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# How important is precautionary labour supply?

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## Abstract

We quantify the importance of precautionary labour supply defined as the difference between hours supplied in the presence of risk and hours under perfect foresight. Using the German Socio-Economic Panel from 2001 to 2012, we estimate the effect of wage risk on labour supply and test for constrained adjustment of labour supply. We find that married men choose on average about 2.8% of their hours of work to shield against wage shocks. The effect is strongest for self-employed, who we find to be unconstrained in their hours choices, but also relevant for other groups with more persistent hours constraints. If the self-employed faced the same wage risk as the median civil servant, their hours of work would be reduced by 4.5%.

**JEL classifications**: C23, D12, E21, J22

## 1. Introduction

This study quantifies the importance of precautionary labour supply, defined as the difference between hours supplied in the presence of risk and hours supplied under certainty. Facing a higher future wage risk, individuals may increase their hours worked in order to insure themselves against bad realizations. Our study provides first empirical evidence for this theoretically predicted phenomenon.

A thorough understanding of labour supply incentives over the life cycle is crucial for understanding household behaviour and is of primary interest for both labour and macroeconomics (Meghir and Pistaferri, 2011). Relevant precautionary labour supply could explain differences in hours worked across occupations or why the self-employed work more hours than employees for a given wage. The extent of precautionary labour supply is key for various policy issues, for instance the optimal design of social security programs.

Our approach allows us to calculate how labour 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.

We find that 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 labour 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 the self-employed faced the same wage risk as the median civil servant, their hours of work would be reduced by 4.5%.

To understand the mechanics behind these results, first consider a standard textbook life-cycle model with exogenous income, where individuals only choose consumption and savings (Deaton, 1992, chap. 6). Here, the precautionary saving motive results from uncertainty in income and *prudence*, i.e., decreasing (absolute or relative) risk aversion (Kimball, 1990).[1] For prudent individuals, lower levels of consumption increase the effect of Jensen's inequality, i.e., the negative effect of risk on expected utility is stronger if consumption is low. Thus, prudent individuals save more in order to defer consumption in the face of future risk (see, e.g., Carroll and Samwick [1998] or Parker and Preston [2005] for empirical evidence). Now consider a model with endogenous labour supply, where labour income uncertainty results from wage risk. Under the plausible assumption that the labour supply elasticity is not strongly negative, increases in the hourly wage rate translate to increases in labour income.[2] Therefore, increases in wage risk also translate into increases in income risk that may amplify the precautionary saving motive. With flexible labour supply, additional savings are achieved not only by reducing consumption, but also by increasing labour supply in a given period. These theoretical predictions are derived in Pistaferri (2003), Low (2005), and Flodén (2006).[3]

The empirical relationship between risk and hours of work has been documented to be positive for self-employed men in the USA (Parker *et al.*, 2005), male employees in the USA (Kuhn and Lozano, 2008), and German and US workers (including self-employed) of both sexes (Bell and Freeman, 2001). For Italy, Pistaferri (2003) finds a small, but economically negligible, effect of subjective wage risk on labour supply. Benito and Saleheen (2013) show that men and women use hours worked to shield themselves against subjectively perceived financial shocks. We contribute to this literature as the first study that quantifies the amount of precautionary labour supply.

1 Formally, a measure of the strength of prudence is defined as $\frac{-u'''(c)}{u''(c)}$, where $u'''(c)$ and $u''(c)$ denote the third and second derivatives of the utility function with respect to consumption. An individual is prudent if $u'''(c) > 0$.

2 An increase in wage rates translates to an increase in income, even if the income effect dominates the substitution effect, as long as the labour supply elasticity is not below –1. To see this, denote the labour supply elasticity by $e_{hw} = \frac{\partial h}{\partial w}\frac{w}{h}$. Abstracting from taxes, labour income is given by $y = hw$, where $h$ denotes hours of work and $w$ the hourly wage. A marginal increase in the hourly wage leads to an increase of labour income by $\frac{\partial y}{\partial w} = \frac{\partial h}{\partial w} w + h$. Substituting $e_{hw}\frac{h}{w}$ for $\frac{\partial h}{\partial w}$, we obtain $\frac{\partial y}{\partial w} = (e_{hw} + 1)h$, which is positive if $e_{hw} > -1$.

3 Other papers study the relationship between uncertainty and labour supply in settings without saving (Block and Heineke, 1973; Eaton and Rosen, 1980a,b; Hartwick, 2000; Menezes and Wang, 2005) and reach ambiguous conclusions.

In addition, we contribute to the empirical literature with several innovations: first, we use an objective measure of wage risk based on net-of-tax income. In our main specification, we measure wage risk as the standard deviation of past hourly individual net wages. For the precautionary motive, net-of-tax income is relevant. Hence, we calculate marginal net wages using the tax-transfer-microsimulation model STSM (*Steuer-Transfer-Simulations-Modell*; see Steiner *et al.*, 2012).[4] Thus, 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 behaviour, as argued, e.g., in Fossen and Rostam-Afschar (2013). Second, we specify a dynamic labour 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 labour supply. Third, we also control the individual probability of unemployment calculated similarly to Carroll *et al.* (2003).[5]

A caveat is that our results are limited to the partial equilibrium case. However, the evidence for the empirical relevance of precautionary labour supply provided in this paper is important to assess the overall effect of wage risk taking general equilibrium effects into account.[6]

The next section describes our data set and construction of the measure of wage risk and probability of unemployment. Section 3 presents our empirical specification and the estimation methods. Section 4 discusses the main results and occupation-specific findings. In Section 5, we quantify the importance of precautionary labour supply. Section 6 shows that the results are robust, and Section 7 concludes.

## 2. Data

Our study uses data from the German Socio-Economic Panel (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 to 2012 and focus on men because the extensive margin plays an important role in women's labour supply decisions. The sample is restricted to prime-age (older than 25 and younger than 56) married men working at least 20 hours to allow comparisons with the canonical labour supply literature, e.g., Altonji (1986) and MaCurdy (1981).[7] 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

4 The STSM is comparable to FORTAX for the UK (Shephard, 2009) or TAXSIM for the USA (Feenberg and Coutts, 1993).

5 Note that we focus solely on labour supply. The joint investigation of precautionary savings using consumption data is beyond the scope of this study.

6 A few papers study labour supply and precautionary considerations in general equilibrium models. Pijoan-Mas (2006) shows that additional hours of work are a quantitatively important smoothing device. Marcet *et al.* (2007) demonstrate that under reasonable parameter configurations, a wealth effect that reduces labour supply may dominate the positive precautionary saving effect on aggregate output documented in Aiyagari (1994) and Huggett (1993).

7 Including workers with less than 20 weekly hours virtually does not affect the results.

less than 80 hours per week. In total, we observe the main wage risk measure for 10,987 data points from 2,488 persons.[8]

## 2.1 Marginal net wage

In a progressive tax system, where $w_{it}^{\text{gross}}$ denotes hourly gross wage, $h_{it} = \bar{l} - l_{it}$ annual hours of work (or equivalently maximum annual leisure $\bar{l}$ minus chosen leisure $l_{it}$), and $T_{it}(w_{it}^{\text{gross}} \times h_{it})$ is a convex function of annual gross income $y_{it} = w_{it}^{\text{gross}} \times h_{it}$ that returns tax liabilities, the marginal net wage is defined as

$$-\frac{\partial NetInc(y_{it})}{\partial l_{it}} = -\frac{\partial\{w_{it}^{\text{gross}} \times [\bar{l} - l_{it}] - T_{it}(w_{it}^{\text{gross}} \times [\bar{l} - l_{it}])\}}{\partial l_{it}}$$
$$= (1 - T'_{it}(w_{it}^{\text{gross}} \times h_{it}))w_{it}^{\text{gross}} = w_{it}. \quad (1)$$

In a standard static labour supply model, individuals' labour supply responds to the marginal net wage, i.e., net-of-tax income per additional time spent on work. The reason is that at the optimum the marginal rate of substitution equals the marginal rate of transformation. The current marginal net wage is the price at which leisure is transformed into consumption in the respective year, i.e., $-w_{it}$ is the slope of the static budget constraint.

To construct the marginal net wage, first we calculate the hourly gross wage $w_{it}^{\text{gross}}$ by dividing annual gross labour 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 (for more information, see Steiner *et al.* [2012]; Jessen *et al.* [2017]). We obtain marginal net wage rates by scaling the gross wage $w_{it}^{\text{gross}}$ with the marginal net-of-tax rate. The marginal net-of-tax rate depends on the household context due to joint taxation and interactions with the transfer system. Define the net-of-tax rate as the net-of-tax income per euro of additional pre-tax income due to an increase in hours of work. Then, the marginal hourly net wage is given by:

$$w_{it} = \text{Net-of-tax rate}_{it} \times w_{it}^{\text{gross}} = \frac{NetInc(y_{it} + \Delta y_{it}) - NetInc(y_{it})}{\Delta y_{it}} w_{it}^{\text{gross}}. \quad (2)$$

$NetInc(y_{it})$ denotes net income given gross income $y_{it}$, and $\Delta y_{it}$ denotes a small increase in gross income. To calculate the net-of-tax rate over time, we increase each person's annual labour income $y_{it}$ marginally in every period.[9] In practice, the procedure to calculate the marginal net wage for a specific individual in a specific period works as follows:

1. Calculate net household income in the status quo using the STSM.
2. Increase the individual's labour income by $\Delta y_{it}$.
3. Recalculate net household income given the counterfactual increase in labour income.
4. Divide the increase in the household's net income by $\Delta y_{it}$ to obtain the marginal net-of-tax rate.

8 Table A.2 in the Online Appendix summarizes the number of observations lost due to each sample selection step.

9 We set $\Delta y_{it} = 2000$ euros, which implies an increase in labour income of about 40 euros per week. This increase ensures that atrocities in the tax-transfer system that can *locally* lead to very high or very low marginal tax rates do not contaminate the results.

5. Multiplying the marginal net-of-tax rate with the individual's gross hourly wage rate yields the marginal net wage.

Thus, while the marginal net wage refers to the individual, the household context is taken into account when calculating it. The procedure is repeated for every individual in every year, taking into account changes in the tax and transfer system or in the household context. 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.[10] 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 eq. (2). This is true if additional hours of work are fully compensated.

## 2.2 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 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 past time for each individual. Only wage observations from the current occupation are used for the construction of the risk 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:

$$\sigma_{w,it} = \sqrt{\frac{1}{\# - 1} \sum_{j=t-\#}^{t-1} \left(\ln \tilde{w}_{ij} - \ln \bar{\tilde{w}}_i\right)^2}, \tag{3}$$

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 labour 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 the Online Appendix in Table B.2.

10 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. Our measure of the hourly wage rate is based on total labour income and hours of work, so it potentially includes hours and income from secondary jobs. Hence, we are agnostic about whether individuals adjust their hours in their first or in a secondary job.

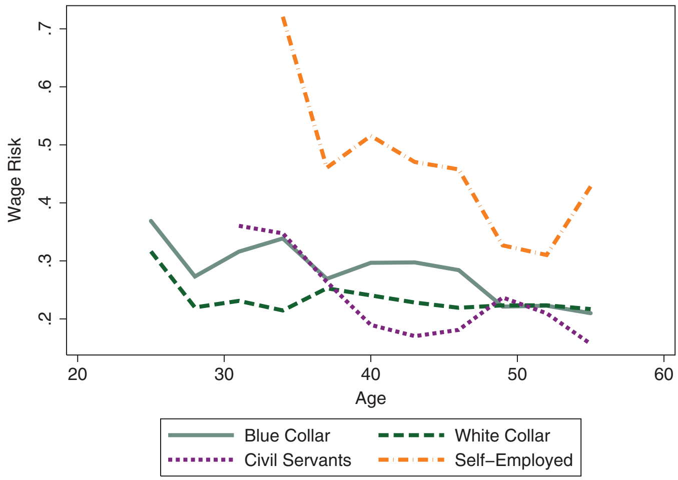


**Fig. 1.** Average net wage risk over the life cycle
*Source:* Authors' calculations.
*Note:* Standard deviations of past marginal net wages for each individual averaged over three years by occupation. We calculate the risk measure for every age for every individual based on past realizations and take the average of this measure over individuals for every age. See eq. (3).

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, e.g. Cunha and Heckman (2008), Navarro (2011), Cunha and Heckman (2016), and 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 the self-employed (see Table A1 for a detailed definition of these variables). 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 Online Appendix.

Figure 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 the 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 significant at the 5% significance level.[11] Blue- and white-collar workers have similar levels of wage risks.

11 We use a two-sample *t*-test with unequal variances to obtain the *p*-values. Test statistics are available from the authors on request.

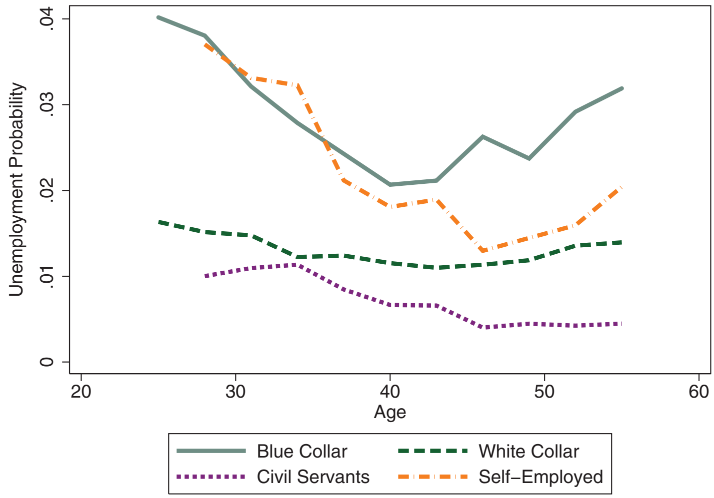


**Fig. 2.** Average unemployment probability over the life cycle
*Source:* Authors' calculations.
*Note:* Predicted probability of unemployment next year for currently working married men averaged over three years by occupation.

Nonetheless, during their thirties and forties, 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 forties.

## 2.3 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).[12] Figure 2 displays how the average unemployment probability evolves over the life cycle for the four occupational groups.[13] 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 the self-employed at younger ages and white-collar workers and the self-employed at older ages. As for the wage risk, we standardize the unemployment probability by its standard deviation for the estimations.

Table 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

12 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, and with education, marital status, and unemployment experience. The heteroskedasticity function includes previous unemployment experience and years of education.

13 As in Fig. 1, only age-occupation combinations with more than 15 observations are displayed.

**Table 1.** Summary statistics

| | Unit | Mean | Std. Dev. | Min | Max | *N* |
|---|---|---|---|---|---|---|
| Labour 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 Labour Income | (Euro) | 3,764.47 | 1,997.75 | 319 | 27,000 | 16,038 |
| Monthly Net Labour Income | (Euro) | 2,458.91 | 1,197.49 | 150 | 15,000 | 16,038 |
| Wage Risk 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 |
| 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 |
| 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 |

*Source:* Authors' calculations.
*Notes:* Data from SOEP (version 30). Sample of married prime-age males; 2001–2012.

hours worked per week, about 42 in our sample. Hourly wages average 22 euros, with average marginal net wages of 12 euros. Hourly wages are constructed by dividing gross monthly labour 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. Labour earnings include wages and salaries from all employment, including training, self-employment income, bonuses, overtime, and profit-sharing.

We use paid hours because an increase in these translates directly into an increase in income. Robustness tests using different measures of hours supplied are reported in Table B.1 in the Online Appendix. The average gross wage risk in our sample is 0.192, which is

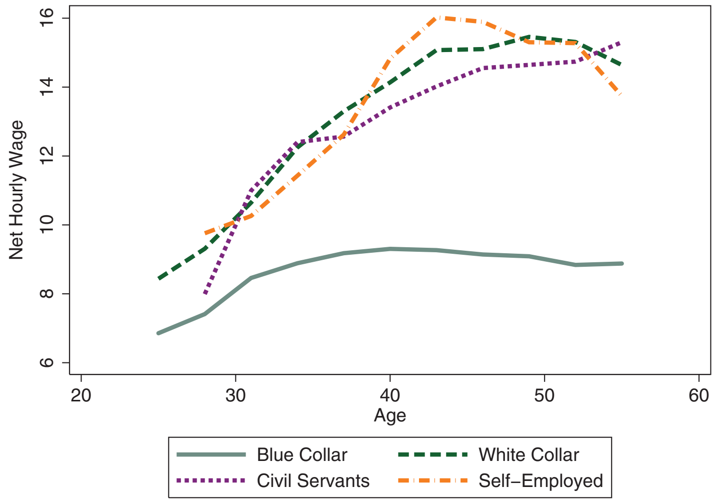


**Fig. 3.** Average marginal hourly net wage over the life cycle
*Source:* Authors' calculations.
*Note:* Marginal net wages for married men averaged over three years by occupation calculated using the STSM.

similar to the average wage risk of 0.21 reported in Parker *et al.* (2005). The last three variables in Table 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 3 shows the evolution of marginal net wages over the life cycle for different occupational groups. Profiles for white-collar workers, civil servants, and the 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 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.

## 3. Empirical strategy

### 3.1 Constrained adjustment of labour supply

We begin the investigation with the following labour supply equation, which is similar to the specification studied in Parker *et al.* (2005):

$$\ln h^*_{it} = \tilde{\beta}_1 \ln w_{it} + \tilde{\beta}_2 X_{it} + \tilde{\beta}_3 \sigma_{w,it} + \omega_{it}, \tag{4}$$

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 $\omega_{it}$ 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 (over-employment) and workers are thus 'off their labour supply curve'. Bryan (2007) uses

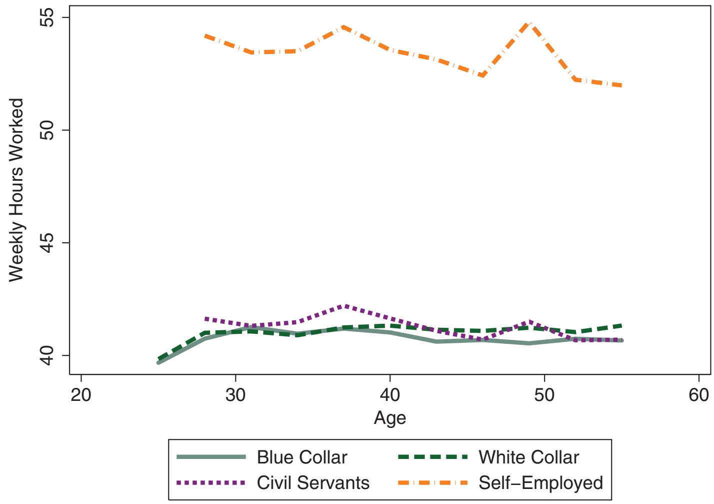


**Fig. 4.** Average weekly hours worked over the life cycle
*Source:* Authors' calculations.
*Note:* Paid hours of work averaged over three years by occupation for married men.

OLS with correction terms from a random effects ordered probit model that determines the probability of being overemployed, unconstrained or underemployed (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 (2013a,b) proposed an index (BB-index) to measure under-employment, i.e., the case that workers would like to work more hours. They find that under-employment has been substantial in the UK labour market recently. Table 1 shows that the BB-Index is positive on average in Germany as well, implying that the average person in the workforce is underemployed.[14] 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, e.g., Robins and West (1980), Euwals (2005), and Baltagi *et al.* (2005):

$$\ln h_{it} - \ln h_{it-1} = \theta(\ln h^*_{it} - \ln h_{it-1}), \qquad 0 < \theta \leq 1. \tag{5}$$

$\theta$ may be interpreted as the speed of adjustment. This speed might be determined by costs to immediately adjust the labour supply to desired hours or habit persistence (see, e.g., Brown 1952). Replace (5) in (4) to obtain the partial adjustment labour supply specification:

$$\ln h_{it} = \alpha \ln h_{it-1} + \beta_1 \ln w_{it} + \beta_2 X_{it} + \beta_3 \sigma_{w,it} + \varepsilon_{it}. \tag{6}$$

This is our empirical labour supply specification. The parameters of (4) can be recovered following the estimation of (6) with $\alpha = 1 - \theta$, $\beta_1 = \theta\tilde{\beta}_1$, $\beta_2 = \theta\tilde{\beta}_2$, $\beta_3 = \theta\tilde{\beta}_3$, and $\varepsilon_{it} = \theta$

14 Following Bell and Blanchflower (2013b), we constructed a variable that measures the probability of being under- or overemployed and included it in $X_{it}$ along with the probability of unemployment as a robustness test in Table B.3 in the Online Appendix.

$\omega_{it}$ (Baltagi *et al.*, 2005).[15] The partial adjustment model nests the classic labour supply equation with $\theta = 1$ as a special case. The short-run labour supply elasticity with respect to wage is given by $SR_{\eta_w} = \beta_1$, and the short-run labour supply elasticity with respect to wage risk by $SR_{\eta_{\sigma w}} = \beta_3$. The corresponding long-run elasticities are $LR_{\eta_w} = \beta_1/(1-\alpha)$ and $LR_{\eta_{\sigma w}} = \beta_3/(1-\alpha)$.

### 3.2 Instrumentation and estimation methods

To estimate our labour supply equation, we need to account for several sources of endogeneity. First, the first difference of the lagged dependent variable is correlated with the error term $\varepsilon_{it}$, 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's (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 labour 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 non-linear tax and transfer system. Therefore, we instrument marginal net wages with the first lag of net labour income. This variable is predetermined during the current period labour supply choices and uncorrelated with the measurement error in current period hours.

## 4. Results

### 4.1 Impact of wage risk on weekly hours of work

Table 2 presents the results of the augmented labour supply equation for different estimators, where the dependent variable is the log of paid hours of work. In all specifications, we control for year dummies, age, age squared, education, the presence of children of different age groups, labour market experience, and a dummy for East Germany. 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 $\alpha$ in eq. (6) is restricted to zero. Columns 4–6 show results for the preferred dynamic specification. Table C.1 in the Online Appendix shows the equivalent table using gross wages instead of marginal net wages. 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 labour supply by 2.8%. The coefficient on unemployment probability is very small and not statistically significant.

15 Note that $\varepsilon_{it}$ might contain an individual time-invariant effect, which is eliminated by first-differencing as in the majority of the estimators used.

**Table 2.** Labour 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 | — | $\text{labinc}_{it-1}$ | $\Delta\text{labinc}_{it-1}$ | $\ln h_{it-2}$, $\Delta\text{labinc}_{it-1}$ | $\ln h_{it-2}, \ldots, \ln h_{it-11}$, $\text{labinc}_{it-1}$ | $\ln h_{it-2}, \ldots, \ln h_{it-11}$, $\Delta\ln h_{it-2}, \ldots, \Delta\ln h_{it-11}$, $\text{labinc}_{it-1}$ |
| 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 |

*Source:* Authors' calculations.
*Notes:* Columns 1–3: Estimation of an immediate adjustment labour supply equation.
Columns 4–6: Estimation of eq. (6) using different estimators.
We use the sample of the dynamic specifications for all estimations.
Robust standard errors clustered at the individual level in parentheses.
*/**/***: Significance at the 10%/5%/1% level.

Column 2 shows results for the pooled 2SLS estimator, where net wage is instrumented with lagged net labour income to overcome the denominator bias.[16] 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.[17] 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. 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 elasticity of $SR_{\eta_w} = 0.16$ and a long-run elasticity of $LR_{\eta_w} = 0.20$. The coefficient of wage risk implies that an increase in wage risk by one standard deviation leads to an increase in hours of work by 2.4% in the short run. 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 over-identification test does not indicate any invalidity in the instruments.

## 4.2 Results by occupations

We expect heterogeneous results across occupational groups regarding the importance of wage risk, especially concerning the 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 3 provides separate results for different occupational groups using the system GMM estimator with the same instruments as in Table 2.[18] As before, the risk measures are normalized by one standard deviation; however, this time not by the overall, but the subsample 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 labour supply

16 We estimate it using the ivreg2 package (Baum *et al.*, 2016).

17 We estimate them using the xtabond2 package (Roodman, 2009).

18 Results obtained using gross wages instead of net wages appear in Table C.2 in the Online Appendix.

**Table 3.** System GMM labour 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 |

*Source:* Authors' calculations.
*Notes:* Estimation of eq. (6) using the SYS-GMM as in column 6, Table 2.
Robust standard errors clustered at the individual level in parentheses.
*/**/***: Significance at the 10%/5%/1% level.

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 in annual hours of 3.66%.[19] The coefficient on the lag of paid hours worked is not statistically significant for the self-employed and civil servants, which makes intuitive 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 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.

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 instruments may be invalid only for blue-collar workers.

Table 4 shows system GMM estimates of the dynamic labour 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, i.e., 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 labour supply is rejected for managers, professionals, technicians, craftsmen, and operatives. An increase in the probability of unemployment corresponds to an increase of hours worked particularly for managers, craftsmen, operatives, and the unskilled.

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

**Table 4.** System GMM labour 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 |

*Source:* Authors' calculations.
*Notes:* Estimation of eq. (6) using the SYS-GMM as in column 6, Table 2.
Robust standard errors clustered at the individual level in parentheses.
*/**/***: Significance at the 10%/5%/1% level.


**Fig. 5.** Reduction in hours of work
*Source:* Authors' calculations.
*Note:* Small circles indicate the percentile rank of individual *i* in the actually 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 $\sigma_{w,it}^{\min}$. Triangles denote the respective long-run hours of work when $\sigma_{w,it}$ is set to $\sigma_{w,it}^{\min}$.

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.

## 5. Importance of precautionary labour supply

With our estimates of the wage risk semi-elasticity, we can quantify the importance of precautionary labour supply in a *ceteris paribus* exercise, similarly to Carroll and Samwick (1998) for precautionary savings.[20] We use the estimates from Table 2 to simulate the resulting distribution of hours if all individuals faced the same small wage risk. We construct this simulated counterfactual $\widehat{h}_{it}$ from the predictions of the dynamic labour supply equation with minimum sample wage risk $\sigma_{w,it}^{\min}$. 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 labour supply and, for the short run, is calculated as

$$\widehat{h}_{SR,it} - h_{it} = -\beta_3 \left( \sigma_{w,it} - \sigma_{w,it}^{\min} \right). \quad (7)$$

Figure 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

20 Precautionary labour supply is likely even more important for singles because spousal labour supply is an additional channel of insurance against wage risk analogous to the added worker effect (Lundberg, 1985) that is not available for singles. However, applying our analysis to singles is difficult because only a small number of individuals in the SOEP are singles over long periods.

**Table 5.** 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 |

*Source:* Authors' calculations.
*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.

observed distribution of hours of work (vertical axis), and $h_i$ indicates the actual hours of work (horizontal axis). The second point $\left(p_i, \widehat{h}_{SR,i}\right)$ keeps the percentile ranking $p_i$ from the observed distribution and indicates the simulated *short-run* value of the hours of work $\widehat{h}_{SR,i}$ when $\sigma_{w,it}$ is set to $\sigma_{w,it}^{\min}$. The third point $\left(p_i, \widehat{h}_{LR,i}\right)$ shows, as before, $p_i$ from the observed distribution and indicates the simulated *long-run* value of the hours of work $\widehat{h}_{LR,i}$ when $\sigma_{w,it}$ is set to $\sigma_{w,it}^{\min}$.

$$\widehat{h}_{LR,it} - h_{it} = -\frac{\beta_3}{1-\alpha}\left(\sigma_{w,it} - \sigma_{w,it}^{\min}\right). \tag{8}$$

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 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 5 reports the labour 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 labour 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 labour supply amounts to 1.16 hours per week on average.

If wage risk was reduced instead to the median wage risk of civil servants, labour 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 labour supply reduction would amount to 4.49%. If the wage risk of all civil servants was reduced to its median, civil servants' labour supply would decrease by 0.74%.[21]

## 6. Robustness

We conduct a wide range of robustness tests, which are reported and described in more detail in the Online Appendix. We repeat the system GMM estimation for our main sample using alternative definitions of hours of work (Table B.1). The impact of wage risk is positive and significant for annual hours, weekly hours as well as desired hours. It is insignificant for contractual hours, likely because contractual hours cannot be adjusted as easily.

In Table B.2, we include a forward-looking risk measure, a risk measure using a five-year rolling window, a measure based on undetrended wages, and a measure using only continuous spells[22]. All measures have a positive and significant effect on hours of work. In addition, it would be interesting to separately analyze individuals who receive performance related bonuses. Since such compensations, e.g. in the form of large, infrequent lump sum bonuses, are often uncertain *a priori*, they may cause a substantial part of labour income risk. Unfortunately, such bonuses are indicated for less than 1% of all observations, making a separate analysis infeasible.

Mastrogiacomo and Alessie (2014) find similar magnitudes of precautionary savings in the Netherlands when using objective or subjective income risk measures. The SOEP does not include subjective expectations that allow us to construct a risk measure, but rather indicators about worries about the personal financial situation. In an additional robustness test reported in Table B.2, we use these as proxies for income risk, but do not find a significant effect. Nonetheless, the coefficient of the preferred risk measure does not change, when additionally controlling for financial worries.

The last two columns in Table B.2 show results for a measure of household risk as well as a measure of individual risk that also uses information from occupation changes. Again, the wage risk measure is positive and significant in both specifications.

To enable comparison with studies that do not use marginal net wages, we provide a full set of results using gross wages instead of marginal net wages. These are reported in Tables C.1 and C.2. The main results are robust to this.

We are grateful to an anonymous referee for pointing out 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. 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.[23] Therefore, our first strategy is to include these additional control variables in the main model. The results are reported in Table B.3. An increase of one unit on the 1 to 10 Likert scale in the preference for hard work leads to a 1% increase in hours of work. A stronger willingness to take risks—in general or in

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

22 I.e., individuals with periods of unemployment in between employment periods or changes of occupation are excluded.

23 However, information is only available for few time periods. The estimation procedure requires that we impute missing observations. Hence, risk attitudes are partially measured after work choices are made.

occupational matters—leads to a significant, but small, increase in hours of work. Controlling for these variables does not change the coefficients of the variables of main interest.

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 (2013a,b) and construct a region-specific indicator for under- or over-employment (see Online Appendix for more information). The sign of the coefficient, reported in the last two columns of Table B.3, is in line with theoretical predictions. People who are more likely to be underemployed on average work slightly less. However, the magnitude is economically not relevant. The main results are highly robust to inclusion and exclusion of these additional control variables.

In case the full set of controls does not capture all potentially omitted variables that affect selection into jobs, we estimate a Heckman (1979) selection correction model for the four occupations, reported in Table B.4. Again, wage risk remains significant and positive except for civil servants. They are the only group for which selection is significant.

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 B.5. 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 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 B.5. 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 behaviour and even constrains precautionary behaviour (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 re-estimate the main specification by occupations including interactions between year indicators and the wage risk measure. Overall, the estimates of the impact of wage risk, reported in Table B.6, 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 the German manufacturing sector made excessive use of short-time work allowance to cushion the effects of the crisis (Burda and Hunt, 2011).

## 7. Conclusion

We quantify the importance of wage risk to explain the hours of work of married men. The analysis is based on the 2001–2012 waves of the German Socio-Economic Panel. 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 expectations. We observe the largest effects of wage risk for the self-employed.

Precautionary labour supply is economically important. Considering a person who works 42 hours per week, precautionary labour supply amounts to about one week per year, or in monetary terms, about 800 euros per year, with a typical net wage rate of 13 euros.

Precautionary labour supply is particularly important for the self-employed, a group that faces average wage risk substantially above the sample mean. This group works 6.17% 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 because labour supply choices of those who have high unemployment probability are constrained by the transfer system. Our results are based on a partial equilibrium exercise. In future research, one could reconcile our insights with structural estimates of general equilibrium models.

## Supplementary material

The SOEP data are confidential but the replication files are available online on the OUP website, as is the online appendix.

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# Appendix

## 1. Definition of variables

**Table A1.** Definition of key variables

| Variable | Definition |
|---|---|
| Paid hours | Sum of contracted hours (see Table A.1 in the Online Appendix) and paid overtime following Euwals (2005). The SOEP provides information on overtime compensation $or_{it}$ in the sense of 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. |
| Wage risk | In a first step, we regress log gross wage growth on age, its square, education, and interactions of these variables to remove variations due to predictable wage growth. In a second step, we obtain the sample standard deviation of all available past detrended log wages for each person, as in Parker *et al*. (2005). This risk measure uses only the variation across time for each individual. |
| Unemployment risk | Questionnaire asks: 'Are you officially registered as unemployed at the Employment Office ("Arbeitsamt")?' We use this information in 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, and with education, marital status, and unemployment experience. The heteroskedasticity function includes previous unemployment experience and years of education. The general ideal follows Carroll *et al*. (2003). |
| Gross wage | Gross income from work last period divided by hours worked in that period. Example for monthly information on income and weekly information on hours of work: Questionnaire asks: 'What did you earn from your work last month?' State 'Gross income, which means income before deduction of taxes and social security' (extra income such as vacation pay or back pay not included, overtime pay included). Wage is gross income last month divided by the product of the weekly hours measure and 4.33 (the average number of weeks per month). |

(Continued)

**Table A1**. Continued

| Variable | Definition |
|---|---|
| Net wage | We increase each person's annual labour income $y_{it}$ marginally (see eq. (2)). We set $\Delta y_{it} = 2000$ euros, which implies an increase in labour income of about 40 euros per week. We calculate net income *NetInc* using the microsimulation model STSM. Jessen *et al.* (2017) present a comprehensive overview of marginal tax rates for different households (for more information, see Steiner *et al*., 2012). |
| Occupation categorizations used in Figs 1–4 and Tables 1, 3, and 5 | |
| (Questionnaire asks: 'What is your current position/occupation? *Please state the **exact** title in German.*') | |
| Blue collar | SOEP definition of semi-trained and trained worker, foreman, team leader |
| White collar | SOEP definition of qualified and high-qualified professionals, managers |
| Civil servants | SOEP definition of low-level, middle-level, high-level, and executive civil service |
| Self-employed | SOEP definition of liberal professions, other self-employed |
| One-digit international standard classification of occupations used in Table 4 | |
| (See http://www.ilo.org/public/english/bureau/stat/isco/isco88/ for more information.) | |

*Source:* Authors' description.

# How important is precautionary labour supply?

Online Supplementary Material

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## A Definition of additional variables and sample restrictions

Table A.1. Definition of key variables

| Variable | Definition |
|---|---|
| Alternative measures of hours of work used in Table B.1 | |
| Annual Hours | This variable is generated by the SOEP. Annual hours worked in the previous year is calculated by adding together the estimated annual hours of full-time, part-time (including marginal employed), vocational training and short-time work. Annual hours of work in each of these four states is calculated by multiplying the average number of hours worked per week by the number of months worked in each of these three states for the previous year and by 4.33 (the average number of weeks per month). |
| Contracted Hours | Questionnaire asks "How many hours per week are stipulated in your contract (excluding overtime)?" |
| Weekly Hours | Questionnaire asks "And how many hours do you generally work, including any overtime?" |
| Desired Hours | Questionnaire asks "If you could choose your own working hours, taking into account that your income would change according to the number of hours: How many hours would you want to work?" |
| Alternative wage risk measures used in Table B.2 | |
| Forward | In a first step, we regress log gross wage growth on age, its square, education, and interactions of these variables to remove variations due to predictable wage growth. In a second step, we obtain the sample standard deviation of all available *future* detrended log wage realizations for each person. |
| Five Years | First step as above. Second step: We obtain the sample standard deviation of up to *five* past detrended log wage realizations for each person. |
| Undetrended | Sample standard deviation of all available past log wage realizations for each person as observed. |
| Cont. Spells | First step as above. Second: Periods of employment that are not interrupted by periods of unemployment or changes between occupations. In the specification using only continuous employment spells, individuals with periods of unemployment or other occupations in between employment periods are dropped. |
| Subj. Risk | Questionnaire asks "How concerned are you about the following issues?" "Your own economic situation". Possible answers are "Very concerned", "Somewhat concerned", and "Not concerned at all". |
| Household Risk | Analogous to wage risk: First, net household income minus individual net labour income is detrended. Second the standard deviation of past individual realizations is calculated. |

*Continued on next page*

| Variable | Definition |
|---|---|
| BB-Index | The Bell-Blanchflower underemployment index is defined following Bell and Blanchflower (2013a,b) as $$u_{BB} = \frac{U\bar{h} + \sum_k h_k^U - \sum_j h_j^O}{U\bar{h} + \sum_i h_i},$$ where $U$ is the number of unemployed, $\bar{h}$ 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 preferred reduction in hours, which are aggregated over all workers $j$ who desire to work more. $\sum_i h_i$ is the sum of actual hours of work over all workers. |

*Source:* Authors' description.

Table A.2. 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:* Authors' calculations.

## B Robustness of results

Table B.1 shows our preferred specification (System GMM) 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 specify 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. 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 B.1. 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* (6) *using the SYS-GMM as in column 6, Table* 2.
Robust standard errors clustered at the individual level in parentheses.
*/**/***: Significance at the 10%/5%/1% level.
*Source:* Authors' calculations.

Table B.2 shows our preferred specification (System GMM), but with eight different 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). In principle, a measure of wage risk based on information from the whole life span could be calculated. However, such a measure would not provide sufficient variation to identify the coefficient of the risk measure. 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 labour 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 labour 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 behaviour. 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.

Table B.2. 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 (6) using the SYS-GMM as in column 6, Table 2.*
Robust standard errors clustered at the individual level in parentheses.
*/**/***: Significance at the 10%/5%/1% level.
*Source:* Authors' calculations.

We are grateful to an anonymous referee for pointing out 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. 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 B.3. 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 (2013a,b) and construct a region-specific indicator for under- or over-employment. The Bell-Blanchflower underemployment index (BB-index) is defined as

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

where $U$ is the number of unemployed, $\bar{h}$ 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 over-employment, i.e., people in the labour force on average wish to work less hours. As shown in Table 1 the value of the index is 2.7 percent on average for our sample. Column 6 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 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.

Table B.3. 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* (6) *using the SYS-GMM as in column 6, Table* 2.
Robust standard errors clustered at the individual level in parentheses.
*/**/***: Significance at the 10%/5%/1% level.
*Source:* Authors' calculations.

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 excluded variables have strong explanatory power in the first stage. In particular, father's education and occupation are significant at least at the five percent level in all specifications. The results are reported in Table B.4. 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 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 an increase in the probability of unemployment leads to a 3.5%-decrease in hours of work. The reason 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.

Table B.4. 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 labour 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.

*/**/***: Significance at the 10%/5%/1% level.

*Source:* Authors' calculations.

Given that the self-employed and civil servant samples are older on average than the main work force, we repeat the analysis by occupations including only individuals aged at least 35 to make the comparison between occupations easier. The results are reported in columns 4-7 in Table B.5. 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 3. This shows that the differences between occupations are not driven by differences in age.

In the final column of this table, we show results obtained for the main sample, but including transfer recipients. This group is dropped from the main analysis because institutional insurance through the transfer system is likely to play a much higher role than precautionary behaviour and even constrains precautionary behaviour (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.

Table B.5. 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* (6) *using the SYS-GMM as in column 6, Table* 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.

*/**/***: Significance at the 10%/5%/1% level.

*Source:* Authors' calculations.

Table B.6. Time-varying 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* (6) *using the SYS-GMM as in column 6, Table* 2.
Robust standard errors clustered at the individual level in parentheses.
*/**/***: Significance at the 10%/5%/1% level.
*Source:* Authors' calculations.

It is interesting to check if the impact of wage risk differed over the years and in particular during the 2008-09 crisis. Therefore, in Table B.6 we report results for the four occupational group, where the uncertainty measure is multiplied with year dummies. This allows to assess how the impact of wage risk evolved over the years. For the self-employed, white collar and blue collar workers, the effect of wage risk is positive and statistically significant in many years. The effect is never significant for civil servants. When looking at the crisis and its aftermath, i.e., 2008-2010, the effect is particularly strong for the self-employed and white collar workers. In 2010, the coefficient for the self-employed was 0.120. A similar pattern is not observable for blue collar workers. Overall, the estimates of the impact of wage risk are less precise due to less observations for a given year.

## C Results using gross wages

Table C.1. 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 | — | $\text{labinc}_{it-1}$ | $\Delta\text{labinc}_{it-1}$ | $\ln h_{it-2}$, $\Delta\text{labinc}_{it-1}$ | $\ln h_{it-2}, \ldots, \ln h_{it-11}$, $\Delta\text{labinc}_{it-1}$ | $\ln h_{it-2}, \ldots, \ln h_{it-11}$, $\Delta \ln h_{it-2}, \ldots, \Delta \ln h_{it-11}$, $\Delta\text{labinc}_{it-1}$ |
| 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 labour supply equation.
Columns 4-6: Estimation of equation (6) using different estimators.
We use the sample of the dynamic specifications for all estimations.
Robust standard errors clustered at the individual level in parentheses.
*/**/***: Significance at the 10%/5%/1% level.
*Source:* Authors' calculations.

Table C shows the equivalent of Table 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 2.

Similarly, Table C.2 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 C.2. 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 (6) using the SYS-GMM as in column 6, Table 2.
Robust standard errors clustered at the individual level in parentheses.
*/**/***: Significance at the 10%/5%/1% level.
*Source:* Authors' calculations.

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