There And Back Again: Women's Marginal Commuting Costs*

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Abstract: The authors estimate how female and male workers trade off higher wages against a shorter commuting distance in a model of job search. Their analysis relies on an administrative dataset from West Germany, and uses a statistical method that can flexibly account for unobserved differences between individuals as well as for the delays involved in finding one's preferred job. Their findings show that childless women are willing to give up €0.27 per day (0.4% of the daily wage) to reduce their commute by one kilometre. Men have a similar willingness to accept longer commutes in return for higher wages as childless women. However, women's willingness to pay more than doubles after having a child, which helps explain the wage gap between mothers and others. Further analysis using a matched sample of married mixed-sex couples indicates that couples often prefer that wives commute shorter distances than their husbands.

Keywords: Commuting, marginal willingness to pay for job attributes, on-the-job search, Cox relative risk model, partial likelihood estimation, gender and parenthood in job search models, heterogeneity in job mobility, gender wage gap

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There is widespread recognition that commuting is not only a negative job attribute, but also that there are remarkable gender differences in commuting distance, pre-dating the Covid-19 pandemic (Auspurg and Schönholzer 2013). In our data, we find that women's commuting distances in West Germany are on average 18% lower than men's. At the same time, the raw gender gap in daily earnings stands at 33%. Qualitatively similar results are found for the UK (Manning 2003*a*; Petrongolo and Ronchi 2020) and France (Le Barbanchon, Rathelot, and Roulet 2021), but compared to other high-income countries, the (West-)German unadjusted gender and motherhood wage gap (Grimshaw and Rubery 2015) are exceptionally high. The issue is therefore particularly salient in this context.

In this paper, we bring the two stylized facts of large gender gaps in earnings and commuting distances together to study gender differences in the willingness to trade higher wages for a lower commuting distance. Workers who are able and willing to commute further have access to more and potentially higher-paid jobs. Therefore, gender differences in willingness to accept a longer commute in return for a higher wage can contribute to wider disparities in labour market outcomes between men and women, and between mothers and non-mothers.[1] The opportunity cost of commuting time is likely higher for women, and in particular mothers and those with other caregiving responsibilities. Hence, a better understanding of differences in willingness to trade off commuting distances against wages can help evaluate policies designed to narrow gender and motherhood gaps in the labour market. These include policies directly reducing the need to commute, which may be easier to implement post-pandemic, as well as policies that can enable or encourage mothers to commute further.

Focusing on the extreme case of West-Germany, we choose a modelling approach that fits its labour market setting. We select and extend the estimation approach of Gronberg and Reed (1994) and Van Ommeren, Van Den Berg, and Gorter (2000), which sidesteps a key source of bias in traditional (*hedonic*) estimates of willingness to pay for job attributes (Hwang, Mortensen, and Reed 1998) in labour markets where it takes time and effort to find one's preferred job. Estimating willingness to pay from a search process from unemployment requires additional information (see Le Barbanchon, Rathelot, and Roulet 2021) and a different set of assumptions. We use a rich and long panel of employed non–university educated workers in West Germany to estimate female and male workers' willingness to accept lower wages in return for a shorter commuting distance. We choose a statistical method (*stratified partial likelihoods*) that, in combination with our long panel, allows us to flexibly account for

[1] For a general discussion of the gender and motherhood wage gap, see Blau and Kahn (2017) and Petrongolo and Ronchi (2020). For a more specific discussion of the motherhood wage gap, see Adda, Dustmann, and Stevens (2017), Angelov, Johansson, and Lindahl (2016), Cortés and Pan (2023), Kleven, Landais, and Søgaard (2019) and Lundborg, Plug, and Würtz Rasmussen (2024).

unobserved heterogeneity.[2] We condition on a large number of additional job characteristics in addition to wages and distance, including detailed occupation and industry indicators. Furthermore, we extensively model heterogeneity in willingness to pay, allowing it to vary by gender, the wage, and the number and age of children throughout. In additional analyses, we study the role of part-time work, economic geography, housing costs as well as relative wage and commuting position in the household. In this way we do not treat willingness to pay for a job attribute as a single fixed parameter, but one that varies both across individuals and over time.[3]

Only a handful of papers have studied related issues so far. In the most closely related study, Le Barbanchon, Rathelot, and Roulet (2021) recently utilised administrative data on unemployed workers. Building upon elicited preferences of unemployed workers in France and information on accepted job offers in combination with assumptions derived from a job search model, they identify willingness to pay to reduce commuting distance with a focus on gender differences. They find that unemployed women value a shorter commuting distance 20% more than unemployed men and this can explain around 10-14% of the residual wage gap.

We complement their analysis in a number of ways. The French and German contexts differ sharply in terms of gender attitudes. This make it important to account for gender-specific wage offer distributions and job offer arrival rates in the German context. Thus we use a different set of assumptions. Our approach allows for gender-specific wage offer distributions and job offer arrival rates, which comes at the cost of, for example, assumptions regarding layoff probabilities. Furthermore, we do not rely on stated preferences and the assumptions necessary to anchor a set of two stated preferences. Additionally, their main analysis focuses on workers who were made redundant or otherwise became unemployed involuntarily, whereas we focus on non-university educated employed workers. Lastly, we exploit the length of our panel to focus on characterising the heterogeneity of willingness to pay, including by child age, with identification coming from observed job-to-job transitions.

Petrongolo and Ronchi (2020) use data on realised job–to–job transitions in the UK that are observed with some approximation to study the willingness to pay to reduce commuting

[2]Gronberg and Reed (1994); Van Ommeren, Van Den Berg, and Gorter (2000); Dale-Olsen (2006*a*) and Russo, Van Ommeren, and Rietveld (2012) could not take this approach due to data restrictions and Borghorst, Mulalic, and Van Ommeren (2022) do not use a flexible hazard rate model, thus not accounting for duration dependence.

[3]Lundberg (2023) discusses the malleability of preferences as a major way forward in gender economics. Although we estimate willingness to pay as a preference parameter, differences may ultimately be rooted in different constraints, in the spirit of Akerlof and Kranton (2000).

distance. They adjust the job search model taken from an early version of Le Barbanchon, Rathelot, and Roulet (2021) and find a gender gap of 15% in willingness to pay. As the estimated willingness to pay is very low, the difference contributes little to the gender wage gap. However, they argue that their estimate is a lower bound on the willingness to pay. As Petrongolo and Ronchi (2020) uses a version of the paper of Le Barbanchon, Rathelot, and Roulet (2021), our analysis builds upon a different set of assumption, together with more detailed data on job durations and family formation, in addition to taking account of unobserved heterogeneity.

Borghorst, Mulalic, and Van Ommeren (2022) – developed simultaneously to and independently from our work – use a linear probability model to estimate job–to–job mobility using administrative data from Denmark. They observe annual, rather than daily, wages and a restricted set of firm level characteristic, which is why they instrument the wage using a leave–one–out approach. In this part of the analysis, they only allow a restrictive set of model parameters to vary with regards to gender.[4] The main emphasis of the paper is on a purely hedonic wage model with which they estimate that 3.6% of the residual gender wage gap is due to differences in compensation for commuting.[5]

In earlier work based on cross-sectional data, Manning (2003*b*) produced some of the first evidence that mothers' wages react more strongly to commuting distance. Van Ommeren and Fosgerau (2009) used a job search model with strong functional form assumptions and find no significant gender differences. Hirsch, König, and Möller (2013) and Lundborg, Plug, and Rasmussen (2017) also find indirect evidence for a relationship between commuting preferences and the gender/motherhood wage gap, as well as Albanese, Nieto, and Tatsiramos (2022) for the employment gap.[6]

Gutierrez (2018)(building on work such as White 1986; Black, Kolesnikova, and Taylor 2014) explicitly model residential location decisions, relying on a monocentric model of the city with a gradient of wages and housing costs. They find that among mixed-sex married couples in the United States, one-tenth of the gender pay gap among childless workers conditioning on age and years of education, and more than a fifth of the motherhood pay gap, are explained by commuting.

[4]For example, their wage offer distribution is not allowed to differ, nor is the wage effect on the job leaving probability.

[5]The validity of the hedonic wage model heavily relies on neoclassical labour market model assumptions, in stark contrast to an on–the–job search model. Using a hedonic model will lead to biased results not only in the presence of worker and firm heterogeneity but also in markets with search frictions (see Gronberg and Reed (1994) and Hwang, Mortensen, and Reed (1998)).

[6]Other early contributions include Rouwendal and Rietveld (1994) and Rouwendal (1999). Consistently with our findings, their results point to the crucial role of children in explaining commuting preferences.

In the wider literature on trade-offs between wages and job attributes, studies such as Reed and Dahlquist (1994); Deleire and Levy (2004); Felfe (2012) and Dale-Olsen (2006*a*) examine gender differences in willingness to pay for attributes including workplace safety, type of tasks, promotion opportunities and type of schedule. They use conditional logit models or duration models in a similar spirit to ours. Additionally, our results for married mixed-sex couples speak to the literature on breadwinner norms (Bertrand, Pan, and Kamenica 2013), which highlights the costs couples are willing to incur to avoid a wife out-earning her husband.

In the labour search literature Bowlus (1997); Bowlus and Grogan (2009), differences in the job search process of male and female workers have been formalised in a number of different ways: as differences in job offer arrival rates, in wage offer distributions, in job destruction rates, or in parameters governing exits into non-participation. As discussed above, our model allows for all these differences. Differences in the marginal willingness to pay to reduce commuting in our model arise from different instantaneous utility functions over wages and commuting distances between men and women.

Estimating marginal commuting cost from a partial labour market search model such as ours or Le Barbanchon, Rathelot, and Roulet (2021) requires an exogeneity assumption for residential location, in exchange for much richer modelling of job choices. Exogenous residential location does not necessarily mean it needs to be fixed, but our method is sensitive to endogeneous residential relocation during a job spell. Low residential mobility is consistent with a lack of re-optimisation during a job spell. The estimated rate of household residential mobility in Germany over a period of two years is estimated to be just over ten percent, substantially lower than the UK rate and only about half of the US rate (Sánchez and Andrews 2011). Moreover, residential mobility increases with educational attainment (ibid and references therein) and our sample excludes university graduates. We therefore work with a relatively immobile sample. In order to take better account of compensation in the form of lower housing prices, we condition on local rental prices and interact them with wages in heterogeneity analysis.

Data

Our data comes from administrative social security records[7]. Our analysis sample consists of the inflow into regular full- or part-time employment[8] between January 1st, 2000 and December 31st, 2013. Employment spells, including wages, are recorded at a workplace level in days. Observations are treated as censored if the spell continues after December 31st, 2013, or if a person reaches the age of 55 (in order to prevent retirement decisions influencing job leaving decisions). We restrict the analysis to West German workers, primarily in order to reliably identify the first birth for women.[9] In additional analysis focused on possible household interactions in commuting choices, we study a sample of mixed-sex married couples. Using geo-coded data, we match pairs of individuals who reside at the same geographical point, share the same last name and have an age difference of fifteen years or less.[10]

Wage income above a threshold is not subject to national insurance contributions and therefore top-coded. To mitigate bias arising from this selection, we restrict our analysis to workers without a university degree[11], whose wages are more likely to be recorded without top-coding. This exclusion also supports our assumption of an exogenous residential location and a predictable covariate process (i.e. that does not depend on future values), since they are less likely to negotiate over wages individually (Hall and Krueger 2012).

Following the literature, including Le Barbanchon, Rathelot, and Roulet (2021), we measure commuting as Euclidean distance between postcode area centroids, which closely correlates with travel time in Germany. The country's geography and limited public transport use for commuting (12% in 2008 (Follmer et al. 2010)) contribute to this close relationship. We condition on regional structure, and exclude observations with distances above 100 km (seven times the mean for men).

Our estimates take as given any adaptations such as the option to work from home as given (analogous to other job attributes in the literature, e.g., mitigation of injury hazards for Dale-Olsen 2006*a*). Even in 2023, the share of employees in Germany 'usually' working from

[7] Specifically, the Institute for Employment Research's *Integrated Employment Biographies*.

[8] Excluding apprenticeships, marginal employment, self-employment, and lifetime civil service appointments. For more details, see Appendix (Section).

[9] Since East German workers' records are only available from the early 1990s, it is difficult to distinguish first from subsequent births to East German women during our sample period. This problem is exacerbated by typical birth spacing patterns around reunification.

[10] For details on the matching process, see Goldschmidt et al. (2014).

[11] 26% of working-age women in Germany held a tertiary qualification towards the end of our sample period. (OECD 2014). Employer-reported education information during job spells is considered reliable and not subject to underreporting (Fitzenberger, Osikominu, and Völter 2005).

home was just 11% (Statistisches Bundesamt Destatis), much below the self-employed. This highlights that even post-pandemic, commuting remains an important feature of the labour market. It is therefore preferable to estimate willingness to pay among a population where working from home plays an especially limited role, so that our estimates are more easily interpretable as willingness to pay to reduce *commuting*. During our sample period (Brenke 2014), only 8% of employees in Germany 'primarily' or even just 'occasionally' worked from home (where the latter may supplement rather than replacing workplace attendance), and the share was highest in high-skilled occupations outside the focus of our analysis.

We analyse changes in the willingness to pay to reduce commuting before and after childbirth, identified by exits into mandatory maternity leave for women (Müller and Strauch 2017). Our sample of non-graduate women are unlikely to become mothers before entering the labour market. We are unable to identify fathers using the same method. However, in our household sample we are able to place the birth of children into the work history of men and investigate whether men's willingness to pay to reduce commuting varies after childbirth.[12]

Methodology

Our analysis builds upon an on-the-job search model extended to two-dimensional jobs (wage w and commuting distance d), closely following Van Ommeren, Van Den Berg, and Gorter (2000) and Gronberg and Reed (1994), showing that voluntary job transitions identify marginal willingness to pay for continuous job attributes without requiring additional data, such as information on rejected job offers in the case of search from unemployment (see Online Appendix A.1). Two important assumption underlie this model. Firstly, residential location is exogenous to the search process. This assumption is well-suited to our context, as we discussed in the opening section. Secondly, in addition to voluntary job transitions, employment spells end for exogenous reasons at a fixed rate .

Based on such a model, Gronberg and Reed (1994) showed that the instantaneous marginal rate of substitution or marginal willingness to pay for a job attributes can be expresses as the ratio of the marginal derivatives of the hazard rate to end a job. Such an equation lends itself to estimation. Note that we choose to focus here on the marginal willingness to pay to *reduce* commuting distance which is defined as:

[12] Adoptions cannot be identified with this approach. There were 6,373 adoptions in Germany in the first year of our observation period (compared to 766,999 births), with the number falling after that (Statistisches Bundesamt Destatis).

$$-\frac{\frac{\partial\theta(w,d)}{\partial d}}{\frac{\partial\theta(w,d)}{\partial w}} = -\frac{\frac{\partial u(w,d)}{\partial d}}{\frac{\partial u(w,d)}{\partial w}} \tag{1}$$

A duration model is the most direct and intuitive empirical implementation of the job search model and the resulting hazard rate. The daily frequency of our dataset closely approximates a continuous process. Job spells can and do start, and wages and other job characteristics can and do change, at any point during a month. The failure event is a job ending for any reason, including voluntary job-to-job transitions, layoffs and labour market exits. We observe job spells that may either be followed by another job spell or by a spell of missing data. Missing data might reflect unemployment, but also full-time education or other periods in self-employment, out of labour market or abroad.

There are two approaches in the literature to dealing with the (ubiquitous) difficulty in unambiguously identifying voluntary job transitions, even in survey data. Gronberg and Reed (1994), Dale-Olsen (2006*b*) and Van Ommeren and Fosgerau (2009) ignore the layoff rate under the assumption that involuntary exits are constant and exogenous. In contrast, papers such as Van Ommeren, Van Den Berg, and Gorter (2000) have information on unemployment transitions and treat these transitions as independently censored.

We follow the majority of the literature for two main reasons. Layoffs are rare in Germany, with an OECD job protection score of 2.5 in the late 2000s (higher than the OECD average, and much higher than the UK and the US) for regular contracts (OECD 2008). We conduct a sensitivity analysis treating job spells that are followed by non-employment for more than 30 days as censored, mirroring the approach of Van Ommeren, Van Den Berg, and Gorter (2000).[13]

We rely on exits from a job to identify willingness to pay. Some women may be prevented from entering the labour market by high commuting costs. If their willingness to pay is even larger in absolute terms, we estimate a lower bound (in absolute terms) for the willingness to pay of all women in the economy.

As explained in the opening section, a number of structural sources for different outcomes by gender in the search model have been suggested, such as different job offer arrival rates or different wage offer distributions. The fundamental source of differences in, for example,

[13]Another option would be to directly model the hazards of leaving a job for other reasons in a competing risk framework. Identification of such a model is much more difficult (Van Den Berg 2001), and examples including some form of amenity (Bonhomme and Jolivet 2009) do not easily translate to our application.

the wage offer distribution is outside the scope of the model and could include gender differences in productivity, bargaining power vis-à-vis the employer, or taste-based discrimination. However, none of these would affect marginal willingness to pay as a function of wage.[14]

Since the hazard rate depends on (w, d) only through the instantaneous utility $u(w, d)$, the structural source of differences in marginal willingness to pay across the (w, d)-plain has to be differences in the instantaneous utility function. The search environment, as described for example by the job offer arrival rate and the distribution of wages, is allowed to differ in many other ways for men and women, or indeed between individual workers of the same gender. These differences are captured firstly by our large number of covariates and secondly by the very flexible way we model the individual-level baseline hazard.

As for the underlying causes of differences in instantaneous utility (which, in turn, generate differences in marginal willingness to pay), a common assumption in the literature ever since Becker (1981) is that women's non-market time is more productive than men's. This could be the case because they remain responsible for the bulk of household and child-rearing tasks. Additional explanations include differences in access to a car as suggested by Best and Lanzendorf (2005), or in the disutility from travelling due to differences in taste or perceived safety. Moreover, these differences might arise due to social norms, following Akerlof and Kranton (2000), who argue that preferences may mostly be internalized norms in cases where group identity prescribes certain behaviours. Such if social norms dictate that mothers should be nearby, for example to attend school-related meetings or events or to be available in case of emergencies, the opportunity cost of time spent commuting would be higher for women than for men and for mothers than for non-mothers. In light of these considerations we allow for the interpretation that differences in preferences might ultimately be rooted in different constraints.

Exploiting the panel dimension of our data, we go beyond the existing literature to flexibly account for unobserved individual-level heterogeneity using Stratified Partial Likelihoods (Ridder and Tunalı 1999). This does not require proportionality of baseline hazards of different individuals, and coefficients are identified using within-individual variation, meaning that unobserved heterogeneity could affect hazards differently at different points in the job spell. It also allows for the inclusion of time-varying variables. Whilst OLS estimates of wage premia or linear probability models of job changes have been augmented with fixed effects

[14] It is common to model the marginal willingness to pay to reduce commuting as a function of the wage. Such a specification is justified by higher opportunity costs of commuting time for high-wage workers. In our application we use a more flexible form than the more typical log-form, guided by goodness-of-fit tests and a desire to better capture nonlinearities in preferences.

(e.g. Duncan and Holmlund 1983; Villanueva 2007; Borghorst, Mulalic, and Van Ommeren 2022) and duration models have used shared-frailty terms to capture scalar unobserved heterogeneity (Van Ommeren, Van Den Berg, and Gorter 2000), data limitations have prevented previous work from using this within-worker variation in a stratified partial likelihood model.

We use a proportional hazards specification of the form

$$\theta_{ij}(t|\mathbf{X}) = \theta_j(t)\exp(\mathbf{X_{ij}}(t)'\beta)$$

for a worker j in job i with baseline hazard θ_j and (time-varying) covariate vector $\mathbf{X}_{ij}(t)$.

Ridder and Tunalı (1999) argue that censoring might not be independent if analysis time returns to zero at the start of each new spell, as in our analysis. An interaction between censoring at the end of the observation period and the timing of failure in an earlier spell within the same group could then affect the types of job spells observed. This illustrates the main limitation of the stratified approach, namely that like any fixed-effects method, it cannot accommodate heterogeneity that changes within individuals across observations. To address this concern and check the sensitivity of our results with respect to censoring, we have also estimated the model on a sample where the censoring date is brought forward by two years. Our main results are unaffected.

Functional form choices for log relative risk face a trade-off: Linear and log-linear relative risk specifications are tractable and produce estimates of marginal willingness to pay to reduce commuting distance that are easy to interpret and to compare to previous work. However, they may oversimplify a complex relationship. Since the goal of this analysis is estimating a marginal cost of commuting, we prioritise finding a well-fitting specification for the effects of the wage over other covariates (most of which are sets of binary variables anyway).

We use fractional polynomials to find the best functional form for the wage (details in the Online Appendix). Based on these results, we use a two-term, linear and quadratic form for the wage. We let commuting distance enter the specification linearly to keep the estimate tractable and interpretable. Dimensions of heterogeneity in willingness to pay such as regional structure enter as dummies to ensure flexibility and produce willingness to pay estimates for interpretable groups.

We specify a stratified partial likelihood model with a log relative risk that is linear in commuting distance and linear plus quadratic in the daily wage. This specification yields a marginal willingness to pay that depends non-linearly on the wage. The hazard rate can be expressed as

$$\theta_{ij}(t|\mathbf{X}) = \theta_j(t)\exp(\beta_{w_1} wage_{ijt} + \beta_{w_2} wage_{ijt}^2 + \beta_d distance_{ijt} + \beta_{\mathbf{z}} f(\mathbf{Z_{ij}(t)})) \tag{2}$$

where the control vector $\mathbf{Z}$ includes the worker-level variables of age (linear and squared), and (sets of) dummies for full-time work, unskilled occupation, occupational field and a dummy for a worker's first job. Moreover, we include regional GDP growth, local unemployment rates and local settlement structure (core cities, urban and rural areas, subdivided into a total of nine categories). We also include an interaction between each of the nine settlement structure dummies and a dummy for zero distances, allowing for a discontinuity in willingness to pay at the lower bound. This addresses potential bias from different behaviours at the lower bound caused by different sizes of postcode areas in rural and urban areas. For women, we include time-varying dummies switching to one at the first and second birth as well as for the youngest child becoming older than 12 years. As this information is constructed from information on (mandatory) maternity leaves (Müller and Strauch 2017), we are unable to reliably identify childbirth in men's biographies in the main sample. Functional forms for age, local unemployment and growth are chosen using a fractional polynomials routine. All variables except age and the dummy for a worker's first job are time-varying. Due to the potential multicollinearity with job duration and therefore with the baseline hazard, we measure age at the beginning of the job spell. In addition, we include a dummy for a worker's first-ever job, i.e. if the worker enters the job with zero work experience.[15] The level of observation in our data is a span, or national insurance record.

Plugging the baseline functional form (2) into equation (1) gives us marginal willingness to pay to reduce commuting as

$$MWP = -\frac{\beta_d}{\beta_{w_1} + 2\beta_{w_2} wage_t} \tag{3}$$

Estimation and Results

The summary statistics are presented in Table 1. On average, men's daily earnings are 33% higher than women's and their commutes are 18% longer. Over ninety percent of men, but only just over 60% of women, work full time and the share of jobs in unskilled occupations

[15]We decided to use this specification after experimenting with a number of functional forms for experience prior to the start of the job, aiming for flexibility to capture genuinely important heterogeneity without demanding too much of the data in identifying the closely related effects of time-by-individual via the baseline hazard in addition to age and experience.

is higher among women than men. The Online Appendix A.2 gives details on variable definitions, corrections applied and rules for inclusion in the sample. In addition, Table 2 compares summary statistics for childless women with those for all mothers. Mothers earn lower daily wages and have shorter commutes than childless women and are less likely to be in full-time jobs.

Table 1: Job-level summary statistics, baseline estimation sample

	Women		Men	
	Mean	Std. Dev.	Mean	Std. Dev.
Daily Wage, 25th percentile	40.3		58.4	
Daily Wage, mean	62.2	30.1	82.7	34.5
Daily Wage, 75th percentile	79.4		102.2	
Euclidean distance in km	11.5	14.0	13.6	15.8
Age at start of job	35.9	10.4	35.2	10.1
Full-time work	0.61	0.48	0.91	0.26
Unskilled job	0.13	0.33	0.10	0.29
Major cities	0.28	0.45	0.27	0.44
Urban areas	0.50	0.50	0.50	0.50
Rural areas	0.22	0.41	0.23	0.42
Child(ren) present	0.40	0.49		
Child(ren) over age 12	0.20	0.39		
Observations	6,433,713		6,876,548	
Jobs	2,435,009		2,679,887	
Persons	968,607		1,027,065	

The sample consists of an inflow of workers into regular full- or part-time employment from 2000–2013 in West Germany (IEB). Values are weighted by length of time they are observed within the job. Wages are measured at constant prices (base year 2013). In this baseline estimation sample, the birth of children cannot be observed in the employment history of men.

The baseline estimation (Table 3) implies daily marginal commuting costs for childless women of €0.27 at the mean wage per km distance to work, or 0.44% of the daily wage. At the 25th percentile of the wage, the figure is €0.22 (0.54%), and €0.34 (0.43%) at the 75th percentile (see Table 4). Figure 1 plots the marginal willingness to pay across a range of wages for women and men.

Men's estimated marginal willingness to pay per kilometre is 15% higher than that of childless women when evaluated at the respective mean wages for each gender. However, this largely reflects men's higher wages. Men and childless women have a similar willingness to pay over a range of wages (see Figure 1), they deviate once the wage increases beyond 80

Table 2: Job-level summary statistics, mothers and childless women

	Childless women		Mothers	
	Mean	Std. Dev.	Mean	Std. Dev.
Daily Wage, mean	65.4	30.5	57.6	28.8
Euclidean distance in km	12.2	14.6	10.5	13.0
Age at start of job	32.4	10.4	41.1	7.88
Full-time work	0.73	0.43	0.45	0.48
Unskilled job	0.12	0.32	0.14	0.34
Major cities	0.31	0.46	0.24	0.43
Urban areas	0.48	0.50	0.52	0.50
Rural areas	0.21	0.41	0.23	0.42
Child(ren) present	0	0	0.99	0.10
Child(ren) over age 12	0	0	0.48	0.49
Observations	2,346,626		1,658,076	

The sample consists of an inflow of workers into regular full- or part-time employment from 2000–2013 in West Germany (IEB). Values are weighted by length of time they are observed within the job. Wages are measured at constant prices (base year 2013). Presence of children is also time-varying, leading to values below one for the average of the children's variable.

Euro per day. Once the wage of childless women exceeds this threshold, their willingness to pay to reduce commuting rises above the one of men.

We find that marginal commuting cost jumps by 130% upon the birth of a woman's first child. This large increase supports the hypothesis that women's higher commuting costs are related to the time cost of non-market work, particularly childcare. Fathers are not legally required to interrupt their market work upon childbirth and are much less likely to do so. Therefore, an analogous analysis for men using exits to parental leave is not possible. However, we return to this question when analysing a matched sample of married couples in Section .

Both an increased marginal effect of commuting distance on the job mobility hazard and a decreased marginal effect of the wage contribute to the increase in women's marginal commuting cost upon childbirth (Table 3). In the context of our model, this implies that the marginal utility of a higher wage has decreased relative to that of a shorter commute . This pattern is consistent with new mothers specialising in non-market work.

Family composition also affects women's job mobility directly. After a first birth, the hazard of leaving a job declines. The direct, additional effect of a second child is much smaller and insignificant. The youngest child reaching the age of twelve is associated with

Table 3: Baseline estimation: Cox partial likelihood model of exits from a job, stratified by individual

	Women		Men	
Age	-.0074*	(.003)	.0553***	(.0028)
Square root	-.0733*	(.035)	-1.02***	(.0324)
Full time	.219***	(.0037)	.167***	(.0059)
First child	-.392***	(.018)		
Second child	-.0183	(.0225)		
Youngest > 12 yrs	-.0808***	(.02)		
Wage	-.0275***	(2.8e-04)	-.0249***	(2.1e-04)
Squared	9.3e-05***	(1.7e-06)	7.6e-05***	(1.1e-06)
Distance	.0044***	(1.4e-04)	.0039***	(9.7e-05)
Child × Wage	.0054***	(4.8e-04)		
Child × Wage squared	-2.2e-06	(3.1e-06)		
Child × Distance	.0025***	(2.7e-04)		
2nd child × Wage	-.0032***	(6.3e-04)		
2nd child × Wage squared	1.3e-05**	(4.2e-06)		
2nd child × Distance	1.6e-04	(3.6e-04)		
Older child × Wage	.0022***	(5.7e-04)		
Older child × Wage squared	-2.6e-05***	(3.8e-06)		
Older child × Distance	-.0018***	(3.3e-04)		
...	...	...	...	
Observations	6,433,713		6,876,548	
Jobs	2,435,009		2,679,887	
Persons	968,607		1,027,065	

Stratified Cox partial likelihood model, additional controls: unskilled occupation dummy, occupational field, regional structure (9 types), establishment size (5 dummies for the 10th, 25th, 50th, 75th and 90th percentile, respectively), local unemployment rate, local GDP growth. Zero work experience, and zero distance in each type of region (urban to rural), are captured by separate dummies. In this baseline estimation sample, the birth of children cannot be observed in the employment history of men. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

an additional reduction in job mobility, but the impact of distance on the hazard reduces. This is reflected in the marginal willingness to pay to reduce commuting partly bouncing back for mothers of older children, consistent with fading time pressures from non-market work as children grow more independent.

Older workers of both genders have lower job mobility, as we would expect.[16] Both men

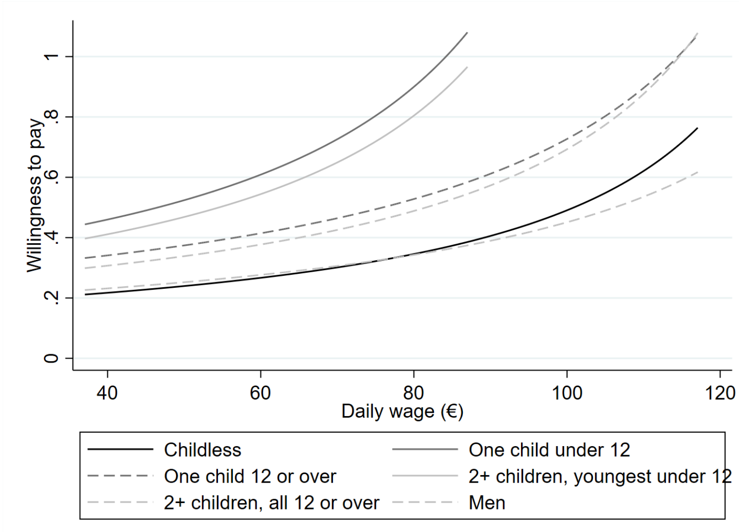
[16]The estimated hazard ratio of a thirty-year old compared to a twenty-year-old worker is 0.624 for men and 0.865 for women; however, note that these are not directly comparable since the estimate for women is conditional on the birth of children, but the one for men is not.

Table 4: Baseline estimation: Marginal willingness to pay to reduce commuting (in Euro per day per km)

	Low wage		Mean wage		High wage	
Childless	.218	(.007)	.274	(.0087)	.343	(.0109)
One child u12	.462	(.018)	.631	(.0242)	.888	(.0354)
One child over 12	.342	(.02)	.424	(.0244)	.524	(.0305)
2+ children, youngest u12	.413	(.0211)	.564	(.0283)	.794	(.0414)
2+ children, all over 12	.308	(.0236)	.387	(.0292)	.484	(.0369)
Men	.243	(.0062)	.316	(.0081)	.416	(.0109)

Results are based on baseline estimation of the stratified Cox partial likelihood models of Table 3 and using Equation 3. Full-time and part-time willingness to pay evaluated at the overall average daily wage by gender. In this baseline estimation sample, the birth of children cannot be observed in the employment history of men. Low and high wage are the 25th and the 75th percentile of overall daily wages by gender, respectively. Standard errors in parentheses.

Figure 1: Women's and men's marginal willingness to pay at different wages (in Euro per km)



Note: Confidence intervals are not displayed in order to maintain clarity, standard errors are reported in Table 4.

and women have higher rates of job mobility in areas of higher economic growth. Full-time workers are more likely to transition with a bigger effect for women, likely reflecting more career progression. Jobs in unskilled occupations have higher hazard rates.[17]

The results of Le Barbanchon, Rathelot, and Roulet (2021) lie in the range of our results

[17]Recall that the effect of any time-invariant worker attributes such as initial educational attainment is captured by the individual baseline hazard.

but exhibit less strong variation with regard to gender and family status. We find a lower willingness to pay for men, of €0.32 per km at the mean daily wage vs their estimate of €0.43 at their mean wage.[18] Our estimates are higher for particular groups of women, especially women with young children. We estimate their willingness to pay as €0.63, that is a 150% increase in willingness to pay compared to men at mean wage of women. In contrast, they find a 22% gender difference with little variation by family status, except for married with children where they find no gender difference). They do not allow their effect to vary by the age of children, which might explain part of the discrepancy in results. But more traditional gender norms in West Germany compared to France could plausibly play an important role here.[19]

Our results are also comparable to earlier results in the Netherlands. Russo, Van Ommeren, and Rietveld (2012) estimate a willingness to pay of €0.49 for a mixed-gender sample of employees of Amsterdam's Vrije Universiteit, and Van Ommeren, Van Den Berg, and Gorter (2000) estimate 0.4 Guilders or €0.18[20] at the mean wage for men.

We estimate a model that allows willingness to pay to vary by the area's settlement structure, distinguishing between core cities, urban areas and rural areas. We report our estimates of willingness to pay in the different types of areas in Panel A of Table 5. For the remainder of the paper the estimation results of the Cox models can be found in the Online Appendix as indicated below the tables for the willingness to pay.

For childless women, willingness to pay varies little across the different types of regions. Evaluated at the overall mean wage, it is slightly higher in urban areas than either in core cities or rural areas. But once a child is born the variation becomes large. For all groups of mothers, willingness to pay to reduce commuting is highest in rural areas, somewhat lower in urban areas, and by far the lowest in core cities. The ordering is the same for men, but the differences are smaller.

The presence of a child interacts with the regional structure and distance, i.e. the marginal value of a shorter commuting distance increases by more after the first birth in

[18] When expressing their results in terms of mean daily wage.

[19] The 2008 European Values Study (EVS 2008) found that 58% of respondents in Germany agreed that a child is likely to suffer if their mother works, compared to 39% in France (notably, the gap has closed in the most recent wave of the survey, after the period we analyse in our work; both countries are relatively middle-of-the-pack compared to the range between countries such as Sweden, Norway and Finland at around 20% agreement, and Italy and Turkey at more than 75%). Moreover, 77% of respondents in Germany, but 89% of respondents in France, agreed that sharing household chores is important for a successful marriage or partnership, and 35% of respondents in France agreed that marriage is an outdated institution, higher than any other country in the study, including Germany (27%, and slightly lower among women).

[20] Adjusting for CPI growth since 2000, this would be €0.24 in 2013, the end of our observation period.

Table 5: Marginal willingness to pay to reduce commuting (in Euro per day per km), by region type and part-time status

A: By region type	Cities		Urban areas		Rural areas	
Childless women	.251	(.0161)	.297	(.0126)	.263	(.0173)
One child u12	.437	(.0419)	.68	(.0325)	.77	(.0511)
One child over 12	.243	(.0516)	.41	(.0382)	.573	(.0785)
2+ children, youngest u12	.393	(.0412)	.601	(.0337)	.665	(.0477)
2+ children, all over 12	.221	(.05)	.4	(.0342)	.479	(.0469)
Men	.286	(.0156)	.318	(.0109)	.346	(.0174)
B: By part-time status, women only						
	Full time		Part time			
Childless	.272	(.0095)	.318	(.0194)		
One child u12	.688	(.0317)	.521	(.0284)		
One child over 12	.38	(.0323)	.333	(.0282)		
2+ children, youngest u12	.625	(.0354)	.484	(.0295)		
2+ children, all over 12	.357	(.0353)	.493	(.0395)		

Results are based on baseline estimation of the stratified Cox partial likelihood models of Table A1 and Table A2 in the Online Appendix and using Equation 3. In this baseline estimation sample, the birth of children cannot be observed in the employment history of men. Each willingness to pay evaluated at the overall average daily wage by gender. Standard errors in parentheses.

areas outside of core cities. Better provision of childcare (Connelly 1992; Del Boca 2002; Del Boca and Vuri 2007) could partly explain this. Unfortunately, we do not have data on childcare availability during the the first years of our time period, and shortening the panel affects the performance of our methodology. However, it is well documented that core cities offer more hours of childcare (Statistisches Bundesamt Destatis).[21] Relatedly, core cities also likely offer a greater variety of job bundles, including flexible working hours or workplace-based childcare. This reduces the need to 'triangulate' between home, work and childcare and makes it easier to find a good fit. Regional differences in willingness to pay persist for mothers of older children, but are reduced. This could reflect better public transport in core cities, which allows older children to be more independent.

We allow for interactions of the wage, distance, presence of children and part-time status and report the resulting marginal willingness to pay separately for part-time and full-time working women in Table 5, panel B. There are significant differences, but all are driven by

[21] Even in the cities with the highest rates (Frankfurt and Heidelberg), the share of under-threes in full-time public childcare was only slightly above 25%. The share of under-threes enrolled in any formal childcare in Germany as a whole (with higher rates in East Germany raising the average) was below that in France, the UK or Spain, and below the OECD-29 average during this period (OECD 2016).)

differential marginal effects of the wage and its interactions, rather than of the commuting distance. Part-time working women have a higher willingness to pay before the birth of a first child as a share of the average wage. This is because, across most of the range of daily wages, part-time working women have a lower marginal utility of the wage than full-time working women. Note that this comparison uses a common mean wage across both groups; if we instead evaluate marginal willingness to pay at full-time and part-time workers' respective mean daily wages, it is lower for women who work part-time (€.245) than for those who work full-time (€.301). The jump upon the birth of the first child (as a share of a constant, overall mean daily wage) is smaller for part-time working mothers. Again, this is driven by changes in the effect of the wage, specifically by the marginal effect of the wage diminishing for full-time working women after childbirth, but less so for part-time working women. One interpretation of this pattern is that reducing commuting and part-time work are substitutable margins of adjustment after childbirth.

We estimate an alternative, non-stratified model to quantify the role of unobserved heterogeneity in our setting. While we defer detailed results to the Online Appendix (section A.4), the estimated effect of wages on the hazard to leave a job is consistent with attenuation bias. In contrast, the effects of commuting distance and first birth are biased upward when unobserved heterogeneity is not accounted for, and consequently, so is the estimated marginal willingness to pay for women. Willingness to pay is also biased upward for men, but the impact is much smaller (Online Appendix Table A8. This indicates that unobserved heterogeneity is, unsurprisingly, correlated with motherhood, commuting distance and job mobility.

Following the discussion on the ambiguity of identifying voluntary job transition we conducted a sensitivity analysis where we treat job spells that are followed by non-employment for more than 30 days as censored, mirroring the approach of Van Ommeren, Van Den Berg, and Gorter (2000). The estimates of marginal willingness to pay to reduce commuting for men are largely similar (Online Appendix ,Panel C, Table A8, for detailed coefficients see the Online Appendix, Table A7). One exception is a smaller estimated willingness to pay for men with high wages compared to our baseline approach. In contrast, the estimates for childless women increase compared to the baseline approach. Otherwise, the results for women are in the same order of magnitude or slightly higher, resulting in a smaller but still remarkable increase of the willingness to pay upon the first birth (79%). The smaller increase could reflect the censoring mechanism, which treats jobs of women who exit the labour force voluntarily as censored. This might disproportionately affect women with a high willingness to pay to reduce commuting. We therefore prefer our baseline specification. Nevertheless, the results of this sensitivity analysis confirm our main qualitative results of a large and significant increase in the willingness to pay to reduce commuting upon childbirth.

To test the assumptions around independent censoring, we estimated a specification censored two years earlier. We discuss results from this specification, as well as a specification where we add housing cost, in the Online Appendix (section A.4). These results confirm the main conclusions of our analysis.

Endogeneity of residential location is a concern in our model. Workers may accept a job with a long initial commuting distance if they anticipate moving closer to their place of work in the future. This would lead us to understate their willingness to pay to reduce commuting distance. To address this problem, we tested specifications which move residential moves forward to the beginning of the job spell during which they occurred, exclude all job spells that include a residential move and censor job spells at the time of move (see Online Appendix Tables A9, A10 and A11). Surprisingly, the marginal willingness to pay to reduce commuting decreases when moving residential moves forward, indicating that the moves occurred mainly for other reasons than the trade-off between wages and commuting. Excluding jobs spells where a move occurred is also an imperfect solution: On the one hand, it excludes some spells that do not violate exogeneity if women and men move for reasons other than to reduce their own commuting distance. On the other, an *ex-post* fixed residential location is not a sufficient condition for exogeneity and it might produce a selective sample. When running our model on this restricted sample the estimated MWPs for women increase compare to the baseline model, whereas the MWP for men reduces slightly, indicating that we estimate a lower bound for women with our baseline approach. The censored sample confirms this conclusion, as the estimated MWP's lie in the middle of the baseline and stayer sample.

Married Couples' Sample

In our main dataset, we observe individual workers only. However, geo-coded data permits the identification of a subset of married couples, allowing us to shed light on household-level factors. The algorithm treats two people as a married couple if their (geo-coded) addresses match, they share a surname and are a man and a woman with an age gap of less than 15 years. Details of the algorithm and the circumstances under which individuals may be misclassified are given in Goldschmidt et al. (2014). It is very unlikely that two individuals who are not a couple are classified as one, but many actual married couples are missed. A comparison with the German microcensus suggests that the method identifies between 25 and 30 percent of all married couples in which the husband is 65 or younger. We work with married couples identified by this method in the universe of data.

Table 6: Job-level summary statistics, household estimation sample

	Women		Men	
	Mean	Std. Dev.	Mean	Std. Dev.
Daily Wage, 25th percentile	37.0		75.4	
Daily Wage, mean	57.8	28.6	97.7	34.0
Daily Wage, 75th percentile	73.2		117.1	
Distance in km	10.2	12.7	13.6	15.2
Age at start of job	38.6	7.95	40.7	7.84
Full-time work	0.46	0.49	0.94	0.22
Unskilled job	0.16	0.36	0.080	0.26
Major cities	0.22	0.41	0.22	0.41
Urban areas	0.54	0.50	0.54	0.50
Rural areas	0.24	0.43	0.24	0.43
Child(ren) present	0.61	0.49	0.62	0.48
Child(ren) over age 12	0.28	0.44	0.30	0.45
Partner's earnings	34,009	20469	19999	11894
Observations	1,938,346		1,489,065	
Jobs	680,118		570,418	
Persons	273,709		269,792	

The sample consists of an inflow of workers into regular full- or part-time employment from 2000–2013 in West Germany in the married couple sample. Values are weighted by length of time they are observed within the job. Wages are measured at constant prices (base year 2013). In this married couple's sample, the birth of a child identified in the employment history of a married women is transferred to her husband.

Given the construction of the married couples' sample, differences in descriptive statistics compared to our main sample (Table 6 vs Table 1) are to be expected. Women and men in the married couples' sample are older, are more likely to have children and to live in urban areas. Women are more likely to work part-time, earn a lower wage and commute less; men earn a higher wage and commute approximately the same distance.

When using the same baseline model specification in our married couples' sample, we also find some differences in willingness to pay compared to the original sample (Table 7 compared to Table 4). Evaluated at the respective mean daily wages (which are lower for women in the married couples' sample), childless women's willingness to pay is more than a third higher in the married couples' sample. The increase upon childbirth, however, is more moderate, leaving willingness to pay for mothers of one young child very similar in both samples. Distance becomes a more important determinant of leaving a job for women in both samples when they have children, with a similar-sized effect. However, willingness to pay increases by more in the main sample because the importance of the wage diminishes at the same time, which is not the case in the married couples' sample. Changes in the role of both wage and distance when the child gets older, and when there is a second child, are similar in both samples.

A key advantage of the married couples' sample is that it allows us to separately estimate willingness to pay for childless men and fathers. In absolute terms, men's willingness to pay in the married couples' sample is above that of their counterparts in the main sample for any subgroup, reflecting their higher earnings. Married men see a much smaller increase in their marginal willingness to pay than married women when a couple's first child is born (see Figure 2). Evaluated at their mean wage, the increase is just 12%. Children reaching the age of twelve reduces fathers' willingness to pay in the married couples' sample, as does a second child being born.[22] For these fathers the marginal willingness to pay returns close to the one of childless men in the married couples' sample, and of all men in the baseline sample. Once all children are above the age of twelve, the marginal willingness to pay drops further in the married men's sample.

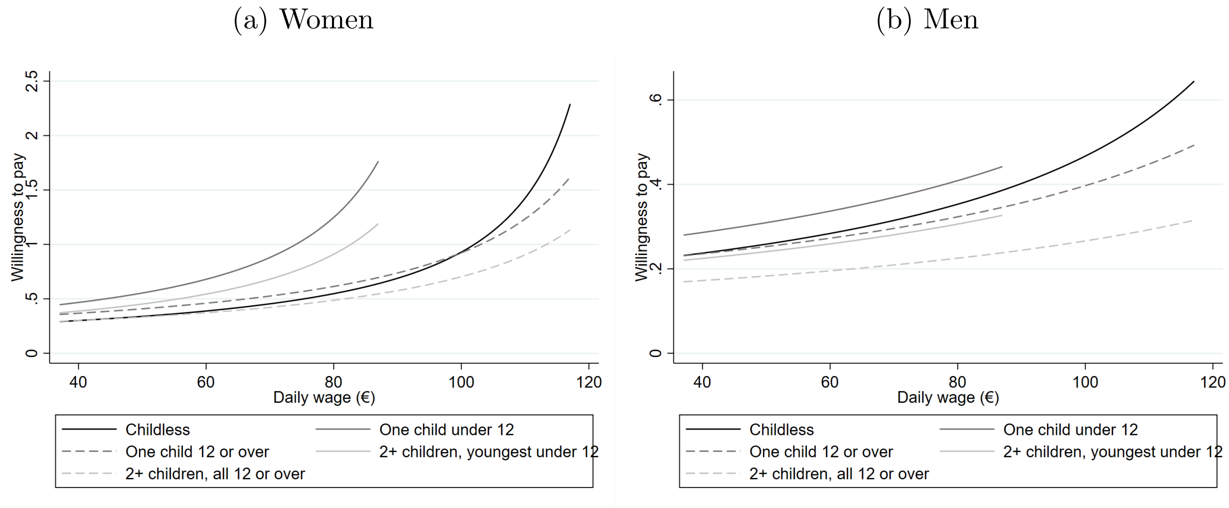
Table 7: Marginal willingness to pay to reduce commuting, married couples' sample

	Low wage		Mean wage		High wage	
Women						
Childless	.292	(.021)	.383	(.0271)	.496	(.0352)
One child u12	.449	(.0302)	.667	(.0444)	1.03	(.0721)
One child over 12	.358	(.033)	.455	(.0411)	.566	(.0514)
2+ children, youngest u12	.373	(.0319)	.534	(.0451)	.778	(.0677)
2+ children, all over 12	.294	(.0368)	.368	(.0454)	.452	(.0557)
Men						
Childless	.334	(.0304)	.451	(.0414)	.645	(.0652)
One child u12	.39	(.0327)	.505	(.0419)	.679	(.0593)
One child over 12	.31	(.0322)	.387	(.0397)	.493	(.0515)
2+ children, youngest u12	.294	(.0399)	.364	(.0489)	.46	(.063)
2+ children, all over 12	.218	(.0431)	.261	(.0512)	.315	(.0621)

Results are based on baseline estimation of the stratified Cox partial likelihood model of Table A3 and using Equation 5. In this married couple's sample, the birth of a child identified in the employment history of a married women is transferred to her husband. Low and high wage are the 25th and the 75th percentile of overall daily wages by gender in the married couples' sample, respectively. Standard errors in parentheses.

[22]Notably, the latter effect is *not* driven by an increased marginal effect of the wage (as might be expected if a second child places increased pressure on the family's budget), but by a significant reduction in the effect of commuting distance.

Figure 2: Married workers' marginal willingness to pay at different wages



In the final specification, we further exploit the household matching in the married couples' sample to allow the impact of a worker's own daily wage and commuting distance to depend upon their relative position within the couple.[23] Time-invariant differences between individuals will be captured by the baseline hazard, so the interaction effects are identified through changes in relative earnings within couples over time, rather than differences in relative permanent earnings potential.

The resulting willingness to pay to reduce commuting from this specification can be found in Table 8. When considering the relative wage position a clear pattern emerges, which is similar for both men and women. Willingness to pay is highest for the lower earner within the couple, resulting from a high sensitivity of job mobility to the wage. While this is consistent with differential preferences of couples as well as within-couple specialisation, the increase with the arrival of the first child remains very small, even for men who are the secondary earner (3-6%).

Looking at relative commuting distance, marginal willingness to pay to reduce commuting distance is lowest for women whose husbands commute further than they do. If the husband does not commute or commutes less than his wife, the wife's willingness to pay is higher. The pattern is different for men: If their wife commutes less than they do, the effect of distance on job mobility is *greater* than if their wife does not commute at all. In an even starker contrast, for husbands whose wife has a longer commute than they do, the effect of distance is reversed and they are *less* likely to leave a job if it has a longer commuting distance. This

[23]The spouse's relative wage is based on annual earnings and thus varies at the annual level, to abstract from very short-term fluctuations

is reflected in willingness to pay, with men whose wives have a longer commuting distance than they do having a high willingness to pay to *increase* their own commuting distance.

Thus we find strong evidence for the avoidance of non-traditional household commuting patterns; this echoes findings in the literature on relative earnings and breadwinner norms (Bertrand, Pan, and Kamenica 2013). Women try to reduce their commuting distance if their husband is commuting less than them, and men are even willing to reduce their wage in order to commute longer distances in case their wife commutes further. The sample design, consisting of mixed-sex couples where neither partner is degree-educated and that have chosen a shared last name, may mean that some couples with more progressive gender attitudes than the average are excluded from the analysis.

Table 8: Marginal willingness to pay to reduce commuting by spouse's relative earnings and commuting distance

	Spouse					
	not commuting		commuting less		commuting more	
Wife's willingness to pay, husband earning less						
Childless	.367	(.0323)	.371	(.014)	.168	(.0228)
One child u12	.571	(.0409)	.576	(.0217)	.338	(.0299)
One child over 12	.399	(.0366)	.403	(.0194)	.19	(.0268)
2+ children, youngest u12	.425	(.0365)	.429	(.0219)	.229	(.0279)
2+ children, all over 12	.291	(.0347)	.295	(.0221)	.113	(.0273)
Husband earning more						
Childless	.481	(.0425)	.486	(.0188)	.22	(.03)
One child u12	.792	(.0572)	.799	(.0311)	.469	(.0417)
One child over 12	.532	(.0491)	.538	(.0263)	.254	(.0358)
2+ children, youngest u12	.555	(.0479)	.56	(.029)	.3	(.0365)
2+ children, all over 12	.37	(.0442)	.375	(.0283)	.144	(.0347)
Husband's willingness to pay, wife earning less						
Childless	.336	(.0378)	.465	(.0152)	-1.27	(.0363)
One child u12	.358	(.0374)	.483	(.0198)	-1.2	(.0383)
One child over 12	.323	(.0396)	.447	(.0217)	-1.22	(.0398)
2+ children, youngest u12	.335	(.0419)	.462	(.0269)	-1.25	(.0455)
2+ children, all over 12	.3	(.0452)	.426	(.0304)	-1.27	(.0488)
Wife earning more						
Childless	.413	(.0465)	.571	(.019)	-1.56	(.0457)
One child u12	.437	(.0459)	.589	(.025)	-1.47	(.0493)
One child over 12	.394	(.0484)	.545	(.0273)	-1.49	(.0515)
2+ children, youngest u12	.41	(.0516)	.566	(.0338)	-1.53	(.0598)
2+ children, all over 12	.367	(.0555)	.521	(.0379)	-1.56	(.0643)

Results are based on baseline estimation of the stratified Cox partial likelihood model of Table A4 and using Equation 5. In this married couple's sample, the birth of a child identified in the employment history of a married women is transferred to her husband. All willingnesses to pay evaluated at the overall average wage by gender. Standard errors in parentheses.

Conclusion

Our key result from a large administrative dataset of non-university educated workers in West Germany is a substantial motherhood gap in marginal commuting cost. In additional analysis of a sample of married mixed-sex couples, we find that – in sharp contrast to mothers – fathers' willingness to pay to reduce commuting distance increases only slightly after the birth of a couple's first child.

When taking relative wages within the couple into account, we find that both men and women have a higher willingness to pay to reduce commuting distance if their partner earns more than they do. This is consistent with a specialisation or preference mechanism. In contrast, our results for relative commuting distance within couples are not gender-symmetric. In couples where the husband commutes less compared to his wife, the wife's willingness to pay to reduce commuting distance is higher, potentially leading to an equalisation of commuting distances. Even more strikingly, men seem to actively avoid commuting shorter distances than their wives. In cases where they do, they are actually willing to give up wages in order to increase, rather than decrease, their commuting distance. As the partner who commutes less is typically in charge of more household tasks, this is consistent with actively avoiding a non-traditional distribution of household tasks.

Furthermore, our estimates indicate that differences in wages and commuting distances between men and childless women are unlikely to arise from differences in the marginal willingness to pay to reduce commuting distance in our setting: Willingness to pay is very similar over a large range of wages. Other explanations are likely to be more important. These include differences in the job offer distribution and wage distribution (potentially reflecting discrimination), occupational sorting, or differences in productivity. Our approach conditions out differences in the job offer and wage distribution while simultaneously conditioning on occupation and productivity-related characteristics, which means that we can credibly isolate the role of willingness to pay to reduce commuting distance. This does mean that on the flip side, we cannot disentangle these other individual factors.

In additional analyses on our main sample of individuals, we find that childless women have a similar marginal willingness to pay in core cities, urban and rural areas, but mothers in core cities have a much lower marginal willingness to pay than their counterparts in urban or rural areas. Heterogeneity analysis by part-time status points towards part-time work and reduced commuting distance being substitutable margins of adjustment after child birth. We also consider the role of the housing market, finding that willingness to pay is slightly lower in areas where housing costs are high, but the differences are small and mostly insignificant.

To put the gap in marginal willingness to pay in the context of the motherhood pay gap, consider a woman employed at mothers' mean wage and mothers' mean commuting distance. To increase her commuting distance to the sample mean of childless women, a childless woman would need to be compensated by a wage increase of €0.47 (as a linear approximation of the willingness to pay at the mean wage), whereas a mother with a child under 12 needs to be compensated by €1.07. This difference amounts to about 8% of the raw motherhood wage gap, which is quite substantial.[24] Thus, our empirical results in combination with this back-of-the-envelope calculations indicate that commuting preferences are an important contributor to wage penalties for mothers relative to childless women and men.

A range of policy levers could reduce the cost of commuting, including improved transport infrastructure, high-quality internet infrastructure enabling working from home, and regulation or incentives encouraging firms to support hybrid working. Our results suggest that these could particularly benefit mothers of young children, and within that group, mothers outside of core cities whose willingness to pay to reduce commuting distance is particularly high. Relative to other well-known drivers of motherhood wage gaps, such as occupational sorting, commuting costs are likely to be relatively amenable to policy intervention. The potential of these policies to reduce inequalities is under-discussed relative to other objectives of, for example, investment in transport infrastructure. Other policy levers that address commuting costs for the specific groups currently bearing the highest cost are also promising: for example, access to childcare with opening hours that allow for longer commutes could be an important factor in underserved areas, such as outside of core cities.

However, results from a sample of non-university educated married couples highlight the limits of such interventions. In that sample, we find evidence of a crucial role for reference points and a perceived norm that husbands 'should' commute more than their wives do. We note that, since our matching algorithm relies on a shared surname, the sample likely excludes some more socially progressive couples. But nevertheless, in this context, interventions that aim to reduce the direct, economic cost of commuting - for example, through faster and more reliable public transport connections - may not be as effective at closing gender and motherhood gaps as we would hope if women face a socio-psychological cost of non-compliance with a norm to commute less than their husbands.

[24]Our measure of willingness to pay is a local one and can only approximate infra-marginal differences such as the one between mothers' and childless women's average wages. We have chosen to use the raw wage gap as it is not a priori clear which covariates should be included in order to calculate the residual wage gap. For example (Le Barbanchon, Rathelot, and Roulet 2021)) calculate a contribution of the gender differences in the willingness to pay to the residual gender wage gap of 10–14%, depending on which covariates are included.

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A Online Appendix

A.1 Model

In this section, we outline an on-the-job search model extended to two-dimensional jobs, closely following Van Ommeren, Van Den Berg, and Gorter (2000). Voluntary job transitions identify marginal willingness to pay for continuous job attributes without requiring additional data, such as information on rejected job offers in the case of search from unemployment.

Consider an employed worker in a job with wage w and commuting distance d, who receives alternative job offers (w^*, d^*) drawn from a distribution $F(w^*, d^*)$ according to a Poisson process with arrival rate λ. Thus, an important assumption underlying the model is that residential location is exogenous to the search process. This assumption is well-suited to our context, as we discussed in the previous section.

In addition to voluntary job transitions, employment spells end for exogenous reasons at rate δ. The expected discounted stream of utility from accepting job offer (w, d) over the life course is

$$\begin{aligned}\rho R(w,d) = u(w,d) + \theta \int\int \max\{0, R(w^\star, d^\star) - R(w,d)\} dF(w^*, d^*) \\ + \delta(U - R(w,d))\end{aligned} \tag{4}$$

where ρ is a discount parameter and U is the expected present value of unemployment. In other words, lifetime utility is composed of an instantaneous component, a continuation value in case of a job switch and another continuation value in case of exogenous job loss.

The optimal strategy, as in the one-dimensional job case treated by Mortensen (1986), is myopic. The reason for this is that lifetime utility R depends on (w, d) only through instantaneous utility $u(w, d)$ and there are no transaction costs. Intuitively, whereas in a model without on-the-job search, a worker may hold out for a better offer, in this case a worker has nothing to lose by accepting a job offer. She will still have an equal chance of receiving a better offer on the job.

Therefore, the worker pursues a reservation utility strategy: She accepts all job offers which offer a higher instantaneous utility than her present job, since the future stream of job offers is not affected by the job currently held. Formally, the set of job offers that are acceptable (i.e., strictly preferred to the current job) is

$$\varsigma(w,d) = \{(w^*,d^*)|u(w^*,d^*) > u(w,d)\}$$

This search and decision process leads to the following specification for the hazard rate from a job (w,d):

$$\theta(w,d) = \delta + \lambda \int_{\varsigma(w,d)} dF(w^*,d^*) = \delta + \lambda(1 - F_u(u(w,d)),$$

i.e. the rate of exit from a job is given by the rate of exogenous exits into unemployment, plus the product of the rate of arrival of alternative offers and the probability that the offer will induce the worker to switch jobs. The second expression follows by substituting the above characterisation for the set of acceptable job offers, with F_u denoting the c.d.f. of $u(w,d)$.

As stated before, lifetime utility in this model depends on the wage and the commuting distance only through instantaneous utility. Therefore, the partial derivative of the hazard rate with respect to the wage w can be expressed as

$$\frac{\partial\theta(w,d)}{\partial w} = \frac{\partial\theta(w,d)}{\partial u(w,d)}\frac{\partial u(w,d)}{\partial w}$$

Clearly, an analogous statement holds for the derivative with respect to the commuting distance d.

This, in turn, gives us the equality stated by Gronberg and Reed (1994): the instantaneous marginal rate of substitution or marginal willingness to pay for a job attributes is equal to the ratio of the marginal derivatives of the hazard rate:

$$\frac{\dfrac{\partial\theta(w,d)}{\partial d}}{\dfrac{\partial\theta(w,d)}{\partial w}} = \frac{\dfrac{\partial\theta(w,d)}{\partial u(w,d)}\dfrac{\partial u(w,d)}{\partial d}}{\dfrac{\partial\theta(w,d)}{\partial u(w,d)}\dfrac{\partial u(w,d)}{\partial w}} = \frac{\dfrac{\partial u(w,d)}{\partial d}}{\dfrac{\partial u(w,d)}{\partial w}} \tag{5}$$

Unlike other commonly studied job attributes, we would expect commuting distance to generate disutility. For ease of interpretation, we therefore choose to focus on the marginal willingness to pay to *reduce* commuting distance which is defined as:

$$-\frac{\frac{\partial\theta(w,d)}{\partial d}}{\frac{\partial\theta(w,d)}{\partial w}} = -\frac{\frac{\partial u(w,d)}{\partial d}}{\frac{\partial u(w,d)}{\partial w}} \tag{6}$$

Thus, the negative of the ratio of the marginal derivatives of the hazard rate gives us the pay that a worker is willing to give up in case commuting distance is reduced by one km.

Regional Labour Market Conditions As an extension to their basic model, Van Ommeren, Van Den Berg, and Gorter (2000) discuss the inclusion of business cycle effects in the model.[25] They would affect the rate of arrival of job offers λ and/or the distribution $F(w, d)$ from which wage offers are drawn. Realistically, not only macroeconomic conditions at the national level should affect these two structural parameters of job search, but regional trends could also enter into the hazard rate.

In our empirical specification, we therefore include dummies for a typology of the local settlement structures, as well as local unemployment and growth rates to reflect regional labour market conditions. In a sensitivity analysis, we additionally control for local rent levels, as they are only available for subsamples of years.[26] Note that due to our exogeneity assumption of the residential location, we define the local labour markets around the residential location, which conserves the stationarity of the decision problem.

A.2 Variable Definitions

Occupational Characteristics The dataset contains two **occupation** variables, a 3-digit variable based on the 1988 classification (Bundesanstalt für Arbeit 1988), and a 5-digit variable based on the most recent classification (Bundesagentur für Arbeit 2011). Since the observation window ends in 2013 and re-coding of older observations to the 2010 system is not error-free, the older variable is likely to be more accurate. However, the 2010 classification combines a horizontal (occupation) and a vertical (skill level) dimension. We recover the skill level information, which is absent from the older variable at the available aggregation level.

[25] Van Ommeren, Van Den Berg, and Gorter (2000) also discuss a number of other assumptions of the basic on-the-job-search model that can be relaxed with the basic result remaining valid, for example endogenous search effort.

[26] We have also experimented with indices counting regular employment relations in the individual's county of residence and in neighbouring districts, in her occupational field, by gender. The intuition is that the individual is likely to receive offers to work in her own profession, as well as in other professions within the same occupational field, which are defined with respect to similarity of tasks performed and skills required. Results were unaffected by the inclusion of different local labour market indicators.

For the horizontal occupation information, we match the 1988 information to 53 task-based *occupational fields*, as defined by the Institute for Vocational Education and Training.

Unskilled is the lowest of four skill levels, characterised as *un- or semiskilled activity* with simple or routine tasks of little complexity, where formal vocational training is not usually required.

Regional Characteristics We match the individual data to the Federal Institute for Research on Building, Urban Affairs and Spatial Development (BBSR)'s classification of 9 types of districts. They are based on administrative districts, but differ from them where administrative divisions group structurally different areas into one unit. We include dummies for the type of area and interactions with a dummy for distances of zero (workers who live and work in the same postcode area) to account for the larger geographical size of postcode areas in sparsely populated regions. Moreover, we estimate a separate willingness to pay for three broader types of area, proposed by the institute as characterising city-periphery relationships:

- Core cities
- Districts with a predominantly urban character ("urban areas")
- Districts with a predominantly rural character and rural areas ("rural areas")

For more detailed information on the classification, see Görmar and Irmen (1991) or the institute's online information (Bundesinstitut für Bau-, Stadt- und Raumforschung 2006).

Rental cost The proxy for rental cost is also provided by the Federal Institute for Research on Building, Urban Affairs and Spatial Development. It is based on asking prices for flats gathered from online platforms and newspapers, using the following criteria:

- pure rental prices with no heating or other utilities included
- non-furnished flats between 40 and 130 square metres
- the ad is displayed for no more than six months
- some additional filters to exclude implausible levels and changes

The providers suggest that their measure is likely to omit some flats offered by very large housing companies, particularly in Berlin and Hamburg, who use their own information channels. It is also likely to omit some flats in rural areas which are only advertised on local notice boards or find a new tenant by word of mouth. Actual rent paid may be slightly lower in areas of low demand where prospective tenants are able to negotiate a lower price.

Other details

Definition of employment spells The self-employed, civil servants and workers in marginal employment (*geringfügig Beschäftigte*, who are exempt from contributions) are not covered by the data or not covered in a consistent way throughout the period, and thus excluded from the analysis. As these types of work are structurally different from regular employment especially with regard to mobility, excluding them also provides a more homogeneous sample. In addition, we exclude apprenticeships and jobs within the context of an active labour market programme, and jobs with a wage or mobility subsidy, since the observed wage and/or commuting distance do not adequately describe the worker's decision problem in these cases. For consistency, we also exclude jobs which switch back and forth between regular and marginal or sponsored employment.

Treatment of missing data A typical job spell used in the model consists of a number of spans, which correspond to national insurance records. There is at least one record per year, plus additional records in case of changes in the employee's data, e.g. a change in the wage. If the **wage** was coded as zero or missing in a span, but valid wage information was available in another span within the same aggregate job spell (a continuous employment at the same firm), we extend the valid wage information to the missing observation.

Residential Moves The residential location recorded in the data corresponds to the end of the span. It hard to gauge the extent to which employers proactively register their employees' changes in residential location with national insurance other than when they make their standard yearly report, or at the end of a job spell. It is plausible that at least some employers simply wait until the next regular entry is due, so we might observe residential moves with a certain delay.

Children We use two dummies for the birth of the first and second child, respectively. These dummies then stay at one. We additionally use dummies to capture the youngest child in the family reaching the age of twelve, to reflect differences in the time constraints of parenting younger versus older children. The timing of births is identified using the routine set out in Müller and Strauch (2017). This is based on exits from employment into the mandatory part of maternity leave. Since entry into maternity leave triggers a national insurance notification, this point is already the end of a span. We do not treat women who are on maternity leave as at risk of a transition (or in other words, maternity leave is a "stop the clock" period from the point of view of the hazard model).

Treatment of tied events Estimating a Cox model in continuous time means that ties arise only as a consequence of imprecise measurement, not as a true feature of the data-generating process. To handle them, we use the Breslow approximation (Breslow 1974; Peto 1972). It calculates the partial likelihood assuming that both individuals recorded to fail at the same time are in the risk sets at each other's failure times. This approximation introduces a bias of the coefficients towards zero, but it is the least computationally demanding and performs well if ties are not too frequent (Kalbfleisch and Prentice 2002, p. 105).

Method of fractional polynomials This method runs through a set of functions, and applies a formal deviance criterion to choose the best form. The available functions are degree-1 and -2 additive combinations of natural logarithms, fractional and integer powers from the set $\{-2, -1, -0.5, 0, 0.5, 1, 2, 3\}$. Evaluation of alternative specifications uses comparative measures based on the log partial likelihood, such that a higher-degree functional form is adopted if it leads to a significant change in the transformed likelihood. For a detailed discussion of fractional polynomials including an application to a Cox model, see Royston and Altman (1994).

Sample Construction

We start with a 10% sample of all individuals with a national insurance number, going back to 1975.

Employment spells are included in the sample if:

- they are part of the inflow sample starting on January 1st, 2000. The data is right-censored on December 31st, 2013.
- they last for more than 60 days. Temporary workers whose contracts last less than two

months are usually not liable to pay full social security contributions, which should preclude their inclusion in the sample. Spells of under two months could be due to exceptions in the national insurance treatment, early firings, miscoded part-time work, or misreported dates, which are difficult to disentangle. Moreover, the optimisation process underlying short-term job location may differ substantially from the one related to long-term job mobility decisions and temporary residential relocations are likely to not appear in the data, which makes distance calculations unreliable. Therefore, spells of under two months are dropped.

- the implied monthly wage is within the limits that make a worker liable to pay national insurance contributions (*Geringfügigkeitsgrenze* and *Beitragsbemessungsgrenze*). Due to different timings of reports, wage information in some spells which are not actually subject to contributions was included in the original dataset.

- they are not overlapped by a spell in registered unemployment or an active labour market programme [27], a mobility-related subsidy or retirement. Small overlaps of up to three days are tolerated. Individuals are observed as registered unemployed if they are eligible for top-up unemployment benefits to close the gap between low earnings and the subsistence level. In this case, the wage paid by the employer is not the wage actually perceived by the worker, who faces a wage distribution that is truncated at the legal minimum subsistence level. A similar distortion of the wage-commuting trade-off arises in the case of a mobility subsidy. Participants in active labour market programmes, on the other hand, do not choose their place of work, and their behaviour can therefore not be adequately reflected in the model. Therefore, these cases are not included in the sample. Selection into standard (i.e. non-subsidised) employment is not addressed here.

- the individual is never recorded as having a university degree, with certain corrections applied. To avoid complications arising from the decision to return to education, we do not include employment spells before university graduation. We do not know if individuals acquire a university degree after the end of the observation window. Eight to ten years after leaving vocational education, this is unlikely to apply to many individuals.

- they belong to a job identified as the main job at that time (more details below).

[27] a programme to support the long-term unemployed, publicly sponsored employment, or a seasonal or temporal work placement organised by the employment agency

- they are not part of a seasonal work pattern, i.e. the worker does not return to the same employer without an intervening spell at a different firm. Spells with the same employer with gaps of up to a week are considered part of a single job to avoid misinterpreting administrative delays to contract renewal as seasonal work. This does not apply to leaves of absence for maternity or illness, which are treated as "stop the clock" periods during which a worker is not at risk of a transition.

We exclude spells where either the place of work or the place of residence was missing or invalid, or where an individual was recorded as living in, or a firm recorded as being located in, two or more different zip code areas at once, since no valid commuting distance can be determined in those cases.

Treatment of overlapping employment spells Overlapping spells present a challenge to the model, since neither the theory nor the empirical model allows for an agent to be in two states at once. To keep the model tractable, we make the simplifying assumption that individuals have one main job, and mobility behaviour in any other jobs is not reflected in the model. Cases where no clear hierarchy of parallel jobs can be determined are excluded.

Multiple job spells with different employers at the same time Overlapping job spells of the same individual with different employers are excluded, except in the following cases:

- **Transitional overlap:** If the overlap is less than two weeks, both spells are included, with the transition assumed to occur at the start of the overlapping period.
- **Short temporary jobs:** If one and only one of the jobs lasts for less than a year and the other one is at least three times as long, the longer spell is considered the main job and included in the sample.
- **Part-time jobs:** If one of the jobs is full-time whereas the other one is part-time, the full-time spell is considered the main job and included in the sample

The three criteria are hierarchical, i.e. we first check for transitional overlap, then for temporary jobs, then for part-time jobs.

Multiple spans with the same employer Spans are records, i.e. within-job observations. In the case of overlap between multiple spans, the outcome - job mobility - is unaffected, and the only question is which values of time-varying covariates are valid at which point in time. Pairs of these spells were split. The span created from the overlapping spans has the covariates of the two original spans if they are non-contradictory. Otherwise, the covariate is set to missing. In the case of conflicting wage information, if the difference is less than 5%, the mean is used. [28]

A.3 Full Cox models referred to in the body of the paper

[28] Browsing the data where spells overlap suggest that while some probably refer to changing wages, others appear to refer to bonuses instead, which would imply that the true wage is the sum of both recorded wages. Separating the two cases would involve (more) arbitrary cut-offs. Since less than 1 % of spells are affected, so no attempt at this is made. In the rare case of triple or greater multiple overlaps which only affects about 1 in 2000 spells, the overlapping portions were dropped without any corrections to the covariates.

Table A1: Estimation by region type: Cox partial likelihood model of exits from a job, stratified by individual

	Women		Men	
...				
First child	-.317***	(.0307)		
Second child	-.016	(.0225)		
Youngest > 12 yrs	-.168***	(.0408)		
Wage	-.0271***	(4.5e-04)	-.0233***	(3.5e-04)
Squared	8.9e-05***	(2.6e-06)	6.9e-05***	(1.9e-06)
Distance	.004***	(2.6e-04)	.0034***	(1.8e-04)
Child × Wage	.0046***	(8.1e-04)		
Squared	-4.6e-06	(5.0e-06)		
Child × Distance	.0012*	(5.4e-04)		
2nd child × Wage	-.0032***	(6.3e-04)		
Squared	1.3e-05**	(4.2e-06)		
2nd child × Distance	1.0e-04	(3.6e-04)		
Older child × Wage	.0036**	(.0011)		
Squared	-2.8e-05***	(7.0e-06)		
Older child × Distance	-.0024**	(7.4e-04)		
Urban × Wage	-5.8e-04	(5.7e-04)	-.0018***	(4.3e-04)
Squared	4.8e-06	(3.4e-06)	3.7e-06	(2.3e-06)
Rural × Wage	-.0012	(7.7e-04)	-.0045***	(5.5e-04)
Squared	1.4e-05**	(4.8e-06)	3.0e-05***	(3.0e-06)
Urban × Distance	7.2e-04*	(3.2e-04)	7.6e-04***	(2.3e-04)
Rural × Distance	4.0e-05	(3.7e-04)	5.6e-04*	(2.6e-04)
Urban × Child × Wage	1.3e-04	(9.9e-04)		
Squared	8.3e-06	(6.2e-06)		
Rural × Child × Wage	.0029*	(.0013)		
Squared	-5.1e-06	(8.4e-06)		
Urban × Older child × Wage	-.0015	(.0013)		
Squared	2.9e-08	(8.6e-06)		
Rural × Older child × Wage	-.0022	(.0017)		
Squared	2.2e-06	(1.1e-05)		
Urban × Child × Distance	.0014*	(6.3e-04)		
Rural × Child × Distance	.0018*	(7.0e-04)		
Urban × Older child × Distance	3.9e-04	(8.7e-04)		
Rural × Older child × Distance	.0012	(9.5e-04)		
...	...	...	...	...
Observations	6,433,713		6,876,548	
Jobs	2,435,009		2,679,887	
Persons	968,607		1,027,065	

Stratified Cox partial likelihood model, additional controls: unskilled occupation dummy, occupational field, regional structure (9 types), establishment size (5 dummies for the 10th, 25th, 50th, 75th and 90th percentile, respectively), local unemployment rate, local GDP growth. Zero work experience, and zero distance in each type of region (urban to rural), are captured by separate dummies. In this baseline estimation sample, the birth of children cannot be observed in the employment history of men. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table A2: Estimations by part-time status: Cox partial likelihood model of exits from a job, stratified by individual: Women

Age	-.00923**	(.003)
Square root	-.05258	(.03507)
Second child	-.03447	(.02261)
2nd child × Wage	-.00215***	(.00064)
2nd child × Wage squared	5.6e-06	(4.2e-06)
2nd child × Distance	.00025	(.00036)
First child	-.4602***	(.0226)
Youngest > 12 yrs	-.1077***	(.02846)
Wage	-.02772***	(.00032)
Squared	9.3e-05***	(1.8e-06)
Distance	.00439***	(.00015)
Child × Wage	.00983***	(.00058)
Child × Wage squared	-3.2e-05***	(3.6e-06)
Child × Distance	.00268***	(.00033)
Older child × Wage	-.00054	(.00076)
Older child × Wage squared	-5.7e-06	(4.7e-06)
Older child × Distance	-.00269***	(.00044)
Part time	-.1765***	(.01981)
PT × Child	.2285***	(.03033)
PT × Older child	-.04989	(.03714)
PT × Wage	-.0024***	(.00063)
PT × Wage squared	4.1e-05***	(4.8e-06)
PT × Distance	-.00013	(.00028)
PT × Wage × Child	-.01117***	(.00095)
PT × Wage squared × Child	5.2e-05***	(7.0e-06)
PT × Distance × Child	-.00057	(.00049)
PT× Wage × Older child	.00661***	(.00114)
PT× Wage squared × Older child	-3.0e-05***	(8.2e-06)
PT × Dist × Older child	.00197**	(.00061)
...	...	...
Observations	6,433,713	
Jobs	2,435,009	
Persons	968,607	

Stratified Cox partial likelihood model, additional controls: unskilled occupation dummy, occupational field, regional structure (9 types), establishment size (5 dummies for the 10th, 25th, 50th, 75th and 90th percentile, respectively), local unemployment rate, local GDP growth. Zero work experience, and zero distance in each type of region (urban to rural), are captured by separate dummies. In this baseline estimation sample, the birth of children cannot be observed in the employment history of men. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table A3: Baseline estimation, married couples' sample: Cox partial likelihood model of exits from a job, stratified by individual

	Women		Men	
Age	-.0168*	(.0079)	.0081	(.012)
Square root	.0868	(.0958)	-.534***	(.151)
Full time	.228***	(.0073)	.222***	(.0209)
First child	-.25***	(.0382)	.266***	(.0763)
Second child	-.112**	(.0386)	.0717	(.0878)
Youngest > 12 yrs	-.0785*	(.0323)	.0542	(.071)
Wage	-.0267***	(7.2e-04)	-.0282***	(.001)
Squared	1.0e-04***	(4.4e-06)	8.7e-05***	(5.3e-06)
Distance	.0055***	(3.8e-04)	.005***	(4.4e-04)
Child × Wage	-2.0e-04	(1.0e-03)	-6.8e-04	(.0015)
Child × Wage squared	2.6e-05***	(6.4e-06)	-3.7e-06	(7.4e-06)
Child × Distance	.0022***	(5.6e-04)	.0013*	(6.4e-04)
2nd child × Wage	-.0034**	(.0011)	.0022	(.0016)
2nd child × Wage squared	8.8e-06	(7.4e-06)	-1.3e-05	(7.8e-06)
2nd child × Distance	-2.7e-04	(6.6e-04)	-.0016*	(7.1e-04)
Older child × Wage	.004***	(9.7e-04)	3.3e-04	(.0013)
Older child × Wage squared	-4.7e-05***	(6.7e-06)	-7.3e-06	(6.0e-06)
Older child × Distance	-.0017**	(5.8e-04)	-.0011	(5.8e-04)
Observations	1,938,346		1,489,065	
Jobs	680,118		570,418	
Persons	273,709		269,792	

Stratified Cox partial likelihood model, additional controls: unskilled occupation, occupational field, regional structure (9 types), establishment size (5 dummies for the 10th, 25th, 50th, 75th and 90th percentile, respectively), local unemployment rate, local GDP growth. Zero distance in each region-type (urban to rural) and zero work experience are captured by separate dummies. In this married couple's sample, the birth of a child identified in the employment history of a married women is transferred to her husband. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table A4: Interaction between own and spouse's wage and commuting distance, married couples' sample: Cox partial likelihood model of exits from a job, stratified by individual

	Women		Men	
Age	-.0111	(.0079)	-1.4e-04	(.0122)
Square root	.034	(.0961)	-.487**	(.153)
Full time	.231***	(.0073)	.217***	(.021)
First child	-.264***	(.0384)	.287***	(.0769)
Second child	-.0942*	(.0388)	.135	(.0883)
Youngest > 12 yrs	-.0409	(.0325)	.0588	(.0714)
Wage	-.027***	(9.0e-04)	-.0301***	(.0012)
Squared	8.9e-05***	(7.1e-06)	9.3e-05***	(7.7e-06)
Distance	.006***	(8.2e-04)	.002	(.0011)
Child × Wage	9.0e-04	(.001)	.0014	(.0015)
Child × Wage squared	1.3e-05*	(6.5e-06)	-1.4e-05	(7.5e-06)
Child × Distance	.0022***	(5.6e-04)	8.0e-04	(6.5e-04)
2nd child × Wage	-.0042***	(.0011)	.0018	(.0017)
2nd child × Wage squared	1.6e-05*	(7.6e-06)	-1.1e-05	(7.9e-06)
2nd child × Distance	-2.7e-04	(6.6e-04)	-.0019**	(7.2e-04)
Older child × Wage	.0023*	(9.8e-04)	1.9e-05	(.0013)
Older child × Wage squared	-3.3e-05***	(6.8e-06)	-5.1e-06	(6.0e-06)
Older child × Distance	-.0017**	(5.8e-04)	-.0012*	(5.9e-04)
Partner commutes: Less × own distance	-3.9e-04	(7.8e-04)	.0034**	(.0011)
More × own distance	-.004***	(9.1e-04)	-.0227***	(.0013)
Partner earns: Less × own wage	-.0026***	(5.6e-04)	-.0011	(6.8e-04)
More × own wage	-.004***	(5.4e-04)	.0052***	(7.3e-04)
Less × own wage squared	1.0e-05	(5.8e-06)	6.6e-06	(5.6e-06)
More × own wage squared	6.7e-05***	(5.7e-06)	1.6e-05**	(6.2e-06)
Observations	1,938,346		1,489,065	
Jobs	680,118		570,418	
Persons	273,709		269,792	

Stratified Cox partial likelihood model, additional controls: unskilled occupation, occupational field, regional structure (9 types), establishment size (5 dummies for the 10th, 25th, 50th, 75th and 90th percentile, respectively), local unemployment rate, local GDP growth. Zero distance in each region-type (urban to rural) and zero work experience are captured by separate dummies. In this married couple's sample, the birth of a child identified in the employment history of a married women is transferred to her husband. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

A.4 Sensitivity analyses regarding censoring, stratification and relocations

Table A5: Early censoring and non-stratified specification of Cox partial likelihood model of exits from a job: Women

	Censored in 2011		Non-stratified	
Age	-.0212***	(.0038)	.0034*	(.0016)
Square root	-.0927*	(.0445)	-.147***	(.0185)
Full time	.221***	(.0045)	.207***	(.0024)
First child	-.352***	(.0215)	-.445***	(.0111)
Second child	.051	(.0275)	.0025	(.0136)
Youngest > 12 yrs	-.0928***	(.0245)	-.0548***	(.0124)
Wage	-.0262***	(3.3e-04)	-.0234***	(1.5e-04)
Squared	8.6e-05***	(2.0e-06)	9.0e-05***	(8.6e-07)
Distance	.0044***	(1.6e-04)	.0051***	(8.4e-05)
Child × Wage	.0052***	(5.7e-04)	.0084***	(3.0e-04)
Child × Wage squared	-2.4e-07	(3.7e-06)	-2.4e-05***	(1.9e-06)
Child × Distance	.0022***	(3.2e-04)	.002***	(1.9e-04)
2nd child × Wage	-.0034***	(7.7e-04)	8.4e-05	(4.0e-04)
2nd child × Wage squared	1.2e-05*	(5.1e-06)	-2.3e-06	(2.6e-06)
2nd child × Distance	4.3e-04	(4.4e-04)	-7.4e-04**	(2.5e-04)
Older child × Wage	.002**	(7.0e-04)	8.8e-04*	(3.6e-04)
Older child × Wage squared	-2.7e-05***	(4.6e-06)	-1.6e-05***	(2.3e-06)
Older child × Distance	-.0016***	(3.9e-04)	-.002***	(2.3e-04)
Vocational education			-.215***	(.003)
...	...	...	...	
Observations	5,221,072		5,249,126	
Jobs	1,996,493		1,855,318	
Persons	876,896		830,117	

Stratified Cox partial likelihood model, additional controls: unskilled occupation dummy, occupational field, regional structure (9 types), establishment size (5 dummies for the 10th, 25th, 50th, 75th and 90th percentile, respectively), local unemployment rate, local GDP growth. Zero work experience, and zero distance in each type of region (urban to rural), are captured by separate dummies. In this baseline estimation sample, the birth of children cannot be observed in the employment history of men. Standard errors in parentheses. Due to the censoring or missing data in the larger set of covariates the number of observations are reduced compared to Table 3. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

We present results from a non-stratified model in Table A8 (Panel A) for comparison with the baseline model (Table 4). The stratified model is our preferred specification. However, we provide this comparator to show that the differences are quantitatively important and to provide some evidence on the direction of bias introduced by unobserved heterogeneity, which is a priori unclear. The interaction of the baseline and the parametric component of the

Table A6: Early censoring and non-stratified specification of Cox partial likelihood model of exits from a job: Men

	Censored in 2011		Non-stratified	
Wage	-.0229***	(2.5e-04)	-.0286***	(1.3e-04)
Squared	6.9e-05***	(1.3e-06)	1.0e-04***	(6.2e-07)
Distance	.0039***	(1.1e-04)	.004***	(6.4e-05)
Age	.0632***	(.0036)	.0267***	(.0015)
Square root	-1.25***	(.0412)	-.342***	(.0173)
Full time	.152***	(.0074)	.239***	(.0048)
Vocational education			-.195***	(.0026)
...	...	...	...	
Observations	5,547,244		5,532,232	
Jobs	2,205,614		2,005,685	
Persons	932,028		880,146	

Stratified and non-stratified Cox partial likelihood model, additional controls: regional structure (9 types), establishment size (5 dummies for the 10th, 25th, 50th, 75th and 90th percentile, respectively), local unemployment rate, local GDP growth. Zero distance in each region-type (urban to rural) and zero work experience are captured by separate dummies. The non-stratified estimation also includes educational level, a full set of year dummies and dummies for (groups of) nationalities. Standard errors in parentheses. In this estimation sample, the birth of children cannot be observed in the employment history of men. Due to the censoring or missing data in the larger set of covariates the number of observations are reduced compared to Table 3. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

hazard in the Cox model is multiplicative. Therefore, unobserved heterogeneity attenuates covariate effects through dynamic sorting *if observed and unobserved determinants of the hazard are uncorrelated* (Ridder and Tunalı 1999, and the references therein). We add educational attainment, year dummies and dummies for groups of nationalities to this model, effects of which are absorbed by the individual-level baseline hazard in our main model. The estimated effect of wages on the hazard to leave a job is consistent with attenuation bias (see Appendix Tables A5 and A6). In contrast, failure to account for unobserved heterogeneity generates an upward bias in the effects of commuting distance and first birth, and consequently, the estimated marginal willingness to pay for women. Willingness to pay is also biased upward for men, but the impact is much smaller.

This indicates that, unsurprisingly, unobserved heterogeneity is correlated with motherhood, commuting distance and job mobility. For example, part of the effect of the first birth on the transition hazard is explained by unobserved heterogeneity: Giving birth makes a job transition less likely, but women who give birth would also have experienced lower job

Table A7: Cox partial likelihood model of exits from a job, censored if non-employment for more than 30 days is observed stratified by individual

	Women		Men	
Age	-.0426***	(.004)	-.184***	(.004)
Square root	.473***	(.0477)	2.07***	(.0474)
Full time	.0975***	(.0052)	.102***	(.0093)
First child	-.857***	(.0278)		
Second child	.165***	(.0348)		
Youngest > 12 yrs	.255***	(.027)		
Wage	-.0249***	(4.0e-04)	-.0249***	(3.2e-04)
Squared	6.3e-05***	(2.3e-06)	5.9e-05***	(1.6e-06)
Distance	.0059***	(1.8e-04)	.0051***	(1.4e-04)
Child × Wage	.0089***	(7.2e-04)		
Child × Wage squared	-2.1e-05***	(4.6e-06)		
Child × Distance	7.8e-04*	(3.9e-04)		
2nd child × Wage	-.0025**	(9.3e-04)		
2nd child × Wage squared	2.0e-05***	(6.0e-06)		
2nd child × Distance	8.6e-05	(5.3e-04)		
Older child × Wage	-.003***	(7.6e-04)		
Older child × Wage squared	4.7e-06	(4.9e-06)		
Older child × Distance	-6.9e-04	(4.4e-04)		
...				
Observations	6,433,713		6,876,544	
Jobs	2,435,009		2,679,885	
Persons	968,607		1,027,063	

Results are based on baseline estimation of the stratified Cox partial likelihood model using Equation 3, but censored if a nonemployment spell of more than 30 days is observed. Stratified Cox partial likelihood model, additional controls: unskilled occupation dummy, occupational field, regional structure (9 types), establishment size (5 dummies for the 10th, 25th, 50th, 75th and 90th percentile, respectively), local unemployment rate, local GDP growth. Zero work experience, and zero distance in each type of region (urban to rural), are captured by separate dummies. In this baseline estimation sample, the birth of children cannot be observed in the employment history of men. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

mobility ('settled down') in any case.[29] The interaction effect of a child with the wage is also overestimated when unobserved heterogeneity is unaccounted for. This supports our argument that for this application, stratification is a more appropriate technique to deal with unobserved heterogeneity than a frailty method, which would rely on an assumption of uncorrelated observed and unobserved heterogeneity. Note also that in this specification,

[29]This is consistent with the argument on earnings in a recent working paper by Lundborg, Plug, and Würtz Rasmussen (2024).

Table A8: Marginal willingness to pay to reduce commuting (in Euro per day per km): censored and non-stratified specifications

	Low wage		Mean wage		High wage	
A: Censored in 2011						
Childless	.229	(.0084)	.284	(.0103)	.351	(.0128)
One child u12	.471	(.0226)	.64	(.03)	.893	(.0436)
One child over 12	.357	(.025)	.435	(.0298)	.526	(.0365)
2+ children, youngest u12	.428	(.0262)	.576	(.0346)	.793	(.0495)
2+ children, all over 12	.331	(.029)	.406	(.0349)	.495	(.043)
Men	.261	(.0079)	.337	(.0101)	.438	(.0135)
B: Non-stratified specification						
Childless	.313	(.0055)	.414	(.0071)	.555	(.0096)
One child u12	.734	(.0206)	1.05	(.0282)	1.57	(.0454)
One child over 12	.507	(.0202)	.647	(.0248)	.829	(.032)
2+ children, youngest u12	.651	(.0262)	.909	(.0353)	1.33	(.0543)
2+ children, all over 12	.429	(.0268)	.54	(.0327)	.678	(.0413)
Men	.238	(.0039)	.338	(.0055)	.507	(.0083)
C: Censoring by non-employment for more than 30 days						
Childless	.297	(.0099)	.345	(.0112)	.395	(.0128)
One child u12	.528	(.032)	.618	(.0359)	.715	(.0418)
One child over 12	.392	(.0261)	.453	(.0292)	.517	(.0334)
2+ children, youngest u12	.501	(.0399)	.627	(.0484)	.783	(.0618)
2+ children, all over 12	.377	(.0344)	.46	(.0411)	.558	(.0502)
Men	.255	(.0073)	.292	(.0082)	.329	(.0092)

Low and high wage are the 25th and the 75th percentile of overall daily wages by gender, respectively. Standard errors in parentheses. Panel C treats a job spell as censored if followed by a non-employment spell of more than 30 days. In this baseline estimation sample, the birth of children cannot be observed in the employment history of men. For full Cox model estimates, see the Appendix A5, A6, and A7.

men's marginal willingness to pay is 18% lower than childless women's when evaluated at their respective mean wages. This is in contrast to the main specification, where men's willingness to pay is 15% *higher*.

We also estimate a model on the sample censored in 2011, two years before the main sample, to see how sensitive our results are to the censoring pattern (see also Tables A5 and A6). This is a good test in order to investigate whether the assumption of independent censoring is appropriate (see Ridder and Tunalı (1999)). Coefficients in this specification are very similar to our main specification. The differences in estimated willingness to pay are small, and insignificant for all groups of women (see Panel B, Table A8).

Table A9: Cox partial likelihood model of women's exits from a job: specifications to test sensitivity to relocations

	(1)		(2)		(3)	
Age	-.0048	(.0032)	.0097**	(.0035)	.002	(.0033)
Square root	-.522***	(.0372)	-.826***	(.0406)	-.636***	(.0391)
Full time	.206***	(.0039)	.203***	(.0042)	.206***	(.0041)
First child	-.631***	(.0195)	-.541***	(.0209)	-.601***	(.0203)
Second child	-.118***	(.0243)	-.0589*	(.0256)	-.0885***	(.0251)
Youngest > 12 yrs	-.0198	(.0212)	-.003	(.0222)	-.013	(.0218)
Wage	-.0269***	(3.0e-04)	-.0236***	(3.2e-04)	-.0266***	(3.1e-04)
Squared	9.1e-05***	(1.8e-06)	8.2e-05***	(2.0e-06)	9.3e-05***	(1.9e-06)
Distance	.0034***	(1.4e-04)	.0048***	(1.6e-04)	.0043***	(1.5e-04)
Child × Wage	.0086***	(5.1e-04)	.006***	(5.5e-04)	.0082***	(5.4e-04)
Child × Wage squared	-1.7e-05***	(3.3e-06)	-7.5e-06*	(3.6e-06)	-1.6e-05***	(3.5e-06)
Child × Distance	.0018***	(2.8e-04)	.0019***	(3.1e-04)	.0024***	(3.0e-04)
2nd child × Wage	-.0023***	(6.8e-04)	-.0032***	(7.2e-04)	-.0027***	(7.0e-04)
2nd child × Wage squared	9.3e-06*	(4.5e-06)	1.5e-05**	(4.7e-06)	1.1e-05*	(4.6e-06)
2nd child × Distance	8.9e-05	(3.9e-04)	-3.0e-05	(4.2e-04)	6.5e-05	(4.1e-04)
Older child × Wage	.0017**	(6.0e-04)	.002**	(6.4e-04)	.002**	(6.2e-04)
Older child × Wage squared	-2.2e-05***	(4.0e-06)	-2.4e-05***	(4.2e-06)	-2.3e-05***	(4.1e-06)
Older child × Distance	-.0013***	(3.4e-04)	-.0018***	(3.7e-04)	-.0017***	(3.6e-04)
Observations	6,433,713		5,228,317		5,731,503	
Jobs	2,435,009		2,202,806		2,429,885	
Persons	968,607		908,040		967,173	

Standard errors in parentheses, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. (1) anticipates relocations, using the end-of-job distance as the whole job spell's commuting distance. (2) drops all job spells with a relocation. (3) censors at relocation.

Table A10: Cox partial likelihood model of men's exits from a job: specifications to test sensitivity to relocations

	(1)		(2)		(3)	
Age	.0311***	(.0029)	.0154***	(.0031)	.031***	(.003)
Square root	-1.11***	(.034)	-1***	(.0363)	-1.12***	(.0353)
Full time	.184***	(.0064)	.187***	(.0068)	.189***	(.0066)
Wage	-.0243***	(2.2e-04)	-.0229***	(2.4e-04)	-.0245***	(2.3e-04)
Squared	7.6e-05***	(1.2e-06)	7.3e-05***	(1.3e-06)	7.9e-05***	(1.2e-06)
Distance	.0037***	(1.0e-04)	.0042***	(1.1e-04)	.0039***	(1.1e-04)
Observations	6,876,546		5,617,178		6,132,872	
Jobs	2,679,885		2,442,336		2,672,275	
Persons	1,027,063		966,644		1,025,047	

Standard errors in parentheses, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. (1) anticipates relocations, using the end-of-job distance as the whole job spell's commuting distance. (2) drops all job spells with a relocation. (3) censors at relocation.

Table A11: Marginal willingness to pay, sensitivity to relocations

	Low wage		Mean wage		High wage	
(1) Anticipate relocations						
Childless	.176	(.0073)	.221	(.0092)	.278	(.0115)
One child u12	.425	(.0229)	.577	(.0306)	.806	(.0441)
One child over 12	.317	(.0248)	.389	(.0299)	.474	(.0368)
2+ children, youngest u12	.382	(.0272)	.519	(.0364)	.723	(.0523)
2+ children, all over 12	.287	(.0299)	.355	(.0365)	.438	(.0454)
Men	.203	(.0058)	.249	(.007)	.303	(.0085)
(2) Drop spells with relocations						
Childless	.282	(.0098)	.358	(.0123)	.455	(.0158)
One child u12	.583	(.0281)	.812	(.0383)	1.18	(.0597)
One child over 12	.434	(.0298)	.539	(.036)	.667	(.0455)
2+ children, youngest u12	.495	(.031)	.694	(.0429)	1.02	(.0673)
2+ children, all over 12	.367	(.0339)	.467	(.0423)	.595	(.0548)
Men	.245	(.0068)	.301	(.0082)	.367	(.01)
(3) Censor at relocation						
Childless	.225	(.0082)	.286	(.0104)	.363	(.0132)
One child u12	.553	(.0256)	.765	(.0346)	1.1	(.0529)
One child over 12	.42	(.0274)	.524	(.0333)	.652	(.0422)
2+ children, youngest u12	.488	(.0293)	.676	(.0399)	.972	(.0609)
2+ children, all over 12	.372	(.0322)	.469	(.0399)	.593	(.0512)
Men	.214	(.0061)	.265	(.0075)	.325	(.0092)

Standard errors in parentheses, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. (1) anticipates relocations, using the end-of-job distance as the whole job spell's commuting distance. (2) drops all job spells with a relocation. (3) censors at relocation.

A.5 Heterogeneity by housing cost

We add information on housing costs to a subsample of the data (from 2004 to 2013). The confidence intervals for willingness to pay for this restricted sample overlap with the ones for the full sample for the baseline specification (Appendix Table A12). We then allow willingness to pay to differ between more and less expensive rental housing markets by interacting dummies for terciles of housing cost with the wage, and with child indicators for women (Table A14).[30] If workers with a higher willingness to pay to reduce commuting distance also lived in areas with higher housing costs, we could be understating their commuting cost since part of it is paid in the housing market. There is evidence that this is somewhat true for men, although the differences are modest: willingness to pay is €0.31 in the most expensive areas, compared to €0.27 in the most affordable areas. The differences we find for childless women are even smaller and statistically insignificant. Coefficients on the interactions of wages and housing costs with child indicators are imprecisely estimated and differences in willingness to pay for mothers by housing costs are not significant. Overall, there is little evidence that sorting into more or less expensive areas by willingness to pay is an economically important pattern.

[30]Interactions of housing cost terciles with distance are insignificant throughout when added.

Table A12: Estimation by housing cost: Cox partial likelihood model of exits from a job, stratified by individual: women

	Baseline		Full model	
Age	-.0793***	(.0049)	-.0797***	(.0049)
Square root	.35***	(.0572)	.353***	(.0573)
Full time	.209***	(.0048)	.209***	(.0048)
First child	-.326***	(.0245)	-.307***	(.0356)
Second child	.0597*	(.0301)	.0579	(.0302)
Youngest > 12 yrs	-.174***	(.0266)	-.177***	(.0266)
Wage	-.0293***	(3.8e-04)	-.03***	(6.6e-04)
Squared	9.9e-05***	(2.4e-06)	1.0e-04***	(4.3e-06)
Distance	.0046***	(1.8e-04)	.0046***	(1.8e-04)
Child × Wage	.0042***	(6.5e-04)	.0037***	(1.0e-03)
Child × Wage squared	4.3e-06	(4.2e-06)	1.0e-05	(6.6e-06)
Child × Distance	.0023***	(3.6e-04)	.0023***	(3.6e-04)
2nd child × Wage	-.0036***	(8.4e-04)	-.0036***	(8.4e-04)
2nd child × Wage squared	1.5e-05**	(5.5e-06)	1.4e-05*	(5.5e-06)
2nd child × Distance	8.6e-05	(4.7e-04)	9.0e-05	(4.7e-04)
Older child × Wage	.0036***	(7.6e-04)	.0036***	(7.6e-04)
Older child × Wage squared	-3.6e-05***	(5.1e-06)	-3.7e-05***	(5.1e-06)
Older child × Distance	-.0014**	(4.2e-04)	-.0014**	(4.2e-04)
Rent Tercile: 2nd			-.0415	(.0302)
3rd tercile			-.0272	(.0334)
2nd tercile × Child			-.0425	(.0415)
3rd tercile × Child			-.024	(.0459)
2nd tercile × Wage			2.4e-04	(8.1e-04)
2nd tercile × Wage squared			2.7e-06	(5.2e-06)
3rd tercile × Wage			.0013	(8.5e-04)
3rd tercile × Wage squared			-7.3e-06	(5.3e-06)
2nd × Child × Wage			.0013	(.0012)
2nd × Child × Wage squared			-1.1e-05	(7.8e-06)
3rd × Child × Wage			4.4e-04	(.0012)
3rd × Child × Wage squared			-6.8e-06	(8.1e-06)
Observations	4,014,943		4,014,943	
Jobs	1,758,875		1,758,875	
Persons	770,225		770,225	

Stratified Cox partial likelihood model, additional controls: unskilled occupation dummy, occupational field, regional structure (9 types), establishment size (5 dummies for the 10th, 25th, 50th, 75th and 90th percentile, respectively), local unemployment rate, local GDP growth. Zero work experience, and zero distance in each type of region (urban to rural), are captured by separate dummies. In this baseline estimation sample, the birth of children cannot be observed in the employment history of men. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table A13: Estimation by housing cost: Cox partial likelihood model of exits from a job, stratified by individual: men

	Baseline		Full model	
Wage	-.026***	(2.8e-04)	-.0288***	(4.6e-04)
Squared	7.2e-05***	(1.6e-06)	8.2e-05***	(2.5e-06)
Distance	.004***	(1.3e-04)	.004***	(1.3e-04)
Age	-.0228***	(.0047)	-.0214***	(.0047)
Square root	-.432***	(.0539)	-.449***	(.054)
Full time	.161***	(.0074)	.159***	(.0074)
Rent Tercile: 2nd			-.17***	(.0255)
3rd tercile			-.259***	(.0279)
2nd tercile × Wage			.003***	(5.7e-04)
3rd tercile × Wage			.0049***	(6.0e-04)
2nd tercile × Wage squared			-1.0e-05***	(3.1e-06)
3rd tercile × Wage squared			-1.6e-05***	(3.3e-06)
Observations	4,277,887		4,277,887	
Jobs	1,899,285		1,899,285	
Persons	812,973		812,973	

Stratified Cox partial likelihood model, additional controls: unskilled occupation dummy, occupational field, regional structure (9 types), establishment size (5 dummies for the 10th, 25th, 50th, 75th and 90th percentile, respectively), local unemployment rate, local GDP growth. Zero work experience, and zero distance in each type of region (urban to rural), are captured by separate dummies. In this baseline estimation sample, the birth of children cannot be observed in the employment history of men. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table A14: Marginal willingness to pay to reduce commuting (in Euro per day per km), by housing cost

	Low rent		Medium rent		High rent	
Childless	.347	(.0147)	.366	(.0153)	.351	(.0146)
One child u12	.882	(.053)	.899	(.053)	.824	(.0472)
One child over 12	.541	(.0389)	.674	(.115)	.574	(.0881)
2+ children, youngest u12	.772	(.0564)	.785	(.0571)	.728	(.0522)
2+ children, all over 12	.491	(.0455)	.497	(.0461)	.468	(.0433)
Men	.265	(.0087)	.289	(.0095)	.309	(.0103)

Results are based on baseline estimation of the stratified Cox partial likelihood models of Table A12 and A13 and using Equation 3. In this baseline estimation sample, the birth of children cannot be observed in the employment history of men. All willingnesses to pay evaluated at the overall average daily wage by gender. Standard errors in parentheses.