TABLE IV
GENDER EFFECT ON THE RESERVATION WAGE AND MAXIMUM ACCEPTABLE COMMUTE IN THE UNITED STATES

	Log ResW (1)	Log max commute (2)
Female	−0.0889***	−0.258***
	(0.0168)	(0.0365)
Mean: males	\$20.13	46.8 min
Observations	3,662	3,918
R-squared	0.625	0.186

Notes. Sample: Survey of Unemployed Workers in New Jersey (see Krueger and Mueller 2016). The table reports regression coefficients of a female dummy on the log of the hourly reservation wage (column (1)) and on the log of the maximum acceptable commute (column (2)). For the sake of comparability to Table 1 in Krueger and Mueller (2016), the sample is restricted to the first interview of each worker. Controls are the same as in column (3) of Table 1 in Krueger and Mueller (2016) (except for nonpublicly available administrative data on UI and past wage levels). Controls include age groups, education groups, potential experience and its square, marital and couple status, number of children, ethnicity and race, previous household income, spouse employment, savings, liquidity access, previous job characteristics (full-time, tenure, and its square), unemployment duration, severance payments received, stated risk preferences, patience proxy and declaration unit for reservation wages. Survey weights are used. Standard errors are robust. *** $p < .01$.

job seekers. This suggests that focusing on unemployed workers is informative about gender differences in job preferences of the whole working population.

Overall, this section has provided evidence of substantial gender gaps in the reservation wage and reservation commute, as well as similar gaps in accepted commute and wage. All gaps grow wider with age and family size, suggesting that labor supply adjusts differently for men and women over their working life cycle. We hypothesize that these gender gaps are partly driven by gender differences in commute valuation. Women have a higher willingness to pay for a shorter commute than men, which translates into a lower reservation wage and commute and results ultimately in a lower reemployment wage and commute. In the next section, we provide estimates of the gender differences in commute valuation.

IV. GENDER DIFFERENCE IN COMMUTE VALUATION

The aim of this section is to quantify the gender gap in willingness to pay for a shorter commute. Commute valuation is identified from the joint distributions of the reservation wage and commute and of the accepted wage and commute. This is not straightforward, as it requires assumptions about what job seekers understand when they declare their reservation wage and maximum acceptable commute. We first introduce a job search model that allows us to be explicit and to formalize these choices.

IV.A. A Search Model Where Commuting Matters

We consider a random job search model where commuting matters (Van Den Berg and Gorter 1997). The instantaneous utility of being employed in a job with log-wage $w = \log W$ and commute τ is given by $u(W, \tau) = \log W - \alpha\tau$. The parameter α measures the willingness to pay for a shorter commute and may differ between men and women. This is the key preference parameter we want to identify. It can be thought of as an individual preference/cost parameter or as a reduced-form parameter that is the outcome of household bargaining on gender task specialization.

Job matches are destroyed at the exogenous rate q. While unemployed, workers receive flow utility b and draw job offers at the rate λ from the cumulative distribution function of log-wage and commute H. The job search model admits a standard solution, which is summarized in the following Bellman equation for the unemployment value U:

$$rU = b + \frac{\lambda}{r+q}\int_0^\infty \int_0^\infty \mathbf{1}_{\{w-\alpha\tau > rU\}}(w - \alpha\tau - rU)dH(w, \tau),$$

where r is the discount rate.

Job seekers accept all jobs that are such that $w - \alpha\tau > rU$. For a job next door, that is, when $\tau = 0$, the reservation log-wage is $\phi(0) = rU$. For a commute τ, the reservation log-wage is: $\phi(\tau) = rU + \alpha\tau$. This allows us to define a reservation log-wage curve:

$$\phi(\tau) = \phi(0) + \alpha\tau.$$

The reservation log-wage curve follows the indifference curve in the log-wage/commute plane with utility level rU. Note that the slope of the reservation log-wage curve is the parameter α, so that identifying the reservation curve yields the willingness to pay for a shorter commute. Replacing rU by $\phi(0)$ in the Bellman equation, we obtain the solution for the intercept of the reservation log-wage curve:

$$\phi(0) = b + \frac{\lambda}{r+q}\int_0^\infty \int_{\phi(0)+\alpha\tau}^\infty (w - \phi(0) - \alpha\tau)dH(w, \tau). \tag{1}$$

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This solves the model. For the sake of completeness, we express the average commute and log-wage in the next job, $E(\tau^n)$ and $E(w^n)$:

(2) $$E(\tau^n) = \frac{1}{p}\int_0^\infty \int_{\phi(0)+\alpha\tau}^\infty \tau dH(w,\tau)$$

(3) $$E(w^n) = \frac{1}{p}\int_0^\infty \int_{\phi(0)+\alpha\tau}^\infty w dH(w,\tau),$$

where $p = \int_0^\infty \int_{\phi(0)+\alpha\tau}^\infty dH(w,\tau)$ is the probability of accepting a job offer.

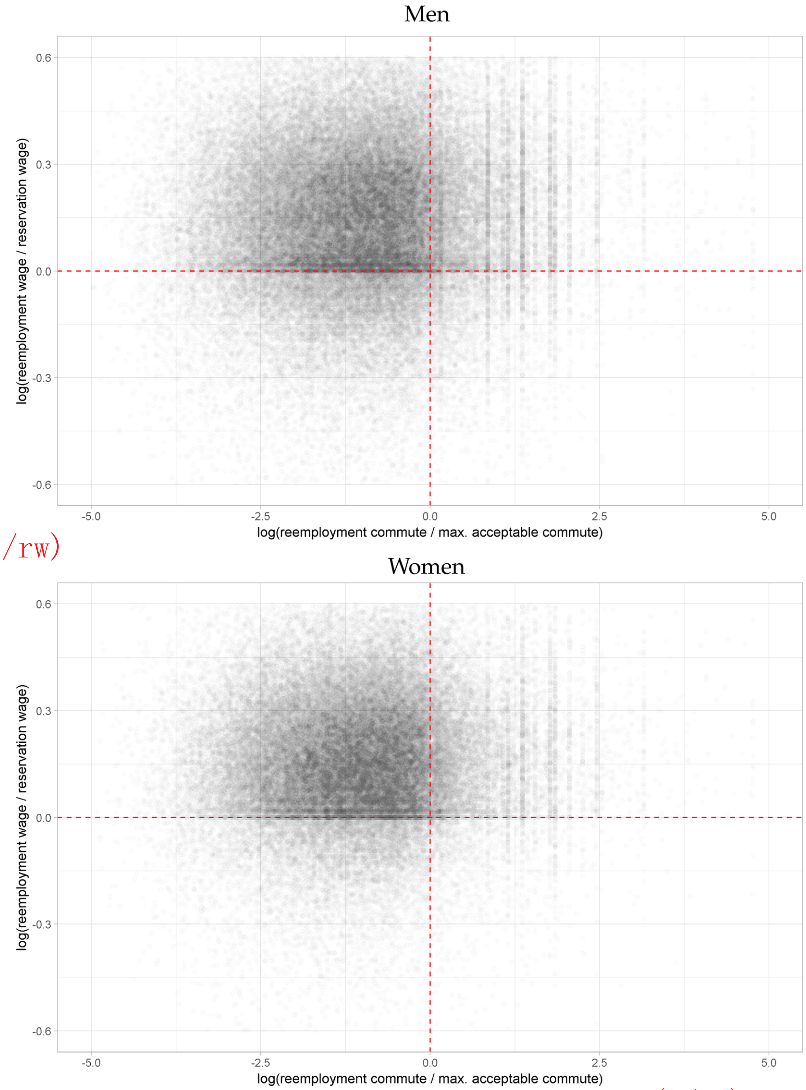
IV.B. *Identifying the Commute Valuation*

To identify the parameter α, the willingness to pay for a shorter commute, we need to relate the search criteria measures to variables in the model. The PES question about the reservation wage does not explicitly anchor the commute dimension. Symmetrically, the question about the maximum acceptable commute does not specify the wage to consider. Without further information, we may consider two main interpretations:

- Interpretation 1: Job seekers answer a pair (τ^*, ϕ^*) of job attributes that lies on their reservation wage curve, so that $\phi^* = \phi(0) + \alpha\tau^*$.
- Interpretation 2: Job seekers report the reservation wage $\phi(0)$ corresponding to the minimum possible commute (0) and the reservation commute $\phi^{-1}(\bar{w})$ corresponding to the largest wage they could get, $\bar{w}$.

Interpretation 2 differs from Interpretation 1 in that it implies that workers do not accept jobs that are both close to their reservation wage and close to their maximum acceptable commute (see Online Appendix Figure C7 for an illustration of these two interpretations). Figure V shows the joint density of reemployment wage and commute, relative to the reservation wage and commute, for men (upper panel) and women (lower panel). By construction, the plot is restricted to workers finding jobs.[12] Consistent with the job search model, most of the density mass is in the upper

12. We convert the maximum commuting time for those who declare in minutes into kilometers, assuming that average commuting speed is 35 km/hour.



ln(wage/rw) >0

ln(d/rd)<0

FIGURE V

Characteristics of Next Job Relative to Search Criteria for Men (Upper Panel) and for Women (Lower Panel)

The figure plots the joint density of the log reemployment wage and commute in deviation from the reservation wage and commute. The vertical dashed red line shows where the reemployment commute equals the maximum acceptable commute. On the horizontal dashed red line, the reemployment wage equals the reservation wage. When job seekers report their reservation commute in minutes, we convert their answers in kilometers using a speed equal to 35 km/h.

left quadrant: workers accept jobs paying more than their reservation wage and closer to home than their reservation commute. Importantly, we do not observe the missing mass predicted by Interpretation 2 in the bottom right corner of the upper left quadrant, where the accepted jobs are both just above the reservation wage and just below the maximum acceptable commute. This is true for both men and women. Figure V provides suggestive evidence in favor of Interpretation 1. We adopt Interpretation 1 in our main analysis, and we provide a robustness analysis under Interpretation 2 in Online Appendix A. In Online Appendix A, we also consider a variant of Interpretation 2 (denoted Interpretation 2 bis), where job seekers report the reservation wage $\phi(\tau_{25})$ corresponding to the first quartile of potential commute and the reservation commute $\phi^{-1}(w_{75})$ corresponding to the third quartile in the potential wage distribution.[13]

To identify the reservation log-wage curve, we leverage the theoretical insight that accepted job bundles are above the reservation wage curve in the commute/wage plane. As a consequence, the frontier of the convex hull of accepted jobs draws the indifference curve delivering the reservation utility. This result holds under some regularity conditions for the job offer distribution. The job offer probability density function must be bounded from below, so there is no region of the commute/wage plane where the acceptance strategy is degenerate and thus less informative.

The identification strategy of the WTP for a shorter commute α proceeds in two steps. First, under Interpretation 1, reservation curves pass through the point where the job bundle equals the declared reservation wage and maximum acceptable commute. This yields one first point of the reservation wage curve. The second step amounts to rotating potential reservation wage curves around the declared reservation job bundle and choosing the reservation curve most consistent with the acceptance strategy of the job search model. We then identify the average slope of the reservation curve by minimizing the sum of squared distances to the reservation curve of accepted bundles that are observed below the reservation curve. We discuss in Section IV.C how classical measurement error and other mechanisms may generate accepted jobs below the reservation wage curve in our data.

Figure VI illustrates the identification strategy. In the log-wage-commute plane, we plot the jobs accepted by 10 workers

13. We thank a referee for suggesting this third interpretation. Note that the argument also makes Interpretation 1 more likely than Interpretation 2 bis.

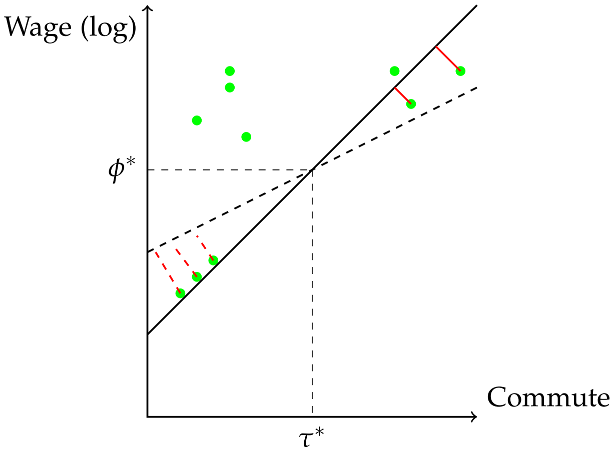


FIGURE VI

Estimation Strategy for the Slope of the Reservation Log-Wage Curve in the Log-Wage-Commute Plane

The figure illustrates the estimation strategy for the slope of the indifference curve in the log-wage-commute plane. Jobs accepted by workers with reported reservation wage ϕ^* and reservation commute τ^* are drawn as green dots. Under Interpretation 1, reservation wage curves go through the (τ^*, ϕ^*) job. We draw two potential reservation wage curves: the solid and the dashed lines. There are three accepted jobs below the dashed line, while there are only two below the solid line. Moreover, jobs below the dashed line are further away from the dashed line (distances in red and dashed) than jobs below the solid line are distant from the solid line (distances in red and solid). Our estimation strategy chooses the solid line as the reservation wage curve.

with the same reported reservation wage ϕ^* and reservation commute τ^*. Under Interpretation 1, the reservation wage curve goes through (τ^*, ϕ^*). We draw two potential reservation wage curves: the solid and dashed lines. There are three accepted jobs below the dashed line, while there are only two accepted jobs below the solid line. Moreover, jobs below the dashed line are further away from the dashed line than jobs below the solid line are distant from the solid line. In practice, the estimator minimizes the number of accepted jobs that are observed below the reservation curve, weighting more the jobs that are further away from the reservation curve. The estimation strategy picks up the solid line. Note that the identification strategy does not require any assumptions on the exact position of the declared reservation job bundle on the reservation curve: it can be anywhere on the curve.

We now define the estimator in formal terms. We denote (τ_i, w_i) the pair of commute and wage accepted by individual i, (τ_i^*, ϕ_i^*) her declared reservation strategy, and $d_{\alpha,\tau_i^*,\phi_i^*}(\tau_i, w_i)$ the distance of the job bundle (τ_i, w_i) to the reservation curve of slope α passing through (τ_i^*, ϕ_i^*). We use as a norm the Euclidean distance between the job bundle and its projection on the reservation line. We further denote $\mathcal{B}_\alpha$ the set of accepted job bundles below the reservation curve ($\mathcal{B}_\alpha = \{i | w_i < \phi_i^* + \alpha(\tau_i - \tau_i^*)\}$). We define the following estimator of the slope α:

$$\hat{\alpha} = argmin_\alpha \sum_{i \in \mathcal{B}_\alpha} p_i \left(d_{\alpha,\tau_i^*,\phi_i^*}(\tau_i, w_i)\right)^2, \tag{4}$$

where p_i are individual weights that we define to make sure that the distribution of covariates of men matches that of women. We compute p_i using inverse probability weighting (Hirano, Imbens, and Ridder 2003). In a first step, we estimate a logit model of being a woman using as covariates the controls X_i from the main gender gap regressions. These include worker characteristics (age, education, family status, and work experience), previous job characteristics (past wage, past commute, part-time, labor contract, and occupation) and fixed effects for past industry, commuting zone, and separation year. Using the estimated logit model, we predict the probability to be a female $\hat{p}(X_i)$. In a second step, we define the weights for men as $p_i = \frac{\hat{p}(X_i)}{1-\hat{p}(X_i)}$. We run the estimation of α separately for women and men.

Last, we restrict the estimation to non–minimum wage workers. The job acceptance strategy of minimum wage workers is degenerate, as there exists a commute threshold such that minimum wage jobs with commute below this threshold yield more than the reservation utility. We select all job seekers declaring a reservation wage at least 5% above the minimum wage. This represents 45.8% of our sample. We verify that our main results from Section III hold in the non–minimum wage workers sample (see Online Appendix Tables D3, D5, and D9). Online Appendix Table D3 shows that the gender gaps in search criteria are similar in this sample, with the gap in reservation wage being one percentage point greater, as expected. We verify the robustness of our results to alternative definitions of the non–minimum wage worker sample.

TABLE V

ELASTICITY OF WAGE WITH RESPECT TO COMMUTE ALONG THE RESERVATION WAGE CURVE

		Without children		With children	
	All (1)	Single (2)	Married (3)	Single (4)	Married (5)
Women	0.148***	0.141***	0.165***	0.148***	0.156***
	(0.0045)	(0.0061)	(0.015)	(0.013)	(0.010)
Men	0.121***	0.111***	0.126***	0.114***	0.141***
	(0.0046)	(0.0053)	(0.014)	(0.013)	(0.010)
Gender gap	0.027***	0.031***	0.039*	0.034*	0.015
	(0.0073)	(0.0072)	(0.020)	(0.018)	(0.015)
Obs.	75,071	38,593	8,670	6,756	21,074

Notes. This table presents estimates of the elasticity of the wage with respect to commute along the reservation wage curve. Estimation minimizes the criteria in equation (4). We restrict the sample to job finders and to non–minimum-wage workers who declare a reservation wage at least 5% above the minimum wage. In column (2), we further restrict the sample to singles without children; in column (3), to married individuals without children; in column (4), to single parents; and in column (5), to married parents. We use inverse probability weighting to balance the covariates of women and men. Bootstrapped standard errors are in parentheses. *** $p < .01$, * $p < .1$.

IV.C. Commute Valuation Estimates

Consistent with Figure V, we consider the log of wages and commutes: we estimate the elasticity along the indifference curve rather than the parameter α directly. Table V presents our elasticity estimates for women in the first row and for men in the second row. The third row shows the gender gap. In column (1), we pool all non–minimum wage workers. The elasticity of wages with respect to commute distance is 0.15 for women and 0.12 for men. The gender gap is positive and statistically significant at the 1% level. This confirms that the disutility associated with commute is larger for women than for men. In columns (2) to (5), we split the sample by marriage status and family size. We find that the elasticity increases slightly with household size, but the gender difference remains around the same level, without any statistically significant differences across subgroups. In Online Appendix Table D12, we report the estimates, separately for the Paris region and the rest of France. The gender gap in commute valuation is smaller in Paris than in the rest of France, but the difference is not statistically significant.

1. Interpreting the Magnitude of the Commute Valuation Estimates. Table V shows that gross monthly wages (FTE) must be increased by 12% to compensate men for a doubling in the commuting distance. Given the average commute of 18.6 kilometers and the average monthly wage of €2,018, an increase of 18.6 kilometers has to be compensated by an increase of €242 (= 0.12 * 2,018) of the monthly wage. The monthly compensating differential for one extra kilometer is about €13. Assuming that full-time employees commute 22 days a month on average (excluding weekends), the daily compensating differential amounts to 59 cents ($= \frac{13}{22}$). How does it compare with the opportunity cost of the time spent commuting? For an increase of one kilometers in the home-work distance, workers spend 3.4 minutes more time commuting per day (assuming an average commuting speed of 35 km/hour). Workers in our sample have an hourly rate of €13.2, which translates into 22 cents a minute. Consequently, the compensating differential for men is 0.8 times the hourly wage ($= \frac{59}{3.4\times22}$). For women, with an elasticity of 14.8%, we obtain a compensating differential of 0.98 times the hourly wage. These estimates of compensating differentials belong to the range of estimates in the literature. Mulalic, Van Ommeren, and Pilegaard (2014) report that estimates of the value of travel time range from 20% to 100% of hourly gross wages (Small 1992; Small, Winston, and Yan 2005; Small and Verhoef 2007; Small 2012).

2. Robustness. In Online Appendix Table D13, we show the robustness of the elasticity estimates to other definitions of minimum wage workers and find similar elasticities and gender gaps in commute valuation. However, when we include minimum wage workers in the estimation sample, the gender gap in commute valuation is significantly lower and statistically significant at the 10% level only. This is expected because minimum wage workers have a degenerate wage offer distribution. Online Appendix Table D14 shows some other robustness tests of the elasticity estimates. Column (1) does not use inverse probability weighting to balance the male and female sample on covariates. Column (2) restricts the sample to workers who declare their maximum commute in kilometers. Column (3) excludes workers with a large deviation between the accepted commute and the reservation commute, for whom nonlinearities are a potential concern. In column (4), we adopt another minimization criteria: the number of accepted bundles below the reservation wage curve (without weighting them by their distance to the curve). Results are robust to these changes of specification. In column (5), we restrict the estimation sample to

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individuals who worked full-time in their previous job. The gender difference in elasticity is smaller when we hold constant the past hours worked, but still significant. This suggests that gender differences in commute valuation come on top of potential gender differences in hours flexibility.

In Online Appendix A, we adopt alternative interpretations of the reported reservation job (ϕ^*, τ^*) (Interpretation 2 and 2 bis). We find again that women have a significantly higher willingness to pay for a shorter commute than men: 23.8% higher under Interpretation 2 and 15.1% higher under Interpretation 2 bis (see Online Appendix Tables A1 and A2).

3. Accepted Job Bundles below the Reservation Wage Curve? Several mechanisms may explain why we observe accepted job bundles below the reservation wage curve. First, it could be due to measurement error in reservation or accepted job attributes. In Online Appendix Table D14, column (6), we add white noise to the data and show that our results are robust to measurement error, with some attenuation bias. This suggests that our main estimate is a lower bound of gender gaps in WTP for a shorter commute. Nonstationarity in job search behaviors is a second possible mechanism, as we pin down the reservation wage curve using reservation job attributes declared at the beginning of the spell. We find that the share of workers who accept jobs that are above their reservation wage curve is 3 percentage points lower for workers who have one more year of unemployment (from an initial share of 83%). This makes duration dependence a marginal contributor to points below the reservation wage curve. The existence of other job amenities is a third possible mechanism. Assuming that workers declare their reservation job attributes conditional on other amenities being at their average, they may accept jobs below the reservation wage curve when amenities are high. As long as the mechanism generating accepted jobs below the reservation wage curve is independent of wage and commute offers, the WTP estimator in equation (4) is still valid, as our simulations related to measurement error suggest. In Section VI.B, we propose an alternative estimation of the gender gap in commute valuation, based on an empirical model of application choice. This approach is robust to unobserved nonwage job amenities that are potentially correlated with wages and commute and provides similar estimates of the gender gap in WTP.

In this section, we have showed that women have a 22% higher willingness to pay for a shorter commute ($\frac{0.027}{0.121} = 0.223$, see Table V column (1)). This result comes from a new (to the best of our knowledge) identification strategy that leverages unique data on job search criteria. The identification strategy mostly relies on the form of the utility function when employed and on the reservation strategy embedded in standard job search models. Specifically, the commute valuation parameter is separately identified from the other model parameters, as long as the job offer distributions are not degenerate. This is worth noting, as an alternative hypothesis supporting the gender gaps documented in Section III could be that men and women do not draw job offers from the same distributions when unemployed (even if they had similar jobs before unemployment). Even in this case, our result on gender differences in willingness to pay for a shorter commute still holds. Next we draw the implications of the gender differences in commute valuation for the gender wage gap.

V. Implications for the Gender Wage Gap

Because women must be compensated more than men to accept far-away jobs, they are more likely to work close to home in jobs that pay relatively less. To what extent do gender differences in commute valuation contribute to the gender wage gap? To quantify this, we first calibrate the job search model above, using the previous estimate of the WTP for a shorter commute. Second, we perform counterfactual simulations where we shock this commute valuation parameter.

V.A. Calibration of the Job Search Model

We calibrate the model, restricting our sample to non–minimum wage workers on which we have estimated the WTP for a shorter commute, α. We proceed as follows.

First, we calibrate r such that the yearly discount rate is 12% (following Van Den Berg 1990) and the match destruction rate q is equal to the inverse of the length of jobs in the subsample of interest (for the median job seeker, a job spell lasts 12 months). Second, we observe in the data the pair (τ^*, ϕ^*), which is a point on the reservation curve, and the previous section yields an estimate of the commute valuation α. We can build the full reservation curve; in particular we deduce $\phi(0) = \phi^* - \alpha\tau^*$.