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**Why Do So Few Women Work in New York  
(And So Many in Minneapolis)?  
Labor Supply of Married Women across U.S. Cities**

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# WHY DO SO FEW WOMEN WORK IN NEW YORK (AND SO MANY IN MINNEAPOLIS)? LABOR SUPPLY OF MARRIED WOMEN ACROSS U.S. CITIES

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**ABSTRACT.** This paper documents a little-noticed feature of U.S. labor markets—very large variation in the labor supply of married women across cities. We focus on cross-city differences in commuting times as a potential explanation for this variation. We start with a model in which commuting times introduce non-convexities into the budget set. Empirical evidence is consistent with the model’s predictions: Labor force participation rates of married women are negatively correlated with the metropolitan area commuting time. Also, metropolitan areas with larger increases in average commuting time in 1980-2000 had slower growth in the labor force participation of married women.

JEL: J21, J22, R23, R41.

Keywords: female labor supply, local labor markets, commute time, non-convex budget sets

## INTRODUCTION

Women’s labor supply has, for good reason, been the object of extensive empirical study. After all, the dramatic rise in female labor force participation that occurred over the past 60 years in the U.S. (and in many other countries) has been the most visible and important shift in the labor market. Also, women’s labor supply is often the margin of adjustment in households’ responses to policy shifts, e.g., changes in the taxation of household income or welfare entitlement programs, and thus holds the key to proper policy evaluation.

Although many empirical studies of female labor supply have been conducted, it appears that an interesting, potentially important feature of the U.S. markets has gone largely unnoticed: There is wide variation in female labor supply across metropolitan areas in the United States. Consider, for example, one large group of women: married non-Hispanic white

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women aged 25 to 55 with a high school level education (found in the 2000 U.S. Census). In Minneapolis 79 percent of such women were employed, while in New York the proportion is only 52 percent.

The cross-city variation in female labor supply within the U.S. that we document in this paper is as large as the well-known and widely studied variation across OECD countries in female employment rates. For instance, by way of rough comparison, one might look at employment rates among women with “upper secondary education” from selected OECD countries: United Kingdom, 80 percent; Sweden, 78 percent; Netherlands, 74 percent; France, 71 percent; Canada, 69 percent; U.S., 66 percent; Italy, 64 percent; Japan, 59 percent; and Germany, 52 percent. (These statistics, for women aged 25 to 64, are from OECD, 2007.) In an effort to make sense of international comparisons, analysts typically focus on policy differences across countries (in paid parental leave, marginal taxes, employment protection, welfare benefits, etc.).<sup>1</sup> Such policy differences, of course, are much smaller across locations in the U.S. than OECD countries. The cross-city variation in female labor supply in the U.S. is apparently generated by characteristics of the local markets themselves.

Furthermore, while the labor supply of women has increased substantially in all cities in the U.S. over the past 60 years, there have been big differences in these cities in the timing and magnitude of the increase. Figure 1 illustrates, for 1940 through 2000, the well-known large increases in the labor supply of married non-Hispanic white women generally, and shows also how different the paths are for two particular urban locations, New York and Minneapolis. In 1940 the labor supply among women was lower in Minneapolis than in New York, but the subsequent growth in female labor supply was much more rapid in Minneapolis than in New York, leading to the large disparities observed in 2000. These results are especially interesting in light of the ongoing discussion about the possibility that the U.S. labor market has now achieved a “natural rate” of female labor force participation.<sup>2</sup>

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<sup>1</sup>Ruhm (1998), for example, focuses on the impact of paid parental leave policy on female labor supply in nine European countries. More generally, a large literature compares labor policy differences in the U.S., Canada, and Europe to explain differences in labor market outcomes. Nickell (1997), Card et al. (1999), Freeman and Schettkat (2001), and Alesina et al. (2005) are just a few examples.

<sup>2</sup>Many authors have documented the fact that female labor force participation slowed considerably in the mid-1990’s, and leveled off in the 2000s, e.g., Blau et al. (2002), Blau and Kahn (2000), and Juhn and Potter (2006). Goldin (2006) points to the importance of considering different age groups separately (rather than simply looking at aggregate measures of female labor supply), noting that for some groups of women “a plateau ... was reached a decade and a half ago.” Looking at variation across local labor markets brings an additional dimension of complexity. Should we expect that cities with low rates of labor force participation will continue to experience an increase in female labor supply until they reach the national average, or are there reasons to expect that some markets have a lower “equilibrium” participation rate than others?

The goals of this paper are to carefully document the cross-city variation in married women's labor supply across U.S. labor markets, to explore potential economic explanations for observed cross-city variation in married women's labor supply, and to examine the implications for the study of female labor supply generally. We believe that many factors are at play in producing the large observed local variation in female labor supply across the U.S., but, we argue, one explanation stands out: Married women, particularly married women with young children, are very sensitive to commuting times when making labor force participation decisions.

In building our argument about the importance of commuting cost, we start with the theory of labor supply when there is a fixed cost of participation (i.e., commuting time). The introduction of a fixed cost of participation introduces non-convexities into the budget constraint. This complication is easily handled in a one-period case for a one-person household: Assuming leisure and consumption are normal, and assuming also that initially the individual is at an "interior solution," an increase in the fixed cost reduces both leisure and labor supplied, up to a threshold at which the individual moves to a "corner solution" of supplying zero labor. Matters are more interesting in a model in which a two-person household takes a "collective" approach to labor supply. In this case, increases in the commute time can induce one partner (traditionally the wife) to move out of the labor force while inducing the other partner (the husband) to work longer hours.

As mentioned above, there are many studies of women's labor supply. Blundell et al. (2007) and Blundell and MaCurdy (1999) provide valuable discussions of key issues in this literature, and Killingsworth and Heckman (1986) overview earlier results. Most studies use national data, with results aggregated at the national level, and no attention is given to the possibility of meaningful local variation. A small body of work in economic geography does provide some evidence about cross-location variation in labor supply (e.g., Odland and Ellis (1998) and Ward and Dale (1992)), but this work does not seek to provide an explanation for the observed variation. In particular, we know of no work that posits the importance of fixed commuting costs for explaining local labor supply and then evaluates predictions empirically. We carry out such an analysis in five additional sections:

Section 1 provides the basic facts about the city-specific employment rates of non-Hispanic white married women in 50 large U.S. metropolitan areas from 1940 through 2000 using Public Use Samples of the U.S. Census. We document significant variation across cities in

current levels of women's employment, and also substantial variation across cities in the magnitude and timing of the increase in female labor supplied over the past 60 years.

Section 2 is a discussion of economic forces that might serve as potential explanations for the observed cross-cities variation in women's labor supply. We argue that the variation in observed employment rates are unlikely to be due primarily to differences across cities in labor *demand*.

Section 3 contains the primary economic contribution of this study. We develop an argument about the effect of cross-city differences in commuting times (owing, for example, to differences in congestion across cities) for labor force participation. Our model allows us to examine the effect of commuting time on individuals' and households' labor supply.

Section 4 presents empirical evidence concerning the predictions of the model. The cross-sectional evidence indicates that in cities with longer average commuting times, female labor force participation rates are lower. Women with young children are particularly sensitive to longer commute. We try a simple IV strategy (using location of birth as an instrument) in an effort to deal with potential endogeneity of location, and find that this does not change our key empirical findings. Also, results are similar for a specification that looks at differences over time (from 1980 to 1990 and 1990 to 2000). These results are all consistent with the theory presented in Section 3.

Finally, Section 5 provides a conclusion and discussion of directions for future research.

## 1. DIFFERENCES IN LABOR SUPPLY ACROSS LABOR MARKETS

This study focuses on the labor supply of married non-Hispanic white women who live in the 50 largest Metropolitan Statistical Areas (MSAs) in the United States. The focus on married women is motivated by the fact that these women are most responsible for the large changes in female labor supply that have been observed over the past several decades (see, e.g., Juhn and Potter (2006)). Women in racial and ethnic minorities are excluded to avoid the complications of dealing with additional dimensions to the analysis, and because sample sizes for these groups are much smaller. The sample is restricted to individuals aged 25 to 55.

For much of our analysis we rely on the 2000 Census 5 percent Public Use Micro Sample (PUMS).<sup>3</sup> We also exploit comparable data from 1940 through 1990 (except 1960, owing to the lack of MSA identifiers for that year).<sup>4</sup>

The PUMS data provide information on employment status. The three main categories are employed, unemployed, and not in the labor force. For the most part, the analysis below looks at the “employment rate” as the measure of labor force participation in a local labor market, thereby including women who are reported as unemployed with those who are out of labor force.<sup>5</sup> Included in the sample are women in the armed forces; they constitute 0.1 percent of the sample in 2000.

We begin by estimating participation rates for each of the 50 MSAs, for selected years, 1940 through 2000. Results are given for women in the largest educational category—women with a high school diploma. The results are sorted by participation rates in 2000, from lowest to highest. The variation evident in the statistics is striking. In 2000 the participation rates of high school educated women vary from just 52 percent in New York City to 79 percent in Minneapolis. Similarly wide variation is evident in other years as well; for instance, in 1970 MSA-specific participation rates varied from 30 percent to 59 percent. There are also substantial differences in the growth of married women’s labor supply over time. For example, from 1940 through 2000, the participation rate increased by 36 percentage points in New York but by 64 percentage points in Minneapolis (as we have seen in Figure 1).

We examined the extent to which cross-city differences in the age distribution of women account for the observed labor force participation rates and found that this matters very little.<sup>6</sup> We also constructed tables similar to Table 1 separately for women without children, women with older children, and women with young children. The summary is reported in Table 2. In each case, we found big differences across cities in labor force participation.

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<sup>3</sup>Individuals with imputed values are excluded from the analysis. The resulting sample is quite large, 423,300 observations.

<sup>4</sup>For details of the data see Ruggles et al. (2010).

<sup>5</sup>The PUMS defines unemployed persons as those without a job and looking for a job. It is difficult to know the extent to which a married woman might indicate that she is looking for a job if she intends to return to work at some point, but is not actively currently seeking a job. In any event, none of the conclusions below are altered if unemployed women are instead included in the labor force. (Unemployed women are only 1.5 percent of the sample; the average unemployment rate is only 2 percent.)

<sup>6</sup>In particular, we try “standardizing” using the national age distribution. First, the age distribution  $f(a)$  is calculated for all the women in a sample. Then in each MSA  $j$ , for each age  $a$  the employment rate  $E_j(a)$  is found. Finally, a standardized participation is obtained for MSA  $j$  rate:  $E_j = \sum_{a=25}^{55} E_j(a)f(a)$ . This makes virtually no difference for Table 1.

For example, in 2000, among high school graduate women with children younger than 5, participation rates varied from 29 percent (New York) to 68 percent (Minneapolis).

The data indicate comparably wide variation in an alternative measure of labor supply—annual work hours, which can be computed as a product of individuals’ reported “weeks worked last year” and “usual weekly hours.” For instance, in 2000 this measure of labor supply varied from 862 hours worked per year (New York) to 1,456 hours (Minneapolis).

Similarly, we repeated the analysis for a second large educational group, women with college degrees. Participation rates of college educated women are generally higher than for women with high school diplomas, but we found significant cross-MSA differences in labor force participation for these women as well. For example, participation rates of college educated women with children range from 65 percent (Honolulu) to 86 percent (Albany).

## 2. POSSIBLE EXPLANATIONS

It is likely that many factors are at play in producing the large variation across cities in observed married women’s labor force participation. However, none of the most obvious explanations seems to be key. In particular, we focus initially on three factors that intuitively might influence participation: local housing prices, local wages, and the local unemployment rate (which might be an indicator of local *demand* for labor).

To set our discussion, we examine MSA-level regressions, separately for high school- and college-educated married women, in which local labor force participation is the dependent variable. As independent variables we have a measure of local housing costs (based on a quality-adjusted housing price index developed by Chen and Rosenthal (2008)), local wages (which we calculate using wages of *single* women, i.e., women who are not in the analyses), and the local unemployment rate for men.<sup>7</sup> Results are reported in Table 3.<sup>8</sup>

One might expect that married women are more likely to work in relatively expensive cities, if only because their income is crucial to pay for high housing costs.<sup>9</sup> Our regressions do not indicate that this is the case. Moreover, we see in the first column in Table 1 that expensive cities such as New York, Los Angeles, and San Francisco are actually close to the

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<sup>7</sup>We thank Stuart Rosenthal for providing the data.

<sup>8</sup>For all regressions, we also used the ACCRA city cost of living index as an alternative to the Chen-Rosenthal housing price index, using the 46 MSAs for which the ACCRA index is available. Key results were virtually unchanged. For a description of how ACCRA (American Chamber of Commerce Research Association) cost of living index is constructed see [www.coli.org](http://www.coli.org).

<sup>9</sup>For instance, Fortin (1995), using data from the 1986 Canadian Family Expenditures Survey, shows that labor supply of some married women is affected by mortgage commitments.

top of the list as cities with *low* participation; there is a negative correlation between female participation rate and our housing price index ( $-0.42$ ).<sup>10</sup>

We might also expect that women's labor decisions are influenced by local wages. In fact, in our regression the local wage rate for women—as indicated by the local wage of single women (who generally have high labor force participation)—does not appear to have a large impact on local labor force participation.<sup>11</sup>

As we have emphasized, the cross-MSA variation is very large. Differences across cities are also quite persistent across decades. For example, the correlation between the 1990 MSA-level participation rates and 2000 MSA-level participation rates is 0.82, and comparable correlation coefficients are 0.82 for 1980–1990, and 0.81 for 1970–1980. These large and persistent differences are likely not due to local demand shocks. Having said that, transitory local shocks surely do matter, and in our regressions we include the local unemployment rate for men as a way of examining this possibility.<sup>12</sup> As expected, married women's labor force participation is lower when male unemployment is relatively high. It is helpful to keep this feature of the markets in mind when conducting explorations of other potential explanations.

Yet another contributing factor to variation in local female labor supply is the cost of child care.<sup>13</sup> We make no attempts here to directly evaluate cross-city differences in child-care costs on the labor supply of mothers with young children. But there are good reasons to believe that child-care costs are at best an incomplete part of the story. First, participation rates and average hours worked vary greatly across cities even among married women *without* children—women who presumably are little affected by differences in child-care costs.

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<sup>10</sup>Johnson (2009) finds that housing prices are positively correlated with female participation rate in a larger sample of over 200 MSAs, but not for the 50 largest MSAs that we consider.

<sup>11</sup>If instead we simply look at the relationship between observed wages for *married* women (among women who work) and the participation rate, we find a correlation coefficient of  $-0.46$  (significant at the 0.01 level). That is, higher wages are associated with *lower* participation rates. Of course, given the selection process of women into the labor market, interpretation is difficult. For instance, this negative correlation would be expected if women who work are disproportionately drawn from the high end of the talent pool.

<sup>12</sup>Women's unemployment rates are mechanically related to participation rates as defined in this study, so as an alternative we use unemployment rates constructed for men as a measure of local labor demand. In particular, we compute unemployment rates of white men with high school diploma or above aged 25 to 55 years old in each of the 50 MSAs (using the usual definition, i.e., the proportion of people in the labor force who are unemployed). To account for differences in demographic distributions across cities, the unemployment rates are standardized using the national distribution of age and education. These standardized unemployment rates for men in the 50 MSAs vary between 1.2 percent and 4.4 percent in 2000.

<sup>13</sup>The evidence on the magnitude of labor supply elasticity with respect to child care costs is mixed and the range of elasticity estimates is wide. (See Blau (2003) for a review of this literature.) Baker et al. (2005) analyze the impact of Quebec's subsidized child care program on labor supply of mothers, finding a significant but relatively small impact.



Second, across the 50 cities in the study, labor supply measures for women *with* children and women *without* children are positively correlated; the correlation is 0.80 (significant at the 0.01 level) for women with a high school diploma. It appears that the same city-specific factors affect both married women with children and without children.

A study of “power couples” by Costa and Kahn (2000) potentially offers some clues about cross-MSA variation in married women’s labor supply. In that paper, college educated couples are shown to be concentrated heavily in large metropolitan areas. If college educated women disproportionately locate in large cities, the average participation rate in these cities would be higher than in smaller cities. This argument, however, does not help explain the large differences in female labor supply that exist even among large cities. Nor is the argument helpful in understanding why labor force participation varies so widely across cities among women with a high school level of education.

Acemoglu et al. (2004) document that in 1950 there was substantial cross-state variation in female labor supply, stemming from military mobilization for World War II, and work by Fernández et al. (2004) indicates that there could be hysteretic effects. In particular, they find that wives of men whose mothers worked are more likely to work. Thus, local history can matter. With that in mind we note that the correlation in MSA-level differences in women’s labor force participation, i.e., correlations in columns of Table 1, are as follows: 2000 outcomes have a 0.86 correlation with 1990 outcomes, a 0.49 correlation with 1980 outcomes, a 0.40 correlation with 1970 outcomes, and a 0.11 correlation with 1950 outcomes. Recent patterns are highly correlated with each other, but are not highly correlated with patterns observed after World War II.

In this study we turn to a simple economic explanation that differs from initially appealing ideas mentioned above; we look at the role of commuting costs that vary across cities.

### 3. A MODEL OF LABOR SUPPLY WITH COMMUTING TIME

Among the important systematic ways in which cities differ are traffic patterns, congestion, and the resulting length of the commute to work. Commuting time can be viewed as a fixed time cost and/or monetary cost of going to work. Oi (1976), in a classic paper, introduces the idea that commuting time might play an important role in family labor supply decisions, including the joint decision about where to live. Cogan (1981) presents an analysis of labor supply with fixed money and time costs of labor market entry, arguing that fixed costs “are of prime importance in determining the labor supply behavior of married women.” The general

idea is developed in the urban economics literature, motivating a small empirical literature. For example, Gordon et al. (1989) show that women have shorter commute times than men regardless of income, occupation, or marital status.

Existing theoretical work does not focus on the role of fixed costs in joint household labor supply, nor does it focus on the dynamic setting. As we show below, these issues may be crucial for understanding the role of fixed costs in labor force participation and hours worked. Also, we know of no research that uses cross-city variation in commuting times as a way of learning about the potential importance of fixed costs in labor supply decisions.

As a starting point to motivate our empirical work, we present a simple theory of fixed cost and labor supply. The analysis proceeds in two parts. First, we consider a one-person one-period case. We then turn to a formulation in which labor is supplied by a two-person household in a “collective model” of intrafamilial decisions along the lines of Chiappori (1988) and Chiappori (1992).<sup>14</sup>

**3.1. A Static One-Person Model.** Consider the standard labor supply model with the addition of a commuting cost, which we assume here is a time cost. Let  $c > 0$  denote the time cost of commuting incurred whenever hours of work are positive. In a one-period one-person model, we can write the budget constraint and time constraint, respectively, as

$$(1) \quad F = N + wT = pX + wL + wIc, \quad \text{and}$$

$$(2) \quad T \geq L + Ic,$$

where  $T$  is the endowment of time,  $N$  is non-labor income,  $p$  is the price of the consumption good  $X$ ,  $w$ , the wage, is the implicit price of leisure  $L$  or commuting,  $I$  is an indicator function equal to 1 when  $L < T$  and equal to 0 when  $L = T$ , and then  $F$  is full income.

For the specification of preferences, we will assume that the agent has a twice differentiable, strictly concave utility function in which both the consumption good and the leisure are normal. (The assumption of strict concavity provides the agent with an incentive to smooth her consumption over time when we move to a multiple-period model.) In such a world, the objective of the agent is to maximize utility,  $u(X, L)$ , subject to the constraint that leisure may never be larger than the time endowment and the budget constraint, which, from (1)

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<sup>14</sup>An alternative interpretation, which might be of independent interest concerns the impact of commuting time on lifetime labor supply in a dynamic setting with 2 periods (which generalizes easily to  $n$  periods).

and (2), can be written

$$(3) \quad N + w(T - L - Ic) = pX.$$

From a technical standpoint, the only difficulty is that the indicator function  $I$  makes the budget constraint nonconvex.

To solve this problem, for any prices and nonlabor income, we can simply solve the problem under two regimes: (i) the agent pays commuting costs and is free to work or not and (ii) the agent does not pay the commuting cost and cannot work. Each of the two regimes provides us with a standard convex budget set, which in turn allows us to apply the theorems of concave optimization. The agent then selects the regime that provides the higher utility.

The problem facing the agent when she does not commute (and therefore cannot work) is simple: She merely sets consumption equal to  $X = N/p$ . In solving the problem when the agent *does* commute, necessary conditions are

$$(4) \quad u_L(X, L) \geq \lambda w,$$

$$(5) \quad u_X(X, L) = \lambda p, \quad \text{and}$$

$$(6) \quad N + w(T - L - Ic) = pX.$$

As we have said, the agent solves each of these problems and chooses the outcome that yields higher utility.

We depict the indirect utility function that results from each of the two optimizations problems in Figure 2, which shows the relationship between the indirect utility functions and full income  $F$ . Because leisure is a normal good, we know that the *no commute/no work* indirect utility,  $V_{nc}$ , must cut the *commute/work* indirect utility,  $V_c$  from below. Let  $F^*$  depict full income such that the agent is indifferent between working and not.<sup>15</sup> To the right of  $F^*$ , the agent chooses not to work, and to the left of  $F^*$  the agent works. To develop some intuition for the agent's choices, consider behavior at  $F = F^*$ . If the agent works, she consumes more but pays both the cost of commuting and the cost of foregone leisure due to work. If she chooses not to work, she accrues additional leisure and does not need to pay the commuting costs, but she must reduce her consumption.

Comparative statics are easy in this set up. Consider, for example, an increase in commuting costs  $c$ . The impact on the indirect functions is depicted in Figure 3, which shows the

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<sup>15</sup>We are looking for a non-trivial solution, so we do not consider a case when  $V_{nc} > V_c$  for all positive values of  $F$ .

comparative statics for an increase in commuting cost to  $c' > c$ . If the agent is initially not commuting (i.e., if her income initially exceeds  $F^*$ ), the increase, of course, has no impact on her utility. If she initially is working, the utility of working must decline. The critical value of full income is now smaller. An increase in commuting cost, of course, never induces a non-working agent to join the labor force, but it can induce her to withdraw from the labor force.

We have one final observation in the static model. Conditional on staying employed, individuals facing an increased commuting cost will work fewer hours. Figure 4 illustrates the logic: When commuting cost is  $c_0$ , the choice is  $(L^*(c_0), X^*(c_0))$ . After an increase in  $c$ , the new budget constraint is  $BC_1$ . Since leisure and consumption are both normal, the new optimal point must lie on a segment of  $BC_1$  between points  $y$  and  $z$ . Thus the decrease in leisure  $L$  is less than  $x - y = \Delta c$ . Since the decrease in leisure is less than the increase in commuting time, work hours decline.

**3.2. Labor Supply with a Two-Person Problem.** The one-person model provides insight, but it misses a key feature of an intertemporal or two-person problem. When working more than one period, an agent can move financial resources between periods to smooth consumption. The same is true of Chiappori's collective household model; as Chiappori (1992) shows (see Proposition 1), the collective approach can formally be modeled as a case in which each person maximizes own utility subject to a transfer between individuals (endogenously determined). Here we set up a case with two people, which is formally the same as an individual maximizing over two periods. To simplify notation, we suppose that both individuals in the household face the same wage.

With two individuals, total income becomes

$$\sum_{i=1}^2 (wT + N_i) = F_1 + F_2.$$

where  $F_i$  is full income for individual  $i$ . The non-convex budget set is now

$$(7) \quad F_1 + F_2 = \sum_{i=1}^2 (wT + N_i) = \sum_{i=1}^2 (pX_i + wL_i + wL_i c),$$

and the objective function is

$$(8) \quad U = \sum_{i=1}^2 u(X_i, L_i),$$

which assumes a symmetric collective model of household maximization.<sup>16</sup>

Our approach to the utility maximization problem is to exploit separability of the utility function and make this a “two-stage budgeting problem.” Let  $y_i$  be a (possibly negative) transfer from one person to another in the household.<sup>17</sup> In the first stage, conditional on the price, the wage, and a given allocation of income, the household chooses the optimal consumption bundle in the same manner as in a one-person problem. Let  $V_c(F_i, w, p)$  denote the indirect utility function if the commuting cost is incurred and  $V_{nc}(F_i, p)$  denote the indirect utility function when there is no commuting cost, where  $F_i \equiv y_i + N_i + wT$ . Now define

$$(9) \quad V^*(F, w, p) = \max\{V_c(F, w, p), V_{nc}(F, p)\}.$$

The function  $V^*(F, w, p)$  is continuous in its variables, but, of course, is not differentiable at the crossing point.

For the second-stage, the household’s problem is:

$$(10) \quad \max_{y_i} \left\{ \sum_{i=1}^2 V^*(F_i, w, p) \right\}$$

subject to the budget constraint

$$\sum_{i=1}^2 F_i = \sum_{i=1}^2 (wT + N_i + y_i) = \sum_{i=1}^2 (wT + N_i).$$

The utility function is continuous in  $F_i$  and the budget set is convex so the function has an optimum. There may not be a unique optimum because the household may be indifferent between having the husband work full time and the wife stay at home or the wife work full time and the husband stay at home. We follow a convention of assuming the wife is the one staying at home if one person does not work.

While the explicit derivation of the optimal conditions is tedious, armed with our insights from the static model this is a simple problem to solve. If a household’s nonlabor income is sufficiently high, neither individual in the household will work. Conversely, if nonlabor income is sufficiently low, both may choose to work. In each of these cases, the analysis parallels the static case.

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<sup>16</sup>A generalization would place differing weights on  $u$ , but this does not alter any of the analysis that follows, and complicates notation.

<sup>17</sup>Of course,  $y_1 = -y_2$ .

More interesting, however, is the case where the household finds it optimal to have one person work but not the other. To see why, consider Figure 5. In Figure 5 we see that the two indirect functions create a nonconcave objective function, and this in turn has an important impact on participation. Consider a household with full income  $F^*$  for each individual, such that  $V_c(F^*, w, p) = V_{nc}(F^*, p)$ . If these individuals were on their own, each would be indifferent between working or not working. However, given full income  $2F^*$  in the *two-person* model, one person would work while the other would not. The ability of individuals within the household to reallocate resources between themselves essentially “convexifies” the budget set. Note that our argument holds for households with full income that differ modestly from  $2F^*$ ; a household with joint full income sufficiently close to  $2F^*$  will optimally have one person work while the other one does not.

Comparative statics in the two-person model are not particularly difficult. As in the static model, an increase in  $c$  can induce individuals to withdraw from the labor market, e.g., in this case it could cause households with two earners to become a household with one earner.

In contrast to the one-person model, in the two-person model an increase in the commuting cost can result in the husband supplying *more* hours in a period in which he works. To see how this occurs, consider Figure 6. In Figure 6, the household initially has just enough full income  $F^*$  to be indifferent between both individuals working or one individual working.<sup>18</sup> Now if there is a small increase in the commute cost, the household unambiguously prefers to have one person working. Relative to the outcome with both are working, when they move to having one person only work, full income declines. Given that leisure is normal, the husband will consume less leisure, i.e., work longer hours. ( Since the consumption good is also normal, consumption of course declines as well.) As Heim and Meyer (2004) note, the presence of nonconvexities in the budget constraint may result in the agent exhibiting discontinuous behaviors similar to an agent with nonconvex preferences (but conventional convex budget sets).

To summarize, in both a one-person and two-person model we have an unambiguous prediction about the impact of commuting cost  $c$  on labor force participation: An increase in  $c$  can never cause an increase in labor force participation, but it can cause a decrease. For example, with a two-person household an increase in commute time can induce one individual to withdraw from the labor market.

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<sup>18</sup>They can work both periods, and each receive  $V_c(F^*, w, p)$ . Or one can work, and have full income  $F^* - \delta$ , while the other does not, but has full income  $F^* + \delta$ .

In a cross section of cities in which traffic congestion and commuting times vary, we therefore expect, all else equal, to see lower labor force participation rates in cities with high commuting times. In our model, we cannot predict with assurance whether the husband or wife will withdraw from the labor force. But as we have noted, as an empirical matter it is more likely to be the wife. Also, if the woman has a comparative advantage in such home-related activities as bearing children or caring for an infant child, this reinforces the likelihood that the woman stays home in a household that has one earner. The economic logic is simple: congestion can increase commute times and this increases the value of being at home relative to the value of working. This effect is likely to be especially pronounced for women with young children. We therefore expect labor supply of these women to be especially sensitive to changes in commuting time.<sup>19</sup>

With these ideas in mind, we turn to an empirical investigation of the relationship between commuting times across U.S. cities and labor supplied.

#### 4. AN EMPIRICAL ANALYSIS OF LABOR SUPPLY AND COMMUTING TIME

**4.1. MSA-Level Labor Force Participation and Commuting Time.** To begin the empirical analysis, average commuting time is computed for each of the 50 MSAs. The Census asks respondents about how long it takes them to get to work; this number is multiplied by two to obtain the daily commute measure. For each MSA, average daily commuting time is estimated for working married women, women with children, and married men. These results are summarized in Table 4. Women on average have shorter commute than men, and women with children commute even less. For men, commuting to work and back takes on average at least 43.5 minutes a day (in Dayton) and can be as much as 76.1 minutes a day (in New York). For women, the daily commute varies between 38.6 and 63.4 minutes a day on average, depending on the MSA. In short, the cross-city differences in average daily commute are substantial. The cities with the longest commute times are New York, Washington, D.C., and Chicago. Dayton, Oklahoma City, and Buffalo have the shortest average commute times in the sample of 50 MSAs.

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<sup>19</sup>In terms of our model, one can think of the value of “leisure” (which includes such non-leisurely activities as child care) as being highest for women with children. A more formal way to proceed is to allow individuals to allocate time among work, commuting, leisure, and also home production. Kolesnikova (2007) follows this route, showing that individuals who have a fixed time commitment to home production (e.g., women who are committed to the care of young children at home) will indeed have labor supply decisions that are more responsive to the commuting cost.

Clearly, commuting time is important. A married man in the median MSA works approximately 9 hours per day, and typical total daily commutes easily tops 1 hour in many cities.

To examine an effect of commute times on the participation decision of women it is necessary to have a measure of the average commuting time cost in an MSA. Because many women do not work, there is a selection bias in our estimates of their average commute. Thus, women's average commute time cannot be used as such a measure. On the other hand, most white married men do work.<sup>20</sup> Their average commuting time can serve as an indicator of the difficulty of the commute in a city.<sup>21</sup> In what follows this measure is referred to as an "MSA daily commute time."

Our theory above gives an unambiguous prediction about the effects of commuting time on labor force participation, so we focus initially on the cross-MSA relationship between labor force participation by married women and commuting times.

Table 5 reports the results from an MSA-level linear regression in which the dependent variable is the MSA labor force participation rates, and average MSA daily commute time (for men) is the key explanatory variable. The analysis is performed separately for high school educated women and for women with a college degree. Results indicate that higher levels of the MSA daily commute time are generally associated with lower levels of labor force participation rates among married women. The association is stronger for women with a high school level of education. A 1-minute increase in the MSA daily commute time is associated with a 0.3 percentage point reduction in the labor force participation rate of high school educated women. We have included some of the other control variables discussed when we presented Table 3.<sup>22</sup> Among these variables, only the unemployment rate of non-Hispanic white men is found to be statistically significant at conventional levels. Failure to include this latter variable makes virtually no difference to the estimates of the impact of commuting time; there is a near zero correlation between MSA daily commute time and the MSA unemployment rate.<sup>23</sup>

<sup>20</sup>Participation of white married men aged 25 to 55 is above 90 percent in every city in our analysis.

<sup>21</sup>We also used city commute data provided by the Census. The numbers are very close to the estimates reported in Table 4 and produced similar results in our analysis.

<sup>22</sup>As with the analysis reported in Table 3, all regressions were also conducted using the ACCRA city cost of living index instead of the Chen-Rosenthal housing price index. Results were unchanged.

<sup>23</sup>As we discuss above, we do not have measures of child-care costs, which might also be relevant. Cortes and Tessada (2007) consider the possibility that in a city with a high concentration of low-skilled immigrants, who often work as maids and nannies, the cost of day care might be low, which in turn could increase the labor supply of high-skilled American women. They find no such effects except for an increase in work hours



**4.2. Individual Level Analysis of Labor Force Participation.** We next look at individual level data, examining the relationship between the MSA daily commute time and women’s labor force participation decisions. We estimated both a probit model and a linear probability model. Marginal effects were nearly identical, so we focus on the linear model, which is easier to interpret. In our analysis the dependent indicator variable, which equals 1 if the woman participates in the labor force, is multiplied by 100 (so that probability can be discussed in percentage points).

The usual problem when estimating labor supply is that wages are observed only for those who work. The approach here is to include in each regression only women with exactly the same level of education, i.e., women who are likely to have similar levels of market productivity.<sup>24</sup> As we discuss above, the relative value of women’s time spent at home is plausibly higher among women with young children than those without children, which in turn can make these women more sensitive to commute time than other women. To account for this possibility, the sample is also divided into three separate groups: women with children younger than 5 years old, women with children none of whom are younger than 5, and women with no children.

Panel A1 of Table 6 presents the first piece of individual-level empirical evidence concerning the effect of the MSA daily commute time on workforce participation decisions by married women. It indicates that an increase in MSA commute time is associated with a decrease in the probability of being in the labor force for all the groups. The effect is the largest for high school graduates with young children. For them, a 1 minute increase in the MSA commute time is associated with a decrease in the probability of labor force participation of 0.51 percentage points. For high school educated women with children but no young children, the associated decrease in probability of labor force participation is 0.31 percentage points, while for high school educated women with no children the estimated effect is 0.14 percentage points. For women with a college degree the effect of longer average commuting time on labor

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of women with graduate degrees (a group whom we do not study). Nonetheless, to satisfy curiosity, we did include their measure of low-skilled immigrants in our city-level regressions (for the 25 cities for which the Cortes and Tessada measure is available). We similarly find no evidence that low-skilled immigrants increase women’s labor supply. More importantly, our key results about the effect of commuting time were virtually unchanged.

<sup>24</sup>While there is substantial debate on the validity of education as an instrument in this context, in fact education is frequently used as an instrument in estimating labor supply. See, for example, Ham and Reilly (2002), MaCurdy (1981), Altonji (1986), Ham (1986), and Reilly (1994).

force participation decisions is somewhat smaller, as one would expect if the opportunity cost of their time at work is higher than for women with a high school level of education.<sup>25</sup>

Panel A2 repeats our empirical exercise, but now instruments for location with place of birth. The concern is that place of residence is a decision variable, and that the results in Panel A1 might simply reflect that one-earner households sort to places with long commutes. We instrument with state of birth, which is arguably exogenous, and this proves to be a powerful predictor of the city of residence in adulthood. In particular, for each state, we compute the average daily commute in 2000 of white non-Hispanic married men who are 25-55 years old. We then use average commute time in a state of birth as an instrument for MSA daily commute time for each observation.<sup>26</sup> Our key coefficients of interest remain negative and statistically significant, and if anything are larger in absolute value.<sup>27</sup>

In Panels B1 and B2 of Table 6, the model specification includes other factors that affect labor supply decisions, such as the number of children, husband's education, and controls for 5-years age intervals. In addition, an MSA unemployment rate for white men is included to control for local labor market conditions. When these factors are included, the estimated coefficients tend to be slightly smaller (in absolute value) than the ones reported for the first two panels, but the overall pattern is similar.<sup>28</sup>

**4.3. Changes in MSA Labor Force Participation, 1980 to 2000.** In this section we return to the MSA-level analysis, but now we examine data from the 1980 and 1990 Census PUMS. We compute labor force participation rates of white married women in 50 large MSAs in 1980 and 1990. The average MSA daily commuting time is also calculated using 1980 and 1990 Census PUMS data and is used as an explanatory variable.<sup>29</sup> In analyzing 2000 data, neither wage nor housing prices seem to matter much for married women's participation at the MSA level (see Table 5), so we do not include them in our regressions here.

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<sup>25</sup>Johnson (2009) includes the average MSA travel time in his model that estimates effect of house price variation across metro areas on married women's labor supply. He also finds a negative effect of longer commute on labor force participation.

<sup>26</sup>Notice that the first-stage  $F$  statistics are quite large; we do not have a problem with weak instruments.

<sup>27</sup>In all analyses standard errors are obtained using a Huber-White sandwich estimator of the variance using clustering on MSA level.

<sup>28</sup>We include husband's education as a proxy for husband's permanent income, which is the appropriate construct affecting married women's labor supply, as Mincer (1962) argues. In fact, our empirical exercise is similar in spirit to Mincer (1962), who also uses cross-city variation to examine female labor supply. His work does not consider commuting times, and in our work we are not trying to estimate labor supply elasticities.

<sup>29</sup>Unfortunately, we cannot examine earlier decades because we do not have commuting time data.

As was the case in 2000, there is a significant variation across cities in labor force participation rates of white married women in 1980 and 1990. For instance, the participation rates of high school educated women vary from 41 percent in Pittsburgh to 67 percent in Greensboro in 1980 and from 51 percent in New York City to 78 percent in Minneapolis in 1990.<sup>30</sup> As in 2000, this variation is “explained” in part by commuting times: Longer average commute times within an MSA are associated with lower levels of married women’s labor force participation. The results are reported in Table 7.<sup>31</sup>

We next ask if *changes* in commuting time over the decades under study are correlated with MSA-specific trends in women’s labor force participation.

In order to conduct a differences-in-differences analysis, we first calculate the differences in female participation rates between the Census dates for each city. When we undertake this first step, we find that the labor force participation rose about 16 percentage points on average between 1980 and 2000 for high school educated married women and about 14 percentage points for college educated women, with most of the increase occurring between 1980 and 1990. However, the rise in participation is not uniform across MSAs. From 1980 through 2000 high school educated women’s employment rose by only 5 percentage points in San Francisco, while increasing by 24 percentage points in Buffalo. During the 1990s, some MSAs experienced moderate increases in participation rates of married women while in others participation declined (e.g., it fell by 7 percentage points in Honolulu).

Our second step is to look at MSA-specific changes in the average commuting time. We find that between 1980 and 2000 the daily MSA commute increased by about 5 minutes on average.

We then examine a relationship between changes in the MSA-level commute time and changes in women’s participation rates over the same period. To control for changes in labor market demand conditions over the years, changes in unemployment rates of white men are also included (in specification (2) for each regression). Results are presented in Table 8. Panel A shows the results of the linear regression estimation for white married women with a high school diploma. Panel B present the results for women with a college degree. We find a negative relationship between the change in commute time in an MSA and the

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<sup>30</sup>The numbers reported here are averages that are standardized using the sample age distribution. They differ slightly from the results in Table 1.

<sup>31</sup>As was true of the 2000 data, including the MSA unemployment rate of white men in the regression to control for labor demand conditions does not change the effect of a longer commute on the employment rates of white women. Compare columns (1) and (2) for each group in Table 7.

change in the employment rate of married women in that MSA; cities in which commuting time increased most rapidly generally also experienced slower growth in female labor force participation.

We note a final observation. In results not reported here (but available in an earlier working paper version of the paper), we investigated the relationship between MSA commute times and married men's labor supply. We find that in cities with long commute times, men actually work slightly longer hours.<sup>32</sup> This is consistent with our model above: Commute times seem to affect the intensive margin in household labor supply, as married women withdraw from the labor force, but if anything induces marginally more labor supplied from husbands.

## 5. CONCLUDING REMARKS

Using Census Public Use Micro Sample (PUMS) data for 1940 through 2000, we find wide variation in labor market participation rates of white married women in 50 large U.S. metropolitan areas. This wide variation is found in all years and appears for women with different levels of education, as well as for women with and without children. Among a number of explanations for observed cross-city differences in female labor supply one emerges as particularly important: Married women's labor force participation decisions appear to be related to MSAs' commuting times.

While the possibility that fixed employment costs might matter for female labor supply was raised in the literature many years ago (e.g., Oi, 1976, and Cogan, 1981), we know of no work that has explored implications for the cross-city variation in labor supply. With this in mind we have undertaken an analysis that, first, sets out a simple theory of labor supply. In our theoretical exploration, we emphasized that an increase in the commuting cost  $c$  can be expected to reduce labor force participation of married women, while at the same time increasing the number of hours worked by working husbands.

Cross-section evidence is consistent with the theory. We find a negative association between commuting time and women's labor force participation rates in the three decades for which we have commuting time data (1980, 1990, and 2000). Similarly, the negative correlation between commuting times and participation appears also in a differences-in-differences

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<sup>32</sup>OLS estimates are positive and statistically significant, but IV estimates are not statistically significant.

analysis; metropolitan areas that experienced relatively large increases in average commuting times between 1980 and 2000 experienced slower growth of labor force participation of married women.

For women, the effects of commute times are quite large. From Table 5 we see that a 1-minute increase in an MSA's commute time is associated with an approximately 0.3 percentage point decline in the labor force participation of women with a high school education.<sup>33</sup> Table 4 indicates that the difference in commute times, from the shortest-commute MSA to the longest-commute MSA, is 33 minutes. Taking our estimates at face value, this variation might be expected to lead to a 10 percentage point difference in participation across cities. Commute time differences across MSAs thus plausibly "explain" a fair amount of the cross-MSA variation in participation observed in Table 1.

It might be reasonable to ask, indeed, if our estimates are "too big." One might reason that a 30-minute increase in commute time presents an opportunity cost of only \$10 daily for a woman who earns \$20/hour. Could such a modest factor plausibly have such a large impact on behavior? In thinking about that issue it is important to recognize that the cost to a typical mother of living in a high-commute MSA extends far beyond the increased time she spends commuting if she works. As we noted above, in congested cities there will be additional time required for travel to the grocery store, piano lessons, or little league baseball. If she chooses to commute a long distance to work, a mother will likely be farther from her children's school, which can make life difficult when she needs to be available for a parent-teacher conference or a class play. Our estimates are not picking up the pure effect of the commute time of the participation decision, but the total impact of congestion on the participation decision.

One implication of our work concerns the century-long increase in the female labor force participation—the increase in married women's participation from only about 7 percent in 1900 to current rates (near 70 percent). There are doubtless many factors contributing to this trend, many of which have received careful examination in the literature. Little attention has been given, though, to the possibility that part of this trend is due to the reduction in commuting costs, owing to improvements in transportation technology—the expansion of modern public transportation, the introduction and continued improvement in automotive technology, improvements in roads, and so on—and changes in residential patterns.

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<sup>33</sup>We draw a similar inference for the modal group of women, women with older children at home, in the individual-level analysis in Table 6. (See the middle columns in Panel B2.)

Our findings about the wide cross-city variation in the labor force participation of married women also introduce a new dimension to the current discussion about trends in the female labor supply. In particular, these findings complicate discussions about women having reached a “natural rate” of labor force participation. The issue is how close to 1 we can expect this participation rate to be. Goldin (2006), Juhn and Potter (2006), and others show that labor force participation rates depend on a combination of demographic factors such as age, presence of children, education, and race. The “natural rate” of participation is expected to be different for different groups. Our research suggests that the maximum achievable rate of labor force participation for each group would also vary across cities (and also across countries) because of differences in commuting time.

Of course, commuting times in local communities also depend on population density, the resources devoted to transportation, and local planning (e.g., zoning laws that may sometimes serve to isolate residential communities from job locations). Thus, from a public policy perspective, it may be that targeted actions that reduce commuting times would thereby increase labor force participation by women.<sup>34</sup>

Yet another open policy issue concerns the importance of variation in labor supply across cities for tax and welfare policy. It would be interesting to analyze how differences in the time cost of commuting affect labor supply responses to changes in such policies.

Finally, we note that most empirical research in labor economics is conducted at the national level, with little attention given to the possibility that local labor markets differ in important ways. Our research points to the value of work that allows for the possibility of differences across locations in labor supply responses. More generally, there is surely a rich set of interesting issues yet to be examined around the interactions of urban characteristics and labor market outcomes.

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<sup>34</sup>In this sense, the work here is related to the “spatial mismatch hypothesis” literature that typically looks at job accessibility as a determinant of the generally poor employment prospects of minority workers. This idea was first introduced by Kain (1968).

## REFERENCES

- Acemoglu, D., D. Autor, and D. Lyle (2004). Women, War, and Wages: The Effect of Female Labor Supply on the Wage Structure at Midcentury. *Journal of Political Economy* 112(3), 497–551.
- Alesina, A., E. Glaeser, and B. Sacerdote (2005). Work and Leisure in the US and Europe: Why So Different? *NBER Working Paper 11278*.
- Altonji, J. (1986). Intertemporal Substitution in Labor Supply: Evidence from Micro Data. *Journal of Political Economy* 94(3), 176–215.
- Baker, M., J. Gruber, and K. Milligan (2005). Universal Childcare, Maternal Labor Supply, and Family Well-Being. *NBER Working Paper 11832*.
- Blau, D. (2003). Childcare Subsidy Programs. *Means-Tested Transfer Programs in the United States*, Robert A. Moffitt, ed. Chicago: University of Chicago Press.
- Blau, F., M. Ferber, and A. Winkler (2002). *The Economics of Women, Men, and Work*. Prentice Hall.
- Blau, F. and L. Kahn (2000). Gender Differences in Pay. *Journal of Economic Perspectives* 14(4-Fall), 75–99.
- Blundell, R. and T. MaCurdy (1999). Labor Supply: A Review of Alternative Approaches. *Handbook of Labor Economics* 3, 1559–1695.
- Blundell, R., T. MaCurdy, and C. Meghir (2007). Labor Supply Models: Unobserved Heterogeneity, Nonparticipation and Dynamics. *Handbook of Econometrics* 6A, 4667–4775.
- Card, D., F. Kramarz, and T. Lemieux (1999). Changes in the Relative Structure of Wages and Employment: A Comparison of the United States, Canada, and France. *Canadian Journal of Economics* 32(4), 843–877.
- Chen, Y. and S. Rosenthal (2008). Local Amenities and Life Cycle Migration: Do People Move for Jobs or Fun? *Journal of Urban Economics* 64(3), 519–537.
- Chiappori, P. (1988). Rational Household Labor Supply. *Econometrica* 56(1), 63–90.
- Chiappori, P. (1992). Collective Labor Supply and Welfare. *Journal of Political Economy* 100(3), 437–467.
- Cogan, J. (1981). Fixed Costs and Labor Supply. *Econometrica* 49(4), 945–963.
- Cortes, P. and J. Tessada (2007). Cheap Maids and Nannies: How Low-Skilled Immigration Is Changing the Labor Supply of High-skilled American Women. *Working Paper*.

- Costa, D. and M. Kahn (2000). Power Couples: Changes in the Locational Choice of the College Educated, 1940-1990. *Quarterly Journal of Economics* 115(4), 1287–1315.
- Fernández, R., A. Fogli, and C. Olivetti (2004). Mothers and Sons: Preference Formation and Female Labor Force Dynamics. *Quarterly Journal of Economics* 119(4), 1249–1299.
- Fortin, N. (1995). Allocation Inflexibilities, Female Labor Supply, and Housing Assets Accumulation: Are Women Working to Pay the Mortgage? *Journal of Labor Economics* 13(3), 524–557.
- Freeman, R. and R. Schettkat (2001). Marketization of production and the US-Europe employment gap. *Oxford Bulletin of Economics and Statistics* 63(s 1), 647–670.
- Goldin, C. (2006). The Quiet Revolution that Transformed Women's Employment, Education, and Family. *American Economic Review* 96(2), 1–21.
- Gordon, P., A. Kumar, and H. Richardson (1989). Gender Differences in Metropolitan Travel Behaviour. *Regional Studies* 23(6), 499–510.
- Ham, J. (1986). Testing Whether Unemployment Represents Intertemporal Labour Supply Behaviour. *Review of Economic Studies* 53(4), 559–578.
- Ham, J. and K. Reilly (2002). Testing Intertemporal Substitution, Implicit Contracts, and Hours Restriction Models of the Labor Market Using Micro Data. *American Economic Review* 92(4), 905–927.
- Heim, B. and B. Meyer (2004). Work Costs and Nonconvex Preferences in the Estimation of Labor Supply Models. *Journal of Public Economics* 88(11), 2323–2338.
- Johnson, W. (2009). House Prices and Female Labor Force Participation. Working paper.
- Juhn, C. and S. Potter (2006). Changes in Labor Force Participation in the United States. *Journal of Economic Perspectives* 20(3), 27–46.
- Kain, J. (1968). Housing Segregation, Negro Employment, and Metropolitan Decentralization. *Quarterly Journal of Economics* 82(2), 175–197.
- Killingsworth, M. and J. Heckman (1986). Female Labor Supply: A Survey. *Handbook of Labor Economics* 1, 103–204.
- Kolesnikova, N. (2007). Three Essays on Local Labor Markets. Doctoral Dissertation, Carnegie Mellon University.
- MaCurdy, T. (1981). An Empirical Model of Labor Supply in a Life-Cycle Setting. *Journal of Political Economy* 89(6), 1059–1085.



- Mincer, J. (1962). Labor Force Participation of Married Women: A Study of Labor Supply. *Aspects of Labor Economics*, 63–97.
- Nickell, S. (1997). Unemployment and Labor Market Rigidities: Europe versus North America. *Journal of Economic Perspectives* 11(3), 55–74.
- Odland, J. and M. Ellis (1998). Variations in the Labour Force Experience of Women Across Large Metropolitan Areas in the United States. *Regional Studies* 32(4), 333–347.
- Oi, W. (1976). Residential Location and Labor Supply. *Journal of Political Economy* 84(4), 221–238.
- Reilly, K. (1994). Annual Hours and Weeks in a Life-Cycle Labor Supply Model: Canadian Evidence on Male Behavior. *Journal of Labor Economics* 12(3), 460–477.
- Ruggles, S., A. J. T., K. Genadek, R. Goeken, M. B. Schroeder, and M. Sobek (2010). *Integrated Public Use Microdata Series: Version 5.0 [Machine-readable database]*. Minneapolis: University of Minnesota.
- Ruhm, C. (1998). The Economic Consequences of Parental Leave Mandates: Lessons from Europe. *Quarterly Journal of Economics* 113(1), 285–317.
- Ward, C. and A. Dale (1992). Geographical Variation in Female Labour Force Participation: An Application of Multilevel Modelling. *Regional Studies* 26(3), 243–255.

FIGURE 1. Labor Force Participation Trends of Non-Hispanic White Married Women with a High School Degree

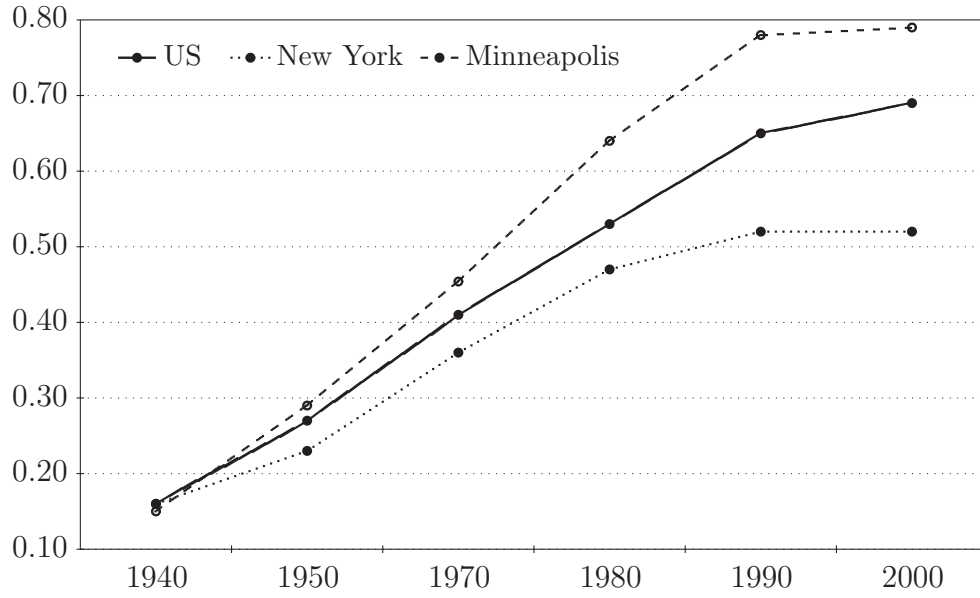


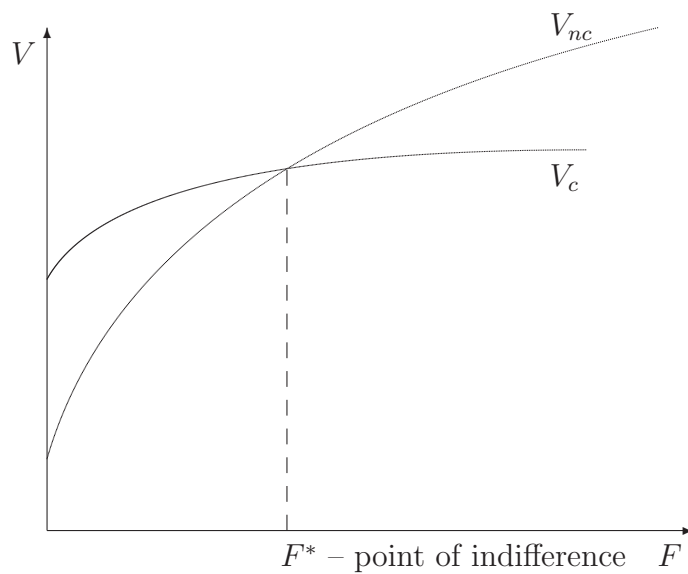
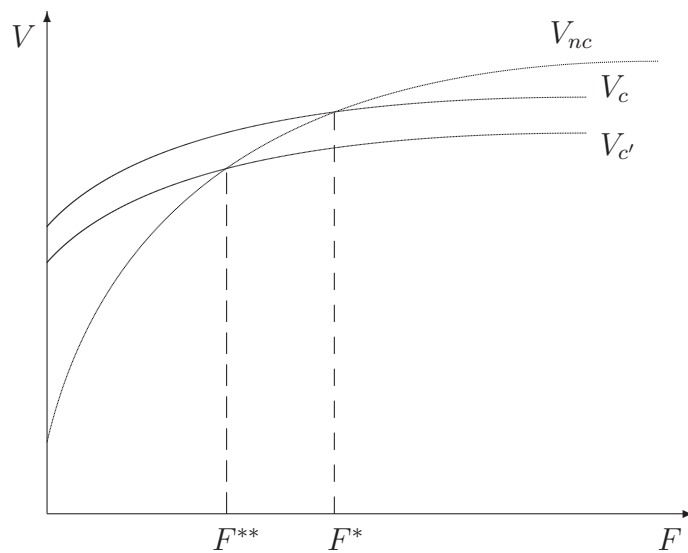
FIGURE 2. Indirect Utility as a Function of  $F$ —Commuting and Not CommutingFIGURE 3. Consequence of an Increase in Commuting Time from  $c$  to  $c'$ 

FIGURE 4. Effect of an Increase in Commuting Time  $c$  on Consumption and Leisure in the Static Model

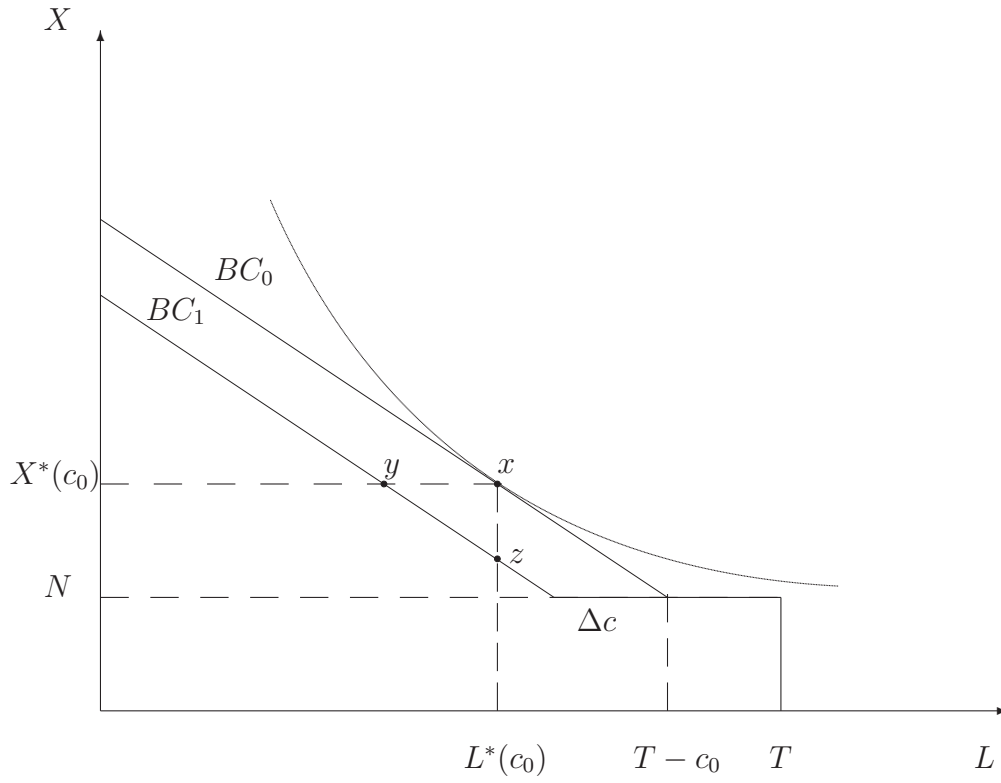


FIGURE 5. “Convexifying” the Objective Function with an Intertemporal Transfer

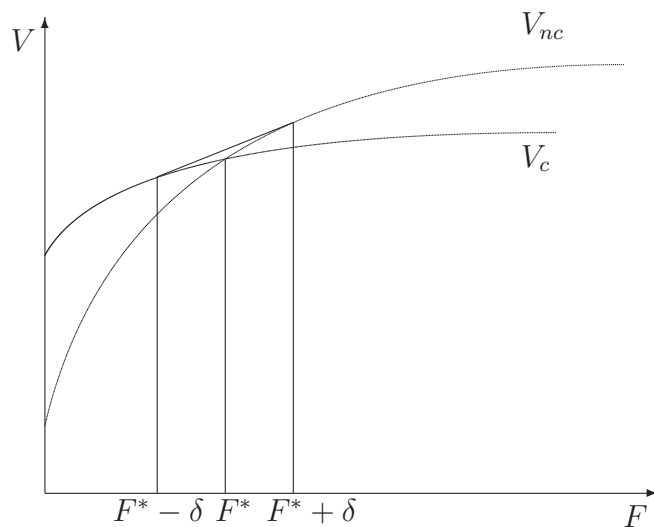


FIGURE 6. Indirect Utility Functions for a Person Who Is Indifferent between Working One Period and Two Periods

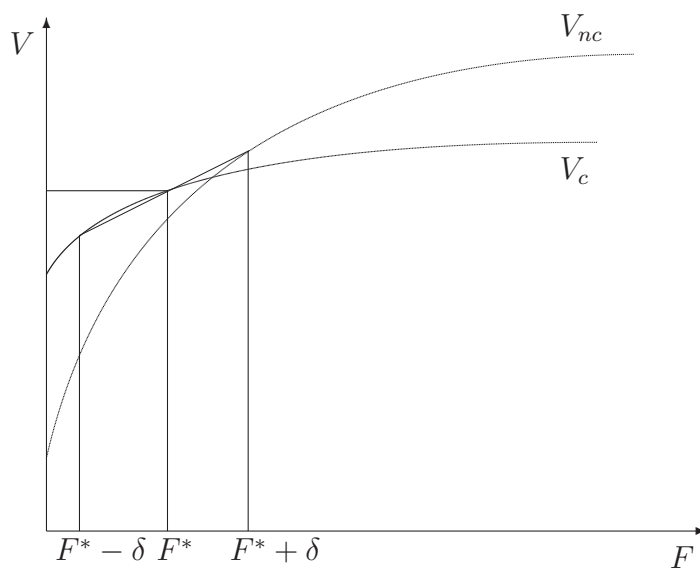


TABLE 1. Labor Force Participation Rates of Non-Hispanic White Married Women with a High School Degree, Census PUMS

MSA	Census Year					
	2000	1990	1980	1970	1950	1940
<b>United States</b>	<b>0.69</b>	<b>0.65</b>	<b>0.53</b>	<b>0.41</b>	<b>0.27</b>	<b>0.16</b>
New York	0.52	0.52	0.47	0.36	0.23	0.16
Honolulu	0.55	0.62	0.46	0.34	—	—
Los Angeles	0.59	0.59	0.54	0.41	0.31	0.21
Miami	0.61	0.65	0.55	0.45	0.29	0.23
Houston	0.61	0.63	0.56	0.37	0.34	0.16
San Francisco	0.62	0.66	0.57	0.42	0.34	0.20
San Diego	0.62	0.62	0.51	0.39	0.29	0.14
San Antonio	0.63	0.60	0.53	0.37	0.37	0.18
Detroit	0.63	0.59	0.45	0.31	0.26	0.17
New Orleans	0.64	0.62	0.44	0.36	0.15	0.20
Sacramento	0.64	0.64	0.54	0.41	0.50	0.20
Birmingham	0.65	0.63	0.53	0.43	0.29	0.16
Pittsburgh	0.65	0.58	0.41	0.30	0.14	0.08
West Palm Beach	0.65	0.65	0.53	0.40	—	—
Portland	0.66	0.66	0.54	0.41	0.31	0.20
Norfolk	0.67	0.61	—	0.40	0.16	0.11
Oklahoma City	0.67	0.68	0.57	0.49	0.51	0.26
Chicago	0.67	0.64	0.53	0.40	0.28	0.17
Las Vegas	0.68	0.63	0.56	0.45	—	—
Phoenix	0.68	0.64	0.54	0.49	0.22	0.20
Seattle	0.68	0.67	0.55	0.41	0.25	0.16
Atlanta	0.68	0.69	0.59	0.44	0.34	0.26
Dallas-Fort Worth	0.69	0.69	0.60	0.50	0.35	0.24
Memphis	0.69	0.66	0.55	0.41	0.40	0.18
Philadelphia	0.69	0.64	0.49	0.35	0.23	0.15
Austin	0.69	0.71	0.65	0.47	0.37	0.23
Orlando	0.69	0.69	0.56	0.38	0.35	—
Tampa	0.69	0.69	0.53	0.43	0.26	0.22
Dayton	0.69	0.66	0.51	0.42	0.41	0.19
Cleveland	0.69	0.65	0.49	0.36	0.25	0.17
Charlotte	0.70	0.73	0.65	0.54	0.25	0.29
Nashville	0.70	0.71	0.60	0.44	0.34	0.34
Salt Lake City	0.70	0.67	0.53	0.44	0.22	0.12
Boston	0.71	0.65	0.54	0.40	0.24	0.11
Buffalo	0.71	0.65	0.47	0.37	0.23	0.12
Baltimore	0.72	0.69	0.54	0.41	0.29	0.18
Indianapolis	0.72	0.70	0.57	0.46	0.30	0.22
St. Louis	0.72	0.67	0.53	0.38	0.25	0.17
Cincinnati	0.72	0.68	0.52	0.35	0.21	0.11
Louisville	0.73	0.68	0.54	0.40	0.31	0.19
Washington	0.73	0.72	0.58	0.45	0.35	0.26
Richmond	0.73	0.73	0.60	0.39	0.35	0.25
Kansas City	0.74	0.69	0.57	0.49	0.31	0.23
Denver	0.74	0.72	0.58	0.47	0.38	0.14
Albany	0.74	0.70	0.56	0.43	0.29	0.15
Columbus	0.74	0.70	0.57	0.39	0.40	0.20
Rochester	0.75	0.68	0.54	0.43	0.31	0.23
Greensboro	0.77	0.76	0.68	0.59	0.37	0.30
Milwaukee	0.78	0.74	0.60	0.46	0.33	0.17
Minneapolis	0.79	0.78	0.64	0.44	0.29	0.15

TABLE 2. Summary of Married Women's Labor Force Participation by Education and Presence of Children in 50 Large MSAs

	With Children Under 5		With Children, None Under 5		No Children	
	High School	College	High School	College	High School	College
lowest MSA	0.29	0.41	0.56	0.65	0.54	0.78
10th percentile	0.44	0.54	0.65	0.7	0.65	0.82
25th percentile	0.48	0.56	0.67	0.72	0.69	0.84
median MSA	0.53	0.6	0.72	0.74	0.73	0.86
75th percentile	0.59	0.65	0.74	0.78	0.75	0.88
90th percentile	0.62	0.67	0.77	0.79	0.78	0.89
highest MSA	0.68	0.71	0.83	0.86	0.82	0.93

Note: Data are from the 2000 Census PUMS. The sample consists of non-Hispanic white married women aged 25 to 55 years with non-imputed data. The unit of observation is the MSA. There are 50 MSAs. The MSA-specific averages are standardized using the sample age distribution.

TABLE 3. MSA-Level Regression: Labor Force Participation of Married Women

N=50	High School	Bachelor's Degree
Quality Adjusted Housing Index $\times 10^6$	-4.90 (3.75)	-4.09 (2.77)
Single Women Wage	-0.002 (0.003)	-0.001 (0.002)
Unemployment Rate	-3.01** (1.02)	-0.343 (0.767)
$R^2$	0.311	0.120
N	50	50

Significance Levels, One-Tailed Tests: \*5%, \*\*1%

Note: Data are from 2000 Census PUMS. The unit of observation is an MSA. Standard errors are in parentheses.

TABLE 4. Summary of Daily Commuting Time in 50 Large MSAs

	Married Men	Married Women	Married Women with Children
lowest MSA	43.5	38.6	37.4
10th percentile	47.0	41.1	40.5
25th percentile	50.4	44.0	43.0
median MSA	54.3	47.4	46.5
75th percentile	58.1	50.6	49.2
90th percentile	63.1	53.4	51.8
highest MSA	76.1	63.4	61.2

Note: Data are from the 2000 Census PUMS. The sample consists of non-Hispanic white married men and women aged 25 to 55 years with non-imputed data. The unit of observation is the MSA. There are 50 MSAs. The MSA-specific averages are standardized by age and education.

TABLE 5. MSA-Level Regression: Labor Force Participation of Married Women (with Commute Time as an Explanatory Variable)

N=50	High School	Bachelor's Degree
Quality Adjusted Housing Index $\times 10^6$	-1.56 (3.58)	-3.08 (2.79)
Single Women Wage	0.12 (0.29)	0.10 (0.22)
Unemployment Rate	-3.52** (0.94)	-0.55 (0.76)
MSA Average Daily Commute $\times 10^{-2}$	-0.33** (0.10)	-0.15* (0.10)
$R^2$	0.440	0.171
N	50	50

Significance Levels, One-tailed Test: \*5%, \*\*1%

Note: Data are from 2000 Census PUMS. The unit of observation is an MSA. Standard errors are in parentheses.



TABLE 6. Individual-Level Regression: Labor Force Participation of Married Women, by Presence of Children and Woman's Education

	With Children Under 5		With Children, None Under 5		No Children	
	H. School	College	H. School	College	H. School	College
<b>Panel A1. Labor Force Participation, OLS Regression</b>						
MSA Commute	-0.51** (0.228)	-0.36** (0.078)	-0.31** (0.114)	-0.20** (0.075)	-0.14 <sup>†</sup> (0.105)	-0.05 (0.057)
$R^2$	0.0063	0.0031	0.0025	0.0011	0.0005	0.0001
<b>Panel A2. Labor Force Participation, IV Regression</b>						
MSA Commute	-0.78** (0.319)	-0.61** (0.241)	-0.62** (0.214)	-0.40* (0.234)	-0.30 <sup>†</sup> (0.213)	-0.19 <sup>†</sup> (0.135)
$R^2$	0.0046	0.0016	0.0000	0.0000	0.0000	0.0000
$F$ (first stage)	13.8	18.9	13.3	17.1	13.9	14.1
<b>Panel B1. Labor Force Participation, using Husband's Education, OLS Regression</b>						
MSA Commute	-0.44** (0.148)	-0.32** (0.083)	-0.25** (0.084)	-0.13* (0.074)	-0.14 <sup>†</sup> (0.098)	-0.03 (0.055)
Number of Children	-7.11** (0.4)	-10.37** (0.4)	-4.21** (0.3)	-5.74** (0.3)		
MSA Unemp. Rate	-4.45** (1.42)	-0.72 (0.82)	-3.63** (0.75)	-1.53* (0.63)	-2.18** (0.82)	-1.08** (0.41)
<i>Controls</i>						
Woman's Age	Yes	Yes	Yes	Yes	Yes	Yes
Husband's Education	Yes	Yes	Yes	Yes	Yes	Yes
$R^2$	0.0376	0.0735	0.0174	0.0474	0.0226	0.0541
<b>Panel B2. Labor Force Participation, using Husband's Education, IV Regression</b>						
MSA Commute	-0.64** (0.249)	-0.69** (0.237)	-0.44** (0.184)	-0.38* (0.225)	-0.24 (0.204)	-0.17 <sup>†</sup> (0.114)
Number of Children	-7.10** (0.4)	-10.48** (0.5)	-4.15** (0.3)	-5.73** (0.3)		
MSA Unemp. Rate	-4.17** (1.47)	-0.41 (1.31)	-3.50** (0.90)	-1.41 <sup>†</sup> (0.94)	-2.12** (0.82)	-0.97* (0.55)
<i>Controls</i>						
Woman's Age	Yes	Yes	Yes	Yes	Yes	Yes
Husband's Education	Yes	Yes	Yes	Yes	Yes	Yes
$R^2$	0.0367	0.0702	0.0165	0.0456	0.0224	0.0531
$F$ (first stage)	19.8	22.7	18.2	19.9	16.3	17.0
N	13,483	26,654	41,889	32,007	22,770	27,507

Significance Levels, One-Tailed Test: †10%, \*5%, \*\*1%.

Note: Standard errors are clustered at MSA level. Data are from the 2000 Census PUMS. The dependent variable is a labor force participation dummy multiplied by 100 (so interpretation is in percentage points). The sample, drawn from the 2000 PUMS, consists of non-Hispanic white married women aged 25 to 55 years with non-imputed data, living in the 50 largest MSAs. The MSA unemployment rate is for white men and is measured in percentage points. Controls for the woman's age are a set of dummy variables for each of the 5-year intervals between 25 and 55 years old. Controls for a husband's education are a set of dummy variables for each education category (less than high school, high school diploma, etc.).

TABLE 7. MSA-Level Regressions: Labor Force Participation of Married Women, 1990 and 1980

	High School Educated		College Educated	
	(1)	(2)	(1)	(2)
<b>Panel A. MSA Labor Force Participation Rates in 1990</b>				
MSA Commute	-0.314** (0.097)	-0.279** (0.086)	-0.112* (0.056)	-0.110* (0.060)
MSA Unemployment		-2.64** (0.674)		-0.15 (0.449)
Intercept	82.38** (4.93)	88.46** (4.60)	79.14** (2.86)	79.49** (3.06)
$R^2$	0.180	0.382	0.076	0.079
N	50	50	50	50
<b>Panel B. MSA Labor Force Participation Rates in 1980</b>				
	High School Educated		College Educated	
	(1)	(2)	(1)	(2)
MSA Commute	-0.231** (0.104)	-0.226** (0.085)	-0.011 (0.082)	-0.010 (0.082)
MSA Unemployment		-27.3** (0.561)		-0.639 (0.540)
Intercept	66.04** (5.21)	73.57** (4.56)	58.61** (4.15)	60.37** (4.39)
$R^2$	0.096	0.403	0.001	0.030
N	49	49	49	49

Significance levels: \* 5% , \*\* 1%.

Note: Data are from 1990 and 1980 Census PUMS. Norfolk is not identified in 1980 Census data. Standard errors are in parentheses. The dependent variable is MSA-specific labor force participation of non-Hispanic white married women. Unemployment rate is for white men. Both unemployment and labor force participation are measured in percentage points.

TABLE 8. Differences-in-Differences Estimation

<b>Panel A. <math>\Delta</math> Labor Force Participation, High School Graduates</b>						
	1980-2000		1980-1990		1990-2000	
	(1)	(2)	(1)	(2)	(1)	(2)
$\Delta$ Commute	-0.51** (0.188)	-0.31* (0.171)	-0.28* (0.167)	-0.25 <sup>†</sup> (0.158)	-0.62** (0.196)	-0.65** (0.177)
$\Delta$ Unemp. Rate		-2.61** (0.649)		-1.11* (0.425)		-1.69** (0.484)
$R^2$	0.135	0.360	0.058	0.180	0.171	0.342
N	49	49	49	49	50	50
<b>Panel B. <math>\Delta</math> Labor Force Participation, College Graduates</b>						
	1980-2000		1980-1990		1990-2000	
	(1)	(2)	(1)	(2)	(1)	(2)
$\Delta$ Commute	-0.63** (0.139)	-0.48** (0.128)	-0.12 (0.177)	-0.08 (0.167)	-0.33 (0.224)	-0.37 <sup>†</sup> (0.201)
$\Delta$ Unemp. Rate		-1.91** (0.485)		-1.20** (0.448)		-1.94** (0.552)
$R^2$	0.299	0.475	0.010	0.142	0.043	0.242
N	49	49	49	49	50	50

Significance levels: †10%, \*5%, \*\*1%.

Note: Data are from 1980, 1990, and 2000 Census PUMS. The unit of observation is an MSA. Norfolk is not identified in 1980 Census data. In each of the regressions, the dependent variable is change of LFP rates of non-Hispanic white married women in each MSA in the indicated time period, measured in percentage points. The independent variable is the change in average MSA daily commuting time of non-Hispanic white married men over the same period. Standard errors are in parentheses.