**Overall elasticity of group one is 0.14.

Table 2.9 reports results for supporters of CDU/CSU, the Green Party and the FDP. As expected, incomes in all groups are higher than in the left-wing sample. This difference is between 151 (group 4) and 1186 Euros (group 5). Compared to the left-wing sample, the expectation is that the working poor will not demand substantially more redistribution. Indeed, the absolute justness social weights for this group are the highest among the working groups. For relative justness, groups 3 to 5 have the highest weights as they are relatively content with their net income.

The analysis by party supporters shows that social weights for absolute justness are roughly decreasing for supporters of the grand coalition, thus corroborating our main findings. Consequently, the results for absolute subjective justness in the main sample seem to be driven mainly by supporters of the grand coalition and independents (see Appendix E). Their preferences for the tax transfer schedule seem to be roughly in line with the concept of minimum absolute sacrifice, for which we find decreasing social weights in the main analysis. If the concept of justness that explains current tax practice and the subjective justness for most people is the concept of minimum absolute sacrifice, the role of welfarist optimal taxation models is not as important as previously assumed.

Our results provide the grounds for future research on the formation of preferences for tax transfer schedules. First, a large scale survey that allows to disentangle single parties or even the wings of parties could be used to confirm our suggestive evidence. Second, it would be interesting to investigate whether the tax design forms subjective justness or vice versa.

Table 2.8.: Resulting Relative Weights for Subjective Justness Concepts for SPD/Left/Green supporters

	I	II	III	IV	V	VI	VII	VIII
Group	Gross	Net	Just Net	Share	η	ζ	Subjective Justness	
	Income	Income	Income				Abs	Rel
0	0	790	795*	0.10	-	-	1	1
1	1256	954	969	0.18	0.07**	0.07**	0.0106	0.0153
2	2354	1618	1634	0.18	0.10	0.08	0.0755	0.3154
3	3075	1978	2003	0.18	0.08	0.09	0.0472	0.2939
4	3710	2331	2423	0.18	0.07	0.07	0.0142	0.1200
5	5598	3338	3353	0.18	0.05	0.08	0.0818	1.4635

Note: German single households; own calculations based on the SOEP and the STSM.
*Just net income for this group is set as explained in the text.
**Overall elasticity of group one is 0.14.

2.4.3. Robustness

In Appendix F, we show the robustness of our results. First, we analyze the robustness of the obtained social weights for absolute justness to different values of γ and δ (tables F.1 and F.2). The result that social weights decline with income is robust to a wide range of calibrations. This shows that the main result is not driven by the parameter choice. Second, we set the intensive and extensive elasticities of all groups to 0.1 and show the results for all concepts of justness (Table F.3). The results are very close to the main results. This shows that slight variations in the elasticities do not change the results substantially.

2.5. Conclusion

In this paper we reconcile a puzzling contrast between current tax transfer practice in many countries and the common approach in the optimal taxation literature. While the literature commonly assumes that the social planner values an additional unit of income for poor households more than an additional unit of income for higher income households, commonly observed high transfer withdrawal rates are only optimal if social weights of the working poor are very small. Therefore, we compare alternative approaches to welfarism and calculate the implied social weights. We formulate the problem of a social planner for three distinct concepts of justness: the welfarist approach, where the social planner maximizes the weighted sum of utility; alternatively, the minimum sacrifice concept where the social planner minimizes the weighted sum of absolute or relative (tax-)sacrifice; and, thirdly, the approach of subjective justness where the social planner minimizes absolute or relative deviations from perceived just

Table 2.9.: Resulting Relative Weights for Subjective Justness Concepts for CDU/CSU/FDP/Green supporters

	I	II	III	IV	V	VI	VII	VIII
Group	Gross Income	Net Income	Just Net Income	Share	η	ζ	Subjective Justness Abs	Rel
0	0	696	701*	0.04	-	-	1	1
1	1423	925	929	0.20	0.07**	0.07**	0.0571	0.1011
2	2541	1697	1742	0.20	0.10	0.09	0.0147	0.0858
3	3284	2147	2162	0.19	0.08	0.09	0.0481	0.4578
4	3861	2352	2389	0.19	0.06	0.06	0.0186	0.2105
5	6784	3812	3843	0.18	0.04	0.09	0.0213	0.6375

Note: German single households; own calculations based on the SOEP and the STSM.
*Just net income for this group is set as explained in the text.
**Overall elasticity of group one is 0.14.

net income. For the concept of subjective justness, we use a SOEP question introduced in the 2015 wave to obtain information about what amount of taxes individuals consider as just. Of course, all approaches maintain budget neutrality and account for labor supply reactions.

Like the existing literature, we find that the 2015 German tax and transfer system implies very low social weights for the working poor according to the welfarist criterion. The social planner values increasing the income for the working poor by one Euro 0.65 times as much as increasing the income of top earners by one Euro. This implies that an additional Euro of consumption for the working poor is valued less than marginal consumption of top income earners.

In contrast, the current tax-transfer practice can be reconciled as optimal and in line with decreasing social weights under the minimum absolute sacrifice criterion, under which the social planner minimizes the sacrifice of individuals. In this case, the social planner is indifferent between imposing a slightly higher sacrifice on the working poor and imposing four times this additional sacrifice on the middle class.

Moreover, we find that the status quo is roughly in line with a social planner minimizing deviations from what taxpayers consider as just. The subgroup analysis by political parties shows that this result is in line with preferences of supporters of those political parties that shaped the tax policy under CDU/CSU and SPD in the years 2013 to 2017 in Germany. Our results suggest that the role of welfarist optimal taxation models is not as important as previously assumed.

Appendix

A. Review of the Positive Optimal Taxation Literature

In a number of papers, researchers use optimal income taxation frameworks that incorporate labor supply responses to obtain "tax-benefit revealed social preferences" (Bourguignon and Spadaro, 2012), i.e., they calculate the social weights under which the current tax and transfer system is optimal. Blundell et al. (2009) apply the Saez (2002) framework to single mothers in Germany and the UK to calculate implied social weights. They find that working mothers with low incomes have low weights compared to the unemployed and most other income groups. For Germany, social weights for working poor single mothers with children under school-age can even become negative, thus implying a non-paretian social welfare function. Bourguignon and Spadaro (2012) apply positive optimal taxation to the French redistribution system. They find negative social weights for the highest income earners and equally for the working poor if participation elasticities are high. In general, social weights for the working poor are much lower than those for the unemployed or the middle class. Bargain et al. (2014) calculate social weights for 17 European countries and the United States. For all analyzed countries, they find the highest social weights for the unemployed and substantially lower weights for the working poor, i.e., the group with the lowest net income apart from the unemployed. In Belgium, France, Germany, the Netherlands, Portugal, the UK, and Sweden the tax-transfer system implies the lowest social weights for this group. Zoutman et al. (2016) show that the 2006 tax-transfer system in the Netherlands, as well as reform proposals by political parties, imply the highest weights for the middle class. Lockwood and Weinzierl (2016) perform inverse optimal taxation for the US from 1979 to 2010. They find that, if the standard welfarist model is correct, either perceived elasticities of taxable income or value judgments have changed considerably over time. This is interpreted as evidence that conventional assumptions of the benchmark model of optimal taxation should be questioned. Immervoll et al. (2007) find that expanding redistribution to the working poor would be very cost effective and would virtually imply no deadweight burden.

B. Optimal Tax Formulae in the General Model

Behavioral reactions imply that h_i changes in case of a change in T_i. Using the product rule, the first order condition with respect to T_i is obtained as

$$-\mu_i h_i \frac{\partial f_i}{\partial c_i} - \sum_{j=0}^{I} \mu_j f_j \frac{\partial h_j}{\partial c_i} = -\lambda \left(h_i - \sum_{j=0}^{I} T_j \frac{\partial h_j}{\partial c_i} \right), \tag{2.12}$$

where λ is the multiplier of the budget constraint. The first order condition with respect to λ is the budget constraint. Reorganizing 2.12 and defining the explicit social weights as $e_i = \mu_i/\lambda$ yields

$$\left(1 - e_i \frac{\partial f_i}{\partial c_i}\right) h_i - \sum_{j=0}^{I} e_j f_j \frac{\partial h_j}{\partial c_i} = \sum_{j=0}^{I} T_j \frac{\partial h_j}{\partial c_i}. \tag{2.13}$$

Rearranging we obtain

$$h_i = h_i e_i \frac{\partial f_i}{\partial c_i} + \sum_{j=0}^{I} e_j f_j \frac{\partial h_j}{\partial c_i} + \sum_{j=0}^{I} T_j \frac{\partial h_j}{\partial c_i}. \tag{2.14}$$

With no income effects, $\sum_{i=0}^{I} \partial h_j / \partial c_i = 0$, i.e. increasing the income of all groups by the same amount has no effect on the choice of groups. Therefore, summing equation (2.14) over all $i = 0, \ldots, I$, one obtains that the redefined social welfare weights are normalized as

$$\sum_{i=0}^{I} h_i e_i \frac{\partial f_i}{\partial c_i} = 1. \tag{2.15}$$

The assumption of no income effects implies that only h_{i-1}, h_i, h_{i+1}, and h_0 react to changes in T_i such that equation (2.13) simplifies to

$$\begin{aligned}\left(1 - e_i \frac{\partial f_i}{\partial c_i}\right) h_i \quad = \quad & T_0 \frac{\partial h_0}{\partial c_i} + T_{i-1} \frac{\partial h_{i-1}}{\partial c_i} + T_i \frac{\partial h_i}{\partial c_i} + T_{i+1} \frac{\partial h_{i+1}}{\partial c_i} \\ & + e_0 f_0 \frac{\partial h_0}{\partial c_i} + e_{i-1} f_{i-1} \frac{\partial h_{i-1}}{\partial c_i} + e_i f_i \frac{\partial h_i}{\partial c_i} + e_{i+1} f_{i+1} \frac{\partial h_{i+1}}{\partial c_i}.\end{aligned} \tag{2.16}$$

Using the assumption that h_i depends only on the difference between the consumption of group i, consumption of the neighboring groups $i-1, i+1$, and group 0 and the fact that $\frac{\partial h_{i+1}}{\partial (c_{i+1} - c_i)} = -\frac{\partial h_i}{\partial (c_{i+1} - c_i)}$, $\frac{\partial h_i}{\partial (c_i - c_0)} = -\frac{\partial h_0}{\partial (c_i - c_0)}$, we can write after rearranging

$$\begin{aligned}\left(1 - e_i \frac{\partial f_i}{\partial c_i}\right) h_i \quad = \quad & (T_i - T_0) \frac{\partial h_i}{\partial (c_i - c_0)} - (T_{i+1} - T_i) \frac{\partial h_{i+1}}{\partial (c_{i+1} - c_i)} + (T_i - T_{i-1}) \frac{\partial h_i}{\partial (c_i - c_{i-1})} \\ & - e_0 f_0 \frac{\partial h_i}{\partial (c_i - c_0)} - e_{i-1} f_{i-1} \frac{\partial h_i}{\partial (c_i - c_{i-1})} - e_{i+1} f_{i+1} \frac{\partial h_{i+1}}{\partial (c_{i+1} - c_i)} \\ & + e_i f_i \left(\frac{\partial h_i}{\partial (c_i - c_0)} + \frac{\partial h_{i+1}}{\partial (c_{i+1} - c_i)} + \frac{\partial h_i}{\partial (c_i - c_{i-1})} \right).\end{aligned} \tag{2.17}$$

Using the definition of the elasticities (2.3) and (2.4) and that $\zeta_i \frac{h_i}{c_i - c_{i-1}} = \frac{\partial h_i}{\partial c_i - c_{i-1}}$, we obtain for each group after reorganizing

$$\begin{aligned}\frac{T_i - T_{i-1}}{c_i - c_{i-1}} &= \frac{1}{\zeta_i h_i}\Bigg\{\left(1 - e_i \frac{\partial f_i}{\partial c_i}\right) h_i - \eta_i h_i \frac{T_i - T_0}{c_i - c_0} + \zeta_{i+1} h_{i+1} \frac{T_{i+1} - T_i}{c_{i+1} - c_i} \\ &\quad + e_0 f_0 \eta_i \frac{h_i}{c_i - c_0} + e_{i-1} f_{i-1} \zeta_i \frac{h_i}{c_i - c_{i-1}} + e_{i+1} f_{i+1} \zeta_{i+1} \frac{h_{i+1}}{c_{i+1} - c_i} \\ &\quad - e_i f_i \left(\eta_i \frac{h_i}{c_i - c_0} + \zeta_{i+1} \frac{h_{i+1}}{c_{i+1} - c_i} + \zeta_i \frac{h_i}{c_i - c_{i-1}}\right)\Bigg\}.\end{aligned} \tag{2.18}$$

Note that, by setting $e_i = 0$, we obtain the Laffer-condition

$$\frac{T_i - T_{i-1}}{c_i - c_{i-1}} = \frac{1}{\zeta_i} + \frac{\zeta_{i+1} h_{i+1}}{\zeta_i h_i} \frac{T_{i+1} - T_i}{c_{i+1} - c_i} - \frac{\eta_i}{\zeta_i} \frac{T_i - T_0}{c_i - c_0}. \tag{2.19}$$

Substituting the equivalent of 2.18 for the next group in 2.18 and simplifying gives

$$\begin{aligned}\frac{T_i - T_{i-1}}{c_i - c_{i-1}} &= \frac{1}{\zeta_i h_i}\Bigg\{\left(1 - e_i \frac{\partial f_i}{\partial c_i}\right) h_i + \left(1 - e_{i+1} \frac{\partial f_{i+1}}{\partial c_{i+1}}\right) h_{i+1} \\ &\quad - \eta_i h_i \frac{T_i - T_0}{c_i - c_0} - \eta_{i+1} h_{i+1} \frac{T_{i+1} - T_0}{c_{i+1} - c_0} + \zeta_{i+2} h_{i+2} \frac{T_{i+2} - T_{i+1}}{c_{i+2} - c_{i+1}} \\ &\quad - (e_i f_i - e_{i-1} f_{i-1}) \zeta_i \frac{h_i}{c_i - c_{i-1}} - (e_i f_i - e_0 f_0) \eta_i \frac{h_i}{c_i - c_0} \\ &\quad - (e_{i+1} f_{i+1} - e_{i+2} f_{i+2}) \zeta_{i+2} \frac{h_{i+2}}{c_{i+2} - c_{i+1}} - (e_{i+1} f_{i+1} - e_0 f_0) \eta_{i+1} \frac{h_{i+1}}{c_{i+1} - c_0}\Bigg\}.\end{aligned} \tag{2.20}$$

Recursive insertion and simplifying gives the I formulae (2.5) that must hold if function (2.1) is optimized.

C. Questionnaire

Figure C.1.: The Question for Justness

67. Is the gross income that you earn at your current job just, from your point of view?

No .... ☐ Yes .... ☐ ➡ *Question **69!***

68. How high would your gross income have to be in order to be just?

Gross: ☐ euros per month Don't know .... ☐

69. Is the net income that you earn at your current job just, from your point of your view?

No .... ☐ Yes .... ☐ ➡ *Question **71!***

70. How high would your net income have to be in order to be just?

Net: ☐ euros per month Don't know .... ☐

Source: Official SOEP Questionnaire

D. Optimal Welfarist Tax Schedule

Table D.1 shows the optimal welfarist tax schedule, where, following Saez (2002), implicit welfare weights are set according to the formula

$$g_i = \frac{1}{\lambda c_i^{0.25}} \tag{2.21}$$

and the shares of income groups are determined endogenously by

$$h_i = h_i^0 \left(\frac{c_i - c_0}{c_i^0 - c_0^0} \right)^{\eta_i}, \tag{2.22}$$

where the superscript 0 denotes values in the status quo. The simulation was done achieving budget neutrality and setting net income of group 0 to the status quo, as a deviation from this is not politically feasible.

E. Resulting Social Weights for Independents

Table E.1 shows results for individuals who do not support any political party.

Table D.1.: Optimal Welfarist Tax Schedule

Group	Gross Income	Net Income	Optimal Net Income	Relative Weight
0	0	625	625	1
1	1137	910	1269	0.838
2	2082	1461	1640	0.786
3	2697	1773	1848	0.763
4	3472	2200	2060	0.742
5	5458	3279	2842	0.685

Note: German single households
Own calculations based on the SOEP and the STSM.

Table E.1.: Resulting Relative Weights for Subjective Justness Concepts for Independents

	I	II	III	IV	V	VI	VII	VIII
Group	Gross Income	Net Income	Just Net Income	Share	η	ζ	Subjective Justness Abs	Rel
0	0	589	594*	0.12	-	-	1	1
1	1050	897	904	0.18	0.08**	0.08**	0.2109	0.4895
2	1948	1379	1429	0.17	0.11	0.07	0.0379	0.2020
3	2551	1694	1730	0.19	0.09	0.07	0.0536	0.4380
4	3325	2111	2176	0.18	0.07	0.07	0.0312	0.3926
5	5270	3199	3329	0.15	0.05	0.09	0.0150	0.4287

Note: German single households; own calculations based on the SOEP and the STSM.
*Just net income for this group is set as explained in the text.
**Overall elasticity of group one is 0.16.

F. Sensitivity checks

Table F.1.: Resulting Relative Weights for Absolute Subjective Justness for Different Values of γ

	I	II	III	IV	V
Group	$\gamma = 1.1$	$\gamma = 1.5$	$\gamma = 2$	$\gamma = 3$	$\gamma = 5$
0	1	1	1	1	1
1	0.5293	0.0415	0.0020	5.6×10^{-6}	6.3×10^{-11}
2	0.4836	0.0249	0.0007	7.2×10^{-7}	1.1×10^{-12}
3	0.4600	0.0198	0.0005	3.1×10^{-7}	2.1×10^{-13}
4	0.4519	0.0175	0.0003	1.7×10^{-7}	6.2×10^{-14}
5	0.4230	0.0129	0.0002	5.5×10^{-8}	6.7×10^{-15}

Note: German single households
Own calculations based on the SOEP and the STSM.

Table F.2.: Resulting Relative Weights for Absolute Subjective Justness for Different Values of δ

	I	II	III	IV	V
Group	$\delta = 0.1$	$\delta = 0.3$	$\delta = 0.5$	$\delta = 0.7$	$\delta = 1$
0	1	1	1	1	1
1	4.4×10^{-6}	1.7×10^{-5}	6.7×10^{-5}	0.0004	0.0020
2	1.2×10^{-6}	5.2×10^{-6}	2.2×10^{-5}	8.8×10^{-5}	0.0007
3	7.8×10^{-7}	3.3×10^{-6}	1.4×10^{-5}	5.7×10^{-5}	0.0005
4	5.8×10^{-7}	2.5×10^{-6}	1.0×10^{-5}	4.3×10^{-5}	0.0003
5	3.2×10^{-7}	1.6×10^{-6}	5.7×10^{-6}	2.4×10^{-5}	0.0002

Note: German single households
Own calculations based on the SOEP and the STSM.

Table F.3.: Resulting Relative Weights for Different Justness Concepts with Elasticities Set to 0.1

	I	II	III	IV	V	VI	VII	VIII	IX	X	XI
Group	Gross Income	Net Income	Just Net Income	Share	η	ζ	Welfarist	Minimum Sacrifice		Subjective Justness	
								Abs	Rel	Abs	Rel
0	0	625	630*	0.11	-	-	1.	1.	1.	1.	1.
1	1137	910	925	0.19	0.1**	0.1**	0.169	0.0020	1.5882	0.0558	0.1173
2	2082	1461	1488	0.17	0.1	0.1	0.321	0.0007	0.9157	0.0595	0.3217
3	2697	1773	1819	0.19	0.1	0.1	0.304	0.0004	0.7778	0.0332	0.2627
4	3472	2200	2242	0.17	0.1	0.1	0.321	0.0003	0.8560	0.0382	0.4678
5	5458	3279	3373	0.18	0.1	0.1	0.293	0.0002	0.9411	0.0157	0.4223

Note: German single households; own calculations based on the SOEP and the STSM.
*Just net income for this group is set as explained in the text.
**Overall elasticity of group one is 0.2.

3. Hours Risk, Wage Risk, and Life-Cycle Labor Supply[1]

3.1. Introduction

The purpose of this paper is to locate the sources of idiosyncratic income risk and attribute to them the strength of their contributions to overall risk. The natural decomposition of labor income is into hours and wages. Knowing which type of shocks contributes more to total income risk is important for the evaluation of potential welfare gains of policy measures aimed at reducing income uncertainty. To that end we use the restrictions implied by a life-cycle model of labor supply and consumption to separate the contributions.

We adopt an augmented specification of the standard life-cycle model of consumption and labor supply, where risk stems from both hours and wage shocks. This is a departure from the extant literature, like Blundell et al. (2016) and Heathcote et al. (2014), who restrict idiosyncratic risk to originate from wages only. One exception is Altonji et al. (2013), who approximate economic decisions of agents in their statistical model of the labor income process. In contrast, our approach is based on lifetime utility maximization and allows to identify several key policy parameters, such as the Frisch and the Marshall elasticity. Moreover, we decompose hours and wage shocks into permanent and transitory components. Permanent wage shocks include the obsolescence of human capital, or the acquirement of new skills. Permanent hours shocks might, e.g., be caused by injuries.

By extending the model to include hours shocks, our analysis is complicated substantially. In our setting hours residuals contain reactions to wage shocks *in addition* to hours shocks. The solution is obtained by utilizing the covariance of hours and wages to estimate the parameter quantifying how wage shocks translate into the marginal utility of wealth. When this parameter increases, shocks have a larger impact, implying less insurance against risk. Accordingly, the parameter is related to estimates of consumption insurance, e.g., Blundell et al. (2008, 2016).[2] However, in contrast to previously estimated

[1]This chapter is based on Jessen and König (2017).

[2]As Cunha et al. (2005) point out, the partial insurance parameter can also be viewed as a test of congruence between the information set of the econometrician and the agent. In terms of welfare and behavior, insurance against and prior knowledge of a shock are equivalent. For instance, the reaction to a fully insured change in wages is the same as the reaction to a change in wages known in advance; it is given by the Frisch elasticity (Heathcote et al., 2014).

consumption insurance parameters, the parameter we estimate is a sufficient statistic for the Marshall elasticity. A major advantage of our approach is to estimate the Marshall elasticity eschewing consumption or asset data, the reliability of which has been hotly debated (Attanasio and Pistaferri, 2016).

Ultimately, the goal of the overarching research agenda is to devise efficacious policies mitigating income risk, with quantification of the sources being the first step. The current approach to the problem of income risk is to reduce the transmission of realized income risk on consumption through various ways of insurance. A prominent example is income taxation, which reduces the transition of gross income risk to net income risk and thus consumption risk (Varian, 1980; Krueger and Perri, 2011; Heathcote et al., 2017). However, an unexplored alternative is attacking the sources of income risk itself: wage and hours risk. Suppose there is the possibility to implement two policies: An increase in income tax progressivity or an extension of health care coverage. Both reduce net income risk. The former does so by reducing the transmission of gross income risk to net income risk, the latter by directly reducing hours risk. To evaluate which policy is preferable, the gains in net income risk reduction on the one hand and the efficiency loss on the other have to be weighed against each other. As a maxim, specific policies are the answers to specific shocks.

The focus of this paper is to gauge the importance of each source of income risk to give an indication which type of risk-alleviating policy - other than insurance - to pursue. If income risk was driven almost entirely by wage risk, devising policies to reduce hours risk would not be a fruitful endeavor. Hence, the aim of this paper is to provide the decomposition of income risk.

In practical terms, we derive a labor supply equation that quantifies the Frisch elasticity, formulate a wage equation, and obtain residuals for hours and wages. We decompose wage risk into permanent (random walk) and transitory (MA(1)) components. Then we separate hours shocks from transitions of wage shocks to hours, obtaining the Marshall elasticity and the permanent and transitory components of hours shocks.

We apply our framework to observations on married men in the US from the Panel Study of Income Dynamics (PSID) over the period 1983 to 1995, since at the end of this period the survey frequency turned bi-annual. In line with the literature, we find that the estimates of permanent and transitory shock variances of wages are close to each other when accounting for measurement error. Our estimate of the Frisch elasticity is 0.28 and our estimate of the Marshall elasticity is -0.70, which is at the lower end of previous findings, a discussion of which is contained in Section 3.5. The result that the permanent variance of hours shocks is virtually zero is robust in all specifications that account for transition of wage shocks to hours. A model abandoning hours shocks fits the data worse and leads to a substantial overestimation - in absolute terms - of the Marshall elasticity.

Our paper is related to studies that decompose total income risk into persistent and transitory components, which derive from ideas by Friedman (1957) and Hall (1978) (see MaCurdy, 1982; Abowd and Card, 1989; Meghir and Pistaferri, 2004; Guvenen, 2007; Blundell et al., 2008; Guvenen, 2009; Hryshko, 2012; Blundell et al., 2016). Abowd and

Card (1989) were pioneers in analyzing the covariance structure of earnings and hours of work. They find that most of the idiosyncratic covariation of earnings and hours of work occurs at fixed wage rates. In contrast, Blundell et al. (2016) restrict the source of risk to stem from wage shocks. We generalize the latter approach and allow for shocks on both wages and hours.

A very closely related paper is Heathcote et al. (2014), who analyze the transmission of wage shocks to hours in a setting where shocks are either fully insurable or not insurable at all (island framework). They derive second hours-wage moments from which they identify variances of shocks as well as the Frisch elasticity of labor supply and risk aversion with respect to consumption. Our study differs in two aspects: First, we assume that there is no perfect insurance market among households, rather shocks are partially insurable as indicated by a consumption insurance parameter similar to Blundell et al. (2008, 2013, 2016). This parameter may differ between individuals. Second, we introduce hours shocks and estimate their variance.

Blundell et al. (2016) estimate the Marshall and Frisch elasticities using hours, income, asset, and consumption data. Similar to them, we allow for partial insurance of permanent wage shocks, but we generalize the approach by allowing for partially insured hours shocks and using hours and income data alone.

3.2. The Model

Individuals maximize the discounted sum of utilities over the lifetime running from t_0 to T:[3]

$$\max_{c_t, h_t} E_{t_0}\left[\sum_{t=t_0}^{T} \rho^{t-t_0} \mathrm{v}(c_t, h_t, b_t)\right], \tag{3.1}$$

where c_t is chosen consumption and h_t chosen hours of work, while b_t contains taste shifters. ρ denotes a discount factor and $\mathrm{v}(\cdot)$ an in-period utility function.

The budget constraint is

$$\frac{a_{t+1}}{(1+r_t)} = (a_t + w_t h_t + N_t - c_t), \tag{3.2}$$

where a_t represents assets, r_t the real interest rate, and N_t non-labor income.

Instantaneous utility takes the additively-separable, constant relative risk aversion (CRRA) form

$$\mathrm{v}_t = \frac{c_t^{1+\vartheta}}{1+\vartheta} - b_t \frac{h_t^{1+\gamma}}{1+\gamma}, \qquad \vartheta < 0, \gamma \geq 0. \tag{3.3}$$

[3] We omit individual-specific subscripts where unambiguous.

An approximation of the first order condition with respect to consumption yields the intertemporal labor supply equation (see MaCurdy (1981), Altonji (1986) and Appendix A):

$$\Delta \ln h_t = \frac{1}{\gamma}\left[-\ln(1+r_{t-1}) - \ln\rho - \eta_t + \Delta \ln w_t - \varsigma \Delta\Xi_t + \Delta \upsilon_t,\right] \tag{3.4}$$

where $\frac{1}{\gamma}$ is the Frisch elasticity of labor supply, w_t is the hourly wage, h_t is hours worked, Ξ_t contains taste shifters, υ_t is a vector of idiosyncratic errors,[4] η_t is a function of the expectation error in the marginal utility of wealth[5], ρ is the discount factor, and r_t is the risk-free real interest rate. γ is identified by estimating equation (3.4) using IV.

The growth rate of wages is determined by human capital related variables X, which contains $\Delta\Xi$, and an idiosyncratic error ω,

$$\Delta \ln w_t = \alpha X_t + \Delta\omega_t \tag{3.5}$$

Idiosyncratic components of hours and wages ω_t and υ_t consist of persistent and transitory components, p_t and τ_t, that follow a random walk and an MA(1)-process respectively, and a measurement error, me_t. E.g., in the case of wages we have:

$$\begin{aligned}
&\omega_{it} = p^{\omega}_{it} + \tau^{\omega}_{it} + \mathrm{me}^{\omega}_{it}\\
&p^{\omega}_{it} = p^{\omega}_{it-1} + \zeta^{\omega}_{it}\\
&\tau^{\omega}_{it} = \theta_{\omega}\epsilon^{\omega}_{it-1} + \epsilon^{\omega}_{it}\\
&\zeta^{\omega}_{it} \sim N\left(0, \sigma^2_{\zeta,\omega}\right), \quad \epsilon^{\omega}_{it} \sim N\left(0, \sigma^2_{\epsilon,\omega}\right)\\
&E\left[\zeta^{\omega}_{t}\zeta^{\omega}_{t-l}\right] = 0, \quad E\left[\epsilon^{\omega}_{t}\epsilon^{\omega}_{t-l}\right] = 0 \quad \forall l \in \mathbb{Z}_{\neq 0}
\end{aligned}$$

Permanent and transitory shocks are uncorrelated. In the case of zero correlation between $\eta_t + \Delta\upsilon_t$ and $\Delta\omega_t$, for example if $\eta_t = 0$, the model parameters $\left(\theta_j, \sigma^2_{\epsilon,j}, \sigma^2_{\zeta,j}\right)_{j\in\{\omega,\upsilon\}}$ are identified through combinations of the autocovariance moments of each shock process.

[4]The term captures for example idiosyncratic taste for work and restrictions in the choice of hours.

[5]$\eta_t = \frac{\varepsilon_{\lambda_t}}{\lambda_t} + \mathcal{O}\left(-1/2(\varepsilon_t/\lambda_t)^2\right)$, i.e., it contains the expectation error of marginal utility of wealth and the approximation error.

3.3. Recovering Labor Supply Elasticities, Wage Shocks, and Hours Shocks

In this section we detail how the labor supply elasticities as well as the standard deviations of permanent and transitory components of wage shocks, ω_t, and hours shocks, υ_t, are recovered in estimation. We abstract from measurement error, which we cover at the end of the section.

Frisch elasticity, hours residuals, and wage residuals — The error term of equation (3.4) is correlated with differenced log wages because wage shocks impact the marginal utility of wealth. To obtain the Frisch elasticity from equation (3.4) we apply instrumental variables based on human capital following MaCurdy (1981). Hours residuals $\Delta\upsilon_t$ are obtained by running IV on differenced log hours using diffenced year, year-of-birth, sex, disability and state dummies as well as an industry and occupational dummy-set, and number-of-kids dummies as covariates. The instruments for differenced log wage are education, education^2, age × education, age × education^2, and the third lag of labor income.

Wage residuals $\Delta\omega_t$ are obtained by estimating equation (3.5), i.e. regressing differenced log wages on the same exogenous regressors as in the hours equation as well as the excluded instruments.

Wage shocks — After recovering $\Delta\omega_t$, all parameters of the autoregressive process, $(\theta, \sigma^2_{\epsilon,\omega}, \sigma^2_{\zeta,\omega})$, are identified through combinations of the autocovariance moments. Label the autocovariance moments by $\Lambda_{\omega,k}$.

$$\Lambda_{\omega,0} = E\left[(\Delta\omega_t)^2\right] = 2\left(1-\theta_\omega+\theta^2_\omega\right)\sigma^2_{\epsilon,\omega}+\sigma^2_{\zeta,\omega}$$

$$\Lambda_{\omega,1} = E\left[\Delta\omega_t\Delta\omega_{t-1}\right] = (1-\theta_\omega)\sigma^2_{\epsilon,\omega}$$

$$\Lambda_{\omega,2} = E\left[\Delta\omega_t\Delta\omega_{t-2}\right] = -\theta_\omega\sigma^2_{\epsilon,\omega}$$

Dividing $\Lambda_{\omega,2}$ by $\Lambda_{\omega,1}$ identifies the parameter θ_ω. Successively, the variance of the transitory shock is identified from $\Lambda_{\omega,1}$ and the variance of the permanent shock from $\Lambda_{\omega,0}$ (see Hryshko, 2012).

Hours shocks — The error term of the labor supply equation contains both taste shocks υ_t and expectation errors η_t. To estimate the variances of $\Delta\upsilon_t$ we need to decompose η_t into parts driven by wage and hours shocks.

Through approximation of the life-time budget constraint, it can be shown that innovations in the marginal utility of wealth are a linear function of permanent income shocks (Blundell et al., 2016), which in our model are given by the term in parentheses:

$$\eta_{it} \simeq \phi^{\lambda}_{it}\left(\left(1+\frac{1}{\gamma}\right)\zeta^{\omega}_{it}+\frac{1}{\gamma}\zeta^{\upsilon}_{it}\right), \quad \phi^{\lambda}_{it} \sim N\left(\mu_{\phi},\sigma^2_{\phi}\right). \tag{3.6}$$

Permanent wage shocks ζ^{ω}_{it} affect income directly (factor 1) and indirectly through labor supply adjustments ($1/\gamma$). Permanent hours shocks ζ^{υ}_{it} affect income only through labor supply adjustments ($1/\gamma$). ϕ^{λ}_{it} measures how the income shocks transfer to η_t, which is in utility units. The parameter is individual-specific since it depends - among other things - on the amount of assets currently held in relation to the total stock of human wealth (see Blundell et al., 2016, p. 396, for the related consumption-insurance parameter). Following Blundell et al. (2016) we make the simplifying assumption that transitory shocks do not impact the marginal utility of wealth. Note that theory predicts that ϕ^{λ}_{it} is positive and thus should follow a distribution with no support on negative values. For small variances relative to the mean, the assumption of a normal distribution can be justified, but we also estimate the model under the assumption that ϕ^{λ}_{it} is lognormally distributed as a robustness test.

The variance of the residuals of the labor supply equation contains both the mean and the variance of ϕ^{λ}_{it} and the variance of the permanent taste shocks, which causes an identification problem. The procedure for wage moments does not carry over and only the variance of transitory shocks can be estimated in the same way:

$$\begin{aligned}\Lambda_{\upsilon,0} = E\left[\left(\eta_t+\Delta\upsilon_t\right)^2\right] = \frac{1}{\gamma^2}\Big(&\gamma^2\sigma^2_{\zeta,\upsilon}+2\gamma^2\left(\theta^2_{\upsilon}-\theta_{\upsilon}+1\right)\sigma^2_{\epsilon,\upsilon}-2\gamma\mu_{\phi}\sigma^2_{\zeta,\upsilon}+\sigma^2_{\phi}\sigma^2_{\zeta,\upsilon}+\gamma^2\,\sigma^2_{\phi}\sigma^2_{\zeta,\omega}\\ &+\mu^2_{\phi}\left(\sigma^2_{\zeta,\upsilon}+(\gamma+1)^2\sigma^2_{\zeta,\omega}\right)+2\gamma\sigma^2_{\phi}\sigma^2_{\zeta,\omega}+\sigma^2_{\phi}\sigma^2_{\zeta,\omega}\Big)\end{aligned} \tag{3.7}$$

Moments $\Lambda_{\upsilon,1}$ and $\Lambda_{\upsilon,2}$ are analogous to their wage process counterparts. To estimate the variance of permanent hours shocks, we need to identify μ_{ϕ} using the contemporaneous covariance of hours and wage residuals:

$$\Lambda_{\omega,\upsilon,0} = E\left[\left(-\eta_t-\Delta\upsilon_t\right)\Delta\omega_t\right] = -\frac{(\gamma+1)\mu_{\phi}\sigma^2_{\zeta,\omega}}{\gamma}. \tag{3.8}$$

This covariance is larger in absolute value for smaller values of γ, which denotes resistance to intertemporal substitution of hours of work, and for larger values of μ_{ϕ}, which denotes the impact of permanent wage shocks on the marginal utility of wealth. If γ goes to infinity, the effect of permanent shocks on income is only mechanical and not through labor supply reactions.

σ^2_ϕ is identified through the cokurtosis moments of the wage and hours residuals. For instance, we can find the following expression, where all additional parameters are identified through wage moments and (3.8):

$$E\left[\left(-\eta_t - \Delta\upsilon_t\right)^2 (\Delta\omega_t)^2\right] - \left(\sigma^2_{\zeta,\omega} + 2\left(\theta^2_\upsilon - \theta_\upsilon + 1\right)\sigma^2_{\epsilon,\omega}\right) E\left[\left(\eta_t + \Delta\omega_t\right)^2\right] = \frac{2(1+\gamma)^2\left(\mu^2_\phi + \sigma^2_\phi\right)\sigma^4_{\zeta,\omega}}{\gamma^2} \tag{3.9}$$

Unfortunately cokurtosis moments are very noisy, hence σ^2_ϕ is only identified with several million observations per cross-section.[6] Therefore, we apply the alternative estimation strategy of calibrating σ^2_ϕ. We find that the estimates of the other structural parameters are robust to a wide range of values of the variance of the transition parameter.

Marshall elasticity — ϕ^λ_{it} is a sufficient statistic for the average Marshall elasticity, the uncompensated reaction to a permanent wage shock.[7]

$$\kappa = \frac{1}{\gamma}\left(1 - \mu_\phi\left(1 + \frac{1}{\gamma}\right)\right) \tag{3.10}$$

The Marshall elasticity is the relevant concept for the evaluation of tax reforms, which are best described as unanticipated, permanent shifts in net-of-tax wages (Blundell and Macurdy, 1999). Using similar considerations as in our study, the Marshall elasticity has been estimated using the covariance of earnings and wages, household sharing parameters, and the ratio of assets to human wealth in Blundell et al. (2016, eq. A2.23). Heathcote et al. (2014) use the covariance of hours and consumption as well as of wages and consumption to estimate the Marshall elasticity. In contrast, we rely only on hours and wage data.

Measurement errors — Following Blundell et al. (2016), we assume that measurement errors are uncorrelated over time, that the variance of the measurement error of hours is $0.23var(h)$, and the variance of the measurement error of wages is $0.13var(w)$, where $var(h)$ and $var(w)$ denote the variances of the residuals of the levels of wages and hours. See Blundell et al. (2016, Appendix 3) for a derivation of these proportions based on findings by Meghir and Pistaferri (2004). As we show in the section 3.5, ignoring measurement errors leads to larger estimates of the variance of the transitory shocks. We adjust the moments described above for measurement errors.

[6] Simulations evidencing this are available upon request from the authors.
[7] See Keane (2011, p.1008) for a discussion of why reactions to permanent shocks equal the Marshall elasticity.

Estimation —We estimate the parameters of the autoregressive processes and the transition of wage shocks by fitting the theoretical moments $\{\Lambda_{\omega,k}, \Lambda_{\upsilon,k}, \Lambda_{\omega,\upsilon,k}\}$ to those of the data. The parameters Θ are identified using the method of minimum distance and an identity matrix as the weighting matrix.[8] The distance function is given by

$$DF(\Theta) = [m(\Theta) - m^d]' I [m(\Theta) - m^d], \tag{3.11}$$

where $m(\Theta)$ indicates theoretical moments and m^d empirical moments. Standard errors of the estimates are obtained by the Delta method. An outline of the entire estimation procedure is detailed in Hryshko (2012).

3.4. The Data

We use annual data on the US from the PSID for the survey years 1983 to 1995. After this point in time the PSID is bi-annual. Our sample consists of working married males aged 28 to 60, who are main earners of their respective households. We eliminate individuals with unrealistically high or low values for annual hours of work or unrealistically high changes in work hours and wages.[9] Table 3.1 shows summary statistics of the main sample. Monetary variables are adjusted to 2005 real values using the CPI-U.

Table 3.1.: Descriptives

	mean	sd	min	max
Age	42.36	8.85	28	60
Annual hours of work	2266.34	436.62	784	3996
Hourly wage	28.80	13.36	4.21	99.30
Number of kids in household	1.43	1.20	0	6
Asset income	2166.48	12213.91	0	601000
N	6544			

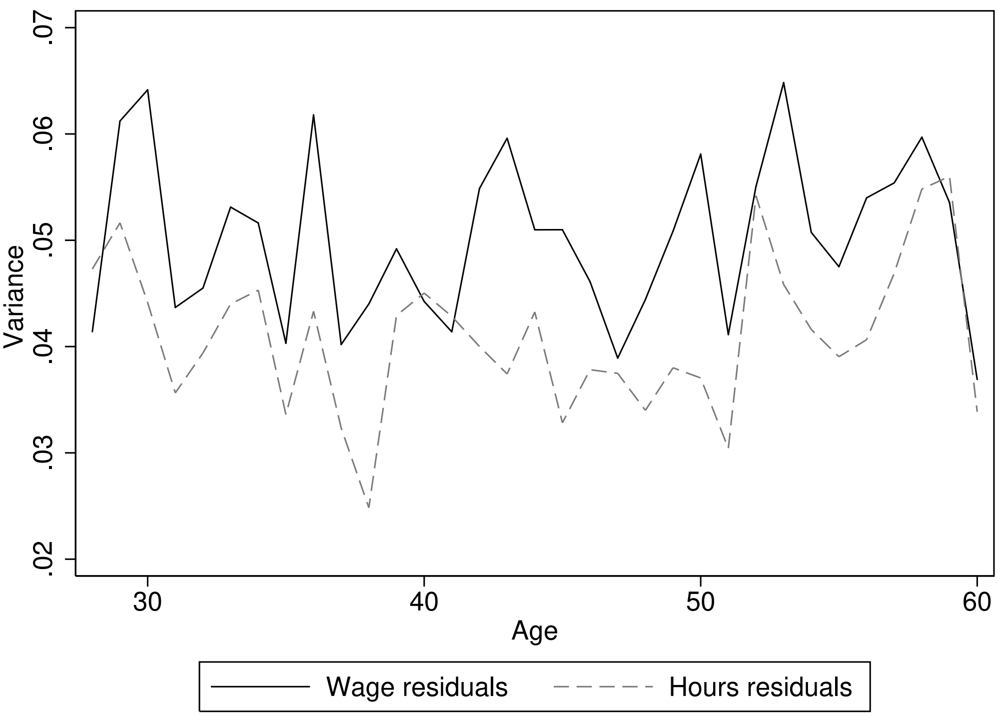
Source: Own calculation based on the PSID.

Figure 3.1 shows the variance of the differenced wage and hours residuals over the life cycle for the main sample. It is apparent that the residuals of hours and wages comove substantially, but not perfectly. This implies that there is substantial transmission between wage shocks and hours, but wage variation is not the end of the story. Rather, a substantial part of the idiosyncratic variation in hours of work results from hours shocks.

[8] Altonji and Segal (1996) show that the identity weighting matrix is preferable for the estimation of autocovariance structures using micro data.

[9] Following Domeij and Flodén (2006), we eliminate individuals with an increase in hours or wages of more than 250 percent, a decrease of more than 60 percent, wages outside the range of 4 to 100, or more than 4000 annual hours of work.

Figure 3.1.: Variance of Differenced Wage and Hours Residuals over the Life-Cycle



Note: Variances of residuals obtained from the regressions of equations (3.4) and (3.5) for the main sample. Cohort effects are not accounted for. *Source:* Own calculation based on the PSID.

Similar series of hours, but without detrending, appear in Kaplan (2012). Kaplan characterizes the pattern of log hours as "strongly U-shaped". We cannot make the same determination for the detrended series.

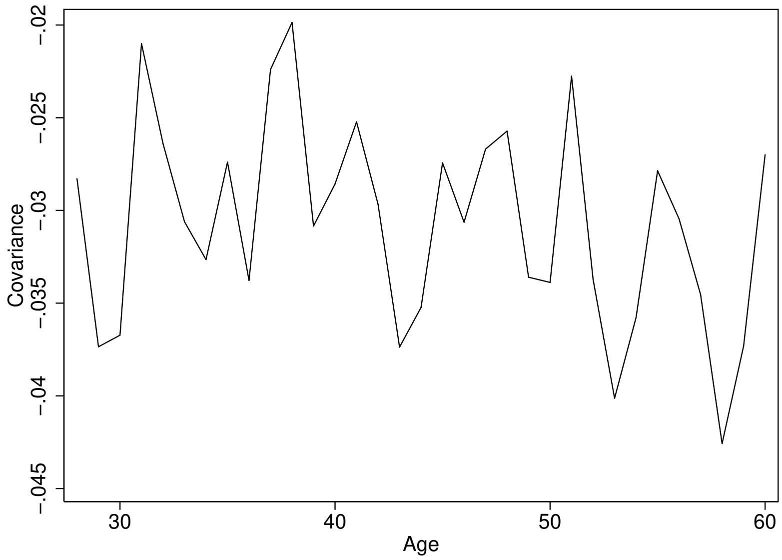
Figures 3.2 and 3.3 show the covariance and the correlation of differenced hours and wage residuals over the life-cycle. As implied by the model, it is substantially negative, as positive wage shocks have a negative impact on the marginal utility of wealth and thus on hours of work residuals. Apart from a weaker correlation at the beginning of working life, no clear pattern is obvious.

From eyeballing both the hours and covariance series, variation in the variance of wages carries over at a roughly constant rate into the covariance of hours and wages.

3.5. Results

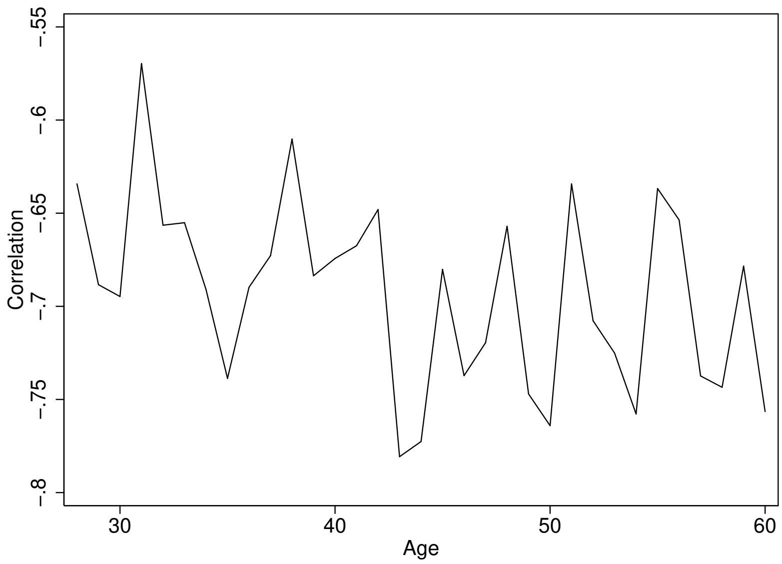
Frisch elasticity — Table 3.2 reports the estimates of the labor supply equation (3.4). In contrast to the most closely related papers (Blundell et al., 2016; Heathcote et al., 2014),

Figure 3.2.: Covariance of Differenced Wage and Hours Residuals over the Life-Cycle



Source: Own calculation based on the PSID.

Figure 3.3.: Correlation of Differenced Wage and Hours Residuals over the Life-Cycle



Source: Own calculation based on the PSID.

we obtain the Frisch elasticity directly through IV estimation and not through covariance moments. The estimated Frisch elasticity for the main sample is 0.28, which is in line with the literature (Keane, 2011). We show results for four sub-samples. The more educated group (II), which excludes high school dropouts, exhibits exactly the same Frisch elasticity as the main sample. Domeij and Flodén (2006) show that liquidity constraints can bias the Frisch elasticity downwards. Therefore we analyze two subsamples that are likely to differ from the main sample in this regard: Individuals with positive asset income (III) are less likely to be credit constrained than others and relatively young individuals (IV) have accumulated less assets on average. The point estimates of the elasticity are as expected: Younger individuals have a slightly lower Frisch elasticity than the main sample and the elasticity of the group with positive asset incomes is higher than the one of the main sample. Finally, we exclude individuals with children in the household (V). Children hinder the adjustment of hours of work and the childless subsample has a higher Frisch elasticity than the main sample.

Table 3.2.: Frisch Labor Supply Equation Estimation

	I	II	III	IV	V
	Main sample	No Drop-outs	Pos. asset income	Age 30-50	No children
$\Delta \ln(wage)$	0.280	0.280	0.352	0.182	0.293
	(0.121)	(0.128)	(0.224)	(0.134)	(0.244)
N	6544	5837	2649	4177	1414
Shea's R^2	0.005	0.005	0.005	0.005	0.006

Standard errors in parentheses
Source: Own calculation based on the PSID.

Standard deviations of wage shocks —Table 3.3 reports the standard deviations of permanent and transitory wage shocks as well as the parameter of shock persistence for the same subsamples as before. The numbers of observations are reported in the previous table. In Columns I-V we account for measurement errors as described previously. Column I shows results for the main sample, in line with Blundell et al. (2016) we find that variances of permanent and transitory shocks are very similar. The standard deviations correspond to variances of about 0.017 for transitory and 0.013 for permanent shocks.[10] While results for all subsamples are similar, transitory shocks seem to play a smaller role for individuals with positive asset income as standard shock sizes and the persistence of transitory shocks are smaller for this group. In contrast, the standard deviation of permanent shocks is slightly larger than for other groups. Column VI shows results not correcting for measurement errors, which leads to larger standard deviations of transitory shocks.

Standard deviations of hours shocks – The estimation procedure relies on the calibration of the variance of ϕ^{λ}_{it}. In Figure 3.4 we show resulting estimates of the standard

[10]We use annual data, Blundell et al. (2016) use bi-annual data and therefore their variance estimates are almost exactly twice as large as ours.

Table 3.3.: Autoregressive Process - Wages

	I	II	III	IV	V	VI
	Main sample	No Drop-outs	Pos. asset income	Age 30-50	No children	With M.E.
θ_{ω}	0.1579	0.1671	0.0806	0.1316	0.1719	0.1337
	(0.0484)	(0.0510)	(0.1325)	(0.0599)	(0.1167)	(0.0417)
$\sigma_{\epsilon,\omega}$	0.1321	0.1299	0.1062	0.1280	0.1290	0.1441
	(0.0078	(0.0083)	(0.0160)	(0.0104)	(0.0226)	(0.0073)
$\sigma_{\zeta,\omega}$	0.1149	0.1146	0.1256	0.1135	0.1023	0.1149
	(0.0081)	(0.0084)	(0.0121)	(0.0101)	(0.0223)	0.0081)

Delta method standard errors in parentheses
Source: Own calculation based on the PSID.

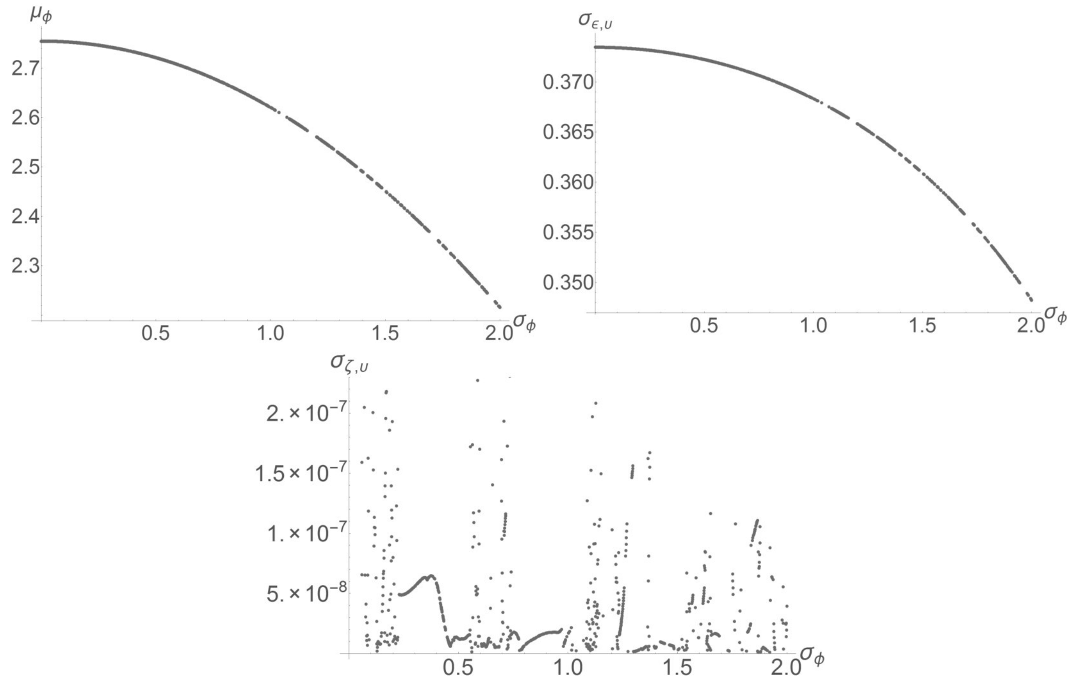
deviation and persistence parameter of transitory hours shocks and the mean of the transition parameter for the main sample using different calibrations. The resulting estimates are not very sensitive to this. Results appear robust to the chosen calibration.

We report estimates for the hours process in Table 3.4. σ_{ϕ} is set to 0.3162 , which corresponds to a variance of 0.1. We find that the permanent variance is essentially zero in all subsamples, making hours shocks entirely transitory. A likely reason is that permanent hours shocks such as serious injuries in practice affect the extensive margin, which is not captured by the model.[11] Transitory hours shocks are common adjustments, due to brief, but pressing needs of, e.g., children or elderly relatives. An example for a positive hours shock is the need to work overtime because of project deadlines or sudden increases in demand, which in the long-run are met with additional hires. In contrast, wage shocks due to obsolescence of human capital or newly acquired skills, can be permanent.

For the main sample the persistence parameter of the transitory process in hours is smaller than the one for the wage process, but not by a large margin. Variation in the persistence parameter along the samples is largely consistent with variation of the persistence parameter for the wage process, except for the young sample (column IV) which has the highest persistence of hours shocks. Children might play a role in raising the persistence parameter for this sample. Changes of estimates of the standard deviations of transitory shocks along the samples lend themselves to intuitive interpretations. In column III we observe a smaller estimate, which is consistent with the notion that the well-situated population in sample III works in less risky jobs. The estimate of the standard devation is larger for the younger population (column IV) than for the main sample. Younger individuals are subject to more hours risk possibly due to the necessity of childcare or care for the elderly, which likely plays into the size of the estimate. The fact that individuals without children (column V) have smaller shock sizes is consistent with this explanation.

[11] Modeling the extensive margin is particularly important when analyzing the importance of shocks for female labor supply.

Figure 3.4.: Parameters of the Hours Process for Differently Calibrated Values of σ_ϕ



Source: Own calculation based on the PSID.

The larger the parameter μ_ϕ, the stronger the transition of permanent shocks to the marginal utility of wealth. We expect households with larger accumulation of wealth relative to human wealth to exhibit smaller values of μ_ϕ. This is largely confirmed in estimation, with the lowest estimate observed in sample III and the highest in sample IV. The former sample is likely to have accumulated a large amount of assets and the latter sample consists of relatively young people, whose human wealth is comparatively large. In column VI, we do not account for the negative correlation in the measurement errors in hours and wages. This leads to a larger estimate of μ_ϕ.

Marshall elasticity—Table 3.5 reports the Marshall elasticity calculated using equation (3.10). It is similar for all subsamples, between -0.60 and -0.70, except for people without children, where it is even more negative. People without children can more readily adjust hours, evidenced by a smaller estimate of the resistance to intertemporal substitution (and thus a larger Frisch elasticity), which drives this different result. Blundell et al. (2016) and Heathcote et al. (2014) both find negative Marshall elasticities for men (-0.08 and -0.35 respectively). Ziliak and Kniesner (2005) report a similarly large negative Marshall elasticity of -0.47. Altonji et al. (2013) report a coefficient that determines "the response to a relatively permanent wage change" of -0.084.

Table 3.4.: Autoregressive Process - Hours

	I	II	III	IV	V	VI
	Main sample	No Drop-outs	Pos. asset income	Age 30-50	No children	With M.E.
θ_{υ}	0.1437	0.1564	0.0597	0.1954	0.1170	0.0745
	(0.0021)	(0.0019)	(0.0073)	(0.0045)	(0.0290)	(0.0018)
$\sigma_{\epsilon,\upsilon}$	0.3730	0.3712	0.2565	0.4134	0.3371	0.4278
	(0.0020)	(0.0022)	(0.0032)	(0.0020)	(0.0160)	(0.0018)
$\sigma_{\zeta,\upsilon}$	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
μ_{ϕ}	2.7413	2.5328	1.9935	3.7669	3.1952	2.9594
	(0.0046)	(0.0047)	(0.0062)	(0.0328)	(0.0390)	(0.0041)
σ_{ϕ}	0.3162	0.3162	0.3162	0.3162	0.3162	0.3162

Delta method standard errors in parentheses
Source: Own calculation based on the PSID.

Our approach to estimating the Marshall elasticity differs from previous works. Blundell et al. (2016) account for taxes. Their results in the no-tax model are virtually the same as the results for responses to before-tax wage changes in the model accounting for taxes. They also perform several sensitivity checks to alterations in the model such as the structure of preferences, whether shock variances vary over time, and the existence of outside insurance. These alterations do not change their estimates substantially. An important conceptual difference is that we do not account for secondary earners. Blundell et al. (2016) find that wage shocks to first and secondary earners are positively correlated and that women have a positive Marshall elasticity, which further amplifies the impact of shocks on the marginal utility of wealth of the household. In our approach, the estimated impact of shocks on the wage of the primary earner includes the correlation with shocks on the secondary earner's wage and labor supply reactions of the secondary earner. Controlling for these factors should lead to an estimated Marshall elasticity closer to zero. A second important difference is that Blundell et al. (2016) use data on assets and human wealth, the Frisch elasticity of labor supply, the price elasticity of consumption, and parameters that are specific to spousal labor supply to calculate the Marshall elasticity. In contrast, we obtain the Marshall elasticity directly from the covariance of hours residuals and wage shocks, using labor supply and wage data alone. We suspect that this is the main source of the discrepancy. Agents might not use the disposable information perfectly, e.g., because they use a rule of thumb (Galí et al., 2004).

Table 3.5.: Marshall Elasticity

	I	II	III	IV	V	VI
	Main sample	No Drop-outs	Pos. asset income	Age 30-50	No children	With M.E.
κ	-0.7011	-0.7012	-0.5960	-0.6292	-0.9177	-0.7792

Source: Own calculation based on the PSID.

Importance of hours and wage shocks — To illustrate the relative importance, we calculate the contribution of hours and wage shocks of the magnitude of one standard deviation on the change in income. A permanent change in log wages of the size of one standard deviation of permanent wage shocks (0.1149) translates into income directly and through the Marshall elasticity. The total impact on log income is thus $(1-0.7011)0.1149 = 0.0343$, i.e., the impact on annual income is 3 percent. For the average annual labor income of 65,270 Dollar, this implies an increase of about 2,281 Dollar.

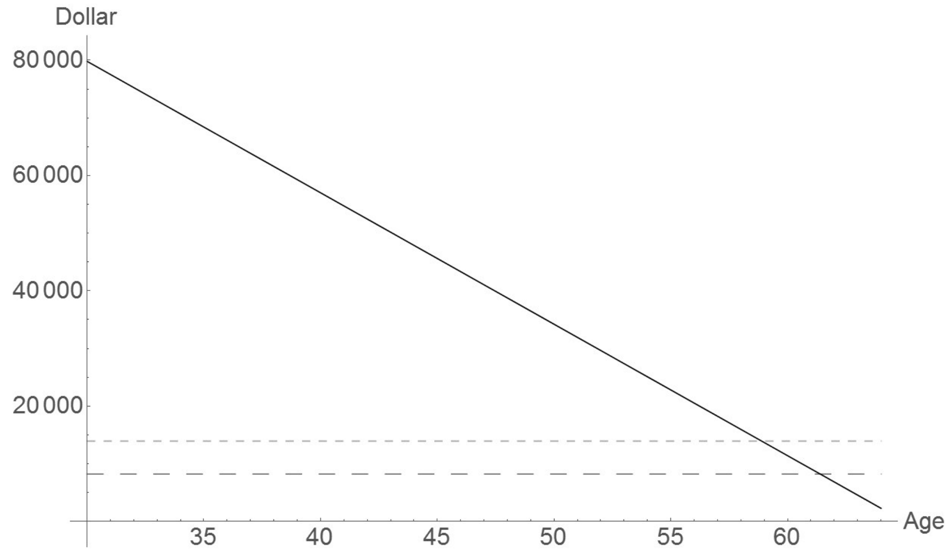
Transitory wage shocks impact labor income directly and through the Frisch elasticity but do not impact the marginal utility of wealth. Their effect on log income is thus given by $1+1/\gamma$. The standard deviation of transitory wage shocks is 0.1321 with the impact on log annual income of a transitory wage shock being $(1+0.28)0.1321 = 0.1691$. This corresponds to about 12,024 Dollar at mean income. The impact of transitory shocks on annual income is substantially larger than the impact of permanent shocks because permanent shocks are cushioned by negative labor supply adjustments.

Transitory hours shocks impact average log income only through the Frisch elasticity. Again, using their standard deviation, the impact of a typical hours shock on log income is 0.1044, which corresponds to about 7,186 Dollar at the mean.

Thus, most of in-period uncertainty is due to transitory wage shocks followed by transitory hours shocks. In comparison, the short-term impact of permanent wage shocks is relatively small. In contrast, the impact on *life-time* income depends on how many years the individual remains in the labor market after the shock. Consider a 42 year old individual, who retires at 65. If the individual receives a typical permanent log wage shock of 0.1306, which corresponds to a permanent shock on annual income of 2,281 Dollar, lifetime income increases by $2,281*(65-42) = 52,453$ Dollar. Thus, at a relatively young age, the impact of permanent wage shocks is substantial. However, as individuals age, the importance of permanent shocks declines. In contrast, the impact of transitory wage shocks is given by $(1+\theta_{\omega})$ times the direct shock, 13,938 Dollar. Equivalently, the impact of transitory hours shocks on lifetime income is 8,218 Dollar. Thus, the impact of transitory hours shocks on income is about 59 percent of the impact of transitory wage shocks and should not be neglected. Figure 3.6 plots the life-time impact of typical shocks, as defined before, on life-time income for different ages of an individual earning the average annual income (Table 3.1) and retiring at age 65. The impacts of the two types of wage shocks (solid black line: permanent shock; short-dashed line: transitory shock) are of roughly equal importance at age 59. Equivalence between permanent wage shocks and transitory hours shocks occurs at age 61. Clearly, depending on the time an individual will remain active on the labor market, permanent shocks may dwarf other types of shocks or be roughly equal to them.

Hours shocks and transition in alternative models — In Table 3.6 we report the autoregressive parameters of hours shocks of the main sample under various restrictions.

Figure 3.6.: Impact of Typical Shocks on Life-Time Income



Note: We calculate the life-time impact of different shocks at mean levels of income (Table 3.1). The black solid line gives the life-time impact on income of a permanent wage shock of one standard deviation at different ages. The short-dashed line gives the impact of a transitory wage shock and the long-dashed line that of a transitory hours shock of one standard deviation. Figures are in real 2005 dollars. *Source:* Own calculation based on the PSID.

In column II, the variance of permanent hours shocks is restricted to zero. The parameter estimates are virtually the same as in the main specification reported again in column I for comparison, were the variance of permanent shocks is estimated to be very close to zero. Thus no major omission occurs when permanent hours shocks are ignored. In column III, both the variance of permanent shocks and the variance of the transition parameter are set to zero. Parameter estimates remain almost unchanged, showing the robustness of the results. In column IV, hours residuals consist only of transitory shocks, i.e., there is no transmission of wage shocks to hours and there are no permanent hours shocks. As expected, this leads to a substantial overestimation of the importance of transitory shocks and their persistence. In column V, only the transmission from wages to hours is restricted to zero. This leads to the estimation of substantial permanent variation in hours, which roughly coincides with the standard deviation of permanent wage shocks times $\mu_\phi(1+1/\gamma)$, lending support to the structure of the labor supply model.

In column VI we report estimates where hours shocks are restricted to zero, which is common in the literature. The implied transition parameter is substantially larger, 4.3128, which implies a Marshall elasticity of -1.2634. A Marshall elasticity smaller than -1 implies that negative permanent wage shocks lead to an *increase* in labor income because the increase in labor supply *over*compensates for the decrease in wages. This seems implausible and is further evidence for the importance of allowing for hours shocks.

The distribution of ϕ_{it}^{λ} has so far been assumed to be normal. However, one might argue that ϕ_{it}^{λ} is drawn from a distribution that has no support on negative values because positive wage or hours shocks should always have a negative impact on the marginal utility of wealth. An obvious choice is the lognormal distribution. Under these distributional assumptions the moments change in a straight-forward manner. The variance of hours residuals is given by:

$$E\left[\left(-\eta_t - \Delta\upsilon_t\right)^2\right] = \frac{1}{\gamma^2}\left(\gamma^2\sigma_{\zeta,\upsilon}^2 + 2\gamma^2\left(\theta_\upsilon^2 - \theta_\upsilon + 1\right)\sigma_{\epsilon,\upsilon}^2 - 2\gamma e^{\mu_{ln(\phi)}+\frac{\sigma_{ln(\phi)}^2}{2}}\sigma_{\zeta,\upsilon}^2 \right. \tag{3.12}$$
$$\left. + e^{2\left(\mu_{ln(\phi)}+\sigma_{ln(\phi)}\right)}\sigma_{\zeta,\upsilon}^2 + (\gamma+1)^2 e^{2\left(\mu_{ln(\phi)}+\sigma_{ln(\phi)}^2\right)}\sigma_{\zeta,u}^2\right)$$

The contemporaneous covariance of hours and wage residuals is

$$E\left[\left(-\eta_t - \Delta\upsilon_t\right)\Delta\omega_t\right] = -\frac{(\gamma+1)e^{\mu_{ln(\phi)}+\frac{\sigma_{ln(\phi)}^2}{2}}\sigma_{\zeta,\omega}^2}{\gamma}. \tag{3.13}$$

The estimates are reported in column VII in Table 3.6 and are qualitatively similar to those of the main model except for a drop in the persistence of transitory shocks. The results appear robust to this change in the distributional assumption.

Below all estimates, we report the value of the distance function. Accounting for the transmission of wage shocks to hours leads to a substantial improvement of the model fit. The value of the distance function is highest for the model without hours shocks, implying the worst fit.

Table 3.6.: AR Hours Estimation in Alternative Models

	I Main sample	II $\sigma_{\zeta,\upsilon}=0$	III $\sigma_\phi = 0$, $\sigma_{\zeta,\upsilon}=0$	IV $\phi = 0$, $\sigma_{\zeta,\upsilon}=0$	V $\phi=0$	VI $\sigma_{\epsilon,\upsilon} = 0$, $\sigma_{\zeta,\upsilon}=0$	VII $\ln\phi \sim N(\mu_{ln(\phi)}, \sigma_{ln(\phi)})$
θ_υ	0.1437	0.1437	0.1443	0.2446	0.1864	0	0.0847
	(0.0021)	(0.0021)	(0.0021)	(0.0014)	(0.0018)	-	(0.0017)
$\sigma_{\epsilon,\upsilon}$	0.3730	0.3730	0.3735	0.4906	0.4134	0	0.4284
	(0.0020)	(0.0020)	(0.0020)	(0.0013)	(0.0020)	-	(0.0017)
$\sigma_{\zeta,\upsilon}$	0.0000	0	0	0	0.3607	0	0.0000
	(0.0000)	-	-	-	(0.0008)	-	(0.0000)
μ_ϕ	2.7413	2.7413	2.7547	0	0	4.3128	2.7022†
	(0.0046)	(0.0046)	(0.0046)	-	-	(0.0118)	(0.0042)
σ_ϕ	0.3162	0.3162	0	0	0	0.3162	0.3162
$DF(\Theta)$	0.1635	0.1635	0.1633	0.2407	0.2230	0.4910	0.1861

† We report the mean of ϕ and not $\mu_{ln(\phi)}$ for comparability.
Delta method standard errors in parentheses
Source: Own calculation based on the PSID.

3.6. Conclusion

We have decomposed idiosyncratic income uncertainty into contributions of transitory and permanent wage and hours shocks. We find that, first, permanent shocks in hours are virtually non-existent. Second, permanent wage shocks play the most important role in determining life-time income uncertainty. Third, transitory wage shocks are more important than transitory hours shocks, but both have an economically significant impact. Our findings regarding wage shocks are very similar to those of the extant literature, but previous papers have ignored hours shocks.

The important conclusion is that wage shocks are the main culprit of life-time income uncertainty. Therefore, policies aimed at reducing income uncertainty should focus on wages. For instance, improvements in health care and other policies aimed at the flexibility of labor supply are of little relevance in this regard. In contrast, policies that reduce wage uncertainty, such as a progressive taxes or policies stabilizing the production capabilities of firms might have stronger welfare effects.

The second main contribution is a new method to estimate the Marshall elasticity. Our method allows to calculate the Marshall elasticity from the covariance of hours and wage shocks. In a first step, we estimated a Frisch elasticity of labor supply of 0.28, which is in line with previous findings (see Keane, 2011), but slightly lower than recent results by Blundell et al. (2016). Our baseline specification implies a Marshall elasticity of -0.70, slightly more negative than most previous estimates. Efforts to reconcile our finding with the extant literature should in a first step focus on the role of the secondary earner, but the probable route for success is to formulate an exhaustive and flexible model of the transition of shocks to the marginal utility of wealth.

3.7. Appendix

A. Derivation of the Labor Supply Equation

We specify the taste shifter $b_t = \exp(\varsigma\Xi_t - \upsilon_t)$. Ξ_t ia a set of personal characteristics. υ_t is an idiosyncratic disturbance with mean zero that captures taste shocks like unexpected changes in childcare needs, sickness, and other unexpected changes in the disutility of labor supply. The residual in the labor supply equation consists of in-period taste shocks and expectations corrections in the marginal utility of wealth due both to wage and hours shocks.

The first order condition of the consumer's problem w.r.t. h_t is:

$$\frac{\partial \mathscr{L}}{\partial h_t} = E_t\left[\left(-b_t h_t^{\gamma}\right) + \lambda_t\left(\frac{\partial w_t}{\partial h_t} h_t + w_t\right)\right] = 0, \tag{3.14}$$

where $\lambda_t = \frac{\partial u(c_t,h_t,b_t)}{\partial C_t}$ denotes the marginal utility of wealth. The Euler equation of consumption is given by

$$\frac{1}{\rho(1+r_t)}\lambda_t = E_t[\lambda_{t+1}]. \tag{3.15}$$

Expectations are rational, i.e., $\lambda_{t+1} = E[\lambda_{t+1}] + \varepsilon_{\lambda_{t+1}}$, where $\varepsilon_{\lambda_{t+1}}$ denotes the mean-zero expectation correction of $E[\lambda_{t+1}]$ performed in period $t+1$. Expectation errors are caused by innovations in the hourly wage residual ω_{t+1} and innovations in taste shocks υ_{t+1}, which, as implied by rational expectations, are uncorrelated with $E_t[\lambda_{t+1}]$. Rational expectations imply that $\varepsilon_{\lambda_{t+1}}$ is uncorrelated over time, so that regardless of the autocorrelative structure of the shock terms, $\varepsilon_{\lambda_{t+1}}$ will only be correlated with the innovations of the shock processes.

Resolving the expectation operator in equation (3.14) and using the definition of w_t yields

$$b_t h_t^\gamma = \lambda_t w_t. \tag{3.16}$$

Taking logs of both sides we arrive at the first structural equation

$$\ln h_t = \frac{1}{\gamma}(\ln\lambda_t + \ln w_t - \ln b_t). \tag{3.17}$$

To find an estimable form for $\ln h_t$, we take logs of (3.15) and resolve the expectation:

$$\ln\lambda_t = \ln(1+r_t) + \ln\rho + \ln(\lambda_{t+1} - \varepsilon_{t+1})$$

A first order Taylor-expansion of $\ln(\lambda_{t+1} - \varepsilon_{t+1})$ gives $\ln(\lambda_{t+1}) + \frac{\varepsilon_{\lambda_{t+1}}}{\lambda_{t+1}}$, leading to the expression

$$\ln\lambda_t = \ln(1+r_t) + \ln\rho + \ln(\lambda_{t+1}) + \frac{\varepsilon_{\lambda_{t+1}}}{\lambda_{t+1}} + \mathcal{O}\Big(-1/2(\varepsilon_{t+1}/\lambda_{t+1})^2\Big). \tag{3.18}$$

Accordingly, when we backdate 3.18, we can remove $\ln\lambda_t$ by first differencing 3.17.

4. How Important is Precautionary Labor Supply?[1]

4.1. Introduction

This study quantifies the importance of precautionary labor supply, defined as the difference between hours supplied in the presence of risk and hours supplied under perfect foresight. Facing a higher future wage risk, individuals may increase their hours worked in order to insure themselves against bad realizations. Our study provides empirical evidence for this theoretically predicted phenomenon. We examine how strongly labor supply adjusts in response to higher wage risk by focusing on the partial equilibrium case similarly to Carroll and Samwick (1998) or Parker and Preston (2005) for consumption.

A thorough intuition of labor supply incentives over the life cycle is crucial for understanding household behavior and is of primary interest for both labor economics and macroeconomics (Meghir and Pistaferri, 2011). Relevant precautionary labor supply could explain differences in hours worked across occupations or why self-employed work more hours than employees for a given wage. The extent of precautionary labor supply is key for various policy issues, for instance the optimal design of social security programs. Our approach allows us to calculate how labor supply would change in partial equilibrium, if self-employed, blue and white collar workers had the same insurance against wage risk as civil servants, for instance through reforms of the social insurance system.

A number of theoretical contributions have studied precautionary labor supply in models with saving (Flodén, 2006; Low, 2005; Pistaferri, 2003). These studies find that individuals facing higher wage risk work more at the beginning of working life in order to accumulate savings. This behavior is governed by the curvature in consumption, i.e. prudence as defined in Kimball (1990), and in leisure of workers' preferences. When leisure is low, not only the marginal utility of leisure is higher, but also the rate at which the marginal valuation rises when leisure falls. This implicates the precautionary motive because of which individuals save more in anticipation of higher future wage risk. With flexible labor supply they do so by consuming less or by working more. The latter concept is precautionary labor supply. Pijoan-Mas (2006) shows that additional hours of work are a quantitatively important smoothing device in a calibration exercise. In our analysis, we abstract from general equilibrium effects which need to be taken into account to assess

[1]This chapter is based on Jessen et al. (2016).

whether the effect of uncertainty on aggregate output is positive or negative. Marcet et al. (2007) demonstrate that under reasonable parameter configurations a wealth effect that reduces labor supply may dominate the positive precautionary saving effect on aggregate output documented in Aiyagari (1994) and Huggett (1993). Prior to this, studies like Block and Heineke (1973); Eaton and Rosen (1980a,b), and Hartwick (2000) predicted theoretically that the relationship between uncertainty and labor supply is positive.[2] Still, the actual importance of precautionary labor supply remains an empirical question.

This paper is one of the few studies that provide empirical evidence on this issue. Pistaferri (2003) finds that the effect of wage risk on labor supply agrees with theoretical predictions, but is economically negligible. This might be due to the fact that Pistaferri (2003) used data collected only every two years for Italy in 1989, 91, and 93. In contrast, we are able to construct growth rates from year to year and to exploit a relatively long time dimension (from 2001 to 2012) of the German Socio-Economic Panel (SOEP).

The relationship between (proxies for) wage risk and hours of work has been documented to be positive for self-employed men in the US (Parker et al., 2005), male employees in the US who work more than 30 hours per week (Kuhn and Lozano, 2008), and for German and US workers (including self-employed) of both sexes (Bell and Freeman, 2001). Benito and Saleheen (2013) show that men and women use hours worked to shield themselves against financial shocks, which the authors define as deviations in the subjective perception of their own financial situation, compared to their expectation from the previous year. We contribute to the literature with several innovations.

First, we specify a dynamic labor supply model that allows for partial adjustment of hours worked. Such a specification reflects constraints in the workers' capacity to adjust immediately to their desired level of labor supply. Our findings reject the immediate adjustment model used in previous work.

Second, we calculate marginal net wages using the tax-transfer-microsimulation model STSM (see Steiner et al., 2012).[3] Therefore, in contrast to the previous studies, we are able to account for partial insurance of wage risk through the tax and transfer system as well as through the social insurance system, which may be an important determinant of precautionary behavior, as argued, e.g., in Fossen and Rostam-Afschar (2013). Bell and Freeman (2001) surmise that "[s]ince we have not taken into account differences in the level of social safety nets or taxation [...] our analysis probably understates the effect of inequality in economic rewards on work time". Our results show that this effect is very small.

Third, the result of Pistaferri (2003) that precautionary labor supply is irrelevant might be due to the fact that he used *subjective information on future income* (see also Mastrogiacomo and Alessie, 2014). We examine several measures for wage risk and do not find relevant precautionary labor supply using subjective risk measures either. If

[2] See Menezes and Wang (2005) for a study that predicts a negative effect of increased wage uncertainty on labor supply if the substitution effect dominates the income effect.

[3] The Steuer-Transfer-Mikrosimulationsmodell (STSM) is comparable to FORTAX for the UK (Shephard, 2009) or TAXSIM for the US (Feenberg and Coutts, 1993).

wage risk is—as in our analysis—measured by the standard deviation of past hourly individual net wages, however, precautionary labor supply becomes relevant. Moreover, this does not change if risk from other sources than own wages is included or future wages are used. Most of our measures of wage risk assume—following e.g. Blundell and Preston (1998), Blundell et al. (2008) or Carroll and Samwick (1998) for income—that information unknown to the econometrician is unpredictable for the worker as well.

Fourth, in addition to wage risk and in contrast to previous studies, we investigate the effect of unemployment probability calculated similarly as in Carroll et al. (2003). We find that unemployment probability also increases labor supply, but is quantitatively less important than wage risk.

Finally, we are the first to quantify precautionary labor supply empirically. Individuals in the main sample choose an additional 2.8% of their hours of work to shield against wage shocks, i.e. about one week per year. Precautionary labor supply is particularly important for the self-employed, a group that faces average wage risks substantially above the sample mean. This group works 6.2% of their hours because of the precautionary motive. If self-employed faced the same wage risk as the median civil servant, their hours of work would reduce by 4.5%.

The next section describes our dataset and construction of the measure of wage risk and probability of unemployment. Section 4.3 presents our empirical specification and the estimation methods. Section 4.4 discusses the main results and occupation specific findings. In Section 4.5 we quantify the importance of precautionary labor supply, Section 4.6 shows that the results are robust, and Section 4.7 concludes.

4.2. Data

Our study uses data from the SOEP (version 30), a representative annual panel survey in Germany. Wagner et al. (2007) provide a detailed description of the data. We use observations from 2001-2012 and focus on men because the extensive margin plays an important role in women's labor supply decisions.The sample is restricted to married men between 25 and 56 years old and working at least 20 hours to allow comparisons with the canonical labor supply literature, for example, Altonji (1986), and MaCurdy (1981).[4] Further, we drop persons who indicated having received social welfare payments because their hours choices are likely driven by institutional constraints rather than precautionary motives. We restrict our sample to individuals working less than 80 hours per week. In total, we observe the main wage risk measure for 10,987 data points from 2,488 persons.[5]

[4]Including workers with less than 20 weekly hours virtually does not affect the results.

[5]Table A1 in the Appendix summarizes the number of observations lost due to each sample selection step.

Marginal net wage According to economic theory, individuals' labor supply responds to the *marginal* net wage. The reason is that at the optimum the marginal rate of substitution equals the marginal rate of transformation. The marginal net wage is the price at which leisure is transformed into consumption.

To construct the marginal net wage, first we calculate the hourly gross wage w_{it}^{gross} by dividing annual gross labor income y_{it} by annual hours of work h_{it}:

$$w_{it}^{\text{gross}} = \frac{y_{it}}{h_{it}}.$$

We calculate net income using the microsimulation model STSM. Jessen et al. (2017b) present a comprehensive overview of marginal tax rates for different households (for more information, see Steiner et al., 2012). We obtain marginal net wage rates by scaling the gross wage w_{it}^{gross} with the marginal net-of-tax rate. Define the net-of-tax rate as the net of tax income per Euro of additional pretax income due to an increase in hours of work. Then the marginal hourly net wage is given by:

$$w_{it} = \text{Net-of-tax rate} \times w_{it}^{\text{gross}} = \frac{NetInc(y_{it} + \Delta y_{it}) - NetInc(y_{it})}{\Delta y_{it}} w_{it}^{\text{gross}}. \tag{4.1}$$

$NetInc(y_{it})$ denotes net income given gross income y_{it}. To calculate the net-of-tax rate we increase each person's annual labor income y_{it} marginally.[6] In practice, the relevant concept is the net of tax income per additional time spent on work. We assume that this coincides with the marginal net wage as calculated in equation (4.1). This is true if additional hours of work are fully compensated.

For the calculation of hourly wages we use *paid* hours because an increase in these translates directly into an increase in income. To construct paid hours we follow Euwals (2005), accounting for differences in compensation of overtime hours.[7]

Wage Risk We construct measures for both gross and marginal net wage risk. First, in order to remove variations due to predictable wage growth, we detrend log gross wage growth with a regression on age, its square, education, and interactions of these variables, following, for instance, Hryshko (2012). In a second step, we obtain the sample standard deviation of past detrended log wages for each person similarly to Parker et al. (2005). Hence, our risk measure uses only the variation across time for each individual. Only wage observations from the current occupation are used for the construction of the risk

[6] We set $\Delta y_{it} = 2000$ Euro, which implies an increase in labor income of about 40 Euro per week.

[7] The SOEP data provide information on overtime compensation or_{it} in the sense whether overtime was (a) fully paid, (b) fully compensated with time off, (c) partly paid, partly compensated with time off, or (d) not compensated at all. $I(or_{it} = a)$ is an indicator function, in this case indicating that overtime rule (a) applies. We approximate paid hours of work as $h_{it} = hc_{it} + I(or_{it} = a)(ht_{it} - hc_{it}) + 0.5I(or_{it} = c)(ht_{it} - hc_{it})$, where hc_{it} are contracted hours of work and ht_{it} are actual hours of work (Euwals, 2005).

measure such that wage risk is not confounded by occupation choices. Thus at least two (not necessarily consecutive) periods of working in the same occupation are needed to construct the risk measure.

The wage risk measure is given by:

$$\sigma_{w,it} = \sqrt{\frac{1}{\#-1}\sum_{j=t-\#}^{t-1}(\ln \tilde{w}_{ij} - \ln \bar{\tilde{w}}_i)^2}, \tag{4.2}$$

where $\tilde{w}_j$ denotes the detrended (net) wage and # denotes the number of past realizations of wage. The idea behind this measure is that workers use past variations in idiosyncratic wages to form expectations about future risk. As we only use past information, we may treat this measure as exogenous at the moment of the labor supply decision. We denote this measure by $\sigma_{w,it}$. For the estimations, we standardize the risk measure by one standard deviation of the sample used in the regression to facilitate interpretation. We provide robustness tests with different risk measures, such as forward looking, five-year rolling windows, without detrending, using only continuous wage spells, subjective risk measures, other household income risk, and including occupational changes in Section 4.6.

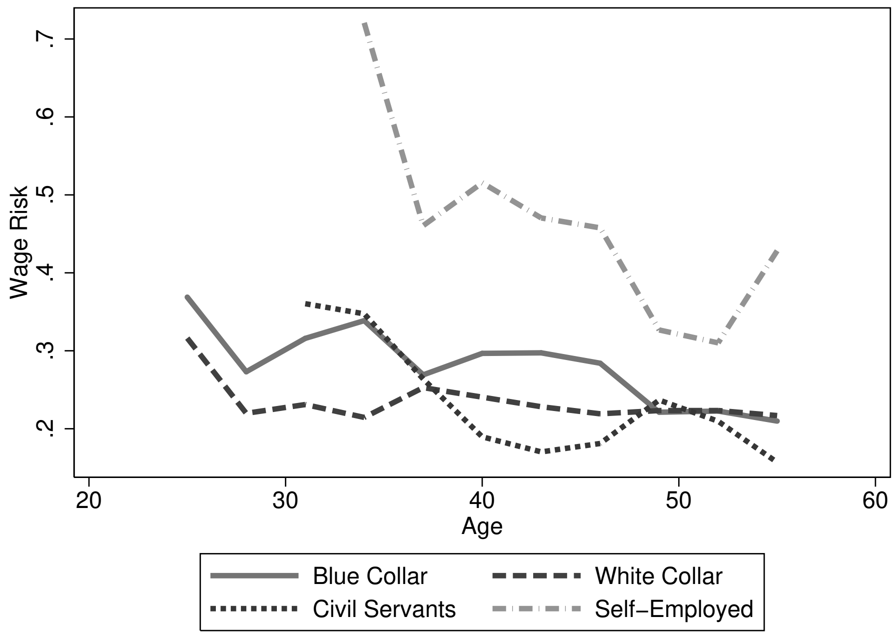
Our measure of wage risk assumes following e.g. Blundell and Preston (1998) or Blundell et al. (2008) that information unknown to the econometrician is unpredictable for the worker as well. Cunha et al. (2005) developed a method that distinguishes information unknown to the econometrician but predictable by the agent from information unknown to both. Applications of this method, see e.g. Cunha and Heckman (2008), Navarro (2011), Cunha and Heckman (2016), Navarro and Zhou (2017), show that equating variability with uncertainty results in overstated risk. To separate the information sets, correlation between choices and future realizations of the stochastic variable may be used.

As in Fossen and Rostam-Afschar (2013), we divide our sample into blue collar workers, white collar workers, civil servants, and self-employed. We are mainly interested in decisions during work life at ages where occupational changes are rare. Nonetheless, we model the selection into occupations as a robustness test in the Appendix.

Figure 4.1 shows how the average net wage risk evolves over the life cycle for each subgroup. We use age groups of three years to obtain a sufficient number of observations for each data point. Only age-occupation combinations with more than 15 observations are displayed, thus the trajectory for self-employed starts at age 35. We find that wage risk decreases slightly over the life cycle for all groups. This is more pronounced for the self-employed. The finding is in line with results in Blundell et al. (2015) who find that income risk decreases over the life cycle in Norway.

As expected, the hourly wages of self-employed workers are more volatile over the entire life cycle than those of employees. At all ages this difference is statistically signifi-

Figure 4.1.: Average Net Wage Risk over the Life Cycle



Note: Standard deviations of past marginal net wages for each individual averaged over three years by occupation. We calculate the risk measure for every age for every individual based on past realizations and take the average of this measure over individuals for every age. See equation (4.2). *Source:* Own calculation based on the SOEP

cant at the 5% significance level.[8] Blue and white collar workers have similar levels of wage risks. Nonetheless, during their 30s and 40s blue collar workers face a statistically significantly higher wage risk than white collar workers. For most age groups, the average net wage risk of civil servants is slightly lower than those of blue collar and white collar workers. This difference is statistically significant at most ages starting in the 40s.

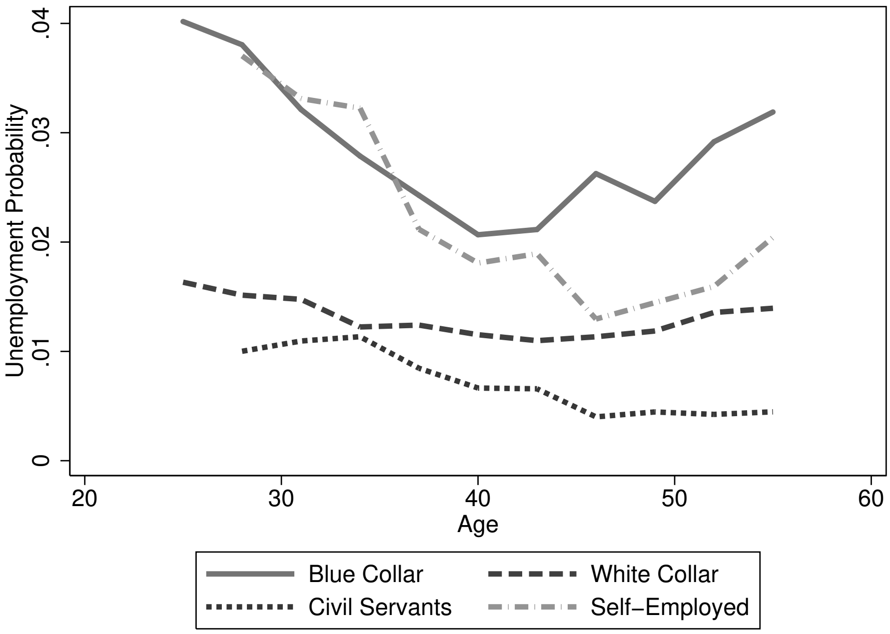
Unemployment Probability The control variable unemployment probability $\Pr_{U,it}$ is the predicted probability to be out of work in the next year. The estimation procedure is similar to the one used by Carroll et al. (2003).[9] Figure 4.2 displays how the average

[8] We use a two-sample t test with unequal variances to obtain the p-values. Test statistics are available from the authors upon request.

[9] We use a heteroskedastic probit model (cf. Harvey, 1976) to estimate the probability of unemployment in the following year conditional on regressors for occupation, industry, region, education, age, age squared, age interacted with occupation as well as with with education, marital status, and unemployment experience. The heteroskedasticity function includes previous unemployment experience and years of education.

unemployment probability evolves over the life cycle for the four occupational groups.[10] Civil servants have the lowest average unemployment probability, followed by white collar workers. For most parts of the life cycle, blue collar workers face the highest average unemployment probability. The mean unemployment probabilities of the occupational groups are statistically significantly different at all ages at the 5% level except for the difference between blue collar workers and self-employed at younger ages and white collar workers and self-employed at older ages. As for the wage risk, we standardize the unemployment probability by its standard deviation for the estimations.

Figure 4.2.: Average Unemployment Probability over the Life Cycle



Note: Predicted probability of unemployment next year for currently working married men averaged over three years by occupation. *Source:* Own calculation based on the SOEP

Summary Statistics Table 4.1 provides weighted summary statistics of the most important variables, including wage risk and unemployment probability measures. In the first row we report the average hours worked per week, about 42 in our sample. Hourly wages average 22 Euro, with average marginal net wages of 12 Euro. Hourly wages are constructed by dividing gross monthly labor incomes by paid hours of work. All monetary variables are converted to 2010 prices using the consumer price index provided by the Federal Statistical Office. Labor earnings include wages and salaries from all

[10]As in Figure 4.1, only age-occupation combinations with more than 15 observations are displayed.

employment including training, self-employment income, and bonuses, overtime, and profit-sharing.

We use paid hours because an increase in these translates directly into an increase in income.[11] The average gross wage risk in our sample is 0.192, which is similar to the average wage risk of 0.21 reported in Parker et al. (2005). The last three variables in Table 4.1 show that our sample has 8.0% self-employed workers, 32.5% blue collar workers, 48.2% white collar workers, and 11.3% civil servants.

Figure 4.3 shows the evolution of marginal net wages over the life cycle for different occupational groups. Profiles for white collar workers, civil servants, and self-employed are very similar with increasing wages until the age of about 45. In contrast, the wages of blue collar workers are lower and exhibit less wage growth. Figure 4.4 shows the same graph for weekly hours of work. This time, the self-employed are the odd ones out working substantially more than the other groups. For all groups average hours worked are relatively constant over the life cycle.

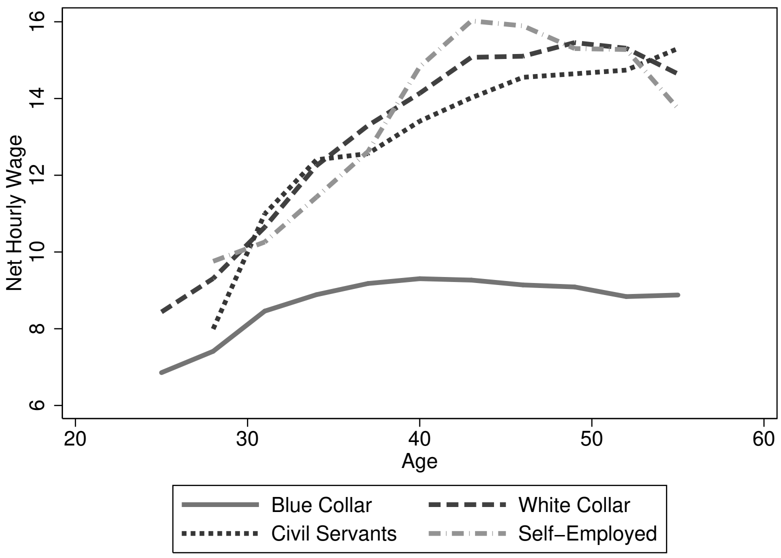
[11] We discuss robustness tests using different measures of hours supplied in Section 4.6.

Table 4.1.: Summary Statistics

	Unit	Mean	Std. Dev.	Min	Max	N
Labor Supply						
Weekly Hours Worked	(h)	42.03	7.3	20	80	16,038
Wages and Incomes						
Hourly Gross Wage	(Euro)	21.96	10.22	2.20	98.06	16,038
Hourly Marginal Net Wage	(Euro)	12.42	6.27	1.04	57.67	16,038
Monthly Gross Labor Income	(Euro)	3,764.47	1,997.75	319	27,000	16,038
Monthly Net Labor Income	(Euro)	2,458.91	1,197.49	150	15,000	16,038
Wage and Unemployment Probability						
Gross Wage Risk	(ln Euro)	0.192	0.196	0	3.539	11,040
Marginal Net Wage Risk	(ln Euro)	0.249	0.224	0	3.354	10,987
Unemployment Probability	(%)	1.4	2.2	0	27.4	16,038
BB-Index	(%)	2.7	4.7	-4.9	16.0	16,038
Demographics and Characteristics						
Age	(a)	43.1	7.5	25	55	16,038
Years of Education	(a)	12.8	2.7	7	18	16,038
Work Experience	(a)	21.5	8.5	0.2	41.2	16,038
Children younger than 3 years	(%)	11.6	32.0	0	100	16,038
Children between 3 and 6 years	(%)	14.5	35.2	0	100	16,038
Children between 7 and 18 years	(%)	45.2	49.8	0	100	16,038
East Germany	(%)	14.5	35.2	0	100	16,038
Type of Work						
Self-Employed	(%)	8.0	27.2	0	100	16,038
Blue Collar	(%)	32.5	46.8	0	100	16,038
White Collar	(%)	48.2	50.0	0	100	16,038
Civil Servant	(%)	11.3	31.7	0	100	16,038
One-Digit International Standard Classification of Occupations (ISCO)						
Managers	(%)	10.7	30.9	0	100	16,038
Professionals	(%)	22.0	41.4	0	100	16,038
Technicians	(%)	20.2	40.2	0	100	16,038
Clerks	(%)	7.7	26.6	0	100	16,038
Service and Sales	(%)	4.5	20.7	0	100	16,038
Craftsmen	(%)	20.9	40.7	0	100	16,038
Operatives	(%)	9.7	29.6	0	100	16,038
Unskilled	(%)	4.3	20.4	0	100	16,038

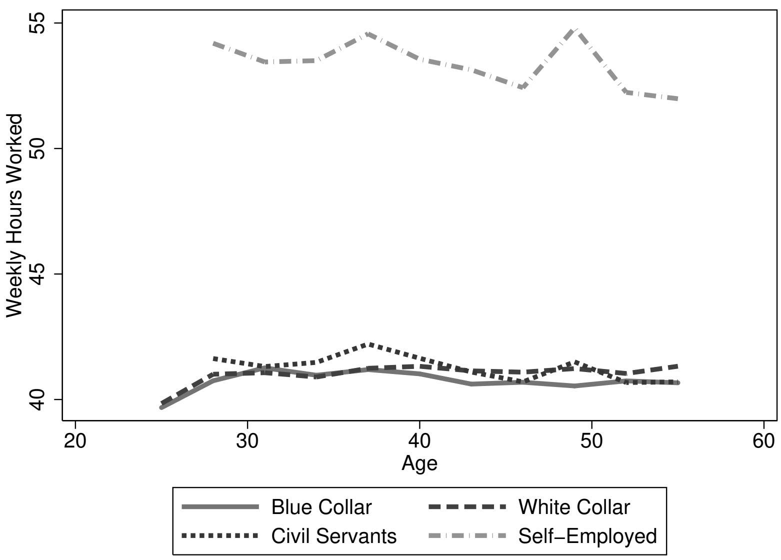
Notes: Data from SOEP (version 30). Sample of married prime-age males; 2001-2012.

Figure 4.3.: Average Marginal Hourly Net Wage over the Life Cycle



Source: Own calculation based on the SOEP

Figure 4.4.: Average Weekly Hours Worked over the Life Cycle



Source: Own calculation based on the SOEP

4.3. Empirical Strategy

4.3.1. Constrained Adjustment of Labor Supply

We begin the investigation with the following labor supply equation which is similar to the specification studied in Parker et al. (2005):[12]

$$\ln h_{it}^{*} = \tilde{\beta}_1 \ln w_{it} + \tilde{\beta}_2 X_{it} + \tilde{\beta}_3 \sigma_{w,it} + \omega_{it}, \tag{4.3}$$

where h_{it}^{*} denotes desired hours of work, w_{it} denotes the marginal net hourly wage, $\sigma_{w,it}$ is a measure of wage risk, X_{it} contains additional controls, and ω_{it} is the residual.

This specification reflects the view that workers in some occupations, in particular those who are not self-employed, work more or less hours than desired. A reason for this might be contractual rigidities or fixed costs of employment like training or social insurance that make short hours of work unprofitable for firms. For manual workers, Stewart and Swaffield (1997) showed that work hours are significantly higher than the desired level (overemployment) and workers thus "off their labor supply curve". Bryan (2007) uses OLS with correction terms from a fist step random effects ordered probit model that determines the probability of being over-employed, unconstrained or under-employed (but not unemployed). He documents that 45% of manual men were constrained in their choices of hours in a given year in the UK. More recently, Bell and Blanchflower (2013b,a) proposed an index (BB-index) to measure the opposite case, i.e. that workers would like to work more hours (under-employment). They find that under-employment has been substantial in the UK labor market recently. Table 4.1 shows that in Germany as well the average person in the work force is underemployed.[13] Hours constraints might be only temporary e.g. if workers may find another job that matches their preferences better. To reflect constraints in the adjustment of hours worked, we explicitly model the dynamics of actual hours choices h_{it} and specify a partial adjustment mechanism employed by, for example, Robins and West (1980), Euwals (2005), and Baltagi et al. (2005):

$$\ln h_{it} - \ln h_{it-1} = \theta(\ln h_{it}^{*} - \ln h_{it-1}), \qquad 0 < \theta \leq 1. \tag{4.4}$$

θ may be interpreted as the speed of adjustment. This speed might be determined by costs to immediately adjust the labor supply to desired hours or habit persistence (see,

[12] Pistaferri (2003) specifies a different labor supply equation, which relies on subjective expectations of future earnings.

[13] Following Bell and Blanchflower (2013b) we constructed a variable that measures the probability of being under- or over-employed and included it in X_{it} along with the probability of unemployment as a robustness test in Table A7 in the Appendix.

e.g., Brown, 1952). Replace (4.4) in (4.3) to obtain the partial adjustment labor supply specification:

$$\ln h_{it} = \alpha \ln h_{it-1} + \beta_1 \ln w_{it} + \beta_2 X_{it} + \beta_3 \sigma_{w,it} + \varepsilon_{it}. \tag{4.5}$$

This is our empirical labor supply specification. The parameters of (4.3) can be recovered following the estimation of (4.5) with $\alpha = 1-\theta$, $\beta_1 = \theta\tilde{\beta}_1$, $\beta_2 = \theta\tilde{\beta}_2$, $\beta_3 = \theta\tilde{\beta}_3$, and $\varepsilon_{it} = \theta\omega_{it}$ (Baltagi et al., 2005).[14] The partial adjustment model nests the classic labor supply equation with $\theta = 1$ as a special case. The short-run labor supply elasticity is given by $SR_{\eta_w} = \beta_1$, and the short-run labor supply elasticity with respect to risk by $SR_{\eta_{\sigma_w}} = \beta_3$. The corresponding long-run elasticities are $LR_{\eta_w} = \beta_1/(1-\alpha)$ and $LR_{\eta_{\sigma_w}} = \beta_3/(1-\alpha)$.

4.3.2. Instrumentation and Estimation Methods

To estimate our labor supply equation, we need to account for several sources of endogeneity. First, the first difference of the lagged dependent variable is correlated with the first difference of the error term ε_{it}, which includes shocks from $t-1$. We follow Anderson and Hsiao (1981) and instrument the lagged difference in the log of hours with the level $\ln h_{it-2}$ (Anderson-Hsiao estimator). In an alternative specification, we exploit additional moment conditions as suggested by Arellano and Bond (1991) and Holtz-Eakin et al. (1988) and apply the two-step difference GMM estimator (DIFF-GMM) with Windmeijer (2005) finite-sample correction. Blundell and Bond (1998) and Arellano and Bover (1995) show that imposing additional restrictions on the initial values of the data generating process and using lagged levels and lagged differences as instruments improves the efficiency of the estimates. We also present the results from this estimator, called the system GMM (SYS-GMM).

Second, marginal net wage rates may be endogenous for two reasons: First, measurement error in hours leads to downward denominator bias in the coefficient of wage rate since the hourly wage is calculated by dividing labor income by the dependent variable hours of work (cf. Borjas, 1980; Altonji, 1986; Keane, 2011). Second, the marginal net wage depends on the choice of hours because of the nonlinear tax and transfer system. Therefore, we instrument marginal net wages with the first lag of net labor income. This variable is predetermined during the current period labor supply choices and uncorrelated with the measurement error in current period hours.

[14]Note that ε_{it} might contain an individual time-invariant effect, which is eliminated by first-differencing as in the majority of the estimators used.

4.4. Results

4.4.1. Impact of Wage Risk on Weekly Hours of Work

Table 4.2 presents the results of the augmented labor supply equation for different estimators, where the dependent variable is the log of paid hours of work. Standard errors are robust and clustered at the individual level. Columns 1-3 show the results for the immediate adjustment specification, i.e. where the adjustment parameter α in equation (4.5) is restricted to zero. Columns 4–6 show results for the preferred dynamic specification. The first column displays results for the pooled OLS estimator. The coefficient of marginal net wage is significantly negative. The main coefficient of interest is the one associated with wage risk. The coefficient of 0.028 indicates that an increase in wage risk by one standard deviation would increase labor supply by 2.8%. The coefficient on unemployment probability is very small and not statistically significant.

Column 2 shows results for the pooled 2SLS estimator, where net wage is instrumented with lagged net labor income to overcome the denominator bias.[15] The sign of the coefficient of net wage becomes positive and the coefficient of wage risk remains significantly positive with a point estimate of 0.036. The unemployment probability becomes significant and the point estimate of 0.020 implies that an increase in unemployment probability by one standard deviation translates into 2.0% more hours worked. Column 3 displays the results obtained with the first difference estimator (FD-IV) with the equivalent instrument for net wages. The wage risk coefficient drops slightly but remains significantly positive. The coefficient of marginal net wage is not robust across estimators.

The partial adjustment specification results appear in columns 4–6 with the Anderson-Hsiao estimator displayed in column 4 and the results for the Difference and System GMM estimators displayed in columns 5 and 6, respectively.[16] The immediate adjustment specification is rejected with all three estimators because of statistically and economically significant point estimates of lagged hours of work between 0.14 and 0.2. For all three dynamic estimators, the coefficients of wage risk and unemployment probability are statistically significant. The magnitude of these effects is similar across all dynamic specifications and close to the results of the immediate adjustment specifications.

[15]We estimate it using the ivreg2 package (Baum et al., 2016a).
[16]We estimate them using the xtabond2 package (Roodman, 2009).

Table 4.2.: Labor Supply Regressions with Alternative Instrumentation Strategies

	OLS	2SLS	FD-IV	Anderson-Hsiao	DIFF-GMM	SYS-GMM
Lag of ln(Hours Worked)				0.155***	0.143***	0.195***
				(0.041)	(0.039)	(0.039)
ln(Net Wage) Risk	0.028***	0.036***	0.010*	0.010*	0.009*	0.024***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.004)
Unempl. Prob.	-0.005	0.020***	0.014**	0.015**	0.013*	0.015***
	(0.006)	(0.006)	(0.007)	(0.007)	(0.007)	(0.004)
ln(Marginal Net Wage)	-0.031***	0.183***	-0.073*	-0.060	-0.062*	0.159***
	(0.009)	(0.019)	(0.039)	(0.041)	(0.034)	(0.019)
Controls	✓	✓	✓	✓	✓	✓
Instruments	—	labinc_{it-1}	$\Delta\text{labinc}_{it-1}$	$\ln h_{it-2}$, $\Delta\text{labinc}_{it-1}$	$\ln h_{it-2},\ldots,\ln h_{it-11}$, $\Delta\text{labinc}_{it-1}$	$\ln h_{it-2},\ldots,\ln h_{it-11}$, $\Delta\ln h_{it-2},\ldots,\Delta\ln h_{it-11}$, $\Delta\text{labinc}_{it-1}$
Observations	8,112	8,112	8,112	8,112	8,112	8,112
AR(1) in FD					0.000	0.000
AR(2) in FD					0.954	0.745
Hansen					0.694	0.368

Notes: Columns 1-3: Estimation of an immediate adjustment labor supply equation.
Columns 4-6: Estimation of equation (4.5) *using different estimators.*
We use the sample of the dynamic specifications for all estimations.
Robust standard errors clustered at the individual level in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Source: Own calculation based on the SOEP

The coefficient on marginal net wage becomes insignificant in the Anderson-Hsiao and even significantly negative in the difference GMM specification. Blundell and Bond (1998) show that the Difference GMM estimator can be heavily downward biased. Therefore, we prefer System GMM. The wage coefficient is estimated with much higher precision using the system GMM estimator yielding statistical significance at the 1% level. This specification implies a short run labor supply elasticity of $SR_{\eta_w} = 0.16$ and a long run elasticity of $LR_{\eta_w} = 0.20$. For the difference and system GMM estimators, autocorrelation and Hansen tests appear below the estimates. The null hypothesis of no autocorrelation of second order cannot be rejected and the Hansen overidentification test does not indicate any invalidity in the instruments.

Table A2 in the Appendix shows the equivalent of Table 4.2 but using gross wages instead of net wages. This facilitates comparison to the extant literature, e.g., Parker et al. (2005), that does not use microsimulation models, but relies on gross wages. The coefficient of gross wage risk is positive and significant at the 1 percent level in three of the specifications. The preferred system-GMM yields similar coefficients for all variables as the system-GMM for net wages in Table 4.2.

4.4.2. Results by Occupations

As argued by Parker et al. (2005), there should be heterogeneity across occupational groups, especially concerning self-employed. To quantify this heterogeneity, we present the results of our preferred specification across the occupational groups introduced above and the International Standard Classification of Occupations of 1988 (ISCO).

Table 4.3 provides separate results for different occupational groups using the system GMM estimator with the same instruments as in Table 4.2. As before, the risk measures are normalized by one standard deviation; however, this time not by the overall, but the sub-sample specific standard deviation. The point estimate of the wage risk coefficient is positive and statistically significant for self-employed, white collar, and blue collar workers, but not statistically different from zero for civil servants. The point estimate is largest for self-employed workers (0.036) and much smaller for white collar (0.010) and blue collar workers (0.007), suggesting the most important role of precautionary labor supply for the self-employed. Note that the result for self-employed is very similar to the one of Parker et al. (2005) where an additional standard deviation of wage risk implies an increase of annual hours of 3.66%.[17]

The coefficient on the lag of paid hours worked is not statistically significant for the self-employed and civil servants, which makes intuitively sense; these two groups are not as severely constrained in their hours choices as regular employees. Blue collar workers (0.226) are more constrained than white collar workers (0.116). This means that if underemployed blue-collar workers desire to work, say, 40 instead of 30 hours per week

[17] This number is obtained by multiplying the coefficient of risk from Model 2 with the reported standard deviation of the wage risk measure.

in Germany, they need about four years to achieve this, while white collar workers need about two years according to our estimates of the speed of adjustment parameter.

Table 4.3.: System GMM Labor Supply Regressions for Occupational Groups

	Self-Employed	White Collar	Blue Collar	Civil Servant
Lag of ln(Hours Worked)	0.109	0.116**	0.226***	0.046
	(0.099)	(0.048)	(0.055)	(0.129)
ln(Net Wage) Risk	0.036***	0.010***	0.007***	-0.007
	(0.012)	(0.003)	(0.003)	(0.007)
Unempl. Prob.	-0.013	0.005	0.009**	-0.001
	(0.014)	(0.004)	(0.004)	(0.005)
ln(Marginal Net Wage)	0.123***	0.133***	0.060***	0.244***
	(0.046)	(0.020)	(0.023)	(0.095)
Controls	✓	✓	✓	✓
Observations	864	5,652	2,987	1,407
AR(1) in FD	0.000	0.000	0.000	0.001
AR(2) in FD	0.688	0.987	0.459	0.286
Hansen	0.213	0.205	0.024	0.298

Notes: Estimation of equation (4.5) *using the SYS-GMM as in column 6, Table 4.2.*
Robust standard errors clustered at the individual level in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Source: Own calculation based on the SOEP

The coefficient of marginal net wage is positive and statistically significant for all groups. It is higher for civil servants than for other occupational groups. As in the estimation using the entire sample, we cannot reject the null hypothesis of no autocorrelation of second order. The Hansen test indicates that the instrument may be invalid only for blue collar workers.

Similarly, Table A3 in the Appendix shows results for the four occupations using gross wages instead of marginal net wages. As for marginal net wages, the wage risk coefficient is significantly positive for self-employed, white collar workers and blue collar workers. The coefficients of all other variables are very similar to the main results.

Table A4 in the Appendix shows system GMM estimates of the dynamic labor supply equation for eight professions grouped according to the ISCO. Each one-digit ISCO group is composed of several of the occupational classifications we used above, that is, some managers are self-employed, some not. Only clerks and operatives appear to be constrained in their hours choices. These constraints are quite persistent. The null hypothesis that wage risk does not affect labor supply is rejected for managers, professionals, technicians, craftsmen, and operatives. An increase in the probability of unemployent corresponds to an increase of hours worked particularly for managers, craftsmen, operatives, and unskilled. The coefficient of marginal net wage is significantly positive for all but clerks, service workers and operatives. Generally, both the coefficients of net wage risk and net wage are of similar magnitude as those obtained in the estimation using the main sample.

4.5. Importance of Precautionary Labor Supply

With our estimates of the wage risk semi-elasticity we can quantify the importance of precautionary labor supply in a ceteris paribus exercise, similarly to Carroll and Samwick (1998) for precautionary savings.[18] We use the estimates from Table 4.2 to simulate the resulting distribution of hours if all individuals faced the same small wage risk. We construct this simulated counterfactual $\hat{h}_{it}$ from the predictions of the dynamic labor supply equation with minimum sample wage risk $\sigma^{\min}_{w,it}$. We use the estimates obtained with the System GMM estimator. We then compare actual hours of work h_{it} observed in the data with their simulated counterfactuals. The difference gives us a measure of the magnitude of precautionary labor supply and, for the short-run, is calculated as

$$\hat{h}_{SR,it} - h_{it} = -\beta_3(\sigma_{w,it} - \sigma^{\min}_{w,it}). \tag{4.6}$$

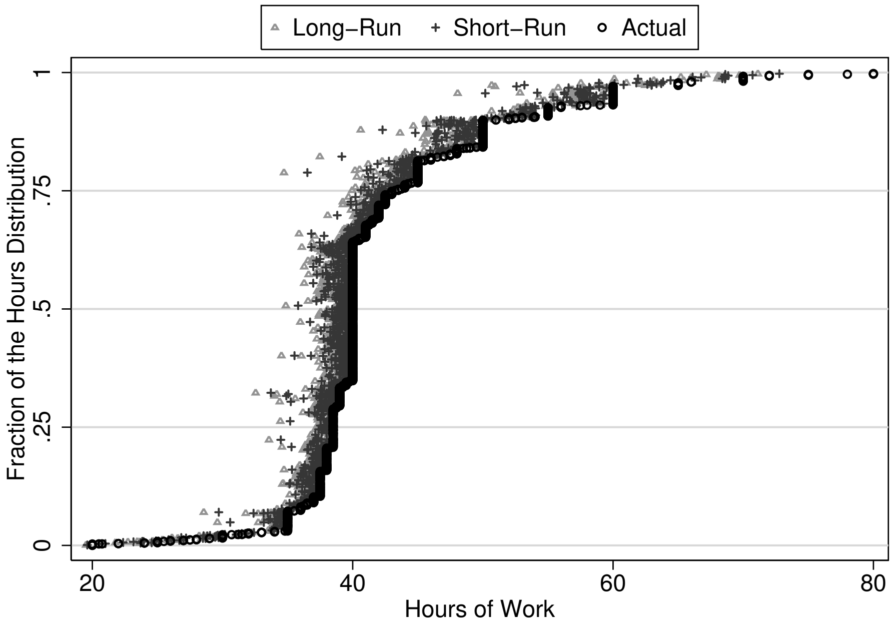
Figure 4.5 shows three points for each individual in the sample in 2011. The first point (p_i, h_i), denoted by a small circle, indicates the percentile rank p_i of individual i in the actually observed distribution of hours of work (vertical axis) and h_i indicates the actual hours of work (horizontal axis). The second point $(p_i, \hat{h}_{SR,i})$ keeps the percentile ranking p_i from the observed distribution and indicates the simulated *short-run* value of the hours of work $\hat{h}_{SR,i}$ when $\sigma_{w,it}$ is set to $\sigma^{\min}_{w,it}$. The third point $(p_i, \hat{h}_{LR,i})$ shows, as before, p_i from the observed distribution and indicates the simulated *long-run* value of the hours of work $\ln \hat{h}_{LR,i}$ when $\sigma_{w,it}$ is set to $\sigma^{\min}_{w,it}$.

$$\hat{h}_{LR,it} - h_{it} = -\frac{\beta_3}{1-\alpha}(\sigma_{w,it} - \sigma^{\min}_{w,it}). \tag{4.7}$$

The short-run simulated hours lie to the left of the actual hours distribution. The horizontal difference between short-run simulated points and observed points indicates the reduction in the number of hours in the short run if wage risk was reduced to the minimum level. The long-run simulated hours lie to the left of both the actual hours

[18] Precautionary labor supply is likely even more important for singles because spousal labor supply is an additional channel of insurance against risk. However, applying our analysis to singles is difficult because only a small number of individuals in the SOEP are singles over long periods.

Figure 4.5.: Reduction in Hours of Work



Notes: Small circles indicate the percentile rank of individual i in the actual observed distribution of hours of work (vertical axis) and the actual hours of work (horizontal axis) in 2011. Plus symbols maintain the percentile ranking from the observed distribution and indicate the simulated short-run *value of the hours of work when $\sigma_{w,it}$ is set to $\sigma^{min}_{w,it}$. Triangles denote the respective long-run hours of work when $\sigma_{w,it}$ is set to $\sigma^{min}_{w,it}$.Source:* Own calculation based on the SOEP

distribution and the short-run simulated points. The horizontal difference between long-run simulated points and observed points indicates the reduction in the number of hours of work in the long-run if wage risk was reduced to the minimum level. The horizontal difference between simulated points in the long- and short-run indicates how much of the adjustment in hours would occur after the immediate reaction to the wage risk reduction.

Table 4.4 reports the labor supply reduction in the short run (columns 1 and 2) and the long-run (columns 3 and 4) if wage risk was reduced to the sample minimum (columns 1 and 3) or the median wage risk of civil servants (columns 2 and 4). In the pooled sample, hours of work would reduce by 2.77% in the long run if wage risk were reduced to the sample minimum. Keep in mind that this is a ceteris paribus exercise neglecting general equilibrium effects. Defining precautionary labor supply as the difference between hours worked in the status quo and in the absence of wage risk and given the average of 42 weekly paid hours of work in our sample, precautionary labor supply amounts to 1.16 hours per week on average.

Table 4.4.: Percentage Reduction for Different Occupations

	Short-Run		Long-Run	
	Perfect Foresight	Civil Servants	Perfect Foresight	Civil Servants
Self-Employed	5.01	3.65	6.17	4.49
Blue Collar	2.17	0.76	2.68	0.94
White Collar	2.03	0.62	2.51	0.77
Civil Servants	2.00	0.60	2.48	0.74
All	2.24	0.84	2.77	1.03

Notes: Simulated percentage reduction in hours of work when reducing wage risk to the sample minimum (perfect foresight) or the median risk faced by civil servants.
Source: Own calculation based on the SOEP

If wage risk was reduced instead to the median wage risk of civil servants, labor supply would decrease on average by 1.03% in the long run. The wage risk of civil servants is below average, therefore this group may be regarded as an important benchmark with particularly low uncertainty. For the self-employed, the long-run labor supply reduction would amount to 4.49%. If the wage risk of all civil servants was reduced to its median, civil servants' labor supply would decrease by 0.74%.[19]

4.6. Robustness

This section discusses the results from various robustness tests. If not indicated otherwise, the results are estimated using the preferred estimator (System GMM). The tables are delegated to the Appendix.

Table A5 shows the main results for four alternative dependent variables. *Annual hours* (column 1) refers to the SOEP-imputed annual hours of work. *Weekly hours*, another variable imputed by the SOEP, is the basis for our main hours worked definition but without adjusting for paid overtime. Respondents are asked directly about *Contracted hours* and *Desired hours*. From a theoretical point of view, desired hours should not be constrained by a partial adjustment mechanism (cf. Euwals, 2005); hence, we use an immediate adjustment model for this specification. Annual hours, weekly hours and desired hours increase with increasing wage risk, while the coefficient for contracted hours is insignificant. The likely reason is that contracted hours cannot be as easily adjusted as actual hours. While still significant and economically important, the coefficient of wage risk in the desired hours specification (0.007) is smaller than in the main specification.

[19]This effect would equal zero if the distribution of wage risk were symmetric for civil servants.

This is not surprising because respondents might understand the question in different ways. Therefore, this measure could be affected by measurement errors, which biases the coefficient towards zero.

Table A6 shows results for eight alternative risk specifications. Column 1 shows the case with a forward looking risk measure, i.e., the standard deviation of future detrended log wages. This is similar to the approach in Feigenbaum and Li (2015). Column 2 uses a five year rolling window for the construction of the wage risk measure. Column 3 shows results obtained using the risk measure constructed using undetrended wages. This measure corresponds to the one used by Parker et al. (2005). Column 4 uses only observations with continuous employment spells, i.e., we drop observations of individuals whose employment is interrupted by periods of unemployment or changes between occupations. Columns 5 and 6 include indicators of subjective risk perceptions (Some Worries, Big Worries), column 7 includes the risk of additional household income as an additional control. This is constructed like our main risk measure, but using net household income minus net labor income of the husband instead of the husband's wage. The coefficient of this risk measure is significant and positive, so this source of risk also leads to precautionary labor supply. In column 8 we construct the wage risk measure using all past wages including those from different occupations than the current one. This increases the number of observations and the coefficient of wage risk substantially. This risk measure includes not only wage risk but also occupational risk and implies that these additional risks cause even more important precautionary behavior. The coefficients of the other regressors change only slightly. The wage risk coefficient is similar as in the main specification and remains statistically significant in all other columns.

It is possible that selection into job types could be driven by risk attitudes and the desire for hard work. If these variables are correlated with risk, this would lead to omitted variable bias. Fuchs-Schündeln and Schündeln (2005) exploit the natural experiment of the German reunification to find that risk-averse individuals self-select into low-risk occupations. Not accounting for this selection mechanism might lead to omitted variable bias. To make sure that our results are robust to such concerns, we employ two strategies, including additional controls and estimating a selection correction model. Fortunately, the SOEP elicits information on both risk preferences and the attitude towards hard work. Therefore, our first strategy is to include these additional control variables in the main model. The results are reported in Table A7. In column 1 we add a variable reporting to what degree respondents agree with the assertion "Success takes hard work" on Likert scale from 1 to 7. As expected, this variable has a positive and significant impact on hours. An increase of 1 on the the Likert scale leads to an increase of 1 percent in hours of work. All other coefficients remain virtually the same. In column 2 we include a control that measures the stated willingness to take risk on a scale from 0 to 10, but do not include the preference for hard work variable. A one unit increase in this variable increases hours of work by 0.3 percent. In column 3 we include both additional control variables. Their coefficients are identical to those reported in the previous columns. The main results are very robust to this variation. In column 4 we report results, where we add a variable that captures the stated willingness to take risks in financial matters on a scale from 0 to 10 in addition to the variable capturing attitudes towards hard work. In column 5 we control for the hard-work variable and a variable capturing stated attitudes towards risks in occupational matters. An increase in the variable capturing attitudes towards occupation risk by one unit leads to an increase in hours of work by 0.4 percent, while the variable for risk attitudes in financial matters is insignificant. Again, the main results do not change.

While we explicitly model hours constraints on the occupational level in our dynamic specification, differences in hours constraints between individuals might still bias our results. Therefore we follow Bell and Blanchflower (2013b,a) and construct a region-specific indicator for under- or overemployment. The Bell-Blanchflower underemployment index (BB-index) is defined as

$$u_{BB} = \frac{U\overline{h} + \sum_k h_k^U - \sum_j h_j^O}{U\bar{h} + \sum_i h_i},$$

where U is the number of unemployed, $\overline{h}$ average hours worked by employed, h^U is preferred additional hours, which are aggregated over all workers k who desire to work more, while h^O is the preferred reduction in hours, which are aggregated over all workers j who desire to work less. $\sum_i h_i$ is the sum of actual hours of work over all workers. We use a variable for desired hours of work in the SOEP to calculate over- and underemployment. In the case that all currently employed workers are satisfied with their hours of work, the BB-index simplifies to the unemployment rate. The higher the value of this index, the more likely it is that workers are underemployed, i.e., wish to work more. Negative values indicate overemployment, i.e., people in the labor force on average wish to work less hours. As shown in Table 4.1 the value of the index is 2.7 percent on average for our sample. Column 6 of Table A7 shows that an increase in the BB-index by 1%-point leads to a decrease in hours of work by 0.001 percent. The sign of the coefficient is in line with theoretical predictions. People who are more likely to be underemployed on average work slightly less, although they potentially want to work more. However, the magnitude is economically not relevant. In Column 7 we include both the BB-index and the general risk preferences variable. The BB-index becomes statistically insignificant, although the reported standard error and coefficient are identical. The reason is that the forth digit after the decimal point differs between the columns. The main results are virtually unchanged. This shows that our main results are highly robust to inclusion and exclusion of these additional control variables.

In addition to these controls, there might be selection into occupations on unobservables. We account for this possibility by estimating a Heckman (1979) selection correction model for each of the four occupations. Indicator variables for the occupation and education of both parents, and spatial planning regions are included only in the selection equation. The results are reported in Table A8. The coefficient of the marginal net wage is biased downwards because we do not instrument it. Moreover, the model omits the dynamic structure of our main estimation. The focus is on the coefficients of wage risk and unemployment risk. Wage risk is positive and statistically significant at the 1 percent level and of the same order of magnitude as in Table 4.3 for the first three occupations. As before, the effect is strongest for the self-employed. The coefficient for civil servants remains insignificant. The effect of the unemployment probability remains the same except for the self-employed, where it indicates that an increase in the probability of unemployment leads to a 3.5%-decrease in hours of work. An explanation for this is that the unemployment probability for the self-employed is also a measure for the deterioration of the business and a decreasing number of orders. In the case of self-employed this is directly related to the number of hours worked. Overall, the results suggest that the main result that increases in wage risk lead to increases in hours of work is not confounded by selection bias.

Given that we do not observe many young self-employed and civil servants in our sample because these occupations are typically chosen by older individuals, we repeat the analysis by occupations including only individuals aged at least 35. The results are reported in Table A9. This makes sure that the comparison is based on common support regarding the life cycle. The results are very similar to those reported in Table 4.3. This shows that the differences between occupations are not driven by differences in age.

We also show results obtained for the main sample, but including transfer recipients in Table A9. This group is dropped from the main analysis because institutional insurance through the transfer system is likely to play a much larger role than precautionary behavior and even constrains precautionary behavior (Hubbard et al., 1995; Cullen and Gruber, 2000; Engen and Gruber, 2001). On the other hand, this group might be subject to more gross wage risk and therefore have stronger precautionary motives. The obtained coefficients of wage risk are virtually unchanged, when this group is included in the estimation sample.

Finally, we reestimate the main specification by occupations including interactions between year indicators and the wage risk measure (Table A10). Overall, the estimates of the impact of wage risk are less precise due to less observations for a given year. Nonetheless, the coefficient is economically and statistically significant for many years except for civil servants, as in the main results. When looking at the crisis known as the Great Recession and its aftermath, i.e., 2008-2010, the effect is particularly strong for the self-employed and white collar workers. A similar pattern is not observable for blue collar workers, which does not surprise, since German manufactures made excessive use of short-time work allowance to cushion the effects of the crisis (Burda and Hunt, 2011).

4.7. Conclusion

We quantify the importance of wage risk to explain the hours of work of married men. The analysis is based on the German Socio-Economic Panel data for 2001 to 2012. We find that workers choose slightly more than an hour per week to shield against wage shocks. These effects are statistically significant for various occupations, but not for civil servants, which is in line with previous studies. We observe the largest effects of wage risk for the self-employed who have typically less coverage by institutional insurance like short term unemployment benefits. Our result for this group is quantitatively similar to previous results by Parker et al. (2005).

Precautionary labor supply is economically important. Considering a person who works 42 hours per week, precautionary labor supply amounts to about one week per year or in monetary terms, about 710 Euro per year, with a typical net wage rate of 13 Euro. If all workers faced the same risk as the median civil servant, hours worked would decrease on average by 1% in the long run. Precautionary labor supply is particularly important for the self-employed, a group that faces average wage risk substantially above the sample mean. This group works 6.2% of their hours because of the precautionary motive. Our findings suggest that unemployment probability also plays a statistically significant role, but is quantitatively less important than wage risk.

Appendix

Table A1.: Sample Restrictions for the Main Sample

Full sample: 416,241 person years	*Eliminated*	*Remaining*
Incomplete interviews	9,829	406,412
Drop if female	207,407	199,005
Drop if not married	55,457	143,548
Drop if younger than 26 or older than 55 in each year	86,223	57,325
Drop if in military or agriculture	2,155	55,170
Drop if transfer recipients	6,806	48,364
Drop if very low hours worked	495	47,869
Drop if unrealistic hours changes	115	47,754
Drop if unrealistic wage changes	670	47,084
Drop if without net wage or risk	36,097	10,987
After first differencing, drop if no available IVs	2,875	8,112

Source: Own calculation based on the SOEP

Table A2.: Comparison of Specifications, Gross Wages

	OLS	2SLS	FD-IV	FD-IV	DIFF-GMM	SYS-GMM
Lag of ln(Hours Worked)				0.173***	0.153***	0.189***
				(0.039)	(0.037)	(0.033)
ln(Gross Wage) Risk	0.044***	0.051***	0.002	0.002	0.002	0.036***
	(0.004)	(0.005)	(0.004)	(0.005)	(0.005)	(0.004)
Unempl. Prob.	-0.003	0.013***	0.005	0.005	0.005	0.008**
	(0.004)	(0.004)	(0.005)	(0.005)	(0.005)	(0.003)
ln(Marginal Gross Wage)	-0.081***	0.130***	0.000	0.012	-0.003	0.112***
	(0.010)	(0.015)	(0.023)	(0.026)	(0.025)	(0.016)
Controls	✓	✓	✓	✓	✓	✓
Instruments	—	labinc_{it-1}	$\Delta\text{labinc}_{it-1}$	$\ln h_{it-2}$, $\Delta\text{labinc}_{it-1}$	$\ln h_{it-2},\ldots,\ln h_{it-11}$, $\Delta\text{labinc}_{it-1}$	$\ln h_{it-2},\ldots,\ln h_{it-11}$, $\Delta\ln h_{it-2},\ldots,\Delta\ln h_{it-11}$, $\Delta\text{labinc}_{it-1}$
Observations	11,276	11,276	11,276	11,276	11,276	11,276
AR(1) in FD					0.000	0.000
AR(2) in FD					0.193	0.100
Hansen					0.708	0.238

Notes: Columns 1-3: Estimation of an immediate adjustment labor supply equation.
Columns 4-6: Estimation of equation (4.5) *using different estimators.*
We use the sample of the dynamic specifications for all estimations.
Robust standard errors clustered at the individual level in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Source: Own calculation based on the SOEP

Table A3.: Occupational Groups, System GMM, Gross wages

	Self-Employed	White Collar	Blue Collar	Civil Servant
Lag of ln(Hours Worked)	0.132**	0.161***	0.197***	0.015
	(0.064)	(0.048)	(0.040)	(0.127)
ln(Gross Wage) Risk	0.019**	0.013***	0.010***	-0.005
	(0.009)	(0.003)	(0.003)	(0.007)
Unempl. Prob.	-0.019	0.007*	0.011***	0.002
	(0.014)	(0.004)	(0.003)	(0.005)
ln(Marginal Gross Wage)	0.082**	0.115***	0.055***	0.226**
	(0.034)	(0.018)	(0.021)	(0.093)
Controls	✓	✓	✓	✓
Observations	1,328	6,755	5,414	1,512
AR(1) in FD	0.000	0.000	0.000	0.001
AR(2) in FD	0.244	0.159	0.953	0.302
Hansen	0.916	0.146	0.052	0.582

Notes: Estimation of equation (4.5) *using the SYS-GMM as in column 6, Table 4.2.*

Robust standard errors clustered at the individual level in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Source: Own calculation based on the SOEP

Table A4.: System GMM Labor Supply Regressions for ISCO Groups

	Managers	Professionals	Technicians	Clerks	Service and Sales	Craftsmen	Operatives	Unskilled
Lag of ln(Hours Worked)	0.135	0.111	-0.054	0.429***	0.016	0.046	0.323***	0.327
	(0.093)	(0.076)	(0.105)	(0.142)	(0.125)	(0.068)	(0.090)	(0.262)
ln(Net Wage) Risk	0.025***	0.027***	0.021***	0.005	0.012	0.022***	0.034***	0.016
	(0.008)	(0.007)	(0.008)	(0.003)	(0.010)	(0.006)	(0.013)	(0.019)
Unempl. Prob.	0.019**	0.007	0.007	-0.008*	0.000	0.019***	0.012*	0.015*
	(0.009)	(0.006)	(0.007)	(0.004)	(0.010)	(0.007)	(0.006)	(0.008)
ln(Marginal Net Wage)	0.187***	0.299***	0.174***	0.043	0.057	0.191***	0.092	0.162*
	(0.059)	(0.051)	(0.041)	(0.027)	(0.059)	(0.044)	(0.066)	(0.085)
Controls	✓	✓	✓	✓	✓	✓	✓	✓
Observations	1314	3007	2197	797	398	1985	880	332
AR(1) in FD	0.000	0.000	0.000	0.000	0.084	0.000	0.001	0.017
AR(2) in FD	0.496	0.259	0.712	0.720	0.451	0.351	0.107	0.765
Hansen	0.703	0.042	0.366	0.466	0.526	0.303	0.062	0.393

Notes: Estimation of equation (4.5) *using the SYS-GMM as in column 6, Table 4.2.*
Robust standard errors clustered at the individual level in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Source: Own calculation based on the SOEP

Table A5.: Alternative Hours Definitions

	Annual Hours	Weekly Hours	Contracted Hours	Desired Hours
Lag of ln(Hours Worked)	0.114	0.110	0.205**	
	(0.075)	(0.070)	(0.081)	
ln(Net Wage) Risk	0.024***	0.020***	-0.001	0.007**
	(0.004)	(0.004)	(0.001)	(0.003)
Unempl. Prob.	0.012**	0.018***	0.001	0.015***
	(0.006)	(0.005)	(0.003)	(0.004)
ln(Marginal Net Wage)	0.218***	0.215***	0.032***	0.144***
	(0.024)	(0.023)	(0.008)	(0.018)
Controls	✓	✓	✓	✓
Observations	11,034	10,845	8,739	10,768
AR(1) in FD	0.000	0.000	0.000	0.000
AR(2) in FD	0.475	0.139	0.726	0.929
Hansen	0.514	0.547	0.810	

Notes: Estimation of equation (4.5) *using the SYS-GMM as in column 6, Table 4.2.*
Robust standard errors clustered at the individual level in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Source: Own calculation based on the SOEP

Table A6.: Alternative Risk Definitions

	Forward	Five years	Undetrended	Cont. Spells	Subj. Risk	Subj. & Wage	Household Risk	With Occ. Changes
Lag of ln(Hours Worked)	0.223***	0.200***	0.192***	0.226***	0.187***	0.195***	0.171***	0.157***
	(0.049)	(0.039)	(0.039)	(0.044)	(0.041)	(0.041)	(0.042)	(0.034)
ln(Net Wage) Risk	0.020***	0.019***	0.023***	0.013***		0.021***	0.013**	0.088***
	(0.003)	(0.003)	(0.004)	(0.003)		(0.004)	(0.005)	(0.013)
Unempl. Prob	0.010***	0.009***	0.009***	0.008***	0.012***	0.011***	0.007**	0.009***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.004)	(0.003)	(0.003)	(0.002)
ln(Marginal Net Wage)	0.154***	0.156***	0.160***	0.158***	0.158***	0.164***	0.107***	0.164***
	(0.022)	(0.019)	(0.019)	(0.020)	(0.021)	(0.020)	(0.030)	(0.015)
Some Worries					0.016	0.055		
					(0.042)	(0.043)		
Big Worries					-0.086	-0.044		
					(0.076)	(0.075)		
ln(Net Household Inc.) Risk							0.061**	
							(0.031)	
Controls	✓	✓	✓	✓	✓	✓	✓	✓
Observations	5,675	8,089	8,112	6,614	8,101	8,101	8,014	15,544
AR(1) in FD	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
AR(2) in FD	0.577	0.835	0.800	0.776	0.425	0.318	0.870	0.498
Hansen	0.233	0.111	0.614	0.014	0.408	0.614	0.521	0.366

Notes: Estimation of equation (4.5) *using the SYS-GMM as in column 6, Table 4.2.*
Robust standard errors clustered at the individual level in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Source: Own calculation based on the SOEP

Table A7.: Additional Control Variables

	I	II	III	IV	V	VI	VII
Lag of ln(Hours Worked)	0.196***	0.198***	0.199***	0.200***	0.203***	0.195***	0.198***
	(0.040)	(0.039)	(0.040)	(0.041)	(0.041)	(0.039)	(0.040)
ln(Net Wage) Risk	0.021***	0.021***	0.020***	0.021***	0.020***	0.021***	0.020***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Unempl. Prob.	0.009***	0.009***	0.010***	0.009***	0.009***	0.010***	0.010***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
ln(Marginal Net Wage)	0.154***	0.156***	0.149***	0.151***	0.147***	0.158***	0.151***
	(0.019)	(0.019)	(0.018)	(0.019)	(0.019)	(0.019)	(0.019)
Success Takes Hard Work	0.010***		0.010***	0.010***	0.010***		0.010***
	(0.002)		(0.002)	(0.002)	(0.002)		(0.002)
General Risk Preference		0.003**	0.003***				0.003**
		(0.001)	(0.001)				(0.001)
Financial Risk Preference				-0.001			
				(0.001)			
Occupational Risk Preference					0.004***		
					(0.001)		
BB-Index						-0.001*	-0.001
						(0.001)	(0.001)
Controls	✓	✓	✓	✓	✓	✓	✓
Observations	7,862	8,109	7,859	7,686	7,653	8,112	7,859
AR(1) in FD	0.000	0.000	0.000	0.000	0.000	0.000	0.000
AR(2) in FD	0.884	0.604	0.709	0.770	0.807	0.764	0.725
Hansen	0.280	0.312	0.149	0.324	0.204	0.297	0.252

Notes: Estimation of equation (4.5) *using the SYS-GMM as in column 6, Table 4.2.*
Robust standard errors clustered at the individual level in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Source: Own calculation based on the SOEP

Table A8.: Two-step Heckman Selection Correction Model

	Self-Employed	White Collar	Blue Collar	Civil Servant
ln(Net Wage) Risk	0.033***	0.016***	0.006*	-0.010
	(0.011)	(0.003)	(0.004)	(0.006)
Unempl. Prob.	-0.035***	0.006	0.009**	0.001
	(0.010)	(0.004)	(0.004)	(0.008)
ln(Marginal Net Wage)	-0.100***	-0.024***	-0.050***	-0.296***
	(0.017)	(0.008)	(0.010)	(0.022)
Inverse Mills Ratio	-0.004	-0.003	0.012	0.026*
	(0.024)	(0.012)	(0.010)	(0.015)
Observations	4,758	4,758	4,758	4,758

Notes: Estimation of the immediate adjustment labor supply equation using the two-step Heckman selection model. Exclusion restrictions are: Indicator variables for the occupation and education of both parents, and spatial planning regions.

Standard errors in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Source: Own calculation based on the SOEP

Table A9.: Variations of the Sample I

	All, age> 34	SE, age> 34	WC, age> 34	BC, age> 34	CS, age> 34	Incl. TR
Lag of ln(Hours Worked)	0.200***	0.105	0.129***	0.210***	0.018	0.201***
	(0.040)	(0.102)	(0.050)	(0.065)	(0.137)	(0.038)
ln(Net Wage) Risk	0.023***	0.036***	0.010***	0.009***	-0.004	0.023***
	(0.004)	(0.012)	(0.003)	(0.003)	(0.008)	(0.004)
Unempl. Prob.	0.010***	-0.015	0.005	0.008**	-0.001	0.015***
	(0.003)	(0.015)	(0.005)	(0.004)	(0.005)	(0.004)
ln(Marginal Net Wage)	0.162***	0.125***	0.135***	0.069***	0.257***	0.156***
	(0.019)	(0.048)	(0.021)	(0.025)	(0.096)	(0.018)
Controls	✓	✓	✓	✓	✓	✓
Observations	7,547	830	5,216	2,539	1,337	8,660
AR(1) in FD	0.000	0.000	0.000	0.000	0.001	0.000
AR(2) in FD	0.627	0.667	0.890	0.434	0.244	0.854
Hansen	0.255	0.204	0.345	0.057	0.299	0.248

Notes: Estimation of equation (4.5) *using the SYS-GMM as in column 6, Table 4.2. SE: Self-employed; WC: White collar, BC: Blue collar, CS: Civil servants; TR: Transfer recipients.*
Robust standard errors clustered at the individual level in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Source: Own calculation based on the SOEP

Table A10.: Year-Specific Effects

	Self-Employed	White Collar	Blue Collar	Civil Servant
Lag of ln(Hours Worked)	0.103	0.117**	0.229***	0.058
	(0.097)	(0.048)	(0.056)	(0.123)
ln(Net Wage) Risk × year				
2003	0.041**	0.007	0.000	0.012
	(0.018)	(0.009)	(0.009)	(0.026)
2004	0.011	0.011	0.013	-0.046
	(0.022)	(0.011)	(0.010)	(0.039)
2005	0.032	0.041***	0.047***	-0.020
	(0.026)	(0.013)	(0.015)	(0.037)
2006	0.044**	0.026	0.004	-0.013
	(0.020)	(0.016)	(0.011)	(0.032)
2007	0.063***	0.020*	0.035***	-0.032
	(0.022)	(0.011)	(0.011)	(0.038)
2008	0.060*	0.026**	0.027**	-0.013
	(0.031)	(0.012)	(0.013)	(0.017)
2009	0.076**	0.031**	0.017	-0.001
	(0.030)	(0.012)	(0.014)	(0.022)
2010	0.120***	0.043***	0.020	-0.022
	(0.028)	(0.016)	(0.020)	(0.048)
2011	0.040	0.040***	0.025	0.030
	(0.034)	(0.012)	(0.017)	(0.040)
Unempl. Prob.	-0.007	0.003	0.002**	-0.000
	(0.006)	(0.002)	(0.001)	(0.005)
ln(Marginal Net Wage)	0.119***	0.133***	0.061***	0.243***
	(0.041)	(0.020)	(0.023)	(0.092)
Controls	✓	✓	✓	✓
Observations	864	5,652	2,987	1,407
AR(1) in FD	0.000	0.000	0.000	0.001
AR(2) in FD	0.666	0.954	0.390	0.331
Hansen	0.229	0.227	0.027	0.312

Notes: Estimation of equation (4.5) *using the SYS-GMM as in column 6, Table 4.2.*
Robust standard errors clustered at the individual level in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Source: Own calculation based on the SOEP

5. What is Wrong with the CRRA Euler Equation for Labor Supply?[1]

5.1. Introduction

The Frisch elasticity of labor supply is a key policy parameter that is crucial for the analysis of various economic issues such as business cycles or reforms of the tax and transfer system. A commonly applied estimation method relies on the approximation of the Euler equation derived from a life cycle model of labor supply and consumption (see, among many others, Altonji, 1986). In this paper we apply this method to estimate the Frisch elasticity for married men in the US. In line with the extant literature, we estimate a Frisch elasticity of .35.

In a next step we show how the parameters of relative risk aversion, and the higher order risk attitude prudence with respect to leisure impact labor supply. This is done by deriving the second-order Taylor approximation of the Euler equation for labor supply. Studies on life-cycle labor supply commonly assume that individuals maximize separable expected utility of the constant relative risk aversion (CRRA) form. Expected utility ties attitudes to risk to intertemporal substitution. In the context of labor supply, this implies a direct, testable relationship between precautionary labor supply, i.e., labor supply as a response to risk, and the Frisch elasticity of labor supply. While the estimate for the Frisch elasticity of labor supply obtained from the second-order approximation is similar to the one obtained from the first-order approximation commonly estimated in the literature, we find that the assumptions of CRRA expected utility are too restrictive. In particular, precautionary labor supply is less important than predicted by expected utility. A recent, burgeoning literature studies the impact of wage risk on labor supply (Pistaferri, 2003; Low, 2005; Parker et al., 2005; Flodén, 2006; Pijoan-Mas, 2006; Jessen et al., 2016). This literature puts a focus on precautionary behavior, which arises if workers are prudent.

The test of the model relies on the estimation of parameters that govern intertemporal substitution and attitudes towards variability in leisure. These parameters explain an important part of patterns of intertemporal labor supply (see the references above). As a third contribution, we describe how labor supply, hourly wage, wage risk, and the variability of leisure growth evolve over the life cycle. We show the results for the US using the panel study of income dynamics (PSID). Married men in the US work on average most from age 30 to 40 and enjoy increasingly more leisure with higher wages after age

[1]This chapter is based on Jessen and Rostam-Afschar (2017).

45. No clear pattern is apparent for the average variance of leisure growth over the life cycle, while the average variance of wage growth declines sharply at the beginning of the life-cycle and is roughly constant from age 30.

The next section derives the empirical specification from theoretical considerations, Section 5.3 briefly describes the data and the construction of key variables. Section 5.4 presents estimation results, and Section 5.5 concludes.

5.2. Theoretical Considerations and Empirical Strategy

5.2.1. Optimal Behavior

Individuals $i = 1, \ldots, N$ maximize expected utility by choosing consumption and leisure for each period t of life which ends in T.[2] Utility is additively separable across time periods and within periods. β is a discount factor.

$$\max_{c_t, l_t} E_{t_0} \left[\sum_{t=t_0}^{T} \beta^{t-t_0} u(C_t, L_t) \right],$$

subject to

$$A_{t+1} = (1 + r_{t+1})[A_t + (\bar{L} - L_t) W_s - C_t].$$

Real hourly wages W_t are uncertain and uninsurable, r_t is the real interest rate, C_t denotes consumption of the composite numeraire good, A_t the amount of assets held at the start of period t and $\bar{L}$ total annual time endowment, which is spent on leisure L_t or work.[3]

Assuming an interior solution for leisure, the first order conditions of the maximization problem are the Euler equation for consumption, where λ_t denotes marginal utility of wealth, and the condition that the marginal rate of substitution equals the marginal rate of transformation:

$$\frac{\partial u(C_t, L_t)}{\partial C_t} = \lambda_t = \beta(1 + r_{t+1}) E_t \left[\frac{\partial u(C_{t+1}, L_{t+1})}{\partial C_{t+1}} \right], \tag{5.1}$$

$$\frac{\partial u(C_t, L_t)}{\partial L_t} = W_t \frac{\partial u(C_t, L_t)}{\partial C_t}.$$

[2] We omit henceforth the individual index i.

[3] We describe the model in terms of leisure instead of the equivalent hours of work in order to stress the similar interpretation of the curvature parameters of the utility function with respect to consumption of the composite numeraire good and consumption of leisure.

We substitute to obtain the Euler equation of labor supply

$$\frac{\partial u(C_t, L_t)}{\partial L_t}\frac{1}{W_t} = \beta(1+r_{t+1})E_t\left[\frac{\partial u(C_{t+1}, L_{t+1})}{\partial L_{t+1}}\frac{1}{W_{t+1}}\right]. \tag{5.2}$$

5.2.2. Approximations of Optimal Behavior

To obtain an equation that is linear in its estimation parameters, we take a multivariate *first-order* Talyor approximation of the marginal utility of leisure per hourly wage in $t+1$, $\frac{\partial u(C_{t+1}, L_{t+1})}{\partial L_{t+1}}\frac{1}{W_{t+1}}$, around L_t and W_t.

$$\frac{\partial u(C_{t+1}, L_{t+1})}{\partial L_{t+1}}\frac{1}{W_{t+1}} \approx u_L\frac{1}{W_t} - u_L\frac{1}{W_t}(\Delta \ln W_{t+1}) + u_{LL}\frac{1}{W_t}(\Delta L_{t+1}). \tag{5.3}$$

Deriving an estimation equation from the first order approximation leads to higher order terms contained in the error term. Using instruments that are uncorrelated with higher order error terms allows a consistent estimation. Alternatively, the multivariate *second-order* Talyor approximation around L_t and W_t is applied:

$$\begin{aligned}\frac{\partial u(C_{t+1}, L_{t+1})}{\partial L_{t+1}}\frac{1}{W_{t+1}} &\approx u_L\frac{1}{W_t} - u_L\frac{1}{W_t}(\Delta \ln W_{t+1}) + u_{LL}\frac{1}{W_t}(\Delta L_{t+1}) \\ &+ u_L\frac{1}{W_t}(\Delta \ln W_{t+1})^2 + \frac{1}{2}u_{LLL}\frac{1}{W_t}(\Delta L_{t+1})^2. \\ &+ u_{LL}\frac{1}{W_t}\Delta \ln W_{t+1}\Delta L_{t+1}.\end{aligned} \tag{5.4}$$

For a similar approximation including consumption terms see Low (2005).[4] These terms are zero in our case because we assume separable utility.

The terms on the right-hand side have economic intuition as they include preference parameters that determine intertemporal substitution. In particular, workers are risk averse with a negative second derivative (concave utility). The degree of risk aversion is measured by the parameter of relative risk aversion with respect to leisure which is defined as $-L\frac{u_{LL}}{u_L}$ (Pratt, 1964).

A positive third derivative (convex marginal utility) defines prudence measured by the parameter of relative prudence with respect to leisure as $-L\frac{u_{LLL}}{u_{LL}}$ (Flodén (2006); Kimball (1990)). If a worker is only risk averse but not prudent, say, because the third derivative is zero as in the quadratic utility case, she will receive utility losses with higher

[4]Similarly, e.g. Ludvigson and Paxson (2001); Carroll (2001); Dynan (1993) apply a *univariate* second-order Taylor approximation to the consumption Euler equation.

risk but not change her decisions; whereas a prudent worker takes action to alleviate bad realizations, i.e. her savings increase in response to an increase in risk. The reason is that for prudent workers not only the marginal utility of leisure is higher when leisure is low, but also the *rate* at which the marginal valuation rises when leisure falls is greater when leisure is low than when it is high.

5.2.3. Estimation Equations under CRRA Utility

Assume constant relative risk aversion (CRRA) utility with

$$u(C,L) = \frac{C^{1-\gamma}}{1-\gamma} + \frac{L^{1-\rho}}{1-\rho}, \tag{5.5}$$

In this case, the parameter of relative risk aversion is $-L\frac{u_{LL}}{u_L} = \rho$ and the parameter of relative prudence is $-L\frac{u_{LLL}}{u_{LL}} = \rho + 1$.

Starting with the first-order approximation, we divide expression (5.3) by u_L and substitute the approximation back into equation (5.2). After backdating, resolving expectations, rearranging, and adding taste shifters X_t we obtain our estimation equation[5]

$$\Delta \ln L_t = \frac{1}{\rho}(r_t - \delta) - \frac{1}{\rho}\Delta \ln W_t + \psi X_t + \varepsilon_t. \tag{5.6}$$

The error term contains higher order terms of the approximation and taste shocks. Approximation around the expectation error of λ_t yields the same estimation equation. This equation was used, e.g., in MaCurdy (1981), Altonji (1986), Domeij and Flodén (2006), and Peterman (2016) for labor supply, and in the context of consumption, e.g., in Browning and Lusardi (1996) and Gruber (2013). The first two terms on the right hand side of this equation capture the effect of impatience, $r < \delta$ (Deaton, 1991), and of the expectation of the change in wages. If workers are patient, i.e. $r > \delta$, leisure increases with time; if wages increase, leisure decreases. $1/\rho$ is the Frisch elasticity of labor supply. To capture time variation of the interest rate, we include year dummies in our empirical specification.

We analogously obtain from the second-order approximation, expression (5.4)

$$\Delta \ln L_t = \frac{1}{1-\Delta \ln W_t}\left\{\frac{1}{\rho}(r_t - \delta) - \frac{1}{\rho}\Delta \ln W_t + \frac{1}{\rho}[\Delta \ln W_t]^2 + 1/2(1+\rho)[\Delta \ln L_t]^2\right\} + \psi X_t + \varepsilon_t. \tag{5.7}$$

The effect of uncertainty is captured by the quadratic terms: An indirect precautionary effect is that the variability in wage leads to leisure being deferred. A direct precautionary effect is that increased variability in leisure leads to leisure being deferred.

[5]Note that $\frac{\Xi_{t+1}-\Xi_t}{\Xi_t} \approx \Delta \ln \Xi_{t+1}$ and that for small values of δ and r, $1 - \frac{1}{\beta(1+r)} = 1 - \frac{1+\delta}{1+r} \approx r - \delta$ with $\beta = \frac{1}{1+\delta}$.

Thus the approximation suggests that precautionary motives lead individuals to work more today, and to work less in the future in the presence of wage uncertainty and variability in leisure (Low, 2005). Instead of a univariate approximation commonly estimated, we apply a multivariate approximation following Low (2005). Therefore, our specification includes a cross-term, namely $\frac{1}{1-\Delta \ln W_t}$. The estimation of Equation (5.7) allows for a direct test of the underlying model based on the coefficients of wage growth and leisure and wage variablity.

5.2.4. Estimation Issues

We need to instrument the estimation equations (5.6) and (5.7) because of measurement error, and because higher order terms of the approximation are omitted. We assume these not to be serially autocorrelated. Hourly wages are constructed by dividing annual labor income by hours. If hours are measured with error, this introduces a negative correlation between wage and hours (denominator bias, see Borjas, 1980; Altonji, 1986; Keane, 2011). Instruments must be uncorrelated with taste shocks contained in ε_t. A widely used instrument for wage growth is based on the lagged change in *directly asked* hourly wage as instrument for wage growth, which is available in the PSID.

Unfortunately, the direct wage measure is only available for a subset of workers. Therefore, we use the lagged constructed hourly wage as instrument. If the measurement error in hours is uncorrelated across periods, this instrument is not prone to the denominator bias. As an alternative instrument, we use lagged labor income which contains information that individuals use to form expectations about wage growth and is valid even if the measurement error in work hours is correlated over time.

For the second-order equations, we require additional instruments because the quadratic and cubic terms might be correlated with higher order moments in the error term (approximation bias) and because they might be measured with error. Similarly to Dynan (1993), we argue that lags of variability of leisure and wage growth as well as of income from other sources than labor of the household head influence the time-series properties of an individual's wage and leisure. For instance, workers with high wage risk in the past may have higher variability of wage growth in the current period. These differences in second moment patterns will affect leisure decisions and are likely to produce differences in the variability in leisure growth.

Our estimation equations include year dummies to capture time variation of the interest rate. For the second equation these year dummies are transformed with the cross terms given in equation (5.7). We instrument them with their lags modified with labor income instead of the wage.

The identification of approximated Euler equations has been debated (Blundell et al., 2007) because it is difficult to find valid and relevant instruments (Carroll, 2001; Ludvigson and Paxson, 2001; Alan et al., 2012). Attanasio and Low (2004) examine this using a Monte Carlo analysis in a wide variety of settings and conclude that under specific conditions linearization is unlikely to bias the results in a serious way. Moreover, Domeij and Flodén (2006) show that the approximation bias is relatively small for the estimation of labor supply equation (5.6).

5.3. Data

We use data from 1971 to 1996 from the Panel Study of Income and Dynamics (PSID, Panel Study of Income Dynamics (2016)), a representative household survey for the US. We restrict our analysis to the years prior to 1997 because from this year onwards, the data were only collected biannually. We restrict our analysis to married males aged 26 to 55 to ensure comparability to the previous labor supply literature and because the extensive margin plays an important role for female labor supply.[6] We drop households with unrealistic values or jumps for annual work hours or wages. In particular, following Domeij and Flodén (2006) we drop observations with increases in hourly wages or annual work hours by at least 350 percent or decreases by at least 60 percent. Moreover, we drop observations with more than 5000 annual hours.

Table 5.1 shows summary statistics of the variables that we include in our specifications. The first variable is the constructed hourly wage, $\frac{INC}{H}$, i.e. individual annual labor income divided by individual annual hours worked. W_{direct} denotes the directly reported hourly wage. Its mean is lower than that of the constructed hourly wage, because the directly reported wage is only available for a subset of workers. Following Domeij and Flodén (2006), annual leisure is constructed by subtracting the annual work hours from 5000. The following variables are the additional regressors included in the regression estimations obtained from second-order approximations, measures for wage risk and variability of leisure.

[6]Only 2.7 percent of all observations of married men aged 25 to 60 are unemployed in the PSID. We could have used the linear probability model to construct a selectivity-correction variable to be included in the second-step estimation of the conditional hours equation. Since this would have required the choice of some exclusion restrictions which, in the present context, are difficult to substantiate convincingly, we chose to estimate without explicit selectivity correction under the assumption that selection bias is negligible.

Table 5.1.: Summary Statistics

	Mean	SD	Min	Max
INC/H	30.01	26.51	1.92	934.34
W_{direct}	21.91	8.25	1.25	141.96
L	2669.973	469.902	8.000	3499.000
$[\Delta \ln INC/H]^2$	0.06	0.13	0.00	1.50
$1/2[\Delta \ln L]^2$	0.03	0.26	0.00	15.83
INC	69352.10	61722.78	3375.56	1719188.42
$year$			1971	1996
Observations	25740			

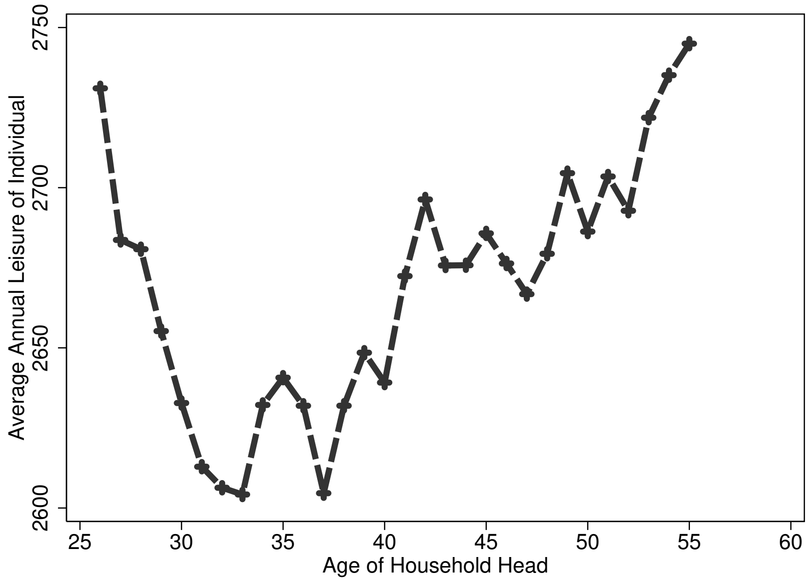
Note: All monetary variables are measured in 2005 Dollar. L is measured in annual hours.

Source: Authors' calculations based on the PSID (1969-1996).

Before turning to the empirical analysis, it is instructive to describe the growth pattern of hours and wages over the life cycle. Figures 5.1 and 5.2 show the age profiles of annual hours of leisure and hourly wages for our estimation sample. We do not account for cohort effects. Leisure follows a checkmark or u-shape over the life cycle. As described above, precautionary behavior can lead to an increase in leisure over time and is thus a potential part of the explanation for this pattern. The wage profile is increasing and concave.

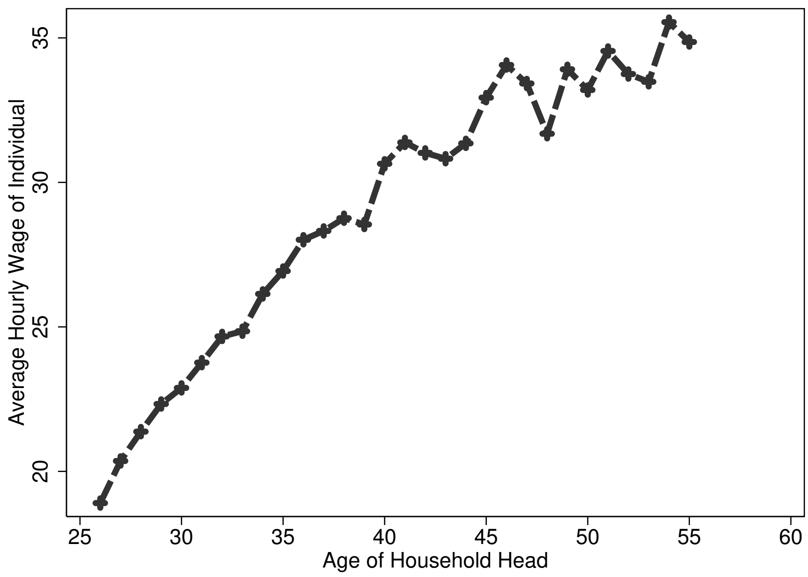
Figure 5.3 shows the path of the variability of leisure divided by two as in the estimation equation (5.7). There is substantial variation in this measure over the life-cycle with no clear trend. Figure 5.4 shows the variability of wages. Wage risk is decreasing for young workers and then roughly constant over the life cycle.

Figure 5.1.: Average Annual Leisure (5000-Hours of Work) of Married Men over Age



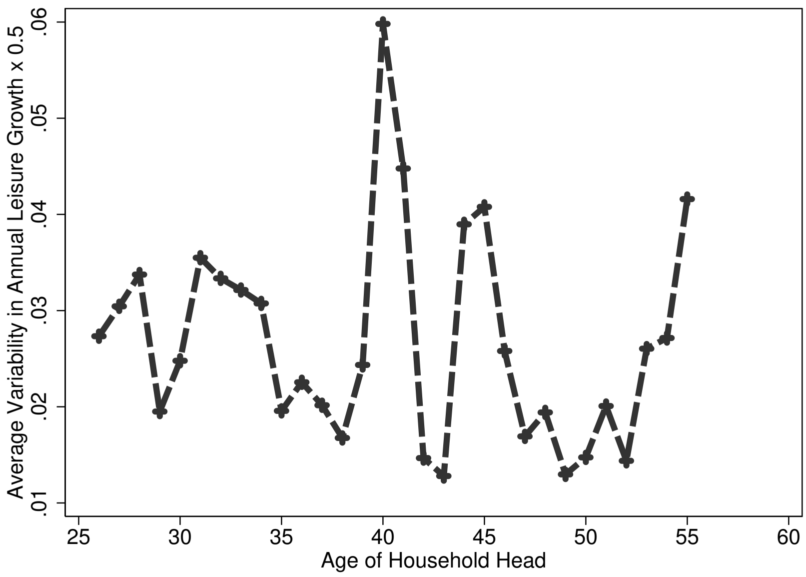
Source: Own calculations based on the PSID (1971-1996).

Figure 5.2.: Average Hourly Wage in 2005 PPP USD of Married Men over Age



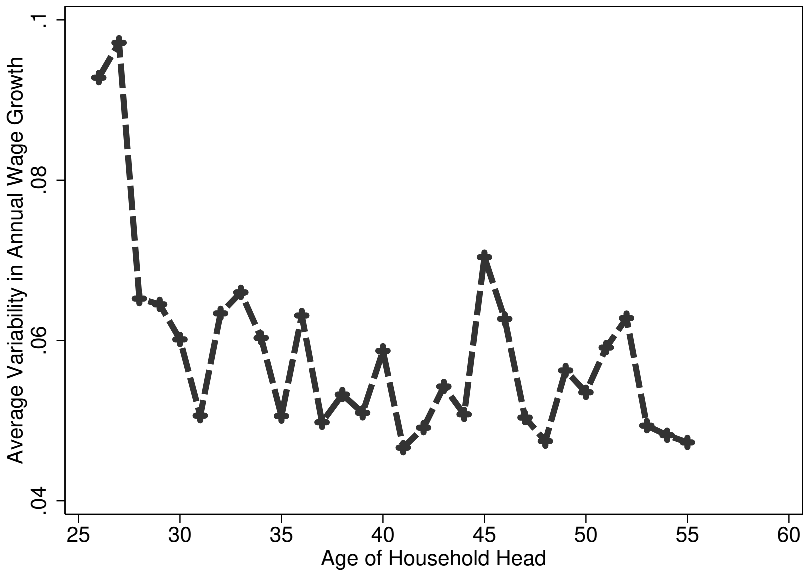
Source: Own calculations based on the PSID (1971-1996).

Figure 5.3.: 0.5×Average Variability in Annual Leisure Growth of Married Men over Age



Source: Own calculations based on the PSID (1971-1996).

Figure 5.4.: Average Hourly Wage Risk of Married Men over Age



Source: Own calculations based on the PSID (1971-1996).

5.4. Estimation Results

Table 5.2 displays results for the estimation of equation (5.6) using PSID data.[7] The displayed coefficient is $-1/\rho$, the Frisch elasticity of leisure. This estimation is directly comparable to the literature, but it does not allow to test the theoretical restrictions.

Table 5.2.: Regression of Growth Rate of Leisure (First-Order Approximation)

	I	II	III	IV
	OLS	II $W_{\text{t-1,direct}}$	III $\frac{inc_{t-1}}{h_{t-1}}$	IV inc_{t-1}
$\Delta\ln(INC/H)$	0.320***	0.224	0.354***	-0.353***
	(0.012)	(0.188)	(0.021)	(0.053)
N	26049	7979	20629	20629
Hansen's J statistic		3.593	0.343	1.370
p-value (Hansen's J)		0.058	0.558	0.242

Estimation Equation: Results from estimation of $\Delta \ln L_t = \text{Constant} + \frac{1}{\rho}(r_t - \delta) - \frac{1}{\rho}\Delta \ln INC/H_t + X_t + \varepsilon_t$. Dependent variable is the growth rate of leisure (5,000 - hours).
Control variables: year dummies.
Instruments: Except in column I, the independent variable is instrumented. In column II ($W_{\text{t-1,direct}}$), the instrument is the lagged level and the lagged difference of the log of the directly asked hourly wage. In column III ($\frac{inc_{t-1}}{h_{t-1}}$), instruments are identical except that they are based on hourly wage constructed as the fraction of income and hours. In column IV (inc_{t-1}), instruments are identical except that they are based on labor income.
Inference: Cluster robust standard errors are in parentheses, significance levels are * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.
Source: Own calculations based on the PSID.

The first column shows the estimates of an OLS estimation. The coefficient has a positive sign, which contradicts theory. This shows that the use of instruments makes a great deal of sense because of the denominator bias, which leads to an upward bias of the coefficient.

Column II reports results obtained using instruments based on the directly reported wage using the limited information maximum likelihood (LIML) estimator, which is approximately median unbiased and better centered than the 2SLS estimator (Staiger and Stock, 1997). Due to the small number of observations, the coefficients are not statistically significantly different from zero. Column III shows results using instruments based on lags of the constructed wage. The obtained coefficients are highly significant and not in line with theory, which suggests that the measurement error in hours—which

[7] We use Stata using the ado-files ivreg2 (Baum et al., 2016b) and estout (Jann, 2007) to generate the results in all tables.

leads to the denominator bias—is correlated over time. The last column shows our preferred estimates which are obtained using instruments based on the lag of labor income. We estimate a Frisch elasticity of labor supply of 0.35, which implies that the parameter ρ equals 2.9. This is in line with estimates reported in Blundell and Macurdy (1999), Blundell et al. (2007) and Keane (2011).

For the IV estimators we report Hansen's J statistic of the test of overidentifying restrictions and the corresponding p-value (Hansen, 1982). For all estimates the null hypothesis that the instruments are valid is not rejected at the five percent level, for our preferred specification it is not rejected at the ten percent level.

Table 5.3.: Regression of Growth Rate of Leisure (Second-Order Approximation)

	I OLS	II IV $W_{\text{t-1,direct}}$	III IV $\frac{inc_{t-1}}{h_{t-1}}$	IV IV inc_{t-1}
$\Delta \ln(INC/H)$	0.313***	0.557	0.279	-0.278***
	(0.013)	(0.772)	(0.302)	(0.079)
$[\Delta \ln(INC/H)]^2 \frac{1}{1-\Delta \ln INC/H}$	-0.317***	-0.169	-0.786*	0.194
	(0.013)	(1.230)	(0.333)	(0.102)
$[\Delta \ln L]^2 \frac{1}{2(1-\Delta \ln INC/H)}$	0.013	-0.046	0.516*	0.581*
	(0.012)	(0.268)	(0.247)	(0.246)
N	25740	5849	16217	16217
Hansen's J statistic		7.966	4.077	8.802
p-value (Hansen's J)		0.241	0.666	0.185
χ^2_{CRRA}	303.162	0.412	4.276	14.934
p-value$_{\text{CRRA}}$	0.000	0.814	0.118	0.001

Estimation Equation: Results from estimation of $\Delta \ln L_t = \text{Constant} + \frac{1}{1-\Delta \ln INC/H_t} \left\{ \frac{1}{\rho}(r_t - \delta) - \frac{1}{\rho} \Delta \ln INC/H_t + \frac{1}{\rho}[\Delta \ln INC/H_t]^2 + 1/2(1+\rho)[\Delta \ln L_t]^2 \right\} + X_t + \varepsilon_t$. Dependent variable is the growth rate of leisure (5,000 - hours).
Control variables: year dummies.
Instruments: all independent variables are instrumented. In column II (IV $W_{\text{t-1,direct}}$), instruments are the lagged level and the lagged difference of the log of the directly asked hourly wage, the lagged level and the lagged difference of the log of other household income from all sources, and the lagged level and the lagged difference of the square of log leisure. Instruments for modified year dummies are their lags. In column III (IV $\frac{inc_{t-1}}{h_{t-1}}$), instruments are identical except that they are based on hourly wage constructed as the fraction of income and hours. In column IV (IV inc_{t-1}), instruments are identical except that they are based on labor income.
Inference: Cluster robust standard errors are in parentheses, significance levels are * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.
Source: Own calculations based on the PSID.

Table 5.3 displays results of the estimation of equation (5.7) using the OLS estimator as well as IV using different sets of instruments corresponding to those used in the estimations reported in Table 5.2.[8] The first reported coefficient is the Frisch elasticity, the second denotes labor supply reactions to wage risk and the third coefficient shows reactions to variability in leisure.

Recall that according to equation (5.7) the effect of wage growth on leisure growth should be $-1/\rho$ according to theory, while the effects of variability in wage and leisure should be $1/\rho$ and $1+\rho$ respectively. That is, the first two coefficients should have the same magnitude but opposite signs. We test the theoretical restriction on the coefficients of the three regressors (where the last one includes the scaling factor 1/2), i.e. we test if the three transformed coefficients are significantly different from one another. The null hypothesis is that of a Wald tests of nonlinear restrictions. $H_{\text{CRRA}}: -1/(-1/\hat{\rho}) = 1/(1/\hat{\rho}) = (1+\hat{\rho})-1$ or $\hat{\rho} = \hat{\rho} = \hat{\rho}$. The χ^2 statistic and p-value for this test are reported below the estimates

For OLS the magnitudes of the first two regressors are virtually the same, but the signs contradict theory. The magnitude of the coefficient of the variability of leisure is smaller than it should be according to theory. The null hypothesis that the theoretical restriction holds is resoundingly rejected.

The second column show results using instruments based on the lag of the directly reported hourly wage. Using the directly reported wage as IV yields very imprecise estimates which do not allow to reject the theoretical restriction. Column III shows results using IVs based on the lagged constructed wage. As for the first-order approximation, this yields coefficients whose sign contradicts theory.

Column IV shows the preferred estimation using IVs based on lagged labor income. The Frisch elasticity is similar to the one obtained using the first-order approximation. This suggests the commonly used estimation equation based on the first-order approximation yields a good estimate of the Frisch elasticity. In line with theory the coefficient of wage variability is of similar magnitude but has the opposite sign. However, the coefficient of leisure variability is just .58. Based on the point estimate of wage growth, theory predicts a magnitude of 4.6. The positive coeffcent shows that precautionary behavior has a role, but it is less important than it would be under expected CRRA utility. The restriction of the theory is rejected with a p-value of 0.001.

5.5. Summary and Conclusions

Deriving estimation equations from a theoretical model, we have studied the impact of wage uncertainty and variability in leisure on labor supply. Our specification extends

[8]We have experimented with using further lags as instruments, however, their explanatory power is weak and the coefficients become insignificant. Adding further controls (age, state dummies) did not alter the results substantially.

the conventional one and allows to test implications of alternative utility functions. In particular, we derive testable restrictions for the key policy parameters of risk aversion and prudence with respect to leisure. We apply the method to the widely used CRRA utility function.

In a first step, we estimate the commonly estimated labor supply estimation based on the first-order approximation of the Euler equation. We find that the Frisch elasticity for US married men is 0.35.

The Frisch elasticity is a key parameter to explain labor supply patterns over the life cycle. We document these patterns and find that in the beginning of economic life average work hours increase and peak between 30 and 40. From then onwards, leisure increases steadily towards retirement age. The hourly wage is concave over the life cycle. It increases most profoundly until about the age of 45. While there is no clear trend in the variability of leisure, wage risk is decreasing for young workers and then roughly constant over the life cycle.

Using the extended estimation specification based on the second-order approximation, we find the restrictions for the parameter of risk aversion with respect to leisure are not rejected. As predicted, the coefficients of wage growth and wage risk are of a similar magnitude. Moreover, the estimated Frisch elasticity is close to the one obtained using the first-order approximation. Thus, omitting higher order terms does not bias this estimate by much. However, the restriction that the parameter of prudence with resepect to leisure is directly tied to the parameter of intertemporal substitution is rejected.

One potential reason for this is that CRRA is too restrictive because it ties the intertemporal elasticity of substitution to risk aversion and prudence. Another restriction is the assumption of intratemporal separability of consumption and leisure. This restriction has been shown to be important by Ziliak and Kniesner (2005). Our result complements, among others, Crossley and Low (2011), who find with a different approach that theoretical restrictions of CRRA utility for *consumption* are inconsistent with features of the data.

Therefore alternative avenues should be explored. For instance, repeating a test like in our analysis based on alternative utility functions such as Cobb-Douglas (see Domeij and Flodén (2006) and Low (2005) for an investigation of this utility function for labor supply) or more general preference structures like Epstein and Zin (1989, 1991) (see Yogo (2004) for an application) could be a fruitful endeavor.

General Conclusion

Main Findings

This dissertation focuses on several issues related to labor supply and the (re)-distribution of income. All chapters use micro data to answer the research question at hand.

Chapter 1 quantifies the contribution of tax-transfer reforms and changes in conditional wage rates to the overall increase in income inequality in Germany from 2002 to 2011. I find that policy changes have led to a decrease in income inequality according to several popular inequality measures. The reason for this is the increase in the generosity of transfers. Labor supply reactions have partly offset this inequality decreasing effect because increased transfers have dampened incentives to work. Wage changes conditional on individual characteristics have led to inequality decreases as well. The result implies that the overall increase in income inequality is entirely due to changes in the population that are not explicitly modeled. These include changes in the composition and size of households.

Chapter 2 generalizes an optimal taxation model to allow for non-welfarist aims of the social planner. In particular, we allow for the social planner's objective function to depend on individual's characteristics that do not enter the utility function. The exercise is motivated by the puzzling fact that if a social planner maximizes a weighted sum of utilities, high transfer withdrawal rates in many countries are only optimal if social weights for the working poor are very low. In contrast, optimal taxation studies usually assume that the social planner puts more weight on low income than on higher income households. The main finding of this chapter is that the German tax and transfer schedule is in line with decreasing social weights if the social planner minimizes absolute sacrifice, i.e., a function that depends on individual's tax liability.

While Chapters 1 and 2 were concerned with income inequality and redistribution, income inequality is usually accompanied by its similarly evil twin brother, income uncertainty, whose interaction with labor supply is the focus of the next three chapters.

In Chapter 3, we find that, in contrast to permanent wage shocks, permanent hours shocks are virtually nonexistent for married males in the US. A likely reason for this finding is that permanent hours shocks such as serious injuries in practice affect the extensive margin, which is not captured by the model. Modeling the extensive margin would be especially important for females. The second finding is that transitory wage shocks are more important than transitory hours shocks in terms of income, but both shocks play an important role. At mean income, a negative transitory wage shock of one standard deviation implies a reduction of annual income of 12,024 Dollar, while an hours shock of one standard deviation leads to a change in annual labor income of 7,186

Dollar. The third finding is that the Marshall elasticity for married males is negative, -0.7. The result is obtained using a sufficient statistic, the effect of permanent wage shocks on labor supply.

In Chapter 4 we estimate a dynamic labor supply equation to quantify the importance of precautionary labor supply in Germany, defined as the difference between hours worked under wage uncertainty and under perfect foresight. We find that married men in Germany on average choose about 2.8% of their hours of work to shield against wage shocks. The effect is strongest for self-employed, but also relevant for other groups. If the self-employed faced the same wage risk as the median civil servant, their hours of work would reduce by 4.5%.

In Chapter 5 we estimate the second order Taylor approximation of the CRRA Euler equation for labor supply for married males in the US. The estimate of the Frisch elasticity from the second order approximation is similar to the one obtained from the first order approximation commonly estimated in the literature starting with MaCurdy (1981), about 0.3. However, the first order approximation does not tell the whole story. While agents use precautionary labor supply to cushion against wage risk, they do so to a lesser extent than predicted by CRRA utility. Therefore the model is rejected.

Future Research

The results obtained in the chapters of this dissertation suggest more research in various directions. Chapter 1 finds that changes in conditional wage rates and the tax-transfer system do not explain the increase in income inequality in Germany from 2002 to 2011. The microsimulation-based decomposition method could be combined with the reweighting technique by DiNardo et al. (1996) – as in Herault and Azpitarte (2016) for Australia – in order to shed some light on the interaction of changes in household structure and characteristics on one side and changes in the tax-transfer system on the other side. Peichl et al. (2012) find that changes in household structure and employment patterns explain 78 percent of the increase in gross income inequality in Germany from 1991 to 2007, but only 22 percent of the increase in net income inequality. It would be interesting to investigate whether *changes* in the tax-transfer system have offset the inequality increasing changes of household structures. Moreover, future research should investigate the impact of tax-transfer reforms on life-time income and consumption inequality. In order to fully capture the effects of reforms on labor supply incentives, this task would require the simulation of labor supply over the entire life-cycle.

The finding in Chapter 2 that the current German tax-transfer system is not optimal under welfarism and social weights that decrease with income provokes further questions: Does this finding hold over the life-cycle or only in a one-period model? If it holds, how are the ideas of justness underlying the design of the tax-transfer system formed?

Chapter 3 finds that both wage shocks and hours shocks are important sources of labor income uncertainty. A natural way to build on this finding is to investigate, what policy measures could alleviate uncertainty stemming from either source. A second finding is a larger negative Marshall elasticity than found in comparable papers such as Heathcote et al. (2014) and Blundell et al. (2016). The likely reason is that our estimate is based on a sufficient statistic using labor supply data alone, while the methods of the other two papers require the use of consumption data. It would be worthwhile to reconcile the methods and investigate further how permanent wage shocks are insured.

In Chapter 4 we find that precautionary labor supply is important. While this finding is robust for several measures of wage uncertainty, it is not clear, how individuals actually form views about wage uncertainty. It would be worthwhile to model the process how individuals update their beliefs about wage uncertainty. The approach could draw from Guvenen (2007), who, in a setting of heterogeneous income processes, models how individuals update their beliefs about their individual income process as more and more information becomes available. The lesson that net wage uncertainty matters for labor supply is important for the design of the tax and transfer system. While the optimal taxation literature has considered the potentially negative labor supply incentive effect of increased insurance in abstract models, e.g., Eaton and Rosen (1980a), simulations of proposed tax-transfer reforms should also incorporate this relationship.

The finding in Chapter 5 that the life-cycle labor supply model with CRRA utility is rejected, provokes two obvious research questions. First, how big is the mistake that one makes when assuming CRRA utility? Second, how can more general models of utility optimization, in particular preferences that loosen the tie between intertemporal substitution and risk aversion be incorporated in the life-cycle labor supply model? I am currently working on these questions.

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