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Cities

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COMMUTING TIME AND THE GENDER GAP IN LABOR MARKET PARTICIPATION *

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ABSTRACT: This paper investigates the contribution of increasing travel times to the persistent gender gap in labor market participation. In doing so, we estimate the labor supply elasticity of commuting time from a sample of men and women in US cities using microdata from the Census for the last 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 minutes increase in commuting decreases the probability of married women to participate in the labor market by 4.6 percentage points. In contrast, the estimated effect on men is small and statistically insignificant. We also find that women with children and immigrant women originating from countries with more gendered social norms respond the most to commuting time variations. This evidence suggests that the higher burden of family responsibilities supported by women may magnify the negative effect of commuting on their labor supply. From our findings, we conclude that the increasing trend in travel times observed in the US and in many European countries during the last decades may have contributed to the persistence of gender disparities in labor market outcomes.

JEL Codes: R41, J01, J16, J22

Keywords: Commuting time, labor supply, gender roles, family responsibilities, city shape

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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 reached a plateau since the early 2000s. The unequal distribution of the family burden and the persistence of gender norms that reinforce the role of women as main caregivers are a promising candidate to account for the remaining gender gaps in the labor market (Kleven et al., 2019a,b).

In this paper we propose a complementary explanation to account for the persistence of gender inequality, namely, the asymmetric effect of commuting costs on the labor supply of men and women. High commuting costs will never induce an individual to join the labor force but it may discourage participation. For example, in a two-member household, the presence of long commutes may foster specialization by family members in either market or home production to avoid paying the cost of going to work twice (Black et al., 2014). We argue that the gendered distribution of household tasks and the prevalence of social norms that make more costly for men to stay home may have contributed to the stagnation of female labor 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 of a full-time worker was 45 minutes. By 2016, it had increased to 54 minutes (i.e. a 26% higher). Also in 2016, 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 a very undesirable activity for workers (Kahneman et al., 2004; Clark et al., 2019), detrimental to 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 labor supply elasticity of commuting time 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 home or market production in the presence of costly commuting. The model also predicts who withdraws from the market on the basis of differences in labor and home productivity, and the presence of social norms about the role of men and women in society. Consistently 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 for women monotonically increases with the number of children, suggesting that family responsibilities are important in explaining within-household specialization in the presence of long commutes. In contrast, we do not find significant differences in the response across skill

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

groups, indicating that differences in labor market productivity are not responsible for the gender asymmetry in the results. Finally, we focus on a sample of immigrant women in the US. We report larger estimates of the commuting time elasticity among those originating from countries with more traditional gender views. This evidence is consistent with the prevalence of social norms that reinforce the male breadwinner model and contribute to the persistence of gender inequality in a context of increasing commuting costs.

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 the individual labor supply, we follow Harari (2016) and use the shape of cities as an exogenous source of variation. We focus on city compactness measured by how close the shape of the city is to 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 regularity to identify the effect of travel times on labor supply decisions (Angel et al., 2010). The main threat to our identification strategy is that a city feature correlated with its shape might have a direct effect on individuals' labor supply. To take into account this concern, 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 have recently moved to a city. 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 vs married women without children).

A few recent studies have already identified a role for commuting costs in explaining the remaining 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 vicinity to workplaces. This is consistent with the view that women attribute a higher value to short commutes than men. For France, Le Barbanchon et al. (2019), using a job search model where commuting matters, estimate that gender differences in the willingness to commute explain about a 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 Black et al. (2014) as it documents 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 geographical accidents that shape cities. Our IV estimates indicates that the effect of commuting is larger than the suggested by OLS. Second, we show that the effect on women increases with the family burden and is stronger among immigrant women originating from countries with more gendered social norms. In contrast, we do not find evidence that wage differentials can explain the gender asymmetry in the response to commuting costs. We conclude that the unequal distribution of family responsibilities and the presence of social norms about the role of men and women in society explain why gender neutral commuting costs affect men and women differently and contribute to the persistence of the gender inequality in the labor market.

The remainder of the paper is organized as follows. Section 2 presents a theoretical framework for the labor force participation decision 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 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. The decision of which member exits depends on: i) differences in productivity in the market and at home and ii) the presences of gendered social norms about the role of men and women in society. We present the model and discuss the main results below and defer to Appendix A the model solution.

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 ($1 = h_j + l_j + k$), 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, time at home 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 is decreasing with commuting costs in both cases, but the slope is more negative when the two 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 symmetric 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, an $\alpha_m > \alpha_f$ 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 favor 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 in 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 city in the 2000 Census (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 focus on measures at the city-level. The first row shows the summary statistics for commuting times. We follow Black et al. (2014) and compute the city-average two-way commute time using the information reported in the Census that asks directly about door-to-door travel time in minutes.[3] As Black et al. (2014) we restrict the sample to white male workers to measure commuting as this group has the highest employment rate. In 2000, the average commuting time is 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, city income, share of employment in manufacturing, share of public employment, share of people below the poverty threshold, share of college educated and the gender wage gap. These variables are computed from the US county and City Data Book (CCDB) in 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 men and women. The first raw displays the descriptives for labor force participation, followed by other indicators of the intensive margin of the labor supply: number of weekly hours worked, part-time employment, an indicator for working long hours (more than 50 hours per week) and the probability of working in an occupation with a high concentration of part-time employment (i.e. an occupation at the top 10th or 25th percentile of the distribution of part-time employment across all occupations). According to the descriptives in the table, the gender gap in participation in 2000 was almost 10 percentage points (80% for men and 70% for women). Part-time employment was also much more prevalent among women than men (17% versus 2%). Men also worked on average more hours 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 shows the descriptive statistics for the individual controls included in estimation: age, spousal income, presence of children, having a college degree and race.

Finally, we employ the World Value Survey (WVS) to measure the gender role attitudes in the country of origin of US immigrants. To increase the number of countries in the sample, we pool the different surveys in the WVS conducted between 2000 and 2011. Following Alesina et al. (2013) 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 *agree* or *strongly agree* with the statements. Figure 2 indicates a clear positive correlation in the responses to the two statements. It also shows a substantial degree of heterogeneity across countries. Accordingly, the

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

[3]Notice that this definition includes all modes of transportation.

share 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 about 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 costs 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 one if individual i living in MSA c participates in the labor market and zero otherwise. The explanatory variable $commuting_c$ is the average two-way commuting time 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: spousal income (in logs), an indicator for the presence of children, age, race and educational attainment dummies.[4] X_c are control variables at the MSA level such as the % of employment in the manufacturing sector, the % of public employment, median household income (in logs), the gender wage gap, 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 reversed causality bias pushing the OLS estimates towards zero. Second, sorting of individuals across cities also represents a threat 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 the 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,

[4] Educational attainment dummies corresponds to the 12 categories of completed education defined in the Census.

2016). 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 measures of city compactness proposed by Angel et al. (2010) and also employed in Harari (2016) 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).[5] Accordingly, the *proximity* index measures the average distance to the CBD.[6] Note that this measure emphasizes commuting trips to 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.[7] 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 building 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.[8] 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. Notice that by construction the normalized proximity and cohesion indices are uncorrelated with

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

[6] A few MSAs have more than one principal city. In these cases, we compute the popultion weighted average *proximity* index.

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

[8] 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.

the size and the density of the city. A value of the normalized index close to 1 means that the city index is close to the optimal EAC index (i.e. circular city). Lower values, instead, indicate that the shape of the city is less circular. Figure 3 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. 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. 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 tends 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 4 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 restrict our sample to 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.[9] Table 2 shows the results for the probability of having a college degree (columns 1 and 2), the probability of being a power couple (columns 3 and 4), the number of children (columns 5 and 6) and the probability of being married (columns 7 and 8). The results indicates that none of the normalized index is correlated with the observed individual characteristics considered, and support the validity of the proposed identification strategy.

[9] 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).

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. Table 3 shows the OLS estimates of the empirical model in equation 1. Columns 1 to 4 focus on women and columns 5 to 8 on men. 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 % of employment in the manufacturing sector, the % of workers employed in the public sector, median household income (in logs) and the gender wage gap.

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 minutes increase in travel time leads to a 1.8 percentage points decrease in the probability of married women to participate in the labor market. The estimates for men are much smaller in magnitude and only statistically significant in the last specification. As discussed in Section 4, reversed causality and sorting are likely to bias the OLS estimate towards zero. Accordingly, the estimates in Table 3 should be interpreted as a lower bound of the effect of commuting costs on labor force participation.

5.2 IV estimates

To deal with the 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), a one standard deviation increase 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 widely accepted by practitioners (Angrist and Pischke, 2008). Instead, 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 size 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 Table 5. Panel A reports the estimates based on the normalized proximity index while Panel B those of the cohesion index. As in Table 3, columns 1 to 4 show the results for married women and columns 5 to 8 for married men. For women, the estimates in all columns are larger (in absolute value) than the corresponding OLS estimates. Also, the point estimates remains 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 minutes increase in commuting decreases the probability for a married women to participate in the labor market by 4.6 percentage points. In 2000, the participation rate of prime age women was 73%, and 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 is mostly concentrated on women. Despite being insignificant, the magnitude and sign of the estimated coefficient suggests that longer commutes may also negatively affect the participation decision 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).

Table 6 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 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 decision of women, while the effect, if any, is much smaller on men. According to the point estimates in Table 5, if commuting times in the US had remained at the 1980 level (a 26% lower than in 2000), the labor force participation of married women would have been 3.7 percentage points higher, which represents about 40% of the current gender gap in participation.[10]

5.3 Effects at the intensive margin

Now we turn the analysis to the intensive margin of the labor supply. We focus on the number of weekly hours worked, the decision to work part-time and that of working long-hours. We define part-time work as working less than 35 hours during 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

[10]The current gap is at 9pp – 85 vs 76, according to 2017 ACS data.

working in a typical part-time occupation as defined in Section 3.

We are aware that the results at the intensive margin can not be interpreted as causal, as the decision to participate in the labor market is clearly affected by our variable of interest. For example, it may be that only the most talented and motivated women get a job in high commuting locations. This would bias our estimates at the intensive margin towards zero. However, we do find these results informative about the effects of commuting on the labor market beyond the decision to participate.

Table 7 presents the results at the intensive margin. Columns 1 and 2 report the OLS estimates and 3 to 6 the IV estimates.[11] In estimation, the sample is restricted to married women aged 25 to 55 working 52 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.[12]

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 minutes increase in commuting decreases hours worked per week between 0.58 and 0.72. The probability of working part-time also increases by 2.4 percentage points when commute increases by 10 minutes. In contrast, there is no effect on the probability of working long hours.[13] Note that the results on part-time employment and the number of hours worked do not decrease when occupation fixed effects are included in estimation. 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.

5.4 Robustness checks

In this section we conduct a number of empirical exercises to validate the robustness of our previous findings. The results in this section are obtained using the normalized proximity index as instrument. The corresponding results for the normalized cohesion index, which are qualitatively very similar, are presented in Appendix B (Tables A1 to A4).

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 estimation. To further explore the possibility that sorting can 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

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

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

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

or city of birth, but it reports the state of birth and records the recent migration histories. With this information, we can estimate the model on the sample of individuals who were born in the state where they live and have not changed residence during the last 5 years. The results for this sample are reported in Table 8. These estimates are very similar in magnitude and significance to those in Table 5 and provide additional evidence that sorting across MSA 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 particiaption. 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 9. The point estimates are positive and statistically significant: a 10 minutes increase in commuting increases the gender gap in participation by about 6 percentage points. This result is similar to the findings in Table 5 and reinforces 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 differences between 1980 and 2000. This specification allows us to control for unobserved MSA characteristics which are time invariant and may have heterogenous effect across genders. Table 10 presents the IV estimates of the following model for women:

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

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 shapes in US cities between 1980 and 2000 are limited and, thus, we instrument $\Delta commuting_c$ with the cross-sectional normalized proximity index described in Section 4. Table 10 presents the results. Accordingly, a 1 minute increase in commuting times reduces the labor supply fo married women between 0.79 and X, depending on the specification. This estimated effect are close to that in Tables 5, reinforcing the robustness of our previous findings.

Finally, we estimate the effect of commuting on labor supply at different points in time. Specifically, we employ the 1990 Census, the 1980 Census and the 2006-2011 ACS samples described in Section 3. The estimation results are reported in Table 11. Columns 1 and 3 employ data for the period 2006-2011 and for 1990, respectively. The point estimates are very similar in magnitude to the one obtained when using the 2000 Census (column 2). The effect for 1980 in column 4 is much smaller and statistically insignificant. One possible explanation is that with commuting times being shorter in 1980, commuting was a less relevant factor to explain labor force participation.

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

5.5 Mechanisms

The previous results uncover an important gender asymmetry in the individual responses to commuting costs. The theoretical framework in Section 2 suggest two possible mechanisms that can account for the stronger response by 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 and avoid paying the cost of going to work twice. Second, the presence of a gendered social norm 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 response to commuting costs across genders we conduct an heterogeneity analysis. We first estimate the model in equation 1 on a sample of individuals with different levels of family responsibilities proxied by the number of children in the households. The estimates are reported in Table 12. Columns 1 to 7 present the results for women and columns 8 to 14 for men. The effect for single women with no children is much smaller than our baseline estimates and statistically insignificant (column 1). Among married women, the effect is smallest for those without children: a 10 minutes increase in commuting reduces participation by 3.35 percentage points (column 2). In the presence of children the effect becomes larger: a 10 minutes increase in commuting decreases the probability to participate by 4.25 percentage points for those with 1 child (column 3). This effect increases up to 5.8 percentage points among those with 3 or more children (column 4 and 5). The estimates in column 6 shows that the effect is magnified in the presence of young children. Accordingly, a 10 minutes increase in commuting times decreases the working probability of mothers with children younger than 5 by 6.7 percentage points. This effect is smaller for women who have children older than 5 years old (column 7). For men, columns 8 to 14 display a negative coefficient but much smaller in magnitude and statistically insignificant for most of the groups. These findings suggest that the higher burden of childcare costs supported by women could be partly responsible for their larger response to commuting times.

Next, we explore whether 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 that proxy for differences in productivity. In column 1 of Table 13, we estimate the model on couples where both members have a college degree (i.e. power couples). Column 2 estimates it on couples where none of the members have college (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 couples when only the wife has college education (i.e. part power couples women). The point estimates of the labor supply elasticities are very similar across different samples. This evidence suggests that differences in productivity, or the presence of a gender wage gap, are not a major driver of the different response by men and women to commuting costs.

Finally, we investigate the implications of culture for the gender asymmetry in our findings. We conduct an epidemiological analysis that mirrors that in Fernández and Fogli (2009), and focuses on a sample of foreign-born individuals living in the US. The idea is that cultural attitudes regarding the role of women in society are transmitted across generations and individuals carry these values when moving to a new country. By comparing immigrants from different countries living in the US, one can isolate the role of culture on individual decisions.

As explained in Section 3, we measure gender role attitudes using the World Value Survey. Following Alesina et al. (2013), 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" and "Men make better political leaders than women". 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 14 displays the results when traditional views are proxied 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". To ease interpretation, commuting time and gender attitudes have been demeaned. Thus, the coefficient of commuting represents its impact at average values of gender attitudes. Note that the interaction term also needs to be instrumented. We instrument commuting times and its interaction with the proxy for gender roles with the normalized proximity index and its iteraction term with the corresponding gender roles proxy. [14]. 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 attitudes' measure and commuting times. This interaction term is negative in both columns, suggesting that the effect of commuting times on the labor force participation of married women is exacerbated by gender norms. Figure 5 shows the implied marginal effects for observed values in gender attittudes. These results are consistent with the hypothesis that social norms may lead women to stay home and take care of the family and other housework when commuting time increases.

The findings in this section allow us to conclude that differences in labor market productivity

[14] Sanderson and Windmeijer (2016) first-stage statistics for models with more than one endogenous variables are provided at the bottom of Table 14

do not seem to be driving the asymmetric gender response to commuting costs. In contrast, we find that the presence of family responsibilities reinforced by the existence of gendered social norms that make more costly for men to stay home seem responsible for the stronger response by women to increases in commuting times.

6 Conclusions

This paper investigates the impact of commuting costs on the labor supply decisions of men and women. We uncover an important gender asymmetry. Namely, women with larger family responsibilities respond the most to increases in travel times, while the effect is non-significant among men and unmarried women. We extend the model in Black et al. (2014) to investigate the potential mechanisms driving these results. The model predicts that differences in home and market productivity and/or the presence of traditional gender norms about the role of men and women in society may explain the differential effect of travel times across genders. In our empirical analysis, we do not observe important differences in the response of family members with different levels of education, suggesting that productivity gaps are not likely to explain the gender asymetry in our results. In contrast, using a sample of female immigrants living in the US, we do find that those originating from countries with more traditional views are more responsive to changes in commuting costs. We interpret this finding as evidence that the presence of a social norm that makes more costly for men to stay home may explain the larger response by married women with more family responsibilities. Our findings are relevant for policy makers, as they identify a potential cause for the stagnation of female labor force participation 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. Thus reducing commuting costs by investing, for instance, in transport infrastrucutre and urban planning can facilitate female participation in the labor market and promote gender equality.

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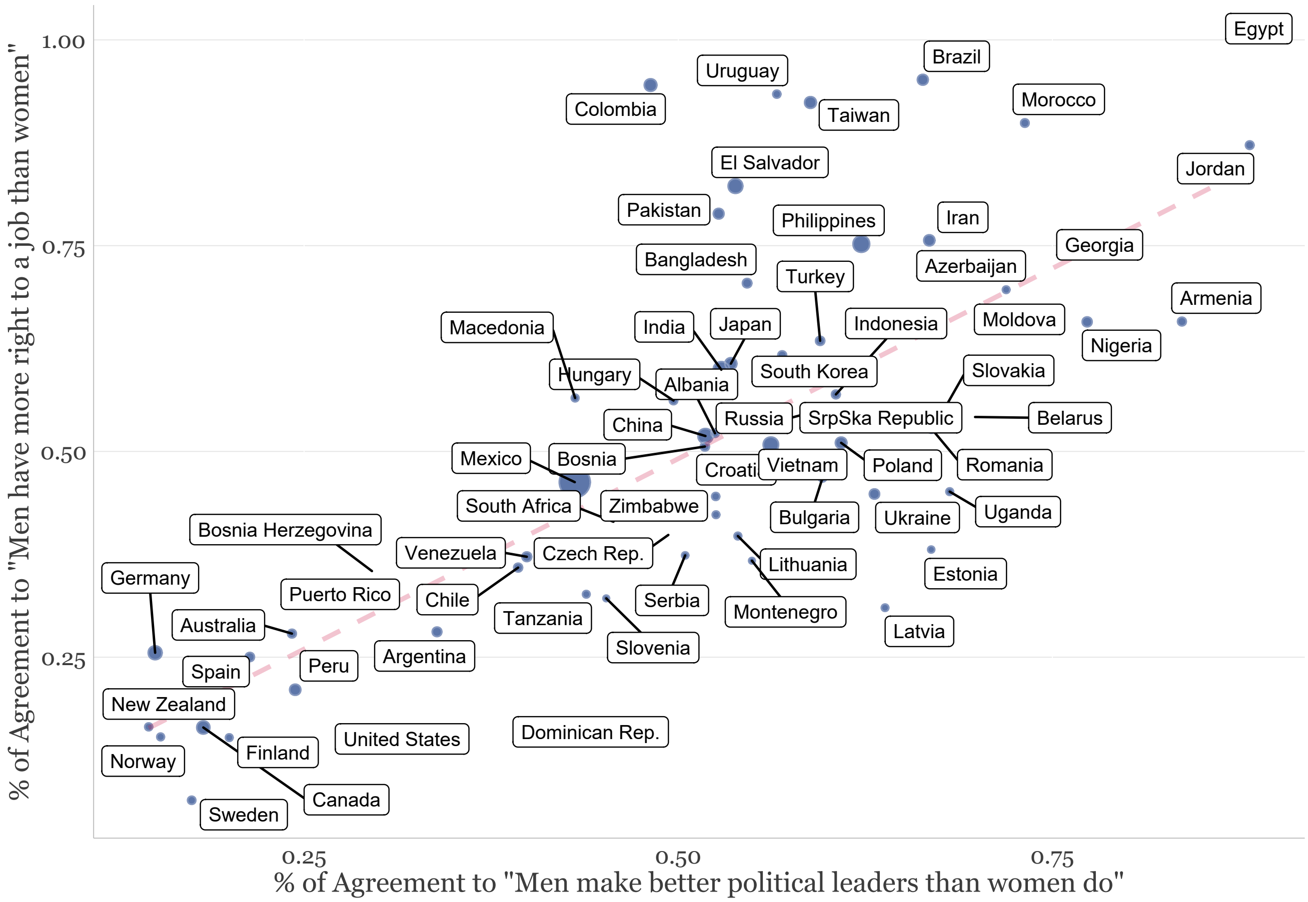
Figures

Figure 1: Commuting Costs and Household Labor Supply

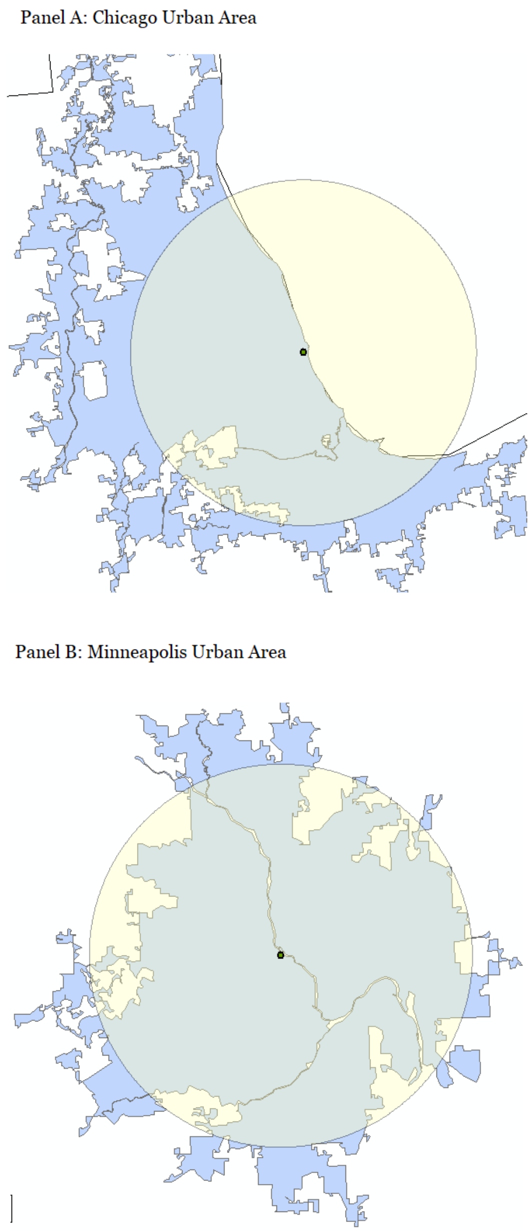
Notes: This figure displays 3 different scenarios of home and market specialization. Utility is the sum of individual utilities in the household. The black line represents the household utility when both members work. The dotted line represents the utility when only the husband works, while the dashed line represents the utility when only the wife works. 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.

Figure 2: Gender Role Attitudes in the World Value Survey



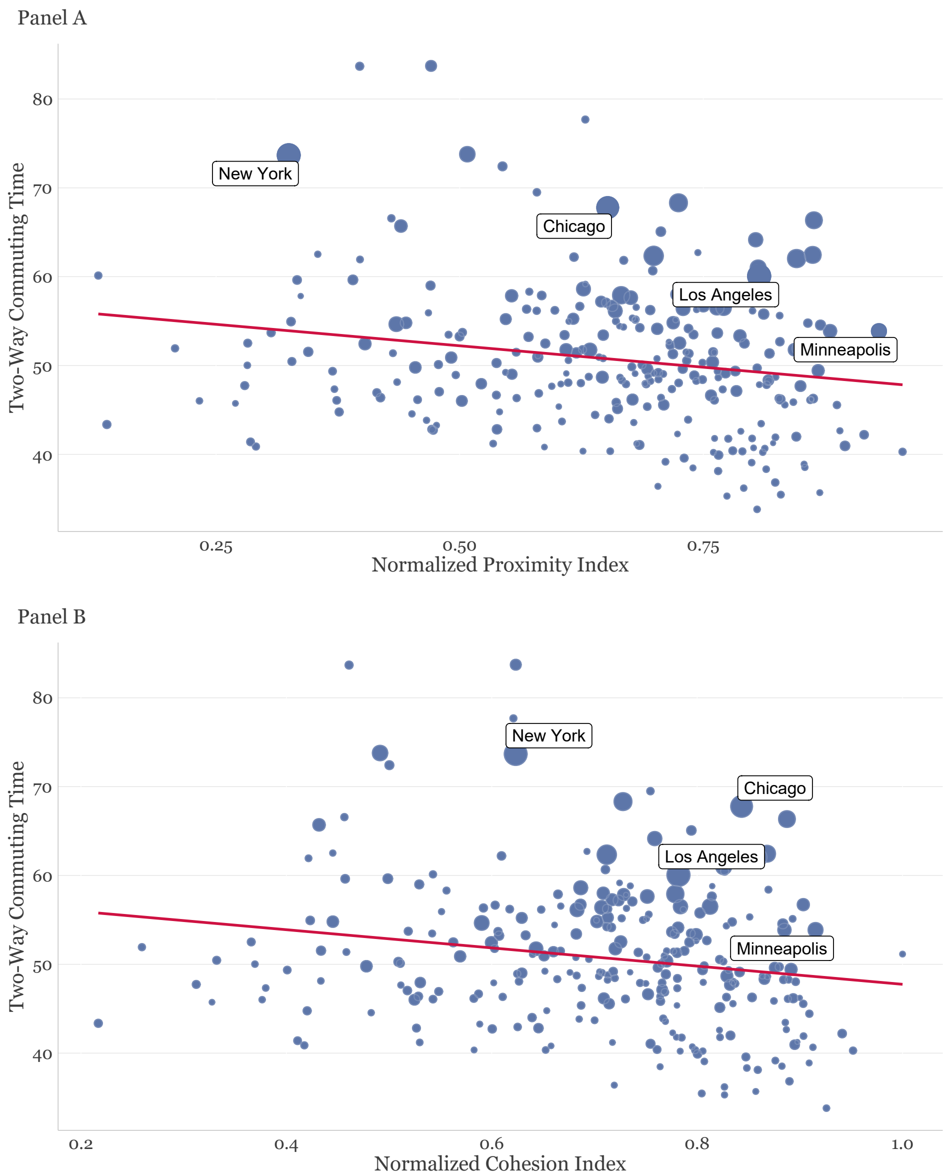
Note: The graph displays the percentage of respondents in each country that *agree* or *strongly agree* with the statement: "When Jobs are scarce, men should have more right to a job than women" (y-axis) and "Men make better political leaders than women do"(x-axis). Each point represents a country. The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

Figure 3: City Shape and Equivalent Area Circle (EAC)



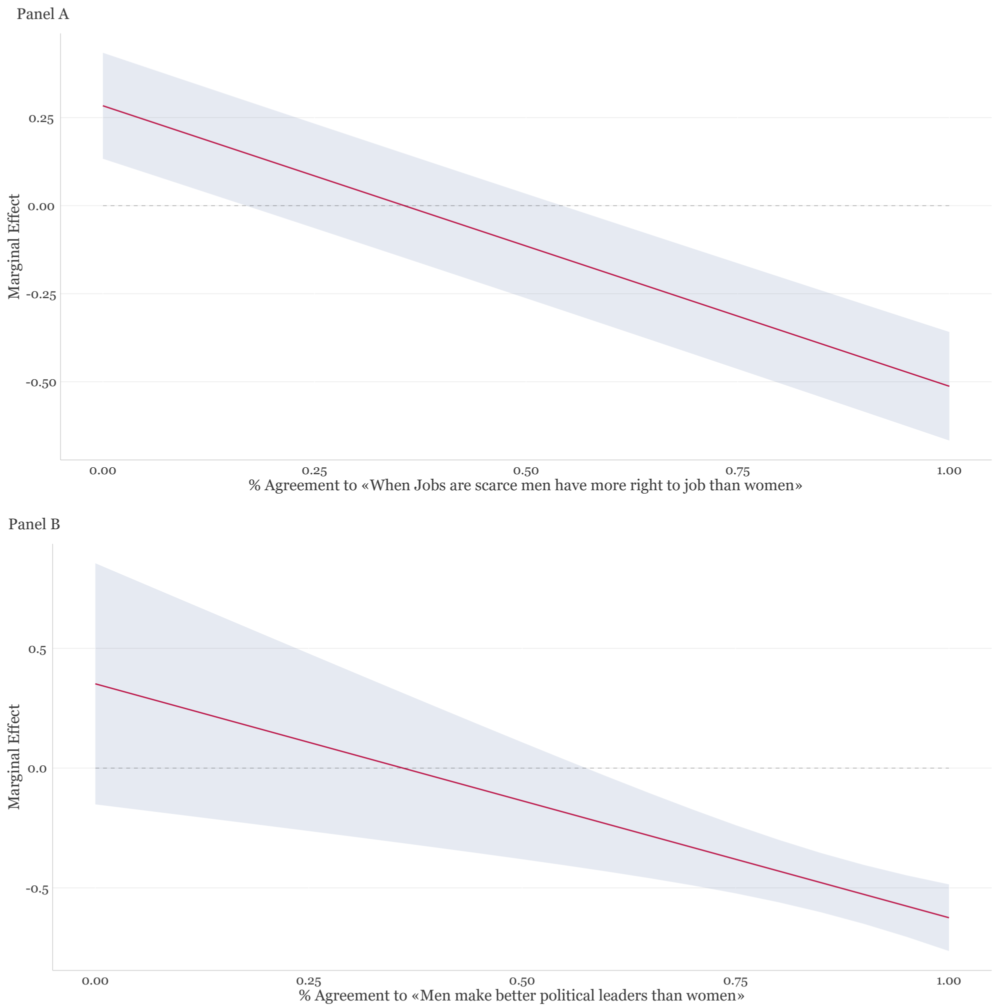
Note: The dark area represents the city actual shape. The circle contains the city area (i.e. Equivalent Area Circle, EAC). The point represents the city Central Business District. The values of the normalized proximity and cohesion indices for Chicago are 0.695 and 0.843, respectively and 0.930 and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

Figure 4: 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 white men in each MSA. The size of the circle represents the population size. Source: 5% 2000 US Census IPUMS sample.

Figure 5: 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 14) 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.

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 City 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
% Poor	272	0.13	0.04	0.06	0.35
% College education	272	0.11	0.03	0.04	0.23
Gender wage gap	272	16,486	3,458	7,405	29,450
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
Children<5	1,382,904	0.23	0.42	0	1
College	1,382,904	0.3	0.46	0	1
White	1,382,904	0.71	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
Children<5	1,349,163	0.25	0.43	0	1
College	1,349,163	0.33	0.47	0	1
White	1,349,163	0.7	0.46	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 defined as two-way door-to-door travel time to work in minutes for 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: Data in Panel A are obtained from the US County and City Data Book (CCDB). Data in Panel B from the 2000 Census. Data in Panel C form the National Historical Geographic Information System (NHGIS) in 2000.

Table 2: City Shape and Individual Characteristics of Migrants – OLS

	(1)	(2)	(3)	(4)
	College Education	Power Couple	n°Children	Married
	Married Women	All	Married Women	All Women
nProximity	-0.005	0.009	0.037	-0.003
	(0.024)	(0.022)	(0.050)	(0.012)
nCohesion	0.006	0.017	-0.005	-0.010
	(0.036)	(0.031)	(0.076)	(0.018)
Observations	236,434	484,519	236,434	399,632

Notes: The sample includes married men and women between 25 and 55 years old who changed MSA of residence in the last 5 years. Each column corresponds to a different dependent variable (i.e. the probability of having a college degree (column 1), the probability of being a power couple (2), the number of children (3) and the probability of being married (4). The rows present the estimated coefficient for different regressions on the normalized proximity index (nProximity) and the normalized cohesion index (nCohesion). All regressions include controls at the individual level (age, education and race dummies and spouse income in logs), controls at the MSA level (% employment in manufacturing, % employment in the public sector, MSA median household income in logs, the gender wage gap and population and its squared term in logs) and regional dummies. Standard errors clustered at the MSA level 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 – OLS

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Married Women				**Married Men**			
Commuting	-0.082	-0.063	-0.124***	-0.176***	-0.021	0.010	-0.008	-0.049***
	(0.093)	(0.052)	(0.026)	(0.025)	(0.072)	(0.040)	(0.019)	(0.015)
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 for 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 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, MSA median household income in logs and gender wage gap. Standard errors clustered at the MSA level 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***		-0.099***	
	(0.037)		(0.025)	
nCohesion		-0.168***		-0.106***
		(0.046)		(0.034)
Observations	1,382,904	1,382,904	1,382,904	1,382,904
F-stat (Excl. Inst.)	24.46	13.56	16.72	9.82
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 two-way commuting time for white male workers at the MSA level 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 in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 5: Effect of Commuting Time on Labor Force Participation – 2SLS

Panel A– Instrument: nProximity	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Married Women				**Married Men**			
Commuting	-0.301	-0.394**	-0.392***	-0.465***	-0.131	-0.133	-0.064	-0.065
	(0.242)	(0.191)	(0.121)	(0.113)	(0.163)	(0.120)	(0.060)	(0.068)
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
F-Stat (Excl. Instr.)	19.740	12.836	13.158	16.270	19.395	12.529	12.741	15.467
Panel B– Instrument: nCohesion								
	Married Women				**Married Men**			
Commuting	-0.491**	-0.444*	-0.357***	-0.388***	-0.321**	-0.226	-0.113	-0.127
	(0.207)	(0.235)	(0.137)	(0.117)	(0.134)	(0.147)	(0.0760)	(0.0863)
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
F-stat (Excl. Instr)	13.558	7.453	7.579	9.828	13.269	7.339	7.416	9.535
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 for 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. In Panel A the instrument is the normalized proximity index while in Panel B it is the normalized cohesion index. F-stat is the value of the statistic for the test of excluded instruments. Standard errors clustered at the MSA level in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 6: Effect of City Shape on Labor Force Participation – Reduced Form Estimates

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Married Women				**Married Men**			
nProximity	0.050**	0.052***	0.052***	0.046***	0.022	0.018	0.008	0.006
	(0.034)	(0.017)	(0.009)	(0.009)	(0.025)	(0.014)	(0.007)	(0.007)
nCohesion	0.083***	0.056***	0.045***	0.041***	0.054***	0.029*	0.014	0.013
	(0.029)	(0.022)	(0.014)	(0.011)	(0.02)	(0.017)	(0.009)	(0.009)
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. The sample is restricted to married women and men between 25 and 55 years old. Individual and MSA controls as defined in Table 3. Standard errors clustered at the MSA level in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 7: Effect of Commuting Time on the Intensive Margin, Married Women – 2SLS

Outcome	(1)	(2)	(3)	(4)	(5)	(6)
	OLS Estimates		IV Estimates			
Hours Worked	-0.809	-0.826	-5.801**	-6.840**	-6.405*	-7.202*
	(0.600)	(0.573)	(2.501)	(2.671)	(3.804)	(3.925)
Part-time	0.027	0.030	0.225**	0.243**	0.227	0.238*
	(0.026)	(0.024)	(0.107)	(0.106)	(0.146)	(0.144)
Long Hours	-0.003	0.001	-0.019	-0.060	-0.075	-0.108
	(0.012)	(0.011)	(0.053)	(0.052)	(0.079)	(0.081)
Part Time Occupation (10%)	-0.019	–	-0.037	–	-0.036	–
	(0.022)	–	(0.078)	–	(0.088)	–
Part Time Occupation (25%)	-0.018	–	0.007	–	0.036	–
	(0.022)	–	(0.068)	–	(0.078)	–
Observations	602,811	602,811	602,811	602,811	602,811	602,811
F-stat (Excl. Instr)	–	–	19.011	19.086	11.446	11.454
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 at 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 for white male workers divided by 100. The sample is restricted to married women between 25 and 55 years old who worked 52 weeks during 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 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 in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 8: Effect of Commuting Time on Labor Force Participation of married women, Non-Movers Sample – 2SLS (Proximity index)

	(1)	(2)	(3)	(4)
Commuting	-0.353***	-0.345***	-0.439***	-0.427***
	(0.150)	(0.135)	(0.137)	(0.091)
Observations	395,689	395,689	395,689	395,689
F-stat (Excl. Instr)	22.195	14.407	14.846	27.027
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 for white male workers divided by 100. The sample is restricted to married women between 25 and 55 years old, born in the same state as they currently live and who did not change residence during the last 5 years. 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. The estimates for the cohesion index are in Table A1. F-stat is the value of the statistic for the test of excluded instruments. *** p<0.01, ** p<0.05, * p<0.1.

Table 9: 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***	0.915***
	(0.292)	(0.229)	(0.294)	(0.295)
Observations	272	272	272	272
F-stat (Excl. Instr)	16.663	28.495	26.208	17.556
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 two-way commuting time at the city level for white male workers. Data comes 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. The estimates for the cohesion index are in Table A2. 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 10: Effect of Commuting Time on the Labor Force Participation of Married Women – First Differences, 2SLS (Proximity Index)

	(1)	(2)	(3)	(4)
ΔCommuting	-0.799**	-0.830***	-0.772**	-0.992**
	(0.327)	(0.308)	(0.311)	(0.532)
Observations	238	238	238	238
F-stat (Excl. Instr)	10.067	10.302	9.683	4.978
Population	NO	YES	YES	YES
Individual Controls	NO	NO	YES	YES
MSA Controls	NO	NO	NO	YES

Notes: The dependent variable is the percentage point change in married women labor force participation between 2000 and 1980. Δ Commuting is the change in commuting time between 1980 and 2000. 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 3. Individual Controls are collapsed and differenced averages from the same individual controls as in Table 3: Share of college education, mean age and share of white people. MSA controls include income growth rate, the change in the share of poor, the change in gender wage gap, the change in population and the change in the share of college graduates between 1980 and 2000. Estimations are based on the normalized proximity index. The estimates based on the cohesion index are in Table A3. 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 11: Effect of Commuting Time on the Labor Force Participation of Married Women – 2SLS (Proximity Index)

	(1)	(2)	(3)	(4)
	2006-2011	2000	1990	1980
Commuting	-0.560***	-0.465***	-0.477***	-0.135
	(0.197)	(0.113)	(0.138)	(0.147)
Observations	1,394,373	1,382,904	1,156,111	1,076,704
F-stat (Excl. Instr)	8.82	16.270	16.718	22.089

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 for white male workers divided by 100. The sample is restricted to married women 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. The results for the cohesion index are presented in Table A4. 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

Table 12: 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	(8) Single	(9) w/o Children	(10) 1 Child	(11) 2 Children	(12) $\geq$ 3 Children	(13) Children <5yo	(14) Children >5yo
	Married Women							**Married Men**						
Commuting	-0.149	-0.335***	-0.425***	-0.585***	-0.580***	-0.670***	-0.377***	-0.075	-0.068	-0.148***	-0.052	0.018	-0.074	-0.069
	(0.107)	(0.103)	(0.115)	(0.137)	(0.186)	(0.177)	(0.106)	(0.062)	(0.156)	(0.052)	(0.057)	(0.070)	(0.0621)	(0.047)
Observations	458,578	403,069	335,857	402,550	241,428	403,069	658,380	648,343	399,091	319,676	394,655	235,741	340,311	609,761
F-stat (Excl. Instr)	18.644	17.174	17.806	16.347	12.211	14.786	16.172	17.410	15.319	17.238	16.076	12.024	14.547	15.820

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 for 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 in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 13: 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***	-0.417***	-0.681***	-0.443***
	(0.127)	(0.109)	(0.214)	(0.119)
Observations	293,063	760,698	201,318	127,825
Adjusted R-squared	0.022	0.039	0.014	0.015
F-test (Excl. Instr)	19.374	13.943	14.862	18.535
Panel B: **Married Men**				
Commuting	-0.124	0.019	-0.135***	0.081
	(0.085)	(0.081)	(0.048)	(0.291)
Observations	282,803	729,392	156,879	180,089
F-stat (Excl. Instr)	19.498	13.267	15.477	14.051

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 for 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 in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 14: Effect of Commuting Time on the Labor Force Participation of Immigrant Married Women – 2SLS

LFP married women interacted with	(1) *When jobs are scarce...*	(2) *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 for 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. 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) in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

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:

$$y^{bothwork} = (1-k)\Big(\frac{w_m - w_f}{2}\Big) \tag{A1}$$
$$c_j^{bothwork} = \alpha_j(1-k)\Big(\frac{w_m + w_f}{2}\Big) \tag{A2}$$
$$l_j^{bothwork} = (1-\alpha_j)(1-k)\Big(\frac{w_m + w_f}{2w_j}\Big) \tag{A3}$$

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.

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 amount of consumption and leisure of the working spouse as well as the transfer is determied by:

$$c_m^{oneworks} = \frac{\alpha_m}{1+\alpha_f} w_m(1-k) \tag{A4}$$
$$l_m^{oneworks} = \frac{1-\alpha_m}{1+\alpha_f}(1-k) \tag{A5}$$
$$c_f^{oneworks} = \frac{\alpha_f}{1+\alpha_f} w_m(1-k) \tag{A6}$$

Commuting costs also decrease the consumption and the amount of time 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:

$$\frac{d\left(u_m(c_m^{bothwork}, l_m^{bothwork}) + u_f(c_f^{bothwork}, l_f^{bothwork})\right)}{dk} = \frac{-2}{1-k} \tag{A7}$$

$$\frac{d\left(u_m(c_m^{oneworks}, l_m^{oneworks}) + u_f(c_f^{oneworks}, l_f^{oneworks})\right)}{dk} = \frac{-1-\alpha_f}{1-k} \tag{A8}$$

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 turn to 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 A4 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 A5 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 low wage 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 high α_j person works. As for the utility derived from time at home, it turns out that the utility loss experienced by the woman if she works, $\alpha_f(ln(1) - ln(l_f^{oneworks}))$, is larger than the utility loss experienced by the man if he works, $\alpha_m(ln(1) - ln(l_m^{oneworks}))$. Hence, if $\alpha_m > \alpha_f$, the woman will stay at home as they both enjoy higher consumption levels and the cost of spending less time at home is lower for the man.

Appendix B

Table A1: Effect of Commuting Time on the Labor Force Participation of Married Women. Non-Movers Sample – 2SLS (Cohesion index).

	(1)	(2)	(3)	(4)
Commuting	-0.407**	-0.435**	-0.458***	-0.392***
	(0.188)	(0.192)	(0.176)	(0.138)
Observations	395,689	395,689	395,689	395,689
Region Dummies	NO	YES	YES	YES
Individual Controls	NO	NO	YES	YES
MSA Controls	NO	NO	NO	YES
F-test	14.079	7.716	7.743	13.735

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 for white male workers, divided by 100. The sample is restricted to married women between 25 and 55 years old, born in the same state as they currently live and who did not change residence during the last 5 years. All regressions include log population and its squared term as controls. Individual controls are: age, education level and race dummies and the log of spouse income. MSA controls include: the share of employment in the manufacturing sector, the share of employment in the public sector, the MSA median income and gender wage gap. Estimations are based on the normalized cohesion index. F-test is the value of the statistic for the test of excluded instruments. Standard errors clustered at the MSA level in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A2: Effect of Commuting Time on the Gender Gap in Labor Force Participation – 2SLS (Cohesion Index)

	(1)	(2)	(3)	(4)
Commuting	1.140***	0.980***	0.859***	1.005***
	(0.352)	(0.268)	(0.319)	(0.349)
Observations	272	272	272	272
Population	NO	YES	YES	YES
Individual Controls	NO	NO	YES	YES
MSA Controls	NO	NO	NO	YES
F-Stat (Excl. Instr.)	13.189	24.975	24.277	16.794

Notes: The dependent variable is the ratio between male and female labor force participation at the city level. Commuting is the average two-way commuting time at the city level for white male workers. Data comes 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: Share of college education, mean age and share of white people. The estimations are based on the normalized cohesion index. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A3: Effect of Commuting Time on the Labor Force Participation of Married Women–First Differences 2SLS (Cohesion Index)

	(1)	(2)	(3)	(4)
ΔCommuting	-0.839*	-0.803*	-0.825*	-0.842*
	(0.344)	(0.318)	(0.383)	(0.425)
Observations	238	238	238	238
Population	NO	YES	YES	YES
Individual Controls	NO	NO	YES	YES
MSA Controls	NO	NO	NO	YES
F-Stat (Excl. Instr.)	9.133	9.438	6.496	5.913

Notes: The dependent variable is the percentage point change in married women labor force participation between 2000 and 1980. Δ Commuting is the change in commuting time between 1980 and 2000. 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 3. Individual Controls are collapsed averages from the same individual controls: Share of college education, mean age and share of white people. Estimations are based on the cohesion proximity index. F-test 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 – 2SLS (Cohesion index).

	(1)	(2)	(3)	(4)
	2006-2011	2000	1990	1980
Commuting	-0.502**	-0.388***	-0.387***	-0.208
	(0.251)	(0.118)	(0.128)	(0.181)
Observations	1,394,373	1,382,904	1,156,111	1,075,956
F-test	4.115	9.826	9.567	12.372

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 for white male workers, divided by 100. The sample is restricted to married women between 25 and 55 years old. All regressions include log population and its squared term as controls; individual controls are: age, education level and race dummies and the log of spouse income; MSA controls: the share of employment in the manufacturing sector, the share of employment in the public sector, the MSA median income and gender wage gap and regional dummies. Estimations are based on the normalized Cohesion index. F-test 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

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2016/1, Galletta, S.: "Law enforcement, municipal budgets and spillover effects: evidence from a quasi-experiment in Italy"
2016/2, Flatley, L.; Giulietti, M.; Grossi, L.; Trujillo-Baute, E.; Waterson, M.: "Analysing the potential economic value of energy storage"
2016/3, Calero, J.; Murillo Huertas, I.P.; Raymond Bara, J.L.: "Education, age and skills: an analysis using the PIAAC survey"
2016/4, Costa-Campi, M.T.; Daví-Arderius, D.; Trujillo-Baute, E.: "The economic impact of electricity losses"
2016/5, Falck, O.; Heimisch, A.; Wiederhold, S.: "Returns to ICT skills"
2016/6, Halmenschlager, C.; Mantovani, A.: "On the private and social desirability of mixed bundling in complementary markets with cost savings"
2016/7, Choi, A.; Gil, M.; Mediavilla, M.; Valbuena, J.: "Double toil and trouble: grade retention and academic performance"
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2016/9, Guio, J.; Choi, A.; Escardíbul, J.O.: "Labor markets, academic performance and the risk of school dropout: evidence for Spain"
2016/10, Bianchini, S.; Pellegrino, G.; Tamagni, F.: "Innovation strategies and firm growth"
2016/11, Jofre-Monseny, J.; Silva, J.I.; Vázquez-Grenno, J.: "Local labor market effects of public employment"
2016/12, Sanchez-Vidal, M.: "Small shops for sale! The effects of big-box openings on grocery stores"
2016/13, Costa-Campi, M.T.; García-Quevedo, J.; Martínez-Ros, E.: "What are the determinants of investment in environmental R&D?"
2016/14, **García-López, M.A; Hémet, C.; Viladecans-Marsal, E.:** "Next train to the polycentric city: The effect of railroads on subcenter formation"
2016/15, Matas, A.; Raymond, J.L.; Dominguez, A.: "Changes in fuel economy: An analysis of the Spanish car market"
2016/16, Leme, A.; Escardíbul, J.O.: "The effect of a specialized versus a general upper secondary school curriculum on students' performance and inequality. A difference-in-differences cross country comparison"
2016/17, **Scandurra, R.I.; Calero, J**.: "Modelling adult skills in OECD countries"
2016/18, **Fernández-Gutiérrez, M.; Calero, J**.: "Leisure and education: insights from a time-use analysis"
2016/19, Del Rio, P.; Mir-Artigues, P.; Trujillo-Baute, E.: "Analysing the impact of renewable energy regulation on retail electricity prices"
2016/20, Taltavull de la Paz, P.; Juárez, F.; Monllor, P.: "Fuel Poverty: Evidence from housing perspective"
2016/21, Ferraresi, M.; Galmarini, U.; Rizzo, L.; Zanardi, A.: "Switch towards tax centralization in Italy: A wake up for the local political budget cycle"
2016/22, Ferraresi, M.; Migali, G.; Nordi, F.; Rizzo, L.: "Spatial interaction in local expenditures among Italian municipalities: evidence from Italy 2001-2011"
2016/23, Daví-Arderius, D.; Sanin, M.E.; Trujillo-Baute, E.: "CO2 content of electricity losses"
2016/24, Arqué-Castells, P.; Viladecans-Marsal, E.: "Banking the unbanked: Evidence from the Spanish banking expansion plan"
2016/25 Choi, Á.; Gil, M.; Mediavilla, M.; Valbuena, J.: "The evolution of educational inequalities in Spain: Dynamic evidence from repeated cross-sections"
2016/26, Brutti, Z.: "Cities drifting apart: Heterogeneous outcomes of decentralizing public education"
2016/27, Backus, P.; Cubel, M.; Guid, M.; Sánchez-Pages, S.; Lopez Manas, E.: "Gender, competition and performance: evidence from real tournaments"
2016/28, Costa-Campi, M.T.; Duch-Brown, N.; García-Quevedo, J.: "Innovation strategies of energy firms"
2016/29, Daniele, G.; Dipoppa, G.: "Mafia, elections and violence against politicians"
2016/30, Di Cosmo, V.; Malaguzzi Valeri, L.: "Wind, storage, interconnection and the cost of electricity"

2017

2017/1, González Pampillón, N.; Jofre-Monseny, J.; Viladecans-Marsal, E.: "Can urban renewal policies reverse neighborhood ethnic dynamics?"
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2018

2018/1, Boadway, R.; Pestieau, P.: "The tenuous case for an annual wealth tax"
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2018/3, Daniele, G.; Galletta, S.; Geys, B.: "Abandon ship? Party brands and politicians' responses to a political scandal"
2018/4, Cavalcanti, F.; Daniele, G.; Galletta, S.: "Popularity shocks and political selection"
2018/5, Naval, J.; Silva, J. I.; Vázquez-Grenno, J.: "Employment effects of on-the-job human capital acquisition"
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2018/10, Grossi, L.; Nan, F.: "The influence of renewables on electricity price forecasting: a robust approach"
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2018/12, van den Bergh, J. C.J.M.; Angelsen, A.; Baranzini, A.; Botzen, W.J. W.; Carattini, S.; Drews, S.; Dunlop, T.; Galbraith, E.; Gsottbauer, E.; Howarth, R. B.; Padilla, E.; Roca, J.; Schmidt, R.: "Parallel tracks towards a global treaty on carbon pricing"
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2018/14, Firmino, J.: "Class composition effects and school welfare: evidence from Portugal using panel data"
2018/15, Durán-Cabré, J. M.; Esteller-Moré, A.; Mas-Montserrat, M.; Salvadori, L.: "La brecha fiscal: estudio y aplicación a los impuestos sobre la riqueza"
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2018/22, Rice, C.; Vall Castelló, J.: "Hit where it hurts – healthcare access and intimate partner violence"
2018/23, Ramos, R.; Sanromá, E.; Simón, H.: "Wage differentials by bargaining regime in Spain (2002-2014). An analysis using matched employer-employee data"

2019

2019/1, Mediavilla, M.; Mancebón, M. J.; Gómez-Sancho, J. M.; Pires Jiménez, L.: "Bilingual education and school choice: a case study of public secondary schools in the Spanish region of Madrid"
2019/2, Brutti, Z.; Montolio, D.: "Preventing criminal minds: early education access and adult offending behavior"
2019/3, Montalvo, J. G.; Piolatto, A.; Raya, J.: "Transaction-tax evasion in the housing market"
2019/4, Durán-Cabré, J.M.; Esteller-Moré, A.; Mas-Montserrat, M.: "Behavioural responses to the re)introduction of wealth taxes. Evidence from Spain"
2019/5, Garcia-López, M.A.; Jofre-Monseny, J.; Martínez Mazza, R.; Segú, M.: "Do short-term rental platforms affect housing markets? Evidence from Airbnb in Barcelona"
2019/6, Domínguez, M.; Montolio, D.: "Bolstering community ties as a means of reducing crime"
2019/7, García-Quevedo, J.; Massa-Camps, X.: "Why firms invest (or not) in energy efficiency? A review of the econometric evidence"
2019/8, Gómez-Fernández, N.; Mediavilla, M.: "What are the factors that influence the use of ICT in the classroom by teachers? Evidence from a census survey in Madrid"
2019/9, Arribas-Bel, D.; Garcia-López, M.A.; Viladecans-Marsal, E.: "The long-run redistributive power of the net wealth tax"
2019/10, Arribas-Bel, D.; Garcia-López, M.A.; Viladecans-Marsal, E.: "Building(s and) cities: delineating urban areas with a machine learning algorithm"
2019/11, Bordignon, M.; Gamalerio, M.; Slerca, E.; Turati, G.: "Stop invasion! The electoral tipping point in anti-immigrant voting"

2020

2020/1, Daniele, G.; Piolatto, A.; Sas, W.: "Does the winner take it all? Redistributive policies and political extremism"
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