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Understanding the Congestion Child Penalty: Can Remote Work Attenuate It?

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Abstract

We study the impact of commuting time on the labor supply of women with children in the United States. We show that congestion and long commutes discourage the labor force participation of mothers with low levels of education, while they do not change the probability that college graduate mothers participate in the workforce. We provide evidence that unequal access to remote work can explain the heterogeneous effects of congestion. Because work-from-home arrangements are more prevalent among college-educated workers, they shield highly educated mothers from the adverse impact of congestion on their labor supply. We also show that cohabiting with a partner who has access to remote work is associated with a reduction in the congestion-induced mobility constraints faced by non-college educated mothers. Last, we show that cultural differences across education groups and the lack of affordable childcare options do not explain why long commuting time disproportionately affects the labor force participation rate of non-college graduate mothers.

JEL Codes: R41, J16, J22, J32
Keywords: Congestion, commuting time, child penalty, labor supply, labor force participation, work from home, remote work.

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1 Introduction

The secular rise in women's labor force participation is one of the most striking phenomena that characterized the evolution of labor markets over the course of the 20th Century (Goldin, 2006). In the United States, the labor force participation rate of prime-age women grew from 35% in 1948 to 74% in 1990 (Department of Labor, 2025). Since the late 1990s, however, women's labor supply plateaued (Albanesi, 2023; Blau & Kahn, 2013), contributing to a slow-down in the convergence of labor market outcomes between male and female workers.

As gender convergence stalled over the last few decades, a growing body of literature emphasized the role of persistent gender differences in constraints to workers' labor supply in determining gender imbalances in labor market outcomes (Olivetti, Pan, & Petrongolo, 2024). Men and women, for example, differ in their geographical mobility and in their willingness to commute (Caldwell & Danieli, 2024; Faberman, Mueller, & Şahin, 2017; Le Barbanchon, Rathelot, & Roulet, 2021). Such differences may increase with parenthood, as women are more likely than men to give up non-local employment and to decrease their commute after the birth of a first child (Albanese, Nieto, & Tatsiramos, 2022; Borghorst, Mulalic, & van Ommeren, 2024).

In this context, long commute times represent barriers to women's labor supply whose impact may be shaped by work-family tradeoffs. Using US census microdata, Black, Kolesnikova, and Taylor (2014) and Farré, Jofre-Monseny, and Torrecillas (2023) show that commuting times and congestion negatively affect the labor force participation of married women, and increasingly so when children are present. Furthermore, Moreno Maldonado (2022) shows that long commuting times contribute to explain why, in the United States, the labor force participation rate of women with children is lower in large metropolitan areas than in small cities.[1]

Building upon this literature, in this paper we document and explain why significant heterogeneity exists between mothers with different levels of education in the impact of

[1]Recent work by Kleven, Landais, and Leite-Mariante (2025) also documents that child penalties are stronger in more urbanized areas. The term "child penalty" refers to the worsening in women's labor market outcomes compared to men's following the birth of a child (Kleven, Landais, & Søgaard, 2019). The large body of literature documenting the hampering effect of parenthood on women's labor supply includes, Adda, Dustmann, and Stevens (2017); Angelov, Johansson, and Lindahl (2016); Cortés and Pan (2023); Goldin, Kerr, and Olivetti (2024), among others.

congestion on their labor supply (a phenomenon we refer to as the "congestion child penalty"). First, we show that, while the probability that college graduate mothers participate in the workforce is virtually unaffected by congestion, long commute times have a large negative effect on the labor force participation rate of mothers with no college education. Second, we show that the availability of remote work in jobs typically performed by highly educated workers can explain why the extensive-margins labor supply of college educated mothers is not sensitive to commuting time. Finally, we establish that alternative explanations are unlikely to account for the heterogeneity in the congestion child penalty that we document. In particular, we show that neither differences in cultural attitudes towards gender roles between college and non-college graduate mothers, nor the lack of affordable childcare services explain the disproportionate impact of congestion on the labor supply of non-college educated mothers.

We use data from the American Community Survey (ACS) between 2005 and 2019 to study the effect of congestion on mothers' labor supply, and complement the ACS with data from the American Time Use Survey (ATUS) (2005-2019) to show that access to remote work attenuates the congestion child penalty.

To estimate the impact of congestion on mothers' labor force participation, we exploit the within-year, within-region variation in the probability of participating in the labor force among mothers living in metropolitan areas (MSAs) characterized by different average commuting times. We explicitly account for the possibility that MSA-specific commuting times and women's labor force participation may be jointly determined. To do so, we instrument congestion using two instruments proposed by Duranton and Turner (2011) and originally used to estimate the impact of road supply on traffic: the MSA-specific length of railroads in 1898, and the MSA-specific planned length of highways in 1947. This method is complementary to the IV approaches of Farré, Jofre-Monseny, and Torrecillas (2023), who rely on the shape of cities to instrument commuting times, and of Black, Kolesnikova, and Taylor (2014), who instrument the MSA-specific commuting time through commuting times in women's states of birth.

Our main results show that small increases in commuting time of approximately 3.5 minutes per commute decrease the labor force participation rate of non-college educated mothers by up to 2.6 percentage points, while not affecting the labor force participation rate of college graduate women with children. We corroborate these results through several robustness checks using different identification strategies, empirical specifications, samples, and data.

We then study whether differential access to remote-work arrangements between workers in different education groups can explain the disproportionate effect of long commuting times on the labor force participation rate of non-college graduate mothers. Intuitively, as employees who telework spend less time commuting (Ji, Oikonomou, Pizzinelli, Shibata, & Tavares, 2024; Pabilonia & Victoria Vernon, 2022), remote work arrangements may allow workers with children to reconcile paid work and family responsibilities (Bratsberg & Walther, 2025; Sherman, 2020; Woods, 2020), and to remain in the labor market without paying the monetary and time costs of congestion. Thus, the option of working from home in jobs typically performed by highly educated workers can shield college graduate mothers from the negative impact of congestion on their labor supply. Mobility constraints, instead, may become binding in the presence of children for non-college educated women who seldom have access to remote work arrangements (Albanesi & Kim, 2021), especially in local areas characterized by long commute times.

Using ACS and ATUS data, we show that access to remote work is in fact associated with a reduction in the congestion child penalty. First, we show that, among non-college graduate mothers, the negative impact of congestion on the probability to participate in the labor force is almost twice as large for women cohabiting with partners working jobs with limited access to remote work than for women cohabiting with partners in highly flexible occupations. As we show that men who work remotely dedicate more time to housework and care activities, it is possible that they help ease the mobility constraints that would otherwise limit their female partners' ability to commute and work.[2]

Second, we show that, among college graduate mothers, the probability of participating in the labor force is unaffected by commuting times whether their partners work either flexible or inflexible jobs. We provide evidence that access to remote work likely protects the labor force status of highly educated mothers in congested areas. In fact, we document that college graduate women are themselves more likely than women with lower

[2] It is not surprising to document that men who work from home devote more time to housework and care for other household members. Using German data to test a household labor supply model with gender-specific commuting times, Carta and De Philippis (2018) find that exogenous shocks that increase married men's commuting times lead to a decrease in the time they spend providing childcare.

levels of education to be employed in jobs that can be performed remotely, work more daily hours from home and are more likely to work fully remotely. In addition, college graduate women rely on remote work more intensely (work more hours remotely, and are more likely to work fully remotely), in highly congested metropolitan areas. Finally, in highly congested MSAs, college graduate women spend more time per day working while providing care to other household members as a secondary activity. This suggests that, in highly congested areas, remote work may represent the means through which women with access to this type of flexibility reconcile their paid work with housework and care provision. Among female workers with no college degree, who are mostly employed in inflexible jobs, reliance on remote work does not change with congestion.

These results are key in interpreting the asymmetry in the impact of long commutes on the labor supply of mothers with different levels of education. As mothers without a college degree have limited opportunities to access remote-work arrangements, congestion likely induces some of them to leave the labor force, while exposing labor force participants to the time cost of long commutes, especially if their partners also work inflexible jobs. Access to remote work, instead, may attenuate the work-family tradeoff among college graduate mothers, thus limiting the negative impact of congestion on their extensive-margin labor supply, irrespective of the characteristics of their partners' jobs.

Other mechanisms may explain our findings. First, studying a subsample of migrant workers in the US, Farré, Jofre-Monseny, and Torrecillas (2023) provide evidence that commuting times have especially detrimental effects on the labor supply of women with more conservative cultural attitudes towards gender roles, while Moreno Maldonado (2022) provides evidence suggesting that households with strong preferences for stay-at-home mothers may be more likely to self-select into highly congested areas. If attitudes towards gender roles differ across education groups, and if low-education mothers share increasingly conservative beliefs towards gender roles in more congested areas, then women's cultural background may explain our finding that congestion mostly affects low-education mothers. Implementing an epidemiological approach whereby we proxy US-born women's cultural background through the female labor force participation rate in their state of birth in 1990, we find that mothers of all levels of education with supposedly more conservative beliefs towards gender roles are less likely to participate in the labor force. At the same time, the effect of congestion on women's labor supply does not change with women's cultural background.[3]

Second, Araujo, McBride, and Sandler (2025) find that the lack of affordable market-provided childcare services has negative effects on the labor force participation of mothers in the US, while Cortés and Tessada (2011) show that low-education migrant workers providing services that substitute for home-production increase the intensive-margin labor supply of highly educated women. If childcare services are increasingly unaffordable in highly congested areas, especially for women with low income-potential, then our findings that congestion mostly affects low-education mothers may be explained by the unobserved correlation between childcare affordability and congestion. To study whether this is the case, we use information on the local presence of low-education migrant workers together with information from the National Database of Childcare Prices collected by the Women's Bureau of the Department of Labor, and instrument the local affordability of childcare services using the historical flows of migrants from different countries as in Cortés and Tessada (2011) and Cortés and Pan (2019). Consistent with Araujo, McBride, and Sandler (2025) findings, we show that mothers of all levels of education are less likely to participate in the labor force in areas with more expensive childcare services. Yet, controlling for the affordability of childcare does not change our main results that commuting times primarily affect the labor force participation rate of mothers with no college education.

Our findings contribute to the ongoing debate regarding the possibility that structural changes in the organization of work, and the shift toward remote work induced by the COVID-19 pandemic, may be beneficial for women's labor force participation (Albanesi, 2023; Alon, Doepke, Olmstead-Rumsey, & Tertilt, 2020). Our findings also contribute to the literature studying the multifaceted effects of alternative work arrangements on women's labor market outcomes. On the one hand, the provision of work arrangements that enhance work-life balance may indirectly increase the gender wage gap. As first theorized by Goldin and Katz (2011), several contributions found evidence that gender differences in earnings tend to increase due to the wage gains associated with inflexible and long work hours (Cortés & Pan, 2019; Gicheva, 2013; Goldin, 2014), and due to women's

[3] Studying second-generation women living in the US, Fernández and Fogli (2009) were the first to use women's place-of-birth characteristics to study the effect of cultural traits related to gender norms on their labor market outcomes.

strong willingness to pay for work arrangements such as schedule flexibility (Wiswall & Zafar, 2018) and telework (Maestas, Mullen, Powell, von Wachter, & Wenger, 2023; Mas & Pallais, 2017). On the other hand, benefits and work arrangements that reconcile work-family tradeoffs may contribute to slack constraints to women's labor supply and enhance their labor force attachment, especially in the presence of children. Bloom, Liang, Roberts, and Zhichun (2015) provide evidence that the availability of telework increases workers' satisfaction and decreases their quit rates; Aksoy, Bloom, Davis, Marino, and Ozguzel (2025) show that the adoption of remote-work arrangements can foster women's employment; Bratsberg and Walther (2025) find that remote work can increase fertility, most prominently among women employed in otherwise highly inflexible jobs.

Our results suggest that remote work can alleviate mothers' mobility constraints and reduce the time costs of commuting, thus encouraging their labor force participation, especially in highly congested metropolitan areas. The benefit of remote work, however, may only be experienced by women employed in certain industries, and in jobs that can be performed remotely, often employing mostly highly educated workers (Bartik, Cullen, Glaeser, Luca, & Stanton, 2020; Buckman, Barrero, Bloom, & Davis, 2025). To the extent that women with lower levels of education remain disproportionately represented in inflexible occupations often requiring in-person contact and interactions (Albanesi & Kim, 2021), their labor supply may remain exposed to the mobility constraints induced by work-family tradeoffs, especially in congested areas. Such constraints can be mitigated if, as our findings suggest, partners with access to remote work use their work flexibility to provide housework and care for other household members.

The paper is organized as follows. In Section 2, we describe the data we use and how we select the sample of interest. In Section 3, we describe our primary empirical strategy and show that long commuting times disproportionately affect the extensive-margin labor supply of non-college graduate mothers. We also show that this result is robust to several different specifications and persists after accounting for possible differences in selection across geographical areas between college and non-college graduate women. Section 4 contains our analysis of the role of remote work arrangements in mitigating the negative impact of congestion on the extensive-margins labor supply of mothers. In Section 5, we rule out that our results can be explained by cultural differences between college and non-college graduate women or by the unobserved relationship between congestion and the availability of affordable market-provided childcare services. Section 6 concludes.

2 Data, sample selection and congestion measures

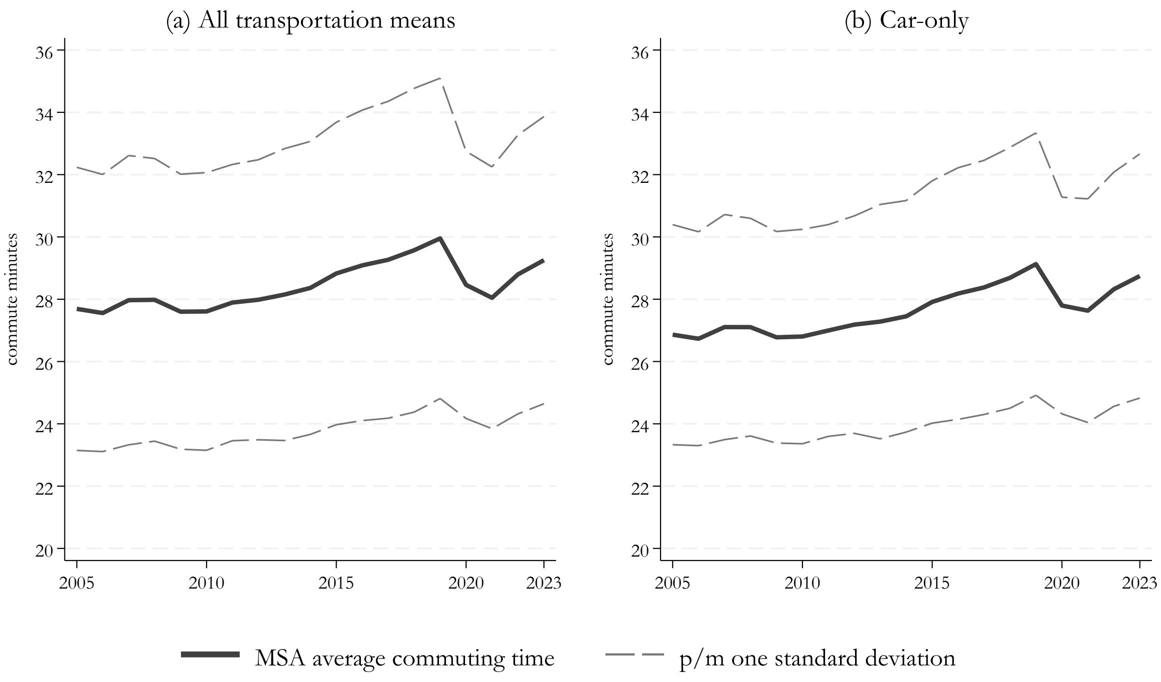
We use American Community Survey (ACS) data and American Time Use Survey (ATUS) data from IPUMS (Flood, Sayer, Backman, & Chen, 2025; Ruggles et al., 2025) for the years 2005 to 2019. Using the ACS, we construct a sample of women who are either household heads or heads' spouses, who are between 18 and 44 years old, and who co-habit with a male partner. For women with children we further limit the sample to those whose first childbirth occurred at age 18 or above, and whose eldest child living in the household is at most 17 years old. We exclude individuals who are currently enrolled in school and individuals with missing information on any demographic or geographical variable of interest. We also drop couples in which the age difference between partners is larger than 15 years. We further restrict all samples to individuals in metropolitan areas for which we have reliable congestion data. We follow Duranton and Turner (2011) and define time-constant metropolitan statistical areas (MSAs) using the 1990 Census Bureau delineation of metropolitan areas. We associate individuals in the ACS to 1990 metropolitan areas using their county of residence, and merge this information to the metropolitan-area specific information collected by Duranton and Turner (2011) regarding railroad development in 1898 and highway-development plans in 1947.

We follow similar criteria to construct a sample of women cohabiting with male partners using ATUS data. The ATUS includes a randomly selected set of individuals from households that have completed their participation in the Current Population Survey (CPS). Respondents are then asked to report their primary activities during the 24 hour period from 4:00 AM the day before the interview to 4:00 AM the day of the interview. For all individuals in our sample, we construct six different time-use categories: sleep, paid work, leisure and self-care, household activities and care activities, travel, missing or unknown. We further split travel-related activities into three subgroups: travel related to paid work (commuting), travel related to household activities or to care for others, travel related to leisure or self-care.

For each metropolitan area, we use ACS data to construct time-varying characteristics of local labor markets by computing the year-specific industrial composition at the state level, the year-MSA specific population, the share of employed 16-to-64 year old men in, respectively, professional occupations and executive occupations, the average occupation score of college graduate employed men, the average occupation score of employed men without a college degree, the average real labor income earned by, respectively, college graduate employed men and employed men without a college degree.

We use the MSA-specific and year-specific average commuting time of 16-to-64 year old employed men who commute by car, cab or motorbike as our main measure of metropolitan area commuting times and congestion. Calculating MSA-specific commuting times excluding individuals who use public transportation, commute by bike, or walk to their workplace provides a clean measure of congestion. This measure is not confounded by the possibly high commuting times of individuals who choose to commute using alternative means of transportation, and who arguably contribute to decreasing congestion rather than increasing it.

Figure 1: Cross-MSA average commuting time by year



Notes: American Community Survey 2005-2023. The figures represent the time trends in the cross-MSA average in commuting times among employed men who travel, respectively, using any transportation means (panel a) or using car, cab or motorbike only (panel b). The year-specific cross-MSA averages are weighted using the MSA-specific population. The dashed lines in the figure represent trends in the cross-MSA average commuting times plus and minus one standard deviation.

Figure 1 represents the time trends in the cross-MSA average commuting time. In panel (a), each MSA-specific commuting time is calculated using observations on all commuting employed men, while panel (b) represents the commuting time of employed men using car, cabs or motorbikes only. The latter is our preferred measure of congestion.

As shown in panel (a), the cross-MSA average commuting time of all men increases between 28 minutes per commute and 30 minutes per commute between 2005 and 2019, while dropping in 2020. One cross-MSA standard deviation in commuting time measures approximately 4 minutes. As shown in panel (b), the cross-MSA commuting time of men using cars, cabs or motorbikes rises from approximately 27 minutes in 2005 to 29 minutes in 2019 and drops thereafter, with a standard deviation of approximately 3.5 minutes every year.

3 Empirical analyses

3.1 Commuting times and mothers' labor force participation

We estimate the relationship between congestion and mothers' labor force participation using the following regression model

$$y_{imt} = \alpha + \beta z_{mt} + x'_{imt}\gamma + \zeta_r + \eta_t + \varphi_{rt} + u_{imt} \tag{1}$$

Where y_{imt} is a dummy variable taking value 1 if i, a woman with at least one child, participates in the labor force in year t, and z_{mt} is a year-specific standardized measure of the average commuting time among employed men who commute by car, cab or motorbike in the MSA (m) where woman i lives. The coefficient β captures the percentage-point change in the labor force participation of mothers associated with a one-standard-deviation increase in MSA-specific commuting time. The vector x_{imt} includes the following characteristics of women and of their partners: own and partner's age dummies, dummies for number of children and for eldest child's age, own and partner's race and ethnicity, couple's marital status, a dummy indicating whether woman i lives in her place of birth, i's birthplace, i's and partner's citizenship status, i's and partner's education, i's non-labor income, and partner's employment status, 1-digit occupation, usual work hours (in logs), (log) real labor earnings (in constant 1999 $). Additional controls include the following time-varying MSA m characteristics: average real labor income of, respectively, college graduate prime-age men and non-college graduate prime-age men in

MSA m in year t, share of prime-age men employed in, respectively, professional and executive occupations in MSA m in year t. The regression further includes region (ζ_r), year (η_t) and region-year (φ_{rt}) fixed effects.

As noted by Farré, Jofre-Monseny, and Torrecillas (2023), the OLS estimator of β can be affected by selection bias and by reverse-causality bias. The sign of selection bias is unclear. On the one hand, women with strong labor force attachment may self-select into bustling, highly congested metropolitan areas; on the other hand, long commuting times may attract couples with strong preferences for non-working mothers towards congested areas. Reverse causality bias should instead attenuate the estimated effect of congestion on women's labor force participation. Increases in women's labor force participation in a certain metropolitan area should arguably contribute to increases in congestion. To address concerns that reverse-causality and selection biases may strongly affect our results, we estimate the impact of congestion on mothers's extensive-margins labor supply using an instrumental variable approach. In particular, we instrument the MSA-specific congestion levels using two instruments proposed by Duranton and Turner (2011) to estimate the causal effect of road supply on vehicle traffic: the MSA-specific kilometers of railroad development in 1898 (the railroad instrument), and the MSA-specific kilometers of highway development planned in 1947 (the highway instrument).

Table 1 reports the estimated β coefficients from different specifications. Column (1) shows the impact of congestion on mothers' labor force participation estimated via OLS, in columns (2) and (3) we instrument MSA-year-specific congestion using, respectively, the railroad instrument and the highway instrument, in column (4) we use both instruments. The coefficients in panel (a) are estimated on all women in our sample, while panels (b) and (c) show the estimated effects of congestion on the labor force participation of, respectively, mothers without a college degree and college graduate mothers.

The results in Table 1, panel (a) support results previously found by Black, Kolesnikova, and Taylor (2014) and Farré, Jofre-Monseny, and Torrecillas (2023) using alternative methods. We find that mothers' labor force participation declines by 1.1-to-1.8 percentage-points with a 3.5-minute (one standard deviation) increase in MSA commuting time.[4]

[4]Farré, Jofre-Monseny, and Torrecillas (2023) find that a 10-minute increase in local commuting time decreases the labor force participation rate of married women by 4 percentage points (Table 1 in their paper). They find a similar effect for women with exactly one child. As we show in Table 1 panel (a), we

Importantly, the results of the cluster-robust IV diagnostic tests reported in panel (a) corroborate the validity of the instrumental variable strategy that we implement. Based on the value of the Kleibergen-Paap Chi-Square statistic, we can reject that our IV model is not identified, while the value of the Kleibergen-Paap Wald-F statistic supports the relevance of the instruments we use. The p-value of the Hansen J test reported in column (4) suggests that we cannot reject the hypothesis that the excluded instruments are uncorrelated with the error term in regression 1 and satisfy the exclusion restriction, at the conventional 5% significance level.

Crucially, panels (b) and (c) in Table 1 show that congestion has highly heterogeneous effects on different groups of women. Among mothers without a college degree, one standard-deviation increase in congestion decreases the probability to participate in the labor force by 1.8-to-2.6 percentage points. Among college graduate mothers, there does not appear to be any robust relationship between MSA-specific commuting times and extensive-margins labor supply.

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find that a 3.5 increase in commuting time decreases the labor force participation rate of all women with children by 1.1 to 1.76 percentage-points. This implies that a 10-minute increase in commuting time decreases the labor force participation rate of women with children by 3.15 to 5.03 percentage points, which is in line with the estimates in the existing literature.

Table 1: Effect of one std. dev. increase in MSA commuting time on mothers' LFP

	(1) Region FE	(2) Reg. FE/IV Railway	(3) Reg. FE/IV Highway	(4) Reg. FE/IV Both
(a) All	-0.0110***	-0.0107^{+}	-0.0176**	-0.0150**
	(0.0029)	(0.0064)	(0.0056)	(0.0057)
Obs	990703	990703	990703	990703
R-squared	0.145	0.098	0.098	0.098
Under-ID (K–P LM) Chi-Sq.		20.88	17.33	21.13
Weak-ID (K–P rk Wald F)		15.31	60.38	30.77
Hansen J p-value				0.121
Andreson-Rubin F p-value		0.153	0.012	0.038
DWH p-value		0.953	0.170	0.506
(b) No College	-0.0180***	-0.0198**	-0.0257***	-0.0234***
	(0.0032)	(0.0073)	(0.0068)	(0.0065)
Obs	564468	564468	564468	564468
R-squared	0.141	0.088	0.088	0.088
Under-ID (K–P LM) Chi-Sq.		21.98	17.93	22.40
Weak-ID (K–P rk Wald F)		15.90	51.52	26.06
Hansen J p-value				0.323
Andreson-Rubin F p-value		0.022	0.004	0.011
DWH p-value		0.753	0.220	0.411
(c) College	-0.0000	0.0025	-0.0048	-0.0027
	(0.0034)	(0.0068)	(0.0065)	(0.0064)
Obs	426235	426235	426235	426235
R-squared	0.148	0.114	0.114	0.114
Under-ID (K–P LM) Chi-Sq.		18.53	16.88	19.53
Weak-ID (K–P rk Wald F)		13.63	78.35	43.48
Hansen J p-value				0.096
Andreson-Rubin F p-value		0.698	0.471	0.099
DWH p-value		0.567	0.278	0.796

Notes: American Community Survey 2005-2019. Sample selection is described in Section 2. Panels (a), (b) and (c) report the β coefficient estimated separately for all mothers, mothers without a college degree, and mothers with a college degree. Coefficients in column (1) are estimated via OLS, coefficients in columns (2), (3) and (4) are estimated via 2SLS. All models include all control variables listed in Section 3 and weight observations using person weights. Standard errors are clustered at the MSA level. Columns (2), (3) and (4) report cluster-robust tests for instruments relevance (Kleibergen-Paap Chi Square, Kleibergen-Paap Wald F), weak-instrument-robust significance of the coefficient of interest (Anderson-Rubin F p-value), and p-value for the test of endogeneity of the instrumented regressor (Durbin-Wu-Hausman test p-value). Column (4) adds the p-value for the overidentification Hansen J test for the exogeneity of the excluded instruments (railroads 1898, highway 1947). p-value < 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

3.2 Robustness

In this section we show the results of several robustness exercises corroborating the evidence provided in the previous section that congestion mostly affects the labor force participation of mothers without a college degree.

3.2.1 Underlying differences in family composition

The stronger effect of congestion on the labor force participation of mothers with relatively low levels of education may be due to underlying differences across education groups in family structure and geographical location. In Table A1 in the Online Appendix, we show that significant differences exist between college and non-college graduate women along these lines. Compared to college graduate women, women without a college education are less likely to cohabit with college-graduate and employed male partners, are less likely to be married to their cohabiting partner, are less likely to live in large metropolitan areas, are more likely to have kids and have their first child at a younger age. To reduce concerns that any of these difference mechanically determine any of the results explained in the previous section, we control for individual and partner's characteristics and for the number and age of any household children in all the regression models that we estimate.

To further exclude that differences in the average characteristics of college and non-college graduate mothers may determine any of the results we find, we re-estimate regression 1 via OLS and IV separately for college graduate mothers and for mothers without a college degree in four different subsamples: married mothers, mothers with one child only, mothers whose eldest child is younger than 6 years-old, mothers not living in the 20 largest metropolitan areas, mothers with college graduate partners. As shown in Table 2, congestion appears to disproportionately affect the labor force participation of mothers with no college education within all of these subsamples of women.

Table 2: Congestion and mothers' LFP for different subsamples

	(a) No College		(b) College	
	(1) Region FE	(2) Reg. FE/IV Both	(3) Region FE	(4) Reg. FE/IV Both
(1) Married only	-0.0180***	-0.0236***	-0.0001	-0.0026
	(0.0033)	(0.0068)	(0.0035)	(0.0064)
Obs	524088	524088	422444	422444
Weak-ID (K–P rk Wald F)		25.77		41.41
Hansen J p-value		0.345		0.095
(2) One child	-0.0139***	-0.0195**	0.0007	-0.0035
	(0.0028)	(0.0061)	(0.0038)	(0.0064)
Obs	162009	162009	132276	132276
Weak-ID (K–P rk Wald F)		27.48		38.96
Hansen J p-value		0.377		0.487
(3) Eldest child $<$ 6 y.o.	-0.0182***	-0.0179**	0.0038	0.0005
	(0.0032)	(0.0069)	(0.0037)	(0.0066)
Obs	150346	150346	182836	182836
Weak-ID (K–P rk Wald F)		29.23		40.52
Hansen J p-value		0.476		0.246
(4) No Top-20 MSAs	-0.0135***	-0.0300^{+}	0.0038	0.0090
	(0.0034)	(0.0178)	(0.0025)	(0.0122)
Obs	255937	255937	159902	159902
Weak-ID (K–P rk Wald F)		6.57		12.43
Hansen J p-value		0.518		0.306
(5) Coll. grad. partner	-0.0082*	-0.0107	0.0001	-0.0028
	(0.0037)	(0.0067)	(0.0035)	(0.0064)
Obs	117193	117193	417471	417471
Weak-ID (K–P rk Wald F)		33.57		42.03
Hansen J p-value		0.216		0.087

Notes: American Community Survey 2005-2019. Sample selection is described in Section 2. Rows (1)-to-(5) report the β coefficient estimated separately for married mothers, mothers with one child, mothers whose eldest child is younger than 6 years-old, mothers not living in any of the 20 largest metropolitan areas, mothers married to college-graduate partners, among non-college graduate women (panel a) and college women (panel b) respectively. Coefficients in column (1) are estimated via OLS, coefficients in columns (2) are estimated via 2SLS using both Duranton and Turner (2011) instruments. All models include all control variables listed in Section 3 and weight observations using person weights. Standard errors are clustered at the MSA level. Columns (2) and (4) report cluster-robust tests for instrument relevance (Kleibergen-Paap Wald F) and the p-value for the overidentification Hansen J test for the exogeneity of the excluded instruments. 2SLS results in row (4) for non-college mothers should be interpreted cautiously given that the weakness of the excluded instruments cannot be rejected (see Weak-ID Wald F). p-value $<$ 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

3.2.2 Selection

In this section we provide evidence that unobserved differences between education groups in women's sorting across geographical areas do not drive our findings. We deem this analysis particularly relevant. The estimation results in Tables 1 and 2 show that there is no relationship between congestion and the extensive-margins labor supply of college graduate mothers. While the results are robust to implementing a 2SLS estimation, the Hansen J test p-value in panel (c) of Table 1 may cast some doubts regarding the exogeneity of the excluded instruments in our analysis of college graduate mothers. This may imply that the IV strategy that we implement may not fully address the endogenous selection of highly educated women across geographical areas. Extensive evidence, however, suggests that this selection exists.

First, as employment opportunities and local labor demand shocks are an important driver of within-US migration (Diamond, 2016), it is plausible that areas with higher labor force participation are also characterized by more congestion and longer commuting times. Second, even conditional on observable characteristics, it is possible that women who self-select into highly congested metropolitan areas have a lower idiosyncratic marginal utility cost from working and commuting compared to women residing in less congested cities. The most urbanized areas are characterized by the availability of non-tradable amenities that attract mostly young and highly educated workers (Albouy & Faberman, 2024; Couture & Handbury, 2020). If these workers tend to be strongly attached to the labor market irrespective of whether they have children, their selection into congested metropolitan areas may attenuate the negative effect of commuting times on the labor supply of highly educated mothers. In Section B in the Online Appendix we use a simple neoclassical labor supply model with commuting costs to formalize this argument.

To reduce concerns that we fail to capture the negative effect of congestion on college-graduate mothers due to strong underlying selection bias, we estimate the following regression

$$y_{imt} = \alpha + \beta c_{imt} + \gamma z_{mt} + \delta c_{imt} \times z_{mt} + x'_{imt}\rho + \zeta_m + \eta_t + \varphi_{rt} + u_{imt} \tag{2}$$

Where y_{imt} is a dummy variable taking value 1 if woman i in year t participates in the labor force, c_{imt} is a dummy variable taking value 1 for mothers and 0 for women without children, and z_{mt} is the standardized measure of the MSA average commuting time. The coefficient β captures the percentage-point difference in the labor force participation rate between mothers and women without children in city m and year t, while γ represents the percentage-point change in the labor force participation of women without children associated with one standard deviation increases in the year-specific MSA commuting time. The main coefficient of interest, δ, measures the degree of heterogeneity in the labor force participation difference between mothers and women without children across MSAs characterized by different levels of congestion.

The regression model includes year fixed effects (η_t), region-year fixed effects (φ_{rt}), and the full set of controls listed in Section 3. Furthermore, we include different MSA fixed effects, ζ_m, for mothers and women without children, thus we allow geographical sorting to differ between these two groups of women.

This empirical strategy exploits the within-MSA variation in congestion over time to identify the differential impact of commuting time on the labor supply of mothers by comparing it to the labor supply of women without children who live in the same geographical area. The underlying intuition is that comparing the labor supply of different groups of women with equal education levels within MSAs could more closely estimate the causal effect of congestion on the labor force participation of mothers. This identification strategy requires the assumption that, even if women with children and women without children of the same level of education sort across metropolitan areas within a region for different reasons, those differences should remain roughly constant over time within the same metropolitan area. Under this assumption, δ identifies the impact of congestion on mothers' labor force participation rate. Columns (1) and (2) of Table 3 report the γ and δ coefficients of regression 2 estimated separately for women with different levels of education via OLS.[5]

The results confirm the finding that congestion exerts a negative impact on the probability that mothers with no college education participate in the labor force, while not affecting the labor force participation rate of college graduate mothers.

[5]We cannot estimate the model via IV since the Duranton and Turner (2011) instruments do not vary within MSA over time.

Table 3: Congestion and LFP gap between mothers and women without children

	(1) No College	(2) College
Z(MSA CT)	0.0078	-0.0063
	(0.0063)	(0.0053)
Z(MSA CT)*Mother	-0.0135*	0.0125
	(0.0065)	(0.0087)
Obs	707970	594051
R-squared	0.151	0.160
Diff = 0 p-value	0.030	

Notes: American Community Survey 2005-2019. Sample selection is described in Section 2. Panels (a) and (b) report the γ and δ coefficients in regression 2 estimated via OLS separately for women without a college degree and for women with a college degree. All models include full set of control variables listed in Section 3, including region $\times$ year fixed-effects, and weight observations using person weights. Different MSA fixed effects for mothers and women without children are also included. Column (1) reports the p-value of the test: $H_0 : \delta_{nocol} = \delta_{col}$ v $H_1 : \delta_{nocol} \neq \delta_{col}$. The test is performed comparing columns (1) and (2) coefficients. Standard errors are clustered at the MSA level. p-value < 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

3.2.3 Additional robustness exercises

In the Online Appendix we show the results of further robustness exercises corroborating Table 1 evidence. Table A2 in the Online Appendix shows that our main results are unaffected when the impact of congestion is estimated in a unique regression for all women with children, and allowed to be heterogeneous across education groups. Table A3 shows that results are unaffected when individual weights are not used. Table A4 shows that our main results are robust to the use of different sets of control variables. Finally, Tables A5 and A6 show the results of estimating regression 1 on ATUS data, and further corroborate our finding that commuting times are especially detrimental to the labor force participation of mothers with lower levels of education.

4 Mechanism: remote work

In this section we show that access to remote work can attenuate the congestion child penalty and explain why long commuting times are especially detrimental to the labor force participation of non-college graduate mothers.

We study two channels through which remote work can reduce the negative effect of commuting times on mothers' labor supply. First, women cohabiting with male partners employed in jobs that can be performed remotely may face weaker mobility constraints. For example, if these men use their remote work arrangements to devote more time to housework and care for other household members, their female partners whose jobs require them to commute may be able to remain in the workforce in spite of the time cost of long travel times. Second, women whose jobs enable them to work remotely may be able to switch to telework in metropolitan areas characterized by long commuting times, reducing their exposure to congestion. This may boost their participation in the workforce in highly congested metropolitan areas, especially if the presence of children exacerbates their mobility constraints.

Below, we show that cohabiting with a male partner employed in a highly flexible occupation is associated with a reduction in the congestion child penalty for non-college graduate mothers. Although college graduate women are more likely to cohabit with men employed in jobs that can be performed remotely, the probability that they participate in the workforce is unaffected by congestion irrespective of their partners' remote work access. We show that in fact, highely educated women are themselves more likely than non-college graduate women to be employed in jobs that can be performed remotely. Moreover, college graduate women rely more intensely on remote work in highly congested areas. The substitution of remote work for in-person work can shied college graduate women from the congestion child penalty.

4.1 Partner's remote work

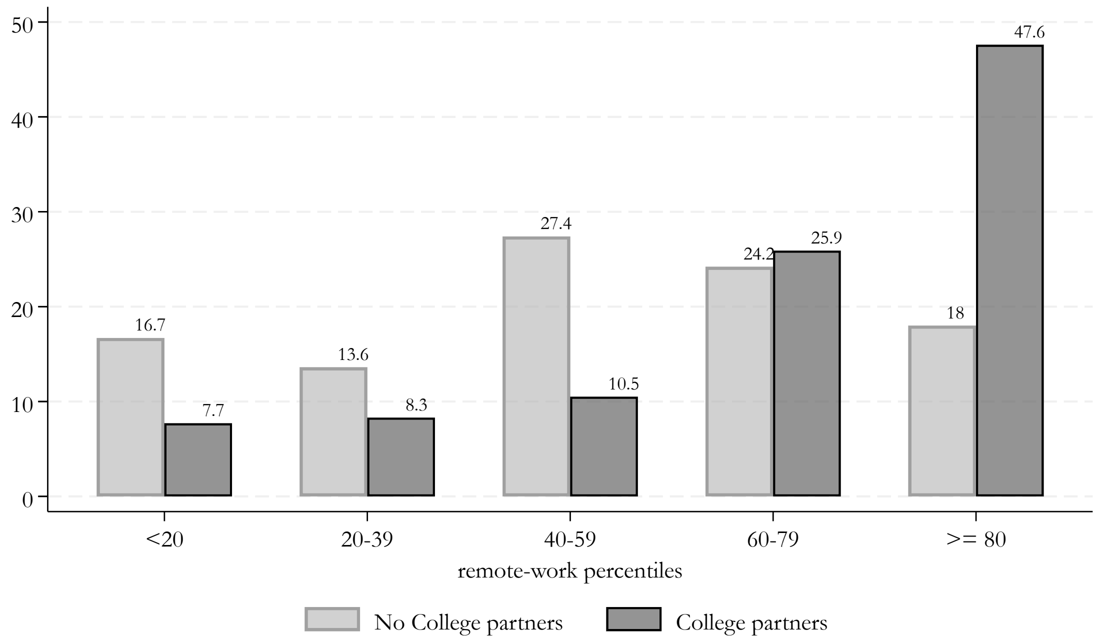
Figure 2 shows the distribution of non-college and college graduate employed male partners of 18- to 44-year-old women in the 2005-2019 ACS across 3-digit occupations classified by percentile of remote-work incidence. We calculate the incidence of remote work within each occupation between 2005 and 2019 as the share of prime-age men who work from home among the total number of men employed in the occupation. As shown in Figure A1 and in Table A7 in the Online Appendix, the incidence of remote-work varies across occupations from less than 1% to more than 30%, and it is highly concentrated in the top-quintile of its cross-occupation distribution. Table A8 in the Online Appendix further shows that the 20 occupations with the highest remote-work incidence include primarily managerial and professional occupations, typically employing highly educated workers.[6]

Figure 2 shows that college graduate male partners of women in our ACS sample are disproportionately represented in high-remote-work-incidence occupations. Specifically, 48% of them are employed in occupations in the top quintile of the remote-work incidence distribution, while only 16% of them are employed in occupations in its bottom two quintiles. Only 18% of male partners with no college education, instead, are employed in jobs at the top of the remote-work-incidence distribution, while 31% are employed in occupations in the bottom two quintiles of the remote-work-incidence distribution.

As shown in Figure A2 in the Online Appendix, classifying men according to their female partner's education confirms that college graduate women are more likely to cohabit with men employed in high-remote-work incidence occupations, compared to women with lower levels of education. This is unsurprising, as Table A1 in the Online Appendix shows that 98% of college graduate women in our sample cohabit with similarly educated partners, while only 19% of women without a college degree cohabit with college graduate men, consistent with well-known patterns of educational assortative mating (Eika, Mogstad, & Zafar, 2019).

[6]It is possible that our measure of occupation-specific incidence of remote work underestimates the extent to which different jobs are performed remotely. In the American Community Survey, work from home is recorded for employed individuals who report to have not commuted in the previous week. This suggests that our variable may only capture remote work among individuals who work fully remotely, who may represent only a portion of individuals who have at least some degree of workplace flexibility. The fact that the ACS question may capture "fully remote" work has been noted by Buckman, Barrero, Bloom, and Davis (2025) as well. Our variable, however, appears to reasonably represent the cross-occupation hierarchy in remote work incidence. Remote work, in fact, became suddenly popular especially in high-skill occupations with the 2020 COVID pandemic. Its rapid adoption was facilitated by the fact that the technologies and infrastructures needed for work-from-home implementation were already available for use (Barrero, Bloom, & Davis, 2023). As shown in Table A8 in the Online Appendix, the occupations with the highest incidence of fully-remote work in 2005-2019 largely overlap with the occupations with the highest incidence of fully-remote work measured in 2021-2023.

Figure 2: Occupational distribution of employed male partners by remote-work quintile



Notes: American Community Survey 2005-2019. The figure represents the distribution of employed male cohabiting partners of 18- to 44-year-old women across 3-digit occupations (1990 Census classification) classified in five quintiles of remote-work incidence. Male partners are split in two sub-groups according to whether they hold a college degree (dark bars) or not (light bars). All statistics reported in the figure are weighted using partner-specific individual weights. For each 3-digit occupation, the incidence of remote work is calculated as the share of prime-age men working in that occupation between 2005 and 2019 who report to not usually commute as they work from home.

Time-use diaries from the 2005-2019 samples of the American Time Use Surveys provide further evidence that college graduate men cohabiting with female partners are more likely to work from home compared to men with lower levels of education.

Using the ATUS time-use diaries, we calculate the total number of hours worked by a person in the previous day, and identify work-from-home hours through information on the location where work activities were performed. We then calculate the share of employed men who worked at least one-half hour from home in the previous day (1/2 h remote work), the share of daily work hours worked from home in the previous day (% Remote Work hours), and the share of men who worked 100% of their previous-day work hours from home (Full day Remote Work). We report the incidence of remote work calculated in the ATUS separately for college-graduate and for non-college graduate men cohabiting with female partners (panel a), and for male partners cohabiting with college graduate and non-college graduate women (panel b), in Table 4. The table clearly shows that remote work is substantially more widespread among college graduate men, the majority of whom cohabit with highly educated women.

The table also includes the share of previous-day hours, and the total previous-day hours, devoted by men to care for other household members and to housework. Highly educated men allocate more time to within-household care activities than men with lower levels of education.

Table 4: Remote work among working men cohabiting with female partners

	(1) No College	(2) College	(3) Diff =0 p-value
		(a) Own education	
Hours worked	8.337	7.869	0
1/2 h remote work	.113	.302	0
% RW hours	.077	.202	0
Full day RW	.056	.147	0
% Care/housework hours	.084	.104	0
Care/housework hours	2.001	2.476	0
		(b) Partner's education	
Hours worked	8.33	7.901	0
1/2 h remote work	.118	.287	0
% RW hours	.077	.196	0
Full day RW	.055	.143	0
% Care/housework hours	.083	.103	0
Care/housework hours	1.988	2.468	0

Notes: American Time Use Survey 2005-2019. The sample includes 18- to 44-year-old men cohabiting with female partners, who are not enrolled in school, are employed, and worked in the previous day according to their time-use diaries. The table shows the average hours worked in the previous day, the share of men working at least one-half hour remotely in the previous day, the share of work hours worked remotely, the share of men who worked fully remotely in the previous day, the share of daily hours devoted to care and housework, and the total number of hours devoted to care and housework, all reported separately by level of education. In panel (a) men are classified in two education groups according to their own level of education. In panel (b) men are classified in two education groups according to their partner's level of education. All statistics are computed using time-use individual weights. P-values for t-tests for differences in means and proportions across education groups are reported in column (3).

We use the ATUS to provide additional evidence that the difference along education lines in care and housework time among working men is linked to the difference in the availability and use of remote work arrangements across education groups. Table 5 shows the selected coefficient of five linear regression models whose dependent variable is the time that men spent providing care or housework in the previous day (in minutes). All models control for men's number of children, day of the week in which the time-use diary was recoded, whether the diary-day is a holiday, whether an individual's partner is a college graduate, as well as year and region fixed effects. As shown in column (1), college graduate men spend approximately 12.60 minutes more per day providing care or performing housework than non-college graduate men. This difference, however, is driven by the higher incidence of remote work among college graduate men. In panel (a), we include an indicator taking the value 1 if a man worked at least 1/5 of their total work time from home in the previous day, and 0 otherwise (row 2, columns 2 and 3), and an interaction of this variable with the college dummy (row 3, column 3). Once we control for remote work, the difference in care and housework time between college and non-college graduate men becomes small and not statistically significant. We obtain similar results in panel (b), where we instead measure remote work using a variable indicating whether a man spent at least 50% of their work hours working remotely in the previous day, and its interaction with the college dummy (column 5).

Table 5: Care/housework time, by education and remote work

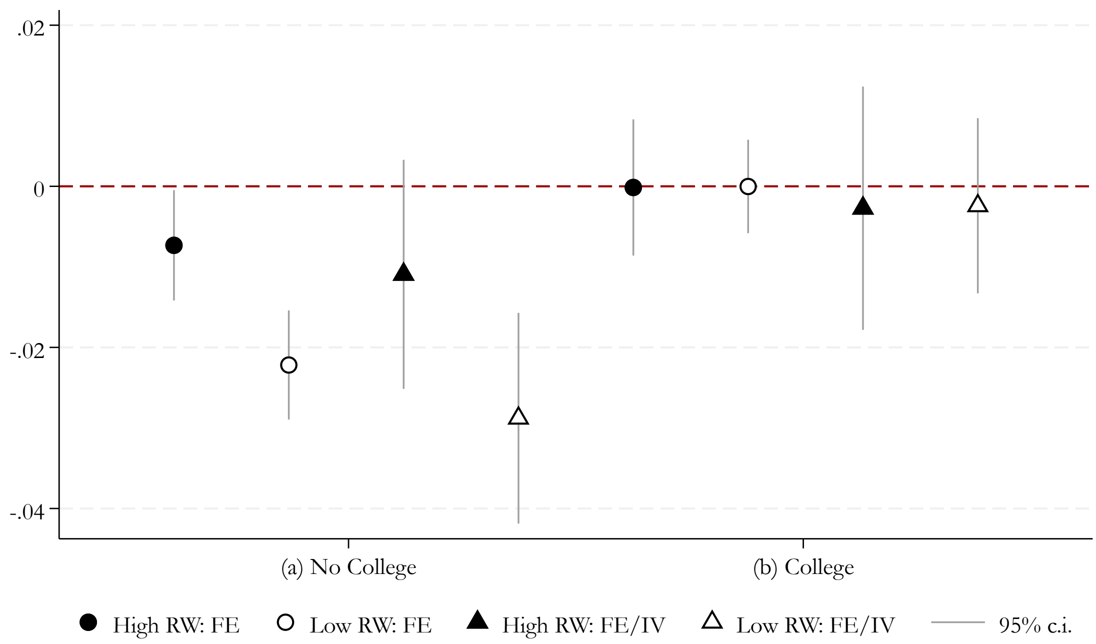
		(a) Some RW		(b) Half-day RW	
	(1)	(2)	(3)	(4)	(5)
College	12.58**	4.63	4.80	5.38	4.58
	(4.31)	(4.18)	(4.18)	(4.15)	(4.12)
Remote Work		92.27***	92.90***	104.45***	101.15***
		(6.00)	(9.97)	(6.52)	(10.69)
College * (Remote Work)			-1.03		5.45
			(11.89)		(12.78)
Obs.	7435	7435	7435	7435	7435
R-squared	0.083	0.141	0.141	0.150	0.150

Notes: American Time Use Survey 2005-2019. Sample selection is described in Table 4 notes. The table shows selected coefficients from five linear regression models whose dependent variable is the time that an individual spent providing care to their household members or housework in the previous day. This includes the time spent traveling for care or housework tasks and excludes leisure and self-care. Control variables include weekday dummies, a variable indicating whether the diary day was a holiday, dummies for the number of own kids in the household, a variable indicating whether the individual's partner is a college graduate, year and region fixed effects. The first row shows the coefficient on the college graduate indicator. In panel (a) row-2 variable takes the value of 1 if a man worked at least 1/5 of their total work time from home in the previous day, and 0 otherwise and in panel (b) row-2 variable takes the value of 1 if a man worked at least half of their total work time from home. Row three adds an interaction term between this measure of remote work and the college graduate indicator. ATUS individual weights are applied. Robust standard errors are shown in parentheses. p-value < 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

These findings support the hypothesis that the time-constraints associated with within-household care activities may be less binding for women who cohabit with men working remotely. This additional support may enable women to remain in the labor force upon having children irrespective of the commuting costs induced by traffic and congestion.

To quantify the extent to which partner's job flexibility may be associated with a reduction in mothers' mobility constraints, we estimate regression 1 using ACS data both via OLS and via 2SLS separately for mothers with no college degree and for college graduate mothers. We restrict the sample to women cohabiting with employed men, and further split each education-specific group of women according to whether their male partners work in an occupation in the top quintile of the remote-work-incidence distribution (high remote-work incidence) or not (low remote-work incidence). In addition to the control variables listed in Section 3, we further control for the following partner's 3-digit occupation characteristics: its one-digit occupation class, the incidence of long work hours, and its occupation income-score percentile.

Figure 3: Congestion and mothers' LFP by partner's occupation remote-work incidence



Notes: American Community Survey 2005-2019. The figure represents the β coefficient in regression model 1 estimated via OLS (circles) and via IV (triangles) using Duranton and Turner (2011) instruments, separately for non-college (panel a) and college graduate (panel b) women with children cohabiting with men in high-remote-work incidence occupations (black symbols) and with men in low-remote-work incidence occupations (white symbols). All regression models include all control variables listed in Section 3 and the additional characteristics of partners' occupation listed in Section 4.1. All regressions are weighted using individual weights, and standard errors are clustered at the MSA level. The figure includes 95% confidence intervals. The estimated coefficients, standard errors, and results of IV diagnostic tests are reported in Table A9 in the Online Appendix, which also reports the β coefficients estimated in a model including MSA fixed effects.

The estimated effects of congestion on mothers' labor force participation for each group of women are reported in Figure 3 together with 95% confidence intervals. Among mothers with no college degree, the negative effect of congestion on the probability of participating in the workforce is especially strong among women cohabiting with men employed in inflexible occupations. Within this group, the labor force participation rate decreases by 2.2 (OLS estimate) to 2.8 (IV estimate) percentage points with an increase in MSA-specific commuting times of around 3.5 minutes (one standard deviation). Among mothers with low levels of education who cohabit with male partners employed in jobs with high incidence of remote work, labor force participation is less responsive to congestion. Within this group, the labor force participation rate declines by a marginally significant 0.7 to 1.1 percentage-point with one standard deviation (3.5 minute) increase in MSA-specific commuting time.[7]

Interestingly, the labor force participation of college graduate mothers does not appear to change with congestion, irrespective of their partners' job flexibility. In the next section we argue that the possibility of working from home may enable college graduate mothers to stay in the labor force in spite of the time-cost of commuting induced by congestion, irrespective of the level of flexibility in their partners' job.

4.2 Women's remote work

We use both ACS and ATUS data to study whether and to what extent the availability of remote-work arrangements shields college graduate mothers' labor supply from the negative impact of congestion.

Using ACS data, we calculate two measures of remote-work incidence among employed women: the share of women employed in occupations with high remote-work incidence,

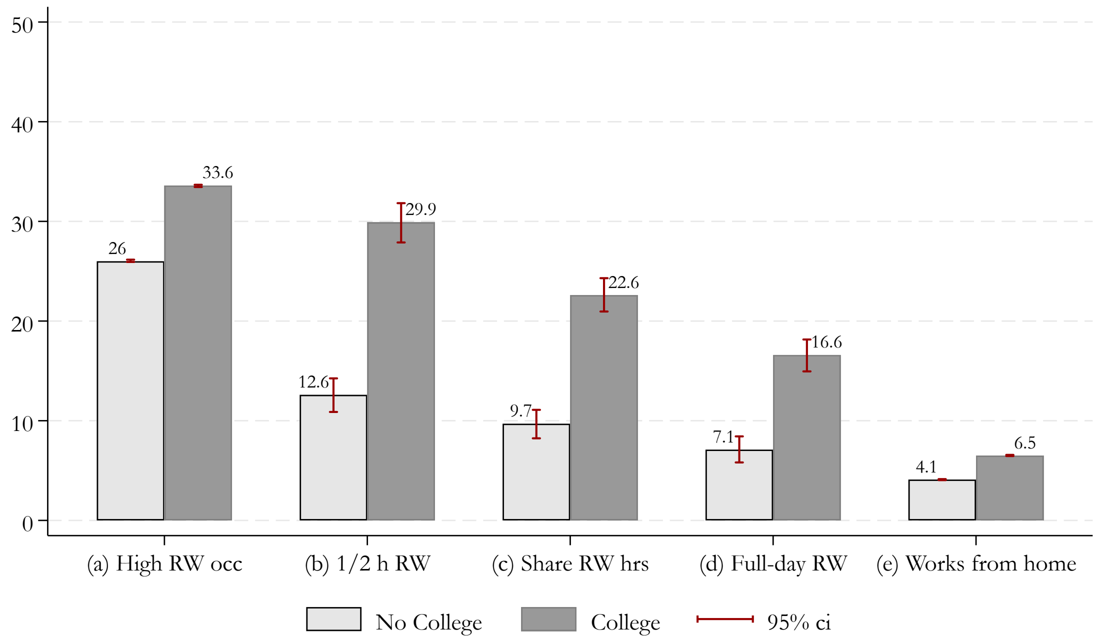
[7] All the regression models that we estimate control for the income-score of women's partners' occupations and for their labor income, as well for the share of men working long hours within the occupation. Yet, one may wonder whether the heterogeneity in the effect of congestion across groups of women defined by the incidence of remote-work in their partner's occupation may reflect other characteristics of those jobs, or of men who perform them. To corroborate the results shown in Figure 3, Figure A3 in the Online Appendix shows that the effect of congestion on the labor force participation of mothers does not change across groups characterized by higher or lower partners' occupation income-score.

and the share of women who work fully remotely. The first measure is computed using the share of women in the top quintile of the remote-work-incidence distribution of occupations defined in the previous section. The second measure is computed using the share of employed women who report their usual transit time to be 0 minutes because they work from home. These two measures capture, respectively, the extent to which remote work is available to women based on the characteristics of their jobs, and the extent to which women use remote-work arrangements to their full potential.

As described in the previous section, ATUS diaries enable us to compute additional measures of remote-work use: the share of employed individuals who worked at least one-half hour from home in the previous day, the share of work hours that individuals worked from home in the previous day, as well as the share of employees working fully remotely (100% of work hours) in the previous day.

We plot the incidence of remote work among women with different levels of education in Figure 4. All measures of remote work show that women with a college degree have a higher incidence of remote work than non-college educated women. In panel (a) we show that working college graduate women are significantly more likely than women with no college degree to be employed in occupations with high remote-work incidence. Almost 34% of college graduate women are employed in occupations at the top-quintile of the remote-work incidence distribution, while 26% of women with no college degree are employed in those jobs. Panel (b) shows that college graduate women (29.9%) are in fact more likely than women with lower levels of education (12.6%) to have worked from home for at least half an hour the previous day. Looking at overall remote work hours, panel (c) shows that college graduate women spent 22.6% of their previous-day work hours working from home, while women with no college degree worked remotely 9.7% of their previous-day work hours. This difference reflects both the large difference in the extensive-margin use of remote work reported in panels (a) and (b), and differences in the intensive-margin use of remote work. In fact, in panel (d) we show that 16.6% of employed college graduate women worked fully remotely in the previous day, while 7.1% of employed women with no college degree worked fully remotely in the previous day. Finally, in panel (e) we show that 6.5% of working college graduate women do not usually commute as they work from home, while only 4.1% of women without a college degree report to usually work remotely.

Figure 4: Incidence of remote-work arrangements among working women



Notes: Panels (a) and (e): American Community Survey (ACS) 2005-2019. Panels (b), (c) and (d): American Time Use Survey (ATUS) 2005-2019. Sample selection described in Section 2. The figure represents the share (%) of 18- to 44-year-old employed women with a college degree (dark bars) and without a college degree (light bars) who work remotely, according to different definitions of remote work. Panels (a) through (e) all present statistics by education level. Panel (a) represents the share of ACS women who are employed in high-remote-work (RW) occupations. Panel (b) shows the share of ATUS women who report to have worked at least 1/2 hour remotely in the previous day. Panel (c) shows the share of hours worked remotely by ATUS women during the previous day. Panel (d) shows the share of women who report to have worked fully remotely in the previous day. Panel (e) shows the share of ACS women who do not usually commute because they work from home. All statistics are computed using individual weights.

The graph shows remarkable differences along education lines both in the availability of remote work and in women's use of this flexible work arrangement. College graduate women appear to have substantially more flexibility than non-college graduate women in their workplace location. This fact suggests that non-college graduate women may be substantially more exposed to the congestion-induced time costs of commuting than college graduate women. Consistent with this argument, we further show that the gap between college and non-college graduate women in remote-work use is higher in highly-congested metropolitan areas.

In Figure 5 we plot the gap in remote-work use between college graduate women and non-college graduate women in low-congestion metropolitan areas and high-congestion metropolitan areas, estimated using ATUS data. The gaps are estimated controlling for region and year fixed-effects, for time-varying MSA characteristics and for both the women's and their partners' characteristics. The figure shows an interesting pattern.

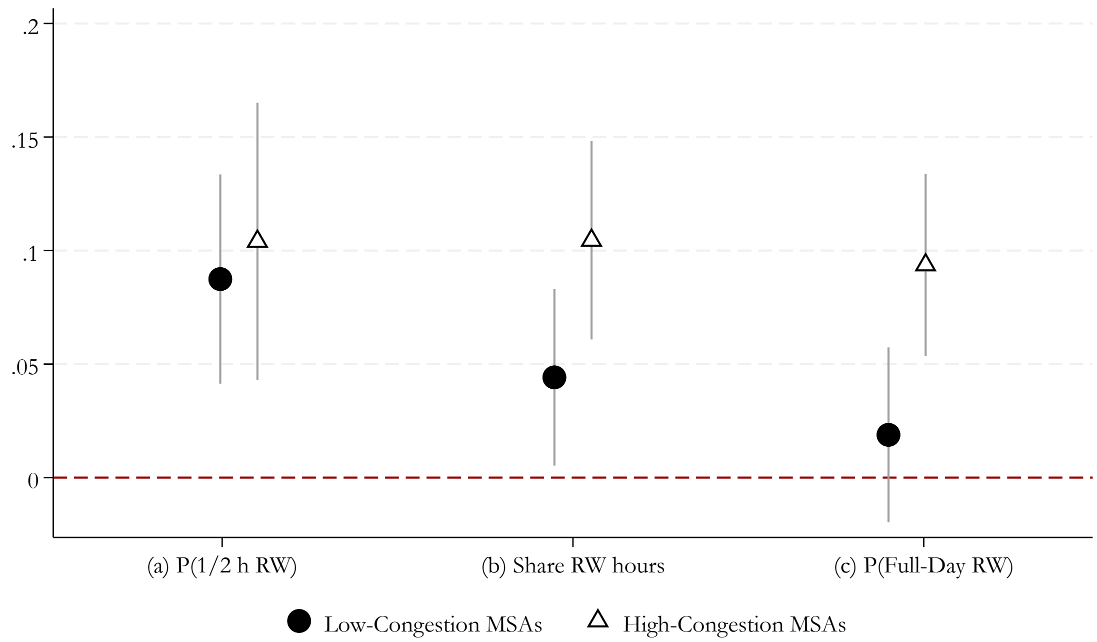
First, the adjusted gap in the probability of working from home at least 1/2 hour in the previous day, shown in panel (a), is virtually identical across metropolitan areas characterized by different levels of congestion. Both in high-congestion and low-congestion areas, college graduate women are around 10 percentage-points more likely to have spent at least half an hour working from home in the previous day compared to non-college educated women. The similarity across different geographical areas in the gap between college and non-college employed women in the probability of spending any amount of time working from home may be a reflection of the difference in the occupations that college and non-college graduate women typically perform.

However, panels (b) and (c) show that the reliance on remote work changes with congestion among college graduate women at the intensive margin. Panel (b) shows that in low-congestion areas, the share of hours worked remotely by college graduate women is less than 5 percentage-points higher than the share of hours worked remotely by mothers without a college degree. In high-congestion areas, the gap between college and non-college graduate women in the share of work hours worked from home in the previous day is as large as 10 percentage points. Furthermore, panel (c) shows that while the probability of having worked fully remotely in the previous day does not differ between college graduate women and women without a college degree in low-congestion areas, college graduate women are approximately 9 percentage-points more likely to have worked fully remotely in the previous day compared to non-college graduate women in highly congested metropolitan areas.

We provide additional support for these findings in the Online Appendix. In panel (d) of Table A10, we show that using ACS data to estimate the adjusted gap between college and non-college graduate women in the probability of working fully remotely delivers similar results: college graduate women are more likely to usually work remotely than non-college graduate women in highly congested areas only. We further show in Table A11 that the results in Figure 5 are driven by the higher incidence of remote work among college graduate women in highly congested areas. Within this group, in fact, the incidence of remote work among employed workers is significantly higher in high-congestion areas than in low-congestion areas.

These results are key to understanding why the time cost of commuting mostly affects low-education mothers, while not having any systematic impact on the labor force participation of college graduate mothers. Women in the latter group, in fact, appear to more intensely rely on remote work arrangements available to them, in areas where commutes are especially time consuming. This flexibility may enable college graduate women to remain in the labor force even as the presence of children may exacerbate their mobility constraints, particularly in highly congested metropolitan areas.

Figure 5: Gap between college and non-college graduate women in RW by congestion



Notes: American Time Use Survey 2005-2019. Sample selection is described in Section 2. The sample is restricted to women who report in their time-use diaries to have worked in the previous day. The figure plots the coefficients capturing the gap in the probability of working remotely between college and non-college graduate women, estimated separately for women living in high-congestion areas (white triangles) and in low-congestion areas (black circles). The groups are defined as follows. In every year, we split women in percentiles based on the average commuting time in their metropolitan area of residence. We then create two groups depending on whether women live in areas whose local commuting time is at most as high as the year-specific cross-MSA median (low) ot above the median (high). The dependent variable in panel (a) is a dummy variable taking value 1 if a woman worked at least 1/2 hour from home in the previous day. The dependent variable in panel (b) is the share of work hours worked remotely in the previous day. The dependent variable in panel (c) is a dummy taking value 1 if a woman worked 100% of her work-hours remotely in the previous day. All regressions control for the log of woman's previous-day work hours, a cubic function of the woman's age, her citizenship and marital status, her race and ethnicity, whether she has children, and her male partner's age, race, ethnicity, education, and labor earnings (in logs). The regressions also control for time-varying characteristics of metropolitan areas described in Section 3. Region and year fixed effects are also included. All regressions are weighted using ATUS individual weights, and standard errors are clustered at the MSA level. The figure also includes 95% confidence intervals. The coefficients and standard errors are reported in Table A10 in the Online Appendix.

Table 6 provides additional evidence in line with the hypothesis that the availability of remote work enables mothers with access to this benefit to reconcile work with family responsibilities in areas characterized by high time-costs of commuting.

The table shows the gap between college and non college graduate mothers in the daily amount of time allocated to different activities in low-congestion areas (columns 1 and 3) and in high-congestion areas (columns 2 and 4). The gap is estimated among all women in panel (a) and among working women in panel (b). All gaps represent the coefficient of a dummy variable indicating whether a woman has a college degree, in a regression model whose dependent variable is the time (in minutes) allocated in the previous day to the activity indicated in each row. All regression models control for the characteristics of women and of their partners', for time-varying characteristics of women's metropolitan areas of residence, and region and year fixed effects.

As shown in line 1, college graduate mothers work more per day than mothers without a college degree, especially in highly congested areas. The gap, which is as large as 41.35 minutes per day in highly congested areas, is explained by the higher labor force participation rate of college graduate mothers. Considering only working individuals, non college graduate mothers work at least as much, per day, as mothers with a college degree. The gap between college and non-college graduate mothers in the daily amount of time spent commuting (line 2) follows a similar pattern. If anything, working mothers without a college degree spend more time commuting than college graduate mothers, but this difference is not statistically significant.

The workdays of college graduate mothers also appear more flexible than the workdays of mothers without a college degree, especially in highly congested areas. As shown in line (3), in highly congested areas, college graduate mothers spend 16 more minutes per day working remotely than mothers with no college degree. This gap increases to almost 22 minutes per day when looking only at working women. No differences along education lines, instead, emerge in remote-work time among mothers living in low-congestion areas (columns 1 and 3).

College graduate women appear to be using remote work options to allocate more time to care and housework in highly congested metropolitan areas. In the bottom half of Table 6, we examine differences in time spent on housework and care for others (line 4), time spent working while providing care as a secondary activity (line 5) and time spent on leisure (line 6). Interestingly, the education gap in remote-work time among working mothers in highly congested areas is similar in magnitude to the gap in the daily amount of time spent in care and housework activities (line 4, column 4), and to the gap in the daily amount of time spent working while providing care to others as a secondary activity (line 5, column 4). Among working mothers living in highly congested areas, college graduates spend 21 more minutes than non college graduates in care and housework. While not statistically significant, the education gap in care-and-housework time is reversed when non-working women are included in the estimating sample (line 4, columns 1 and 2).

The estimated education gaps in the daily amount of time spent working while providing care for others (line 5) corroborate the hypothesis that college graduate women may rely on remote work in areas characterized by long commuting times to reconcile work and family obligations. Among working mothers living in highly congested areas, college graduate women spend 19 minutes more than non-college graduate women per day working while providing care for others. The magnitude of this gap also roughly corresponds to the education gap among working mothers in remote-work time.

If the possibility of providing care while working remotely is less likely to be available for non-college graduate women, then long commuting times would be more likely to push non-college graduate women out of the labor force in the presence of children. College graduate mothers, instead, may be more likely to remain in the workforce and substitute remote work for in-person work. This choice, however, may itself come at a cost. As line 6 shows, college graduate women in the latter group devote systematically less time to self-care and leisure than women with lower levels of education, and especially so when working in highly congested areas.

Table 6: Remote work and women's time use

	(a) All		(b) Working	
	(1)	(2)	(3)	(4)
	Low	High	Low	High
(1) Time working (min.)	27.26**	41.35***	-10.84	-3.53
	(8.76)	(9.86)	(9.83)	(10.29)
R-squared	0.200	0.209	0.167	0.158
Diff $\neq$ 0 (H0: Diff=0) p-value	0.254		0.604	
(2) Time commuting (min.)	1.94+	3.90**	-1.34	-0.49
	(1.14)	(1.26)	(1.99)	(2.64)
R-squared	0.103	0.123	0.079	0.100
Diff $\neq$ 0 (H0: Diff=0) p-value	0.227		0.789	
(3) Time work from home (min.)	-0.04	16.06***	-6.79	21.70*
	(3.92)	(3.90)	(7.97)	(7.97)
R-squared	0.063	0.059	0.099	0.092
Diff $\neq$ 0 (H0: Diff=0) p-value	0.003		0.008	
(4) Time care (min.)	-5.03	-14.48	4.62	21.21*
	(8.01)	(8.67)	(9.59)	(9.99)
R-squared	0.124	0.125	0.129	0.156
Diff $\neq$ 0 (H0: Diff=0) p-value	0.400		0.215	
(5) Time work and care (min.)	0.37	12.92**	-4.97	19.02**
	(3.52)	(3.66)	(7.30)	(6.60)
R-squared	0.027	0.044	0.047	0.074
Diff $\neq$ 0 (H0: Diff=0) p-value	0.012		0.013	
(6) Time leisure (min.)	-17.47**	-20.52**	1.19	-10.64*
	(6.68)	(6.07)	(8.12)	(4.85)
R-squared	0.144	0.148	0.143	0.163
Diff $\neq$ 0 (H0: Diff=0) p-value	0.738		0.213	
Obs.	4228	3976	1599	1421

Notes: American Time Use Survey 2005-2019. Sample selection described in Section 2. Panel (b) sample is restricted to women who report to have worked in the previous day. Samples in columns 1 (2) and 3 (4) include women living in low-congestion (high-congestion) metropolitan areas. These groups are constructed as described in Figure 5. Each line plots the coefficient associated to a dummy variable indicating whether a woman has a college degree or not, in a regression whose dependent variable is the time (in minutes) devoted in the previous day to the line-specific activity. In line 1 the dependent variable is the daily time spent working for pay. The other dependent variables are: the time spent commuting for work (line 2), the time spent working from home (line 3), the time devoted to housework or to care for others (line 4), the time spent working while providing care as a secondary activity (line 5) and the time devoted to leisure and self-care (line 6). All models include the control variables listed in Figure 5 notes. All models are weighted using ATUS individual weights. Robust standard errors are shown in parentheses. p-value $<$ 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

5 Alternative mechanisms

It is possible that alternative mechanisms can explain the stronger impact of long commuting times on the labor supply of low-education mothers. In this section, we provide evidence suggesting that our results are not driven by underlying differences in culture and gender norms between women with different levels of education, or by the unobserved relationship between congestion and the availability and affordability of childcare services.

5.1 Culture and gender norms

Farré, Jofre-Monseny, and Torrecillas (2023) provide evidence suggesting that gendered social norms, which typically associate family care responsibilities with women, may explain why commuting costs decrease the labor force participation of married women while not impacting married men's labor supply. To do so, they show that congestion has especially negative effects on the labor force participation of immigrant women in the US who were born in countries with more conservative attitudes towards gender roles.

We follow a similar epidemiological approach to study whether differences in gender norms between college and non-college graduate women may explain our results. Specifically, we restrict our ACS sample to women born in the United States, and use the state-of-birth labor force participation of prime-age women in 1990 as a proxy of women's cultural attitude towards gender roles. We then split women into two groups, those who were born in states with below-average women's labor force participation in 1990 ("low labor supply") and those born in states with above-average women's labor force participation in 1990 ("high labor supply").[8]

We estimate a version of regression 1 that includes the low labor supply indicator defined above, its interaction with the MSA-specific commuting-time variable, and several variables capturing the time-varying industrial structure of women's state of residence. We suppose that women born in states with lower female labor force participation may hold more traditional or conservative beliefs regrading gender roles and be less likely to

[8]This approach is based on the intuition of Fernández and Fogli (2009), who use the labor force participation and fertility in the countries of origin of second-generation American women to study the effects of culture on women's labor-market and family-formation decisions.

participate in the labor force compared to those born in states with historically high female labor supply. Furthermore, if congestion reinforces gender norms, the impact of congestion should be larger in magnitude for women born in states with low women's labor force participation.

We estimate the model through OLS and 2SLS, separately for college graduate mothers and for mothers with no college degree. All models include the full set of controls and fixed effects described in Section 3.

Table 7: Congestion and mother's LFP, by 1990 female LFP in women's state-of-birth

	(a) No College		(b) College	
	(1)	(2)	(3)	(4)
	Reg FE	Reg FE/IV	Reg FE	Reg FE/IV
Z(MSA CT)	-0.0137***	-0.0253***	-0.0027	-0.0025
	(0.0030)	(0.0060)	(0.0033)	(0.0065)
Low LS	-0.0173***	-0.0244***	-0.0122**	-0.0086+
	(0.0044)	(0.0061)	(0.0043)	(0.0051)
Z(MSA CT)* Low LS	-0.0013	0.0091	0.0028	-0.0011
	(0.0049)	(0.0086)	(0.0031)	(0.0046)
Obs	374519	374519	314020	314020
R-squared	0.097	0.092	0.137	0.126
Under-ID (K–P LM) Chi-Sq.		31.16		29.82
Weak-ID (K–P rk Wald Chi-Sq)		15.65		20.12
Hansen J p-value		0.683		0.374
Andreson-Rubin F p-value		0.001		0.687
DWH p-value		0.041		0.434

Notes: American Community Survey 2005-2019. Sample includes US-born women selected as described in Section 2. The table shows selected estimated coefficients from a modified version of regression 1 that includes the following two variables in addition to the control variables listed in Section 3: a dummy variable taking the value 1 if a woman was born in a US state where the labor force participation of prime-age women in 1990 Census data was below the year-specific cross-state average ("Low LS"); an interaction term between the state-of-birth LFP dummy and the MSA-specific congestion measure ("Z(MSA CT)* Low LS"). We also account for differences across states of residence in industrial structure by controlling for the state-year-specific shares of prime-age men employed in different 1-digit industries. The model is estimated separately for non-college graduate women (panel a) and for college women (panel b) via OLS (columns 1 and 3) and via 2SLS (columns 2 and 4) using both Duranton and Turner (2011) instruments and interactions between the instruments and the "Low LS" dummy. All models include region, year, and region-times-year fixed effects, and are weighted using individual weights. Standard errors are clustered at the MSA level. The table also includes results of all IV diagnostic tests described in table 1 notes. p-Value $<$ 0.1 (+), 0.05 (*), 0.01 (**), 0.001 (***).

Table 7 shows that, irrespective of their level of education, mothers born in states with low women's labor supply are themselves less likely to be in the labor force. Congestion, however, does not disproportionately affect the labor force participation of women with supposedly more traditional cultural attitudes. The coefficient of the interaction term between congestion and the low-labor supply indicator is small in magnitude, not statistically significant in all specifications, and not always negative in sign. Finally, explicitly accounting for the relationship between cultural background and labor supply behavior of women does not change the finding that congestion has a negative impact on the labor supply of mothers with no college education, while not affecting the probability that college graduate mothers participate in the workforce. These results suggest that the stronger effect of congestion on low-education women is not driven by supposed unobserved differences in attitudes towards gender roles between women with different levels of education. As shown in Table A12 in the Online Appendix, we reach similar conclusions when estimating versions of regression 1 that include an alternative proxy of culture and gender norms.

5.2 Childcare affordability

Childcare prices tend to be higher in larger and more congested metropolitan areas (Moreno Maldonado, 2022). If market-provided childcare is more expensive in denser urban areas, and therefore childcare becomes increasingly unaffordable with congestion, it could be that high childcare prices, rather than long commuting times, induce mothers with low earning-potential or in low-income households to drop out of the labor force. If this alternative mechanism explains why congestion has an especially strong impact on low-education mothers, then controlling for the MSA-specific affordability of childcare services in regression 1 should induce a substantial decline in the magnitude and statistical significance of the estimated impact of congestion on the labor force participation of mothers without a college degree.

To study whether our results are driven by some unobserved correlation between congestion and childcare affordability, we estimate a version of regression 1 that controls for the MSA-specific share of low-education immigrant workers in the local labor force, and interpret this variable as a proxy of childcare affordability. Cortés and Tessada (2011) show that immigrant workers with low levels of education are highly concentrated in jobs that provide substitutes for household production (such as childcare), and that a stronger presence of immigrants in local labor markets increases the work-hours and the probability of working long-hours of high-income women. We follow this intuition and control for the city-specific immigrant labor force as a measure of the local availability of childcare services. We use public-use data from the 2000 census to proxy the MSA-specific share of migrant workers in the 2000-2010 decade, and data from the 2010 five-year ACS survey to proxy for the MSA-specific share of migrant workers in the 2011-2019 decade.

Table 8: Congestion and mother's LFP - accounting for childcare prices

	(a) No College		(b) College	
	(1)	(2)	(3)	(4)
	Reg FE	Reg FE/IV	Reg FE	Reg FE/IV
Z(MSA CT)	-0.0101**	-0.0217*	-0.0010	0.0004
	(0.0031)	(0.0093)	(0.0029)	(0.0073)
Obs	364611	364611	307380	307380
R-squared	0.095	0.087	0.126	0.112
Under-ID (K–P LM) Chi-Sq.		18.20		18.73
Weak-ID (K–P rk Wald F)		13.20		15.37
Hansen J p-value		0.951		0.182
Andreson-Rubin F p-value		0.000		0.066
DWH p-value		0.144		0.505

Notes: American Community Survey 2005-2019. Sample includes US-born women selected as described in Section 2. The table shows the effect of congestion on the labor force participation of mothers from a version of regression 1 that controls for the share of low-education immigrants in each MSA labor force. We construct the variable as described in Cortés and Tessada (2011). To ensure that the variable is constructed with a sufficient number of observations, the share of immigrants within the labor force of each MSA is computed using the 2008 5-year public-use ACS for all years following 2009, and using the 1% public-use 2000 census data for all years between 2005 and 2009. The modified 1 is estimated via OLS (columns 1 and 3) and via 2SLS (columns 2 and 4) separately for non-college (panel a) and college educated (panel b) mothers. In column-2 and column-4 models, commuting time (congestion) is instrumented using both Duranton and Turner (2011) instruments, while the share of immigrants in the local labor force is instrumented using the Cortés and Tessada (2011) instrument. The latter variable is the number of low-education immigrants in the local area predicted by the past distribution of immigrants from different countries of origin. Details on the instrument construction can be found in Cortés and Tessada (2011). We construct the instrument using the 5% state public-use census data. Standard errors clustered at the MSA level. p-Value $<$ 0.1 ($^+$), 0.05 (*), 0.01 (**), 0.001 (***).

We estimate the augmented version of regression 1 through both OLS and 2SLS. To address the potential endogeneity in the allocation of migrants across metropolitan areas, we also follow Cortés and Tessada (2011) and Cortés and Pan (2019) in instrumenting the current share of migrants in each MSA-specific labor market through the concentration of immigrants predicted by the cross-MSA distribution of past flows of immigrants from different countries. We construct this instrument using data from the public-use 1990 census.

We report the estimated effect of congestion on mothers' labor supply in Table 8, and show that controlling for the supply of market-provided childcare services through the share of immigrant workers in local labor markets does not affect our main results. This evidence should reduce concerns that the negative impact of congestion on the labor force participation of low-education mothers is a mechanical effect of some underlying relationship between congestion and childcare affordability.

As an additional check that the unobserved variation in childcare affordability across metropolitan areas does not drive the asymmetry in the effect of congestion on the labor force participation of college educated and non-college educated mothers, we use county-level data from the National Database of Childcare Prices collected by the Women's Bureau of the US Department of Labor. The data are available for years 2008 to 2022 and contain year-specific information regarding the median cost of (full-time) family-based childcare services and of center-based childcare services for children of different ages (infant, toddler, preschool) in each county.

We use the county codes for identifiable counties in our ACS sample of women to merge each county to the year-county-specific information on childcare prices. We then classify metropolitan areas into two groups: high-cost MSAs are areas that include at least one county in the top-50 percentiles of the year-specific distribution of childcare costs; low-cost MSAs are metropolitan areas that do not include any county in the top-50 percentiles of the year-specific distribution of childcare costs. We construct a separate indicator variable for local childcare costs for each type of childcare for which information is available in the National Database of Childcare Prices.

In Table 9 we show the results we obtain by estimating regression model 1 including all the controls listed in Section 3 plus the indicator that classifies MSAs in two groups based on the price of infant family-based childcare. We estimate the regression model separately for mothers without a college degree and for college graduate mothers. The table shows both the estimated impact of congestion on mothers' participation in the labor force and the estimated relationship between childcare affordability and participation in the labor force. In columns (1), (2), (4) and (5) we estimate regression 1 via OLS. In columns (1) and (4) we include the dummy for high childcare cost. In columns (2) and (5), instead, we predict the probability that the cost of childcare in each MSA is high by regressing the high-cost dummy variable on the Cortés and Tessada (2011) instrument. In columns (3) and (6) we control for the same predicted variable included in regressions (2) and (5), and instrument commuting times using the Duranton and Turner (2011) instruments.

Using the MSA-specific 1990 flows of low-education immigrants to predict whether the cost of childcare is high in a certain MSA in a given year should address the clear endogeneity of local childcare prices, which are arguably jointly determined with local mothers' labor supply. In places with a high participation of mothers in the labor force, in fact, the demand for market-provided childcare services is likely higher, thus inducing a higher cost of childcare in the area. Furthermore, higher childcare prices may reflect higher wages in childcare services, a predominantly female sector, which may encourage women's labor force participation through channels that are not related to the possibility of outsourcing within-household care work to the market. Measuring the affordability of childcare services in a local area as a function on the share of immigrants in its labor market predicted by the historical cross-MSA allocation of migrants from different countries, should ensure that the cross-MSA variation in childcare prices reflects heterogeneity in childcare-services supply, thus in their availability and affordability.

Consistent with this intuition, we find that the labor force participation of mothers of all levels of education is lower in areas where the childcare prices predicted by the historical allocation of migrants from different countries are high (columns 2-3 and 5-6 in Table 9). We find similar results whether estimating the model via OLS or via 2SLS. In line with recent findings by Araujo, McBride, and Sandler (2025), these results suggest that the lack of affordable childcare options is associated with labor-supply declines for women with children.

While high childcare costs negatively affect mothers' labor supply, controlling for the affordability of childcare services does not substantially affect the magnitude and significance of the negative impact of congestion on non-college educated mothers. The estimated effects of congestion reported in columns 2 and 3 of Table 9 are only slightly smaller in magnitude than the baseline coefficients reported in Table 1 (columns 1 and 4, panel b), but remain economically and statistically significant. Controlling for childcare affordability, a 3.5-minute (one standard deviation) increase in local area commuting times lead to a 1.36-to-1.95 percentage-point decline in the probability that mothers without a college degree participate in the workforce. As far as college graduate mothers are concerned, accounting for childcare affordability does not alter the finding that their labor force participation rate does not substantially change with congestion.

Tables A13 to A15 in the Online Appendix show that the results we obtain are unaffected when constructing the childcare-affordability indicator using the cost of center-based childcare for infants or the cost of (family-based or center-based) childcare for toddlers.

Table 9: Congestion and mother's LFP - accounting for family-based infant care cost

	(a) No College			(b) College		
	(1) Reg. FE Raw	(2) Reg. FE Pred.	(3) Reg. FE IV.Pred.	(4) Reg. FE Raw	(5) Reg. FE Pred.	(6) Reg. FE IV.Pred.
Z(MSA CT)	-0.0185***	-0.0136***	-0.0195*	-0.0014	0.0024	-0.0023
	(0.0041)	(0.0038)	(0.0098)	(0.0041)	(0.0039)	(0.0081)
High childcare cost	0.0282**	-0.2518**	-0.1967+	-0.0304**	-0.2114**	-0.1636+
	(0.0106)	(0.0759)	(0.1127)	(0.0100)	(0.0740)	(0.0985)
Obs	372380	372380	372380	311257	311257	311257
R-squared	0.131	0.131	0.123	0.132	0.132	0.126
(K–P LM) Chi-Sq.			27.90			30.82
(K–P rk Wald F)			17.50			30.29
Hansen J p-value			0.675			0.102

Notes: American Community Survey 2008-2019. Sample includes a subsample of women selected as described in Section 2 for years in which childcare-cost information is available in the National Database of Childcare Prices of the DOL Women's Bureau. The table shows selected estimated coefficients from a modified version of regression 1 that includes a dummy variable taking the value 1 for MSAs containing at least one county where the cost of full-time family-based infant care is high (in the top-50 percent of the year-specific cross-county distribution of childcare cost). The regression includes the raw dummy variable in columns (1) and (4). Models in columns (2) and (3), (5) and (6), control for the probability that an MSA includes at least one high-childcare-price county predicted using the Cortés and Tessada (2011) instrument. In columns (3) and (6) the model is estimated via 2SLS using both Duranton and Turner (2011) instruments. All regressions are weighted using individual weights, and standard errors are clustered at the MSA level. p-Value $<$ 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

6 Conclusion

In this paper, we show that the negative effect of congestion on women's labor supply first documented in the US by Black, Kolesnikova, and Taylor (2014) and by Farré, Jofre-Monseny, and Torrecillas (2023) disproportionately concerns mothers without a college degree. For college graduate mothers, the probability of participating in the workforce is not affected by congestion, while long commuting times have a large, negative effect on the labor force participation rate of mothers with no college education. Among women in this group, a 3.5-minute increase in average city-specific commuting time leads to a decline in mothers' labor force participation of up to 2.6 percentage points. This implies that a 10-minute increase in commuting time decreases the labor force participation rate of mothers without a college degree by up to 8 percentage points.

We then show that the availability of remote work in jobs typically performed by highly educated workers can explain why the labor supply of non-college mothers is especially sensitive to commuting times. The possibility to work from home shields college graduate mothers from the mobility constraints induced by congestion, which become increasingly binding in the presence of children for non-college educated women with limited access to this type of workplace flexibility, especially if their partners also work inflexible jobs.

Our results suggest that remote work can alleviate mothers' mobility constraints and reduce the time costs of congestion, thus encouraging their labor force participation. The benefit of remote work, however, may only be experienced by women employed in certain industries, and in jobs that can be performed remotely, often employing mostly highly educated workers (Bartik, Cullen, Glaeser, Luca, & Stanton, 2020; Buckman, Barrero, Bloom, & Davis, 2025). To the extent that women with lower levels of education remain disproportionately represented in high-contact, inflexible occupations often requiring in-person interactions (Albanesi & Kim, 2021), instead, their labor supply may remain exposed to the mobility constraints induced by work-family tradeoffs, especially in congested areas. Such constraints can be mitigated if partners with access to remote-work use their work flexibility to provide housework and care for household members.

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Supplementary Material

Online Appendix

to

Understanding the congestion child penalty: can remote work attenuate it?

Not for Publication

A Additional tables and figures

A.1 Tables

Table A1: Marriage, family formation and geographical location - Women by education

	(1) No College	(2) College	Diff. = 0 p-value
Married	89.8	95.5	0.000
College-grad partner	19.3	98.0	0.000
Employed partner	94.8	97.7	0.000
Has child	79.7	71.6	0.000
Age first birth	25.2	29.7	0.000
N. children	2.1	2.0	0.000
Has two or more children	71.2	68.5	0.000
Age youngest child	5.4	4.2	0.000
Has one K12-age child	42.1	31.9	0.000
Moved from place of birth	55.9	62.4	0.000
Lives in Top-20 MSAs	54.7	63.0	0.000

Notes: American Community Survey 2005-2019. Sample selection is described in Section 2. All differences in means are estimated using individual weights and heteroskedasticity-robust standard errors. For categorical variables, the numbers in the table represent percentages belonging to the category indicated in the corresponding row. The characteristics of mothers and of children are computed for the subsamples of college and non-college graduate women with at least one child.

Table A2: Effect of one std. dev. increase in MSA commuting time on mothers' LFP - All women with heterogeneity in effect of congestion by education

	(1) Region FE	(2) Reg. FE/IV Railroad	(3) Reg. FE/IV Highway	(4) Reg. FE/IV Both
Z(MSA CT)	-0.0165***	-0.0187**	-0.0227***	-0.0211***
	(0.0032)	(0.0067)	(0.0063)	(0.0061)
Z(MSA CT)*College	0.0153***	0.0210***	0.0150**	0.0164**
	(0.0029)	(0.0063)	(0.0051)	(0.0050)
Obs	990703	990703	990703	990703
R-squared	0.145	0.098	0.098	0.098
Under-ID (K–P LM) Chi-Sq.		20.94	17.15	21.03
Weak-ID (K–P rk Wald F)		7.69	28.32	14.60
Hansen J p-value				0.114
Andreson-Rubin F p-value		0.003	0.018	0.002
DWH p-value		0.602	0.315	0.507

Notes: American Community Survey 2005-2019. Sample selection is described in Section 2. Z(MSA CT) is the standardized value of MSA-specific commuting time in year t. In every year, the cross-MSA average of Z(MSA CT) equals 0 and has a standard deviation of 1. The coefficients in column (1) are estimated via OLS, the coefficients in columns (2), (3) and (4) are estimated via 2SLS. All IV columns use instruments interacted with the college dummy variable. All models include all control variables listed in Section 3. Standard errors are clustered at the MSA level. IV diagnostic tests described in Table 1 notes. p-Value $<$ 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

Table A3: Effect of congestion on mothers' LFP - Unweighted

	(1) Region FE	(2) Reg. FE/IV Railroad	(3) Reg. FE/IV Highway	(4) Reg. FE/IV Both
(a) No College	-0.0173***	-0.0204**	-0.0254***	-0.0235***
	(0.0030)	(0.0069)	(0.0065)	(0.0063)
Obs	564468	564468	564468	564468
R-squared	0.138	0.092	0.092	0.092
Under-ID (K–P LM) Chi-Sq.		21.80	17.93	22.26
Weak-ID (K–P rk Wald F)		15.83	54.79	26.61
Hansen J p-value				0.349
Andreson-Rubin F p-value		0.016	0.002	0.006
DWH p-value		0.579	0.175	0.311
(b) College	-0.0003	0.0002	-0.0067	-0.0047
	(0.0034)	(0.0062)	(0.0059)	(0.0059)
Obs	426235	426235	426235	426235
R-squared	0.145	0.117	0.117	0.117
Under-ID (K–P LM) Chi-Sq.		17.46	15.84	17.76
Weak-ID (K–P rk Wald F)		13.31	78.64	40.07
Hansen J p-value				0.045
Andreson-Rubin F p-value		0.973	0.279	0.089
DWH p-value		0.897	0.110	0.566

Notes: American Community Survey 2005-2019. Sample selection is described in Section 2. Panels (a) and (b) report the β coefficient estimated separately for all mothers, mothers without a college degree, and mothers with a college degree. Coefficients in column (1) are estimated via OLS, coefficients in columns (2), (3) and (4) are estimated via 2SLS. All models include all control variables listed in Section 3. Standard errors are clustered at the MSA level. IV diagnostic tests described in Table 1 notes. p-Value < 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

Table A4: Effect of congestion on mothers' LFP - Different sets of control variables

	(1) No controls	(2) Own characteristics	(3) Partner's characteristics	(4) MSA characteristics
(a) All	-0.0204***	-0.0117***	-0.0074**	-0.0110***
	(0.0043)	(0.0026)	(0.0027)	(0.0029)
Obs	990703	990703	990703	990703
R-squared	0.009	0.122	0.145	0.145
(b) No College	-0.0294***	-0.0158***	-0.0131***	-0.0180***
	(0.0044)	(0.0030)	(0.0030)	(0.0032)
Obs	564468	564468	564468	564468
R-squared	0.013	0.121	0.141	0.141
(c) College	-0.0114**	-0.0071**	0.0006	-0.0000
	(0.0039)	(0.0026)	(0.0028)	(0.0034)
Obs	426235	426235	426235	426235
R-squared	0.008	0.115	0.148	0.148

Notes: American Community Survey 2005-2019. Sample selection is described in Section 2. Panels (a), (b) and (c) report the β coefficient estimated separately for all mothers, mothers without a college degree, and mothers with a college degree. All coefficients are estimated via OLS, with year, region and year $\times$ region fixed effects. The control variables included in the models are listed in Section 2. Specifically, column (1) only includes fixed effects, column (2) adds women's characteristics and dummies for number of children and for the eldest child's age, column (3) includes partner's characteristics in addition to own characteristics, column (4) adds time-varying MSA characteristics as well as own characteristics and partner's characteristics. All regressions are weighted using individual weights, and standard errors are clustered at the MSA level. p-Value < 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

Table A5: Effect of congestion on mothers' LFP - ATUS (18-44 y.o.)

	(a) Reg. FE		(b) Reg. FE/IV Both	
	(1)	(2)	(3)	(4)
	No Col	Col	No Col	Col
Z(MSA CT)	-0.0318**	-0.0144	-0.0609**	-0.0347
	(0.0104)	(0.0095)	(0.0219)	(0.0222)
Obs.	4070	4134	4070	4134
R-squared	0.119	0.115	0.117	0.114
Diff < 0 (H0: Diff=0) p-value	0.075		0.201	
Under-ID (K–P LM) Chi-Sq.			17.21	14.91
Weak-ID (K–P rk Wald F)			18.93	19.48
A-R F p-value			0.015	0.000
Hansen J p-value			0.954	0.155
DWH p-value			0.120	0.043

Notes: American Time Use Survey 2005-2019. Sample selection is described in Section 2. Coefficients in columns (1) and (2) are estimated via OLS, coefficients in columns (3) and (4) are estimated via 2SLS using both Duranton and Turner (2011) instruments. All models include all control variables listed in Section 3. Standard errors are clustered at the MSA level. Because the regressions do not include any variable from time-use diaries, time-use weights are not used. IV diagnostic tests described in Table 1 notes. p-Value < 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

Table A6: Effect of congestion on mothers' LFP - ATUS (18-54 y.o.)

	(a) Reg. FE		(b) Reg. FE/IV Both	
	(1)	(2)	(3)	(4)
	No Col	Col	No Col	Col
Z(MSA CT)	-0.0243*	-0.0101	-0.0624**	-0.0237
	(0.0098)	(0.0075)	(0.0207)	(0.0184)
Obs.	5495	5585	5495	5585
R-squared	0.100	0.096	0.097	0.095
Diff < 0 (H0: Diff=0) p-value	0.103		0.082	
Under-ID (K–P LM) Chi-Sq.			16.84	15.21
Weak-ID (K–P rk Wald F)			18.69	22.24
A-R F p-value			0.001	0.002
Hansen J p-value			0.595	0.195
DWH p-value			0.043	0.036

Notes: American Time Use Survey 2005-2019. Sample selection, described in Section 2, is modified to include women between 18 and 54 years-old. Coefficients in columns (1) and (2) are estimated via OLS, coefficients in columns (3) and (4) are estimated via 2SLS using both Duranton and Turner (2011) instruments. All models include all control variables listed in Section 3. Standard errors are clustered at the MSA level. Because the regressions do not include any variable from time-use diaries, time-use weights are not used. IV diagnostic tests described in Table 1 notes. p-Value < 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

Table A7: Partner's remote work by partner education and occupation

	(1) All	(2) No College	(3) College	(4) Diff. = 0 p-value
Occ. in 5th RW Quintile	8.08	6.57	8.59	0
Occ. in 4th RW Quintile	3.02	2.33	3.61	0
Occ. in 3rd RW Quintile	1.79	1.58	2.29	0
Occ. in 2nd RW Quintile	1.04	.83	1.35	0
Occ. in 1st RW Quintile	.71	.58	.97	0

Notes: American Community Survey 2005-2019. The table shows the shares of male partners of women in the sample in Section 2, who are employed in different quintiles of the remote-work incidence distribution of occupations and work from home. Shares of remote-working male partners are calculated separately for all male partners (column 1), non-college male partners (column 2) and college graduate male partners (column 3). P-values for differences in proportions between columns (2) and (3) are reported in column (4). All statistics are computed using partner-specific individual weights.

Table A8: Top-20 Occupations by remote-work incidence

	Pctile 2005-19	Pctile 2021-23	RW share 2005-19	RW share 2021-23
Writers and authors	100	100	0.307	0.571
Sales engineers	100	99	0.233	0.537
Farmers (owners and tenants)	100	75	0.229	0.183
Art makers	99	95	0.181	0.396
Business and promotion agents	99	92	0.183	0.330
Photographers	99	83	0.186	0.254
Management analysts	98	99	0.179	0.520
Musician or composer	98	85	0.145	0.273
Managers of properties and real estate	98	80	0.141	0.223
Real estate sales occupations	97	82	0.135	0.252
Art/entertainment performers and rel.	97	81	0.122	0.234
Child care workers	97	71	0.129	0.154
Farm managers (no hort. farms)	97		0.122	
Editors and reporters	96	96	0.115	0.430
Designers	96	93	0.111	0.365
Actors, directors, producers	96	90	0.110	0.314
Economists, market, survey res.	95	97	0.109	0.438
Manag. and specialists in marketing, ad., p.r.	95	95	0.107	0.388
Salespersons, n.e.c.	95	87	0.108	0.283
Advertising and related sales jobs	94	96	0.101	0.421

Notes: American Community Survey 2005-2019. The table shows the 20 occupations (3-digit 1990 Census classification) with highest remote-work incidence between 2005 and 2019. Occupation-specific remote-work incidence is calculated as the share of prime-age men employed in the occupation between 2005 and 2019 who report to not usually commute due to working from home. Statistics are calculated using individual weights. The table also shows remote-work incidence within the occupations calculated between 2021 and 2023.

Table A9: Congestion and mothers' LFP by partner's occupation remote-work incidence

	(a) High RW		(b) Low RW	
	(1) R FE	(2) R FE/IV	(3) R FE	(4) R FE/IV
(a) No College	-0.0073*	-0.0109	-0.0222***	-0.0288***
	(0.0035)	(0.0073)	(0.0034)	(0.0067)
Obs	128539	128539	409053	409053
R-squared	0.136	0.101	0.152	0.089
Under-ID (K–P LM) Chi-Sq.		22.01		22.56
Weak-ID (K–P rk Wald F)		31.59		25.26
Hansen J p-value		0.581		0.291
Andreson-Rubin F p-value		0.349		0.002
DWH p-value		0.628		0.347
(b) College	-0.0001	-0.0027	-0.0000	-0.0024
	(0.0043)	(0.0077)	(0.0029)	(0.0055)
Obs	201868	201868	215372	215372
R-squared	0.150	0.113	0.150	0.120
Under-ID (K–P LM) Chi-Sq.		17.65		20.79
Weak-ID (K–P rk Wald F)		47.43		37.77
Hansen J p-value		0.612		0.033
Andreson-Rubin F p-value		0.868		0.005
DWH p-value		0.712		0.436

Notes: American Community Survey 2005-2019. The figure represents the β coefficient in regression model 1 estimated via OLS and via IV using Duranton and Turner (2011) instruments, separately for non-college (row a) and college graduate (row b) women with children cohabiting with men in high-remote-work incidence occupations (panel a) and with men in low-remote-work incidence occupations (panel b). All regression models include all control variables listed in Section 3 and the additional characteristics of partners' occupation listed in Section 4.1. Columns (3) and (6) exclude region fixed effects and include MSA fixed effects. All regressions are weighted using individual weights, and standard errors are clustered at the MSA level. p-Value < 0.1 ($^+$), 0.05 (*), 0.01 (**), 0.001 (***).

Table A10: Gap between college and non-college graduate women in remote work incidence

	(1) Low-congestion MSAs	(2) High-congestion MSAs
	ATUS	
(a) 1/2 hour RW	0.0875***	0.1041**
	(0.0233)	(0.0300)
(b) % RW hours	0.0442*	0.1045***
	(0.0197)	(0.0215)
(c) Full day RW	0.0188	0.0937***
	(0.0195)	(0.0197)
Obs	1882	1677
R-squared	0.117	0.184
	ACS	
(d) Works from home	0.0021	0.0079***
	(0.0019)	(0.0020)
Obs	480332	464542
R-squared	0.015	0.013

Notes: Panels (a), (b) and (c): American Time Use Survey 2005-2019. Panel (d): American Community Survey 2005-2019. Sample selection is described in Section 2. The sample is restricted to working women report in their time-use diaries to have worked in the previous day. Dependent and control variables in rows (a), (b) and (c) are described in Figure 5 notes. Panel (d) model estimates the gap in the probability of usually working remotely between college and non-college graduate women in ACS data, in low-congestion MSAs and in high-congestion MSAs. The model includes the same control variables as models (a), (b) and (c). All regressions are weighted using individual weights, and standard errors are clustered at the MSA level. p-Value < 0.1 ($^+$), 0.05 (*), 0.01 (**), 0.001 (***).

Table A11: Remote work incidence among college and non-college graduate women in high- and low-congestion areas

	(1) Low-cong. MSAs	(2) High-cong. MSAs	(3) Diff =0 p-value
		(a) No College	
Hours worked	7.236	7.366	.411
1/2 h remote work	.079	.088	.55
% RW hours	.103	.086	.337
Full day RW	.056	.036	.079
		(b) College	
Hours worked	7.051	7.118	.662
1/2 h remote work	.193	.228	.053
% RW hours	.192	.232	.036
Full day RW	.098	.121	.068

Notes: American Time Use Survey 2005-2019. Sample selection is described in Section 2. Panel (a) shows different measures of the average incidence of remote work among non-college educated women in low-congestion (column 1) and high-congestion (column 2) areas. Panel (b) shows the same statistics for college graduate women. In both panels, the first line indicates the group-specific average amount of daily hours worked in the previous day. The second line indicates the share of women who, in the previous day, worked at least one-half hour from home. The third line indicates the share of work hours worked remotely in the previous day. The fourth line indicates the share of women who worked fully remotely in the previous day. In both panels, column 3 shows the p-value of tests for the statistical difference between the line-specific averages (or proportions).

In Table A12 we estimate the effect of congestion on mothers' labor force participation controlling for alternative measures of gender attitudes. Specifically, we classify women's state of birth in three groups based on how abortion access changed following the 2022 overturn by the United States Supreme Court of the Roe v. Wade decision of 1973. US states followed different paths following the 2022 decision of the US Supreme Court. Some states implemented laws to grant or protect abortion access, while in other states access to abortion was restricted if not prohibited. The Center for Reproductive Rights (2025) is tracking abortion access across US states. We hypothesize that states with historically more conservative or traditional attitudes towards gender roles may have been less likely to implement abortion-access expansions or protections after 2022. We follow the Center for Reproductive Rights (2025) classification of US states to divide US-born women in our ACS sample in three groups: women born in States where abortion access was expanded or protected after 2022, women born in States where abortion access was not protected after 2022, women born in States with restricted abortion access following 2022. We include dummies for non-protected or restricted abortion access in regression 1, together with their interaction with the MSA-specific commuting-time variable. We estimate the model through OLS and 2SLS separately for non-college mothers and for mothers with a college degree.

The estimated coefficients of interest of the expanded regression 1 are reported in Table A12. The coefficients of dummies for states of birth where abortion access was not protected or restricted after 2022 are negative for mothers of all levels of education. The effect of congestion, however, does not change between women born in states where abortion access was expanded or protected after 2022 and women born in states with restricted abortion access. The gap in the effect of congestion on labor supply between college and non-college mothers is also unaffected by the inclusion of variables proxying women's gender-role attitudes using abortion access in their state of birth. These results further suggest that underlying differences in culture and gender norms should not be the primary reason why congestion mostly impacts the labor supply of non-college educated mothers.

Table A12: Congestion and mother's LFP by state-of-birth abortion access

	(a) No College		(b) College	
	(1)	(2)	(3)	(4)
	Reg FE	Reg FE/IV	Reg FE	Reg FE/IV
Z(MSA CT)	-0.0171***	-0.0292***	-0.0015	-0.0036
	(0.0034)	(0.0059)	(0.0034)	(0.0063)
Not protected access	-0.0041	-0.0163	-0.0119+	-0.0132
	(0.0096)	(0.0128)	(0.0070)	(0.0091)
Restricted access	-0.0071+	-0.0119*	-0.0039	-0.0050
	(0.0042)	(0.0051)	(0.0034)	(0.0042)
Z(MSA CT)* Not protected	0.0005	0.0192	0.0101+	0.0113
	(0.0070)	(0.0183)	(0.0052)	(0.0091)
Z(MSA CT)* Restricted access	0.0063+	0.0121+	-0.0020	-0.0008
	(0.0037)	(0.0065)	(0.0025)	(0.0041)
Obs	374519	374519	314020	314020
R-squared	0.097	0.092	0.137	0.126
Under-ID (K–P LM) Chi-Sq.		34.25		34.11
Weak-ID (K–P rk Wald Chi-Sq)		10.09		14.81
Hansen J p-value		0.041		0.405
Andreson-Rubin F p-value		0.000		0.610
DWH p-value		0.041		0.948

Notes: American Community Survey 2005-2019. Sample includes US-born women selected as described in Section 2. The table shows selected estimated coefficients from a modified version of regression 1 that includes the following two additional variables: two dummies indicating states of birth where abortion access was unprotected and restricted abortion access based on state-level legislation following the 2022 overturn of Roe v. Wade; interactions between the abortion-access dummies and the congestion variable. The model is estimated separately for non-college graduate women (panel a) and for college women (panel b) via OLS (columns 1 and 3) and via 2SLS (columns 2 and 4) using both Duranton and Turner (2011) instruments and interactions between the instruments and the abortion-access dummies. All models include region, year, and region-times-year fixed effects, and are weighted using individual weights. Standard errors are clustered at the MSA level. The table also includes results of all IV diagnostic tests described in table 1 notes. p-Value $<$ 0.1 ($^+$), 0.05 (*), 0.01 (**), 0.001 (***).

Table A13: Congestion and mother's LFP - accounting for center-based infant care cost

	(a) No College			(b) College		
	(1) Reg. FE Raw	(2) Reg. FE Pred.	(3) Reg. FE IV.Pred.	(4) Reg. FE Raw	(5) Reg. FE Pred.	(6) Reg. FE IV.Pred.
Z(MSA CT)	-0.0191***	-0.0135***	-0.0194*	-0.0011	0.0024	-0.0023
	(0.0041)	(0.0038)	(0.0098)	(0.0042)	(0.0039)	(0.0081)
High childcare cost	0.0314*	-0.3266**	-0.2553+	-0.0207*	-0.2743**	-0.2123+
	(0.0124)	(0.0984)	(0.1461)	(0.0098)	(0.0961)	(0.1278)
Obs	372435	372435	372435	311281	311281	311281
R-squared	0.131	0.131	0.123	0.132	0.132	0.126
(K–P LM) Chi-Sq.			27.93			30.83
(K–P rk Wald F)			17.51			30.30
Hansen J p-value			0.675			0.102

Notes: American Community Survey 2008-2019. Sample includes a subsample of women selected as described in Section 2 for years in which childcare-cost information is available in the National Database of Childcare Prices of the DOL Women's Bureau. The table shows selected estimated coefficients from a modified version of regression 1 that includes a dummy variable taking value 1 for MSAs containing at least one county where the cost of full-time center-based infant care is high (in the top-50 percent of the year-specific cross-county distribution of childcare cost). The regression includes the raw dummy variable in columns (1) and (4). Models in columns (2) and (3), (5) and (6), control for the probability that an MSA includes at least one high-childcare-price county predicted using the Cortés and Tessada (2011) instrument. In columns (3) and (6) the model is estimated via 2SLS using both Duranton and Turner (2011) instruments. All regressions are weighted using individual weights, and standard errors are clustered at the MSA level. p-Value $<$ 0.1 (+), 0.05 (*), 0.01 (**), 0.001 (***).

Table A14: Congestion and mother's LFP - accounting for family-based toddler care cost

	(a) No College			(b) College		
	(1) Reg. FE Raw	(2) Reg. FE Pred.	(3) Reg. FE IV.Pred.	(4) Reg. FE Raw	(5) Reg. FE Pred.	(6) Reg. FE IV.Pred.
Z(MSA CT)	-0.0186***	-0.0136***	-0.0195*	-0.0012	0.0024	-0.0023
	(0.0041)	(0.0038)	(0.0098)	(0.0041)	(0.0039)	(0.0081)
High childcare cost	0.0256*	-0.2696**	-0.2106^{+}	-0.0297**	-0.2263**	-0.1751^{+}
	(0.0102)	(0.0812)	(0.1207)	(0.0093)	(0.0793)	(0.1055)
Obs	372380	372380	372380	311257	311257	311257
R-squared	0.131	0.131	0.123	0.132	0.132	0.126
(K–P LM) Chi-Sq.			27.90			30.82
(K–P rk Wald F)			17.50			30.29
Hansen J p-value			0.675			0.102

Notes: American Community Survey 2008-2019. Sample includes a subsample of women selected as described in Section 2 for years in which childcare-cost information is available in the National Database of Childcare Prices of the DOL Women's Bureau. The table shows selected estimated coefficients from a modified version of regression 1 that includes a dummy variable taking value 1 for MSAs containing at least one county where the cost of full-time family-based toddler care is high (in the top-50 percent of the year-specific cross-county distribution of childcare cost). The regression includes the raw dummy variable in columns (1) and (4). Models in columns (2) and (3), (5) and (6), control for the probability that an MSA includes at least one high-childcare-price county predicted using the Cortés and Tessada (2011) instrument. In columns (3) and (6) the model is estimated via 2SLS using both Duranton and Turner (2011) instruments. All regressions are weighted using individual weights, and standard errors are clustered at the MSA level. p-Value $<$ 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

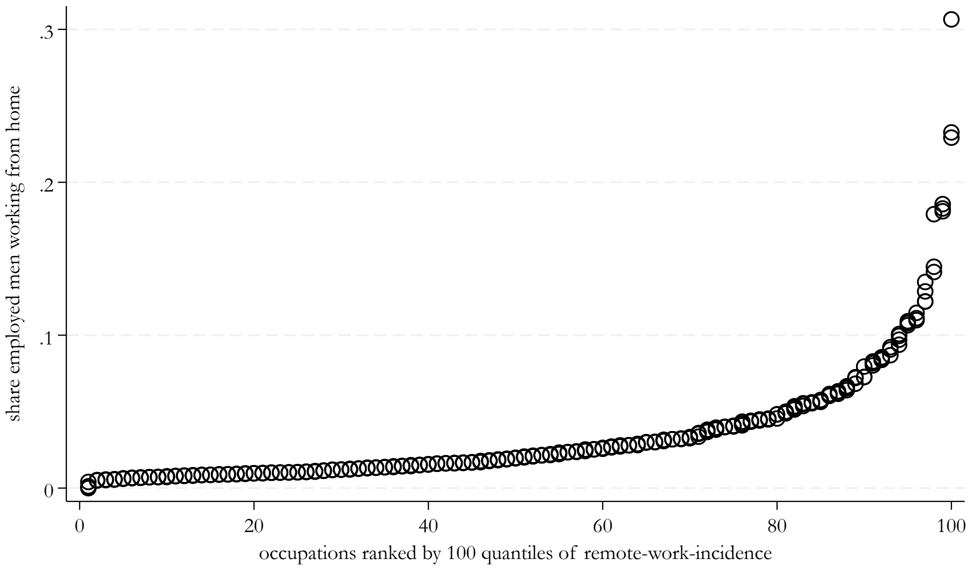
Table A15: Congestion and mother's LFP - accounting for center-based toddler care cost

	(a) No College			(b) College		
	(1) Reg. FE Raw	(2) Reg. FE Pred.	(3) Reg. FE IV.Pred.	(4) Reg. FE Raw	(5) Reg. FE Pred.	(6) Reg. FE IV.Pred.
Z(MSA CT)	-0.0190***	-0.0135***	-0.0194*	-0.0011	0.0024	-0.0023
	(0.0041)	(0.0038)	(0.0098)	(0.0041)	(0.0039)	(0.0081)
High childcare cost	0.0322**	-0.3339**	-0.2610^{+}	-0.0214*	-0.2804**	-0.2171^{+}
	(0.0118)	(0.1006)	(0.1494)	(0.0089)	(0.0983)	(0.1307)
Obs	372435	372435	372435	311281	311281	311281
R-squared	0.131	0.131	0.123	0.132	0.132	0.126
(K–P LM) Chi-Sq.			27.93			30.83
(K–P rk Wald F)			17.51			30.30
Hansen J p-value			0.675			0.102

Notes: American Community Survey 2008-2019. Sample includes a subsample of women selected as described in Section 2 for years in which childcare-cost information is available in the National Database of Childcare Prices of the DOL Women's Bureau. The table shows selected estimated coefficients from a modified version of regression 1 that includes a dummy variable taking value 1 for MSAs containing at least one county where the cost of full-time center-based toddler care is high (in the top-50 percent of the year-specific cross-county distribution of childcare cost). The regression includes the raw dummy variable in columns (1) and (4). Models in columns (2) and (3), (5) and (6), control for the probability that an MSA includes at least one high-childcare-price county predicted using the Cortés and Tessada (2011) instrument. In columns (3) and (6) the model is estimated via 2SLS using both Duranton and Turner (2011) instruments. All regressions are weighted using individual weights, and standard errors are clustered at the MSA level. p-Value $<$ 0.1 ($^{+}$), 0.05 (*), 0.01 (**), 0.001 (***).

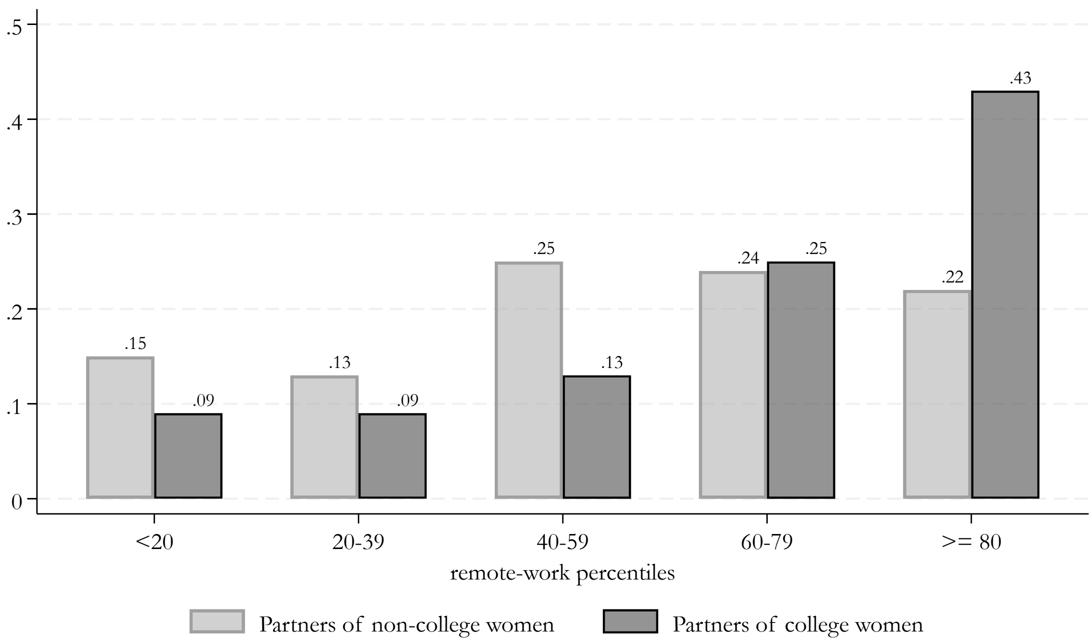
A.2 Figures

Figure A1: Remote-work incidence among employed men by occupation



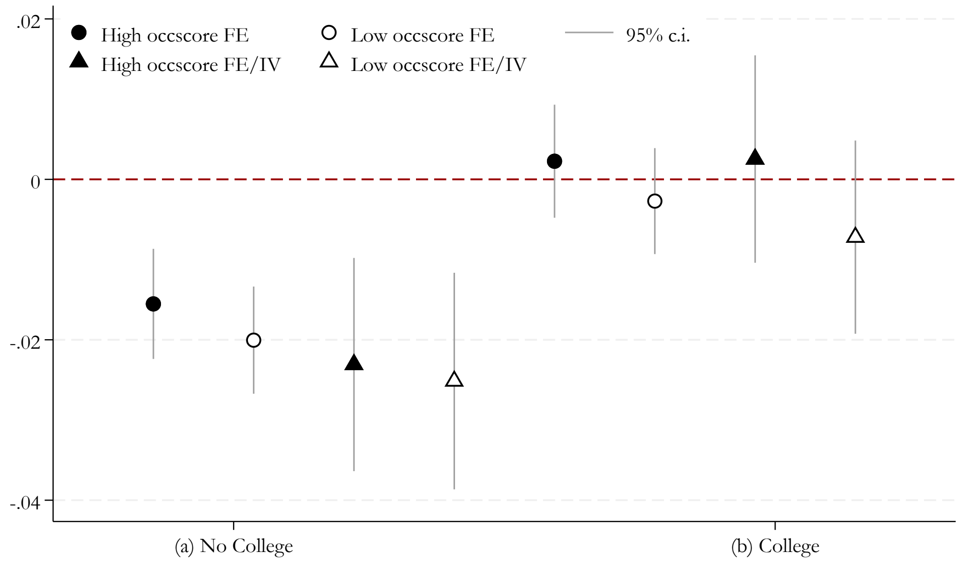
Notes: American Community Survey 2005-2019. The figure plots the incidence of remote work by 3-digit occupation (1990 Census classification). Occupations are ranked by percentiles of their remote-work incidence. Remote-work incidence by occupation is calculated as the share of prime-age men employed in the occupation between 2005 and 2019 who report to not usually commute because they work from home.

Figure A2: Occupational distribution of employed male partners by remote-work quintile



Notes: American Community Survey 2005-2019. The figure represents the distribution of employed male cohabiting partners of 18- to 44-year-old women across 3-digit occupations classified in five quintiles of remote-work incidence. Male partners are split in two sub-groups according to whether their female partner holds a college degree (dark bars) or not (light bars). All statistics reported in the figure are weighted using individual weights. For each 3-digit occupation, the incidence of remote work is calculated as the share of prime-age men working in that occupation between 2005 and 2019 who report to not usually commute as they work from home.

Figure A3: Effects of congestion on mothers' LFP by partner's occupation score



Notes: American Community Survey 2005-2019. The figure represents the β coefficient in regression model 1 estimated via OLS and via IV using Duranton and Turner (2011) instruments, separately for non-college (panel a) and college graduate (panel b) women with children cohabiting with men in the top quintile of the occupation income-score distribution (black symbols) and with men in the bottom four quintiles of the occupation income-score distributon(white symbols). All regression models include all control variables listed in Section 3, and the occupation-specific incidence of long work-hours and of remote-work. All regressions are weighted using individual weights, and standard errors are clustered at the MSA level. The figure includes 95% confidence intervals.

B Neoclassical labor supply model with congestion

B.1 Set-up and implications

The hypothesis that women with stronger labor force attachment (or lower marginal utility cost of working) are likely to self-select in more congested areas is consistent with the implications of a neoclassical labor supply model that incorporates exogenous commuting times. Because the regressions we estimate compare mothers to women without children conditional on individual characteristics (including marital status) and household characteristics (including partners' education and income, if any) we rely on an individual labor supply model that highlights the roles of commuting times and of the idiosyncratic utility cost of working in determining individuals' labor supply decisions. The model builds upon the unitary household labor supply model with commuting time developed by Ji, Oikonomou, Pizzinelli, Shibata, and Tavares (2024).

Consider an individual i who derives utility from their consumption y and who incurs a convex utility cost when working in the labor market or commuting. Consumption depends on i's non-labor income, n, and on their labor income wh^w, where w is the wage rate offered to worker i in their local labor market, and h^w is the number of hours that i works for pay. Let the time constraint for individual i be $h^w + h^h + \tau = 1$, where h^h represents the number of hours spent neither working nor commuting. Let τ indicate the average commute time in the metropolitan area where i lives. Using a quasi-linear utility function to disregard income effects, i's can take the following form

$$u\left(y, h^w\right) = n + wh^w - \phi_i \frac{\left[h^w + \tau\right]^{1+\gamma}}{1+\gamma} \tag{3}$$

The individual chooses h^w to maximize their utility, so that $\frac{\partial u}{\partial h^w} = 0$. This first order condition implies

$$h^{w*} = \left(\frac{w}{\phi_i}\right)^{1/\gamma} - \tau \tag{4}$$

Where $1/\gamma$ is the wage-elasticity of i's labor supply h^{w*}.

Given i's optimal choice of work-hours, i's labor force participation condition can be derived. Individual i participates in the labor market, choosing $h^{w*} > 0$ if

$$w > \phi_i \tau^\gamma \equiv w^r \tag{5}$$

Individual i participates in the labor force if the wage they would receive for every work hour compensates them for the marginal utility cost of their first work hour, which includes the idiosyncratic marginal utility-cost component, ϕ_i, and the utility cost due to commuting, τ. $\phi_i \tau^\gamma$ is worker i's reservation wage, w^r.

The model has two main implications. First, a ceteris paribus increase in commuting times τ increases workers' reservation wages, thus exerting a negative impact on labor force participation of any worker for whom $\phi_i > 0$. Longer commutes diminish the marginal utility of every work-hour, representing a time cost for which workers are not compensated. Second, to the extent that the presence of children increases the relative value of non-market time, mothers (m) should have a higher marginal utility-cost of working compared to women without children (o), $\phi_m > \phi_o$. Hence, an equal increase in congestion τ should have a more negative impact on the labor force participation of mothers than on the labor force participation of women without children.

While the second implication of the model is consistent with the empirical evidence we provide, the first implication appears at odds with the finding that, among women without children, labor force participation is (at least) equal across MSAs characterized by different levels of congestion, conditional on MSA fixed-effects and on several time-varying proxies of MSA-specific labor demand, hence on the average wages w offered to women in different metropolitan areas.

This result can be explained if women with lower idiosyncratic utility cost of working, ϕ, select into more congested areas. Keeping fixed labor demand, that is, the distribution of wages, $F(w)$, in high commuting-time (H) areas and low commuting-time (L) areas, so that $F^H(w) = P^H(W \leq w) = P^L(W \leq w) = F^L(w)$, the equality in labor force participation between women without children in different metropolitan areas, $P^H(w > w^r_H) = P^L(w > w^r_L)$, implies that women without children living in different areas have similar reservation wages. Given ϕ, the reservation wage in high-commuting time areas should be higher to compensate women for longer commutes ($\tau_H > \tau_L$). If reservation wages are equal, however, it must be that $\phi^H < \phi^L$.

It is worth noting that the implication that women with relatively low idiosyncratic utility cost of working and commuting are likely to self-select in high commuting-time metropolitan areas can hold even if, conditional on the MSA characteristics noted above, high commuting-time labor markets offer higher wages compared to low commuting-time areas.

Formal proofs are provided in Section B.2 below.

The implications of the model suggest that selection bias may lead us to underestimate the actual impact of congestion on the labor force participation of mothers ($\delta + \gamma$) and of women without children (γ).

B.2 Proofs of model implications

Proposition 1 Under the assumptions of the model in Section B.1, a ceteris paribus increase in commuting time τ increases workers reservation wages w^r thus causing a decline in labor force participation.

A worker's reservation wage is $w^r = \phi\tau^\gamma$. Hence, $\frac{\partial w^r}{\partial \tau} = \gamma\phi\tau^{\gamma-1} > 0$.

Proposition 2 Under the assumptions of the model in Section B.1, a ceteris paribus increase in commuting time τ has a stronger negative effect on the labor force participation of workers with higher marginal utility cost of working ϕ.

$\frac{\partial w^r}{\partial \tau} = \gamma\phi\tau^{\gamma-1} > 0$. Hence, $\frac{\partial^2 w^r}{\partial \tau \partial \phi} = \gamma\tau^{\gamma-1} > 0$

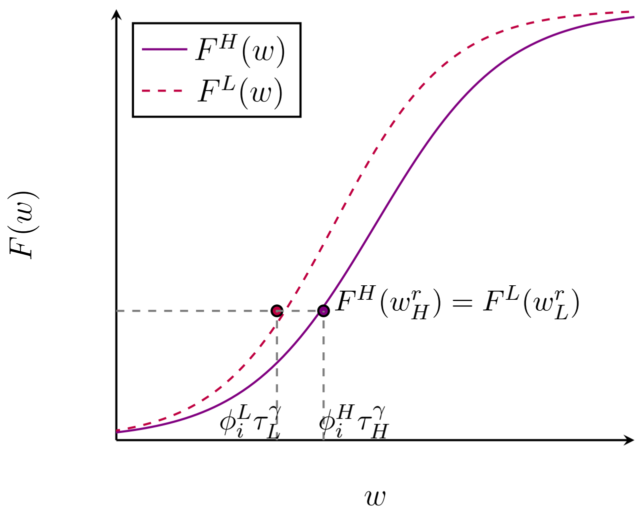
Proposition 3 Under the assumptions of the model in Section B.1, keeping constant the characteristics of labor demand and of workers across metropolitan areas, the labor force participation of workers is equal across areas characterized by different commuting times if workers with low marginal utility cost of working select in metropolitan areas with high commuting times (τ_H).

The share of workers participating in the labor force in a metropolitan area is the share of workers who are offered a wage w higher than their reservation wage, that is $P^J(w > w^r_J)$, where $J = H$ in high-commuting-time areas and $J = L$ in low-commuting-time areas.

$$
\begin{aligned}
P^H(w > w^r_H) &= P^L(w > w^r_L) \\
1 - F^H(w^r_H) &= 1 - F^L(w^r_L) \\
F^H(w^r_H) &= F^L(w^r_L) \\
w^r_H &= w^r_L \\
\phi^H_i &= \left(\frac{\tau_L}{\tau_H}\right)^\gamma \phi^L_i < \phi^L_i
\end{aligned}
\tag{6}
$$

Proposition 4 Under the assumptions of the model in Section B.1, it is possible that $\phi^H_i < \phi^H_i$ even if the distribution of wages in high-commuting-time areas first order stochastically dominates (FOSD) the distribution of wages in low-commuting-time areas as illustrated in figure A4.

Figure A4: Labor force participation, wage distribution and reservation wages in high and low congestion areas assuming heterogeneity in labor demand.



$$
\begin{aligned}
P^H(w > w_H^r) &= P^L(w > w_L^r) \\
1 - F^H(w_H^r) &= 1 - F^L(w_L^r) \\
F^H(w_H^r) &= F^L(w_L^r) \\
F^L(w_H^r) &> F^L(w_L^r) \text{ assuming } F^H \text{ FOSD } F^L \\
w_H^r &> w_L^r \\
\phi_i^H &> \left(\frac{\tau_L}{\tau_H}\right)^\gamma \phi_i^L < \phi_i^L \qquad (7)
\end{aligned}
$$

This inequality does not exclude that $\phi^H < \phi^L$.