POLITICAL ECONOMY RESEARCH INSTITUTE
PERI
University of Massachusetts Amherst

The Interaction of Metropolitan Cost-of-Living & the Federal Earned Income Tax Credit: One Size Fits All?

Jeffrey P. Thompson
& Katie Fitzpatrick

August 2009

POLITICAL ECONOMY RESEARCH INSTITUTE

Gordon Hall
418 North Pleasant Street
Amherst, MA 01002

Phone: 413.545.6355
Fax: 413.577.0261
peri@econs.umass.edu
www.peri.umass.edu

UMASS

WORKINGPAPER SERIES

Number 204

The Interaction of Metropolitan Cost-of-living & the Federal Earned Income Tax Credit: One Size Fits All?

Jeffrey P. Thompson
Assistant Research Professor
Political Economy Research Institute
Gordon Hall
418 N. Pleasant St., Suite A
Amherst, MA 01002
United States
jthompson@peri.umass.edu

Katie Fitzpatrick
Economist
Economic Research Service, USDA
kfitzpatrick@ers.usda.gov

Abstract:

This paper explores the interaction between the federal Earned Income Tax Credit (EITC) and the cost-of-living faced by single mothers. After the 1993 EITC expansion, we identify up to a 10 percentage point increase in labor force participation for single mothers in the lowest cost areas but no discernable response in the highest cost areas. We conclude that the EITC's welfare-enhancing properties are undermined by the interaction of the program's fixed national rules and geographic variation in wages and cost-of-living. In addition, our findings suggest that the EITC does little to reduce joblessness in many of the nation's largest cities.

We would like to thank Stacy Dickert-Conlin, Andrew Hanson, Jeffrey Kubik, Tim Smeeding, Tracy Gordon, Gary Engelhardt, Chris Rohlfs, and Perry Singleton for helpful comments and guidance. We would also like to thank Yong Chen and Stuart Rosenthal for use of their quality-adjusted rental cost data. The views expressed in this paper are solely those of the authors and should not be construed in whole or in part to any branch of the federal government. Any remaining errors are our own.

JEL Codes: H24, H31, I31, J22, R23
Keywords: EITC; Cost-of-Living; Tax reform; Labor supply

" Among the 122 large cities…the average EITC (for all earners) in 2003 ranged from roughly $1,200 in Cambridge, MA, to $2,284 in McAllen, along the Texas-Mexico border."

"The New Safety Net: How the Tax Code Helped Low-Income Working Families During the Early 2000s" (Berube, 2006)

Introduction

The federal Earned Income Tax Credit (EITC) is a wage subsidy available to lower-income, working families. Since its inception in 1975, major expansions in 1986, 1993, and 2001 contributed to large increases in the size of the benefit and the number of potential beneficiaries. By 2008, the EITC was worth up to $4,800 and families with earnings as high as $38,000 qualified for some credit.[1]

Policymakers intend for the EITC to reward work by altering the labor supply incentives of the potentially eligible. Estimates of the effect of the EITC on labor supply consistently find large positive effects on the decision to work but no effect on the decision of how much to work. Previous studies, however, fail to adequately address the influence of geographic differences in both wages and the cost-of-living. Cost differences make the credit more (or less) valuable across geographic areas. With a nationally uniform *benefit structure*, the EITC is more valuable in a geographic area with a low cost-of-living relative to an area where the cost-of-living is high.

In addition, the nationally uniform *eligibility rules* effectively treat equivalent workers differently across geographic areas because, although net wages may equalize across areas for specific worker-types, gross wages vary considerably.[2] EITC eligibility based on gross income, therefore, results in variation in EITC benefits across geographic areas. In general, low-skilled workers in high-cost areas earn higher gross wages and are more likely to end up on the phase-

[1] A small credit for very low-income childless workers was added in 1993.
[2] Albouy (2008) – discussed further below – examines the broader issue of the economic consequences of a nationally uniform federal income tax code in the face of regional differences in wages and cost-of-living.

out portion of the EITC schedule or off the schedule completely than similar workers in low-cost areas. For example, a single mother working full-time as a janitor in a high-cost city (Cambridge, MA) may have earnings that place her on the phase-out portion of the credit. In contrast, with a lower wage in a low-cost city (McAllen, TX), her annual earnings would place her on the phase-in portion. The single mother in this example would earn different credit amounts depending on whether she lived in the high-cost city or the low-cost city.

Local costs are critical to analyzing the EITC because earnings and the bundle of goods and service available for purchase with earnings are realized in specific local labor markets. We address differences across geographic areas by including a measure of location-specific prices – housing costs of the Metropolitan Statistical Area (MSA) – to examine heterogeneous effects of the EITC across geographic areas. Using the 1993 EITC expansion, we find that the effect of the EITC on labor supply depends on the housing costs in the worker's local labor market. The EITC contributed to an approximately 10 percentage point increase in the participation of single women in the lowest-cost areas. We find no evidence of a participation effect for single women in the highest-cost metropolitan areas, where nearly 40 percent of the population lives. We find some evidence of differences in the hours of work decision across cost areas, but these results are not robust to the selection of our sample or to the choice of our reform period.

This paper proceeds as follows: Section II discusses how the EITC affects labor supply decisions, the theory behind wage differences across local areas, and the relevant literature on the EITC. Section III provides our data and methodology. Section IV provides our estimates for participation and hours worked decision. Section V discusses the implications of the findings, and Section VI concludes.

Section II: Theory and Literature

The EITC and Labor Supply

The structure of the EITC (displayed in Figure 1) includes a “phase-in,” “plateau,” and “phase-out” region. Earnings in the phase-in region receive a constant rate subsidy, up to the maximum credit. Earnings in the plateau region receive the maximum credit. Once earnings reach the phase-out region the credit decreases at a constant rate for each additional dollar of earned income until the credit is completely eliminated.

< Figure 1 approximately here>

In the standard static model of labor supply, the EITC shifts out the budget constraint and provides unambiguously positive incentives on labor force participation.[3] However, this shifted budget constraint also creates EITC-induced kinks. As a result, the impact of the EITC on hours worked is ambiguous owing to negative income effects (assuming leisure is normal good) over the entire schedule but substitution effects that vary across the regions of the credit.

The phase-in region contains positive substitution effects that encourage additional hours of work by increasing the hourly return to work; no substitution effect exists in the plateau region; a negative substitution effect in the phase-out region reduces the hourly return to work. As a result, the net effect varies across regions: in the phase-in region, the net effect is theoretically ambiguous while in the plateau and phase-out regions the net effect is unambiguously negative.[4] Thus, the overall effect of the EITC on hours becomes an empirical

[3] This is the case for single women. Married couples eligible for the EITC face more complex participation decisions.
[4] For incomes greater than the phase-out region, the EITC may induce a taxpayer to reduce her hours to receive a credit.

question which depends on the distribution of beneficiaries across the schedule and the relative magnitudes of the income and substitution effects.

Regional Differences in the Cost-of-living

The EITC is expected to have different impacts on labor supply across MSAs due to variation in the cost-of-living, particularly the considerable geographic variation in housing costs. The causes and consequences of this geographic variation have been the subject of considerable interest, both in the economics literature and in policy debates.[5] In fact, a National Academy of Science (NAS) commissioned study recommended that the federal poverty threshold be adjusted to reflect differences in housing costs and other prices across geographic areas (Citro and Michael, 1995).[6] The NAS study showed that wages were higher in areas with a high cost-of-living.

We find empirical support for this relationship between wages and the cost-of-living. We use 1990 quality-adjusted annual rental cost data provided by Chen and Rosenthal (2008) to measure the cost-of-living.[7] Using the Current Population Survey (CPS) Outgoing Rotation Group data for 1990-1995, we estimate average hourly wages of single, female household heads ages 18 to 49 by deciles of the MSA quality-adjusted housing costs. In Table 1, we show estimates for selected industry and occupations that employ the greatest numbers of single

[5] See Rosen (1979), Roback (1988), and Hoynes (2000).
[6] These recommendations to adjust the poverty threshold to differences in housing costs or other price differences across geographic areas were not ultimately adopted due to a variety of reasons, including measurement problems, lack of data, and political constraints. The NAS did conclude, however, that "the available data suggest that areas with higher prices are also areas with higher income levels: for example, a cost-of-housing index that we calculated for states correlates highly with state median family income." (Citro and Michael, 1995; 184)
[7] Chen and Rosenthal construct their measure by estimating a hedonic regression controlling for structural characteristics of housing units in each MSA from the 1990 Census. From these estimates, Chen and Rosenthal report housing costs for each MSA relative to the mean, ranging from $3,785 below the mean to $6,152 above the mean. For ease in interpretation, we transform Chen and Rosenthal's measure into a positive value for all MSAs by adding $4,000 to each value. The new range of quality adjusted rent, which we refer to as our rental costs, is $215 to $10,152.

women. Wages for nursing aids was $5.51 in the lowest-cost decile and $8.32 in the highest decile; average wages in eating and drinking establishments were $4.12 in the lowest decile and $5.54 in the highest. After controlling for demographic and labor market characteristics, a regression of hourly wages on average rent yields a coefficient of 0.00031, suggesting $1,000 in higher quality-adjusted annual rents is associated with $.31 higher hourly wages (which translates to $645 in annual earnings for full-time, full-year workers). Separate regressions by occupation and industry groups (included in Table 1) yield similar results.

< Table 1 approximately here >

This relationship between local prices on wages is established in several models, including those by Roback (1988), who extends the classic Rosen-Roback hedonic model to include heterogeneous labor inputs, and more recently by Black, et al. (2007).[8] With two skill-types and variation across cities in amenities, both of these models show that if low-skilled types have a lower willingness to pay for amenities – whether or not the amenity is productive – the low-skilled must receive a higher wage in a high-amenity city to have the same utility level across cities.[9] The willingness of high-skill types to pay for the amenity raises rent in the high-amenity cities, and higher rent must be offset through higher wages if low-skill types are also to reside in high-amenity cities. Equilibrium sorting implies that wages, and therefore incomes, will differ across cities for the same skill types and that low-skill types will have higher wages in high-cost regions.

[8] Black et. al. are specifically concerned with the variation in the returns to education, contrasting earnings of college graduates with high school graduates, but the logic of the model applies to wages as well.
[9] An additional assumption is that the different labor inputs are not perfect substitutes.

The EITC, Labor Supply, and Cost-of-living

Geographic variation in wages implies that low-skilled workers will face different EITC treatment based on where they live. To show this empirically, in the period before the EITC expansion of 1993, we examine the incomes of employed single women relative to the EITC schedule for different MSAs in Table 2. [10] The top panel of Table 2 contains estimates for all single women while the bottom panel displays estimates for single women with a high school degree or less. In the early 1990s, 12.6 percent of single females in MSAs in the lowest quarter of the rental cost distribution have incomes too high to be on the EITC schedule compared to 32.1 percent in the top quarter of MSAs. For low-educated women, the figures are 6.3 percent and 18.3 percent, respectively. In addition, eligible workers in low-cost areas are more likely to fall in the phase-in and plateau regions of the credit, where the benefit is larger.

< Table 2 approximately here>

We can unambiguously predict that an expansion of the EITC will have a greater impact on labor force participation in low-cost areas than in high-cost areas for three reasons. The first two of these reasons follow directly from our estimates in Table 2. Workers in low-cost areas have lower wages. These lower wages result in annual incomes that are more likely to make these workers income-eligible for the EITC. Secondly, once income-eligible for the credit, workers in low-cost areas are more likely to receive larger benefits because their incomes are more likely to place them on the phase-in or plateau region rather than the phase-out region. Finally, related to the variation that exists in wages, geographic variation also exists in housing prices. As a result, any given nominal benefit will have different purchasing power across metropolitan areas. EITC benefits to a worker in a low-cost area have greater purchasing power –

[10] We measure adjusted gross from the prior year income information in the March CPS for 1990 to 1993.

and, thus, these benefits are a greater real incentive – than the same nominal benefit to a worker in a high-cost area.

It is less clear how variation in MSA costs will affect the decision of how many hours to work because of offsetting income and substitution effects. With a larger share of low-skilled workers directly impacted by the EITC, hours in low-cost areas should be more responsive overall to the policy change. And with a larger share of workers located in the phase-in region of the credit, low-cost areas should be more likely to have positive responses on the hours of work decision. At the same time, however, the income effect should be greater in low-cost areas because the nominal benefit has greater purchasing power. Ultimately, the mix of incentives faced by workers on different portions of the credit makes it difficult to make strong predictions about responses in hours of work to the EITC. In short, we expect the hours worked decision to be less responsive to the cost-of-living than the participation decision.

Previous EITC Literature

A large literature studying the labor supply response to the EITC, fully reviewed in Hotz and Scholz (2003) and Eissa and Hoynes (2006), emerged after the pioneering work of Eissa and Liebman (1996). Eissa and Liebman examine the EITC expansion in the Tax Reform Act of 1986 (TRA86) with a difference-in-difference analysis. Because only families with children could receive the credit, Eissa and Liebman use single mothers as the treatment group and single, childless women as the control group. They find that the 1986 expansion of the EITC increased the labor force participation of single mothers by 2.8 percentage points relative to single women without children. Depending on their specification, they estimate no change or a small, positive change in the hours worked of single mothers relative to single women without children.

The findings of Eissa and Liebman's difference-in-difference approach are largely consistent with other approaches. A large increase in participation is found using a variety of econometric methodologies, samples, and expansion periods: a panel dataset of California welfare recipients (Hotz, et al., 2006); models including welfare use (Grogger, 2003); simulation studies (Dickert, et al., 1995; Scholz, 1996); and structural modeling with extensive controls for all tax and benefit changes over the 1984 to 1996 period (Meyer and Rosenbaum, 2001). In contrast, almost no study finds a substantial change in the hours worked of recipients.[11] Eissa and Liebman (1996) and Eissa and Hoynes (2006) posit a number of reasons for the inability of the EITC to influence the hours worked by a recipient: labor market norms and institutions which allow for only part-time or full time work, measurement error, and a lack of knowledge about the exact structure of the EITC.

To our knowledge, no examination of the labor supply response to the EITC rigorously considers cost-of-living differences across geographic areas. Meyer and Rosenbaum (2001) control for the state cost-of-living in their structural model, but they do not report estimates for this variable, nor do they interact it with their tax change variable. Other EITC work that examines geographic variation focuses on take-up of the credit and suggests that urban areas have lower utilization rates than other areas (Maynard and Dollins, 2002; Berube and Tiffany, 2004; Hirasuna and Stinson, 2004). In short, few analyses consider geographical differences in the EITC despite differences in participation and average credit size across state and metropolitan areas, and the discussion by advocates for the working poor of the variation in average credit amounts across metropolitan areas (Berube, 2006).

[11] The one exception we are aware of is Wu (2005), which shows different effects on the phase-in and phase-out regions of the EITC schedule, which cancel out to produce no overall effect on hours.

Section III: Methodology and Data Sources

Estimation Strategy

We consider the EITC expansion included in the Omnibus Budget Reconciliation Act of 1993 (OBRA93), which increased the maximum credit, extended EITC eligibility to those with higher incomes, and created a small credit for childless workers. These EITC increases were implemented in steps from 1994 through 1996 by adjusting five credit parameters, details of which are contained in Table 3. Potential recipients faced more generous benefits in 1994, 1995, and 1996 as a result of OBRA93.

< Table 3 approximately here >

We use the familiar difference-in-difference estimator to measure how an affected group (low-educated, single mothers) changes its labor supply relative to an unaffected group (low-educated, single women without children). We choose this sample for several reasons. Low-educated workers are more likely to have earnings in the credit range; single parents are the largest group of workers eligible for the EITC; women almost always head single parent families, and; unmarried individuals allow us to avoid intra-household bargaining decisions that affect married individuals. While OBRA93 extended EITC eligibility to those without children, the credit is quite small and available only to those extremely low incomes. Our identifying variation comes from group differences in tax schedules faced by single mothers and single women without children. For identification, we require that differential trends in labor force participation and hours of work do not exist between single mothers and single women without children.[12]

Unlike previous work, we allow for heterogeneous effects across local areas by interacting our cost-of-living measure by the difference-in-difference estimator. Our coefficient of interest is the heterogeneous effect of the EITC across metropolitan areas. This is not the standard triple-difference estimator because the addition of the cost-of-living variable does not provide us an additional control group. Instead, it allows us to explore differential responses across areas.

Data

The data we use are from the 1990 through 1995 monthly CPS.[13] The CPS is a monthly survey of approximately 50,000 households which provides current demographic, labor market, geographic, and income information for responding households. We construct tax units from the sample by matching children age 18 and under, as well as full-time students age 19 to 24, to their mothers. We limit our sample to single (never married, widowed, or divorced) women who are heads of tax units, ages 16 to 49. In our main results, we further limit our sample to those with a high school degree or less. We drop the self-employed, as well as unpaid agriculture workers, and those with negative unearned income. We drop from the sample those who report attending school full-time and those who report an illness or disability that prohibits work.

[12] Later we do explore the robustness of our findings using single women with more than one child as our treatment group and single women with only one child as the control group.
[13] We do not include summer months (June, July, and August) in our data because the geographical variables are not available in June, July, or August 1995 as a result of the CPS redesign. We do not include data from 1996 because of the work mandates that were associated with welfare reform legislation in 1996.

For each tax unit, we merge on unemployment rates in each MSA and an MSA cost-of-living measure.[14] For those tax units residing outside of an MSA, we merge on the state's non-MSA value for unemployment rates and cost-of-living. Our cost-of-living measure is the 1990 quality-adjusted housing costs provided by Chen and Rosenthal (2008). Chen and Rosenthal construct their cost measure by estimating a hedonic regression controlling for structural characteristics of housing units in each MSA and state non-MSA from the 1990 Census. From these estimates, Chen and Rosenthal report housing costs for each MSA and non-MSA relative to the mean, ranging from $3,785 below the mean to $6,152 above the mean. For ease in interpretation, we transform Chen and Rosenthal's measure into a positive value for all MSAs by adding $4,000 to each value. The new range of quality-adjusted rent, which we refer to as our rental costs, is $215 to $10,152.[15] Our use of geographic variation forces us to drop observations without a basic geographic identifier (MSA or state non-MSA) because we cannot assign unemployment rates or rental costs.

Table 4 presents summary statistics of the characteristics of our full sample, as well as our treatment and control groups in Columns 1 through 3 and across cost-of-living areas in Columns 4 through 7. Overall, our sample of childless women is more likely to have received a high school degree than our sample of single mothers. Single mothers are more likely to be nonwhite and live in MSAs with slightly lower average rental costs. Single mothers have much lower levels of labor force participation but, conditional upon working, their earnings, hours of work and unemployment rates are similar.

< Table 4 approximately here >

[14] Details on the creation of MSAs that are consistent over the 1990 to 1995 period are included in Appendix A.
[15] The full listing of MSA and state non-MSA rental costs is available from the authors.

Looking across metropolitan areas, higher-cost areas have more single women who have received their high school degree. The highest-cost areas are much more likely to have implemented a waiver to the state's Aid to Family with Dependent Children (AFDC) program. Despite the work mandates associated with welfare waivers, the highest-cost areas have lower levels of labor force participation. Conditional upon working, the difference in wages across local areas is nearly two dollars: hourly earners in the lowest quarter have average hourly wages of $5.99 while average hourly wages are $7.76 in the highest quarter. Similarly, conditional upon working, average weekly earnings are $70 higher in the highest quarter than in the lowest.

Other than differences in wages and earnings across areas, women in different quarters of the rental cost distribution appear roughly comparable in the number of children they have. Conditional on having any children, the number of children a woman has is not associated with the cost-of-living. Mothers in the lowest quarter of rental costs have, on average, 1.81 children. In the highest quarter of rental costs, mothers have on average 1.87 children. With these small differences, we expect that mothers in different areas would not qualify for different EITC benefits based solely on their demographic characteristics. Differences in EITC eligibility arise from differences in incomes.

Section IV: Results

Participation Estimates

We estimate how the effect of the EITC on labor force participation differs across local areas with the probit equation:

$$(1)\quad \Pr(LFP = 1) = \Phi\,(\alpha + \beta Z + \gamma_0 \text{ treatment} + \gamma_1 \text{ post} + \gamma_2 \text{ (treatment*post)} \\ + \gamma_3 \text{ (treatment*post*cost)} + \gamma_4 \text{ cost})$$

Our dependent variable, LFP, is a dichotomous variable equal to 1 if the respondent reported working last week and 0 if not. The difference-in-difference estimator, γ_2, measures how low-educated, single mothers change their labor force participation relative to low-educated, single women without children after 1993.[16] Our main coefficient of interest, γ_3, measures the heterogeneous effect of the EITC across local areas. Our independent variables (Z) control for observable differences between our treatment and control groups, as well as covariates associated with labor force participation. These include age, age squared, number of preschool age children, number of dependents,[17] the number of dependents squared, an indicator for more than one child, race, MSA unemployment rate, and educational attainment. We also control for the month of implementation of AFDC policy waivers. Standard errors are clustered at the MSA level. All reported estimates from the participation equations are the mean marginal effects.[18]

We present estimates of the mean marginal effects from our probit regressions in Table 5. We first estimate the effect of the EITC on labor market participation similar to prior work. We find that low-educated single mothers increased their employment rate by 4.7 percentage points relative to low-educated single women without children as a result of the 1993 expansion, in Column 1 of Table 5. This estimate is larger than the roughly three percent participation increase estimated by Meyer (2002) and Meyer and Rosenbaum (2001) for the 1993 expansion. Both studies, however, do not limit their sample by education. When we expand the sample to include

---

[16] Technically the interaction terms in a probit model are not straightforward to interpret. The coefficient on the interaction terms does not simply capture the marginal effect, but also includes additional terms that are conditional on the interacted variables as well as any other independent variables. We also performed Linear Probability Models (LPM) in addition to probit models. Our LPM results (not reported here) are similar to our probit results. We chose to report results from probit regressions for ease of comparison with other estimates in the literature. We also used the inteff procedure, described in Ai, et al. (2004), to obtain correct marginal effects (and standard errors) for the difference-in-difference variable in the probit equation. These results were nearly identical to the results obtained from calculating the mean marginal probit effects via Gelbach's (2004) margfx procedure, as well as results from LPM models. All are available upon request.

[17] We defined a dependent as a child under the age of 18 or between the ages of 18 and 24 and in school full time.

[18] We employed the margfx command to calculate the mean of the marginal effects, as opposed to calculating the marginal effect evaluated at the mean (Gelbach, 2004).

all levels of education, in Column 4, our estimates are nearly identical. Controls for education, age, race, and the local unemployment rate have the expected sign.

< Table 5 approximately here >

We explore whether the EITC participation effect differs systematically by local areas in Column 2. We begin by creating dichotomous variables for MSA in cost quarters, based on the distribution of quality-adjusted rental costs, omitting the first cost quarter. Interacting these dichotomous variables with the difference-in-difference variable demonstrates that the lowest cost quarter has a 7.3 percentage point increase in labor force participation. The estimate for the second cost quarter implies that the increase in these areas is 2.3 percentage points more than the lowest cost quarter, although the point estimate is insignificant. The third cost quarter implies a slightly lower response than the first cost quarter, although again it is insignificant. The estimate for the highest cost quarter is almost equal in magnitude and opposite in sign to the lowest cost quarter. In sum, the increase in labor force participation in the bottom three quarters of rental costs is 7.3 percentage points while the highest quarter of rental costs shows no response on participation.

To take advantage of the full variation in costs we interact the difference-in-difference estimator with our continuous measure of rental costs in Column 3. The difference-in-difference estimator rises to 10.2 percentage points. However, each $1,000 increase in our quality-adjusted rental costs reduces participation by one percentage point. These results again suggest no effect of the EITC on participation in the highest-cost areas.

< Table 6 approximately here >

The local cost-of-living may systematically impact all covariates associated with labor force participation, such as the cost of child care, conditions in the local labor market, and returns to education. The summary statistics in Table 4 demonstrates some differences in the observable characteristics of individuals in each metropolitan area in education and race. Additionally, the implementation of an AFDC waiver is positively correlated with high-cost MSAs, suggesting that states that implemented a waiver tend to contain high-cost MSAs. Using the distribution of rental costs, we split the sample of low-educated women into quarters by cost-of-living. We further split the highest rent quarter in half (75th to 87th percentile and 88th percentile and above) to determine if differences in MSAs at the upper tail of the distribution drove the lack of a participation effect found in Column 3 in the most expensive areas. We test to determine if we should pool these cost-of-living areas or estimate each area separately. A Wald test strongly rejects pooling (p=0.000).

We rerun our participation equation separately for each of these cost-of-living areas and report the results in Table 6. Overall, the EITC expansion effect is larger in the lower-cost areas than the higher-cost areas. The second (Column 2) and third (Column 3) quarter of costs have the largest and most significant effects: an increase in employment of 6.3 and 7.0 percentage points, respectively. Meanwhile, the first quarter (Column 1) has a slightly smaller response, with a rise in participation of 4.7 percentage points. Above the 75th percentile of rent (Columns 4 and 5), the EITC has no significant effect on participation.

We test whether each of these point estimates are significantly different from each other. We cannot conclude that the estimates in the first three quarters (Columns 1 through 3 of Table 6) are different from one another at the 10 percent significance level. However, virtually all of the point estimates in the first three quarters are significantly different at the 10 percent significance level from the point estimates from the 75th to 87th percentile (Column 4), as well as the point estimate from MSAs above the 87th percentile (Column 5). The one exception to these findings is that we cannot conclude that the point estimate from the lowest quarter of MSAs (Column 1) is statistically different from the point estimate from MSAs above the 87th percentile. The p-value from this Chi-Squared test is 0.24.

Robustness Checks

We perform several tests to explore whether our results are dependent on our methodological considerations, and if the identifying assumptions of the difference-in-difference estimator are valid. First, we expand our sample to all women, regardless of education level for each specification. The difference-in-difference, not including our cost-of-living variable, falls from 4.7 to 3.1 percentage points (Column 4 of Table 5). When we include the heterogeneous effects using dichotomous variables for each cost quarter in Column 5, our heterogeneous effect shows the same pattern as our baseline estimates but with smaller magnitudes. Participation increased by 4.3 percentage points in the lowest quarter, 7.1 points in the second quarter, and 4.3 percentage points in the third quarter. Again, there was no change in participation in the highest rent quarter. Results using the continuous cost-of-living measure (Column 6 of Table 5) imply that areas with the very highest costs (roughly above the 85th percentile) actually had a reduction in employment of single mothers relative to single, childless women as a result of the EITC.

When we split all single women, regardless of education, into quarters of the rental cost distribution and further split the top quarter in half, the same pattern of results is again apparent in Columns 6 through 10 of Table 6. The largest response is in the second quarter of costs (Column 7) with an almost 6 percentage point increase in labor force participation. The lowest quarter (Column 6) and third quarter (Column 8) display a similar response of roughly 3 percentage points. Unlike our sample of low-educated women, women in MSAs in the 75th to 87th percentile of the rental cost distribution also show a labor force participation response of 3 percentage points (Column 9). In MSAs above the 87th percentile (Column 10), there is no response in labor force participation.

Next, we check the robustness of our cost measures with two different measures of housing costs: Housing and Urban Development (HUD) Fair Market Rent data from 1990 and median rent data from the 1990 Census.[19] Both measures suggest the same magnitude and pattern of results as our quality-adjusted rental cost data.[20] (The distribution of these two alternative housing cost measures, compared with our positive quality-adjusted measure is included in Appendix B.)

Finally, we test the validity of the identifying assumption in the difference-in-difference estimator – that the policy change under consideration is the only group and time-varying factor (outside of the additional covariates) impacting the dependent variable is valid – in several ways. First, we conduct a placebo test by running the same regressions from equation 1 during years when there was no policy change. If the change in the EITC is causing single mothers to increase their labor force participation, we shouldn't see a change in labor force participation in years when the policy is not changing. We limit the sample to observations in 1990 and 1991, treating 1990 as our "pre" period and 1991 as our "post" period (Columns 1 and 2 of Table 7).[21] The

---

[19] The HUD fair market rent data provides estimates the price for a two-bedroom unit from a series of separate regional surveys. The Census median rent data includes all types of rental housing, regardless of the number of rooms. Thus, the Census data may introduce variation in the median rent arising from the mix of types within the rental market while the HUD data controls for the rental size and, to some extent, the quality of the rental housing stock.

[20] These results are not included, but are available on request.

[21] Because of annual inflation-based adjustments, there are small changes to some EITC parameters every year during this period and since. Between 1990 and 1991 there were some additional changes in phase-in and phase-out rates from OBRA90, but these changes were relatively small, especially compared to those contained in OBRA93 (Table 3 and Figure 1.)

difference-in-difference estimates, as well as the heterogeneous effects, from the placebo tests are small and insignificant. In contrast, limiting the sample to 1993 and 1994 we continue to find large and significant results for our variables of interest (Columns 3 and 4 of Table 7).

As an additional robustness check we explore the differential response of families with two or more children compared to those with only one child. Changes to the EITC in the mid-1990s affected both types of families, but the changes were much larger for families with two or more kids. Between 1993 and 1995, the maximum EITC benefit rose 46 percent for one-child families, but it more than doubled for families with two or more kids, going up by 106 percent. Ceteris paribus we expect that larger families will experience greater changes in labor force participation than one-child families. In fact, results from specification (1) using families with two or more kids as the treatment group and families with only one child as the control group do suggest larger responses. Interacting the DD estimator with cost-of-living measures (similar to columns 2 and 3 in Table 5) show increases in labor force participation between 3.6 and 3.0, depending on the specification, but the results narrowly miss significance at standard levels.[22] Limiting our sample size in this way (excluding childless women reduces the sample by more than half) reduces the precision of the estimates, and makes it impossible to re-create the separate DD results by cost group as in Table 7.

Finally, we explore whether the expansion of state-level EITC policies adopted in the mid-1990s could be driving the geographic patterns in labor supply response we observe.[23] (We are already controlling for the adoption of welfare-reform waivers which vary across states and

[22] For the specification interacting the difference in difference coefficient with quarter of the cost-of-living distribution, p=.18. Expanding the sample to bring in single mothers with some college, but no degree, increases the sample size from 24,973 to 36,038, and improves the precision of the results – the standard error falls from .027 to .020 – with only a small impact on the coefficient. These results are available on request.
[23] Data on state EITC policies is from Leigh (2007) Table 2.

impact the labor force participation of single mothers.) To test the impact of these policies, we run additional regressions including a variable to reflect refundable state-level EITCs. Whether the state policies are coded as a simple dummy variable, the state credit rate, or the maximum dollar amount of the state EITC, our coefficients of interest are unaffected.[24]

Hours Results

While participation varies by cost-of-living, as we predicted, we do not have a clear prediction about a change in hours of work arising from the EITC in different MSAs. To estimate the effect on hours worked for those working, we again adopt a difference-in-difference strategy. We use the same covariates as in our participation equation but our dependent variable is hours worked last week. Our equation is:

$$\text{(2) Hours} = \alpha + \beta Z + \gamma_0 \text{ treatment} + \gamma_1 \text{ post} + \gamma_2 \text{ (treatment*post)} \\ + \gamma_3 \text{ (treatment*post*cost)} + \gamma_4 \text{ cost}$$

We drop women who did not report working last week. Our independent variables (Z) are identical to those in the participation equation and include age, age squared, number of preschool-age children, number of dependents, the number of dependents squared, an indicator for more than one child, race, MSA unemployment rate, educational attainment, and the month of implementation of welfare policy waivers. Standard errors are clustered at the MSA level.

We begin with the standard difference-in-difference strategy seen in the literature for our sample of low-educated single women. Like other work, we find no effect of the EITC on the hours worked per week in column 1 of Table 8. However, when we look at heterogeneous effects

[24] These results are not included, but are available on request. The coefficients on the state EITC covariates are uniformly negative, suggesting that states with refundable EITC policies in the mid-1990s were those with lower rates of female labor force participation.

in each cost quarter (column 2) of the distribution of housing costs with the lowest-cost quarter serving as the omitted group, we do begin to find significant responses in the lowest and highest-cost quarters. In the lowest-cost quarter, single mothers increased their hours of work by 1.3 hours per week, relative to single women without children. In the highest-cost quarter, single mothers reduced their hours of work by 0.6 hours per week, relative to single women without children.

< Table 8 approximately here >

These results demonstrate that in different MSAs women may be facing different incentives from the EITC. As in our earlier example of janitors in different cities, working single mothers in the lower-cost areas are more likely to have annual earnings that place them in the phase-in portion of the credit. The estimates suggest that in the lowest-cost areas the substitution effect outweighs any negative income effect created by the credit structure. In contrast, a working single mother in the highest-cost areas is more likely to face the high marginal tax rates arising from the phase-out of the credit. In these areas, the substitution and income effects work in concert, reducing the labor supply of working single mothers.

In Column 3, we use a continuous measure of housing costs interacted with the difference-in-difference estimator. The difference-in-difference estimator rises to an increase in weekly hours of 1.7 hours. However, each $1,000 increase in our quality adjusted rental costs reduces weekly hours by 0.4. Thus, the very highest rental cost areas behaved differently than other areas. Single mothers living in MSAs at or above the 85th percentile of costs reduce their hours of work in response to the EITC.

As in our participation estimates, we again divide the sample into four quarters based on the rental cost distribution and divide the highest quarter in half. We report the estimates from these regressions in Columns 1 through 5 of Table 9. Our results are less robust when we run sub-samples separately, possibly because of the small sample sizes for each estimate. However, we find the same pattern of results: our difference-in-difference estimates changes from a positive signed coefficients in Columns 1 through 3 to negative signed coefficients in Columns 4 and 5 when the subsample changes from below the 75th percentile and subsamples above the 75th percentile.

< Table 9 approximately here >

Although the coefficients seem small, the labor supply responses at the lower tail of the cost-of-living distribution are not trivial. For a full-time single mother in a low-cost area working full-year, our estimates suggest that there was an increase of 68 to 86 hours of work per year. In a high-cost area, single mothers may have reduced their annual hours of work between 20 to 30 hours. Additionally, our results demonstrate the need for estimating the effect in the local labor market in understanding the labor supply response to the EITC. The effect of the EITC on the intensive margin is dependent on the local wages facing potential recipients.

Robustness Checks

We check the robustness of the hours worked results to ensure our results are not driven by our methodology and that the indentifying assumptions are valid. Overall, the hours worked estimates are less robust than our participation estimates. Expanding our sample to include all employed single women, regardless of education level, provides estimates that are smaller and less significant in Columns 4 through 6 of Table 8 and Columns 6 through 10 of Table 9. This suggests that our findings of the hours response in the directions predicted are a result of selecting our sample on the low-educated. We also consider the performance of the regressions using alternative measures of housing costs. The results do not change when using either the HUD Fair Market Rent or the median rental data from the 1990 Census.

We test the identifying assumption of the difference-in-difference estimator by performing placebo tests. As in our participation results, we create two sub-samples 1990-1991 and 1993-1994 where the first year in each is our "pre" period and the second period is the "post" period. In this case, our 1990-1991 subsample and our 1993-1994 subsample each have small and insignificant results. These tests also suggest that our results are not robust.

We also estimate our equation with a Heckman selection model to correct for selecting our sample on those working last week.[25] The results from Heckman models (Table 10) provide still weaker support for the influence of cost-of-living on the impact of the EITC on hours worked. Columns 1 through 3 include second stage results of the Heckman model for low-educated women. The simple difference-in-difference (Column 1) is small and not significant. The regression including an interaction between the difference-in-difference and quarters of the rent distribution, however, does show a positive (though not significant) effect on hours worked in the lowest quarter of the rent distribution (0.2 hours per week), and a negative and significant (nearly one hour per week) effect in the highest-cost quarter. The interaction between the DD variable and a continuous measure of rent (Column 3) and the separate regressions by cost group (Columns 4 through 8) all have signs suggesting the same pattern of positive impacts on hours in

[25] The second stage of the Heckman model excludes the following variables that were included in the previous equations: the number of children under 18, the number of children under 18 squared, and an indicator for the presence of a second child.

lower-cost regions and negative impacts in higher-cost areas. None of these coefficients, however, are significant.

< Table 10 approximately here >

The lack of robustness to our hours worked results is not surprising. Almost no study has found robust effects on the hours of work decision. This could be either because of measurement error, the lack of continuous hours of work choices for low-income workers, or lack of knowledge by recipients as to how the number of hours worked translates into EITC eligibility. While we find that there is some hours worked response in our sample of low-educated single women, our precision is limited by our small sample sizes of in each MSA.

Section V: Discussion

Our results suggest that the EITC has had little impact on the labor supply of low-income single mothers in the highest cost-of-living areas such as Boston, New York City, Los Angeles, and San Francisco. Although the EITC is an important transfer to low-income *workers* in high-cost areas, the incentive is apparently insufficient to induce *non-working* single mothers in those areas to work rather than rely on the social safety net. This result is potentially a source of concern for two reasons. First, although the high-cost areas where the EITC produces no labor supply response account for 13 to 25 percent of MSAs, they represent as much as 40 percent of the total population. Second, these high-cost areas include many of the large metropolitan regions that are widely believed to have serious problems with poverty and joblessness. Whether the size of the credit is insufficient to overcome the fixed costs of work in higher cost-of-living areas, or the nationally fixed eligibility rules are incompatible with the local wage structure, or some other reason, the EITC seems to be unsuccessful at changing the labor market decisions of low-skilled non-workers in these areas.

Our findings also raise concerns regarding the welfare and efficiency impacts of the EITC. Since its inception one argument in support of the EITC has been its efficiency-enhancing properties. By offsetting relatively high taxes on the labor of low-paid workers and the steep marginal tax rates faced by those contemplating leaving public assistance, it reduces distortions in behavior (Ventry, 2001 and Hoffman and Seidman, 1990). Indeed, recent work by Eissa et al. (2008) finds that the EITC has improved welfare. Eissa et al. study a series of EITC reforms, including the reform contained in OBRA93, and evaluate welfare gains by contrasting the EITC, taking into account the interactions with other tax and transfer programs, to a lump-sum benefit. The ultimate welfare gains of the EITC result from welfare improvements along with extensive margin outweighing welfare losses along the intensive margin and those caused by the use of distortionary taxes to finance the benefit. [26]

While the welfare improvements of the EITC depend on responses along the extensive margin, our findings suggest that there is no such response in high-cost areas. Moreover, employed single mothers in high-cost areas may have reduced their hours of work in response to the policy. In sum, our findings of no change in participation and possibly fewer hours worked for those working imply welfare losses in high-cost areas. In low-cost areas, however, large increases in employment, and possible increases in hours worked, may have produced even

[26] Eissa, et al. (2008) show that calculating the welfare gains of the EITC depend on correctly measuring labor supply responses on both the intensive and extensive margins. The response along each margin is related to a different tax wedge, and impacts welfare in opposite directions. As the EITC lowers the average tax rate, employment increases, which generates a host of positive public budget externalities and increases welfare. The change in marginal tax rates, which influence the intensive margin, varies across the schedule. Overall, changes in the intensive margin are found to be welfare decreasing, as hours of work reductions (and related negative public budget externalities) along the phase-out region swamp increases along with phase-in region.

larger welfare improvements than those suggested by Eissa et al. (2008). If a large portion of the country experiences welfare losses because the program rules and benefits are not compatible with the local labor market, there would appear to be considerable room for improvement.

The imbalance in the value of the EITC between low- and high-cost regions may cause additional welfare losses, not considered by Eissa et al. (2008), by creating incentives for low-skilled workers to relocate from high-cost to low-cost metropolitan areas. Under a spatial equilibrium with geographic differences in the cost-of-living, gross wages will vary across areas for given worker types. Their real wages, however, should be equal. A major reform to the EITC, which is based on gross wages, will disturb that equilibrium and provide an incentive for lower-income households to relocate to low-cost regions. In particular, households may seek to move to a lower-cost area to realize a similar after-tax income but fewer hours devoted to work. Albouy (2008) explores similar incentives arising from federal income tax deductions and shows that the size of these distortions can be considerable. In future work, we plan to examine if the EITC induced low-skilled workers in high-cost areas to migrate to low-cost areas.

If policymakers intend to alter the labor supply decisions of low-skilled women, these conclusions are cause for concern. The appropriate policy remedy, however, is not clear. The most obvious solution to the problem identified in this paper is to determine EITC eligibility and benefit levels based on "real" (cost-of-living adjusted) dollar amounts. While straight-forward, this approach is not without problems. The EITC is already complicated to claim, which may contribute to errors in claiming the credit or reduced participation rates (Holtzblatt and McCubbin, 2004). Introducing regional differences in the federal credit could exacerbate these problems as well as stir-up political opposition among perceived "losers," an unfortunate side-effect for a program that has enjoyed considerable broad-based political support.

States could play an important role in addressing geographic imbalances. Although not very widespread during the period we study, state-level EITCs have become increasingly common.[27] By 2007, twenty one states (including the District of Columbia) had adopted refundable EITCs to supplement the federal policy and three additional states had non-refundable state EITC policies (Levitis and Koulish, 2007). While some of the higher-cost states have adopted relatively generous credit programs – the state EITC is set at 30 percent of the federal benefit in New York and 35 percent in Washington, D.C. – in most, it remains a small share of the federal credit. Many high-cost states, including California, Connecticut, and Hawaii, lack refundable EITCs. Furthermore, no state has modified its credit to adjust for cost differences within a state, which can be substantial.

Local EITCs represent another promising opportunity for addressing the issues associated with cost-of-living differences. While only a few localities have adopted supplemental EITC policies, they have been implemented in high-cost areas: New York City, San Francisco, and Montgomery County, MD (Holt, 2006). In two of these cases, the size of the local benefit is noteworthy; the local benefit in San Francisco is 16 percent of the federal credit, and in Montgomery County, MD it is 20 percent. The local credit in New York City, however, is set at only five percent of the federal EITC benefit. The merit of local EITCs is that they can provide a benefit that is more targeted to high-cost areas, but it is far from clear that local or regional governments have sufficient resources to fund EITC programs adequate to overcome the hurdles imposed by high cost-of-living.

[27] In the mid-1990s only five states had refundable EITC policies in place: Maryland, Minnesota, New York, Vermont, and Wisconsin. Our analysis generally ignores these state policies, which were quite small at the time, and with many of the highest cost areas in California, Connecticut, Hawaii, New Jersey, and Massachusetts, the pattern of state EITCs shows little relationship to a state's cost-of-living. By 2007, however, twenty states and the District of Columbia had adopted refundable EITCs to supplement the federal policy. Three additional states had EITC policies that were not refundable.

Further, insufficient purchasing power of the federal EITC benefit in high-cost areas is only part of the cost-of-living problem. Unless eligibility rules reflect local wage levels, fewer workers will be impacted in high-cost areas, workers will be treated differently by the policy depending on where they live, and incentives to relocate will remain.

Section VI: Conclusion

The federal EITC affects the labor decisions of the potentially eligible. We replicate the estimates in the literature of the intensive and extensive labor supply effects of the EITC using the 1993 expansion as a natural experiment. We contribute to the literature by taking into account local price differences. We find that the credit has differential effects across geographic areas, particularly for the participation decision. The effects of the EITC on labor market participation of single women are greatest in lower-cost areas. We demonstrate that estimates of the labor supply response to the EITC that do not account for the specific prices and local labor markets of potential beneficiaries will not fully capture the behavioral response.

We suggest that the welfare gain from the 1993 expansion is distributed unevenly across metropolitan areas. In fact, metropolitan areas with the very highest costs may have experienced a welfare loss for each EITC dollar spent while low-cost areas overwhelmingly benefited from the credit. Improved policy targeting to populations that did not benefit from the 1993 expansion may be necessary to address geographic imbalances.

APPENDIX A. Data Description

MSAs over time in the CPS

Cost-of-living and unemployment rates in this study are based on information in MSAs and non-MSA areas. There is one non-MSA area for every state, except New Jersey. Any observations which do not report a basic geographic identifier (MSA or non-MSA) are dropped from the data.

MSA definitions, which are based on at least 50,000 persons residing in a geographic area, were updated following the 1990 Census and implemented in the CPS in 1994. To link geographic units in 1994 and 1995 to the equivalent geographic units before 1994 requires constructing consistent geographic definitions over this period. The major changes to the MSA definitions include: 1) new Primary Metropolitan Statistical Areas (PMSAs) were created from several MSAs and tracking the separate MSAs that were discontinued: 2) collapsing of multiple adjacent MSAs into a single MSA; 3) creation of new MSAs out of previously non-MSA areas, and: 4) downgrading areas from MSA to non-MSA.

To make the geography in the CPS consistent between 1990 and 1995 we had to make two basic changes. For previously distinct MSAs that were consolidated into PMSAs or larger MSAs, we applied the new consolidated definition on the early 1990s geography. Any MSAs that were created out of, or returned to, non-MSA regions were classified as non-MSA in all years.

Also during this period there was a major sample redesign in the CPS. In addition to the adoption of the new MSA definitions, and changes to some of the basic questions, the CPS also changed its sample frame. Some MSAs were dropped from the survey, while others were added. Areas that were either added to or dropped from the survey are included in the non-MSA region; added regions are included in non-MSA in the mid-1990s, while dropped regions are included in non-MSA in the early-1990s when they are still in the survey.

Twenty-three MSAs are excluded in Chen and Rosenthal's data, but are included in the CPS. These few MSAs are assigned the quality-adjusted rent measure of the closest substitute – based on geographic proximity, median household income, total population, unadjusted median home price, and unadjusted median rent from the 1990 Census. MSAs with missing values (and the 'donor' MSA in parenthesis) include: Burlington, VT (from Manchester, NH); Charleston, WV (from Huntington-Ashland); Columbus, GA-AL (from Macon, GA); Evansville-Henderson, IN-KY (from Louisville, KY); Fargo-Moorhead, ND-MN (from St. Cloud, MN); Fitchburg-Leominster, MA (from Worcester, MA); Fort Smith AR-OK (from Fayetteville-Springdale, AR); Fort Walton Beach, FL (from Pensacola and Jacksonville, FL); Gadsden, AL (from Florence, AL); Huntsville, AL (from Decatur, AL); Lake Charles, LA (from Beaumont-Port-Arthur, TX); Laredo, TX (from McAllen-Edinburg-Pharr-Mission, TX); Lawton, TX (from Lubbock, TX); Naples, FL (from Palm Beach-Boca-Delray, FL); Panama City, FL (from Pensacola, FL); Portland, ME (from Manchester, NH); Portsmouth-Rochester, NH-ME (from Lawrence-Haverill MA/NH); Poughkeepsie/Dutchess, NY (from Orange, NY); Sioux City IA-NE (Waterloo-Cedar Falls, IA); Sioux Falls, SD (from Wichita, KS and Tulsa, OK); Tallahassee, FL (from Tampa-St.Petersburg-Clearwater, FL); Topeka, KS (from Kansas City, MO-KS); Wheeling, WV-OH (from Pittsburgh-Beaver Valley, PA).

Appendix B: Distribution of Housing Cost Data

	Quality-Adjusted Rental Costs	HUD Fair Market Rent: 2 Bedroom Apartment	Census Median Rent
1st Decile	1,190	425	371
2nd	1,656	449	401
3rd	2,066	477	421
4th	2,449	506	460
5th	2,775	554	501
6th	3,131	602	525
7th	3,877	635	533
8th	4,973	703	626
9th	6,771	747	671
10th	9,920	930	855

Note: The first column provides the (positive) quality-adjusted rental measures in 1990 from Chen and Rosenthal (2008). Chen and Rosenthal report quality adjusted rent in each MSA relative to the mean. To ensure all rental values are positive, we transform their data by adding $4,000 to each value. The middle column provides the 1990 HUD fair market rent values for each MSA. The HUD data provides the price for a two-bedroom unit from a series of separate regional surveys. The final column provides median rent data from the 1990 Census. The Census data includes all types of rental housing, regardless of the number of rooms. Thus, quality-adjusted rent best controls for the quality of the housing stocks by taking into account all housing characteristics. The Census data may introduce variation in the median rent arising from the mix of types within the rental market while the HUD data controls for the rental size and, to some extent, the quality of the rental housing stock.

References

Ai, Chunrong, Norton, Edward C., Wang, Hua, 2004. Computing Interaction Effects and Standard Errors in Logit and Probit Models. Stata Journal 4 (2), pp. 154-167.

Albouy, David Y., 2008. The Unequal Geographic Burden of Federal Taxation. NBER Working Paper, vol. 13995.

Berube, Alan, 2006. The New Safety Net: How the Tax Code Helped Low-Income Working Families During the Early 2000s. The Brookings Institution.

Berube, Alan, Tiffany, Thacher. 2004. The 'State' of Low-Wage Workers: How the EITC Benefits Urban and Rural Communities in the 50 States. The Brookings Institution.

Black, Dan, Kolesnikova, Natalia, Taylor, Lowell, 2007. Earnings Functions when Wages and Prices Vary by Location. St. Louis Federal Reserve BankWorking Paper, 2007-031A.

Chen, Yong, Rosenthal, Stuart, 2008. Local Amenities and Life Cycle Migration: Do People Move for Jobs or Fun?" Journal of Urban Economics, forthcoming.

Citro, Constance F., Michael, Robert T., 1995. Measuring Poverty: A New Approach. Washington, D.C.: National Academy Press.

Dickert, Stacy, Houser, Scott, Scholz, John Karl. The earned income tax credit and transfer programs: a study of labor market and program participation. In: J. Poterba, Editor, Tax Policy and the Economy (9th ed.) (1995), pp. 1–50.

Eissa, Nada, Liebman, Jeffrey B. 1996. Labor Supply Response to the Earned Income Tax Credit. Quarterly Journal of Economics 111 (2), 605-637.

Eissa, Nada, Hoynes, Hilary. Behavioral Responses to Taxes: Lessons from the EITC and Labor Supply. In: J. Poterba, Editor, Tax Policy and the Economy (20th ed.) (2006), pp. 74-110.

Eissa, Nada, Kleven, Henrik, Kreiner, Claus, 2008. Evaluation of four tax reforms in the United States: Labor Supply and welfare effects for single mothers. Journal of Public Economics 92 (3-4), 795-816.

Gelbach, Jonah, 2004. Files to calculate mean marginal effects for probit and logit models. http://www.glue.umd.edu/~gelbach/

Grogger, Jeffrey T., 2003. The Effects of Time Limits, the EITC, and Other Policy Changes on Welfare Use, Work, and Income Among Female-Headed Families. Review of Economics and Statistics 85 (2), 394-408.

Hirasuna, Donald P., Stinson, Thomas F. Stinson, 2004. Urban and Rural Differences in State Earned Income Tax Credit Programs: Minnesota's Experience. Rural Poverty Research Center Working Paper, vol. 04-08.

Hoffman, Saul D., Seidman, Laurence S., 1990. The Earned Income Tax Credit: Antipoverty Effectiveness and Labor Market Effects. Kalamazoo, MI: W.E. Upjohn Institute for Employment Research.

Holt, Steve, 2006. The Earned Income Tax Credit at Age 30: What We Know. The Brookings Institution.

Holtzblatt, Janet, McCubbin, Janet, 2004. Issues Affecting Low-Income Filers. In: Aaron, Henry J., Slemrod, Joel (Eds.) The Crisis in Tax Administration. Washington, DC: Brookings Institution, 148-187.

Hotz, Joseph V., Mullin, Charles, Scholz, John Karl, 2006. Examining the Effect of the Earned Income Tax Credit on the Labor Market Participation of Families on Welfare. NBER working paper, vol. 11968.

Hotz, Joseph V., Scholz, John Karl, 2003. The Earned Income Tax Credit. In: Moffitt, Robert (Ed.) Means-Tested Transfer Programs in the United States. Chicago: University of Chicago Press, 141-197.

Hoynes, Hilary, 2000. Local Labor Markets and Welfare Spells: Do Demand Conditions Matter? Review of Economics and Statistics 82 (3), 351-368.

Leigh, Andrew. 2005. Who Benefits from the Earned Income Tax Credit? Incidence Among Recipients, Coworkers and Firms. Research School of Social Sciences Australian National University Working Paper, vol. 494.

Levitis, Jason, Koulish, Jeremy. 2007. A Majority of States with Income Taxes Have Enacted State Earned Income Tax Credits. Center on Budget and Policy Priorities.

Maynard, Mike, Dollins, David, 2002. Participation in the Earned Income Tax Credit Program for Tax Year 1996. Research Report. SBSE Research.

Meyer, Bruce D, 2002. Labor Supply at the Extensive and Intensive Margins: The EITC, Welfare, and Hours Worked. American Economic Review 92 (9), 373-79.

Meyer, Bruce D., Rosenbaum, Dan T, 2001. Welfare, the Earned Income Tax Credit, and the Labor Supply of Single Mothers. Quarterly Journal of Economics 116 (3), 1063-1112.

Roback, Jennifer, 1988. Wages, Rents and Amenities: Differences between Workers and Regions. Economic Inquiry 26 (1), 23-41.

Rosen, Sherwin, 1979. Wage-based Indices of Urban Quality of Life. In Mieszkowski, Peter, Straszheim, Mahlon (Eds.): Current Issues in Urban Economics. Baltimore: Johns Hopkins Univ. Press, 74-104.

Scholz, John Karl, 1996. In-Work Benefits in the United States: The Earned Income Tax Credit. The Economic Journal 106: 156-169.

Ventry, Dennis J., 2001. The Collision of Tax and Welfare Politics: The Political History of the Earned Income Tax Credit. In Meyer, Bruce D., Holtz-Eakin, Douglas (Eds.): Making Work Pay: The Earned Income Tax Credit and Its Impact on America's Families. New York: Russell Sage Foundation, 15-66.

Wu, Ximing, 2005. Labor Supply and Income Effects of the Earned Income Tax Credit and Welfare Programs. Unpublished Manuscript.

Table 1. Average Hourly Wage for Specific Occupations and Industries by Rental Cost Decile (1990-93)

Decile of the Rent Distribution	Quality-Adjusted Rental Costs	Occupations: Cashier	Occupations: Secretary	Occupations: Nursing Aids, Orderlies, Attendants	Industries: Eating & Drinking Places	Industries: Hospitals	Industries: Elementary & Secondary Schools
1st	$1,190	$4.47	$6.40	$5.51	$4.12	$9.09	$6.58
2nd	1,656	4.82	7.02	5.59	4.25	9.89	6.65
3rd	2,066	4.94	6.98	6.20	4.39	9.82	8.07
4th	2,449	4.83	7.39	6.07	4.17	10.30	7.02
5th	2,775	5.30	7.59	5.78	4.52	11.00	7.27
6th	3,131	5.22	8.11	6.47	4.39	10.86	7.96
7th	3,877	5.33	8.06	6.45	4.43	11.18	8.27
8th	4,973	5.65	8.62	7.20	4.66	12.07	8.64
9th	6,771	5.90	9.34	7.57	5.13	12.40	9.68
10th	9,920	6.37	10.01	8.32	5.54	14.22	10.68

Coefficient from regression of hourly wages on annual quality-adjusted rent

All Industries & Occupations	Occupations: Cashier	Occupations: Secretary	Occupations: Nursing Aids, Orderlies, Attendants	Industries: Eating & Drinking Places	Industries: Hospitals	Industries: Elementary & Secondary Schools
0.00031***	.00031***	.00024***	.00042***	.00037***	.00019***	.00048***
(.00002)	(.00002)	(.00004)	(.00005)	(.00002)	(.00003)	(.00008)

Note: The first column provides the (positive) quality-adjusted rental measures in 1990 from Chen and Rosenthal (2008). Chen and Rosenthal report quality adjusted rent in each MSA relative to the mean. To ensure all rental values are positive, we transform their data by adding $4,000 to each value. Hourly wage, industry, and occupation data come from the 1990-1993 Current Population Survey (CPS) Outgoing Rotation Groups (ORG). Regression of hourly wages on monthly rent include controls for age, age squared, local unemployment rate, education, industry, Number of Dependents under age 5, number of dependents, welfare reform variables, year effects, and education by industry effects. All regressions are run separately by industry and occupation groups and weighted by the CPS-ORG household weight. Statistical significance is denoted as follows: * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 2. Distribution of Employed Single Women, aged 18 to 49, across EITC Schedule (1990-1993)

	Rental Cost Distribution				
	All Areas	Bottom Quarter	Second Quarter	Third Quarter	Top Quarter
Panel A. All Education Levels					
Phase-in	33.7%	45.0%	36.1%	32.5%	29.0%
Plateau	13.9%	16.8%	16.3%	13.1%	12.4%
Phase-out	26.8%	25.6%	27.1%	28.0%	26.6%
Off Schedule	25.6%	12.6%	20.4%	26.4%	32.1%
Panel B. High School Degree or Less					
Phase-in	39.7%	49.0%	41.9%	38.1%	34.9%
Plateau	17.1%	18.9%	19.3%	16.2%	15.9%
Phase-out	29.2%	25.8%	27.0%	30.6%	31.0%
Off Schedule	14.0%	6.3%	11.7%	15.2%	18.3%

Note: Data from the 1990-1993 March Current Population Survey (CPS) which, in addition to the demographic and labor market information provided in the regular monthly survey, provides additional information on prior year income and labor supply. We assign single females to each region of the schedule based on this prior year income and demographic characteristics. We include only those females who report working non-zero hours in the current year and in the prior year. The metropolitan area costs are based on the distribution of (positive) quality-adjusted rental measures in 1990 from Chen and Rosenthal (2008).

Table 3: EITC Parameters before and after Omnibus Budget Reconciliation Act (OBRA) 1993

	Phase-In Region Credit Rate	Phase-Out Region Credit Rate	Income Amounts for Plateau Region		Ending Income	Maximum benefit
			Begin Plateau	End Plateau		
Workers Without Children						
1990	0	0	0	0	0	0
1991	0	0	0	0	0	0
1992	0	0	0	0	0	0
1993	0	0	0	0	0	0
1994	7.65	-7.65	4,000	5,000	9,000	306
1995	7.65	-7.65	4,100	5,130	9,230	314
1996	7.65	-7.65	4,220	5,280	9,500	323
Workers with 1 Child						
1990	14	-10	6,810	10,730	20,264	953
1991	16.7	-11.93	7,140	11,250	21,250	1,192
1992	17.6	-12.57	7,520	11,840	22,370	1,324
1993	18.5	-13.21	7,750	12,200	23,050	1,434
1994	26.3	-15.98	7,750	11,000	23,755	2,038
1995	34	-15.98	6,160	11,290	24,396	2,094
1996	34	-15.98	6,330	11,610	25,078	2,152
Workers with 2 or more Children						
1990	14	-10	6,810	10,730	20,264	953
1991	17.3	-12.36	7,140	11,250	21,250	1,235
1992	18.4	-13.14	7,520	11,840	22,370	1,384
1993	19.5	-13.93	7,750	12,200	23,050	1,511
1994	30	-17.68	8,425	11,000	25,296	2,528
1995	36	-20.22	8,640	11,290	26,673	3,110
1996	40	-21.06	8,890	11,610	28,495	3,556

Source: Urban-Brookings Tax Policy Center. Dollar amounts unadjusted for inflation.

Table 4: Summary Statistics for Low-Educated, Single Women (Weighted)

	All Single Women	With Dependents	Without Dependents	By Rental Cost Distribution: Lowest Quarter	Second Quarter	Third Quarter	Highest Quarter
Age	31.276	31.575	31.056	31.602	30.901	31.332	31.184
	(9.568)	(7.994)	(10.573)	(9.628)	(9.563)	(0.962)	(9.492)
Less than High School	0.318	0.365	0.282	0.325	0.313	0.290	0.336
	(0.466)	(0.482)	(0.450)	(0.469)	(0.464)	(0.454)	(0.472)
High School Graduate	0.682	0.635	0.718	0.675	0.687	0.710	0.664
	(0.466)	(0.482)	(0.450)	(0.469)	(0.464)	(0.454)	(0.472)
Nonwhite	0.307	0.392	0.245	0.267	0.296	0.344	0.305
	(0.461)	(0.488)	(0.430)	(0.442)	(0.456)	(0.475)	(0.460)
Preschool Children	0.301	0.710	-	0.269	0.293	0.310	0.315
	(0.647)	(0.834)		(0.600)	(0.660)	(0.651)	(0.662)
Number of Dependents	0.779	1.836	-	0.805	0.777	0.777	0.766
	(1.134)	(1.045)		(1.124)	(1.107)	(1.126)	(1.155)
AFDC Waivers	0.311	0.301	0.318	0.237	0.118	0.213	0.495
	(0.463)	(0.459)	(0.466)	(0.425)	(0.323)	(0.410)	(0.500)
MSA Unemployment Rate	0.059	0.060	0.058	0.061	0.054	0.048	0.068
	(0.022)	(0.023)	(0.021)	(0.020)	(0.021)	(0.016)	(0.022)
(Positive) Quality-Adjusted Rental Costs	4,175.05	4,068.14	4,253.78	1,259.13	2,379.26	3,496.50	6,938.16
	(2496.40)	(2436.09)	(2537.05)	(409.84)	(242.78)	(517.79)	(1591.30)
Labor Force Participation	0.599	0.512	0.662	0.601	0.626	0.628	0.565
	(0.490)	(0.500)	(0.473)	(0.490)	(0.484)	(0.483)	(0.496)
Observations	59,708	24,973	34,735	13,347	7,213	15,406	23,742
Number of Dependents for Mothers	1.836	1.836	-	1.811	1.809	1.826	1.870
	(1.045)	(1.045)		(1.011)	(0.994)	(1.032)	(1.092)
Observations	24,973	24,973		5,860	3,029	6,516	9,568
Hours Worked Last Week, if Working	36.828	36.628	36.941	36.806	36.388	37.472	36.460
	(11.352)	(11.525)	(11.251)	(11.816)	(11.048)	(11.625)	(10.921)
Observations	36,877	12,770	24,107	8,134	4,793	10,011	13,939
Hourly Earnings, if Working and Paid	6.938	6.998	6.850	5.992	6.448	6.964	7.756
	(3.108)	(3.044)	(3.131)	(2.368)	(2.806)	(3.122)	(3.420)
Observations	17,010	5,965	11,045	3,996	2,325	4,822	5,867
Weekly Earnings, if Working	302.509	303.505	301.940	264.351	276.348	302.668	335.005
	(190.155)	(195.250)	(187.180)	(170.900)	(165.761)	(172.850)	(2.165)
Observations	35,595	12,335	23,260	7,823	4,631	9,691	13,450

Note: Authors' calculations based on the 1990-1995 monthly Current Population Survey (CPS). All summary statistics are weighted by the CPS household weight. Single women are considered low-educated if they have a high school degree or less. (Positive) quality-adjusted rental cost data are 1990 measures from Chen and Rosenthal (2008). Chen and Rosenthal report quality adjusted rent in each MSA relative to the mean. To ensure all rental values are positive, we transform their data by adding $4,000 to each value.

Table 5. Mean Marginal Effects from Probit Estimates on Labor Force Participation

	Low-educated women			All Women		
	(1)	(2)	(3)	(4)	(5)	(6)
Treatment	0.035	0.037	0.036	0.015	0.018	0.017
	(0.024)	(0.024)	(0.024)	(0.017)	(0.017)	(0.017)
Post	-0.080 ***	-0.079 ***	-0.081 ***	-0.060 ***	-0.057 ***	-0.060 ***
	(0.021)	(0.021)	(0.02106)	(0.013)	(0.013)	(0.013)
Treatment*Post	0.047 ***	0.073 ***	0.102 ***	0.031 ***	0.043 ***	0.063 ***
	(0.010)	(0.018)	(0.015)	(0.006)	(0.011)	(0.010)
Treatment*Post		0.023			0.028	
*2nd Cost Qtr		(0.027)			(0.018)	
Treatment*Post		-0.015			-0.009	
*3rd Cost Qtr		(0.026)			(0.017)	
Treatment*Post		-0.076 ***			-0.041 ***	
*4th Cost Qtr		(0.023)			(0.016)	
Treatment*Post			-0.00001 ***			-0.00001 ***
*Rental Costs			(0.000)			(0.00000)
High School Graduate	0.243 ***	0.242 ***	0.242 ***	0.177 ***	0.177 ***	0.176 ***
	(0.013)	(0.013)	(0.013)	(0.008)	(0.008)	(0.008)
Some College				0.237 ***	0.237 ***	0.237 ***
				(0.008)	(0.008)	(0.008)
College Degree				0.264 ***	0.263 ***	0.263 ***
				(0.007)	(0.007)	(0.007)
Beyond College				0.241 ***	0.240 ***	0.240 ***
				(0.006)	(0.006)	(0.006)
Nonwhite	-0.117 ***	-0.118 ***	-0.118 ***	-0.079 ***	-0.080 ***	-0.079 ***
	(0.013)	(0.013)	(0.013)	(0.008)	(0.008)	(0.008)
Age	0.024 ***	0.024 ***	0.024 ***	0.016 ***	0.016 ***	0.016 ***
	(0.004)	(0.004)	(0.004)	(0.003)	(0.003)	(0.003)
Age Squared	-0.000 ***	-0.000 ***	0.000 ***	-0.000 ***	-0.000 ***	0.000 ***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Number of	-0.114 ***	-0.116 ***	-0.11407 ***	-0.069 ***	-0.070 ***	-0.069 ***
Dependents	(0.029)	(0.030)	(0.029)	(0.022)	(0.023)	(0.022)
Dependents	0.012 ***	0.012 ***	0.012 ***	0.007 **	0.007 **	0.007 **
Squared	(0.004)	(0.004)	(0.004)	(0.003)	(0.003	(0.003)
Preschool Children	-0.088	-0.087	-0.08583	-0.082	-0.082	-0.081
	(0.009)	(0.009)	(0.009)	(0.007)	(0.007)	(0.007)
Second Child	0.033	0.034	0.032	0.011	0.012	0.010
	(0.027)	(0.028)	(0.028)	(0.019)	(0.019)	(0.019)
Unemployment Rate	-1.774 ***	-1.672 ***	-1.761 ***	-1.322 ***	-1.197 ***	-1.321 ***
	(0.262)	(0.275)	(0.261)	(0.156)	(0.168)	(0.156)
AFDC Waiver	-0.001 ***	-0.001 ***	-0.001 ***	0.007 ***	0.006 ***	0.007 ***
	(0.016)	(0.015)	(0.01529)	(0.009)	(0.009)	(0.009)
Rental Costs	-0.000		0.000	-0.000 **		-0.00000
	(0.000)		(0.000)	(0.000)		(0.00000)
Second Cost Quarter		0.009			-0.000	
		(0.020)			(0.014)	
Third Cost Quarter		0.015			0.018	
		(0.018)			(0.011)	
Fourth Cost Quarter		0.020			0.002	
		(0.020)			(0.012)	
Time Trend?	Yes	Yes	Yes	Yes	Yes	Yes
Observations	59,708	59,708	59,708	121,434	121,434	121,434

Note: Data are from the 1990-1995 monthly surveys of the Current Population Survey (CPS). The dependent variable is labor force participation. Rental cost data is the (positive) quality-adjusted rental measures in 1990 from Chen and Rosenthal (2008). Chen and Rosenthal report quality adjusted rent in each MSA relative to the mean. To ensure all rental values are positive, we transform their data by adding $4,000 to each value. Reported coefficient estimates represent the mean marginal effects. All regressions are weighted with CPS household weight. Clustered standard errors are in parentheses. Statistical significance is denoted as follows: * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 6. Mean Marginal Effects from Probit Estimates on Labor Force Participation, by Cost of Living

	Low-educated, Single Women					All Single Women				
	By Rental Cost Distribution Percentiles					By Rental Cost Distribution Percentiles				
	0-25th	25th-50th	50th-75th	75th-87th	Above 87th	0-25th	25th-50th	50th-75th	75th-87th	Above 87th
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Treatment	0.097	0.057	-0.006	0.160 ***	-0.075	-0.010	0.049	0.013	0.065	-0.037
	(0.066)	(0.086)	(0.037)	(0.051)	(0.054)	(0.057)	(0.050)	(0.025)	(0.047)	(0.037)
Post	-0.056	-0.224 ***	-0.080 **	-0.052	-0.019	-0.067 **	-0.147 ***	-0.051 *	-0.075 ***	-0.001
	(0.040)	(0.060)	(0.040)	(0.036)	(0.060)	(0.027)	(0.035)	(0.026)	(0.025)	(0.026)
Treatment*Post	0.047 **	0.063 ***	0.070 ***	0.008	0.035	0.030 **	0.059 ***	0.036 ***	0.031 **	0.011
	(0.020)	(0.024)	(0.017)	(0.021)	(0.028)	(0.012)	(0.015)	(0.013)	(0.013)	(0.016)
High School Graduate	0.293 ***	0.247 ***	0.224 ***	0.288 ***	0.174 ***	0.227 ***	0.178 ***	0.154 ***	0.203 ***	0.137 ***
	(0.024)	(0.023)	(0.028)	(0.031)	(0.021)	(0.017)	(0.014)	(0.017)	(0.019)	(0.014)
Some College						0.272 ***	0.224 ***	0.225 ***	0.260 ***	0.201 ***
						(0.018)	(0.019)	(0.014)	(0.021)	(0.013)
College Degree						0.288 ***	0.221 ***	0.244 ***	0.277 ***	0.260 ***
						(0.013)	(0.013)	(0.012)	(0.018)	(0.017)
Beyond College						0.258 ***	0.224 ***	0.230 ***	0.245 ***	0.228 ***
						(0.017)	(0.015)	(0.010)	(0.012)	(0.017)
Nonwhite	-0.102 ***	-0.105 ***	-0.140 ***	-0.138 ***	-0.081 ***	-0.086 ***	-0.081 ***	-0.087 ***	-0.087 ***	-0.054 ***
	(0.031)	(0.030)	(0.023)	(0.028)	(0.027)	(0.024)	(0.021)	(0.015)	(0.018)	(0.016)
Age	0.021 ***	0.026 ***	0.015 **	0.013 *	0.047 ***	0.013 *	0.018 ***	0.012 **	0.012 **	0.027 ***
	(0.008)	(0.009)	(0.007)	(0.006)	(0.008)	(0.007)	(0.007)	(0.005)	(0.006)	(0.005)
Age Squared	-0.000 **	-0.000 **	-0.000 *	-0.000	-0.001 ***	-0.000 *	-0.000 **	-0.000 **	-0.000 *	-0.000 ***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Number of Dependents	-0.181 **	-0.101	-0.049	-0.297 ***	-0.024	-0.047	-0.081	-0.051 *	-0.158 **	-0.015
	(0.076)	(0.106)	(0.040)	(0.075)	(0.086)	(0.066)	(0.065)	(0.030)	(0.069)	(0.055)
Dependents Squared	0.019	0.012	0.006	0.034 ***	-0.004	0.004	0.010	0.006 *	0.017	-0.003
	(0.012)	(0.015)	(0.004)	(0.010)	(0.011)	(0.010)	(0.010)	(0.004)	(0.011)	(0.008)
Preschool Children	-0.050	-0.098	-0.100	-0.061	-0.118	-0.057	-0.084	-0.081	-0.071	-0.115
	(0.018)	(0.026)	(0.018)	(0.019)	(0.020)	(0.017)	(0.019)	(0.012)	(0.012)	(0.016)
Second Child	0.096 **	0.008	-0.019	0.150 ***	-0.029	0.011	-0.015	-0.005	0.077 *	-0.015
	(0.049)	(0.083)	(0.050)	(0.057)	(0.089)	(0.044)	(0.056)	(0.030)	(0.043)	(0.049)
Unemployment Rate	-1.227 **	-2.221 ***	-1.605 ***	-1.401 ***	-2.862 ***	-1.337 ***	-1.793 ***	-1.318 ***	-0.840 ***	-1.489 **
	(0.600)	(0.645)	(0.491)	(0.317)	(0.874)	(0.297)	(0.358)	(0.344)	(0.202)	(0.595)
AFDC Waiver	0.024 ***	-0.002 ***	-0.061 ***	0.017 ***	0.032 ***	0.024 ***	0.012 ***	-0.014 ***	-0.002 ***	0.019 ***
	(0.023)	(0.039)	(0.013)	(0.020)	(0.035)	(0.015)	(0.025)	(0.010)	(0.014)	(0.021)
Time Trend?	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	13,236	7,077	14,868	10,538	13,989	22,539	14,099	31,014	22,367	31,415

Note: Data are from the 1990-1995 monthly surveys of the Current Population Survey (CPS). The dependent variable is labor force participation. Rental cost data is the (positive) quality-adjusted rental measures in 1990 from Chen and Rosenthal (2008). Chen and Rosenthal report quality adjusted rent in each MSA relative to the mean. To ensure all rental values are positive, we transform their data by adding $4,000 to each value. Reported coefficient estimates represent the mean marginal effects. All regressions are weighted with CPS household weight. Clustered standard errors are in parentheses. Statistical significance is denoted as follows: * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 7. Mean Marginal Effects from "Placebo" Regressions on Labor Force Participation

	Low-Educated Women							
	1990 to 1991				1993 to 1994			
	(1)		(2)		(3)		(4)	
Treatment	0.028		0.029		0.023		0.025	
	(0.020)		(0.020)		(0.059)		(0.059)	
Post	0.011		0.009		-0.052	***	-0.051	***
	(0.008)		(0.008)		(0.011)		(0.011)	
Treatment*Post	-0.015		-0.020		0.031	**	0.044	**
	(0.012)		(0.018)		(0.014)		(0.021)	
Treatment*Post *2nd Quarter			0.016				0.061	*
			(0.025)				(0.033)	
Treatment*Post *3rd Quarter			0.025				-0.015	
			(0.020)				(0.032)	
Treatment*Post* 4th Quarter			-0.016				-0.049	*
			(0.020)				(0.027)	
High School Graduate	0.235	***	0.235	***	0.257	***	0.256	***
	(0.011)		(0.011)		(0.017)		(0.017)	
Nonwhite	-0.126	***	-0.126	***	-0.108	***	-0.109	***
	(0.014)		(0.014)		(0.016)		(0.016)	
Age	0.032	***	0.032	***	0.025	***	0.024	***
	(0.003)		(0.003)		(0.005)		(0.006)	
Age Squared	-0.000	***	-0.000	***	-0.000	***	-0.000	***
	(0.000)		(0.000)		(0.000)		(0.000)	
Number of Dependents	-0.096	***	-0.097	***	-0.086		-0.088	
	(0.022)		(0.023)		(0.073)		(0.074)	
Dependents Squared	0.009	***	0.009	***	0.008		0.008	
	(0.003)		(0.003)		(0.011)		(0.012)	
Preschool Children	-0.119	***	-0.120	***	-0.080	***	-0.079	***
	(0.007)		(0.007)		(0.012)		(0.012)	
Second Child	0.032	*	0.033	*	0.014		0.014	
	(0.018)		(0.018)		(0.052)		(0.052)	
Unemployment Rate	-3.079	***	-2.796	***	-4.068	***	-4.024	***
	(0.545)		(0.572)		(1.005)		(1.012)	
AFDC Waiver					0.002		0.002	
					(0.016)		(0.015)	
Rental Costs	-0.000	*			0.000			
	(0.000)				(0.000)			
Second Cost Quarter			0.010				-0.039	
			(0.017)				(0.026)	
Third Cost Quarter			-0.001				0.007	
			(0.015)				(0.023)	
Fourth Cost Quarter			-0.015				0.018	
			(0.016)				(0.022)	
Month Effects?	Yes		Yes		Yes		Yes	
Obs.	25,762		25,762		23,401		23,401	

Note: Data are from the 1990, 1991, 1993, and 1994 monthly surveys of the Current Population Survey (CPS). The dependent variable is labor force participation. Rental cost data is the (positive) quality-adjusted rental measures in 1990 from Chen and Rosenthal (2008). Chen and Rosenthal report quality adjusted rent in each MSA relative to the mean. To ensure all rental values are positive, we transform their data by adding $4,000 to each value. Reported coefficient estimates represent the mean marginal effects. All regressions are weighted with CPS household weight. Clustered standard errors are in parentheses. Statistical significance is denoted as follows: * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 8: Difference-in-Difference Estimates on Hours worked per Week for Single Women, Conditional on Working

	Low-educated women			All Women		
	(1)	(2)	(3)	(4)	(5)	(6)
Treatment	-0.368	-0.323	-0.327	-0.189	-0.182	-0.219
	(0.872)	(0.874)	(0.880)	(0.615)	(0.617)	(0.613)
Post	-0.777	-0.760	-0.775	-0.071	-0.007	-0.050
	(0.687)	(0.686)	(0.687)	(0.464)	(0.456)	(0.462)
Treatment*Post	0.359	1.254 **	1.726 ***	0.317	0.688 *	0.215
	(0.379)	(0.600)	(0.630)	(0.243)	(0.369)	(0.408)
Treatment*Post *2nd Quarter		-1.121			-1.476 **	
		(0.956)			(0.701)	
Treatment*Post *3rd Quarter		-0.406			-0.288	
		(0.864)			(0.562)	
Treatment*Post* 4th Quarter		-1.900 **			-0.214	
		(0.832)			(0.509)	
Treatment*Post* Rental Costs			-.00035 **			0.000
			(.00015)			(0.000)
High School Graduate	2.306 ***	2.297 ***	2.279 ***	2.302 ***	2.312 ***	2.292 ***
	(0.343)	(0.341)	(0.340)	(0.337)	(0.339)	(0.339)
Some College				2.989 ***	2.970 ***	2.981 ***
				(0.357)	(0.357)	(0.359)
College Degree				4.844 ***	4.833	4.857
				(0.397)	(0.398)	(0.398)
Beyond College				6.465 ***	6.464 ***	6.489 ***
				(0.510)	(0.513)	(0.513)
Nonwhite	-0.330	-0.344	-0.344	-0.331	-0.347	-0.320
	(0.335)	(0.332)	(0.336)	(0.229)	(0.229)	(0.229)
Age	1.102 ***	1.097 ***	1.091 ***	1.083 ***	1.086 ***	1.088 ***
	(0.131)	(0.130)	(0.131)	(0.102)	(0.102)	(0.102)
Age Squared	-0.014 ***	-0.014 ***	-0.014 ***	-0.014 ***	-0.014 ***	-0.014 ***
	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
Number of Dependents	-0.415	-0.408	-0.407	-0.743	-0.741	-0.737
	(0.981)	(0.986)	(0.986)	(0.702)	(0.704)	(0.700)
Dependents Squared	0.009	0.007	0.008	0.021	0.021	0.020
	(0.123)	(0.125)	(0.123)	(0.098)	(0.099)	(0.098)
Preschool Children	-0.120	-0.123	-0.112	-0.255	-0.267	-0.256
	(0.319)	(0.322)	(0.319)	(0.237)	(0.239)	(0.239)
Second Child	-1.032	-1.070	-1.079	-0.336	-0.347	-0.344
	(1.016)	(1.016)	(1.018)	(0.645)	(0.644)	(0.645)
Unemployment Rate	-36.560 ***	-34.12 ***	-35.88 ***	-32.38 ***	-27.64 ***	-30.49 ***
	(7.968)	(9.246)	(8.177)	(5.672)	(6.697)	(5.893)
AFDC Waiver	-0.090	-0.113	-0.045	-0.304	-0.282	-0.197
	(0.327)	(0.350)	(0.340)	(0.268)	(0.283)	(0.274)
Rental Costs	-0.003		0.000	-0.007 *		-0.000 ***
	(0.005)		(0.000)	(0.004)		(0.000)
2nd Cost Qtr		-0.225			0.105	
		(0.664)			(0.477)	
3rd Cost Qtr		0.357			0.308	
		(0.514)			(0.381)	
4th Cost Qtr		0.412			-0.490	
		(0.527)			(0.398)	
Constant	19.047 ***	18.521 ***	18.736 ***	19.706 ***	18.942 ***	19.647 ***
	(2.157)	(2.123)	(2.171)	(1.673)	(1.658)	(1.660)
Time Trend?	Yes	Yes	Yes	Yes	Yes	Yes
Observations	36,877	36,877	36,877	88,440	88,440	88,440

Note: Data are from the 1990-1995 monthly surveys of the Current Population Survey (CPS). The dependent variable is hours worked last week. Rental cost data is the (positive) quality-adjusted rental measures in 1990 from Chen and Rosenthal (2008). All regressions are weighted with CPS household weight. Clustered standard errors are in parentheses.

Table 9: Difference-in-Difference Estimates on Hours worked per Week for Single Women, Conditional on Working, by Cost of Living

	Low-educated, Single Women					All Single Women				
	By Rental Cost Distribution Percentiles					By Rental Cost Distribution Percentiles				
	0-25th	25th-50th	50th-75th	75th-87th	Above 87th	0-25th	25th-50th	50th-75th	75th-87th	Above 87th
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Treatment	-1.928	-2.745	-0.365	3.592 *	1.724	0.156	-0.645	-1.511	3.116 ***	0.937
	(3.530)	(3.112)	(1.267)	(2.107)	(1.577)	(2.033)	(1.701)	(0.955)	(1.153)	(1.537)
Post	1.502	-1.375	-0.176	-5.186 **	-0.149	1.618	0.249	1.080	-2.044 *	-1.308
	(1.316)	(1.526)	(1.381)	(2.052)	(1.246)	(1.257)	(1.120)	(0.865)	(1.111)	(0.916)
Treatment*Post	0.910	0.507	1.344 *	-0.131	-1.751 *	0.641	-0.920	0.681	0.061	0.222
	(0.668)	(1.021)	(0.676)	(0.843)	(1.035)	(0.387)	(0.750)	(0.455)	(0.482)	(0.676)
High School	2.645 ***	3.698 ***	2.920 ***	1.351 *	0.957	2.556 ***	3.794 ***	3.000 ***	1.360 *	1.106
Graduate	(0.727)	(0.760)	(0.664)	(0.780)	(0.812)	(0.723)	(0.766)	(0.667)	(0.774)	(0.860)
Some College						3.229 ***	4.278 ***	3.516 ***	2.785 ***	1.421 *
						(0.662)	(0.879)	(0.667)	(0.864)	(0.743)
College Degree						5.860 ***	5.882 ***	5.770 ***	3.457 ***	3.509 ***
						(0.862)	(1.010)	(0.698)	(1.081)	(0.896)
Beyond						6.964 ***	10.199 ***	6.172 ***	6.807 ***	4.752 ***
College						(0.918)	(1.697)	(0.987)	(0.865)	(0.935)
Nonwhite	-0.570	-1.990 ***	-0.056	0.029	0.270	0.048	-1.081	-0.374	0.077	-0.229
	(0.754)	(0.725)	(0.687)	(0.790)	(0.656)	(0.590)	(0.663)	(0.472)	(0.463)	(0.440)
Age	0.950 ***	1.146 ***	1.060 ***	1.464 ***	1.046 ***	1.184 ***	1.378 ***	0.914 ***	1.416 ***	0.854 ***
	(0.295)	(0.300)	(0.236)	(0.336)	(0.271)	(0.221)	(0.231)	(0.191)	(0.236)	(0.209)
Age Squared	-0.011 **	-0.015 ***	-0.013 ***	-0.019 ***	-0.013 ***	-0.015 ***	-0.018 ***	-0.011 ***	-0.019 ***	-0.010 ***
	(0.004)	(0.004)	(0.003)	(0.005)	(0.004)	(0.003)	(0.004)	(0.003)	(0.003)	(0.003)
Number of	1.012	1.308	0.180	-5.020 **	-2.580	-1.615	-0.971	1.114	-4.060 ***	-2.187
Dependents	(3.832)	(3.805)	(1.534)	(2.315)	(1.917)	(2.094)	(2.215)	(1.068)	(1.241)	(1.986)
Dependents	-0.358	-0.140	-0.074	0.572 **	0.768 **	-0.029	0.278	-0.182 *	0.479 ***	0.312
Squared	(0.526)	(0.509)	(0.116)	(0.218)	(0.358)	(0.297)	(0.333)	(0.094)	(0.175)	(0.397)
Preschool	0.127	0.713	-0.280	-0.235	-0.909	0.072	0.606	-0.272	-0.931 **	-0.841
Children	(0.662)	(0.866)	(0.716)	(0.658)	(0.628)	(0.460)	(0.650)	(0.506)	(0.417)	(0.702)
Second Child	0.824	-1.663	-2.994	2.094	-2.527	1.927	-0.788	-2.549 **	1.504	0.419
	(2.433)	(3.046)	(2.121)	(2.724)	(1.717)	(1.343)	(2.084)	(1.164)	(1.305)	(1.172)
Unemployment	-16.320	-56.280 *	-66.966 ***	-23.960	-29.162	-22.393 *	-28.169	-44.084 ***	-24.016 **	-29.691 *
Rate	(13.987)	(31.613)	(17.007)	(16.810)	(23.261)	(11.888)	(18.010)	(14.389)	(11.543)	(17.520)
AFDC Waiver	0.272	0.073	-0.485	0.110	-0.178	-0.112	-0.015	-0.878 **	-0.247	0.118
	(0.653)	(1.862)	(0.404)	(0.730)	(0.563)	(0.539)	(0.903)	(0.404)	(0.679)	(0.448)
Constant	21.232 ***	18.305 ***	22.408 ***	8.283 *	20.101 ***	18.066 ***	12.796 ***	23.809 ***	11.779 ***	22.350 ***
	(5.355)	(5.213)	(3.806)	(4.430)	(3.747)	(3.807)	(4.150)	(2.996)	(2.958)	(3.594)
Time Trend?	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs.	8068	4704	9660	6418	8027	15614	10632	23481	16285	22428
R^2	0.04	0.07	0.06	0.07	0.05	0.06	0.08	0.05	0.07	0.05

Note: Data are from the 1990-1995 monthly surveys of the Current Population Survey (CPS). The dependent variable is hours worked last week. Rental cost data is the (positive) quality-adjusted rental measures in 1990 from Chen and Rosenthal (2008), who report quality adjusted rent in each MSA relative to the mean. To ensure all rental values are positive, we transform their data by adding $4,000 to each value. All regressions are weighted with CPS household weight. Clustered standard errors are in parentheses. Statistical significance is denoted as follows: * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 10: Estimates from Heckman Models of Hours Worked Last Week for Low-educated Women

	All Areas			By Rental Cost Distribution Percentiles				
				0-25th	25th-50th	50th-75th	75th-87th	Above 87th
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Treatment	-1.053 ***	-1.061 ***	-1.061 ***	-0.599 *	-1.766 **	-1.097 ***	-0.645	-1.372 ***
	(0.193)	(0.193)	(0.194)	(0.312)	(0.711)	(0.350)	(0.462)	(0.416)
Post	-0.747 ***	-0.737 ***	-0.752 ***	-0.581	-1.028	-0.688	-1.152 **	-0.406
	(0.247)	(0.246)	(0.246)	(0.514)	(0.732)	(0.494)	(0.466)	(0.537)
Treatment*Post	-0.171	0.203	0.279	-0.213	0.305	0.303	-0.448	-0.837
	(0.282)	(0.397)	(0.404)	(0.486)	(0.813)	(0.492)	(0.642)	(0.732)
Treatment*Post*2nd Quarter		-0.486						
		(0.601)						
Treatment*Post*3rd Quarter		-0.136						
		(0.542)						
Treatment*Post*4th Quarter		-0.796 *						
		(0.454)						
Treatment*Post*Rental Costs			0.000					
			(0.000)					
Preschool Children	-0.318 *	-0.342 **	-0.323 *	-0.307	-0.038	-0.036	-0.742 *	-0.516
	(0.173)	(0.173)	(0.172)	(0.348)	(0.678)	(0.266)	(0.428)	(0.414)
Age	1.062 ***	1.066 ***	1.062 ***	1.031 ***	1.251 ***	1.024 ***	0.887 ***	1.100 ***
	(0.054)	(0.055)	(0.055)	(0.132)	(0.149)	(0.108)	(0.111)	(0.100)
Age Squared	-0.013 ***	-0.013 ***	-0.013 ***	-0.013 ***	-0.016 ***	-0.013 ***	-0.011 ***	-0.014 ***
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)
High School Grad	2.289 ***	2.319 ***	2.295 ***	2.261 ***	2.009 ***	3.141 ***	1.997 ***	1.751 ***
	(0.220)	(0.219)	(0.217)	(0.471)	(0.540)	(0.407)	(0.592)	(0.404)
Nonwhite	-0.759 ***	-0.769 ***	-0.748 ***	-0.699 *	-2.040 ***	-0.914 ***	-0.764 *	0.082
	(0.163)	(0.164)	(0.164)	(0.378)	(0.494)	(0.231)	(0.459)	(0.322)
Unemployment Rate	-25.44 ***	-23.87 ***	-25.72 ***	-20.30 ***	-37.49 ***	-29.92 ***	-16.88 **	-21.88 **
	(3.43)	(3.82)	(3.55)	(6.22)	(11.97)	(11.47)	(7.34)	(8.56)
Rent Percentile	0.007 **			-0.017	-0.037	0.033	-0.036	-0.020
	(0.003)			(0.024)	(0.040)	(0.026)	(0.046)	(0.048)
AFDC Waiver	-0.011	0.008	0.021	0.408	-0.514	0.023	-0.069	0.262
	(0.199)	(0.215)	(0.207)	(0.479)	(0.611)	(0.303)	(0.557)	(0.353)
Second Cost Quarter		0.035						
		(0.349)						
Third Cost Quarter		0.769 ***						
		(0.250)						
Fourth Cost Quarter		0.592 **						
		(0.250)						
Rental Costs			0.000 *					
			(0.000)					
Constant	18.378 ***	18.080 ***	18.515 ***	18.518 ***	17.875 ***	17.506 ***	24.427 ***	19.854 ***
	(0.973)	(0.998)	(0.963)	(2.120)	(2.834)	(2.444)	(4.827)	(5.058)
Time Trend	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs.	59,708	59,708	59,708	13,236	7,077	14,868	10,538	13,989

Note: Data are from the 1990-1995 monthly surveys of the Current Population Survey (CPS). Results are from the second stage of the Heckman Model. Variables excluded from the second stage of the Heckman Model include Number of Dependents under 18, the Number of Dependents squared, and an indicator variable for the second child. The dependent variable is hours worked last week. Rental cost data is the (positive) quality-adjusted rental measures in 1990 from Chen and Rosenthal (2008). All regressions are weighted with CPS household weight. Clustered standard errors are in parentheses.

Figure 1. EITC Schedule: 1990, 1993, and 1996
Head of Household Filer with 2 Children