Commuting Time and the Gender Gap in Labor Market Participation[*]

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Abstract

In this paper we investigate the contribution of increasing travel times to the persistent gender gap in labor market participation. In doing so, we estimate the effect of commuting times on the labor supply of men and women in the US using microdata from the censuses of the last two decades. To address endogeneity concerns, we adopt an instrumental variables approach that exploits the shape of cities as an exogenous source of variation for travel times. Our estimates indicate that a 10 minute increase in commuting time decreases the probability of married women participating in the labor market by 4.4 percentage points. In contrast, the estimated effect on men is small and statistically insignificant. When exploring potential mechanisms behind the gender asymmetry in our results, we do not find evidence that differences in labor market productivity within couples countribute to the larger penalty of commuting times on women. However, we do find that the negative effect on women increases with the number of children and is larger among those originating from countries with more gendered social norms. Based on this evidence, we conclude that in a context of increasing commuting costs the presence of gender norms that attribute to women the role of main caregivers may prevent gender convergence.

JEL Codes: R41, J01, J16, J22.

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1 Introduction

Gender inequality in labor market outcomes persists in all industrialized countries. Despite important advances, the process of gender convergence seems to have plateaued since the early 2000s. Unequal distribution of the family burden and the persistence of gender norms that reinforce the role of women as the main caregivers are promising candidates to account for the remaining gender gaps in the labor market (Kleven et al., 2019a,b; Bertrand, 2020; Olivetti et al., 2020).

In this paper we propose a complementary explanation for the persistence of gender inequality, namely, the asymmetric effect of commuting costs on the male and female labor supply. High commuting costs will never induce an individual to join the labor force but may discourage participation. For example, in a two-member household, the existence of long commutes may foster specialization by family members in either market or home production to avoid doubling up the cost of going to work (Black et al., 2014). We argue that the gendered distribution of household tasks promoted by the prevalence of social norms that make it more costly for men to stay at home may have contributed to the stagnation of female labor market participation in a context of increasing travel times.

In the US, the cost of going to and from work has increased significantly. In 1980, the average two-way commuting time for a full-time worker was 45 minutes. By 2016, it had increased to 54 minutes (i.e., 20% higher). In the same year, about 20% of commuters spent more than 90 minutes a day traveling to and from work.[1] In Europe, commuting costs are comparable in size and have also increased over time (Gimenez-Nadal and Molina, 2014). Commuting is highly undesirable for workers (Kahneman et al., 2004; Clark et al., 2019), detrimental to their mental and physical health (Roberts et al., 2011; Sandow et al., 2014), and responsible for work absenteeism (Van Ommeren and Gutiérrez-i Puigarnau, 2011).

In this paper, we estimate the effect of commuting time on the decision to participate in the labor market using microdata from the US census and investigate its contribution to the persistence of gender inequality. To guide our empirical analysis, we rely on a parameterized version of the model in Black et al. (2014). In the model, household members specialize in either home or market production in the presence of costly commuting. The model also predicts who will withdraw from the market on the basis of differences in labor and home productivity, and the presence of social norms about the roles of men and women in society. Consistent with the theoretical predictions, we uncover an important degree of heterogeneity in our results. First, we find a large effect of commuting costs on the labor supply of married women, while the effect is small and generally non-significant for men. We also show that the response of women increases monotonically with the number of children, suggesting that family responsibilities are important to explain within-household specialization in the presence of long commutes. **In**

[1] Commuting times are recorded in the US Census since 1980.

contrast, we do not find significant differences in the response of couples with different skill compositions, indicating that gender gaps in labor market productivity are not responsible for the asymmetry in our results. Finally, we focus on a sample of immigrant women in the US. **We document a negative relationship between commuting times and the labor supply of those originating from countries with more traditional gender views. All in all, we conclude that our results support the theoretical prediction that the presence of traditional social norms that reinforce the male breadwinner model, rather than the existence of labor market productivity differentials within couples, can account for the larger penalty of commuting times on women**.

In our empirical analysis, we exploit the variation in commuting times across MSA (Metropolitan Statistical Areas) in the US using the IPUMS data. To identify the causal effect of commuting on individual labor supply we follow Harari (2020) in using the shape of cities as an exogenous source of variation. We focus on city compactness measured by how closely a city's shape resembles a circle. Compactness is determined by the presence of geographical accidents such as steep mountains and water bodies. It has been shown that more compact cities enjoy shorter commuting times and we exploit this feature to identify the effect of travel times on labor supply decisions (Angel et al., 2010). The main drawback of our identification strategy is that a city feature correlated with its shape might have a direct effect on individuals' labor supply. To take this into account, we first normalize our measure of city compactness so that it is uncorrelated to size and density. We also follow the suggestions in Altonji et al. (2005) to validate our identification strategy. First, we show that city shape is uncorrelated with the observable characteristics of individuals **that are associated with labor force participation**. Second, we verify that our IV estimates are unaffected by the inclusion of individual and MSA controls. Finally, we show that our estimates conform to expectations when the model is estimated on subgroups of the population that should be affected differently by commuting times (e.g., single women and married women without children).

A few recent studies have already identified a role for commuting costs in explaining gender differences in labor market outcomes. Using evidence for the UK, Petrongolo and Ronchi (2020) show that men have higher wage returns from voluntary job changes, while women have higher returns in terms of proximity to workplaces. This is consistent with the view that women attribute a higher value to short commutes. For France, Le Barbanchon et al. (2020), using a job search model where commuting matters, estimate that gender differences in the willingness to commute explain about 10% of the gender gap in re-employment wages. For Sweden, Bütikofer et al. (2019) show that women benefit less from transport infrastructures that give access to distant labor markets, negatively affecting the gender wage gap. Moreno-Maldonado (2019) using a quantitative spatial model of households shows that the labor force participation of women with children is lower in big cities due to longer commutes. The paper that is closest to ours is by Black et al. (2014) who document that US cities with longer commutes have lower participation rates among married women.

Our contribution to the existing literature is twofold. We first provide a causal estimate of the effect of commuting time on individuals' labor supply based on an innovative source of exogenous variation that relies on the geographical accidents that shape cities. Our IV estimates indicate that the effect of commuting is larger than that suggested by OLS. **Second, while previous papers have documented the presence of a negative correlation between commuting costs and married females labor supply, our study goes a step further by providing suggestive evidence that this relationship results from the pervasive effect of traditional social norms. We document that the negative effect of commuting costs on females increases with family responsibilities and is stronger among immigrant women originating from countries with more gendered social norms. However, we do not find evidence that differences in labor market productivity within couples can explain the asymmetric response of men and women to commuting times**.

Our results complement alternative explanations that have been proposed to understand the stagnation of gender inequality in the labor market during the last decades. Traditional arguments based on differences in human capital accumulation and the presence of discrimination do not suffice to account for the persistence of gender inequality (Altonji and Blank, 1999). More recently, researchers have identified more promising mechanisms for the lack of gender convergence. For instance, some have argued that very generous family-friendly policies may be detrimental to women's labor market advances (Blau and Kahn, 2013; Cipollone et al., 2014). In addition, other public policies, such as the prevalence of joint taxation, may have a disincentive effect on female labor supply (Bick and Fuchs-Schündeln, 2017). Gender differences in psychological traits such as attitudes towards risk, competition, negotiation, and altruism may contribute to the gender gap in particular labor market contexts (Gneezy et al., 2003; Niederle and Vesterlund, 2007; Croson and Gneezy, 2009). Finally, the persistence of social norms that perpetuate the role of women as main caregivers have been proved to be responsible for a substantial fraction of the remaining gender gaps, particularly among married individuals (Fernández et al., 2004; Fortin, 2005, 2015; Alesina et al., 2013a; Farré and Vella, 2013; Bertrand et al., 2015; Cools and Patacchini, 2019). According to our results, increasing commuting costs may discourage female labor market participation in the presence of traditional social norms that make more costly for men to assume the bulk of family responsibilities.

The remainder of the paper is organized as follows. Section 2 presents a theoretical framework for the labor force participation decisions of household members and guides the empirical analysis. Section 3 describes the data, samples and main variables in the analysis. Section 4 lays out the main empirical specification and the instrumental variables strategy that we use. Results are presented and discussed in Section 5 and some concluding remarks are presented in Section 6.

2 Theoretical framework

To study the relationship between commuting costs and the labor supply of individual household members, we parametrize the model in Black et al. (2014). The main model prediction is that in the presence of high commuting costs it is optimal that one household member withdraws from the labor force. Deciding which member exits depends on: i) differences in productivity in the market and at home and ii) the presence of gendered social norms about the roles of men and women in society. We present the model and discuss the main results below. The model solution is presented in Appendix A.

A household consists of two spouses ($j = m$ for men or f for women). The utility of each spouse is given by $u_j(c_j, l_j) = \alpha_j \ln c_j + (1 - \alpha_j) \ln l_j$, where c_j denotes consumption and l_j denotes time spent at home, which we interpret as domestic work. The parameter α_j reflects the individual preference for consumption over time spent at home. The inequality $\alpha_m > \alpha_f$ is consistent with women being more productive in domestic work or with the presence of a social norm that makes domestic work more acceptable for women. Individuals face a time constraint ($\bar{l} = h_j + l_j + k$), where h_j denotes time spent in market work and k is a fixed commuting cost that can only be avoided by not participating in the labor market. There might be intra-household transfers y, implying that the budget constraints for men and women are $c_m = w_m h_m - y$ and $c_f = w_f h_f + y$. We assume that households maximize the sum of the individual utilities $u_m + u_f$.

Since commuting costs are unaffected by the number of hours worked, it is necessary to solve the model in two steps. First, we obtain the optimal amount of consumption, the time at home (and working hours) and the transfer made when both spouses work ($h_m > 0$ and $h_f > 0$) and when only one does ($h_m > 0$ and $h_f = 0$, or viceversa). Second, we compare the utility levels in each situation to determine if there is specialization (one spouse stays at home) and, when necessary, which spouse will exit the labor market.

The top panel of Figure 1 plots household utility, $u_m(c_m, l_m) + u_f(c_f, l_f)$ as a function of commuting costs when both spouses are identical in terms of wages and preferences. The solid line represents the level of utility when both spouses work while the dotted line represents the utility when only one does. Utility decreases with commuting costs in both cases, but the slope is more steeply negative when both household members work as commuting costs are paid twice. Hence, for some parameter configurations, an increase in commuting costs (k) might induce some households to specialize.

To determine who stays at home when commuting costs are high, we first focus on the role of different productivities in the labor market (e.g. the presence of a gender gap in wages) under the assumption of symmetrical preferences ($\alpha_m = \alpha_f$). An illustration of this case is provided in the second panel of Figure 1. Since both spouses are equal in terms of preferences, it yields higher consumption levels and utility if the spouse with the higher wage works.

Let us now analyze the case where wages are equal but spouses preferences are different or there are gendered social norms. For example, $\alpha_m > \alpha_f$ may reflect a situation where women are more productive in domestic work or the presence of a social norm that supports the male breadwinner model. One example is depicted in the bottom panel of Figure 1. Here, women value time at home relatively more and, as a result, $u_m + u_f$ is higher when they stay home.

From this stylized model we can derive several predictions that will be empirically tested. First, the presence of high commuting costs favors within-household specialization in either market or home production. Second, the model predicts that the presence of a gender wage gap will induce women to specialize in home production. Finally, traditional gender norms will lead women to withdraw from the market when commuting costs increase.

3 Data, sample and variables

In the empirical analysis we employ data from the decennial US censuses and the American Community Surveys (King et al., 2010). The baseline analysis is conducted on the 5% census sample of 2000, which is the last census to record commuting times. To investigate the robustness of our results, we also employ the 5% census metro sample from 1980 and 1990 and the 1% annual samples in the American Community Surveys for the 2007-2011 period. We restrict the analysis to prime-age individuals (25-55 years old), with special emphasis on married couples as the model predictions are specific to two-member households. We only consider individuals living in cities. We use the definition of a city from the census of 2000 (i.e., Metropolitan Statistical Areas, MSA). There are 272 MSAs that comprise about 80% of the US population.[2]

Table 1 provides descriptive statistics for the main variables in the study. Panel A focuses on measures at the city/MSA level. The first row shows the summary statistics for commuting times. Following Black et al. (2014), we compute the city-average two-way commute time using the information reported in the census about door-to-door travel time in minutes, and restrict the sample to **(non-Hispanic)** white male workers as this group has the highest employment rate. In 2000, the average commuting time was about 51 minutes, with a standard deviation of 8, a maximum of 84 and a minimum of 34 minutes. Table 1 also shows summary statistics for population size, median household income, share of employment in manufacturing and share of public employment. These variables are computed from the US county and City Data Book

[2]The average MSA population was around 755 thousand inhabitants in 2000. The smallest MSA is Kokomo (IN) with about 102 thousand inhabitants, the median city is Montgomery (AL) with 333 thousand and the largest MSA is Los Angeles with more than 9.5 millions.

(CCDB) of 2000 and employed as controls in the empirical analysis.

Panel B displays the summary statistics for the controls at the individual level, obtained from the IPUMS in 2000, separately for **married** men and women. The first row displays labor force participation, followed by indicators of the intensive margin of the labor supply: number of weekly hours worked, part-time employment, working long hours (i.e. more than 50 hours per week) and the probability of working in an occupation with a high proportion of part-time employment (i.e. an occupation in the top 10th or 25th percentile of the distribution of part-time employment across all occupations). According to the figures in the table, the gender gap in participation in 2000 was 19 percentage points (89% for men and 70% for women). Part-time employment was much more prevalent among women (17% versus 2%). Men worked more hours on average than women (46.3 versus 39.55) and had a higher probability of working long hours (36% versus 14%). The rest of the rows in the panel show the descriptive statistics for the individual controls included in the estimation: age, spouse income, presence of children, having a college degree and race.

Finally, we employ the World Value Survey (WVS) to measure gender role attitudes in the countries of origin of US immigrants. To increase the number of countries in the sample, we pool results from the WVSs conducted between 2000 and 2011. Following Alesina et al. (2013b) we focus on two statements about the role of men and women in society: "When jobs are scarce, men should have more right to a job than women" and "Men make better political leaders than women do". We compute the percentage of individuals in each country who answered *agree* or *strongly agree* with the statements. Figure A1 indicates a clear positive correlation across responses to the two statements. It also shows a substantial degree of heterogeneity across countries. Accordingly, the level of agreement with the statement "When jobs are scarce, men should have more right to a job than women" varies from 4% in Canada to almost 96% in Egypt. When asked whether "Men make better political leaders than women do", the share varies from less than 10% in Sweden to about 88% in Egypt. The large variation in gender role attitudes across countries should allow us to explore the interaction between culture and commuting cost impacts among US immigrants.

4 Empirical strategy

To study the effect of commuting costs on the labor supply decisions of men and women we estimate the following model:

$$Pr(LaborForce_{ic}) = \beta \ commuting_c + X_i'\lambda + X_c'\gamma + \varepsilon_{ic} \tag{1}$$

where $LaborForce_{ic}$ is an indicator variable that takes value 1 if individual i living in MSA c participates in the labor market and 0 otherwise. The explanatory variable $commuting_c$ is the average two-way commuting time in minutes for working white men in city c. To ease interpretation we divide commuting time by 100. Accordingly, β is the percentage point increase in the probability of participating in the labor market resulting from a one minute increase in travel time.

X_i includes a comprehensive set of individual characteristics: spouse income (in logs), an indicator for the presence of children, age, race and educational attainment dummies.[3] X_c represent control variables at the MSA level such as the percent of employment in the manufacturing sector, the percent of public employment, median household income (in logs), population -and its square- (in logs) and regional dummies. The descriptive statistics for the control variables are displayed in Panels A and B of Table 1.

Despite this rich set of control variables, the OLS estimates in equation 1 might be biased for at least two reasons. First, cities experiencing positive economic shocks might have higher rates of labor force participation which directly impact on congestion and travel times. This may generate a reverse causality bias pushing the OLS estimates towards zero. Second, sorting of individuals across cities also represents a challenge to the OLS estimates. Costa and Kahn (2000) show that high-power couples tend to sort into large cities to better deal with the co-location work problem. To the extent that commuting times are longer in larger cities, the OLS estimates will also be biased towards zero in the presence of sorting.

In order to address endogeneity concerns, we adopt an instrumental variable strategy that exploits the shape of cities as an exogenous source of variation for travel times. City shape is determined by geographical constraints such as water bodies or steep slopes (Saiz, 2010; Harari, 2020). At the same time, the shape of cities is an important determinant of intra-urban commuting costs. A city with a more compact geometry (i.e. a shape closer to that of a circle) will be characterized by shorter within-city trips and more cost-effective transport networks (Bertaud, 2004; Cervero, 2001). We employ the measures of city compactness proposed by Angel et al. (2010) and also employed in Harari (2020) as an instrument for commuting time. We use the US 2000 Urban Area GIS Files from the National Historical Geographic Information System Database (NHGIS) (Steven Manson and Ruggles, 2018) to compute the city shape variables. We overlay a 100x100 meters grid to the shape of the city and compute the following two measures of city compactness:

- $Proximity_c = \sum_i^N \frac{d_{i,CBD}}{N}$, where $d_{i,CBD}$ is the distance between the centroid of each grid cell, i, and the Central Business District (CBD).[4] Accordingly, the *proximity* index measures the average distance to the CBD.[5] Note that this measure emphasizes commuting trips to

[3] **We employ five mutually exclusive race categories: non-Hispanic white, non-Hispanic black, withe-Hispanic, black-Hispanic, Chinese, Japaneses and "others". Education categories are primary or less, high school graduate, college graduate and more than college.**

[4] We use the Central Business District Geocodes dataset from Holian and Kahn (2015).

[5] A few MSAs have more than one principal city. In these cases, we calculate the distance between each centroid and each principal city and then average them according to the principal city population.

the city center.

- $Cohesion_c = \sum_i^N \sum_j^N \frac{d_{ij}}{N(N-1)}, \forall i \neq j$, where $d_{i,j}$ is the distance between grid cell centroids, i and j. The *cohesion* index is the average distance between all pairs of points in the city.[6] Note that this measure implicitly assumes that jobs and residents are homogeneously distributed throughout the city.

These two indices are correlated with city surface as bigger cities will present longer distances between the interior points. To isolate the effect of city shape, we normalize the two measures following the procedure described in Angel et al. (2010). In a first step, we compute the Equivalent Area Circle (EAC) of city c. That is, a circle whose area coincides with that of the city. The rationale for using a circle is that it is the most compact geographical shape (i.e. the distances between the interior points are minimized). In a second step, we compute the proximity and cohesion indices that the EAC of each city would exhibit.[7] Finally, we compute the normalized proximity and cohesion indices as:

- $nProximity_c = \frac{Proximity_{EAC,c}}{Proximity_c}$
- $nCohesion_c = \frac{Cohesion_{EAC,c}}{Cohesion_c}$

We employ the normalized version of the proximity and cohesion index in our empirical analysis. Note that by construction the normalized proximity and cohesion indices are uncorrelated with the size and the density of the city. A normalized index value close to 1 means that the city index is close to the optimal EAC index (i.e., a circular city). Lower values indicate that the shape of the city is less circular. Figure A2 illustrates the actual shape and the EAC of Chicago (Panel A) and Minneapolis (Panel B). Chicago, a city with long commuting times (i.e. average two-way commute in 2000 was 68 minutes), has a non-circular shape as the lake causes a mismatch between the EAC and the current city shape. As a result, the normalized proximity and cohesion indices are low (i.e. 0.635 and 0.843, respectively). In contrast, Minneapolis, a city with short commutes (i.e. the average two-way commute in 2000 was 54 minutes), has a rather circular shape that closely overlaps that of the EAC. In this case, the values of the normalized proximity and cohesion indices are higher (i.e. 0.930 and 0.915, respectively). Panel D in Table 1 displays the descriptive statistics for the two indices.

The proximity and cohesion indices have different underlying assumptions regarding the nature of commuting within cities. While the proximity index considers commuting trips to the city center, the cohesion index assumes that jobs (and homes) are homogeneously distributed within the city.

[6] In MSAs with more than one principal city, we compute the popultion weighted average *cohesion* index.

[7] Specifically, $proximity_{EAC,c} = (2/3) \times r_{EAC,c}$ while $cohesion_{EAC_c} = 0.9054 \times r_{EAC,c}$, where $r_{EAC,c}$ is the radius of the Equivalent Area Circle.

Kahn (2010) has documented that commuting times in the US are a monotonic function of the distance to the CBD in medium-big and small MSAs. However, in metropolitan areas with more than 4 million inhabitants, he finds a tipping point at 7 miles to the CBD. When people live more than 7 miles away from the CBD, commuting time tends to decrease, suggesting that commutes for people living in distant suburbs tend to be more local. Accordingly, in our sample we expect the proximity index to predict commuting time better for small and medium-sized MSAs than for the largest MSAs.

Figure A3 plots the values of the normalized proximity and cohesion indices against commute time. As expected, the figure displays a negative correlation between commuting times and the degree of compactness of a city as measured by the normalized indices.

In our framework, the possibility that dual earner couples choose to sort into more compact cities, with shorter commuting times, poses a threat to identification. To verify the validity of our exclusion restriction, we follow Altonji et al. (2005) and test for sorting on observable characteristics. We first examine men and women who recently moved into the city (i.e. within the past 5 years). **We test if individual characteristics that are important determinants of labor supply are correlated with the shape of the city as captured by the two indices.[8] Panel A in Table 2 shows the results for this sample of movers on the probability of having a college degree, the probability of being a power couple, the number of children, and the probability of being married. The results indicate that none of the used indices is correlated with the observed individual characteristics, supporting the validity of our proposed identification strategy. Since our empirical exercise is conducted on the sample of mover and non-mover residents in an MSA, Panel B conducts the same exercise for the full sample. Finally, Panel C reports the results for the sample of immigrants living in the US that will be used in a final exercise to provide additional evidence on the importance of gender norms in driving our results. In either case, we do not find evidence of a significant correlation between observable individual characteristics and our proposed instrumental variables.**

5 Results

5.1 OLS Estimates

We now turn to our main empirical exercise and estimate the effect of commuting costs on the labor supply of men and women. Panel A in Table 3 shows the OLS estimates of the empirical model in equation 1. Columns 1 to 4 focus on married women and columns 5 to 8 on married men.

[8] Note that the normalized indices that we use are orthogonal to city size. This should alleviate concerns about the fact that power couples sort into large cities (Costa and Kahn, 2000).

The specification in columns 1 and 5 includes only as controls the MSA population and its square (in logs). Columns 2 and 6 add regional dummies, and columns 3 and 7 the individual controls: age, education and race dummies, and spouse income (in logs). Finally, columns 4 and 8 include the MSA controls: the percent of employment in the manufacturing sector, the percent of workers employed in the public sector and median household income (in logs).

For women, the point estimates displayed in columns 1 to 4 suggest a negative relationship between commuting time and labor market participation. The effect is statistically significant at conventional levels in columns 3 and 4 when controlling for individual and MSA characteristics. According to the estimates in our preferred specification in column 4, a 10-minute increase in travel time decreases the probability of married women participating in the labor market by 2 percentage points. The estimates for men are much smaller in magnitude and only statistically significant in the last specification. As discussed in Section 4, reverse causality and sorting are likely to bias the OLS estimate further towards zero. Accordingly, these estimates should be interpreted as a lower bound of the effect of commuting costs on labor force participation.

5.2 IV estimates

To deal with endogeneity concerns we instrument commuting times using the normalized proximity and cohesion indices (see Section 4). Table 4 displays the estimates of the first-stage. The results indicate that both the normalized proximity and cohesion indices are strong predictors of commuting times. In our preferred specification that includes controls at the MSA level (columns 3 and 4), an increase of one standard deviation in the normalized proximity index (0.16) decreases two-way commuting by 16 minutes. For the cohesion index (S.D. of 0.15) the effect is 14 minutes. The F-test of excluded instruments indicates that the proximity index is a stronger instrument than the cohesion index. For the proximity index the F-test is beyond 10, which is the rule-of-thumb standard widely accepted by practitioners (Angrist and Pischke, 2008). For the cohesion index, the value of the F-test is just below 10. This result is consistent with our previous discussion regarding the suitability of the proximity index to predict commuting times in small and medium sized cities, which constitute the majority of cities in our sample. It also suggests that trips to the city center, better captured by the proximity index, are still important in the US despite the important decentralization employment between 1960 and 2000 (Baum-Snow, 2010).

The IV estimates of equation 1 are presented in Panels B and C in Table 3. Panel B reports the estimates based on the normalized proximity index and Panel C those of the cohesion index. For women, the estimates in all columns are larger (in absolute value) than the corresponding OLS estimates. Also, the point estimates remain stable when regional dummies (column 2), individual characteristics (column 3) and MSA controls (column 4) are sequentially included. This stability in the coefficients alleviates concerns about the validity of our identification strategy (Altonji et al., 2005). According to our preferred specification in column 4, a 10-minute increase in commuting time decreases the probability of a married woman participating in the labor market by 4.4 percentage points when the proximity index is employed in estimation (by 3.8 when the cohesion index is used). In 2000, the participation rate of prime-age married women was 70%, implying that the estimated effect represents a 6% decrease relative to the mean.

In columns 5 to 8 we estimate the same models for married men. The point estimates are much smaller and statistically insignificant in most specifications, suggesting that the effect of commuting mostly impacts on women. Despite being insignificant, the magnitude and sign of the estimated coefficient suggests that longer commutes may also negatively affect the participation decisions of men. This result is consistent with a strand of the literature showing that better access to jobs within cities improves labor market performance (Aslund et al., 2010; Gobillon et al., 2011; Andersson et al., 2018).

Panel D in Table 3 reports the estimates of the direct effect of city shape metrics on the labor force participation of married women and men (i.e. the reduced-form estimates). Conforming to expectations, more circular cities, with higher values of the proximity and cohesion indices, are associated with higher rates of female labor force participation. For men, the coefficients on the indices are also positive but statistically insignificant when control variables are included in the estimation.

The evidence presented so far reveals an important gender asymmetry in the effect of commuting costs on individuals' labor supply. Namely, long commutes negatively affect the labor supply decisions of women, while the effect, if any, is much smaller on men. According to the IV estimates in Panel B in Table 3, if commuting times in the US had remained at the 1980 level (45 instead of 54 minutes), the labor force participation of married women would have been 4 percentage points higher, which represents 30% of the gender gap in participation for married individuals in 2000 (i.e. 19 percentage points).

While we stress the role of work related trips as the main mechanism behind our findings, other factors related to the cost of within-city trips could influence labor supply decisions. For example, longer trips for childrearing (e.g. taking children to school) or associated to domestic chores (e.g. grocery shopping) could also negatively affect female labor force participation.[9] **It could also be that commuting is more costly for women than for men, inducing the former group to stay at home. This could occur, for instance, if women are more dependent on public transportation. While we cannot rule out the presence of these alternative mechanisms, they all have in common that mobility costs within the city differently affect the labor market participation of**

[9] Using the American Time Use Survey at the State level for 2005, we find that the length of trips associated with rearing children is uncorrelated with commuting times, suggesting that this channel does not explain the bulk of our results.

women.

5.3 Effects at the intensive margin

Now we turn the analysis to the intensive margin of the labor supply. **According to theory (see equation A4 in the Appendix), conditional on working, women should work fewer hours if commuting costs increase as they reduce the time available to work in the market and at home. To test this prediction, we examine the** number of weekly hours worked, the decision to work part-time and working long hours. We define part-time work as less than 35 hours in a typical week and long hours as working more than 50 hours per week. We also investigate the effect of commuting costs on the probability of working in a typical part-time occupation as defined in Section 3.

We are aware that the results at the intensive margin cannot be interpreted as causal, as the decision to participate in the labor market is clearly affected by commuting times. For example, it may be that only the most talented and motivated women decide to work in high commuting locations. This would bias our estimates at the intensive margin towards zero. Nevertheless, we find these results informative about the effects of commuting on the labor market beyond the decision to participate.

Table 5 presents the results at the intensive margin. Columns 1 and 2 report the OLS estimates and 3 to 6 the IV estimates.[10] The sample is restricted to married women aged 25 to 55, **working a minimum of 35 weeks** during the reference year. Columns 1, 3 and 5 show the results for our preferred specification in terms of control variables. Columns 2, 4 and 6 add occupational fixed effects.[11]

The IV estimates in column 3 to 6 indicate the presence of statistically significant effects. There is evidence that higher commuting costs reduce the number of hours worked: a 10-minute increase in commuting time decreases hours worked per week by between 0.62 and 0.82. The probability of working part-time also increases by 2.4 percentage points in response to a 10 minutes increase in commute. In contrast, there is no effect on the probability of working long hours.[12] Note that the results on part-time employment and the number of hours worked do not decrease when occupational fixed effects are included in the estimation (columns 4 and 6). This suggests that most of the effect occurs within occupations rather than by sorting across occupations. This last result is also consistent with the absence of any effect of commuting on the probability of working in an occupation with a high concentration of part-time employment.[13]

[10] Columns 3 and 4 employ the proximity index as instrument and column 5 and 6 the cohesion one.

[11] Occupational groups at the 3-digit level as in the 1990 Census.

[12] The negative effect of commuting on the intensive margin are consistent with the theoretical predictions of the model outlined in section 6.

[13] The results at the intensive margin for men are reported in Table A1. The only effect that is statistically

5.4 Robustness checks

In this section we conduct a number of empirical exercises to validate the robustness of our previous findings, **focusing on results obtained using the normalized proximity index as instrument. The results using the cohesion index are qualitatively and quantitatively similar.** An important concern for the validity of our results is the presence of sorting across cities. So far we have shown the absence of sorting on the basis of observable characteristics. In Table 2 observable individual characteristics such as education and family responsibilities appeared uncorrelated with the two instruments employed in the estimation. To further explore the possibility that sorting could be affecting our results, we estimate the model in equation 1 on a sample of individuals that are less mobile. Unfortunately, the IPUMS microdata do not provide the county or city of birth, but they do report the state of birth and recent migration histories. With this information we can estimate the model on the sample of individuals who were born in the state where they currently live and have not changed residence in the last 5 years. The results for this sample are reported in Table A2. These estimates are very similar in magnitude and significance to those in Table 3 and provide additional evidence that sorting across MSAs does not seem to be the main driver of our results.

Another concern is the presence of unobservable MSA characteristics that affect labor supply decisions and commuting time simultaneously, such as climate or other city features that may attract a particular type of worker. To address this concern, we estimate a model at the MSA level where the dependent variable is the gender ratio in labor force participation. This specification eliminates all unobservable city characteristics that homogeneously affect male and female labor supply. The results of this alternative specification are presented in Table A3. The point estimates are positive, statistically significant, reinforcing the view that female labor supply is much more responsive to changes in the duration of commutes.

Next we estimate the model in equation 1 in long-term differences between 1980 and 2000. This specification allows us to control for unobserved MSA characteristics which are time invariant and may have heterogenous effects across genders.

$$\Delta FLS_c = \beta_1 \Delta commuting_c + \Delta X_c' \lambda + \eta_c$$

where ΔFLS_c is the percentage point change in female labor supply in city c between 1980 and 2000. $\Delta Commuting_c$ is the increase in commuting time over the same period and $\Delta X_c'$ captures differences in the control variables at the MSA level. Changes in the shapes of US cities between

significant at conventional levels is for the probability of working part-time. The results indicate that longer commutes also increase part-time work among married men although the effect is one fourth of that estimated for women.

1980 and 2000 are limited and thus, we adopt as instrument $\Delta commuting_c$ with the cross-sectional normalized proximity **in 2000**.The results are presented in the bottom panel in Table A4 **while the top panel presents the OLS estimates for completeness**. According to the 2SLS estimates, a 10 minute increase in commuting time reduces the labor supply of married women by 7 to 9 percentage points. **These estimates and those of the gender gap in participation reported in Table A3 are larger in magnitude than the baseline estimates of Table 3. One possible explanation is the use of individual versus aggregated data. While in Table 3 the unit of analysis is the individual in Tables A3 and A4 the unit of analysis is the city, which implicitely assigns a higher weight to individuals living in smaller cities.**

Finally, we estimate the effect of commuting on labor supply at different points in time. Specifically, we employ the 1990 and 1980 censuses and the 2006-2011 ACS samples described in Section 3. The estimation results are reported in Table A5. Columns 1 employs data for the period 2006-2011, column 2 employs data for the period 2000 and replicates our main findings in Table 3, columns 3 uses data for 1990 and column 4 for 1980. The point estimates in column 1 and 3 are very similar in magnitude to the one obtained using the 2000 census (column 2). In contrast, the effect for 1980 in column 4 is much smaller and statistically insignificant, a result that also appears in Black et al. (2014). This suggests that commuting was a less relevant factor in explaining labor force participation back then.

In sum, the previous results suggest that bias from sorting and omitted variables does not seem to be driving our results. It also indicates that the role of commuting as a determinant of individual decisions to participate in the labor market has increased over time.

5.5 Mechanisms

The previous results uncover an important gender asymmetry in individual responses to commuting times. The theoretical framework in Section 2 suggests two possible mechanisms to account for the commuting time penalty on women. First, in a two-member household, differences in home productivity or the presence of a gender wage gap in the labor market may induce women to withdraw from the labor force to avoid paying the cost of going to work twice. Second, the presence of a gendered social norm that makes socially more costly for men assuming the bulk of family responsibilities, may lead women to stay home and take care of the family and other housework when commuting time increases.

To investigate the contribution of these two mechanisms in explaining the different responses to commuting costs across gender we conduct heterogeneity analyses. We first estimate our main model (equation 1) on a sample of individuals with different levels of family responsibilities using as proxy the number of children in the household. The estimates are reported in Table 6. Panel A presents the results for women while Panel B focuses on men. Column 1 displays the results on single women. The point estimate is much smaller than our baseline estimate in Table 3 and statistically insignificant. Among married women the effect is smallest for those without children: a 10-minute increase in commuting time reduces participation by 3.0 percentage points (column 2). In the presence of children the effect becomes larger: a 10-minute increase in commuting time decreases the probability of participation by 4.0 percentage points for those with 1 child (column 3). This effect increases to more than 5 percentage points among those with 2 children (column 4) or more (column 5). The estimates in column 6 show that the effect is magnified in the presence of young children. Accordingly, a 10-minute increase in commuting time decreases the probability of working outside the home for mothers with children younger than 5 by 6.6 percentage points. This effect is smaller for women who have children with 5 or more years of age (column 7). For men, Panel B displays point estimates that are negative but much smaller in magnitude and statistically insignificant for most groups. These findings suggest that the higher burden of childcare carried by women could be partly responsible for their stronger response to commuting times.

Next, we explore whether gender differences in labor market productivity are driving our findings. We estimate the model in equation 1 on a sample of couples with different levels of education to proxy for differences in productivity. Column 1 of Table 7 display the results for couples where both members have a college degree (i.e. power couples). Column 2 is restricted to couples without college education (i.e. low power couples). Column 3 focuses on couples where only the husband has a college degree (i.e. part power couples, men) and, finally, column 4 on those where only the wife has college education (i.e. part power couples, women). The point estimates are very similar across the different samples, suggesting that differences in productivity, or the presence of a gender wage gap, are not a major driver of the asymmetric response by men and women to commuting costs.

To provide additional evidence that gendered social norms may be responsible for the larger commuting time penalty on women, we conduct an epidemiological analysis that mirrors that in Fernández and Fogli (2009). Accordingly, we estimate the labor supply response to commuting times on a sample of foreign-born individuals living in the US.[14] This exercise is motivated by the idea that cultural attitudes regarding the role of women in society are transmitted across generations and individuals retain these values when moving to a new country. Accordingly, by comparing immigrants from different countries living in the US we can assess the role of culture on individual decisions. We are aware that this approach has limitations as comparisons across immigrants from different countries do not only isolate the role of culture but it may also capture individual unobserved characteristics that can be correlated with gender

[14] Differently from Fernández and Fogli (2009) we can not restrict our analysis to second generation immigrants as information on parental origin is only available in the 1970 census and there is no information about commuting times in that census.

norms, such as human capital differences (e.g. work experience or driving skills), or other economic or social features of country of origin (e.g. the level of development). Thus, we will be cautious in interpreting the results of our epidemiological approach.

As explained in Section 3, we measure gender role attitudes in the country of origin using the World Value Survey. Following Alesina et al. (2013b), we compute the percentage of individuals in each country who "strongly agree" or "agree" with the following two statements: "When jobs are scarce, men have more right to a job than women do" and "Men make better political leaders than women do". To formally examine the role of social norms, we estimate the baseline model including, as additional regressors, the measure of gender attitudes in the country of origin (i.e. the percentage of agreement with a traditional statement) and its interaction term with commuting time. Column 1 in Table 8 displays the results when traditional views are represented by the statement "When jobs are scarce, men should have more right to a job than women", while column 2 shows the corresponding results for the statement "Men make better political leaders than women do". To ease interpretation, commuting time and gender attitudes have been demeaned. Thus, the coefficient on commuting times represents its impact at average values of gender attitudes. Note that the interaction term also needs to be instrumented. Accordingly, we instrument commuting times and its interaction term with the proxy for gender roles with the normalized proximity index and its interaction term with the corresponding gender roles proxy.[15] The F-statistics indicate that the two instruments are relevant predictors of the two endogenous variables.

Our coefficient of interest is the interaction term between the measure of attitudes and commuting times. The estimate on this interaction is negative in both columns, suggesting that the effect of commuting times on the labor force participation of married women is magnified by the presence of traditional gender norms. Figure A4 shows the implied marginal effects for observed values in gender attitudes. **To understand the quantitative implications of the results we provide a comparison between Germany and Iran, two countries with marked differences in gender norms as displayed Table A1. Focusing on the results of column 1 (i.e. "When jobs are scarce"), the estimates imply that the effect of interest is close to zero (and not statistically significant) for married German women living in the US, while it is around -0.32 for Iranian women (i.e. a 10 minute increase in commuting times reduces labor supply by 3.2 percentage points). While this evidence supports the theoretical prediction that the larger penalty of commuting times on women responds to the presence of gendered social norms, the presence of other unaccounted features correlated with the country of origin's culture could also mediate gender differences in labor market outcomes.**

The evidence in this section provides some light on the potential mechanisms driv-

[15] Sanderson and Windmeijer (2016) first-stage statistics for models with more than one endogenous variables are provided at the bottom of Table 8.

ing the asymmetric effect of commuting costs across genders. First, we document that the commuting time penalty on married women increases with family responsibilities measured by the number and the age of children. We also report a larger response to commuting times among women originating from countries with more traditional gender norms. In contrast, we do not find evidence that the magnitude of the effect varies with the educational mix of the couple. All in all, we find this evidence supports the theoretical prediction that the larger family burden supported by women and the presence of traditional social norms, and not productivity differentials within couples, are responsible for the different response of men and women to commuting costs.

6 Conclusions

The persistence of traditional gender norms have been identifying as the main driver of the remaining gender differentials in industrialized countries (Kleven et al., 2019a; Olivetti et al., 2020). In this paper we complement this result by providing evidence that, in a context of increasing commuting costs, the presence of pervasive social norms that attribute to women the role of main caregivers in the family may prevent gender convergence to happen any soon.

Using microdata from the US census of the last two decades, we uncover an important gender asymmetry in the labor supply effect of commuting times. **Namely, married women, in particular, those with more and younger children substantially decrease their labor force participation in response to increasing commuting times. In contrast, the labor supply of men and that of married and childless women is almost unresponsive. We also document that immigrant females originating from countries with more traditional gender norms are more responsible to changes in commuting costs. Finally, we do not observe important differences in the magnitude of the effect of commuting costs across couples with different levels of education. We interpret these findings as supportive evidence that the presence of social norms that make it more costly for men to stay home, rather than gaps in labor market productivity within couples, may account for the commuting time penalty on women**.

Our findings are relevant for policy makers, as they identify an important cause of the stagnation in gender convergence observed during recent decades. The increase in congestion and excessive agglomeration in cities may be partly counterbalancing the increase in female labor supply observed since the late 1950s. Accordingly, reducing commuting costs by investing, for instance, in transport infrastructure and urban planning can facilitate female participation in the labor market and promote gender equality in the future.

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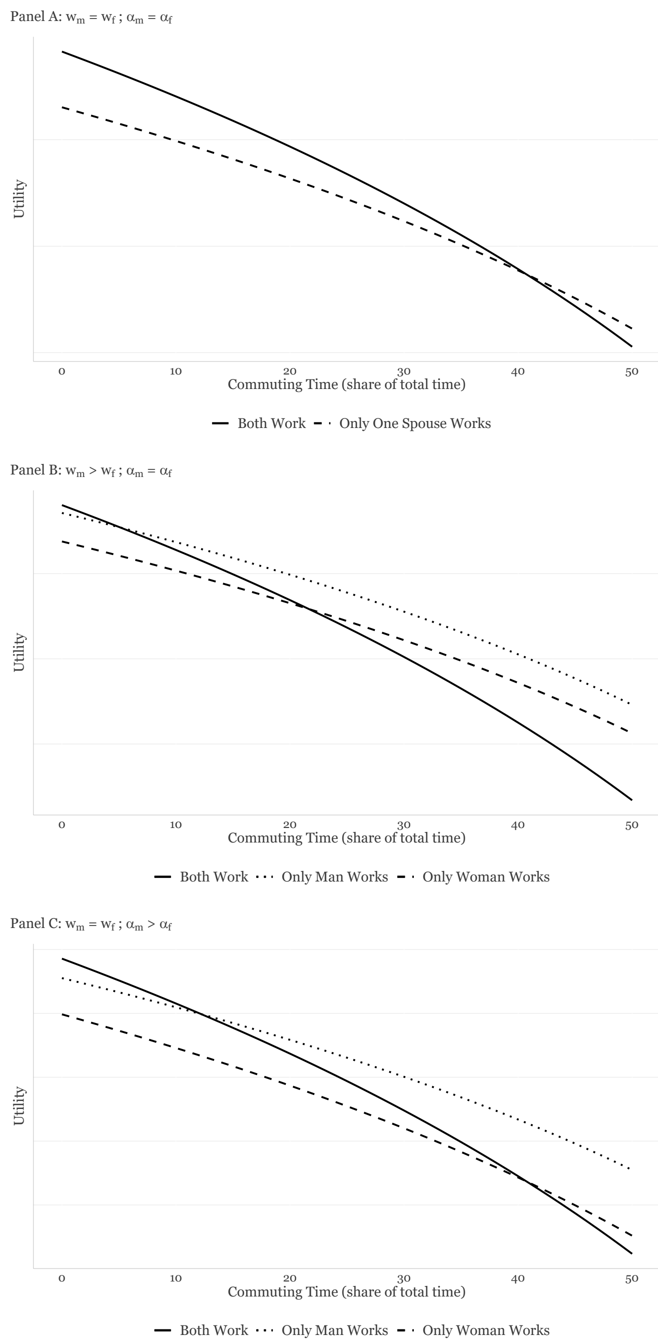
Figures

Figure 1: Commuting Costs and Household Labor Supply



Notes: This figure displays 3 different scenarios of home and market specialization. Utility is the sum of individual utilities in the household. Panel A illustrates a symmetric situation where there is no gender gap in wages and gender roles are egalitarian. Panel B illustrates a situation where there is a wage gap that favors men and gender roles are egalitarian. In Panel C there is no gender wage gap but gender norms favor the male breadwinner culture.

Tables

Table 1: Summary Statistics

Panel A: MSA variables, CCDB 2000	# cities	Mean	Sd	Min	Max
Commuting	272	0.51	0.08	0.34	0.84
Population	272	755,150	1,251,024	101,541	9,519,338
Median household income	272	16,674	2,719	10,650	23,958
% Manufacturing employment	272	0.13	0.08	0.01	0.59
% Public employment	272	0.16	0.06	0.07	0.49
Panel B: Individual variables, IPUMS 2000	# indiv.				
Married Women					
Labor force	1,382,904	0.7	0.46	0	1
Weekly hours worked	610,069	39.55	9.98	1	99
Part-time employment	610,069	0.17	0.37	0	1
Working long hours	610,069	0.14	0.34	0	1
Part-time occupation (25th Pctile)	610,069	0.28	0.45	0	1
Part-time occupation (10th Pctile)	610,069	0.13	0.34	0	1
Age	1,382,904	40.29	8.27	25	55
Spousal Income	1,382,904	46,537	54,378	0	354,000
Children	1,382,904	0.71	0.45	0	1
College	1,382,904	0.3	0.46	0	1
Married Men					
Labor force	1,349,163	0.89	0.31	0	1
Weekly hours worked	912,261	46.3	9.84	1	99
Part-time employment	912,261	0.02	0.15	0	1
Working long hours	912,261	0.36	0.48	0	1
Part-time occupation (25th Pctile)	912,261	0.24	0.43	0	1
Part-time occupation (10th Pctile)	912,261	0.11	0.31	0	1
Age	1,349,163	40.81	8.19	25	55
Spousal Income	1,349,163	19,492	27,165	0	354,000
Children	1,349,163	0.7	0.46	0	1
College	1,349,163	0.33	0.47	0	1
Panel C: City Shape measures, NHGIS 2000	# cities				
nProximity	272	0.67	0.16	0.11	0.95
nCohesion	272	0.71	0.15	0.21	1

Notes: Commuting is the average MSA two-way commuting time in minutes for non-Hispanic white male workers divided by 100. Hours worked are average weekly hours worked in 1999. Part-time employment is defined as working less than 35 hours per week. Working long hours is defined as working more than 50 hours per week. Sources: Sources: Data in Panel A are obtained from the US County and City Data Book (CCDB), in Panel B from the 2000 census and in Panel C from the National Historical Geographic Information System (NHGIS) in 2000.

Table 2: City Shape and Individual Characteristics of the Sample

	(1) College Education	(2) Power Couple	(3) Nº Children	(4) Married
	Married Women	All Married Individuals	Married Women	All Women
Panel A: Movers				
nProximity	-0.017	-0.005	0.018	0.025
	(0.085)	(0.076)	(0.059)	(0.035)
nCohesion	0.064	0.068	-0.031	0.019
	(0.093)	(0.085)	(0.104)	(0.045)
Observations	236,434	484,519	236,434	399,632
Panel B: Full Sample				
nProximity	-0.027	-0.013	-0.079	0.027
	(0.057)	(0.052)	(0.086)	(0.030)
nCohesion	0.048	0.049	-0.07	-0.009
	(0.066)	(0.058)	(0.106)	(0.032)
Observations	1,382,904	2,732,067	1,382,904	2,195,894
Panel C: Immigrants				
nProximity	-0.044	-0.02	0.211	0.078
	(0.057)	(0.051)	(0.165)	(0.053)
nCohesion	-0.018	0.003	0.009	0.039
	(0.092)	(0.077)	(0.241)	(0.059)
Observations	285,275	573,335	285,275	412,773

Notes: Panel A is restricted to individuals aged 25 to 55 who changed MSA of residence in the previous 5 years. Panel B is restricted to individuals aged 25 to 55 who did not change MSA of residence in the previous 5 years. Panel C is restricted to individuals aged 25 to 55 living in the US but born abroad. Each number in the table corresponds to a different OLS coefficient from a regression of a observable individual characteristic on the instrument (nProximity or nCohesion). No additional controls are added in estimation. In column (1) the dependent variable is an indicator for college education; in column (2) an indicator for being a power couple (i.e. where both members have college education); in column (3) the number of children; and in column (4) an indicator for being married. Standard errors clustered at the MSA level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Source: 5% 2000 US Census IPUMS.

Table 3: Effect of Commuting Time on Labor Force Participation

Panel A - OLS	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Married Women				Married Men			
Commuting	-0.082	-0.063	-0.121***	-0.200***	-0.021	0.01	-0.005	-0.046***
	(0.093)	(0.052)	(0.028)	(0.024)	(0.072)	(0.040)	(0.020)	(0.015)
Panel B - Instrument: nProximity								
	Married Women				Married Men			
Commuting	-0.301	-0.394**	-0.356***	-0.442***	-0.131	-0.133	-0.035	-0.024
	(0.242)	(0.191)	(0.116)	(0.102)	(0.163)	(0.120)	(0.060)	(0.072)
F-Stat (Excl. Instr.)	19.740	12.836	13.153	16.645	19.395	12.530	12.741	15.843
Panel C - Instrument: nCohesion								
	Married Women				Married Men			
Commuting	-0.491**	-0.444*	-0.342**	-0.384***	-0.321**	-0.226	-0.104	-0.116
	(0.207)	(0.235)	(0.137)	(0.111)	(0.134)	(0.147)	(0.078)	(0.097)
F-Stat (Excl. Instr.)	13.558	7.452	7.577	9.550	13.269	7.340	7.418	9.265
Panel D- Reduced Form								
	Married Women				Married Men			
nProximity	0.050	0.052***	0.048***	0.046***	0.022	0.018	0.005	0.002
	(0.034)	(0.017)	(0.010)	(0.009)	(0.025)	(0.014)	(0.008)	(0.007)
nCohesion	0.083***	0.056***	0.044***	0.041***	0.054***	0.029*	0.013	0.012
	(0.029)	(0.022)	(0.014)	(0.011)	(0.020)	(0.017)	(0.010)	(0.010)
Observations	1,382,904	1,382,904	1,382,904	1,382,904	1,349,163	1,349,163	1,349,163	1,349,163
Region Dummies	NO	YES	YES	YES	NO	YES	YES	YES
Individual Controls	NO	NO	YES	YES	NO	NO	YES	YES
MSA Controls	NO	NO	NO	YES	NO	NO	NO	YES

Notes: The dependent variable is a binary indicator that takes value 1 if the individual is in the labor force and 0 otherwise. Commuting is the average MSA two-way commuting time in minutes for non-Hispanic white male workers divided by 100. The sample is restricted to married women and men between 25 and 55 years old. Panel A shows OLS estimates. Panel B shows 2SLS estimates using the normalized proximity index as instrument. Panel C shows 2SLS estimates using the normalized cohesion index as instrument. Panel D shows reduced form estimates. All regressions include population and its squared term in logs. Individual controls are age, education level and race dummies and the log of spouse income. MSA controls include: % employment in manufacturing, % employment in the public sector and MSA median household income in logs. F-stat is the value of the statistic for the test of excluded instruments. Standard errors clustered at the MSA level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 4: Effect of City Shape on Commuting Time – First Stage Estimates

	(1)	(2)	(3)	(4)
nProximity	-0.166***		-0.103***	
	(0.037)		(0.025)	
nCohesion		-0.168***		-0.109***
		(0.046)		(0.037)
Observations	1,382,904	1,382,904	1,382,904	1,382,904
F-stat (Excl. Inst.)	19.740	13.558	16.645	9.550
Region Dummies	NO	NO	YES	YES
Individual Controls	NO	NO	YES	YES
MSA Controls	NO	NO	YES	YES

Notes: The dependent variable is the average MSA two-way commuting time in minutes for non-Hispanic white male workers divided by 100. The explanatory variables are the normalized proximity index (first row) and the normalized cohesion index (second row). The sample is restricted to married women between 25 and 55 years old. All regressions include population and its squared term in logs. Columns 3 and 4 include region dummies, individual and MSA controls as defined in Table 3. F-stat is the value of the statistics for the test of excluded instruments. Standard errors clustered at the MSA level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 5: Effect of Commuting Time on the Intensive Margin, Married Women

Outcome	(1)	(2)	(3)	(4)	(5)	(6)
	OLS Estimates		IV Estimates			
Weekly hours worked	-1.000	-1.018	-6.232**	-7.250**	-8.242*	-8.914*
	(0.736)	(0.708)	(2.926)	(3.084)	(4.788)	(4.842)
Part-time	0.039	0.040	0.241**	0.259**	0.235	0.247*
	(0.027)	(0.025)	(0.106)	(0.106)	(0.146)	(0.145)
Long hours	-0.009	-0.005	-0.02	-0.058	-0.074	-0.107
	(0.012)	(0.012)	(0.051)	(0.050)	(0.079)	(0.080)
Part-time occupation (10%)	-0.024	-	-0.039	-	-0.035	-
	(0.022)	-	(0.073)	-	(0.084)	-
Part-time occupation (25%)	-0.024	-	-0.039	-	-0.035	-
	(0.022)	-	(0.073)	-	(0.084)	-
Observations	852,914	852,914	852,914	852,914	852,914	852,914
F-stat (Excl. Instr)			18.759	18.805	10.742	10.740
Instrument			nProximity	nProximity	nCohesion	nCohesion
Occupation Fixed Effects	NO	YES	NO	YES	NO	YES

Notes: Hours worked refers to the regular number of hours worked per week. Part-time is defined as working less than 35 hours per week. Long hours is defined as working more than 50 hours per week. Part-time occupation (10%) and (25%) are indicator variables that equal 1 if the occupation is in the top 10th or 25th percentile of the distribution of part-time employment across all occupations. Commuting is the average MSA two-way commuting time in minutes for non-Hispanic white male workers divided by 100. The sample is restricted to married women between 25 and 55 years old who worked more than 35 weeks of the last year. All regressions include population and its squared term in logs. Individual and MSA controls as defined in Table 3. Occupation fixed effects controls for 1990 census occupational group fixed effects. Columns 3 and 4 use the nProximity index as instrument, while columns 5 and 6 use the nCohesion index as instrument. F-stat is the value of the statistic for the test of excluded instruments. Standard errors clustered at the MSA level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 6: Effect of Commuting Time on Labor Force Participation in the Presence of Family Responsibilities – 2SLS (Proximity Index)

	(1) Single	(2) w/o Children	(3) 1 Child	(4) 2 Children	(5) $\geq$ 3 Children	(6) Children < 5yo	(7) Children > 5yo
Panel A: Married Women							
Commuting	-0.110	-0.299***	-0.401***	-0.569***	-0.548***	-0.659***	-0.359***
	(0.113)	(0.094)	(0.105)	(0.128)	(0.174)	(0.168)	(0.097)
Observations	458,578	403,069	335,857	402,550	241,428	321,455	658,380
F-Stat (Excl. Instr)	19.289	17.797	18.321	16.696	12.243	15.085	16.451
Panel B: Married Men							
Commuting	-0.046	-0.020	-0.123**	-0.024	0.060	-0.018	-0.042
	(0.124)	(0.160)	(0.052)	(0.059)	(0.071)	(0.066)	(0.050)
Observations	648,343	399,091	319,676	394,655	235,741	340,311	609,761
F-Stat (Excl. Instr)	18.107	15.909	17.764	16.411	12.085	14.826	16.120

Notes: The dependent variable is a binary indicator that takes value 1 if the individual is in the labor force and 0 otherwise. Commuting is the average MSA two-way commuting time in minutes for non-Hispanic white male workers divided by 100. All regressions include population and its squared term in logs. Individual and MSA controls as defined in Table 3. Instrument used is normalized proximity. F-stat is the value of the statistic for the test of excluded instruments. Standard errors clustered at the MSA level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 7: Effect of Commuting Time on Labor Force Participation by Educational Characteristics of Couples – 2SLS (Proximity Index)

	(1) Power Couple	(2) Low Power	(3) Part Power (Man)	(4) Part Power (Woman)
Panel A: Married Women				
Commuting	-0.354***	-0.438***	-0.670***	-0.443***
	(0.122)	(0.108)	(0.206)	(0.115)
Observations	293,063	760,698	201,318	127,825
F-stat (Excl. Instr)	20.555	14.103	15.631	19.439
Panel B: Married Men				
Commuting	-0.116	0.026	-0.124***	0.105
	(0.080)	(0.072)	(0.046)	(0.282)
Observations	282,803	729,392	156,879	180,089
F-stat (Excl. Instr)	20.797	13.426	16.404	14.551

Notes: The dependent variable is a binary indicator that takes value 1 if the individual is in the labor force and 0 otherwise. Commuting is the average MSA two-way commuting time in minutes for non-Hispanic white male workers divided by 100. As in Costa and Kahn (2000), *Power couples* are defined as couples in which both spouses have at least college education (+4 years of college education). *Low power couples* are defined as couples in which both spouses have less than college education. *Part power couples* are those in which one spouse has at least college education while the other has not. All regressions include population and its squared term in logs. Individual and MSA controls as defined in Table 3. Instrument used is normalized proximity. F-stat is the value of the statistic for the test of excluded instruments. Standard errors clustered at the MSA level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 8: Effect of Commuting Time on the Labor Force Participation of Immigrant Married Women – 2SLS

	(1)	(2)
LFP married women interacted with	*When jobs are scarce...*	*Men better leaders...*
Commuting (minutes)	-0.034	-0.040
	(0.058)	(0.094)
Commuting × Gender attittudes	-0.796***	-0.972***
	(0.045)	(0.324)
Gender attitudes	0.019	0.036
	(0.078)	(0.077)
Sanderson-Windmeijer F Statistic (Commuting)	26.10	94.30
Sanderson-Windmeijer F Statistic (Interaction term)	4.16	4.76
Observations	472,702	472,702

Notes: The dependent variable is a binary indicator that takes value 1 if the individual is in the labor force and 0 otherwise. Commuting is the average MSA two-way commuting time in minutes for non-Hispanic white male workers divided by 100. Gender attitudes are the share of agreement with each WVS statement by country of origin. Each column uses a different question from the WVS (given in column headers): "When jobs are scarce, men should have more right to a job than women" (1) and "Men make better political leaders than women do" (2). Sample (5% 2000 US census IPUMS and 2006-2011 pooled ACS) is restricted to married women aged between 25 and 55 years who were born outside the US. All regressions include log population and its squared term, year fixed effects, and individual controls, MSA controls and region dummies as in Table 3. Commuting time and gender attitudes have been demeaned to ease interpretation. Instruments used are the normalized proximity and the interaction between the share of agreement with a WVS statement and the normalized proximity index. Sanderson-Windmeijer F-statistic is the value of the statistic for the test of excluded instruments when there are two or more endogenous variables. Two-way clustered standard errors (MSA and country of birth) are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Online Appendix A

In this appendix we detail the solution to the model outlined in Section 2. When both spouses work, the optimal amount of consumption, time at home and the transfer are determined by:

$$y^{bothwork} = \Big(\frac{w_m - w_f}{2}\Big)(\bar{l} - k) \tag{A1}$$

$$c_j^{bothwork} = \alpha_j\Big(\frac{w_m + w_f}{2}\Big)(\bar{l} - k) \tag{A2}$$

$$l_j^{bothwork} = (1 - \alpha_j)\Big(\frac{w_m + w_f}{2w_j}\Big)(\bar{l} - k) \tag{A3}$$

$$h_j^{bothwork} = \Big(1 - (1 - \alpha_j)\frac{w_m + w_f}{2w_j}\Big)(\bar{l} - k) \tag{A4}$$

The optimal transfer is proportional to the wage gap, so that low wage individuals receive a positive transfer and viceversa. The consumption of individual j depends on own preferences (α_j) and the average household wage. Time spent at home depends on own preferences to stay at home, $(1 - \alpha_j)$, and it increases when the individual wage decreases relative to that of the other spouse. Higher commuting costs, k, reduce the time available to work in the market and at home. Accordingly, these costs can be interpreted as a negative income effect[16].

Let us now analyze the optimal outcomes when only one spouse participates in the labor market. Here, we analyze the case in which the woman stays home and, thus, does not incur in commuting costs. Interchanging the gender subscripts gives the solution in which the man stays at home. When the woman stays home, $l_f^{oneworks} = 1$, consumption equals the transfer received from the spouse that participates ($c_f^{oneworks} = y^{oneworks}$). The optimal consumption of both spouses and the leisure and hours worked of the working spouse are determined by:

$$c_m^{oneworks} = \frac{\alpha_m}{1 + \alpha_f} w_m(\bar{l} - k) \tag{A5}$$

$$c_f^{oneworks} = \frac{\alpha_f}{1 + \alpha_f} w_m(\bar{l} - k) \tag{A6}$$

$$l_m^{oneworks} = \frac{1 - \alpha_m}{1 + \alpha_f}(\bar{l} - k) \tag{A7}$$

$$h_m^{oneworks} = \Big(1 - \frac{1 - \alpha_m}{1 + \alpha_f}\Big)(\bar{l} - k) \tag{A8}$$

Commuting costs also decrease the consumption (and working hours) and the amount of time

[16] **We restrict our attention to parametrizations of the model where w_m, w_f and $\bar{l}$ are high enough to ensure levels of consumption and leisure are above unity, which guarantees that these terms remain positive after being logged.**

spent at home by the spouse who works. However, when only one spouse works, the level of consumption and time spent at home depends on the preferences of the two members of the household.

In order to analyze how commuting affects household utility, we differentiate utility with respect to k when both spouses work and when only one does:

$$\frac{d\left(u_m(c_m^{bothwork}, l_m^{bothwork}) + u_f(c_f^{bothwork}, l_f^{bothwork})\right)}{dk} = \frac{-2}{\bar{l} - k} \tag{A9}$$

$$\frac{d\left(u_m(c_m^{oneworks}, l_m^{oneworks}) + u_f(c_f^{oneworks}, l_f^{oneworks})\right)}{dk} = \frac{-1 - \alpha_f}{\bar{l} - k} \tag{A10}$$

The two expressions are negative, indicating that commuting decreases utility. Note that $\alpha_f < 1$ implies that the utility of the household when both spouses work is more sensitive to changes in commuting costs as these are paid twice. This implies that for certain parameter configurations, an increase in k will induce some households to specialize.

We now analyze who will stay home when commuting costs are high and it is not optimal for both spouses to work. We first analyze the role of wages by assuming that preferences are homogeneous ($\alpha_m = \alpha_f$). Equations A5 and A6 indicate that the consumption of the two household members will be higher if the individual who works is the one with the higher wage. The time spent at home is $l_j = 1$ for the spouse who stays and A7 for the working member, implying that the time spent at home does not depend on the wage of the household member that works. As a result, household utility is higher when the lower waged worker stays home.

We now turn to the case in which wages are equal but preferences are not. As explained above, $\alpha_m > \alpha_f$ might reflect women being intrinsically more productive at domestic work or social norms that make staying at home less desirable for men. The consumption for both members is higher when only the man works compared to the case where only the woman works. Hence, both individuals enjoy more consumption when the higher α_j person works.

As for leisure components, the utility that the household obtains if the woman stays at home is:

$$(1 - \alpha_f)\ln \bar{l} + (1 - \alpha_m)\ln\left(\frac{(1 - \alpha_m)}{(1 + \alpha_f)}\left(\bar{l} - k\right)\right) \tag{A11}$$

Conversely, the leisure components of utility if the man stays at home are:

$$(1 - \alpha_m)\ln \bar{l} + (1 - \alpha_f)\ln\left(\frac{(1 - \alpha_f)}{(1 + \alpha_m)}\left(\bar{l} - k\right)\right) \tag{A12}$$

The first utility component is higher when the woman stays home as the value of leisure is higher for the woman and the time endowment is common for both. However, the second component of utility is higher when the man stays home. If the woman works, she works less hours implying that her time devoted to leisure is higher and the value she attaches to leisure is also higher. However, we can show that the first component of utility, i.e. the value of leisure for the spouse staying at home, dominates.

To be more specific, it can be shown that starting from a symmetric equilibria ($\alpha_m = \alpha_f$), the utility difference between the two scenarios is increasing with α_m. To see that, we differentiate the lesisure components of the utility differential with respect to α_m:

$$(1-\alpha_f)\ln\bar{l} + (1-\alpha_m)\ln\left(\frac{(1-\alpha_m)}{(1+\alpha_f)}\left(\bar{l}-k\right)\right) - (1-\alpha_m)\ln\bar{l} - (1-\alpha_f)\ln\left(\frac{(1-\alpha_f)}{(1+\alpha_m)}\left(\bar{l}-k\right)\right) \tag{A13}$$

Which yields:

$$\ln\left(\frac{\bar{l}}{\frac{(1-\alpha_m)}{(1+\alpha_f)}\left(\bar{l}-k\right)}\right) - \frac{\alpha_m+\alpha_f}{1+\alpha_m} = \ln\left(\frac{\bar{l}}{\left(\bar{l}-k\right)}\right) + \ln\left(\frac{(1+\alpha_f)}{(1-\alpha_m)}\right) - \frac{\alpha_m+\alpha_f}{1+\alpha_m} \tag{A14}$$

At the symmetric equilibrium, this expression simplifies to:

$$\ln\left(\frac{\bar{l}}{\left(\bar{l}-k\right)}\right) + \ln\left(\frac{(1+\alpha)}{(1-\alpha)}\right) - \frac{2\alpha}{1+\alpha} > 0 \tag{A15}$$

It turns out this expression is always positive. To see that, we start by focusing on the case in which α is very small. Note that the first two component of equation A15 are positive while the third is negative. When α is very small, the second term in equation A15 becomes:

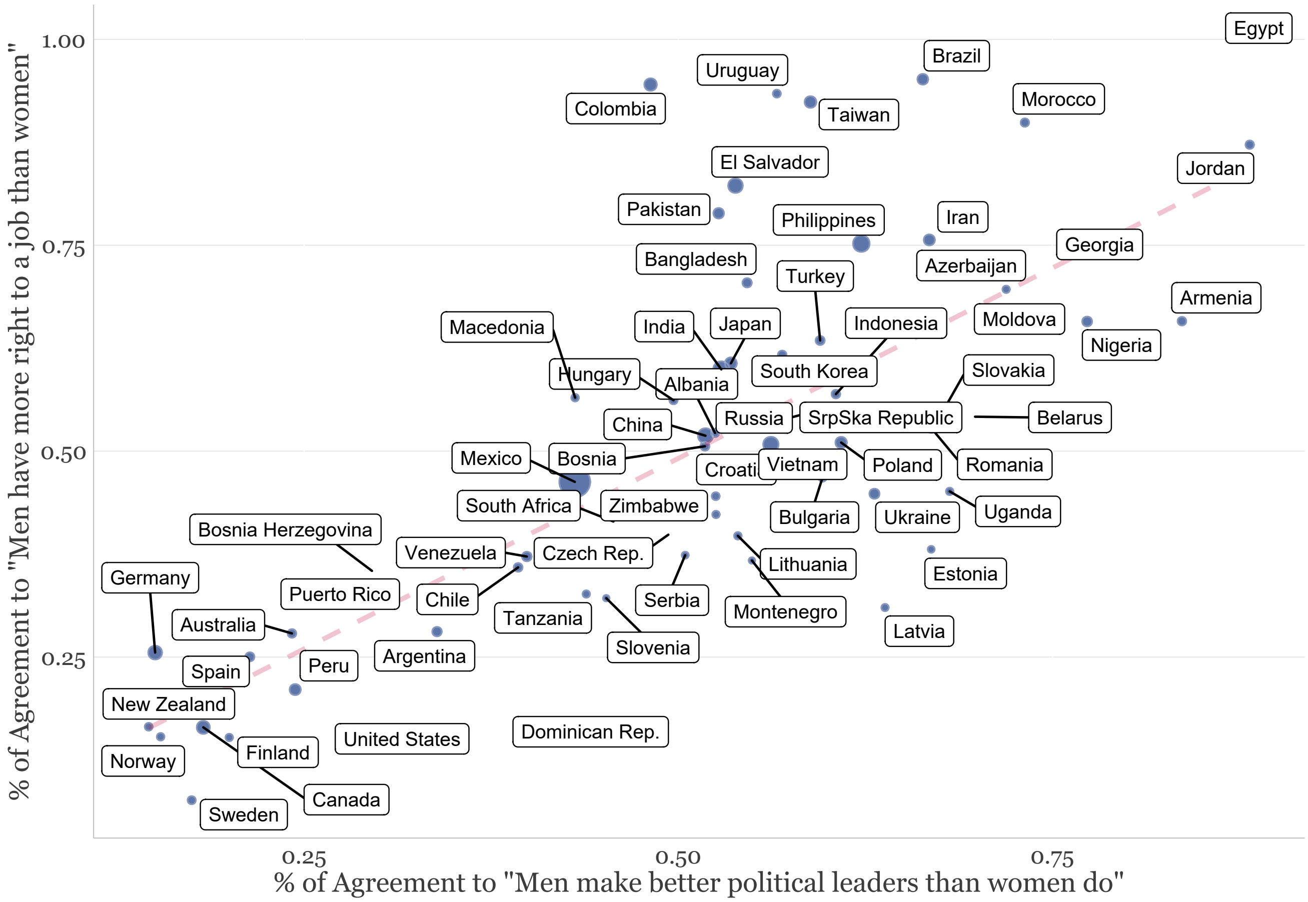
$$\ln\left(\frac{(1+\alpha)}{(1-\alpha)}\right) = \ln\left(1+\frac{2\alpha}{1-\alpha}\right) \approx \frac{2\alpha}{1-\alpha} > \frac{2\alpha}{1+\alpha} \tag{A16}$$

Which dominates the third term and therefore equation A15 will be positive. This result also holds for higher values of α. As α increases, $\ln\left(\frac{1+\alpha}{1-\alpha}\right)$ increases at a faster rate than $\frac{2\alpha}{1+\alpha}$. More specifically, while the first expression grows at the rate $\frac{2}{(1+\alpha)(1-\alpha)}$, the second expression grows at the rate $\frac{2}{(1+\alpha)(1+\alpha)}$. Therefore, for all symmetric equilibria,

an increase in α_m implies that the sum of utilities derived from leisure is higher in the case in which the woman stays home.

Online Appendix B: Figures

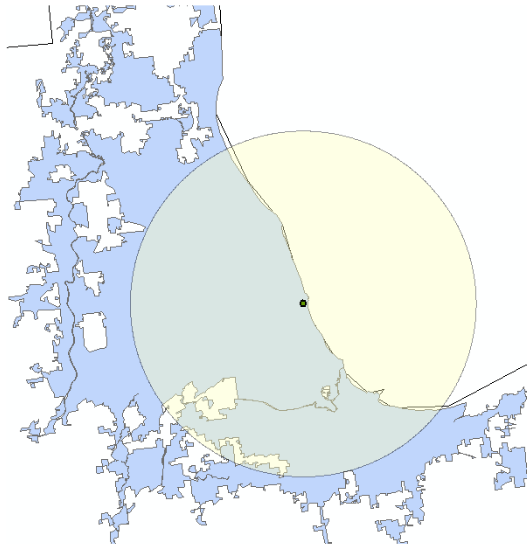
Figure A1: Gender Role Attitudes in the World Value Survey



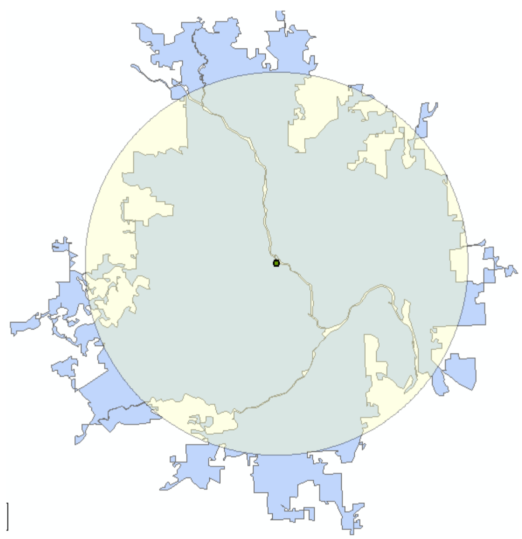
Note: The graph displays the percentage of respondents in each country that *agree* or *strongly agree* with the statement: "When jobs are scarce, men should have more right to a job than women" (y-axis) and "Men make better political leaders than women do"(x-axis). Each point represents a country. The size of the dot illustrates the magnitude of each immigrant group in the US population in 2000. Source: World Value Survey. Several years between 2000 and 2011.

Figure A2: City Shape and Equivalent Area Circle (EAC)

Panel A: Chicago Urban Area

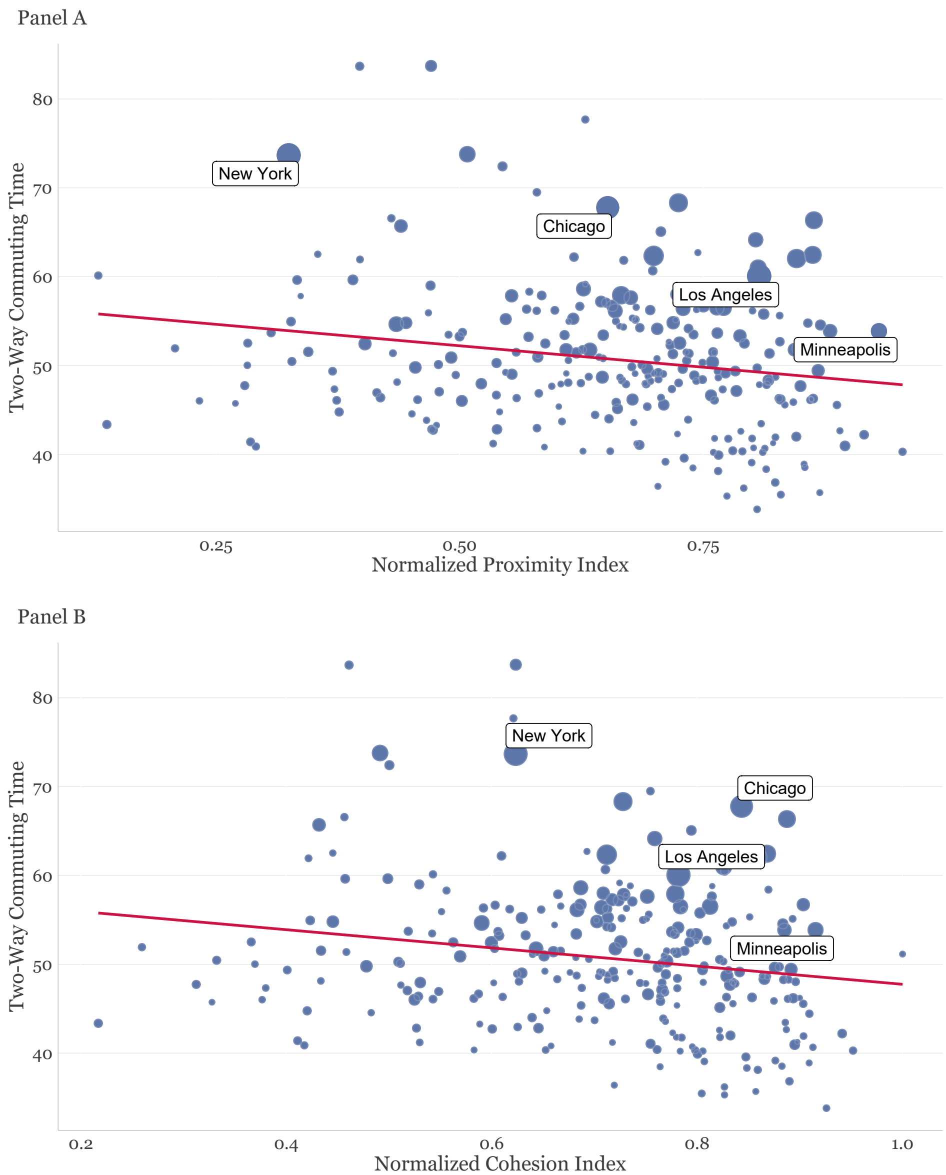


Panel B: Minneapolis Urban Area



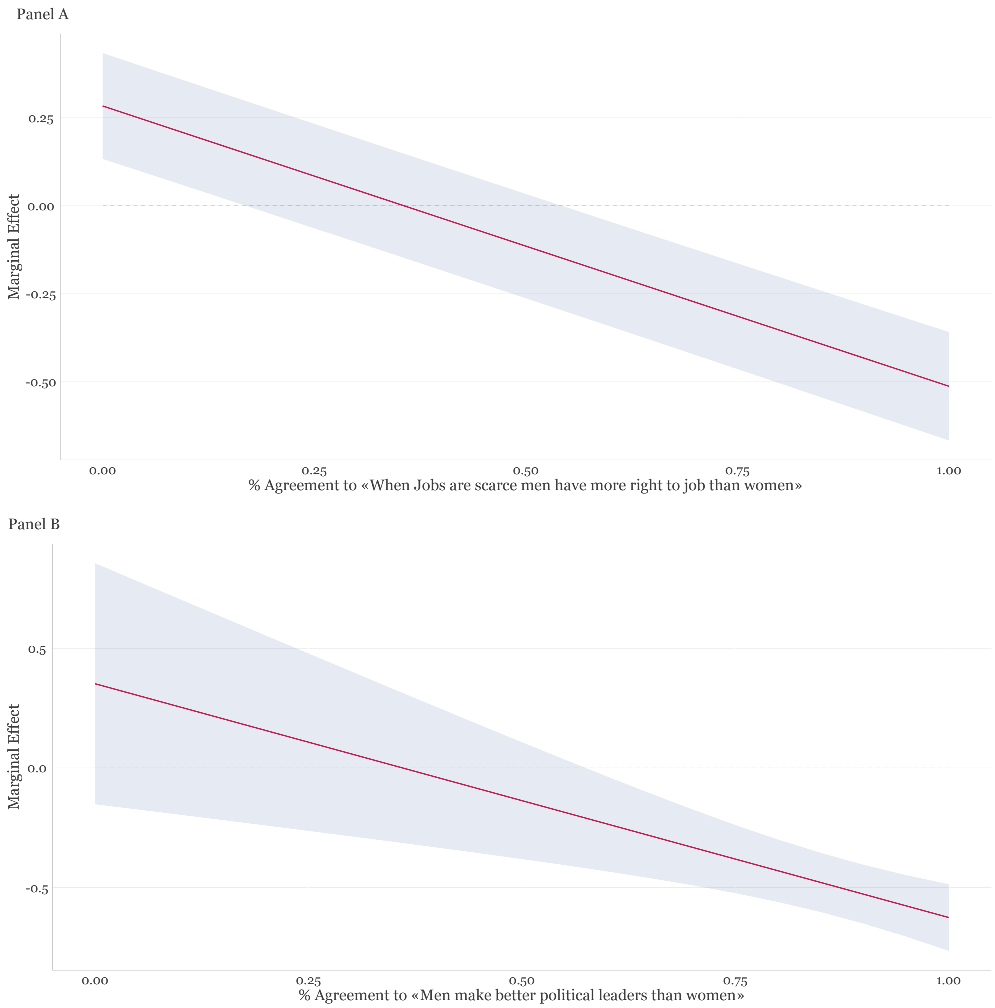
Note: The dark area represents the city's actual shape. The circle contains the city area (i.e., Equivalent Area Circle, EAC). The point represents the city's Central Business District. The values of the normalized proximity and cohesion indices are, respectively, 0.695 and 0.843 for Chicago, and 0.930 and 0.915 for Minneapolis. Source: NHGIS Urban Area Maps, 2000.

Figure A3: Normalized and Cohesion Proximity Index and Commuting Time



Note: The graph plots the value of the normalized proximity (Panel A) and cohesion indices (Panel B) and the average two-way commuting time for working non-Hispanic white men in each MSA. The size of the circle represents the population size. Source: 5% 2000 US census IPUMS sample.

Figure A4: Marginal Effect of Commuting on the Labor Supply of Married US Immigrant Women



Note: This graph plots the marginal effect (implied by estimates in Table 8) of an increase in commuting time on the labor force participation of married women as a function of gender attitudes in the country of origin. Shadowed bands are 95% confidence intervals.

Online Appendix C: Tables

Table A1: Effect of Commuting Time on the Intensive Margin for Married Men

Outcome	(1)	(2)	(3)	(4)	(5)	(6)
	OLS Estimates		IV Estimates			
Weekly hours worked	0.458	0.511	0.521	-1.458	-0.145	-1.428
	(0.549)	(0.423)	(2.480)	(1.702)	(2.906)	(2.233)
Part-time	-0.004	0.001	0.047**	0.045**	0.040*	0.041*
	(0.004)	(0.004)	(0.019)	(0.019)	(0.024)	(0.023)
Long hours	0.011	0.022	0.111	0.028	-0.007	-0.037
	(0.029)	(0.022)	(0.127)	(0.083)	(0.143)	(0.114)
Part-time occupation (10%)	-0.011***	-	0.007	-	0.001	-
	(0.004)	-	(0.013)	-	(0.016)	-
Part-time occupation (25%)	-0.047***	-	0.090*	-	0.001	-
	(0.016)	-	(0.054)	-	(0.016)	-
Observations	902,447	902,447	902,447	902,447	902,447	902,447
F-stat (Excl. Instr)			17.421	17.863	10.162	10.218
Instrument			nProximity	nProximity	nCohesion	nCohesion
Occupation Fixed Effects	NO	YES	NO	YES	NO	YES

Notes: Hours worked refers to the regular number of hours worked per week. Part-time is defined as working less than 35 hours per week. Long hours is defined as working more than 50 hours per week. Part-time occupation (10%) and (25%) are indicator variables that equal 1 if the occupation is in the top 10th or 25th percentile of the distribution of part-time employment across all occupations. Commuting is the average MSA two-way commuting time in minutes for non-Hispanic white male workers divided by 100. The sample is restricted to married men between 25 and 55 years old who worked more than 35 weeks of the last year. All regressions include population and its squared term in logs. Individual and MSA controls as defined in Table 3. Occupation fixed effects controls for 1990 census occupational group fixed effects. Columns 3 and 4 use the nProximity index as instrument, while columns 5 and 6 use the nCohesion index as instrument. F-stat is the value of the statistic for the test of excluded instruments. Standard errors clustered at the MSA level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A2: Effect of Commuting Time on Labor Force Participation of Married Women and Men, Non-Movers Sample – 2SLS (Proximity index)

Panel A: Married Women	(1)	(2)	(3)	(4)
Commuting	-0.353**	-0.345**	-0.396***	-0.394***
	(0.150)	(0.135)	(0.128)	(0.084)
Observations	395,689	395,689	395,689	395,689
F-stat (Excl. Instr)	22.195	14.408	14.824	26.012
Panel B: Married Men				
Commuting	-0.105	-0.126*	-0.095*	-0.072
	(0.087)	(0.077)	(0.049)	(0.047)
Observations	362,215	362,215	362,215	362,215
F-stat (Excl. Instr)	22.106	14.311	14.759	26.122
Region Dummies	NO	YES	YES	YES
Individual Controls	NO	NO	YES	YES
MSA Controls	NO	NO	NO	YES

Notes: The dependent variable is a binary indicator that takes value 1 if the individual is in the labor force and 0 otherwise. Commuting is the average MSA two-way commuting time in minutes for non-Hispanic white male workers divided by 100. The sample is restricted to married men and women between 25 and 55 years old, born in the same state they currently live in and who have not changed residence in the last 5 years. Panel A shows results for married women. Panel B shows results for married men. All regressions include population and its squared term in logs. Individual and MSA controls as defined in Table 3. Estimations are based on the normalized proximity index. F-stat is the value of the statistic for the test of excluded instruments. *** p<0.01, ** p<0.05, * p<0.1.

Table A3: Effect of Commuting Time on the Gender Gap in Labor Force Participation – 2SLS (Proximity Index)

	(1)	(2)	(3)	(4)
Commuting	0.982***	0.894***	0.802***	0.939***
	(0.292)	(0.229)	(0.294)	(0.307)
Observations	272	272	272	272
F-stat (Excl. Instr)	16.663	28.495	26.208	17.548
Population	NO	YES	YES	YES
Individual Controls	NO	NO	YES	YES
MSA Controls	NO	NO	NO	YES

Notes: The dependent variable is the ratio between male and female labor force participation at the city level. Commuting is the average MSA two-way commuting time in minutes for non-Hispanic white male workers divided by 100. Data come from collapsing 5% 2000 US census IPUMS data at the MSA level and the 2000 CCDB. Population includes log population and its square term. Individual and MSA controls as defined in Table 3. Individual controls are collapsed averages from the same individual controls as in Table 3: share of college education, mean age and share of white people. The estimations are based on the normalized proximity index. F-stat is the value of the statistic for the test of excluded instruments. Robust standard errors in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A4: Effect of Commuting Time on the Labor Force Participation of Married Women and Men – First Differences, OLS and 2SLS (Proximity Index)

Panel A: OLS	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Married Women				Married Men			
ΔCommuting	-0.181**	-0.230**	-0.210***	-0.149**	-0.005	-0.019	0.007	-0.003
	(0.086)	(0.090)	(0.071)	(0.069)	(0.056)	(0.062)	(0.039)	(0.037)
Panel B: 2SLS								
	Married Women				Married Men			
ΔCommuting	-0.799**	-0.830***	-0.772**	-1.000**	-0.063	-0.08	-0.012	-0.028
	(0.327)	(0.308)	(0.311)	(0.476)	(0.168)	(0.165)	(0.131)	(0.184)
Observations	238	238	238	238	238	238	238	238
F-Stat (Excl. Instr)	10.067	10.302	9.683	5.097	10.067	10.302	9.683	5.097
Population	NO	YES	YES	YES	NO	YES	YES	YES
Individual Controls	NO	NO	YES	YES	NO	NO	YES	YES
MSA Controls	NO	NO	NO	YES	NO	NO	NO	YES

Notes: The dependent variable is the percentage point change in married female and male labor force participation between 2000 and 1980. Pan Δ Commuting is the change in commuting time between 1980 and 2000. Panel A shows OLS estimates. Panel B shows 2SLS estimates based on the normalized proximity index. Data comes from collapsing 5% 2000 and 1980 US census IPUMS data at the MSA level, the 2000 CCDB and NHGIS. Population is the 1980 log population and its squared term. Individual and MSA controls as defined in Table **??**. Individual controls are collapsed and differenced averages from the same individual controls as in Table **??**: share of college education, mean age and share of white people. MSA controls include income growth rate, the change in the share of poor residents, the change in population and the change in the share of college graduates between 1980 and 2000. F-stat is the value of the statistic for the test of excluded instruments. Robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A5: Effect of Commuting Time on the Labor Force Participation of Married Women and Men – 2SLS (Proximity Index)

	(1)	(2)	(3)	(4)
	2006-2011	2000	1990	1980
	Married Women			
Commuting	-0.546***	-0.442***	-0.445***	-0.137
	(0.182)	(0.102)	(0.137)	(0.170)
Observations	1,394,373	1,382,904	1,156,111	1,076,704
F-stat (Excl. Instr)	10.301	16.645	17.123	21.625
	Married Men			
Commuting	-0.010	-0.024	-0.039	-0.007
	(0.052)	(0.072)	(0.036)	(0.037)
Observations	1,313,411	1,349,163	1,124,872	1,057,073
F-stat (Excl. Instr)	9.906	15.843	16.493	21.061

Notes: The dependent variable is a binary indicator that takes value 1 if the individual is in the labor force and 0 otherwise. Commuting is the average MSA two-way commuting time in minutes for non-Hispanic white male workers divided by 100. The sample is restricted to married women and men between 25 and 55 years old. All regressions include population and its squared term in logs. Individual and MSA controls as defined in Table 3. Estimations are based on the normalized proximity index. F-stat is the value of the statistic for the test of excluded instruments. *** p<0.01, ** p<0.05, * p<0.1. Sample: 5% US 2000, 1990 and 1980 census IPUMS and 2006-2011 Pooled ACS