# Do Women Expect Wage Cuts for Part-Time Work?

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**Abstract**

Wage expectations for full- and part-time employment are key for understanding the labor supply decisions of women. However, whether women expect different wages between part-time and full-time work is not fully understood. Using German survey data, I quantify the expected full-time/part-time wage differential for a representative sample of female workers. I document that women, on average, expect part-time wage penalties of 4-5 percent. Comparing beliefs to selectivity-adjusted estimates of the part-time wage gap indicates that women's expectations are generally realistic. However, variation in beliefs is substantial. Part-time workers and mothers overestimate the wage gains from full-time employment, whereas full-time workers underestimate part-time wage losses.

*Keywords:* expectations, female labor supply, part-time wage gap

*JEL codes:* D84; J22; J31

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## 1. Introduction

Many women will work part-time at some stage of their working career. Although both men and women mostly start out in full-time positions, women, more frequently than men, make adjustments to their labor supply choices. One common pattern is that women sharply reduce working hours after giving birth to their first child, subsequently increase hours as children enter day care or school, and often reduce hours again when elderly relatives are in need of care. Despite this mobility, women remain overrepresented in part-time jobs well beyond the child-rearing years, generating substantial and persistent gender segregation in employment contracts (Petrongolo, 2004), which in turn has significant implications for gender wage gaps (Cortés and Pan, 2019; Goldin, 2014).

One key question that may influence women's labor supply decisions is whether they expect wage changes when moving between full-time and part-time employment. Expectations about the wage impact of working part-time can affect women's willingness to reduce hours. Conversely, beliefs about a full-time wage gain may determine whether part-time working women seek to resume a full-time position. Hence, labor supply choices directly depend on expected wage offers, which may differ between part-time and full-time employment (Blundell et al., 2016; Francesconi, 2002).

Previous studies analyze if women face a wage penalty when they transition between full-time and part-time work, but unless one is willing to impose rational expectations, the expected part-time wage effect of real individuals remains unidentified. If women fail to factor in wage losses where they exist or anticipate wage gains where there are none, they may misjudge the true costs or benefits of switching employment status, which can lead to suboptimal decision making. Moreover, even in the absence of beliefs-biases, understanding the wage expectations of real decision-makers would release us from making unverified assumptions about rationality in standard labor supply models. However, whether women expect different wages in part-time and in full-time employment remains largely unanswered. Stevens et al. (2004), in a survey of British women about various aspects of work-life balance, collected women's beliefs about firm-specific part-time wage gaps. They report that 74% of surveyed women believe their employer pays part-time and full-time working women the same wage. However, Stevens et al. (2004) do not measure women's self-beliefs about their *own* wage change when switching between

full- and part-time work, which is arguably the more relevant criterion in labor supply decisions. Eisenhauer et al. (2021) study women's self-beliefs about future returns to hours worked, hence focusing on the second-order effects of part-time work on wage growth.[2] Breunig et al. (2019) also study the long-run income expectations of households, but do not distinguish between part-time and full-time work. Although beliefs about the long-run effects are important when agents are forward-looking, expectations about the direct or short-run impact of part-time work on wages are of particular interest if agents are either myopic, that is, not forward-looking, or if they are present-biased, thereby prioritizing immediate over future pay-offs. To the best of my knowledge, no existing study analyzes women's expectations about this first-order part-time wage effect. Hence, whether women expect wage cuts for part-time work is not fully understood.

This paper analyzes if women expect wage cuts for part-time work and evaluates if these beliefs are correct. The empirical design consists of two steps. In the first step, I elicit women's expectations about the part-time wage effect, using novel data from the Innovation Sample of the German Socio-Economic Panel (SOEP-IS). The SOEP-IS is a moderately sized survey of German households that is representative of the German population, thus allowing me to study the expectations of employed women across various backgrounds and occupations. To identify expectations about the part-time wage effect, I use a within-subject design, hence comparing the part-time and the full-time scenarios at the individual respondent level. Specifically, I elicit respondents' expected gross income for a hypothetical switch between working full-time at 40 hours and part-time at 20 hours, ceteris paribus, and utilize information on current earnings and agreed working hours to measure the expected part-time wage effect. For robustness, I also check if results differ when working hours are defined to include overtime and when expectations are constructed using probability intervals instead of numerical point estimates.
In the second step, I estimate the wage effect of working part-time econometrically and use these estimates as a benchmark to assess if women's expectations are correct. I build

[2]In the long-term, part-time employment can generate second-order wage effects by causing productivity losses or different rates of promotions that in turn compress wages. Second-order effects can arise as part-time work becomes persistent, creating a 'part-time experience penalty' that goes beyond the direct first-order effect of working part-time in a given period. Unless noted otherwise, this paper refers to the first-order part-time wage effect.

on a vast and active literature that attempts to isolate causal part-time wage effects conditional on worker selection and job segmentation (for a review, see Section 2 as well as Gallego-Granados, 2019).[3] To derive the benchmark part-time wage effect, I estimate Mincerian log wage equations and then use coefficient estimates to isolate conditional part-time wage gaps. I account for the endogeneity of wages and working hours by estimating the wage process conditional on participation and on selection into part-time and full-time employment. First, I use standard 'reduced-form' techniques, including fixed effects regression and ordered-probit Heckman sample selection. Next, I build a static structural discrete choice model of female labor supply and estimate wages jointly with a multinomial logit hours equation, building on Van Soest (1995) and Francesconi (2002). I estimate the model on a sample of German women that is comparable to the SOEP-IS sample, exploiting the longitudinal dimension of the larger German Socio-Economic Panel (SOEP) to model selection. For identification, I use actual transitions between employment states, changes in socio-demographic characteristics, as well as non-linearities in the German tax and benefit system. Throughout all analyses, I use various measures of working hours and part-time status to ensure that results do not depend on a specific hours threshold or on the exclusion of overtime. Using the main specification of the discrete choice model, I compute a selectivity-corrected benchmark estimate of the part-time wage effect and compare this benchmark with the expected part-time wage effect of women constructed in the first step.

I find that women, on average, expect small wage penalties for working part-time. Estimates of the mean expected pay cut for a switch between full-time and part-time employment range between 3.63 and 4.58 percent ($p<0.01$). Moreover, variation in beliefs is substantial. Most women expect near-proportional wages, about 30 percent expect part-time wage cuts exceeding five percent, the remaining 25 percent expect a part-time wage premium. I document that motherhood and current part-time status are significant correlates of heterogeneity in expectations. Mothers and part-time workers expect

[3] Manning and Petrongolo (2008) use reweighting and standard Heckman selection-correction or condition on occupation to eliminate the most important sources of selection bias. Others use simultaneous wage-hours models (Wolf, 2002), random effects with a dynamic multinomial probit component for employment choices (Paul, 2016), or semi-parametric estimations that builds on imputation-based sample selection correction for conditional quantiles (Gallego-Granados, 2019).

notable wage gains from full-time employment, exceeding the population mean in expectations by factor two, approximately. Full-time working women, in contrast, expect no wage loss when switching to part-time. I also find that highly educated women in executive positions, as well as women with basic education, fear stronger wage penalties for part-time work than middle-educated women without managerial responsibilities. Next, I show that women's expectations are realistic on average. Selectivity-corrected estimates of the average part-time wage effect range between 1.13 and 4.74 percent, depending on the specification and, hence, are similar to women's mean expectations. These estimates are also in line with existing literature that finds modest to no significant part-time wage penalties among women, on average (Gallego-Granados, 2019; Paul, 2016; Booth and Wood, 2008; Connolly and Gregory, 2008; Manning and Petrongolo, 2008; Aaronson and French, 2004; Wolf, 2002, , to cite a few). Finally, a comparison of women's beliefs and benchmark estimates within subgroups indicates that mothers and part-time workers overestimate the gains from full-time employment, whereas full-time workers underestimate the negative part-time wage effect. For these subgroups, beliefs are consistent with worker overconfidence: part-time workers believe they would fare considerably better in full-time positions although estimates indicate that they would not; similarly, full-time workers are confident that they wouldn't incur any wage cuts in part-time employment whereas estimates demonstrate that they would. Beyond having important implications for women's labor supply decisions as discussed below, these results clearly show that women's wage expectations are not afflicted by cognitive dissonance - the inclination to report beliefs that justify past choices - which is a commonly voiced concern with subjective data (Bertrand and Mullainathan, 2001), but seems to be less relevant here.

The contribution of this paper is three-fold. First, I show that, despite some exceptions, women have realistic expectations in an important labor market domain. Although the observed wage gaps between part-time and full-time workers are large (10-20 percent), women's expectations about the part-time wage effect are close to the econometrically estimated wage gaps, which are notably smaller (below 5 percent). In times of ever-growing evidence on behavioral anomalies in many domains, it seems that subjective income expectations are less affected by beliefs-biases (Breunig et al., 2019; Zafar, 2011).[4] The risk

[4]Behavioral anomalies can be classified into nonstandard preferences, beliefs, and decision-making (DellaVigna, 2009). Empirical work shows that workers overpredict

of developing beliefs-biases depends on agents' information sets and increases with the complexity of the decision problem. Individuals, on average, have more experience in the domain of labor income, than, for instance, in the financial market domain. Furthermore, expectation formation concerning the short-run is less complex than belief-formation requiring long-run extrapolation. This paper shows that women's expectations about the first-order part-time wage effect are approximately correct.

Second, this paper contributes to the growing literature using stated expectations data to overcome the standard identification problem in revealed preference analysis (Manski, 2004; Zafar, 2011; Wiswall and Zafar, 2016). Specifically, the results presented here provide a behavioral foundation for economists modeling female labor supply. Labor supply models require an assumption about the wage-hours locus and there are competing approaches in the literature. Standard discrete choice models in the tradition of Van Soest (1995) impose constant wage offers for full-time and part-time work, implicitly assuming that decision-makers do not expect a part-time wage effect. Others estimate counterfactual part-time and full-time wage offers, which translates into the implicit assumption that real agents' wage expectations match econometric predictions of the part-time wage effect (Francesconi, 2002; Löffler et al., 2014). So far, the choice between competing models is guided by whether researchers believe a rational agent *should* take into account a part-time wage effect. However, whether agents *actually* expect different wage offers in part-time and in full-time employment is ultimately an empirical question. This paper answers this question for a representative sample of female workers and can serve as a behavioral guideline for researchers striving to capture key features of real agents' decision-processes within economic models of female labor supply.

Third and finally, this paper contributes to our understanding of observed choice patterns in many OECD labor markets, where female part-time shares reach levels of almost 60 percent.[5] On the one hand, the beliefs documented in this paper can rationalize why few

their productivity (Hoffman and Burks, 2017), the unemployed overestimate their job finding probability (Spinnewijn, 2015), and employees overrate the performance of their own company (Oyer and Schaefer, 2005). Agents' systematic mistakes can have implications for welfare analysis (Kőszegi and Rabin, 2007) and behavioral economists have made efforts to augment standard economic models with behavioral elements. However, divergence from standard rational beliefs comes at the expense of model simplicity. Given the appeal of rational expectations for modeling purposes, it is central to uncover under which circumstances the standard model remains appropriate.

[5]The female part-time share reached 57 percent in the Netherlands in 2019, in Germany

women hesitate to reduce hours, for instance to engage in care work. Full-time working women expect no direct wage impact of switching to a part-time position. Hence, they have no wage incentive to prefer full- over part-time employment in a given period. In the short run, this can create an indifference between full- and part-time work, thereby explaining the mobility from full-to part-time employment observed in many female labor markets. On the other hand, women's wage expectations cannot explain why so many women remain in part-time positions over prolonged periods of time. Part-time working women do expect wage gains from full-time work - in fact, they even seem to overestimate these gains. Possibly, expected short-run wage gains are too small to alter labor supply choices, in comparison to other pecuniary and non-pecuniary benefits of part-time employment. In Germany, joint taxation schemes continue to reward, thereby incentivizing, the part-time employment of the secondary earner (Bick and Fuchs-Schündeln, 2017; Steiner and Wrohlich, 2004). Under joint taxation, the expected budget effect of a higher full-time wage rate may be offset by anticipated net losses from tax splitting, which might explain why women remain in part-time employment although they expect wage gains when working full-time. By uncovering women's beliefs about the relative costs of part-time employment, the results from this paper narrow down the set of potential determinants of women's labor supply decisions.

The remainder of this paper proceeds as follows. Section 2 discusses the mechanisms that generate part-time wage gaps and reviews previous empirical findings. Section 4 describes the research design. Section 5 presents the results and section 6 concludes.

and the UK the share is around 36 percent. OECD (2021), Part-time employment rate (indicator). doi: 10.1787/f2ad596c-en (Accessed on 28 January 2021)

## 2. Part-Time Wage Gaps: Mechanisms and Empirical Literature

This section reviews the existing literature on part-time wage effects. There are three broad explanations why part-time workers earn lower wages than full-time workers: First, the jobs they do are different. Second, the workers themselves are different. Third, firms face different costs when employing workers part-time or full-time and compensate by paying different wages to otherwise identical workers in identical jobs. I review these three mechanisms in turn, starting with the last. The section concludes by summarizing empirical estimates of the part-time wage effect from the existing literature.

### *2.1. Firms' Cost Functions*

Firms face fixed costs of labor for recruitment, training, and coordination. One argument proposed by Oi (1962) and later developed by Montgomery (1988) is that these fixed costs pay off less for part-time workers who work fewer hours; hence, firms pay them lower wages to offset relatively higher costs.[6] Goldin (2014) extends this argument, showing that coordination costs can be higher for part-time workers. She finds that returns to both long hours and to particular hours, for instance to being around when colleagues are, increase if coordination and transaction costs are high. Thus, firms reward workers who facilitate smooth workflows and do not cause communication or delegation costs by working irregular hours. Another reason why part-time workers may be less valuable to firms are productivity detriments caused by every-day 'set-up' costs. Set-up costs create rising marginal products of daily working hours until a 'fatigue' threshold, resulting in an S-shaped relationship between hours and productivity, and ultimately, wages (Barzel, 1973; Moffitt, 1984). Others argue that fewer hours may reduce fatigue effects and unproductive 'slack', rationalizing positive part-time premia (Tummers and Woittiez, 1991). Empirically, researchers analyze if firms reward workers with identical characteristics differently in part-time and in full-time employment, which could point to different cost functions for different hours arrangements. However, firms that do face higher costs when employing workers part-time may not always have the wriggle room to adjust hourly wages to match their costs. Worker protection regulation and principles of equal treatment limit their capabilities for cutting wages when their employees reduce

[6] See also Ermisch and Wright (1993) and Wolf (2002) for excellent reviews of the early literature.

hours. Instead, firms may choose to not offer flexible hours for all positions, which explains why part-time and full-time employees often do different jobs.

### *2.2. Job Segmentation*

The segmentation of jobs into those that are exclusively offered on a full-time basis and those that are 'part-time compatible' constitutes a challenge for researchers seeking to identify the true part-time wage penalty. Part-time workers are heavily concentrated in low-wage sectors and in work arrangements with fixed-term contracts, often deemed as marginal or secondary employment (Fernández-Kranz and Rodríguez-Planas, 2011). Wage gains generated by changes in the skill content of work due to technological change also fall exclusively upon full-time workers, spreading wages even further between sectors (Gallego-Granados, 2019; Black and Spitz-Oener, 2010). Notably, the task-level segregation of jobs hurts women more than men and remains a major obstacle toward full 'gender convergence' in pay (Goldin, 2014). Job segmentation also poses an identification challenge for researchers because it generates additional job mobility. If firms do not offer comparable part-time positions, workers seeking to reduce hours may have to change employers or move to a different position within their firm. This involuntary mobility comes with the risk of losing job-specific human capital, or making an occupational downward move; both can in turn compress wages (Manning and Petrongolo, 2008). There is no consensus whether one should include mobility-driven wage losses when estimating the part-time wage effect. An all-encompassing definition of the part-time penalty can include mobility-induced wage cuts, capturing the constraints and costs many workers face when making the transition to part-time in real life. If, instead, the goal is to test some of the theoretical predictions presented above, one needs rich enough data to isolate the part-time wage penalty for identical jobs from wage losses that are due to changes of employer, jobs, or tasks.

### *2.3. Worker Selection*

Finally, part-time workers may earn lower wages because they differ in relevant ways from full-time workers. Some of these differences already exist when people initially select into part-time and full-time jobs, reflecting differences in preferences or constraints. One example is given by Adda et al. (2017), who show that career choices mirror worker differences in desired fertility, with women who have higher preferences for children choosing

family friendly occupations - with the 'amenity' of part-time work - more often.[7] Other differences can arise between previously identical workers, only emerging subsequently, after employment choices have persisted for prolonged periods of time. The most prominent example is the differential accumulation of human capital in part-time and full-time jobs (Blundell et al., 2016). Put differently, it is part-time work itself that makes part-time workers less productive, especially if the choice to work part-time becomes permanent. Work experience appears to be most valuable if it is accumulated in full-time, which can generate second-order effects and create a part-time experience penalty that goes beyond the direct wage effect of working part-time in a given period. In a related argument, Connolly and Gregory (2010) highlight the 'dual nature' of part-time work: a transition to part-time can both harm and support a career, depending on whether the move is permanent, marking the beginning of a "part-time/non-employment cycle," or a temporary "interlude" that maintains labor force attachment and leads back to a full-time career. Last, but not least, institutions and norms also matter. The extent to which part-time and full-time workers differ also depends on the labor force participation of women, in general, and of mothers, in particular. Whether and which women work, in turn, reflects the gender norms of a society and the political environment, such as whether part-time work and non-employment are incentivized through the joint taxation of married couples or if affordable child care reduces the opportunity costs for working mothers (Müller and Wrohlich, 2020). In Germany, the selection of women into part-time has undergone a major transition, going from positive in the early 1990s to null, if not negative, in the 2010s (Gallego-Granados, 2019; Biewen et al., 2018). In the end, regardless if the differences between part-time and full-time workers are inherent, acquired along the way, or induced by institutions - taken together, these constitute the main identification challenge for researchers striving to estimate the part-time effect. A vast literature has emerged around the task of adjusting wage differentials for the non-random distribution of dissim-

[7]The fact that workers accept lower wages for part-time positions is sometimes interpreted within the narrative of *compensating wage differentials*. If workers value the flexibility to reduce working hours, they may accept lower wages for part-time positions in return, especially if part-time jobs are scarce. An alternative reading is that more vulnerable employees face constraints that firms can exploit. Ermisch and Wright (1993) point out that *monopsonistic employers* can exploit the restricted geographic mobility and lower labor supply elasticities of potential part-time workers - often women with young children who face larger spatial constraints - by offering them lower wages.

ilar workers across full-time and part-time jobs. The next section summarizes the main findings from this empirical literature.

### *2.4. Estimated Part-Time Wage Effects in Previous Literature*

Estimates of the part-time wage effect in previous literature differ across countries and over time and depend on the definition of part-time status as well as sample composition, most crucially, whether the sample includes men and workers in marginal part-time contracts (Table 1). I broadly summarize the literature with four main findings: (1) adjusted gaps are substantially lower than raw gaps, in most cases adjustments reduce observed wage gaps by at least half, in some cases they even turn the raw penalty into a part-time premium; (2) adjusted gaps are larger for men than for women, no paper finds part-time premia for men in their main specification; (3) fixed effects estimates yield smaller part-time wage gaps than estimates that do not exclusively use within-individual variation; and (4) the wage exogeneity assumption is problematic, most studies attempt to jointly model wages and hours or correct for selection into part-time. For German females, estimates of the raw wage gap are around 20 percent including marginal part-time and adjusted penalties range from average part-time premia of eight percent (Gallego-Granados, 2019) to penalties of nine percent for short part-time work up to 15 hours per week (Paul, 2016). These findings confirm earlier results for German women by Wolf (2002), who reports lower wages only when working very few hours, but a flat wage-hours locus between 20 - 37 hours. Overall, most empirical studies yield only small first-order part-time wage effects for women, the bulk of observed pay differentials between workers is explained by job segmentation and worker selection into part-time.

Table 1: Part-Time Wage Effects in Previous Literature

| Authors (Year) | Unadjusted | Adjusted | Country/ Years | Data |
|---|---|---|---|---|
| Aaronson and French (2004) | 20-56 | -1-25 | US/ 1968-2000 | PSID, HRS, ORG, CPS |
| Averett and Hotchkiss (1996) | 27-62 | 7-50 | US/ 1989 | CPS |
| Bardasi and Gornick (2008) | -1-22 | -3-18 | Cross Country/ 1994-1995 | LIS |
| Blank (1990) | 19-26 | -17-18 | US/ 1987 | CPS |
| Blundell et al. (2016) | n.r. | 5-8 | UK/ 1991-2008 | BHPS |
| Booth and Wood (2008) | 1-13 | -16 to -10 | Australia/ 2001-2004 | HILDA |
| Connolly and Gregory (2008) | 15-29 | 3-11 | UK/ 1975-2001 | NESPD |
| Ermisch and Wright (1993) | 17 | 8 | UK/ 1980 | WES |
| Fernández-Kranz and Rodríguez-Planas (2011) | 35-38 | 9-14 | Spain/ 1996-2006 | CSWH |
| Fernández-Kranz et al. (2015) | 31-37 | 6-8 | Spain/ 1996-2006 | CSWH |
| Fouarge and Muffels (2009) | n.r. | 0-5 | Cross Country/ 1984-2006 | SEP, BHPS, SOEP |
| Gallego-Granados (2019) | 9-19 | -8-7 | West GER/ 1990-2009 | SOEP |
| Gornick and Jacobs (1996) | 8-28 | -19-23 | Cross Country/ 1986-87 | LIS |
| Hardoy and Schøne (2006) | 5 | 0 | Norway/ 1997-1998 | LLS |
| Hirsch (2005) | 26-49 | -5-18 | US/ 1995-2002 | CPS |
| Manning and Petrongolo (2008) | 25-27 | 3-10 | UK/ 2001-2003 | LFS |
| Matteazzi et al. (2014) | 8-22 | -20-15 | Cross Country/ 2009 | EU-SILC |
| Mumford and Smith (2008) | 12-18 | 0-11 | UK/ 2004 | WERS04 |
| Paul (2016) | 4-32 | -2- 9 | GER/ 1984 - 2011 | SOEP |
| Preston and Yu (2015) | 15-22 | -4-9 | Australia/ 2010 | Australia at Work |
| Simpson (1986) | 31 | 10 | Canada/ 1981 | Survey of Work History |
| Wolf (2002) | n.r. | 0-1 | West GER/ 1995 | SOEP |

*Notes:* Mean or median wage gaps (log difference*100%) rounded to the closest integer value. Negative values denote part-time premia. Some papers include males. All reporting errors are my own. Literature review incomplete, suggestions welcome. The review does not contain results by McGinnity and McManus (2007); Green and Ferber (2005); Montgomery and Cosgrove (1995) due to lack of access. n.r.= not reported.

## 3. Research Design

To measure and assess women's expectations about the part-time wage effect, I follow a two-step approach. In the first step, I estimate women's expectations. For identification, I use a within-subject design. This allows me to compare the part-time *and* the full-time scenarios at the individual respondent level, thereby shutting off an important channel of selection bias because expected wage offers are compared within rather than across individuals (section 3.1). In the second step, I estimate the part-time wage effect econometrically, employing different econometric approaches for robustness, and use these estimates as a benchmark to assess if women's expectations are correct (section 3.2).

### *3.1. Expectations about the Part-Time Wage Effect*

To measure expectations about the part-time wage effect, I utilize a within-subject design that provides information on both the full-time *and* the part-time scenario for the same respondent. Specifically, respondents are asked to consider a hypothetical switch of employment status, based on the following survey instrument if the respondent is employed full-time:[8]

*'Please imagine you were to switch to a part-time job from now on, working 20 hours per week. Please only consider part-time jobs that you could carry-out with your current level of qualification. Which monthly gross income do you expect to earn when working part-time at 20 hours per week?'*

Part-time working individuals are asked a corresponding question that enquires about their expected income when switching to a full-time position at 40 hours per week. I construct each respondent's expected part-time wage effect as follows:

Let $n$ denote the individual respondent who is in one of two states of the world $j$, working either full-time (FT) or part-time (PT), $j_n \in \{FT, PT\}$. Based on $n's$ current working hours, $h_n$, her current monthly gross labor income, $Y_n$ and her reported expected monthly gross labor income for a hypothetical switch between full-time and part-time, $\tilde{Y}_n$, her expected part-time wage effect $\tilde{p}_n$ is given by,

[8]The survey instrument is based on a filter question that splits according to self-reported employment status.

$$\tilde{p}_n = (lw_{FT,n} - lw_{PT,n}) \times 100\% \tag{1}$$

where $lw_{j,n}$ is the natural logarithm of current or expected gross hourly wage, obtained by converting weekly hours to monthly terms,

$$lw_{j,n} = \begin{cases} ln\left(\frac{Y_n}{h_n \cdot \frac{52}{12}}\right) & \text{if } j_n = j \\ \\ ln\left(\frac{\tilde{Y}_n}{h_j \cdot \frac{52}{12}}\right) & \text{if } j_n \neq j \end{cases} \tag{2}$$

with $h_j \in (40, 20)$, as specified in the question on expected income.[9] The sample mean across all individual respondents yields the expected part-time wage effect, $\tilde{p}$.

To keep the survey instrument simple, the question does not specify whether working hours include overtime. Arguably, the more natural reading of the question is in terms of agreed contractual hours, therefore $\tilde{p}$ defines current working hours $h_n$ as agreed hours in the main specification. However, to allow for the possibility that individuals read the question differently and also account for overtime, I specify a second measure of the expected part-time wage effect that uses actual hours, including overtime, for additional robustness.

The question fixes qualification requirements, encouraging respondents to consider the impact of reducing hours in a *ceteris paribus* scenario. This should limit the extent to which people simultaneously account for occupational downgrading, upgrading or job switching. However, the survey instrument is not explicit in this regard and hence one cannot rule out that some individuals do make these kinds of adjustments. Therefore, the derived measure of the expected part-time wage effect potentially encompasses expected mobility-induced wage losses. Consequently, the estimates in this paper can be

[9] In wave 2019 of the SOEP Innovation Sample, the survey instrument is phrased slightly differently and elicits expected hourly wages directly. Hence, the translation from monthly income to hourly wage is only applied for survey waves 2016-2018. To facilitate belief elicitation in wave 2019, individuals are given an individual-specific estimate of their current hourly wage before they report expected counterfactual wages. Elicitation of hourly wages reduces the variability in responses by facilitating the calculation of counterfactual wages, but the change in the wording of the question does not significantly shift point estimates, hence, I pool all survey waves in the empirical analysis.

interpreted as an upper bound of women's expected static wage penalties.[10]

In the main specification, expectations are identified directly from reported Euro amounts. As a sensitivity check, I also construct estimates of the subjective mean probabilistically, using the percent chance respondents expect to earn less than 80 percent and more than 120 percent of their individual-specific point estimate (Appendix C.1).

### *3.2. Estimation of the Part-Time Wage Effect*

To compute the benchmark part-time wage effect, I estimate Mincerian log wage equations and then use coefficient estimates to isolate conditional part-time wage gaps. First, I use 'reduced-form' techniques including fixed effects regression and ordered-probit Heckman sample selection (section 3.2.1), then I build a structural discrete choice model of female labor supply to estimate wages jointly with a multinomial logit hours equation, which allows me to account for the endogeneity of wages and working hours (section 3.2.2).

#### *3.2.1. Reduced-form*

Individual $n$ at time $t$ is observed in one out of three mutually exclusive choice categories $j_n \in \{1, 2, 3\}$, where the choice set $j$ is defined as $1 = FT$ (full-time), $2 = PT$ (part-time), and $3 = OLF$ (out of labor force). Let $w_n$ denote $n's$ wage if $n$ is working at time $t$.[11]

*Combined Wage Equation*

A single combined log wage equation for employed women, $j_n \in \{FT, PT\}$, is given by

$$ln(w_n) = \alpha + \beta \cdot \mathbb{1}(j_n = PT) + Z_n'\gamma + \mu_n + \epsilon_n \tag{3}$$

where $\mathbb{1}(.)$ indicates current employment status $j_n$ and $\beta$ is the parameter of interest, measured conditional on a vector of exogenous covariates, $Z_n$, and an individual-specific fixed effect, $\mu_n$. The vector $Z_n$ collects basic controls for years of education, a quadratic in part-time and full-time work experience (in years), as well as binary indicators for region

[10] In this paper, I take a static perspective and focus on the first-order effect of part-time work on wages. The SOEP-IS survey also elicits individuals' income expectations for the scenario that the hypothetical switch of employment status persists over time, thereby collecting beliefs about dynamic part-time penalties (Eisenhauer et al., 2021).

[11] Time subscript $t$ is dropped for better readability.

(East/West) and immigrant background; if specified broadly $Z_n$ additionally contains occupation major group (1-digit ISCO-88), industry (2-digit NACE), linear and quadratic tenure, as well as binary indicators for firm size ($> 200$), public sector and fixed term contract. In the combined wage equation, the parameter vector $\gamma$, constant $\alpha$ and random shocks $\epsilon_n$ are constrained to be constant across $j$.

*Separate Wage Equations*

Next, parameters and disturbances may vary over $j_n \in \{FT, PT\}$,

$$ln(w_{jn}) = \alpha_j + Z_n'\gamma_j + \mu_{jn} + \epsilon_{jn} \tag{4}$$

yielding separate wage processes for part-time and full-time workers as in standard Blinder-Oaxaca decomposition analyses (Blinder, 1973; Oaxaca, 1973; Jann, 2008). The part-time wage effect can no longer be measured through a constant $\beta$, instead counterfactual full-time wages must be predicted for part-time workers and vice versa, holding fixed individual-specific characteristics (endowments). A part-time wage penalty then unfolds if parameters vary across $j$ such that the 'returns' to identical characteristics differ between part-time and full-time; for instance if the returns to work experience or to having a permanent contract differ across employment states.

*Ordered-Probit Selection Rule*

Estimating the wage equations based on employed women alone can lead to inconsistent coefficient estimates and hence biased part-time wage effects (Heckman, 1979). Non-random choices of part-time and full-time employment generate additional sample selection (Ermisch and Wright, 1993; Paci et al., 1995). To account for the endogeneity of wages and working hours, consider a generalized two-step Heckman estimator as described in Chiburis and Lokshin (2007) with an ordered-probit selection rule,

$$j_n^* = W_n'\delta + u_n \tag{5}$$

where the latent categorical selection variable $j_n^*$ is unobserved, $W_n$ is a vector of individual-specific instruments of observed employment status $j_n \in \{FT, PT, OLF\}$ containing marital status, weekly labor income of woman $n's$ spouse, as well as presence and number of

children. Then augmenting the wage equation by an estimate of the omitted regressor $\lambda_n \equiv E[u_n|j_n, W_n]$ obtained from a first-step ordered probit of $j_n$ on $W_n$ yields consistent estimates of 4, given $W$ satisfies the exclusion restriction $Cov(W_n, \epsilon_{jn}) = 0 \; \forall j_n$.

*3.2.2. Structural Discrete Choice Model*

This section develops a static discrete choice model of female labor supply to model the selection into part-time, full-time, and non-employment more comprehensively. The model incorporates incentives embodied in the German tax and welfare regime that affect employment choices through a classic consumption-leisure trade-off. Separate wage functions for part-time and full-time work are estimated jointly with individual employment choices to relax the exogeneity assumption between wages and hours (Van Soest, 1995; Francesconi, 2002; Löffler et al., 2014). Identification builds on individual transitions between employment states, changes in socio-demographic characteristics, and non-linearities in the tax-benefit system. The model is estimated by full information maximum simulated likelihood (FIMSL). Given the structural parameter estimates, I predict the wage effect of part-time work conditional on selection into employment and into part-time and full-time jobs.

*Model and Estimation*

Let $n$ again denote the individual decision maker. Each decision-maker is assumed to maximize the instantaneous utility index $U(C_{nj}, L_j)$ over the choice set $j_n \in \{FT, PT, OLF\}$ according to a consumption-leisure trade-off, where leisure in choice j is given by $L_j = T - h_j$ with weekly time endowment defined as $T = (24-8)\cdot 5 = 80$ hours and weekly working hours derived from sample medians in the respective choice category. In the main specification based on self-reported part-time status and contractually agreed working hours, sample medians are $h_j \in \{39, 23, 0\}$ s.t. $L_j \in \{41, 57, 80\}$.

The budget constraint is given by

$$C_{nj} = w_{nj} \cdot h_j + s_n - T(w_{nj} \cdot h_j, s_n, k_n) \tag{6}$$

where $C_{nj}$ is consumption, $w_{nj}$ denotes hourly wage, $s_n$ is weekly labor income of the spouse and $T(.)$ is the tax-benefit system that also depends on the number of children, $k_n$. I assume an isoelastic shape of the utility index with constant relative risk aversion

(CRRA),

$$U(C_{nj}, h_j) = \left[ \frac{C_{nj}^{1-\gamma}}{1-\gamma} - (\sum_k \alpha_k' x_{n,k}) \cdot h_j \right] + \epsilon_{nj} \tag{7}$$

where $\gamma$ denotes the coefficient of risk aversion and $\alpha_k$ measures disutility of hours worked, which may vary across decision-makers according to observable taste-shifters $x_{n,k}$ that include binary indicators for the presence of young children in the household and for living in former East Germany. The additive random component $\epsilon_{nj}$ is assumed to be Type I extreme value distributed and independent over n and j (McFadden et al., 1973).

Following Francesconi (2002) and in accordance with the reduced-form approach, I specify separate wage functions for part-time and full-time work to let hourly wages vary over hours categories $j$. Wages are normalized to zero in the non-employment state, $w_{n,OLF} = 0$. Counterfactual wage offers $w_{FT,n}$ and $w_{PT,n}$ are imputed according to full-time and part-time sector specific log-wage equations,

$$ln(w_{jn}) = Z_n' \gamma_j + \zeta_{nj} \tag{8}$$

where the vector $Z_n$ contains a constant and the same set of basic controls as in equation 3, $\gamma_j$ is the corresponding parameter vector. Unobserved sector-specific factors $\zeta_{nj}$ are assumed to be normal and independent of $Z_n$ and across $j$.

I derive net income by simulating income tax payments and transfers for each choice category based on the tax-benefit function $T(.)$, which incorporates detailed features of the 2005 German tax and welfare regime (see Appendix B.1 for specifics). In particular, $T(.)$ covers income tax formulas, social security contributions, solidarity surcharge tax, professional and deductible expenses, child benefits and unemployment transfers, and accounts for the joint taxation of married couples.

Labor supply and wage equations are estimated simultaneously by full information maximum simulated likelihood (FIMSL). Following Van Soest (1995), I integrate out the wage equation prediction errors $\zeta$, which may, if ignored, lead to inconsistent estimates.[12]

[12]Two-step procedures such as Heckman's two-step selection correction that impute $w_n$ by $exp\,(Z_n'\gamma)$ implicitly assume that wages are predicted without error (Van Soest, 1995).

Since the model contains separate wage equations for full-time and part-time sectors, two types of prediction errors must be integrated out (Francesconi, 2002). The resulting log likelihood function is given by equation (11) in Appendix B.2 and is estimated by approximating the expectation terms via Clenshaw-Curtis quadrature. I present measures of internal goodness of fit and the full set of estimation results in Appendices C.4 and C.3. When discussing results, I focus on the predicted part-time wage effect.

*3.2.3. Predicted Part-Time Wage Effect*

Given the 'reduced-form' and structural coefficient estimates from the respective wage functions, I predict full-time and part-time wages $\hat{w}_{FT,n}$ and $\hat{w}_{PT,n}$ for each decision-maker $n$. Analogous to equation (1), these predictions yield $n's$ individual part-time wage effect $\hat{p}_n$,

$$\hat{p}_n = (l\hat{w}_{FT,n} - l\hat{w}_{PT,n}) \times 100\% \tag{9}$$

and by averaging over all employed decision-makers, I obtain the part-time wage effect $\hat{p}$ used as a 'benchmark' to compare with women's average expectations $\tilde{p}$.

## 4. Data

This section presents the two data sets used throughout the empirical analyses, the SOEP and the SOEP Innovation Sample (SOEP-IS). Section 4.1 describes key features of the data, section 4.2 defines the subject pool used for estimation. Appendix A.2 provides additional details on survey administration.

### *4.1. The SOEP and the SOEP Innovation Sample*

The SOEP and the SOEP Innovation Sample (SOEP-IS) are two related longitudinal surveys of private households in Germany, which together constitute the German Socio-Economic Panel.[13] Both the SOEP and the SOEP-IS are annual panel surveys and representative of the German population. Sample design and field processes of both studies are almost identical; in addition, the two surveys share a sizeable part of the questionnaire, using identical wording to facilitate comparability. The SOEP and SOEP-IS differ in longitudinal depth, sample size, and focus; this paper draws on both to combine the strengths of the two data sets. The larger SOEP was launched in 1984 and comprises the life-time trajectories of approximately 15,000 households over more than 30 years. It is widely used by the general scientific community and complements employment and earnings paths with a rich set of household characteristics and socio-demographic indicators. The smaller SOEP Innovation Sample was established in 2011 as a supplement to the SOEP. The SOEP-IS facilitates particularly innovative research by inviting users to submit their own proposals for questions, which must pass a competitive review process before being included on the survey. The 2016 wave of the SOEP-IS introduced an innovative module on income expectations. I utilize survey waves 2016-2019 of this module to estimate expectations about the part-time wage effect for a representative sample of employed women.[14] In addition, I draw on waves 2005-2016 of the larger SOEP study, which outperforms the SOEP-IS in longitudinal depth and sample size. Specifically, I exploit the rich panel dimension of the full SOEP to model the endogeneity of part-time status when estimating the part-time wage effect. Table 2 presents summary statistics

[13] For a jump-start on using the SOEP, see Haisken-DeNew and Frick (2005). For information about the SOEP-IS, see Richter and Schupp (2012); Richter et al. (2015).

[14] The SOEP-IS contains about 4,000 households in total. To ensure a reasonable length of the questionnaire but still accomodate the proposals of as many researchers as possible, the proposed innovative modules are assigned to different subsamples. As a result, each innovative module features only a subset of the full SOEP-IS sample.

for both SOEP and SOEP-IS samples to show the two data sets are comparable and representative of the same population.

*4.2. Sample Definition*

The estimation sample contains all employed women excluding self-employed individuals, pensioners, civil servants, those in training, in the military or in community service. Income expectations in the SOEP-IS were only surveyed for women who are working, but non-working women are included to model labor force participation when estimating the part-time wage effect using the full SOEP. Observations in the top and bottom one percent of the factual and expected income distributions are dropped. Further, I exclude workers in marginal part-time contracts and use self-reported part-time status and contractually agreed working hours excluding overtime in the main specification.[15] I present additional results for hours-based definitions of part-time work and weekly hours including overtime for robustness. The subject pool contains N=959 observations in the SOEP-IS and an unbalanced panel of N=4,707 women in the full SOEP.

Table 2: Composition of SOEP and SOEP-IS

| | SOEP | SOEP-IS | p-value ($\Delta$) |
|---|---|---|---|
| Gross hourly wage (in euros) | 16.27 | 16.94 | 0.18 |
| Agreed wkly hrs. | 31.80 | 30.99 | 0.16 |
| Overtime hrs. per week | 2.64 | 2.77 | 0.59 |
| Education (in years) | 12.69 | 12.66 | 0.84 |
| Age (in years) | 45.10 | 43.21 | 0.02 |
| With children (in percent) | 29.23 | 32.81 | 0.27 |
| Eastern Germany (in percent) | 18.87 | 16.85 | 0.38 |
| Native born (in percent) | 75.36 | 72.68 | 0.44 |
| Public sector (in percent) | 27.09 | 26.10 | 0.70 |
| Firm size $>$ 200 (in percent) | 49.40 | 53.56 | 0.19 |
| Fixed term contract (in percent) | 12.00 | 11.94 | 0.98 |
| Tenure (in years) | 10.84 | 8.94 | 0.01 |
| Manager (in percent) | 4.49 | 3.69 | 0.48 |
| N | 4,707 | 959 | |

*Notes:* Sample averages with population weights. Sample excludes women who are not working, self-employed, in marginal employment, civil servants, pensioners, in training, in community service or in the military. SOEP (2016) and SOEP-IS (2016-19).

[15]Marginal employment in Germany describes contracts with a maximum monthly income of 450 Euros, which are exempted from income tax obligations. These contracts are also known as "Mini-Jobs" or "geringfügige Beschäftigung."

## 5. Results

This section starts by presenting women's expectations (section 5.1) and then provides econometric estimates of the part-time wage effect (section 5.2). Section 5.3 compares women's expectations with estimated part-time penalties to assess if expectations are correct.

### *5.1. Expectations about the Part-Time Wage Effect*

Women, on average, expect small wage penalties for working part-time (Figure 1). Estimates of the mean expected pay cut for a switch between full-time and part-time work range between 3.63 and 4.58 percent ($p<0.01$).

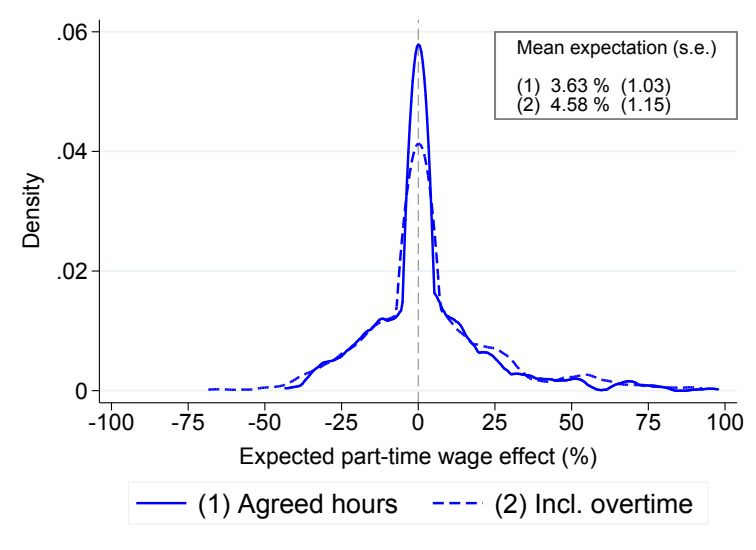



Figure 1: Plot shows the distribution of the expected full-time/part-time wage differential in percent for two definitions of working hours, contractually agreed hours (1, solid line, N=661) and hours including overtime (2, dashed line, N=667). Box shows sample mean with standard errors (s.e.) in parentheses. SOEP-IS (2016-19).

Variation in beliefs is substantial. Most women - about 45 percent - expect only small wage differentials between full- and part-time work of five percent or less (Figure 2). The median woman expects a part-time wage effect of zero. However, 30 percent of women in the sample expect notable part-time wage penalties exceeding 5 percent. At the same time, a significant fraction of women anticipates higher wages in part-time than in full-time: 25 percent of women expect a part-time wage premium of five percent or more. In Appendix C.1, I further show that the distribution of beliefs is similar if expectations are constructed based on probability intervals rather than from direct point estimates.

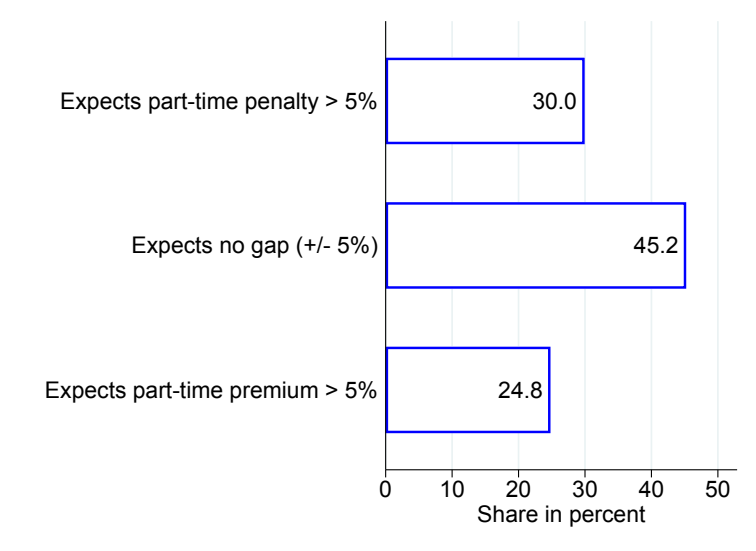



Figure 2: Plot shows the proportion of individuals expecting a part-time wage penalty above 5 percent, the share expecting no sizeable part-time wage effect, and the share expecting a part-time wage premium above 5 percent. Based on contractually agreed working hours, N=661. SOEP-IS (2016-19).

Figures 1 and 2 reveal substantial dispersion in expectations about the part-time effect. To understand this heterogeneity, I disaggregate mean expectations by worker attributes (Tables 3, 4). Descriptive analyses reveal that motherhood and current part-time employment are significant correlates of heterogeneity in expectations. Women with children expect nearly twice as large part-time wage effects than the average woman (6.88 percent, $p<0.05$), almost 40 percent of mothers anticipate a wage cut for part-time work of more than 5 percent. Among women without children, only 25 percent expect a part-time wage cut, expected wage penalties of non-mothers are below average and only marginally significant (2.05 percent, $p<0.1$). A clear asymmetry also arises with respect to current part-time employment status: part-time workers expect to gain more from switching to full-time than full-time workers expect to lose when switching to part-time. Only 23 percent of full-time working women fear a part-time wage loss - in contrast, 36 percent of part-time workers anticipate a full-time wage gain.

To isolate conditional correlations, I further run multivariate OLS regressions of the expected part-time wage effect on joint characteristics, as well as an ordered probit of expecting a part-time premium, no gap or a part-time penalty on joint covariates (Table 5). The correlation with motherhood is robust to the inclusion of worker characteristics in three out of four multivariate specifications. Given that motherhood and part-time

Table 3: Expected Part-Time Wage Effect ($\tilde{p}$) by Socio-Demographic Characteristics

| | Mean (S.E.) | Expects penalty | Expects no gap | Expects premium |
|---|---|---|---|---|
| All women | 3.63 (1.03) | 30.0 | 45.2 | 24.8 |
| Education: Basic | 9.73 (4.12) | 41.0 | 33.3 | 25.6 |
| Education: Intermediate | 1.71 (1.05) | 28.4 | 44.4 | 27.2 |
| Education: Tertiary | 5.65 (2.42) | 28.5 | 53.6 | 17.9 |
| With children | 6.88 (1.74) | 38.9 | 43.3 | 17.8 |
| Without children | 2.05 (1.24) | 25.4 | 46.3 | 28.3 |
| Age < 40y. | 4.08 (1.70) | 32.5 | 40.2 | 27.3 |
| Age > 40y. | 3.36 (1.29) | 28.4 | 48.3 | 23.3 |
| Married | 5.03 (1.38) | 34.6 | 39.1 | 26.3 |
| Single | 0.73 (1.61) | 21.8 | 51.1 | 27.1 |
| Eastern Germany | 1.37 (2.09) | 23.3 | 48.1 | 28.7 |
| Western Germany | 4.18 (1.17) | 31.6 | 44.5 | 23.9 |
| Native born | 3.97 (1.16) | 28.8 | 47.1 | 24.0 |
| Foreign born | 1.26 (2.29) | 31.9 | 41.2 | 26.9 |

*Notes:* SOEP-IS 2016-19. All values in percent. Sample mean of the expected part-time wage effect $\tilde{p_n}$ and the fraction of individuals expecting a part-time wage penalty above 5%, the share expecting no wage effect (+/- 5%) and the proportion expecting a part-time premium above 5%. Results based on self-reported part-time status and contractually agreed working hours. Standard errors (S.E.) clustered at the person-level.

Table 4: Expected Part-Time Wage Effect ($\tilde{p}$) by Job Characteristics

| | Mean (S.E.) | Expects penalty | Expects no gap | Expects premium |
|---|---|---|---|---|
| All women | 3.63 (1.03) | 30.0 | 45.2 | 24.8 |
| Part-time | 6.70 (1.56) | 36.4 | 41.5 | 22.1 |
| Full-time | 0.21 (1.27) | 22.8 | 49.4 | 27.9 |
| Overtime hrs. > 0 | 3.01 (1.28) | 28.9 | 45.5 | 25.6 |
| No overtime hrs. | 4.02 (1.52) | 31.0 | 44.8 | 24.2 |
| Manager | 13.19 (6.42) | 45.2 | 35.5 | 19.4 |
| No manager | 3.16 (1.01) | 29.2 | 45.7 | 25.1 |
| Public sector | 2.21 (1.57) | 23.3 | 52.4 | 24.3 |
| Private sector | 4.25 (1.25) | 32.9 | 42.0 | 25.1 |
| Firm size > 200 | 2.51 (1.27) | 27.9 | 46.8 | 25.2 |
| Firm size < 200 | 5.13 (1.61) | 32.4 | 43.8 | 23.8 |
| Fixed term contract | 9.62 (4.17) | 39.7 | 35.6 | 24.7 |
| Permanent contract | 2.87 (1.03) | 28.5 | 46.5 | 25.0 |
| Tenure > 10y. | 2.00 (1.37) | 23.3 | 54.5 | 22.2 |
| Tenure < 10y. | 4.19 (1.44) | 34.1 | 40.0 | 25.9 |

*Notes:* SOEP-IS 2016-19. All values in percent. Sample mean of the expected part-time wage effect $\tilde{p_n}$ and the fraction of individuals expecting a part-time wage penalty above 5%, the share expecting no wage effect (+/- 5%) and the proportion expecting a part-time premium above 5%. Results based on self-reported part-time status and contractually agreed working hours. Standard errors (S.E.) clustered at the person-level.

employment are two choices that women tend to make coincidentally, the joint inclusion of both variables can compromise precise estimation of the individual determinants. A Wald test on the joint significance of motherhood and part-time status indicates that the two variables are jointly significant in OLS and in ordered probit models that use the full set of covariates (p<0.01). Mothers and part-time working women expect substantially higher wage penalties for part-time work than non-mothers and full-time workers (+3.71 and + 4.42 p.p., respectively).

Table 5: Multivariate OLS and Ordered Probit Estimates of the Expected Part-Time Wage Effect ($\tilde{p}$)

| | (1) OLS | | (2) OLS | | (3) Ordered Probit | | (4) Ordered Probit | |
|---|---|---|---|---|---|---|---|---|
| | Coef. | Std.Err. | Coef. | Std.Err. | Coef. | Std.Err. | Coef. | Std.Err. |
| Education: Basic | 8.64* | (4.59) | 7.69* | (4.36) | 0.21 | (0.21) | 0.16 | (0.21) |
| Education: Tertiary | 4.60 | (2.80) | 3.62 | (2.67) | 0.18 | (0.13) | 0.17 | (0.13) |
| With children | 5.45** | (2.29) | 3.71 | (2.68) | 0.37*** | (0.12) | 0.31** | (0.14) |
| Age > 40y. | 2.65 | (2.15) | 4.08* | (2.27) | 0.16 | (0.12) | 0.21* | (0.13) |
| Married | 4.02* | (2.17) | 2.61 | (2.00) | 0.21* | (0.11) | 0.17 | (0.12) |
| Eastern Germany | −2.93 | (2.62) | −2.16 | (2.66) | −0.24* | (0.14) | −0.22 | (0.14) |
| Native born | 4.06 | (2.84) | 3.99 | (2.83) | 0.09 | (0.15) | 0.10 | (0.16) |
| Part-time | | | 4.42* | (2.36) | | | 0.18 | (0.12) |
| Overtime hrs. > 0 | | | 0.05 | (1.93) | | | −0.04 | (0.11) |
| Manager | | | 12.67* | (7.05) | | | 0.34 | (0.29) |
| Public sector | | | −1.10 | (1.79) | | | −0.09 | (0.11) |
| Firm size > 200 | | | −2.77 | (2.11) | | | −0.05 | (0.11) |
| Fixed term contract | | | 4.22 | (4.24) | | | 0.09 | (0.21) |
| Tenure > 10y. | | | −1.74 | (2.33) | | | −0.10 | (0.12) |
| Constant | yes | | yes | | yes | | yes | |
| Survey year FE | yes | | yes | | yes | | yes | |
| Mean dep. variable | 3.00 | | 3.00 | | | | | |
| $N$ | 535 | | 535 | | 535 | | 535 | |

*Notes:* SOEP-IS 2016-19. Estimates from multivariate OLS regressions of the expected part-time wage effect $\tilde{p_n}$ in percent (Models 1,2) and from multivariate ordered probit regressions where the ordinal dependent variable indicates if $\tilde{p_n}$ is below -5%, between -5 to 5% or above 5% (Models 3,4). Education base category: intermediate education. Based on self-reported part-time status and contractually agreed working hours. Standard errors clustered at the person-level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Multivariate analyses also reveal that women in managerial positions expect notably larger wage effects of part-time work than non-managers (+12.67 p.p., p<0.1). Further, women with basic education anticipate stronger wage penalties than women with intermediate education (+7.69 p.p., p<0.1). Thus, the fear to be penalized in part-time employment seems to concern women at the top as well as those at the bottom of the wage distribution.

*5.2. Estimates of the Part-Time Wage Effect*

Next, I present econometric estimates of the part-time wage effect. Unadjusted part-time wage differentials for women in Germany reach up to 21 percent if samples include marginal part-time workers, but vary between seven and ten percent once marginal part-time is excluded (Table 6, row 1). However, these raw gaps provide no suitable estimates of the causal effect of part-time work on wages due to job segmentation and worker selection (see section 2). Based on reduced-form and structural discrete choice estimation, I derive adjusted estimates of the part-time wage effect that can be compared with women's expectations (Table 6). Column 1 in Table 6 shows the main specification with self-reported part-time status and agreed hours.

Table 6: Estimates of the Part-Time Wage Effect ($\hat{p}$)

| | (1) | | (2) | | (3) | | (4) | |
|---|---|---|---|---|---|---|---|---|
| Part-time wage effect ($\hat{p}$) | Coef. | Std.Err. | Coef. | Std.Err. | Coef. | Std.Err. | Coef. | Std.Err. |
| 1. Unadjusted (raw gap) | 9.67 | (0.80) | 10.17 | (0.78) | 8.85 | (0.79) | 7.17 | (0.77) |
| (A) Combined wage function | | | | | | | | |
| 2. OLS, basic controls | 4.12 | (0.73) | 4.74 | (0.73) | 2.32 | (0.73) | 1.02 | (0.70) |
| 3. OLS, broad controls | 0.31 | (0.57) | 1.15 | (0.58) | −2.75 | (0.57) | −3.70 | (0.56) |
| 4. Fixed effects, basic controls | −3.06 | (0.56) | −1.24 | (0.58) | −8.99 | (0.64) | −10.52 | (0.59) |
| 5. Fixed effects, broad controls | −3.31 | (0.55) | −1.49 | (0.56) | −9.50 | (0.63) | −11.01 | (0.58) |
| (B) Separate FT/PT wage functions | | | | | | | | |
| 6. OLS, basic controls | 3.90 | (0.13) | 4.54 | (0.13) | 2.30 | (0.11) | 2.04 | (0.11) |
| 7. OLS, broad controls | 0.32 | (0.11) | 1.22 | (0.10) | −2.48 | (0.11) | −3.09 | (0.09) |
| 8. Fixed effects, basic controls | 8.12 | (0.21) | 8.97 | (0.20) | 4.74 | (0.17) | 5.62 | (0.16) |
| 9. Fixed effects, broad controls | 5.78 | (0.22) | 6.88 | (0.21) | 4.10 | (0.19) | 4.42 | (0.18) |
| 10. Ordered probit Heckman sample selection (basic controls) | −2.63 | (0.12) | −2.40 | (0.12) | −3.75 | (0.10) | −5.35 | (0.11) |
| (C) Structural estimates | | | | | | | | |
| 11. Discrete choice model | 4.74 | (0.63) | 1.13 | (0.66) | 4.65 | (0.62) | 2.51 | (0.66) |
| *Part-time status* | self-reported | | self-reported | | hours < 30 | | hours < 30 | |
| *Working hours* | agreed hrs. | | incl. overtime | | agreed hrs. | | incl. overtime | |

*Notes:* Table presents estimates of the part-time wage effect $\hat{p}$ in percent. Panel (A) gives the coefficient estimate of a binary part-time indicator from a combined log wage regression on full- and part-time workers. Panels (B) and (C) give the mean difference in full-and part-time log wage predictions obtained after separate full-and part-time log wage regressions. All wage regressions are based on SOEP waves 2005-2016, with a minimum sample size of N=48,603. Predictions reported in Panels (B) and (C) are restricted to employed women sampled in 2016 to obtain part-time wage effects for a sample that is comparable to the SOEP-IS. Basic controls include years of education, linear and quadratic work experience in part-time and full-time, region (Eastern/ Western GER) and immigrant background. Broad controls add occupation major group (ISCO 88, 1 digit), industry (NACE,2 digit), linear and quadratic tenure and indicators for firm size > 200, public sector and fixed term contract. Selection equation for the multinomial choice to work full-time, part-time or not at all (row 10) contains marital status, partner income and presence and number of children. Structural estimates from multinomial logit models with separate wage functions for part- and full-time.

Adjusting the raw wage gap for differences in basic characteristics between part-time and full-time workers by OLS reduces the pay gap to roughly four percent (Table 6, rows 2 and 6). Compositional differences in education and work experience between workers in different contract types explain almost half of the observed wage gap. With a broad set of controls that adjust for job segmentation, including occupational segregation and

differences in sector and tenure, OLS estimation no longer yields significant differences in pay (rows 3 and 7). However, a causal interpretation of these results hinges on the assumption that self-selection into part-time and full-time contracts is independent of unobservables, otherwise estimates may be biased. One solution to this identification problem is to estimate wage regressions with individual fixed effects that control for time-invariant unobservables. The underlying assumption is that, for a given woman, changes in unobserved characteristics do not systematically determine mobility between part-time and full-time contracts. If mobility from full- to part-time contracts is induced by concurrent unobserved productivity shocks, such as deteriorating health, a within-individual comparison based on switchers alone could lead to upward biased results of the part-time penalty. Similarly, if a switch from part-time to full-time involves a reward for women who are becoming more productive at their jobs, and if this happens systematically, estimates of the part-time wage effect may also be upward biased. However, wage equations estimated with individual fixed effects that rely on individual transitions between part-time and full-time employment for identification yield small wage premiums for working part-time of about three percent (rows 4-5). These effects are hence notably smaller than estimates from between-individual comparisons. Thus, the aforementioned systematic differences between unobservables and job mobility are unlikely, even though this cannot be tested directly. One potential explanation for the positive part-time wage premium estimated by fixed effects are heterogeneous effects of working part-time. If switchers constitute a selective sample of employees, under heterogenous effects of working part-time fixed effects estimates that exclusively rely on within-variation of employment status might not be equal to the average part-time penalty in the population of interest. A descriptive comparison of switchers with non-switchers indicates that switchers are representative of the population in terms of education levels but differ in terms of other observable characteristics, such as the number of children they have (Table APP.4). To the extent that switchers represent a positive selection of women who succeeded in negotiating more flexible working hours or who juggle work and child care obligations rather than drop out of the labor force altogether, fixed effects estimates may be downward-biased, thereby yielding positive part-time premiums. Notably, fixed effects regression of separate wage functions for part-time and full-time workers, where identification does not rely on switchers alone, yields a part-time wage penalty of 5.8

percent (row 9). However, these estimates again rely on the assumption that part-time status is not determined by unobservables. To control for selection into the different contracts, I then adopt a Heckman-type ordered-probit selection rule, using marital status, partner income, as well as the presence and number of children as instruments for contract type (row 10). Arguably, the assumptions on the instruments are rather restrictive. Therefore, I estimate part-time wage effects inside a multinomial logit discrete choice model that allows for a more comprehensive mapping of the labor supply decision (row 11). Again, estimates are provided for different definitions of hourly wages and part-time status for robustness. Structurally estimated part-time wage effects are similar across different specifications, albeit slightly smaller if hours include overtime. Estimates of the part-time wage effect range from 1.1 to 4.7 percent and, hence, are comparable in size to reduced-form estimates.

Taken together, although point estimates vary somewhat across specifications, I confirm the findings from the previous literature: among women, the wage penalty for working part-time is positive but modest in size. My main specification based on the discrete choice model yields a mean part-time wage effect of 4.74 percent, which is substantially lower than observable wage gaps between full- and part-time workers. Most of the observable raw wage gap between workers is explained by differences in worker characteristics, job segmentation, and the non-random selection of dissimilar workers into full-time and part-time jobs. This is especially true if workers in marginal employment are excluded.

*5.3. Comparison of Expected and Estimated Part-Time Wage Effects*

This section takes women's expectations and compares them to estimated wage effects from the discrete choice model.

Overall, expectations are well aligned with estimated part-time wage effects (Figure 3). On average, the estimated wage effect of moving between full- and part-time is below 5 percent; moreover, women expect small wage penalties between 4 and 5 percent. In the main specification, estimated wage effects are within the confidence bands around women's expected wage changes, suggesting women's expectations about the wage-hours locus are, on average, realistic. Discrete choice estimates based on hours including overtime are somewhat smaller (below 2 percent), thereby providing a lower bound of the benchmark wage effect. Hence, if anything, women overestimate rather than underesti-

mate the part-time effect.

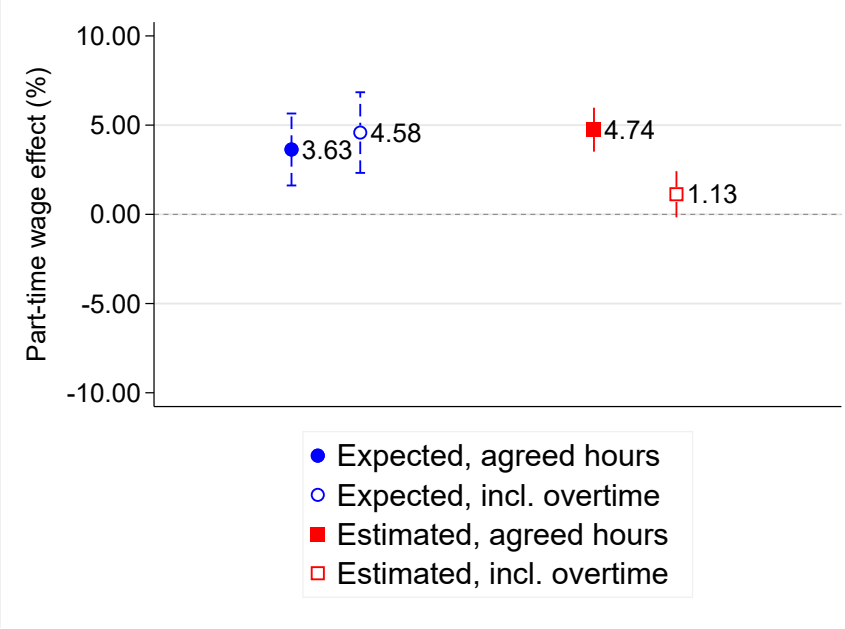
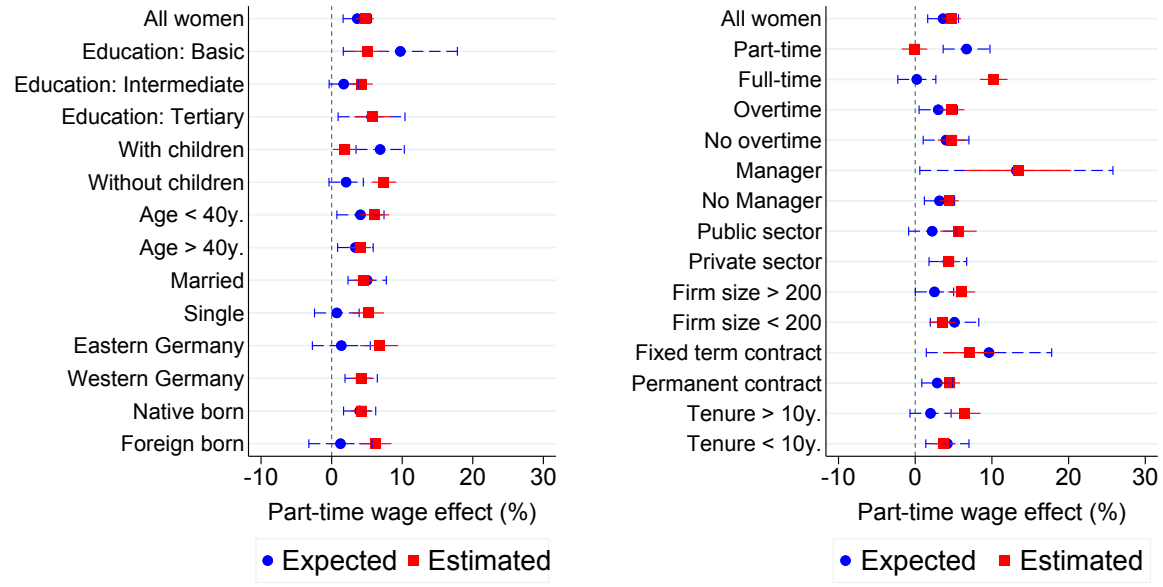



Figure 3: Comparison of expected (blue dashed line) and estimated (red solid line) part-time wage effect. Based on self-reported part-time status, estimates from discrete choice models. Markers indicate mean wage gaps and whiskers 95% C.I.s. SOEP-IS (2016-19) and SOEP (2016).

Next, I compare expected and estimated part-time wage effects within subgroups based on splits by socio-demographic and job characteristics (Figure 4). I illustrate results graphically here and report the corresponding point estimates in Table APP.7 in the Appendix.

One striking dimension of effect heterogeneity is *current employment status*: part-time workers expect wage gains in full-time employment, whereas full-time workers expect proportional wages. Moreover, estimates of the part-time wage effect reveal the reverse: part-time workers are unlikely to gain from switching to full-time, but full-time workers are predicted to face wage losses in part-time. On average, part-time workers have accumulated more part-time work experience, which is not rewarded in full-time positions (see Table APP.5 and also compare Blundell et al. (2016)). For an average part-time worker, the benefits of moving to a full-time job are not immediate but will only materialize once she accumulates additional work experience in full-time. For full-time workers, returns to full-time experience are also lower in part-time, generating wage losses in part-time employment. Interestingly, for these subgroups, beliefs are consistent with worker overconfidence: part-time workers believe they would fare considerably better in full-time




(a) Socio-Demographics

(b) Job Characteristics

Figure 4: Comparison of expected (blue dashed line) and estimated (red solid line) part-time wage effect. Based on agreed hours and self-reported part-time status. Estimates from the discrete choice model. Markers indicate point estimates and whiskers 95% C.I.s. SOEP-IS (2016-19) and SOEP (2016).

positions although estimates indicate that they would not; similarly, full-time workers are confident that they wouldn't incur wage cuts in part-time employment whereas estimates demonstrate that they would. One often voiced concern with subjective data is cognitive dissonance, the propensity to report beliefs that justify past choices (Bertrand and Mullainathan, 2001). Here, the wage expectations of part-time and full-time working women are clearly contradictory to cognitive dissonance: if part-time workers preferred to report beliefs that are consistent with their choice to work part-time, they should not disclose they expect higher wages in full-time employment. Conversely, full-time workers should report strong part-time wage penalties if they wanted to legitimize their choice to work full-time. In contrast to cognitive dissonance, both full-time and part-time working women report beliefs that rationalize changes of employment status, which has important implications for understanding women's labor supply decisions (see the discussion in Section 6).

As noted previously, *motherhood* also significantly correlates with the probability to expect wage cuts for part-time work. Indeed, mothers face significant part-time wage cuts. However, mothers tend to overestimate the size of the part-time wage penalty: on average, they expect a wage penalty of seven percent; in contrast, estimates indicate part-time wage losses of below two percent, on average. The difference between expectations and

predictions is only significant at the 10-percent level due to subsample size limitations in the SOEP-IS. With this caveat in mind, one cautious interpretation of the finding is that mothers may overestimate the part-time wage effect because they internalize the second-order consequences of part-time employment, such as reduced career progression. Women without children, in contrast, appear to underestimate the part-time wage penalty. One potential explanation could be stigma: perhaps, it is socially less accepted if a woman without children opts for a part-time contract and, if she does, she may signal reduced willingness to perform.

*Education* is another interesting source of heterogeneity: women with a basic education and those with a tertiary degree expect significant part-time wage penalties, whereas women with medium education expect near-constant wages. These beliefs are quite realistic: both women with low and with high education levels face larger part-time wage effects than middle-educated women, although differences are imprecisely estimated. The results are also in line with Gallego-Granados (2019), who establishes sizeable part-time wage penalties for women at the bottom of the wage distribution, those who tend to have lower levels of education. I show these women also expect larger wage cuts.

Another group of women who expects realistically large part-time wage effects are *managers*. Women with managerial responsibilities expect 13 percent lower wages in part-time, mirroring estimates of the part-time wage effect for women with managerial tasks. Given the small number of women with executive positions in the SOEP samples, standard errors are considerable, but point estimates are well aligned and strictly positive. Hence, women in executive positions face notable wage cuts when they transition from full- to part-time and they also expect a part-time wage effect. This relates to the job-mobility argument by Manning and Petrongolo (2008) and others: women in positions that are not deemed 'part-time compatible' face an occupational downward move when switching from full- to part-time and, in turn, undergo cuts in hourly pay. Here I show that women who are particularly likely to suffer mobility-induced wage losses also fear these wage cuts most.

Point estimates also show that *public sector employees* and individuals in *larger firms* do not fear part-time work as much as employees in the private sector and in smaller firms, but differences are insignificant and estimates indicate no differential remuneration of part-time work across these groups. Women with *fixed-term contracts* expect

larger part-time wage effects than women with permanent contracts. Point estimates, in line with previous work by Fernández-Kranz and Rodríguez-Planas (2011), confirm that women in these types of secondary work arrangements indeed face larger part-time wage penalties. I do not find an important role for *overtime* hours. Although full-time workers report significantly more overtime in absolute terms - 1.1 hours more per week, on average ($p<0.05$) - when computed in relation to current working hours, part-timers work relatively more overtime, namely roughly one percentage point more ($p<0.05$), rationalizing why part-time penalties effectively do not differ between these groups. Women with higher levels of *firm-specific human capital* (tenure above 10 years) are predicted to lose more by moving to part-time, although they do not expect larger penalties.

Finally, if the sample is split by *region* into former East and West German states, one point worth noting is that although true part-time wage effects do not differ significantly, East German women view part-time work somewhat more favorably, suggesting a role for institutions and norms in shaping the beliefs of individuals. Women in Eastern Germany are significantly more likely to work full-time and they tend to have shorter career breaks when they have children. Easterners do not view part-time work as a career disruption, which may reflect that part-time work is usually a temporary solution. In contrast, a move from full-to part-time is often permanent in Western Germany, partly because child care infrastructure remains less developed. These differences exemplify the 'dual nature' of part-time work discussed by Connolly and Gregory (2010), who argue that spells of part-time employment can be both 'a support' and 'a trap to a successful career,' depending on whether these spells are short 'interludes,' as is the more common pattern in Eastern Germany, or whether part-time work becomes the norm in alternation with periods of non-employment, as is often the case in Western Germany.

## 6. Discussion

This paper documents that women expect small part-time wage penalties of 4 to 5 percent on average, in line with selectivity-corrected estimates of the part-time wage effect. Moreover, dispersion in beliefs is substantial. The majority of women expect proportional wages, but mothers and managers expect sizeable part-time wage cuts. Furthermore, part-time workers anticipate full-time wage gains, but full-time workers expect no part-time wage losses.

These findings shed light on important labor supply decisions that women face over the life-cycle. The beliefs documented here can rationalize why few women hesitate to reduce hours; for instance, to engage in care work. Full-time working women who expect no direct wage impact of switching between full- and part-time employment have no wage incentive to prefer full- over part-time work in a given period. In the short run, this can create an indifference between full- and part-time contracts, thereby explaining the mobility from full- to part-time employment observed in female labor markets across OECD countries.

On the one hand, these results are reassuring from a policy perspective, because many women value part-time work (Wippermann, 2018) and spells of part-time employment can support a woman's career by maintaining labor force attachment (Connolly and Gregory, 2010). However, as part-time work becomes persistent, it can harm workers in the long-run by creating second-order effects. The reasons are manifold and include the lack of transferability of work experience from part- to full-time employment, causing steeper wage growth in full-time (Francesconi, 2002; Blundell et al., 2016), or different rates of promotions in jobs where 'being around when others are' is rewarded disproportionately, ultimately driving up gender wage gaps (Goldin, 2014). Not least, even with similar wage rates, part-time workers accumulate lower pension entitlements across the life-cycle, contributing to gender pension gaps (Frericks et al., 2009; Hammerschmid and Rowold, 2019). Taken together, whilst short interludes of part-time work may be innocuous, a growing consensus is emerging that persistent part-time employment should be avoided.

One interesting insight from this paper is that women's wage expectations seem to be at odds with the persistent part-time employment of women observed in Germany and

elsewhere. Part-time working women expect wage gains from full-time work - in fact they even seem to overestimate these gains - which should create a disincentive to remain in part-time employment over prolonged periods of time. Possibly, expected gross wage gains for full-time work are too small to alter labor supply choices, in comparison to other pecuniary and non-pecuniary benefits of part-time employment. In Germany, joint taxation schemes continue to reward, thereby incentivizing, the part-time employment of the secondary earner (Bick and Fuchs-Schündeln, 2017; Steiner and Wrohlich, 2004). This paper studies women's beliefs about individual gross wages, but women's labor supply decisions are partly based on total net income of the household. Under joint taxation, the expected budget effect of a higher full-time wage rate may be offset by anticipated net losses from tax splitting, which might explain why women remain in part-time employment even though they expect wage gains when working full-time. Additionally, beliefs about the long-term consequences of part-time work might be biased. One limitation of this paper is that it does not incorporate women's expectations regarding the second-order effects of part-time work, which provides another mechanism explaining persistent part-time employment (Eisenhauer et al., 2021). Another explanation is involuntary part-time employment, caused by inflexible work contracts (Manning and Petrongolo, 2008) or by the lack of formal care arrangements (Müller and Wrohlich, 2020).

By uncovering women's beliefs about the relative costs of part-time employment, the results from this paper provide new insights that narrow down the set of potential determinants of women's labor supply decisions.

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# Appendix

## A Data: Further Details

### A.1 Expectations Survey

Below is a description of the survey instrument used to elicit expectations of full-time workers about counterfactual earnings in part-time. Part-time workers are asked to consider the opposite scenario of switching to a full-time position of 40 hours per week.

**Please imagine you were to switch to a part-time job from now on working 20 hours per week. Please only consider part-time jobs that you could carry out with your qualification.**

(a) Which monthly gross income do you expect to earn when working part-time at 20 hours per week?

(b) How likely do you think it is that a part-time position at 20 hours per week yields a gross income of less than X-20% per month?*

(c) How likely do you think it is that a part-time position at 20 hours per week yields a gross income of more than X+20% per month?*

**Please report your answer in percent. 0% means you consider it impossible, 100% means that you are certain. You can use the percent values in between to graduate your answer.* [Note: X is the individual-specific response to (a)]

### A.2 Survey Administration

Sample design and field work of the SOEP and the SOEP Innovation Sample (SOEP-IS) are almost identical. For both surveys, participating households were initially selected through multi-stage random sampling with regional clustering. Face-to-face interviews take place once a year and last approximately 1.5-2 hours. Participants receive small gifts upon completion of each interview, as well as small cash incentives. Households either receive 5 Euros per completed personal interview and 10 Euros per household interview, or they receive a lottery ticket for the charitable TV lottery "Ein Platz an der Sonne" (A place in the sun). Administration of both surveys lies with the German Institute for Economic Research, DIW Berlin, but Kantar Public (formerly TNS Infratest) is responsible for the field work, including software programming, interviewer recruitment, interviewer training, and coordination of interviews.

# B Model: Further Details

## B.1 Tax and Welfare Regime

The model implements details of the 2005 German tax and benefit system based on features of the German Tax and Benefit Microsimulation Model (STSM) described in Steiner et al. (2012) to simulate net income for each employment choice, following three steps: First, I subtract professional and deductible expenses to derive taxable income. Second, I calculate income tax liability by applying tax formulas depending on marital status. Finally, I deduct liabilities from gross income and add transfers to obtain net income.

To obtain taxable income in step one, gross labor income of the household is converted into real terms (base year 2005) and aggregated to annual amounts. For counterfactual choice categories, I derive alternative-specific gross earnings by multiplying hours times the hourly wage rate that is allowed to vary across full-time and part-time choices. I disregard income components from alternative sources such as capital income or income from renting and leasing.[16] Given gross annual real income, I deduct the lump-sum amount of 920 Euros for professional expenses ("Werbungskosten") for all workers. In addition, actual or lump-sum deductible expenses ("Sonderausgaben") are subtracted up to a maximum amount. I simplify this step and consider only the general flat rate amount of 36 Euros ("Pauschbetrag") as well as expenses for social security contributions.[17] Table APP.1 presents detailed information on how deductible expenses are accounted for. For simplicity, the model does not incorporate loss deductions and extraordinary deductable expenses ("aussergewöhnliche Belastungen"). One further simplification I resume to involves the distinction between child allowances that are deducted before applying the tax function and child benefits ("Kindergeld"), which are added afterwards. A more accurate account of the tax-benefit system would conduct a higher-yield test ("Günstigerprüfung") and assign the more favorable rule (Steiner and Wrohlich, 2008). I abstract from this distinction and assume all couples with children receive child benefits.

[16] Since I exclude pensioners and self-employed women, I also disregard income from pensions or self-employment.

[17] I abstract from other deductible expenses such as insurance contributions, alimony payments, church tax, expenses for training, donations and tax consultancy expenses

Table APP.1: Annual Deductable Expenses for Social Security Contributions in 2005

| | Single individual | Married couple |
|---|---|---|
| Minimum ("Vorsorgepauschale") | $0.2 \cdot RV \ + \ min(0.12 \cdot INC, 1900)$ | $0.2 \cdot RV_{HH} \ + \ min(0.12 \cdot INC_{HH}, 3800)$ |
| Actual expenses | | |
| *Bracket 1 ("Diff. Vorwegabzug")* | $max(0, 3068 - 0.16 \cdot INC)$ | $max(0, 6136 - 0.16 \cdot INC_{HH})$ |
| *Bracket 2* | $min(1334, SV - Bracket\ 1)$ | $min(2668, SV - Bracket\ 1)$ |
| *Bracket 3* | $min(667, SV - Bracket\ 1 - Bracket\ 2)$ | $min(1334, SV - Bracket\ 1 - Bracket\ 2)$ |
| Maximum | 2001 | 4002 |

*Notes:* All amounts in Euros and annual terms. RV= old age pension contributions ("Rentenversicherung"). HH= household level. INC= gross income. SV= total social security contributions ("Sozialversicherung"). Old age (RV) contributions deductable up to a correction factor (20% in 2005)

Given taxable income, I obtain income tax liability of the household in step two. In Germany, due to the joint taxation of married couples ("Ehegattensplitting"), singles and married individuals are taxed differently. For singles, income tax formulas are applied directly to individual taxable income. For married couples, total taxable income of the household is first divided by two. Income tax formulas are then applied to half the amount of total taxable household income. The derived tax liability is then doubled to determine overall tax liability of the couple. Table APP.2 contains income tax formulas as well as minimal and maximal marginal tax rates for all available tax brackets. Income is not taxed below an annual allowance of 7,664 Euros and tax rates evolve according to a partially linear rule until a top income threshold of 52,152 Euros after which income is taxed at a constant marginal rate of 42%.

Table APP.2: Income Tax Formula in 2005 (§ 32 a Abs. 1 EStG)

| Zone | Tax bracket | Tax formula | MTR (min) | MTR (max) |
|---|---|---|---|---|
| 1 | $\leq$ 7664 | t=0 | 0 | 0 |
| 2 | 7665-12739 | t=(883.74Y + 1500)Y | 15% | 23.97% |
| 3 | 12740-52151 | t=(228.74Z + 2397)Z + 989 | 23.97% | 42% |
| 4 | $\geq$ 52152 | t=0.42X - 7914 | 42% | 42% |

*Notes:* Income and tax liabilities refer to annual Euro amounts. MTR = marginal tax rate. Y and Z are 1/10000 of excess income over upper bound of the previous bracket. X is taxable income.

In step three, I compute net income by deducting income tax, social security contributions, and the solidarity surcharge ("Solidaritätszuschlag")[18] from gross income and by adding transfers and benefits. I calculate unemployment benefits according to ALG II standard rates ("Regelbedarfssätze") that differ between East and West German regions and by household composition (Table APP.3). Payments are means-tested and individ-

[18]Solidarity surcharge of 5.5% on tax liability accrues for couples (individuals) owing above 1944 (972) Euros annual tax.

uals are only eligible for unemployment transfers if joint household income, including spousal income, is lower than transfer claims and if household assets are below exempted wealth allowances. I simplify the means-test by assuming households are ineligible for social assistance as soon as one spouse has positive labor income. In accordance with the STSM, I do not model payments from unemployment insurance (ALG I).[19] Child benefits are added once for each couple (the first three children receive 154 Euros each, all additional children receive 179 Euros each). I refrain from covering any additional benefits (e.g. allowances for housing, education, widows etc.).

Table APP.3: Unemployment Benefit Standard Rates in 2005 (SGB II/Hartz IV and SGB XII)

| | Single adults (I) | Adults in couples (II) | Youth 14 - 18 (III) | Children < 14 (IV) |
|---|---|---|---|---|
| East | 331 | 298 | 265 | 199 |
| West | 345 | 311 | 276 | 207 |

*Notes:* Monthly allowances per person in Euros.

### B.2 Simulated Log Likelihood Function

If full-time and part-time wages were observed for all individuals, including non-workers, the log-likelihood function would be given by

$$ln(L(\theta)) = \sum^{N} ln\left(\frac{exp(\beta' x_{ni})}{\sum_j exp(\beta' x_{nj})}\right) + \sum^{N}\left\{ln\ \phi\left(\frac{lnw_n - Z_n'\gamma}{\sigma_w}\right) - ln\ \sigma_w\right\} \qquad (10)$$

where the first summand denotes the likelihood contributions from logit choice probabilities over hours choices and the second term gives the likelihood of the wage equation residuals, assuming log-normality, where $\phi(.)$ is the normal density.

Accounting for unobserved wage offers, two types of prediction errors must be integrated out, resulting in the following simulated log-likelihood function:

[19] Individuals who worked in the previous year are, in principle, entitled to payments from unemployment insurance for the first 6 months after becoming unemployed. These payments are not means-tested and replace 60-67% of previous net income. I follow the STSM and assume all unemployed directly apply for unemployment benefits (ALG II).

$$
\begin{aligned}
ln(SL) = & \sum^{FT} ln \left\{ \frac{1}{R} \sum^{R} P^{(r)}_{n,FT} \right\} + \sum^{FT} \left\{ ln\ \phi \left( \frac{lnw^{FT}_n - Z'_n \gamma^{FT}}{\sigma^{FT}_w} \right) - ln\ \sigma^{FT}_w \right\} \\
& + \sum^{FT} \left\{ ln\ \frac{1}{R} \sum^{R} \phi \left( \frac{lnw^{PT,(r)}_n - Z'_n \gamma^{PT}}{\sigma^{PT}_w} \right) - ln\ \sigma^{PT}_w \right\} \\
& + \sum^{PT} ln \left\{ \frac{1}{R} \sum^{R} P^{(r)}_{n,PT} \right\} + \sum^{PT} \left\{ ln\ \phi \left( \frac{lnw^{PT}_n - Z'_n \gamma^{PT}}{\sigma^{PT}_w} \right) - ln\ \sigma^{PT}_w \right\} \\
& + \sum^{PT} \left\{ ln\ \frac{1}{R} \sum^{R} \phi \left( \frac{lnw^{FT,(r)}_n - Z'_n \gamma^{FT}}{\sigma^{FT}_w} \right) - ln\ \sigma^{FT}_w \right\} \\
& + \sum^{OLF} ln \left\{ \frac{1}{R} \sum^{R} P^{(r)}_{n,OLF} \right\} + \sum^{OLF} \left\{ ln\ \frac{1}{R} \sum^{R} \phi \left( \frac{lnw^{FT,(r)}_n - Z'_n \gamma^{FT}}{\sigma^{FT}_w} \right) - ln\ \sigma^{FT}_w \right\} \\
& + \sum^{OLF} \left\{ ln\ \frac{1}{R} \sum^{R} \phi \left( \frac{lnw^{PT,(r)}_n - Z'_n \gamma^{PT}}{\sigma^{PT}_w} \right) - ln\ \sigma^{PT}_w \right\}
\end{aligned}
\tag{11}
$$

where $P^{(r)}_{n,i} = \frac{exp(\beta' x_{ni})}{\sum_j exp(\beta' x^{(r)}_{nj})}$ denotes the simulated logit choice probability from draw $r \in R$. $w^{FT,(r)}_n$ and $w^{PT,(r)}_n \in x^{(r)}_{nj}$ are simulated full-time and part-time wage offers.

A full information maximum simulated likelihood estimator is given by

$$
\hat{\theta}_{FIMSL} = argmax_\theta\ ln(SL), \qquad \theta = (\beta, \gamma, \sigma^{FT}_w, \sigma^{PT}_w)
$$

# C Additional Results

## C.1 Probabilistic Expectations

Here I show that constructing expected part-time wage effects from probability intervals yields similar point estimates as when asking for numerical Euro amounts directly. I use respondents' subjective probabilities for earning less than 80 percent and more than 120 percent of their numeric point estimate to construct individual-specific probability distributions (Figures APP.1, APP.2). I use non-parametric spline interpolation that allows for flexible approximations to individuals' subjective distributions, which outperforms parametric approximations (Bellemare et al., 2012). From these fitted distributions, I derive alternative measures of central tendency (subjective means and medians). Subjective means and medians also center symmetrically around zero, mirroring point estimates

from direct numerical questions (Figures APP.3, APP.4).

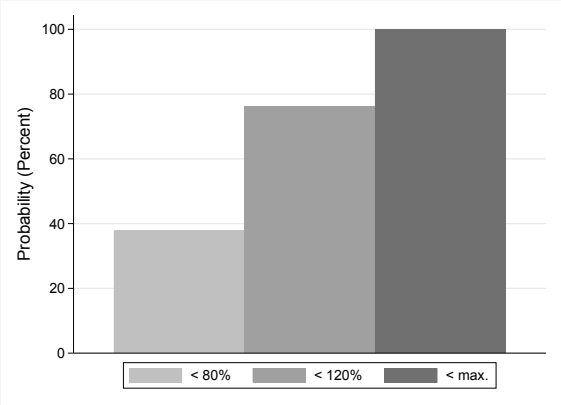



Figure APP.1: Mean discrete subjective cumulative density functions (C.D.F.s) for expected wages based on reported probabilities. SOEP-IS (2016).

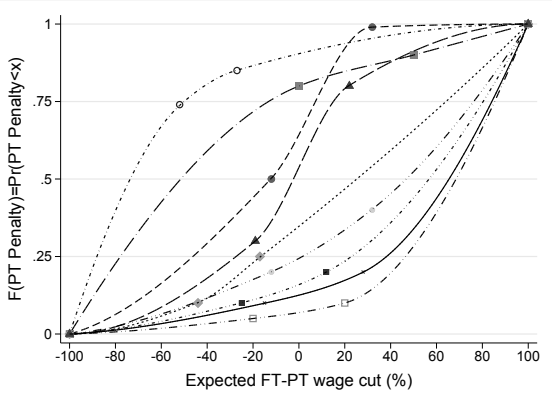



Figure APP.2: Interpolated smooth subjective C.D.F.s for expected part-time wage effect from subjective probabilities and non-parametric piecewise cubic hermite interpolating polynomials for selected individuals. SOEP-IS (2016).

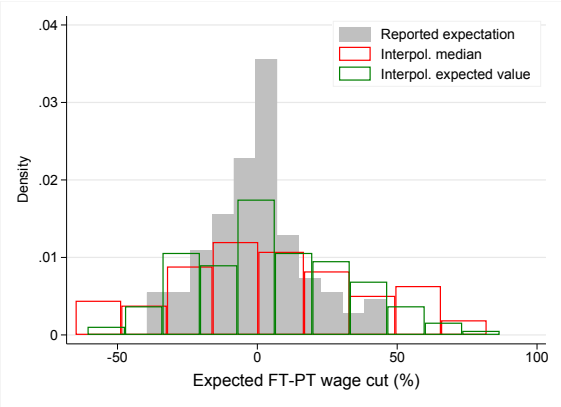



Figure APP.3: Distribution of central tendency, comparing reported point estimates (grey), interpolated medians (red), and interpolated mean (green). Interpolation based on probabilistic questions and non-parametric splines. SOEP-IS (2016).

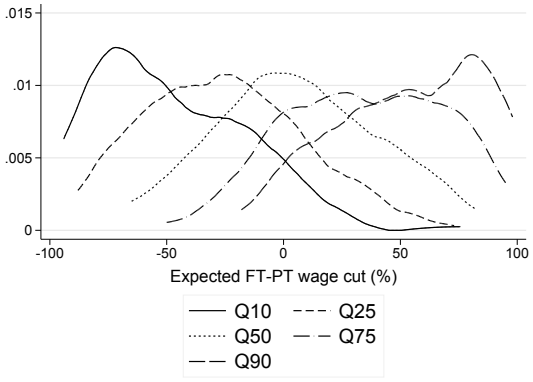



Figure APP.4: Distribution of subjective quantiles, based on probabilistic questions and non-parametric spline interpolation. SOEP-IS (2016).

### C.2 Fixed Effects Identification

Table APP.4 presents summary statistics about the subset of women who switch at least once between part-time and full-time employment during the entire observation period, and compares these women to those who are observed to work either full-time or part-time for the entire observation period. Switchers significantly differ from full-time-only working women and from part-time-only working women in a number of observable characteristics. Most notably, switchers are significantly more likely to have children than full-time-only working women, the share of mothers is highest among part-time-only working women. In terms of productivity differences, switchers have similar education as full-time-only working women and only moderately higher education as part-time-only workers. Fixed effects estimation that uses within-variation in employment status for identification of the part-time wage effect relies on women switching employment status. To the extent that the subset of switchers differs from the population of interest, the results diverge from average treatment effects.

Table APP.4: Composition of switchers, full-time-only workers and part-time-only workers

| | Switcher | Full-time only | Part-time only | Switcher vs. FT *p-value* ($\Delta$) | Switcher vs. PT *p-value* ($\Delta$) |
|---|---|---|---|---|---|
| Gross hourly wage (in euros) | 14.40 | 15.76 | 13.64 | 0.00 | 0.00 |
| Agreed weekly hrs. | 31.88 | 38.77 | 22.42 | 0.00 | 0.00 |
| Overtime hrs. per week | 2.88 | 3.42 | 1.97 | 0.00 | 0.00 |
| Education (in years) | 12.57 | 12.72 | 12.01 | 0.16 | 0.00 |
| Age (in years) | 42.77 | 41.66 | 46.04 | 0.02 | 0.00 |
| With children (in percent) | 0.34 | 0.15 | 0.49 | 0.00 | 0.00 |
| Eastern Germany (in percent) | 0.24 | 0.20 | 0.14 | 0.05 | 0.00 |
| Native born (in percent) | 0.79 | 0.81 | 0.79 | 0.17 | 0.72 |
| Public sector (in percent) | 0.33 | 0.27 | 0.30 | 0.00 | 0.19 |
| Firm size > 200 (in percent) | 0.46 | 0.52 | 0.42 | 0.01 | 0.02 |
| Fixed term contract (in percent) | 0.14 | 0.09 | 0.11 | 0.00 | 0.01 |
| Tenure (in years) | 9.79 | 11.35 | 10.74 | 0.00 | 0.04 |
| Manager (in percent) | 0.01 | 0.02 | 0.01 | 0.02 | 0.01 |
| N | 18,478 | 17,668 | 12,779 | | |

*Notes:* Sample averages with population weights. Sample excludes women who are not working, self-employed, in marginal employment, civil servants, pensioners, in training, in community service or in the military. SOEP (2005-2016).

### C.3 FIMSL Estimation Results

Table APP.5 presents the full set of estimation results of the discrete choice model for different specifications of part-time status and working hours.

Table APP.5: FIMSL Estimation Results of the Discrete Choice Model

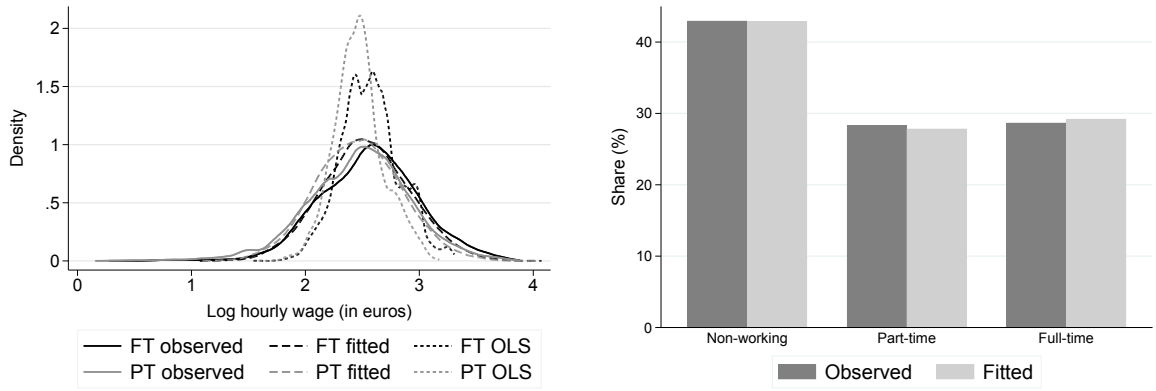
| **PT status: self-reported** | **(1) Agreed hours** | | | | **(2) Incl. overtime** | | | |
|---|---|---|---|---|---|---|---|---|
| | Full-time | | Part-time | | Full-time | | Part-time | |
| *Log wages* | Coef. | Std.Err. | Coef. | Std.Err. | Coef. | Std.Err. | Coef. | Std.Err. |
| Education (years) | 0.091 | 0.001 | 0.090 | 0.001 | 0.091 | 0.001 | 0.087 | 0.001 |
| FT experience (years) | 0.030 | 0.001 | 0.028 | 0.001 | 0.034 | 0.001 | 0.030 | 0.001 |
| FT experience sq. | 0.000 | 0.000 | −0.001 | 0.000 | −0.001 | 0.000 | −0.001 | 0.000 |
| PT experience (years) | 0.000 | 0.001 | 0.018 | 0.001 | −0.003 | 0.001 | 0.016 | 0.001 |
| PT experience sq. | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| East | −0.282 | 0.003 | −0.261 | 0.004 | −0.287 | 0.004 | −0.267 | 0.004 |
| Foreign born | −0.061 | 0.004 | −0.050 | 0.004 | −0.066 | 0.004 | −0.052 | 0.004 |
| Constant | 1.140 | 0.009 | 1.060 | 0.010 | 0.994 | 0.011 | 1.018 | 0.012 |
| Std.Dev. | 0.076 | 0.001 | 0.090 | 0.001 | 0.087 | 0.001 | 0.102 | 0.001 |
| *Hours choice* | | Coef. | Std.Err. | | | Coef. | Std.Err. | |
| Consumption | | 0.124 | 0.003 | | | 0.138 | 0.003 | |
| Hours | | 0.038 | 0.001 | | | 0.036 | 0.001 | |
| Hours × Kids | | 0.041 | 0.001 | | | 0.039 | 0.001 | |
| Hours × East | | −0.018 | 0.001 | | | −0.017 | 0.001 | |
| Log likelihood | | 167219.180 | | | | 186678.8172 | | |
| **PT status: hours-based** | **(3) Agreed hours** | | | | **(4) Incl. overtime** | | | |
| | Full-time | | Part-time | | Full-time | | Part-time | |
| *Log wages* | Coef. | Std.Err. | Coef. | Std.Err. | Coef. | Std.Err. | Coef. | Std.Err. |
| Education (years) | 0.090 | 0.001 | 0.091 | 0.001 | 0.092 | 0.001 | 0.086 | 0.001 |
| FT experience (years) | 0.030 | 0.001 | 0.027 | 0.001 | 0.034 | 0.001 | 0.028 | 0.001 |
| FT experience sq. | 0.000 | 0.000 | −0.001 | 0.000 | −0.001 | 0.000 | −0.001 | 0.000 |
| PT experience (years) | 0.001 | 0.001 | 0.017 | 0.001 | 0.000 | 0.001 | 0.017 | 0.001 |
| PT experience sq. | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| East | −0.280 | 0.003 | −0.261 | 0.004 | −0.279 | 0.004 | −0.277 | 0.005 |
| Foreign born | −0.064 | 0.004 | −0.048 | 0.004 | −0.068 | 0.004 | −0.051 | 0.004 |
| Constant | 1.145 | 0.009 | 1.060 | 0.010 | 0.976 | 0.010 | 1.028 | 0.011 |
| Std.Dev. | 0.077 | 0.001 | 0.090 | 0.001 | 0.093 | 0.001 | 0.098 | 0.001 |
| *Hours choice* | | Coef. | Std.Err. | | | Coef. | Std.Err. | |
| Consumption | | 0.120 | 0.003 | | | 0.149 | 0.003 | |
| Hours | | 0.037 | 0.001 | | | 0.035 | 0.001 | |
| Hours × Kids | | 0.043 | 0.001 | | | 0.041 | 0.001 | |
| Hours × East | | −0.019 | 0.001 | | | −0.021 | 0.001 | |
| Log likelihood | | 167485.728 | | | | 186786.764 | | |

*Notes:* SOEP (2005-2016). Results from full information maximum simulated likelihood (FIMSL) estimation with constant relative risk aversion (CRRA) utility index.

## C.4 Internal Goodness of Fit

I present graphical evidence of model fit (Figure APP.5) and estimated wage elasticities (Table APP.6) for the main specification of the discrete choice model with self-reported part-time status and agreed working hours.

### *C.4.1. Model Fit: Wages and Hours Choices*




(a) Hourly Wages

(b) Hours Choices

Figure APP.5: Goodness of Fit of the Discrete Choice Model

### *C.4.2. Wage Elasticities*

Estimated elasticities for a one percent increase in gross hourly wage for females in the sample are 0.41 percent for working hours and 0.22 percentage points for participation. These elasticities are mostly within the confidence intervals of comparable estimates by Haan (2006), deviations can be explained by differences in sample composition, most notably I include singles whereas Haan (2006) focuses on married couples.

Table APP.6: Labor Supply Elasticities

| | Δ Hours (percent) | | Δ Participation (p.p.) | |
|---|---|---|---|---|
| | Coef. | Std.Err. | Coef. | Std.Err. |
| All women | 0.41 | 0.00 | 0.22 | 0.02 |
| By region | | | | |
| East | 0.34 | 0.00 | 0.25 | 0.04 |
| West | 0.43 | 0.00 | 0.21 | 0.02 |

*Notes: Predicted changes for a 1% increase in gross hourly wage.*

### C.5 Subgroup Comparison of Expected and Predicted Wage Effects

Table APP.7 presents the full set of point estimates corresponding to the graphical evidence in Figure 4 in section 5.3.

Table APP.7: Comparison of expected and estimated wage effects by subgroups

| | Expected | | Estimated | |
|---|---|---|---|---|
| | Mean | S.E. | Mean | S.E. |
| All women | 3.63 | 1.03 | 4.74 | 0.63 |
| Education: Basic | 9.73 | 4.12 | 5.05 | 1.41 |
| Education: Intermediate | 1.71 | 1.05 | 4.19 | 0.83 |
| Education: Tertiary | 5.65 | 2.42 | 5.79 | 1.31 |
| With children | 6.88 | 1.74 | 1.88 | 0.88 |
| Without children | 2.05 | 1.24 | 7.40 | 0.89 |
| Age $<$ 40y. | 4.08 | 1.70 | 6.09 | 1.04 |
| Age $>$ 40y. | 3.36 | 1.29 | 4.02 | 0.79 |
| Married | 5.03 | 1.38 | 4.55 | 0.75 |
| Single | 0.73 | 1.61 | 5.21 | 1.14 |
| Eastern_Germany | 1.37 | 2.09 | 6.73 | 1.38 |
| Western Germany | 4.18 | 1.17 | 4.17 | 0.70 |
| Native born | 3.97 | 1.16 | 4.20 | 0.74 |
| Foreign born | 1.26 | 2.29 | 6.17 | 1.17 |
| Part-time | 6.70 | 1.56 | −0.10 | 0.85 |
| Full-time | 0.21 | 1.27 | 10.23 | 0.92 |
| Overtime | 3.01 | 1.28 | 4.75 | 0.85 |
| No overtime | 4.02 | 1.52 | 4.78 | 0.93 |
| Manager | 13.19 | 6.42 | 13.44 | 3.50 |
| No Manager | 3.16 | 1.01 | 4.44 | 0.64 |
| Public sector | 2.21 | 1.57 | 5.64 | 1.21 |
| Private sector | 4.25 | 1.25 | 4.40 | 0.74 |
| Firm size $>$ 200 | 2.51 | 1.27 | 6.05 | 0.90 |
| Firm size $<$ 200 | 5.13 | 1.61 | 3.53 | 0.88 |
| Fixed term contract | 9.62 | 4.17 | 7.05 | 1.74 |
| Permanent contract | 2.87 | 1.03 | 4.52 | 0.67 |
| Tenure $>$ 10y. | 2.00 | 1.37 | 6.42 | 1.06 |
| Tenure $<$ 10y. | 4.19 | 1.44 | 3.74 | 0.78 |

*Notes:* SOEP-IS (2016-19) and SOEP (2016). Sample means with standard errors (S.E.) of the expected part-time wage effect and the estimated part-time wage effect overal and within subgroups. Results based on self-reported part-time status and contractually agreed working hours. Estimates from the CRRA discrete choice model. Standard errors in SOEP-IS clustered at the person-level.