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DISCUSSION PAPER SERIES

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ABSTRACT

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

We estimate female and male workers' marginal willingness to pay to reduce commuting distance in Germany, using a partial-equilibrium model of job search with non-wage job attributes. Commuting costs have implications not just for congestion policy, spatial planning and transport infrastructure provision, but are also relevant to our understanding of gender differences in labour market biographies. For estimation, we use a stratified partial likelihood model on a large administrative dataset for West Germany to flexibly account for both unobserved individual heterogeneity and changes dependent on wages and children. We find that an average female childless worker is willing to give up daily €0.27 per kilometre (0.4% of the daily wage) to reduce commuting distance at the margin. The average men's marginal willingness to pay is similar to childless women's over a large range of wages. However, women's marginal willingness to pay more than doubles after the birth of a child contributing substantially to the motherhood wage gap. A married mixed-sex couple's sample indicates that husbands try to avoid commuting shorter distances than their wives.

JEL Classification: C41, J13, J16, J31, J62

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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1 Introduction

Work-from-home mandates during the COVID pandemic triggered a renewed public discussion about the costs of commuting. By now it is not only recognized that commuting is a negative job attribute but also that there are remarkable gender differences in commuting distances (pre-dating the pandemic). These gender differences are also reflected in our West German sample. We find that women's commuting distances are 18% lower than men's. This is accompanied by the very well known raw gender gap in wages, which in our sample stands at 33% of daily earnings. Qualitatively similar results are found by Manning [2003a] and Petrongolo and Ronchi [2020] for the UK as well as Barbanchon et al. [2021] for France.

By studying the gender differences in the willingness to trade higher wages for a lower commuting distance in West Germany, we investigate the relationship between these two stylized facts. Because workers who are able and willing to commute further have access to more and potentially higher-paid jobs, gender differences in preferences over wages and commuting can contribute to disparities in labour market behaviour and wages between men and women. In particular, these differences in the willingness to pay might be a contributor to the gender and the motherhood wage gaps, the origins of which are subject to an ongoing lively discussion.[1] The opportunity cost of commuting is likely to be higher for women, and in particular mothers and those with other caregiving responsibilities.

Differences in the willingness to pay to reduce commuting distance might not only explain parts of the gender and motherhood wage gap, but can also help evaluate policies designed to address these issues. These include policies directly reducing the need to commute, such as telecommuting schemes and other alternative forms of workplace organisation. Giving recent experiences with widespread working from home, these policies are now easier to implement. Policies that aim at enabling mothers to commute further are relevant as well.

So far only a few papers have studied similar issues. One recent exception is Barbanchon et al. [2021] who utilize administrative information 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 with a specific gender aspect. They

[1] For a general discussion of the gender and motherhood wage gap, see for example Blau and Kahn [2017] and Cortés et al. [2022] and for more specific discussion of the motherhood wage gap Adda et al. [2017], Lundborg et al. [2017] and Kleven et al. [2019].

find that unemployed women value commuting 20% more than unemployed men and this can explain around 10-14% of the residual wage gap. Petrongolo and Ronchi [2020] uses less informative data than Barbanchon et al. [2021], observing realised job changes only. They adjust the job search model of Barbanchon et al. [2021] but are unable to address unobserved heterogeneity, and find a gender gap of 15% in the willingness to pay to reduce commuting. However, as the estimated willingness to pay is very low, the differences hardly contributes to the gender wage gap. Borghorst et al. [2022] use a linear probability model to estimate job-to-job mobility using administrative data from Denmark. Their paper is based on wages that are only observed on a yearly bases and a restrictive set of firm level characteristic, which is why they instrument the wage. Additionally, using a purely hedonic wage model they estimate that 3.6% of the residual gender wage gap is due to differences in compensation for commuting.

Van Ommeren and Fosgerau [2009] use a combination of a job search model with a short-run model of commuting time for a fixed distance with strong functional form assumptions and find non-significant gender differences only. Based on cross-sectional data, Manning [2003b] produced some of the first evidence that mothers' wages react more strongly to commuting distance. Hirsch et al. [2013] and Albanese et al. [2023] also find indirect evidence for the correlation of commuting preferences and the gender wage and employment gap.[2]

Approaching the topic through the lens of urban economics, Gutierrez [2018] find that among mixed-sex married couples in the United States, one-tenth of the gender pay gap (conditional on age and years of education) among childless workers and more than a fifth of the motherhood pay gap are explained by commuting. Their modelling approach emphasises residential location decisions, relying on a monocentric model of the city with a gradient of wages and housing costs.

This paper contributes to the existing literature in a number of ways. Firstly we analyse the marginal willingness to pay to reduce commuting using a rich and long panel of employed West German individuals. Using employed individuals, Gronberg and Reed [1994] and Van Ommeren et al. [2000] show that on the basis of a partial job search model the willingness to pay for a job attribute can be directly estimated with the aid of a hazard model of leaving a job. Due to data limitations, specifically the need for a long panel, only a few studies have used this approach in order to study non-wage job amenities (see Dale-Olsen [2006a] for safety, Russo et al. [2012] for commuting with a single employer, Borghorst et al. [2022] for commuting in Denmark).

[2] Earlier exceptions are Rouwendal and Rietveld [1994] and Rouwendal [1999]. However their empirical approach relies heavily on functional form assumptions. The results point to the strong influence of children on commuting preferences.

Our approach requires a different set of assumption than for example Barbanchon et al. [2021] and therefore complements their analysis. Secondly, compared to other high-income countries, the German unadjusted gender and motherhood wage gap [Grimshaw and Rubery, 2015] are exceptionally high, making the issue particularly salient in this context. Thirdly, in an additional analysis we assess how strongly the willingness to pay to reduce commuting depends on a spouse's income and commuting distance. To the best of our knowledge we are the first paper with this particular focus.

In addition, compared to existing papers using the same approach we make a number of methodological improvements and extensions. Firstly, we estimate female and male workers' marginal willingness to pay to reduce commuting distance with a flexible Cox Model and take account of unobserved heterogeneity by using a stratified model (neither Gronberg and Reed [1994], Van Ommeren et al. [2000] nor Dale-Olsen [2006a] could take this approach due to short panels, Borghorst et al. [2022] does not use a flexible harzard rate model). Secondly, we model the willingness to pay very flexibly, i.e. varying over gender, wage, and number and age of children. In this way we do not treat willingness to pay for a job attribute as a single preference parameter that is fixed across individuals and time.[3] Also note that, although we estimate willingness to pay as a preference parameter, this does not exclude the possibility that differences between groups are ultimately rooted in different constraints. One example of this would be gender and parenting norms, in the spirit of Akerlof and Kranton [2000] who argue that preferences may mostly be internalized norms in cases where group identity prescribes certain behaviours.

We find a marginal willingness to pay for childless women of €0.27 to reduce commuting distance by one kilometre at their mean daily wage. Willingness to pay more than doubles after the first birth to €0.63. This can explain a substantial part of the raw gender wage gap, particularly for mothers. In contrast using a married mixed-sex couple's sample, husband's willingness to pay to reduce commuting only slightly increases after a first birth. Moreover, our results indicate that husbands try to avoid commuting shorter distances than their wives.

Our paper connects to a number of other strands of literature, as it considers commuting decisions in the context of a search labour market with a focus on gender and motherhood. The paper is part of the work that recognises that augmenting Mincerian wage regression for job amenities usually results in a substantial downward bias of OLS estimation of worker's marginal willingness to pay due to unobserved worker and firm

[3]See Lundberg [2022] who discusses the malleability of preferences as a major way forward with regards to gender economics.

heterogeneity as well as in the presence of search frictions [Hwang et al., 1998, Gronberg and Reed, 1994]. Job search models start from the central premise that more desirable job attributes decrease the probability of quitting in a model with multi-dimensional jobs [Clark, 2001]. Therefore, information on job transitions allows the econometrician to identify the willingness to pay for a continuous job attribute. Estimating willingness to pay from a search process from unemployment, on the other hand, requires additional information [see Barbanchon et al., 2021].

In the labour search literature, 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 job destruction rates, or in parameters governing exits into non-participation. Our model allows for all these differences. Additionally, 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.

In Bowlus [1997] and Bowlus and Grogan [2009], parameter estimates such as job finding and job destruction rates differ significantly across genders. These differences are heterogeneous – in several instances, they even change sign – across education levels, highlighting the importance of interactions of gender with other determinants of labour market behaviour. We exclude university graduates to create a more homogeneous group in terms of education.[4] We also explore various other types of heterogeneity, analysing the role of housing costs as well as differences between urban and rural areas and between full-time and part-time workers.

Recalling the 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 [2006a] examine gender differences in willingness to pay for attributes including workplace safety, type of tasks, promotion opportunities and different work schedules. They use duration models or conditional logit models, and studied the effect of job attributes on job-to-job mobility as well as mobility between labour market states around childbirth. We control for a large number of additional job characteristics (in addition to wages and distance), for example detailed occupation and industry indicators.

By definition, commuting distance arises as the result of decisions in at least two markets – one in the housing market and one in the labour market. In urban economics, commuting distance is studied as a result of residential location decisions for an exogenous wage and job location. In this tradition, compensation for commuting distance is

[4] Note that only 26% of working-age women in Germany held a tertiary qualification towards the end of our sample period. [OECD, 2014]

available in the housing market through lower house prices. Early models of commuting in urban economics considered a household with one worker, often quietly presumed to be male. In an early contribution, White [1986] uses an OLS regression of commuting time on income and demographic variables, with a focus on gender differences, in particular women's shorter commutes, a highly persistent finding across time and space [for Germany, see Auspurg and Schönholzer, 2013]. However, her approach is unable to account for a number of unobserved differences between households classified as male and female-headed, and is vulnerable to simultaneity bias. Models such as Black et al. [2014] extend the traditional framework to accommodate household decision-making.

Estimating marginal commuting cost from a partial labour market search model such as ours 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 specific job spells. 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 has been shown to increase with educational attainment (ibid and references therein) and our sample excludes university graduates. We therefore work with a relatively immobile sample. In order to additionally take compensation in the form of lower housing prices into account, we control for local rental prices and interact them with wages in a sensitivity analysis.

The studies discussed above, whether in the tradition of labour market search or urban economics, focus on longer-term job and residential location choices. There are two estimation strategies in the choice and transportation economics literatures which estimate the value of travel time directly, without modelling job or residential location choices. On the one hand, stated preference methods study hypothetical choices from survey data. The results exhibit quite striking variations, even when similar models are applied [e.g. Calfee et al., 2001, Small et al., 2005]. It could be the case that the treatment of unobserved heterogeneity drives the stark differences in estimates, in which case the problem would be econometric in nature. Another possibility is that stated preference data could simply be an unreliable signal of underlying preferences [Hensher, 2004].

On the other hand, revealed preference methods analyse observed choices nested within the commuting decision, such as mode, route, or vehicle choices. For example, Brownstone et al. [2003] and Lam and Small [2001] both estimate a high marginal willingness to pay for commuting time of more than 70% of the wage rate. They circumvent the problem of biased reporting of willingness to pay, but data with sufficient variation of alternatives is not readily available, and the interpretation often extrapolates far beyond the range of commuting times actually observed. Older studies using revealed preference data also relied on strong assumptions on the shape of the utility function, as Van Ommeren et al. [2000] point out.

Additionally, our results for married mixed-sex couples speak to the literature on breadwinner norms [Bertrand et al., 2013], which highlights the costs couples are willing to incur to avoid a situation where a wife out-earns her husband.

The paper is structured as follows: The first section specifies a partial-equilibrium model of job search with jobs characterised by a wage and a commuting distance. Particular reference is made to differences by gender and the impact of the regional labour market situation, and a tractable estimator of marginal willingness to pay to reduce commuting distance is derived, following Van Ommeren et al. [2000]. We then present and discuss an estimate of marginal commuting cost using a Cox model on an administrative linked employer-employee dataset. We then present result based on a sample of married mixed-sex couples, which we use to study the association between willingness to pay and a partner's wage and commuting distance.

2 A Model of Job Search with Commuting

In this section, we will outline an on-the-job search model extended to two-dimensional jobs, closely following Van Ommeren et al. [2000]. Without additional information, for example on rejected job offers, marginal willingness to pay for non-binary job attributes cannot be recovered from search from unemployment. Voluntary job transitions, on the other hand, do identify this parameter.

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 (see 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 whole of the life course is

$$\rho R(w,d) = u(w,d) + \theta \int\int \max\{0, R(w^\star, d^\star) - R(w,d)\} dF(w^*, d^*) \\ + \delta(U - R(w,d)) \tag{1}$$

where ρ is a discount parameter and U is the expected present value of unemployment. Lifetime utility is thus 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:

$$\frac{\frac{\partial\theta(w,d)}{\partial d}}{\frac{\partial\theta(w,d)}{\partial w}} = \frac{\frac{\partial\theta(w,d)}{\partial u(w,d)}\frac{\partial u(w,d)}{\partial d}}{\frac{\partial\theta(w,d)}{\partial u(w,d)}\frac{\partial u(w,d)}{\partial w}} = \frac{\frac{\partial u(w,d)}{\partial d}}{\frac{\partial u(w,d)}{\partial w}} \tag{2}$$

Commonly the job attribute that is investigated creates positive utility, which is naturally not the case with our attribute, commuting distance. For ease of interpretation, we therefore choose to focus on the marginal willingness to pay to *reduce* commuting distance which is defined as:

$$-\frac{\frac{\partial\theta(w,d)}{\partial d}}{\frac{\partial\theta(w,d)}{\partial w}} = -\frac{\frac{\partial u(w,d)}{\partial d}}{\frac{\partial u(w,d)}{\partial w}} \tag{3}$$

Thus for our purposes, the negative of the ratio of the marginal derivatives of the hazard rate gives the amount of wage that a worker is willing to give up in case commuting is reduced by one km.

Regional Labour Market Conditions As an extension to their basic model, Van Ommeren et al. [2000] discuss the inclusion of business cycle effects in the model.[5] 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. We have also experimented with indices counting regular employment relations in the individual's county

[5] Van Ommeren et al. [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.

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. 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.

Gender As explained in the introduction, 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, 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.[6] 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 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. Notable sources of such heterogeneity could be the job offer arrival rate and the distribution of wages and commuting distances offered.

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 is that women's non-market time is more productive than men's, a classic assumption ever since Becker [1981]. This could be the case because they remain responsible for the bulk of household and child-rearing tasks. 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 commuting time would be higher for women than for men and for mothers than for non-mothers. Other possible 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. In light of these considerations we allow for the interpretation that differences in preferences might ultimately be rooted in different constraints.

[6] The marginal willingness to pay to reduce commuting is commonly modeled 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.

3 Empirical Analysis

3.1 Data

Our data comes from administrative social security records[7] which are more accurate than survey datasets commonly used in studies of commuting. For instance, they avoid problems of recall error in job spell durations and biased self-reporting of wages. We start with a 10% sample of all individuals with a national insurance number, going back to 1975. We then build our analysis sample, consisting of the inflow into regular full- or part-time employment[8] between January 1st, 2000 and December 31st, 2013. Employment spells including the wages are recorded at a workplace level in days. On this basis we can determine the employment duration at the workplace, i.e. the job spell. The observations are treated as censored if the spell last longer than 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 sample to West German workers 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 married mixed-sex couples' sample. To construct this, geo-coded data is used to 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. Details on the matching process can be found in Goldschmidt et al. [2014].

The administrative data records daily wages. However, 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[10], whose wages are more likely to be recorded without top-coding. By excluding the university graduates, we also focus on a geographically immobile sample giving additional justification for the assumption of a exogenous residential location.

Following the majority of the literature, we measure commuting as distance, see for

[7] We use a sample of 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 5).

[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, when women in East Germany often had a first child born before reunification and further children born much later.

[10] Some systematic underreporting of higher education is known to occur in the data. However,(employer-reported) education information during job spells is considered reliable [Fitzenberger et al., 2005]. To minimise bias from underreported education, we smooth education, classifying individuals as university graduates after the first reporting of a university degree, based on one of Fitzenberger et al. [2005]'s correction procedures.

example Barbanchon et al. [2021]. More specifically, our distance variable measures Euclidean distance between postcode area centroids. We argue that in the particular context of Germany, distance and travel time are very closely related. Firstly, Germany's geography generates little heterogeneity in travel time for a given distance. Secondly, we control for regional structure, which captures elements of transport infrastructure. Finally, only 12% of the traffic volume associated with travel to work in Germany used public transport in 2008 [Follmer et al., 2010]. This means that public transport infrastructure is not a central determinant of travel time, again making travel time relatively more homogeneous in space. Additionally, we drop observations with a distance above 100km.

Note, similarly to willingness to pay estimates for other job attributes, our estimates take as given any adaptations that employers and/or workers use to make the attribute distance less onerous, such as the option to work from home [e.g., employers may take measures to mitigate predicted injury hazards both before and after the fact in Dale-Olsen, 2006a]. These adjustment actually seem to be rather minor during our observation window. According to an analysis based on the *Mikrozensus* [Brenke, 2014], only 8% of employees in Germany occasionally or primarily worked from home during our sample period, and the share was highest in high-skilled occupations which normally require a university degree and are not in our analysis. In addition, it is likely that many of the workers in the "occasional work from home" group complete work-related tasks at home outside of working hours, which needn't reduce the number of journeys to their place of work.

One focus of our analysis is the variation of the willingness to pay to reduce commuting distance before, relative to after, the birth of a child. To identify the timing of births in the data, we use a routine due to Müller and Strauch [2017], based on exits from employment into the mandatory part of maternity leave. Particularly for our sample of non-graduate women who are unlikely to have a child before entering the labour market, this is a very reliable way of identifying first births. The way of identifying birth, however, also restricts our focus in the first part of our analysis towards women when it comes to the variation of the willingness to pay after the birth of a child. However, in our household sample we are able to place birth of children into the work history of men and investigate whether men's willingness to pay to reduce commuting varies after childbirth.[11]

[11] Unfortunately, adoptions cannot be identified with this approach.

3.2 Specification and Model Choice

A duration model is the most direct and intuitive empirical implementation of the job search model discussed in Section 2. The daily frequency of our dataset comes close to 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.

In the duration model of job mobility specified in this section, the failure event is a job ending for any reason. These include voluntary job transitions, layoffs and exits from the labour market. Voluntary job transitions and layoffs are not unambiguously distinguishable in our data. We observe job spells that may either be followed by another job spell, or by a spell of missing data which could happen for several reasons, most importantly unemployment. But missing data might also reflect periods of full-time education, other non-participation, time spent abroad and self-employment.

There are two approaches in the literature on how to deal with this fact. The main part of the literature is not able to identify voluntary job transitions unambiguously and ignore the layoff rate under the assumption that involuntary exits are constant and exogenous. Examples start with the seminal paper of Gronberg and Reed [1994] and continue with Dale-Olsen [2006b] and Van Ommeren and Fosgerau [2009]. In contrast, papers such as Van Ommeren et al. [2000] have information on unemployment transitions and treat these transitions as independently censored.

We follow the majority of the literature for two main reasons. In West Germany, layoffs are rare, with an OECD score of job protection of 2.6 (which is higher than the OECD average and much higher than the UK and the US) for regular contracts. Furthermore, we do not have exact information on involuntary job changes. Nevertheless, we conduct a sensitivity analysis where we treat job spells that are followed by non-employment for more than 90 days as censored. In this way, we mirror the approach of Van Ommeren et al. [2000].

A more complete model might 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 than in the present case [Van Den Berg, 2001]. Bonhomme and Jolivet [2009] specify a model where workers are at risk for different events ending a job and let the hazards depend on individual characteristics. However, their model is quite different from ours: they study an objective, continuous latent amenity whose value is compared to a subjective, individual-specific threshold. This threshold varies by observed and unobserved worker characteristics and the comparison determines “good” and “bad” jobs. The authors then estimate this model on categorical attribute data. In contrast, we conceptualise and measure a continuous amenity, which can in principle enter the utility function in any functional form and whose offer distribution is unspecified. Our model can also accommodate a more general form of unobserved heterogeneity.

The Stratified Cox Model Our sample contains multiple job spells per person, enabling us to account for unobserved heterogeneity using the method of Stratified Partial Likelihoods. As Ridder and Tunalı [1999] set out, failure to account for shared unobserved characteristics of a group of spells[12] biases coefficient estimates. The direction of bias is a priori unclear in many applications, including our own.[13]

We go beyond the existing literature to capture these unobserved influences in a very flexible way, under the assumption that they are constant across different jobs spells of the same individual, using Stratified Partial Likelihood estimation. This method allows the baseline hazard to differ across individuals in an arbitrary way. Coefficients are identified using within-individual variation. Whilst OLS estimates of wage premia or linear probability models of job changes have been augmented with fixed effects [e.g. Duncan and Holmlund, 1983, Villanueva, 2007, Borghorst et al., 2022], and duration models have used shared-frailty terms to capture scalar unobserved heterogeneity [Van Ommeren et al., 2000], data limitations have prevented previous work from using this within-person variation in a stratified partial likelihood model. This approach is able to capture any heterogeneity that has the same shape within an individual across jobs and does not require proportionality of baseline hazards of different individuals. This means that unobserved heterogeneity could affect hazards differently at different points in the job spell. In addition we allow variables to change over time, like the presence of children and business cycle effect, which are difficult to capture in OLS approaches.

We rely on exits from a job to identify willingness to pay, meaning that we do not capture the decisions of non-participating women, some of whom may be prevented from entering the labour market by high commuting costs. Assuming that the willingness to pay is even larger in absolute terms for these women, we estimate a lower bound (in absolute terms) for the willingness to pay of all women in the economy.

[12] In our case, all job spells belonging to the same worker.

[13] Coefficients are biased even if unobserved heterogeneity is uncorrelated with the regressors included in the model. If this is the case, and heterogeneity is scalar, then the estimated coefficients will be biased towards zero [Ridder, 1984, Ridder and Tunalı, 1999]. However, if unobserved heterogeneity has an arbitrary form or is correlated with the regressors – as is likely the case in our application – the bias is much more complex.

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 $\mathbf{X}_{ij}(t)$.[14]

Ridder and Tunalı [1999] argue that censoring might not be non-independent under what they call a synchronous observation plan, that is if analysis time returns to zero at the start of each new spell. This is the case in our analysis, where we assume that the baseline hazard is defined in terms of the time elapsed in the current job. The problem could arise because the interaction between censoring at the end of the observation period and the timing of failure in an earlier spell within the same group affects 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 of Covariates The standard Cox model assumes a linear form for the log relative risk, but a number of diagnostic tools are available to determine whether this simple specification fits the data well. There is a clear trade-off in model choice here: 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 explore fractional polynomials to find the best functional form for the wage. This method runs through a pre-determined 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

[14] 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].

$\{-2, -1, -0.5, 0, 0.5, 1, 2, 3\}$. 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].

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, for example the measures of rental housing cost, enter as dummies to ensure flexibility and produce willingness to pay estimates for interpretable groups.

Main Specification 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

$$\theta_i(t, \mathbf{X_t}) = \theta_i(t) \exp(\beta_{w_1} wage_t + \beta_{w_2} wage_t^2 + \beta_d distance_t + \beta_{\mathbf{x}} f(\mathbf{X}(\mathbf{t}))) \tag{4}$$

where the control vector $\mathbf{X}$ includes the worker-level variables of age (linear and squared), and (sets of) dummies for full-time work, unskilled occupation, and occupational field. 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, respectively. 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 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. The level of observation in our data is a span, or national insurance record.[15]

[15] For a stochastic covariate process to be valid in a survival model, it has to be predictable, i.e. determined only by information on the past history of the process itself and its covariates, not their future paths. This could be problematic in the case of the wage, e.g. if workers and employers use inside knowledge of future job offers in wage (re-)negotiations. We focus on workers without a university degree,

The main 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 is higher among women than men. The appendix 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.

3.3 Estimation and Results

Plugging the baseline functional form (4) into equation (3), marginal willingness to pay to reduce commuting is given by

$$MWP = -\frac{\beta_d}{\beta_{w_1} + 2\beta_{w_2}^2 wage_t} \tag{5}$$

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

We find that marginal commuting cost increases 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 childcare work. We derive the information on childbirth from social security records; 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 which could shed additional light on gendered

among whom individual wage negotiations are rare (Hall and Krueger [2012] estimate a "dramatic" positive effect of education level on the probability of individual bargaining for the United States (p. 64)).

allocations of market and non-market work, is not possible with this dataset. However, we return to this question when analysing a matched sample of married couples in Section 3.5. 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. 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 increased specialisation after the birth of the first child, with new mothers specialising in non-market work.

In addition to the impact on willingness to pay for a reduced commuting distance, family composition affects women's job mobility patterns 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 an additional reduction in job mobility, but the impact of distance on the hazard reduces. This is reflected in the marginal willingness to pay for reduce commuting partly bouncing back for mothers of older children, consistent with a reduction in 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 and women have higher rates of job mobility in areas of higher economic growth. Full-time workers of either gender are also more likely to transition, with a bigger effect for women, likely reflecting more career progression for full-time workers. Unskilled jobs, identified using occupation codes, have higher hazard rates. Recall that the effect of any time-invariant worker attributes such as educational attainment at labour market entry will be captured by the individual-level baseline hazard.

The results of Barbanchon et al. [2021] lie in the range of our results but exhibit less strong variation with regard to gender and family status. We find a lower willingness to pay in terms of the daily wage for men, of €0.32 per km at the mean daily wage vs their estimate of €0.43 at their mean wage[17]. In addition, our estimate are higher for particular groups of women. This is especially true, for women with children below the age of 12 years. Here we find a willingness to pay of (€0.63, that is a 150% increase in willingness to pay compared to men at mean wage of women (compared to a 22% increase which they find for all women compared to men, with little variation by family status). Stronger gender norms in West Germany compared to France might play a role here.

[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.

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

Our results are also comparable to earlier results in the Netherlands. For example Russo et al. [2012] estimate €0.49 for a mixed-gender sample of employees of Amsterdam's Vrije Universiteit, or Van Ommeren et al. [2000] estimate 0.4 Guilders or 18 Euro-cents estimated by at the mean wage for men. Here one needs to particularly account for inflation.

The type of settlement structure could affect willingness to pay, in particular via transport infrastructure. There are a wide range of differences between those areas. We therefore 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 Table 5 and 6.

For childless women, there is little variation with regard to settlement structure. Willingness to pay evaluated at the overall mean wage is slightly higher in urban areas than either in core cities or rural areas. Once a child is born the variation in the willingness to pay to reduce commuting becomes large. For all groups of mothers, willingness to pay 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, although the differences are smaller.

A driver of the effect is the interaction between the presence of a child, the regional structure and distance, i.e. the marginal value of a shorter commuting distance increases by more after the first birth in areas outside of core cities. A potential explanation would be differences in the availability of flexible working hours or workplace-based childcare: Core cities offer a greater variety of job bundles. This reduces the need to "triangulate" between home, work and childcare and makes it easier to find a good fit.

A lack of childcare places could also constrain mothers' ability to accept job offers with longer commuting distances that would otherwise exceed the reservation utility. The important role of childcare rationing and cost for female participation is well documented in the literature [Connelly, 1992, Del Boca, 2002, Del Boca and Vuri, 2007]. Unfortunately, we do not have data on the the first years of our time period with regards to childcare provision. However, it is well documented that core cities offer more hours of childcare. For example, a report by the Federal Statistical Office [Statistische Ämter des Bundes und der Länder, 2013] shows that the share of under-threes in full-time public childcare in West Germany is currently highest in the cities of Frankfurt and Heidelberg. Even in these cities, the share was only slightly above 25%, highlighting the sparse provision of full-time childcare for young children in West Germany. The better provision of childcare could therefore explain the lower willingness to pay to reduce commuting in core cities, particularly for children under 12.

As children get older, the regional difference in willingness to pay persists, but is reduced. This could reflect better public transport in core cities, which allows older children to be more independent and rely less on parents to drive them.

We allow for interactions of the wage, distance, presence of children and part-time status in Table 7 and report the resulting marginal willingness to pay separately for part-time and full-time women in Table 8. There are significant differences, but all are driven by differential marginal effects of the wage and its interactions, rather than of 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 keeps the mean wage constant 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 part-time working (€.245) than for full-time working women (€.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 upon childbirth, but less so for part-time working women. This could indicate that reducing commuting and part-time work are substitutable margins of adjustment after childbirth.

3.4 Sensitivity Analysis

For comparison with our baseline stratified model (see Table 3), we estimate a non-stratified model (Tables 9 and 10). 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 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. 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 unobserved heterogeneity is correlated with motherhood, commuting distance and job mobility. Given that preferences over different types of jobs are likely to contribute to sorting into motherhood, this is not surprising. 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 "settled down" and had lower job mobility in any case. 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, 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 9 and 10). 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.

We are able to add information on housing costs to a subsample of the data (the period 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 (Table 12). We then estimate a model that allows for 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 child indicators for women.[18] If those workers with a higher willingness to pay to reduce commuting distance also chose to live in areas with higher housing costs, we could be understating their commuting cost - since part of it is paid in the housing market, not just the labour market. There is evidence that this is somewhat true for men: Although the differences are not large (willingness to pay is €0.31 in the most expensive areas, compared to €0.27 in the most affordable areas), willingness to pay is significantly higher in more expensive areas. The differences we

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

find for childless women are even smaller, as well as statistically insignificant. Mothers' willingness to pay is consistently lower in the most expensive areas compared to those in the middle tercile, but the coefficients on the interactions of wages, child indicators and housing cost indicators are imprecisely estimated, and differences in willingness to pay 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.

Endogeneity of residential location is a concern in our model. Workers may accept a job with a temporarily long commuting distance if they anticipate moving closer to their place of work in the future. This would lead to us understating their willingness to pay to reduce commuting distance. To address this problem, we tested specifications which move all residential moves forward by one year, move them forward all the way to the beginning of the job spell during which they occurred, and which exclude all job spells that include a residential move between postcode areas. This last one is an imperfect solution: On the one hand, it also 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 hand, an *ex-post* fixed residential location is not a sufficient condition for exogeneity. Nevertheless, these specifications exclude the previously described scenario, which would be the greatest cause for concern about biased estimates in this case. Results showed that marginal willingness to pay were similar to the baseline estimate, suggesting that the effect of endogenous residential moves is limited (results available on request).

Following the discussion on the ambiguity of identifying job transition we conducted a sensitivity analysis where we treat job spells that are followed by non-employment for more than 90 days as censored, mirroring the approach of Van Ommeren et al. [2000] (see Table 15). The results with respect to the marginal willingness to pay to reduce commuting for men are very similar. For women we tend to find a higher marginal willingness to pay, except for women with a child aged under 12. Willingness to pay for women after the first birth is lower, resulting in a smaller but still remarkable increase of the willingness to pay upon the first birth (56%). This could reflect the censoring mechanism, which treats jobs of women who exit the labour force for a period as censored. These women are not observed further, and all following job spells are deleted. 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.

3.5 Married Couples' Sample

A worker's choice of a job with a commuting distance and a wage is not taken in isolation: To better understand results such as the increase in women's marginal commuting cost upon the birth of their first child, we would like to be able to analyse choices over jobs at the household level. In and of itself, the data does not allow us to match households. However, a new routine developed at IAB uses geocoded data to identify a subset of the married couples in the data. The algorithm treats two people as a married couple if their (geocoded) 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 will be 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.[19] We work with the universe of married couples identified by this method.

Given the construction of the household sample, differences in descriptive statistics compared to our baseline sample (Table 17 vs. Table 1) are to be expected. Women and men in the household sample are older, are more likely to have children and live in urban areas. Women are more likely to work part-time and have a lower average wage and commute less. Men in the household sample earn a higher wage and commute approximately the same distance, at 13.6 km.

We also find some differences in the willingness to pay compared to the original sample. Evaluated at the respective mean daily wages (which are somewhat 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 (see Table 19). The increase upon childbirth, however, is more moderate and as a result, willingness to pay for mothers of one young child is 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 original sample because the importance of the wage diminishes at the same time, which is not the case in the married women's sample (see Table 18). 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.

[19] Note that all married couples include those where one person is self-employed, a civil servant or out of the labour force and therefore does not appear in the dataset at all, so these couples are impossible to match in this data by definition.

The construction of the married couples' sample also allows us to separately calculate willingness to pay for childless men and fathers. 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 and Figure 3). Evaluated at their mean wage, the increase is just 12%. A second child being born reduce fathers' willingness to pay in the married couples' sample, as does children reaching the age of twelve. For these fathers the marginal willingness to pay becomes again close to the one of childless fathers (of the married couple sample) which is similar to the one of the baseline men's sample. Additionally, it worth pointing out that once two children are above 12 the marginal willingness to pay reduces further in the married men's sample.

In the following specifications we allow the impact of a worker's own daily wage, respectively commuting distance, to depend upon the relative position within the couple. Hereby we want to investigate whether previously found pattern, like avoidance of wifes outearning husband [Bertrand et al., 2013], can also be found in our context of willingness to pay to reduce commuting.

In the specification in Table 20, we allow the impact of a worker's own daily wage to vary depending on whether their spouse earns more or less than they do[20], or whether their spouse was unemployed at the time. Note that 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, e.g., differences in relative permanent earnings potential.

Among both men and women, sensitivity of job mobility to the wage, and thus willingness to pay (Table 21), is highest when they are the lower earner within the couple. 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 (6%) even for men who are the secondary earner in their marriage.

In the specification in Table 22, we allow the effect of workers' own commuting distance to vary depending on how it compares to their spouse's. For women, the impact of a longer commuting distance on job mobility is greatest if their husband does not commute at all (see Table 23), whereas it is mitigated when their husband also commutes, with the total effect smallest if the husband commutes more than his wife. As a result, marginal willingness to pay to reduce commuting distance is highest

[20] This variable is based on annual earnings and thus varies at the annual level, to abstract from very short-term fluctuations

among women whose husbands do not commute, and lowest for women whose husbands commute more than they do.

The pattern is different for men: For men whose wife commutes less than they do, the effect of distance on job mobility is *greater* than for those whose wife does not commute at all; their willingness to pay is actually close to zero. In stark contrast, for men 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 is reflected in willingness to pay, with men whose wives have a longer commuting distance having a high willingness to pay to *increase* their own commuting distance.

Finally, in the specification in Table 24, we allow the impact of *both* the wage and the commuting distance to vary depending on the spouse's relative job characteristics. This largely confirms the previous results (see Table 25): As before, for women, the impact of a longer commuting distance on job mobility is mitigated if their husband commutes more than they do. For men, the effect flips: men with longer commuting distance have a lower hazard of leaving their job if their wife commutes more than they do. This reversal of the distance effect does not occur for women. For women, the resulting willingness to pay is lower when their husband commutes more than they do than if their husband has a shorter commute or none at all. Women who are the primary earner in their marriage also have a lower willingness to pay.

Thus with regard to commuting we find strong evidence for the avoidance of non-traditional household patterns. Women try to reduce commuting in case their partner is commuting less than them, and men are even willing to reduce their wage in order to commuting longer distances, in case their partner commutes more. We find, however, symmetric correlations for both genders when looking at the relative wage position. The results concerning the relative commuting position might be a reflection of traditional breadwinner norms potentially driven by the design of our household sample, consisting of mixed-sex couples that have chosen the traditional joined last name.

4 Conclusion

Our estimates on a large administrative dataset of non-university educated workers in West Germany show evidence for a substantial motherhood gap in marginal commuting cost. In an additional sample of married mixed-sex couples, we find that in 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 a partner's relative wage into account, we find that if their partner earns more than they do, both men and women have a higher willingness to pay to reduce commuting distance which is in line with a mechanism operating via specialization or preferences. However, we do not find such symmetric results with regard to the relative commuting distance. 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 equalization 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, as willingness to pay is very similar over a large range of wages. Other explanations that have been put forward are more likely to be the reason. These include differences in the job offer distribution and wage distribution (potentially indicating discrimination), occupational sorting, or differences in productivity related characteristics. Our approach conditions out job offer and wage distribution while simultaneously controlling for occupation and productivity related characteristics, which means that we can credibly isolate the role of willingness to pay to reduce commuting distance but on the flip side, we cannot pursue this line of reasoning further.

In additional analyses, 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. We also find that the increase in willingness to pay with first birth is smaller but more persistent for part-time working mothers. Considering the housing market, willingness to pay is slightly lower in areas where housing costs are high, but the differences are small and mostly insignificant.

Differences in willingness to pay for job attributes such as commuting distance potentially play an important role in determining motherhood wage gaps. Our measure of willingness to pay is a local one and can only approximate inframarginal differences such as the one between mothers and childless women average wages. Taking this local approximation at face value and extrapolating, our baseline specification suggests that mothers of young children would be willing to give up almost 12% of their daily wage to reduce their commuting distance from the sample mean to zero (based on the mean values of mothers). In contrast, men would only be willing to give up less than five percent of their wage for a change of the same magnitude.

To put the gap in the marginal willingness to pay in the context of the motherhood pay gap, consider a woman employed at the mean mother's wage and mean mother's 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 (as a linear approximation of the willingness to pay at the mean wage) of €0.47, 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. [21] Thus, our empirical results in combination with this back-of-the-envelope calculations indicate that commuting preferences are an important contributor to gender and motherhood wage gaps.

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5 Appendix

5.1 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. 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).

Sample Construction

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 [22], 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).

- 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.

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

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. [23]

[23] 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.

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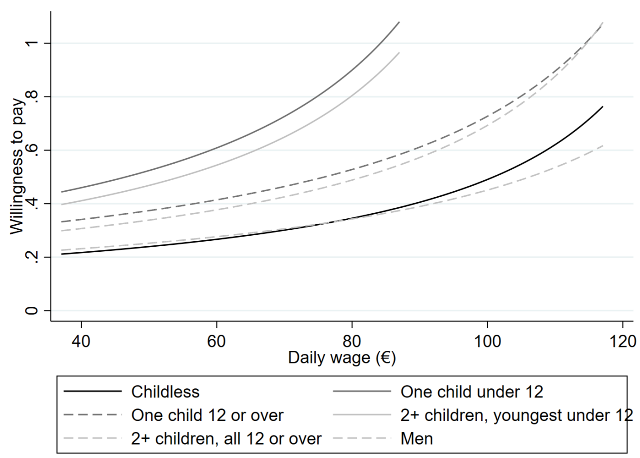
Figure 1: Women's and men's marginal willingness to pay at different wages (in Euro per km)



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

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

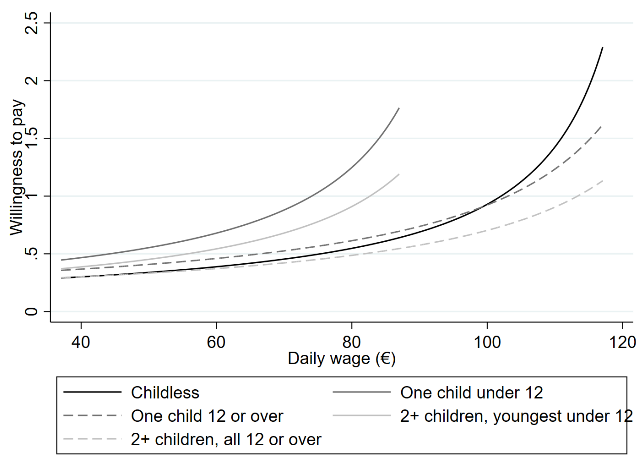
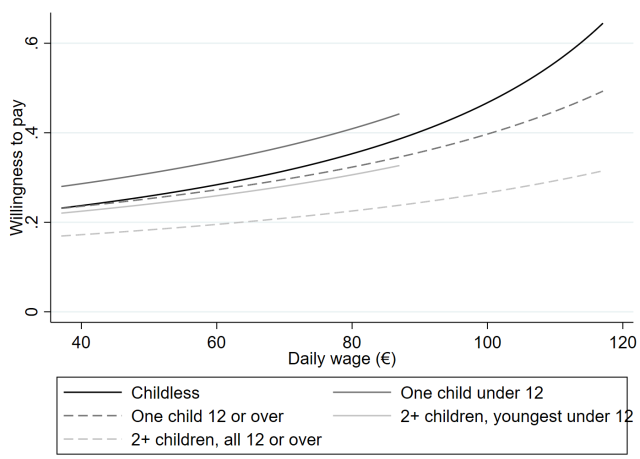


Figure 3: Married men's marginal willingness to pay at different wages



List of Tables

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.

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.

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, 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 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 4: 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 model of Table 3 and using Equation 5. 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.

Table 5: Estimation by region type: Cox partial likelihood model of exits from a job, stratified by individual

	Women		Men	
Age	-.0074*	(.003)	.0553***	(.0028)
Square root	-.0736*	(.035)	-1.02***	(.0324)
Full time	.219***	(.0037)	.167***	(.0059)
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, 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 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 6: Marginal willingness to pay to reduce commuting by region type (in Euro per day per km)

	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)

Results are based on baseline estimation of the stratified Cox partial likelihood model of Table 5 and using Equation 5. 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.

Table 7: Estimations for women by part-time status: Cox partial likelihood model of exits from a job, stratified by individual

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, 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. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 8: Marginal willingness to pay to reduce commuting of women by part-time status (in Euro per day per km)

	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 model of Table 7 and using Equation 5. 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.

Table 9: 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 and nonstratified 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. Due to the censoring or missings in the larger set of covariates the number of observations are reduced compared to Table 3. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 10: 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 nonstratified 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 missings in the larger set of covariates the number of observations are reduced compared to Table 3. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 11: Marginal willingness to pay to reduce commuting (in Euro per day per km): censored and non-stratified specifications

	Low wage		Mean wage		High wage	
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)
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)

Low and high wage are the 25th and the 75th percentile of overall daily wages by gender, respectively. Standard errors in parentheses.

Table 12: 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, 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. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 13: 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, 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 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 14: Marginal willingness to pay to reduce commuting by housing cost (in Euro per day per km)

	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 model of Tables 12 and 13 and using Equation 5. 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.

Table 15: Cox partial likelihood model of exits from a job, censored if non-employment for more than 90 days is observed stratified by individual

	Women		Men	
Age	.0386***	(.0047)	-.0603***	(.0045)
Square root	.991***	(.0557)	2.01***	(.053)
Full time	.08***	(.0058)	.0572***	(.0099)
First child	-.269***	(.0321)		
Second child	.436***	(.0402)		
Youngest $>$ 12 yrs	.0087	(.0302)		
Wage	-.0186***	(4.5e-04)	-.0206***	(3.4e-04)
Squared	4.6e-05***	(2.7e-06)	5.8e-05***	(1.7e-06)
Distance	.0039***	(2.1e-04)	.0036***	(1.5e-04)
Child $\times$ Wage	.0051***	(8.4e-04)		
Child $\times$ Wage squared	-2.5e-06	(5.4e-06)		
Child $\times$ Distance	-9.4e-05	(4.4e-04)		
2nd child $\times$ Wage	-.0017	(.0011)		
2nd child $\times$ Wage squared	1.4e-05*	(6.9e-06)		
2nd child $\times$ Distance	1.5e-04	(6.0e-04)		
Older child $\times$ Wage	4.7e-04	(8.6e-04)		
Older child $\times$ Wage squared	-9.0e-06	(5.6e-06)		
Older child $\times$ Distance	9.0e-04	(4.9e-04)		
Observations	4,955,158		5,272,164	
Jobs	1,888,235		2,095,188	
Persons	960,869		1,014,023	

Results are based on baseline estimation of the stratified Cox partial likelihood model of Table **??** and using Equation 5, but censored if a nonemployment spell of more than 90 days is observed. In this 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.

Table 16: Marginal willingness to pay to reduce commuting when censoring by non-employment for more than 90 days (in Euro per day per km)

	Low wage		Mean wage		High wage	
Childless	.263	(.015)	.304	(.017)	.347	(.019)
One child u12	.382	(.044)	.473	(.053)	.583	(.066)
One child over 12	.461	(.045)	.542	(.051)	.629	(.059)
2+ children, youngest u12	.374	(.056)	.491	(.072)	.651	(.095)
2+ children, all over 12	.449	(.059)	.558	(.072)	.691	(.090)
Men	.259	(.011)	.324	(.014)	.407	(.018)

Results are based on baseline estimation of the stratified Cox partial likelihood model of Table **??** and using Equation 5, but censored if a nonemployment spell of more than 90 days is observed. In this 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.

Table 17: 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 on 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.

Table 18: 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 19: 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 18 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.

Table 20: Interaction between own and spouse's wage, married couples' sample: Cox partial likelihood model of exits from a job, stratified by individual

	Women		Men	
Age	-.0103	(.0079)	-3.1e-05	(.0121)
Square root	.0227	(.0961)	-.486**	(.152)
Full time	.231***	(.0073)	.223***	(.0209)
First child	-.264***	(.0384)	.306***	(.0767)
Second child	-.0939*	(.0388)	.134	(.0881)
Youngest > 12 yrs	-.0399	(.0325)	.0542	(.0712)
Wage	-.0266***	(8.8e-04)	-.0294***	(.0012)
Squared	8.7e-05***	(7.0e-06)	8.9e-05***	(7.6e-06)
Distance	.0056***	(3.8e-04)	.0052***	(4.5e-04)
Child × Wage	9.0e-04	(.001)	.001	(.0015)
Child × Wage squared	1.3e-05*	(6.5e-06)	-1.2e-05	(7.5e-06)
Child × Distance	.0022***	(5.6e-04)	.0014*	(6.4e-04)
2nd child × Wage	-.0042***	(.0011)	.0017	(.0016)
2nd child × Wage squared	1.6e-05*	(7.6e-06)	-1.1e-05	(7.8e-06)
2nd child × Distance	-2.5e-04	(6.6e-04)	-.0018*	(7.1e-04)
Older child × Wage	.0023*	(9.8e-04)	4.0e-05	(.0013)
Older child × Wage squared	-3.3e-05***	(6.8e-06)	-5.2e-06	(6.0e-06)
Older child × Distance	-.0018**	(5.8e-04)	-.0011	(5.8e-04)
Partner earns: Less × own wage	-.003***	(5.2e-04)	-.0017**	(6.4e-04)
More × own wage	-.0044***	(5.0e-04)	.0047***	(6.9e-04)
Less × own wage squared	1.2e-05*	(5.7e-06)	9.5e-06	(5.5e-06)
More × own wage squared	7.0e-05***	(5.6e-06)	1.9e-05**	(6.1e-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. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 21: Marginal willingness to pay to reduce commuting by spouse's relative earnings

	Partner					
	no recorded earnings		lower earnings		higher earnings	
Women						
Childless	.343	(.0245)	.308	(.0214)	.435	(.031)
One child u12	.562	(.0383)	.494	(.0322)	.744	(.0505)
One child over 12	.39	(.0357)	.349	(.0313)	.504	(.0462)
2+ children, youngest u12	.467	(.0396)	.417	(.0346)	.59	(.0502)
2+ children, all over 12	.326	(.0403)	.295	(.0362)	.405	(.0502)
Men						
Childless	.436	(.0427)	.441	(.0395)	1.43	(.192)
One child u12	.498	(.0438)	.503	(.0406)	1.34	(.148)
One child over 12	.384	(.0405)	.388	(.0388)	.926	(.11)
2+ children, youngest u12	.354	(.0486)	.358	(.0475)	.913	(.142)
2+ children, all over 12	.253	(.0505)	.255	(.0502)	.59	(.123)

Results are based on baseline estimation of the stratified Cox partial likelihood model of Table 20 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.

Table 22: Interaction between own and spouse's commuting distance, married couples' sample: Cox partial likelihood model of exits from a job, stratified by individual

	Women		Men	
Age	-.0179*	(.0079)	.0079	(.0121)
Square root	.103	(.0958)	-.533***	(.151)
Full time	.228***	(.0073)	.217***	(.0209)
First child	-.251***	(.0382)	.246**	(.0766)
Second child	-.113**	(.0386)	.0738	(.088)
Youngest > 12 yrs	-.0796*	(.0323)	.0588	(.0712)
Wage	-.0268***	(7.2e-04)	-.0283***	(.001)
Squared	1.0e-04***	(4.4e-06)	8.8e-05***	(5.3e-06)
Distance	.0071***	(7.1e-04)	-3.2e-04	(9.3e-04)
Child × Wage	-2.0e-04	(1.0e-03)	-3.3e-04	(.0015)
Child × Wage squared	2.6e-05***	(6.4e-06)	-5.4e-06	(7.4e-06)
Child × Distance	.0022***	(5.6e-04)	7.6e-04	(6.4e-04)
2nd child × Wage	-.0034**	(.0011)	.0022	(.0016)
2nd child × Wage squared	8.6e-06	(7.5e-06)	-1.4e-05	(7.8e-06)
2nd child × Distance	-2.8e-04	(6.6e-04)	-.0018*	(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)	-.0012*	(5.8e-04)
Partner commutes: Less × own wage	-.0016*	(6.4e-04)	.0055***	(8.5e-04)
More × own wage	-.0053***	(7.7e-04)	-.0214***	(.001)
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 23: Marginal willingness to pay to reduce commuting by spouse's relative commuting distance

	Partner					
	not commuting		commuting less		commuting more	
Women						
Childless	.488	(.0498)	.368	(.0268)	.12	(.0431)
One child u12	.8	(.0692)	.64	(.0432)	.334	(.0593)
One child over 12	.574	(.0623)	.443	(.0405)	.171	(.0544)
2+ children, youngest u12	.644	(.0633)	.513	(.044)	.259	(.0556)
2+ children, all over 12	.468	(.0606)	.358	(.0447)	.128	(.0545)
Men						
Childless	-.029	(.0842)	.469	(.0422)	-1.95	(.0923)
One child u12	.0349	(.077)	.477	(.0425)	-1.67	(.082)
One child over 12	-.0557	(.0726)	.351	(.0401)	-1.63	(.0761)
2+ children, youngest u12	-.104	(.0802)	.323	(.0493)	-1.75	(.0917)
2+ children, all over 12	-.181	(.0792)	.212	(.0516)	-1.7	(.0907)

Results are based on baseline estimation of the stratified Cox partial likelihood model of Table 22 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.

Table 24: 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. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 25: 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 24 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.