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THE INTERACTION OF METROPOLITAN COST-OF-LIVING AND THE FEDERAL EARNED INCOME TAX CREDIT: ONE SIZE FITS ALL?

Katie Fitzpatrick and Jeffrey P. Thompson

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 an 8 percentage point increase in labor force participation for single mothers in the lowest cost areas but no discernible 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 the cost-of-living. In addition, our findings suggest that the EITC does little to reduce joblessness in many of the nation's largest cities.

Keywords: EITC, labor supply, housing costs, cost-of-living differentials

JEL Codes: H31, J22, R23

"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."

Alan Berube (Berube, 2006, p. 1)

I. INTRODUCTION

Policymakers intend for the EITC to reward work by altering labor supply incentives. Estimates of the impact of the EITC on labor supply find large positive effects on the decision to work but no effect on the decision of how much to work (Hotz and Scholz, 2003; Eissa and Hoynes, 2006). Previous studies, however, fail to address the influence of geographic differences in both wages and cost-of-living. Cost-of-living

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differences make the credit more (or less) valuable across geographic areas. With a nationally uniform *benefit structure*, the EITC is more valuable in an 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 for specific worker-types, gross wages vary considerably.[1] EITC eligibility based on gross income, therefore, results in variation in EITC benefits across local areas. In general, low-skilled workers in high-cost areas earn higher gross wages. These workers are more likely to end up on the phase-out portion of the EITC schedule or off the schedule completely than similar workers in low-cost areas. As a result, these workers earn different credit amounts depending on where they live.

Local costs are critical to analyzing the EITC because earnings and bundles of consumption goods and services are realized in specific local labor markets. We address differences across local areas by including a measure of location-specific prices — housing costs in the Metropolitan Statistical Area (MSA) — to examine the effect of the EITC across local areas. Using the 1993 EITC expansion, we find that the labor supply response to the EITC depends on housing costs in the local labor market. In the lowest-cost areas, the EITC expansion resulted in an increase in labor force participation among single mothers of up to 8 percentage points. In contrast, we find no significant increase in participation for single mothers in the highest-cost metropolitan areas, where nearly 40 percent of the population lives. We find some evidence of differences in the number of hours worked across local areas, but the results are not robust.

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 results. Section V discusses the implications of the findings, and Section VI concludes.

II. THEORY AND LITERATURE

A. The EITC and Labor Supply

The structure of the EITC (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. In the phase-out region, the credit decreases with each additional dollar of earned income until completely eliminated.

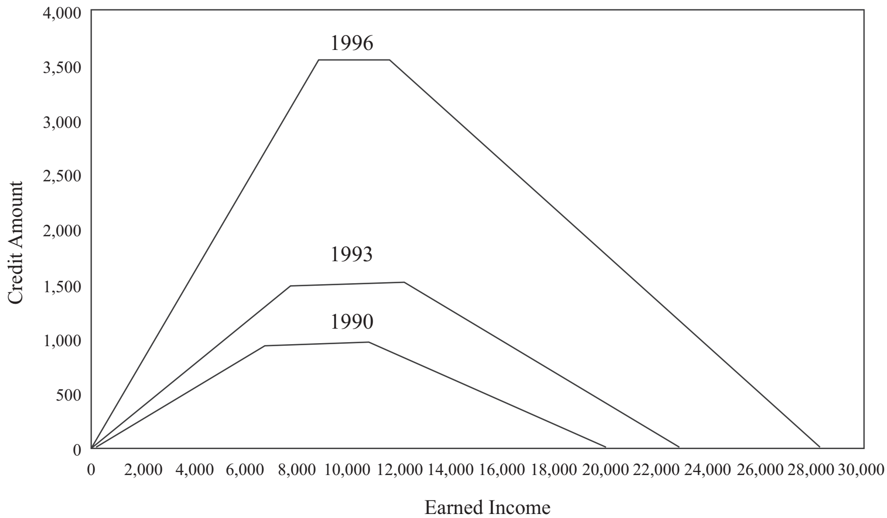
In the standard static model of labor supply, the EITC shifts out the budget constraint and provides unambiguously positive incentives to work.[2] This shift, however, creates EITC-induced kinks which make the expected effect on hours worked dependent on the worker’s placement on the credit schedule. For a worker in the phase-in region, theory

[1] Albouy (2009a) — 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.

[2] This is the case for single women. Married couples face more complex participation decisions.

Figure 1

EITC Schedule: 1990, 1993, and 1996
Head of Household Filer with Two Children



provides no clear prediction due to offsetting income and substitution effects. In contrast, workers in the plateau and phase-out regions of the credit receive a disincentive to work additional hours from the EITC, assuming leisure is a normal good.[3] The overall effect of the EITC on hours, therefore, depends empirically on the distribution of beneficiaries across the schedule and the relative magnitudes of the income and substitution effects.

B. Regional Differences in Wages and the Cost-of-Living

The EITC is expected to have different impacts on labor supply across areas due to variation in wages and the cost-of-living, particularly the geographic variation in housing costs. The causes and consequences of this variation have been the subject of considerable interest, both in the economics literature and in policy debates.[4]

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[3] At income levels above the phase-out region, the EITC may induce a reduction in hours in order to receive a credit.

[4] Examples in the economics literature include Rosen (1979), Roback (1988), Glaeser (1998), Beeson (1991), Hoynes (2000), and Moretti (2009). In policy debates, the best example is the National Academy of Science (NAS) study on the poverty threshold that recommended that the threshold reflect differences in housing and other prices across geographic areas (Citro and Michael, 1995). This recommendation was not ultimately adopted for a variety of reasons, including measurement problems, lack of data, and political constraints.

We find a relationship between wages and cost-of-living, using earnings data from the 1990–1993 Current Population Survey (CPS) and 1990 quality-adjusted rental cost data from Chen and Rosenthal (2008) to proxy for cost-of-living.[5] Panel A of Table 1 provides the average hourly wages of single, low-educated, female household heads (age 18 to 50) in industries and occupations employing the greatest number of single women, in each of five portions of the quality-adjusted rental cost distribution. For nursing aids, wages vary from $5.72 in the lowest-cost area to $7.83 in the highest-cost area, which includes MSAs above the 87th percentile of rental costs.[6] Similarly, wages in eating and drinking establishments vary from $4.30 to $5.61. The estimates in Panel B of Table 1 suggest that, after controlling for demographic and labor market characteristics, every thousand dollar increase in quality-adjusted annual rents is associated with $604 in higher annual earnings. Separate regressions by occupation and industry groups yield similar results.

The relationship between local prices and wages is established in the now-standard spatial equilibrium model developed by Rosen (1979) and Roback (1982), and applied and extended by others, including Beeson and Eberts (1989), Gyourko and Tracy (1989), Rauch (1993), Glaeser and Saiz (2004), Rosenthal and Strange (2004), Shapiro (2006), Albouy (2009a, 2009b), Deitz and Abel (2008), Moretti (2009), and Black, Kolesnikova, and Taylor (2009). The basic intuition of the model is that households (firms) value the bundle of amenities associated with a place in addition to traded and non-traded goods (inputs). With perfect mobility, households (firms) choose their utility maximizing (unit cost minimizing) location, and, in equilibrium, will be indifferent across locations. Households and firms will face an array of possible locations with different combinations of amenities, but compensating differences in wages and rents will leave marginal households and firms equally well-off across their alternatives.

The way in which amenities are capitalized into wages and rents depends on whether the amenity is a *consumption* amenity (valued primarily by households) or a *productive* one (valued primarily by firms). Consumption amenities such as ocean views attract people, raise the supply of labor, and reduce wages. Productive amenities, such as a moderate climate which reduces heating and cooling costs, attract firms, raise the demand for labor, and increase wages. Both types of amenities, however, result in higher rental costs. Consumption amenities boost household demand for housing; productive amenities increase demand for business locations.

These same basic relationships between amenities, wages, and local prices continue to hold in the models that extend the Roback (1982) framework to include heterogeneous

[5] The Chen and Rosenthal data, discussed in more detail later, uses a hedonic regression to determine quality-adjusted housing costs in each MSA.

[6] We split the top quarter of the rent distribution into two groups to make groups that contain roughly the same number of observations and have similar variances. The upper tail of the distribution has considerably higher rental costs than those in the 75th to 87th percentile.

Table 1

Relationship between Wages and Annual Rental Costs for Low-Educated Single Women, by Selected Industries and Occupations

Panel A. Average Hourly Wage for Specific Occupations and Industries by Percentile of Rental Cost Distribution (1990–1993)

			Industries			Occupations		
Percentile of Rental Cost Distribution	Annual Rental Costs	All Industries & Occupations	Eating & Drinking Establishments	Hospitals	Elementary & Secondary Education	Secretaries	Cashiers	Nursing Aide/ Orderly
Below 25th percentile	$2,902	$7.46	$4.30	$7.68	$6.42	$6.99	$4.79	$5.72
25th–50th percentile	$4,013	$7.74	$4.56	$8.32	$7.01	$7.75	$5.03	$6.03
50th–75th percentile	$5,071	$8.11	$4.70	$8.36	$7.80	$8.27	$5.22	$6.50
75th–87th percentile	$6,823	$8.69	$5.14	$9.33	$8.18	$9.06	$6.08	$7.29
Above 87th percentile	$9,719	$9.09	$5.61	$9.80	$8.91	$10.07	$6.16	$7.83

Panel B. Coefficient from Regression of Annual Earnings on Annual Quality-adjusted Rent by Selected Industries and Occupations

		Industries			Occupations		
	All Industries & Occupations	Eating & Drinking Establishments	Hospitals	Elementary & Secondary Education	Secretaries	Cashiers	Nursing Aide/ Orderly
	604.4***	599.2***	658.1**	767.6**	722.8***	306.2***	901.8***
	(79.94)	(59.01)	(154.9)	(343.2)	(142.6)	(93.31)	(204.5)

Notes: The first column in Panel A uses 1990 quality-adjusted annual rental costs based on data from Chen and Rosenthal (2008). Hourly wage, industry, and occupation data in Panel A come from the 1990–1993 Current Population Survey (CPS) Outgoing Rotation Groups (ORG). Panel B uses data from the March CPS from 1990–1993. Regressions of annual earnings on annual quality-adjusted rent in Panel B 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. Low-educated women are defined as those with a high school degree or less. All regressions are run separately by industry and occupation groups and weighted by the CPS household weight. Asterisks denote significance at the 1% (***), 5% (**), and 10% (*) levels.

labor inputs.[7] And, the pattern observed in Table 1, where low-skilled workers are paid more in high cost-of-living locations, is consistent with either productive amenities or luxury consumption amenities.[8] Under either explanation, equivalent low-skill workers will have equal utility across space, with some residing in high-wage, high-cost areas and others in low-wage, low-cost areas.

C. The EITC, Labor Supply, and Cost-of-Living

This geographic variation in wages and cost-of-living implies that low-skilled workers will face different EITC treatment based on where they live. To show this before the 1993 EITC expansion, we use data from the March CPS in Table 2 and compare the incomes of employed single women to the EITC schedule across MSAs. In the early 1990s, Panel A shows that only 7.8 percent of low-educated single females in MSAs below the 25th percentile of rental costs had incomes too high to be on the EITC schedule compared to 24.5 percent in the highest-cost areas. 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 may be larger.

Panel B of Table 2 shows the differences in estimated EITC benefit levels across MSAs. The average benefit in areas below the 25th percentile of rental costs was $732 in 1992, while the average benefit was $703 in the highest-cost areas. Adjusting these benefit levels to reflect real purchasing power suggests even larger differences. Using MSA-level quality-adjusted rent as a proxy for regional price differences, the "real" average EITC benefit in the lowest quartile was between $1,434 and $811, depending on whether all prices or only housing prices vary across areas.[9] Real EITC benefits in areas above the 87th percentile of the rental costs were between $407 and $613. Similarly, the real increases in the maximum EITC benefit were greater in lower-cost areas, as

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7 These models include those of Roback (1988), Beeson (1991), Black, Kolesnikova, and Taylor (2009), and Moretti (2009). With two skill-types and variation in amenities across cities, the models presented by Roback (1988), Black, Kolesnikova, and Taylor (2009), and Lee (2009) show that if low-skill types are imperfect substitutes for high-skill types, and 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 maintain constant utility across cities. The willingness of high-skill types to pay for the amenity raises rent in the high-amenity cities which must be offset through higher wages if low-skill types are to reside in high-amenity cities.

8 For work emphasizing productive amenities, see Mare and Glaeser (2001), Gabriel and Rosenthal (2004), Albouy (2009b), and Moretti (2009); for consumption amenities, see Roback (1988), Black, Kolesnikova, and Taylor (2009), and Lee (2009).

9 The extent to which prices other than housing vary across MSAs is a subject of debate (Rauch, 1993) not resolved here. We calculate these two different estimates to provide a potential range of price variation; variation is much lower when we assume that only housing prices vary. The adjustment which assumes all prices vary across MSAs simply divides the nominal EITC benefit by the ratio of the area average rent level to the national average rent level. The second approach, which assumes that only rent varies, uses an index that adds 0.80 (roughly the non-shelter share of average household total expenditures in 1992 based on the Consumer Expenditure Survey) to 0.2 times the ratio of area rental costs divided by the national average rental costs.

Table 2

Distribution of Low-Educated Employed Single Women, age 18 to 50, across EITC Schedule and EITC Benefit Level by Percentile of Rental Cost Distribution (1990–93)

	All Areas	By Percentile of Rental Cost Distribution: Below 25th Percentile	25th–50th Percentile	50th–75th Percentile	75th–87th Percentile	Above 87th Percentile
Panel A. Region of EITC Schedule						
Phase-in region	38.5%	45.6%	41.6%	37.4%	36.4%	32.2%
Plateau region	16.6%	19.3%	18.3%	15.3%	16.2%	14.6%
Phase-out region	28.6%	27.2%	26.6%	30.0%	29.6%	28.7%
Off the EITC schedule	16.4%	7.8%	13.5%	17.2%	17.8%	24.5%
Panel B. Average 1992 Federal EITC Benefit						
Nominal average EITC benefit	$716	$732	$719	$715	$701	$703
Real average EITC benefit						
Partial price variation		$811	$763	$728	$670	$613
Full price variation		$1,434	$1,007	$785	$570	$407
Panel C. Change in real EITC maximum benefit (1993 to 1995) by number of children and cost-of-living						
One Child	$660					
Partial price variation		$732	$700	$672	$631	$576
Full price variation		$1,293	$924	$724	$537	$382
Two Children	$1,559					
Partial price variation		$1,728	$1,653	$1,587	$1,491	$1,361
Full price variation		$3,054	$2,183	$1,710	$1,269	$902

Notes: Data are from the March Current Population Survey (CPS). The March CPS contains information on prior year income and demographics, which we use to assign single females to each region of the EITC benefit schedule. We include only those females who report working non-zero hours in the current year and in the prior year. The rental cost distribution is based on 1990 quality-adjusted annual rental costs from Chen and Rosenthal (2008). EITC benefit levels in Panel B use the variable in the March CPS calculated by the Census Bureau, based on earnings, income, household composition and other demographic details. The real EITC benefit adjusts the nominal levels in two ways. The first approach assumes that only rent varies, and that all other prices are equal across areas ("Partial Price Variation"). This adjustment uses an index that adds 0.80 (roughly the non-shelter share of average household total expenditures in 1992 based on the Consumer Expenditure Survey) to 0.2 times the ratio of area rental costs divided by the national average rental costs. The first way assumes that all prices vary across areas in the same way as quality-adjusted rent ("Full Price Variation"), and divides the nominal EITC benefit by the ratio of the area average rent level to the national average rent level.

shown in Panel C. For example, if all prices vary across MSAs, then the real increase in maximum EITC benefit for families with two or more children was $3,054 in the lowest-cost areas, compared to just $902 in the highest-cost ones.

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. First, workers in low-cost areas have lower wages and, therefore, have annual incomes more likely to be income-eligible for the EITC. Second, once income-eligible, workers in low-cost areas will, on average, receive larger benefits because their income places them on the upper end of the phase-in or on the plateau region of the schedule, rather than the phase-out region. Finally, EITC benefits to a worker in a low-cost area have greater purchasing power — 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 local costs will affect the hours decision due to 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 to the policy change. With a larger share of workers falling in the EITC phase-in region, low-cost areas should be more likely to have positive responses on the hours 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 regions of the credit makes it difficult to make strong predictions about the hours response to the EITC. In short, we expect the hours worked decision to be less responsive to the cost-of-living than the participation decision.

D. Previous EITC Literature

A large literature on 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. They find that the expansion 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 hours worked.

The findings of Eissa and Liebman are consistent with other methodologies, samples, and expansion period, including a panel dataset of California welfare recipients (Hotz, Mullin, and Scholz, 2006), models including welfare use (Grogger, 2003), simulation studies (Dickert, Houser, and Scholz, 1995; Scholz, 1996), and structural models (Meyer and Rosenbaum, 2001). Almost no study finds a substantial change in the hours worked of recipients, a result Eissa and Liebman (1996) attribute to labor market norms and institutions which allow for only part-time or full time work, measurement error, and/or a lack of knowledge about the exact structure of the EITC.[10]

[10] 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 worked.

To our knowledge, only one other study explores the influence of cost-of-living on the labor supply response to the EITC.[11] In one portion of their analysis, Meyer and Rosenbaum (2000) compare changes in the employment rates of single mothers in low-cost states with single mothers in high-cost states from 1984 to 1996. They conclude that single mothers in low-cost states, as measured by a state-level housing cost index, have larger increases in their labor force participation. While they find the expected sign, they state that the large magnitude of their 4 to 5 percentage point relative increase suggests that the effect is not due solely to costs and taxes.

We extend their analysis with an improved measure of the prices individuals face. Meyer and Rosenbaum's state-level data understate the differences that exist across local areas. In our data, a state-level measure of quality-adjusted rent suggests a difference between high-cost and low-cost states of $3,300 but an MSA-level measure suggests a difference between high-cost and low-cost areas of $10,000.[12] Moreover, intrastate variation is as important as interstate variation. For example, both Oregon and Pennsylvania have state average quality-adjusted rent of approximately $4,700. The difference between high-cost areas and low-cost areas in Pennsylvania is $4,550 and $1,050 in Oregon.

III. METHODOLOGY AND DATA SOURCES

A. 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, steadily increasing benefit generosity in 1994, 1995, and 1996.

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 likely to have earnings in the credit range; single mothers are the largest group of workers eligible for the EITC; and, unmarried women allow us to avoid intra-household bargaining decisions that affect married women. While OBRA93 extended EITC eligibility to those without children, the credit is quite small and available only to those with 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

[11] Eissa and Hoynes (forthcoming) note the distinctive geographic distribution of the EITC at the state level.

[12] The state-wide average quality-adjusted rents here are based on MSA-level rents from Chen and Rosenthal (2008) and population distribution and weights in the 1990–1993 CPS.

women without children. In robustness checks, we estimate models that compare the response of single mothers by family size.

Unlike previous work, we allow for heterogeneous effects across local areas by interacting our cost-of-living measure with the difference-in-difference estimator. Our coefficient of interest is this heterogeneous effect. This is not the standard triple-difference estimator because the addition of the cost-of-living variable does not provide an additional control group. Therefore, we do not interact the treatment and time variables with the cost-of-living variables. Instead, the interaction with the cost-of-living and the difference-in-difference estimator allows us to explore differential responses across areas.

B. Data

We use the 1991–1995 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.[14] For each tax unit, we merge on MSA-level unemployment rates and a cost-of-living measure.[15] For those tax units residing outside of an MSA, we use the state's non-MSA values.

We limit our sample to single (never married, widowed, or divorced) women who are heads of tax units, ages 16 to 50. We drop women with more than a high school degree. We also drop the self-employed, unpaid agriculture workers, and workers 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. Our focus on geographic variation also forces us to drop observations without a basic geographic identifier (MSA or state non-MSA).

An ideal gauge of differences in the cost-of-living across local areas would contain a composite measure of the expenditure an individual would need to maintain a constant utility level when purchasing a basket of goods, services, and local amenities. Such a measure, unfortunately, does not exist. Instead, we measure the local

[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.

[14] While our unit of analysis is the tax unit, in the regression results we continue to use the household weights provided by the Census Bureau. We expect this decision will have little impact on our results, as in most cases the two units are identical. We also find that our results are robust to using weighted or unweighted regressions.

[15] Details on the creation of MSAs that are consistent over the 1990 to 1995 period are available from the authors.

cost-of-living with quality-adjusted annual rental cost data provided by Chen and Rosenthal (2008). While not a perfect measure of the local cost-of-living, it provides a number of advantages. First, because housing expenses are typically the largest budgetary item for a family, the data allow us to account for a large portion of the budgetary needs across local areas. Second, the data allows us to control for differences in housing quality across areas. Finally, assuming the value of local amenities is capitalized into rental prices, the data allows us to control for differences in local characteristics.

The Chen and Rosenthal data are created by estimating hedonic regressions, controlling for the characteristics of housing units in each MSA, with data from the Census. From these estimates, they report housing costs in 1990 for each MSA and non-MSA relative to a national mean. As a result, their range of housing costs range 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 back the mean quality-adjusted annual rent for 1990 ($5,659) to each value. The new range of quality-adjusted rent, which we refer to as annual rental costs, is $1,874 to $11,811.[16]

Table 3 presents summary statistics of the characteristics of our full sample, as well as our treatment and control groups in Columns 1–3 and across the rental cost distribution in Columns 4–8. 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 local costs. Despite living in MSAs with similar unemployment rates, single mothers have much lower levels of labor force participation. Conditional on employment, however, their hours of work are similar.

Comparing across areas, all areas have similar percentages of low-educated single women with a high school degree. Women in the lowest-cost areas are less likely to be nonwhite. 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 tend to have lower levels of labor force participation.

Women in different areas of the rental cost distribution appear roughly comparable in the number of children they have. Mothers in areas below the 25th percentile of rental costs have, on average, 1.82 children. In the areas above the 87th percentile of rental costs, mothers have on average 1.89 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 wage rates and hours worked.

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[16] The full listing of MSA and state non-MSA rental costs is available from the authors.

Table 3

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

	All Women	With Dependents	Without Dependents	By Percentile of Rental Cost Distribution: Below 25^{th} Percentile	25^{th}–50^{th} Percentile	50^{th}–75^{th} Percentile	75^{th}–87^{th} Percentile	Above 87^{th} Percentile
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Age	31.28	31.79	30.89	31.65	30.87	31.39	30.87	31.35
	(9.568)	(8.105)	(10.526)	(9.630)	(9.565)	(9.657)	(9.480)	(9.457)
High school graduate	0.682	0.638	0.716	0.673	0.688	0.710	0.676	0.659
	(0.466)	(0.481)	(0.451)	(0.469)	(0.463)	(0.454)	(0.478)	(0.474)
Nonwhite	0.307	0.393	0.242	0.262	0.298	0.341	0.320	0.304
	(0.461)	(0.488)	(0.428)	(0.440)	(0.457)	(0.474)	(0.467)	(0.460)
AFDC waivers	0.311	0.300	0.319	0.238	0.112	0.188	0.476	0.516
	(0.461)	(0.458)	(0.466)	(0.426)	(0.316)	(0.391)	(0.499)	(0.500)
MSA unemployment rate	0.059	0.060	0.058	0.062	0.054	0.047	0.069	0.066
	(0.022)	(0.023)	(0.021)	(0.020)	(0.020)	(0.013)	(0.031)	(0.016)
Annual rental costs (in $thousands)	5.834	5.733	5.911	2.902	4.013	5.071	6.823	9.719
	(2.496)	(2.433)	(2.540)	(0.401)	(0.249)	(0.462)	(0.619)	(0.939)
Worked last week	0.599	0.517	0.661	0.599	0.627	0.631	0.572	0.564
	(0.490)	(0.500)	(0.473)	(0.490)	(0.484)	(0.483)	(0.495)	(0.496)
Number of dependents, for mothers		1.840		1.819	1.811	1.819	1.859	1.890
		(1.047)		(1.021)	(0.977)	(1.039)	(1.062)	(1.107)
Number of preschool children, for mothers		0.698		0.595	0.683	0.705	0.752	0.757
		(0.832)		(0.774)	(0.865)	(0.831)	(0.867)	(0.826)
Hours worked last week, if working	36.83	36.72	36.89	36.77	36.51	37.59	36.32	36.46
	(11.35)	(11.51)	(11.25)	(11.90)	(10.99)	(11.51)	(11.26)	(10.86)
Observations	59,708	25,457	34,251	13,236	7,077	14,868	10,538	13,989

Notes: Authors' calculations from the 1991–1995 Current Population Surveys (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. Rental costs are based on Chen and Rosenthal's (2008) 1990 quality-adjusted annual housing measures.

IV. RESULTS

A. Participation Estimates

1. Primary Results

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

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

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.[17] 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, the number of dependents squared, an indicator for more than one child, race, MSA unemployment rate, and educational attainment. We also control for the implementation of AFDC policy waivers.[18] Standard errors are clustered at the MSA level.

We present estimates of the mean marginal effects from our probit regressions in Table 4. The first column provides the estimate of the effect of the EITC on labor market participation, without addressing geographical differences in the cost-of-living. 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.[19] Controls for education, age, race, and the local unemployment rate have the expected sign.[20]

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[17] The interaction terms in a probit model are not straightforward to interpret. The coefficient on the interaction does not simply capture the marginal effect, but also includes additional terms that are conditional on the interacted variables as well as other independent variables. We used the "inteff procedure," written for Stata by Ai, Norton, and Wang (2004), to obtain correct marginal effects and standard errors for our difference-in-difference variable in the probit equation. These results were nearly identical to the results obtained from calculating the mean marginal probit effects using Gelbach's (2004) "margfx procedure," as well as results from LPM models.

[18] We do not account for the specific provisions of each waiver. Instead, the waiver variable is intended to correct for the changing options that a single mother faces when choosing whether or not to work.

[19] This estimate is larger than the roughly three percent participation increase estimated by Meyer and Rosenbaum (2001) for the 1993 expansion. We found estimates similar to theirs when we used their sample selection criteria.

[20] The variable *Post* is negative and significant, which is not surprising given the "jobless recovery," which saw falling employment rates even among prime-age white males, following the July 1990 to March 1991 recession.

Table 4

Mean Marginal Effects from Probit Estimates on Labor Force Participation for Low-Educated Single Women

				By Percentile of Rental Cost Distribution				
	All Areas	Rental Cost Dummies	Continuous Rental Costs	Below 25th Percentile	25th–50th Percentile	50th–75th Percentile	75th–87th Percentile	Above 87th Percentile
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Treatment	0.037	0.038	0.037	0.102	0.052	–0.023	0.186***	–0.104*
	(0.029)	(0.030)	(0.029)	(0.069)	(0.106)	(0.049)	(0.059)	(0.061)
Post	–0.080***	–0.080***	–0.081***	–0.057	–0.223***	–0.081**	–0.054	–0.022
	(0.021)	(0.021)	(0.021)	(0.039)	(0.060)	(0.040)	(0.035)	(0.059)
Treatment*Post	0.047***	–0.004	0.122***	0.047**	0.062***	0.070***	0.009	0.036
	(0.010)	(0.020)	(0.018)	(0.020)	(0.024)	(0.017)	(0.021)	(0.028)
Treatment*Post*Below 25th Percentile		0.078***						
		(0.023)						
Treatment*Post* 25th–50th Percentile		0.089***						
		(0.024)						
Treatment*Post* 50th–75th Percentile		0.079**						
		(0.022)						
Treatment*Post* 75th–87th Percentile		–0.005						
		(0.023)						
Treatment*Post*Annual rental costs (in $thousands)			–0.014***					
			(0.003)					
Below 25th Percentile		–0.017						
		(0.024)						
25th–50th percentile		–0.001						
		(0.025)						
50th–75th percentile		–0.006						
		(0.023)						
75th–87th percentile		0.023						
		(0.023)						
Annual rental costs (in $thousands)	–0.003		0.002	–0.004	–0.065	–0.013	–0.022	–0.006
	(0.003)		(0.003)	(0.032)	(0.050)	(0.020)	(0.021)	(0.012)
Observations	59,708	59,708	59,708	13,236	7,077	14,868	10,538	13,989

Note: Authors' calculations from the 1991-1995 CPS. The dependent variable equals one if the respondent reported working last week. Other covariates include age, age squared, number of preschool age children, number of dependents, the number of dependents squared, an indicator for more than one child, nonwhite, MSA unemployment rate, and an indicator if the respondent graduated from high school. Rental costs are based on Chen and Rosenthal's (2008) 1990 quality-adjusted annual housing measures. Reported coefficient estimates represent the mean marginal effects. All regressions are weighted with the CPS household weight. Standard errors, clustered by MSA, are reported in parentheses. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) levels.

We explore whether the EITC participation effect differs systematically by local areas in Column 2. We begin with dichotomous variables representing MSA cost-of-living in each of the five areas, omitting the highest cost area. The point estimates on the heterogeneous effect are 7.8, 8.9, and 7.9 percentage points in the areas below the 25^{th} percentile, areas in the 25^{th} to 50^{th} percentile, and 50^{th} to 75^{th} percentile of rental costs, respectively. These estimates are the pattern of responses that we expected — the lowest cost areas have the larger effects. It is surprising that the point estimates in the highest cost areas are negative, although these estimates are not statistically significant from zero and we cannot rule out a small positive effect.

To use the full variation in costs we interact the difference-in-difference estimator with the continuous measure of costs in Column 3. Each $1,000 increase in annual rental costs reduces participation by one percentage point. While the lowest cost areas experience an 8 percentage point increase, we again find no significant increase in participation in the highest-cost areas.

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.[21] We rerun (1) for each area and report the estimates in Columns 4–8 of Table 4. The results show larger effects in the lower cost areas than higher cost areas. The second (Column 5) and third (Column 6) quarter of costs have the largest and most significant effects: an increase in employment of 6.2 and 7.0 percentage points, respectively. The first quarter (Column 4) has a smaller response, with an increase of 4.7 percentage points. In the highest areas (Columns 7 and Columns 8), the EITC has no significant effect on participation. The participation effect in the upper end of the rental cost distribution is significantly different from the lower cost areas.[22]

2. *Robustness Checks*

We perform several tests to assess the robustness of our findings. We first exploit the larger benefits the 1993 expansion directed at families with two or more children, compared to those with only one child. We find a similar pattern of results, but the results narrowly miss significance at conventional levels.[23] The small sample sizes reduce the precision of our estimates and prevent estimation within each area.

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[21] The summary statistics in Table 4 demonstrate 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. A Wald test strongly rejects pooling ($p<0.01$) of these local areas.

[22] The point estimates for each of the lower three cost areas are not significantly different from one another at the 10 percent level. The MSAs at the 75^{th} to 87^{th} percentile, however, are statistically different from the lower cost areas. The highest cost area is not significantly different from the lower three cost areas. However, when we break up the distribution into quartiles, rather than five areas, the top quartile is significantly different from the other quartiles.

[23] These results are provided in Appendix A. To do this analysis, we expand the sample to single mothers with less than a college degree. These results suggest that only in areas below the 25^{th} percentile of rental costs is there a positive effect on participation.

We also expand our sample to all women, regardless of education level for each specification. Again, our estimates show the same pattern as our baseline estimates but with smaller magnitudes. 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. Both measures suggest the same magnitude and pattern of results as our quality-adjusted rental cost data.[24] In addition, we explored using wages as a way to reflect cost-differences across areas, and regressions interacting MSA-level wages with the difference in difference variable yield very similar results.

Finally, we consider whether state EITC policies adopted in the mid-1990s account for the geographic patterns we observe. We run specifications with a covariate reflecting refundable state EITCs. Whether the state policies are coded as a dummy variable, the state credit rate, or the maximum dollar amount of the state EITC, our coefficients of interest are unaffected.[25]

B. Hours Results

1. Primary Results

While participation varies by cost-of-living, as we predicted, we do not have a clear prediction how the effect of the EITC on hours worked may differ across areas. To estimate the effect on hours worked for those working, we again adopt a difference-in-difference strategy. Our estimating equation is:

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

We use the same covariates as before and estimate (2) with ordinary least squares (OLS).[26] Because our dependent variable is hours worked last week, we drop women who did not report working last week. Standard errors are clustered at the MSA level.

We first estimate (2) without accounting for local cost differences across areas. Similar to other results in the literature, we find no effect of the EITC on the hours worked per week in Column 1 of Table 5. Upon interacting dichotomous variables representing

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[24] The HUD fair market rent data estimates the price of 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 introduces 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.

[25] 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.

[26] Our main results use an OLS specification that does not correct for selection into the labor market. In robustness checks, we estimate a Heckman selection model to account for the endogeneity of the participation decision.

Table 5

Difference-in-Difference Estimates on Hours Worked per Week for Low-Educated Single Women, Conditional on Working

				By Percentile of Rental Cost Distribution				
	All Areas	Rental Cost Dummies	Continuous Rental Costs	Below 25th Percentile	25th–50th Percentile	50th–75th Percentile	75th–87th Percentile	Above 87th Percentile
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Treatment	–1.010	–0.953	–0.999	–4.265	1.016	–1.644	4.839	9.042
	(1.113)	(1.123)	(1.115)	(4.594)	(3.316)	(1.723)	(3.176)	(6.178)
Post	–0.763	–0.712	–0.765	1.630	–1.362	–0.140	–5.219**	–0.199
	(0.689)	(0.686)	(0.691)	(1.337)	(1.523)	(1.381)	(2.064)	(1.222)
Treatment*Post	0.358	–1.464	2.303***	0.940	0.548	1.326**	–0.118	–1.755*
	(0.379)	(0.930)	(0.848)	(0.664)	(0.996)	(0.678)	(0.858)	(1.029)
Treatment*Post*Below 25th Percentile		2.661**						
		(1.079)						
Treatment*Post* 25th–50th Percentile		1.635						
		(1.166)						
Treatment*Post* 50th–75th Percentile		2.484**						
		(1.091)						
Treatment*Post* 75th–87th Percentile		1.679						
		(1.167)						
Treatment*Post*Annual Rental Costs (in $thousands)			–0.352**					
			(0.153)					
Below 25th percentile		–0.528						
		(0.632)						
25th–50th percentile		–0.609						
		(0.745)						
50th–75th percentile		–0.066						
		(0.615)						
75th–87th percentile		–0.459						
		(0.587)						
Annual rental costs (in $thousands)	–0.052		0.064	0.206	–1.344	–0.133	0.119	–0.284
	(0.065)		(0.076)	(0.704)	(1.782)	(0.488)	(0.663)	(0.357)
Observations	36,877	36,877	36,877	8,068	4,704	9,660	6,418	8,027

Note: Authors' calculations from the 1991–1995 CPS. Sample only includes those who report working last week. The dependent variable is number of hours worked last week. Other covariates include age, age squared, number of preschool age children, number of dependents, the number of dependents squared, an indicator for more than one child, nonwhite, MSA unemployment rate, and an indicator if the respondent graduated from high school. Rental costs are based on Chen and Rosenthal's (2008) 1990 quality-adjusted annual housing measures. All regressions are weighted with the CPS household weight. Standard errors, clustered by MSA, are reported in parentheses. Asterisks denote significance at the 1% (***), 5% (**), and 10% (*) levels.

each cost area in Column 2, however, we find significant responses in the lowest cost areas and areas just above the 50th percentile. In areas below the 25th percentile of rental costs, single mothers show a relative increase of 2.66 hours worked per week; in areas between the 50th and 75th percentile of rental costs, single mothers increase their hours by 2.48. In other areas, however, there is no significant effect on hours worked.[27]

In Column 3, we interact the difference-in-difference estimator with the continuous measure of costs. The difference-in-difference estimator rises to an increase in weekly hours of 2.30 hours. However, each $1,000 increase in annual rental costs reduces weekly hours by 0.4. Thus, the very highest cost areas behaved differently than other areas. The predicted effect of EITC on hours worked changes sign at or above the 85th percentile of costs.

Finally, we run (2) within each cost area. These results, in Columns 4–8 of Table 5, do not consistently reach statistical significance, possibly because of the small sample sizes for each estimate. However, the same pattern of results emerges with the difference-in-difference estimates changing from a positive signed coefficient in the lower cost areas to negative signed coefficients above the 75th percentile of the distribution.

In total, these results imply that women may be responding to labor supply incentives in different regions of the credit schedule: in the lowest-cost areas the substitution effect outweighs any negative income effect, while employed women in the highest-cost areas reduce their hours working in response to the high marginal tax rates in the phase-out of the credit. Moreover, the labor supply responses are not trivial, particularly at the lower tail of the distribution. For a full-time single mother working full-year in a low cost area, our estimates suggest an increase of 49 to 62 hours of work per year; in high-cost areas, our estimates suggest a reduction in 76 to 94 hours of work a year.

2. Robustness Checks

Overall, the hours estimates are less robust than the participation estimates. Expanding our sample to include all employed single women, regardless of education level, provides estimates that have the same pattern, although smaller and less precise. Using a Heckman selection model to correct for selecting our sample on those working last week provides still weaker support for the influence of local costs on EITC-induced changes in hours worked.[28] The only robustness check that clearly corroborates our main results is using either HUD Fair Market Rent data or median rental data from the 1990 Census rather than the Chen and Rosenthal data.

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[27] The point estimate for the areas above the 87th percentile of rental costs is negative but misses conventional significance levels (p=0.11).

[28] The second stage of the Heckman model excludes the following variables that were included in the previous equations: the number of dependents, the number of dependents squared, and an indicator for the presence of a second child. The point estimates from the Heckman model suggest a positive impact on hours in low cost areas and negative impacts in high cost areas but no estimate is statistically significant.

The lack of robustness of these results is not surprising. Almost no study has found a strong effect of the EITC on hours worked. This could be for any of the reasons Eissa and Liebman (1996) suggest: measurement error, inability of low-income workers to incrementally change their hours of work, or lack of knowledge by recipients of the structure of the EITC.

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 discernible labor supply response account for 13 to 25 percent of MSAs, these areas represent as much as 40 percent of the total population. Second, these high-cost areas include many of the large metropolitan areas 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 areas with higher local costs, 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.[29]

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; Hoffman and Seidman, 1990). Indeed, recent work by Eissa, Kleven, and Kreiner (2008) finds that the EITC has improved welfare. They 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 from the EITC are due to welfare improvements along the extensive margin outweighing welfare losses along the intensive margin,

[29] The patterns of participation responses by cost-of-living that we observe do not necessarily imply different labor supply elasticities of low-skilled worker in different areas, and are plausibly driven by the cost-of-living differences we describe. The real increase in the maximum benefit following the 1993 policy change was more than three times larger (assuming full price variation) in low-cost areas than in high-cost ones, and the related wage differences across areas result in workers in high-cost areas being less likely to even get the maximum benefit (Table 2). And while the interaction regressions (Columns 2–3) in Table 4 suggest no response in the highest-cost areas, the separate regression by cost area (Columns 4–8) suggest positive, but small, changes.

less the welfare losses arising from the use of distortionary taxes to finance the benefit.[30]

While the welfare improvements of the EITC depend on responses along the extensive margin, our findings suggest that no measurable response occurred 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 significant change in participation and possibly fewer hours worked for those working imply welfare losses in high-cost areas. In low-cost areas, however, large employment increases, and possibly increases in hours worked, may have produced even larger welfare improvements than suggested by Eissa, Kleven, and Kreiner (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 appears to be considerable room for improvement.

The imbalance in the value of the EITC between low- and high-cost areas may cause additional welfare losses, not considered by Eissa, Kleven, and Kreiner (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 (2009a) 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. Policy responses are available at the federal and the state and local levels, although, each approach has some drawbacks. 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. This approach, though, could make claiming EITC more complicated. There are already concerns that the current policy is too complicated, contributing to errors in claiming the credit and possibly reducing

[30] Eissa, Kleven, and Kreiner (2008) show that calculating the welfare gains of the EITC depends 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.

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.

Some economists, notably Glaeser (1998), also express another concern with proposals to adjust transfer payments by local cost of living. If poor people choose to live in high-cost areas (large cities) because of the area's amenities, then the higher costs they face do not imply lower levels of well-being (Glaeser, Kahn, and Rappaport, 2008). If, in fact, poor households in low-cost areas have lower real incomes, they will have a greater marginal utility of income, and welfare can be increased more by boosting transfer payments in low-cost areas (Glaeser, 1998). Glaeser does, however, note some conditions under which adjustment is desirable. If local amenities and income are complements, as in the luxury amenity stories of Black, Kolesnikova, and Taylor (2009) and Lee (2009), then adjusting transfer payments for local costs is welfare maximizing, as the marginal utility of income will be higher in high-cost regions. Indexing also raises welfare more when it is targeted at poor families with children — the same families that receive the greatest EITC benefits — because these families are more risk averse. Finally, as noted by Kaplow (1995) in considering whether to adjust taxes or benefits to reflect local costs, potential welfare gains need to be considered alongside the efficiency losses resulting from migration from high-cost to low-cost areas if the benefit levels are not adjusted for local costs.[31]

State and local governments could play a potentially important role in addressing geographic imbalances, but this approach is also not without problems. Although not very widespread during the period we study, state-level EITCs have become increasingly common.[32] By 2009, twenty-two states (including the District of Columbia) adopted refundable EITCs to supplement the federal policy and two states had non-refundable state EITC policies (Williams, Johnson, and Shure, 2009). Washington State is the most recent to adopt a refundable EITC and is the first state without an income tax to do so. But, 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

[31] The EITC population and program may be sufficiently different from the traditional public welfare program to justify indexation, despite these concerns. For one thing, as shown in Table 1, even for low-skilled workers there is a considerable wage premium for living in high-cost areas, suggesting that urban consumption amenities are not the predominant location factor. Because it is a work incentive program, the welfare implications of adjusting EITC benefits for local costs may be different than for traditional transfer payments. The EITC arguably increases welfare by overcoming existing disincentives to work and reducing other public benefit payments, as described by Eissa, Kleven, and Kreiner (2008). If high-cost areas have higher wages, higher public transfer benefit levels, and higher rates of utilization of transfer programs, then the employment increase resulting from EITC indexation will likely be welfare enhancing.

[32] 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.

40 percent in Washington, DC — 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 adjusted for intrastate cost differences, which can be substantial.

While only a few localities have supplemental EITC policies, they have been implemented in high-cost areas: New York City, San Francisco, and Montgomery County, MD (Holt, 2006). In one case, the size of the local benefit is noteworthy; the benefit in Montgomery County it is set equal to the state's refundable matching rate, currently 25 percent. The credit in New York City, however, is only five percent of the federal EITC benefit. The merit of local EITCs is that they can provide a benefit targeted to high-cost areas, but it is far from clear that local governments have sufficient resources to fund EITC programs adequate to overcome the hurdles imposed by high cost-of-living. In fact one local EITC initiative, in Denver, was suspended due to insufficient resources, and another, in San Francisco, has been seriously scaled back.[33]

Further, insufficient purchasing power of the federal benefit in high-cost areas is only part of the cost-of-living problem facing the EITC. 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.

VI. CONCLUSION

The large literature looking at the labor supply effects of the EITC overlooks geographical differences in the cost-of-living. We contribute to the literature by specifically accounting for local price differences in the labor supply response. Using the 1993 EITC expansion as a natural experiment, we find that the credit has differential effects across geographic areas, particularly for the participation decision. We conclude that the effects of the EITC on the labor supply of single mothers 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.

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[33] A local credit was adopted in Denver, Colorado in 2002, but was allowed to expire just two years later as the funding source — TANF block grants — was insufficient to sustain the credit. The matching rate for the San Francisco credit was initially set at 10 percent of the federal credit, but in 2007 was converted to a flat benefit of $100 per filing family.

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Appendix A

Mean Marginal Effects from Probit Estimates on Labor Force Participation for Single Mothers with Less than a College Degree

	All Areas	Rental Cost Dummies	Continuous Rental Costs
	(1)	(2)	(3)
Two or more children	–0.120***	–0.120***	–0.120***
	(0.008)	(0.008)	(0.008)
Post	–0.056**	–0.051**	–0.056**
	(0.022)	(0.029)	(0.022)
Two or more children*Post	0.009	–0.002	0.035
	(0.012)	(0.028)	(0.029)
Two or more children*Post*Below 25th Percentile		0.061*	
		(0.032)	
Two or more children*Post* 25th–50th Percentile		–0.026	
		(0.046)	
Two or more children*Post* 50th–75th Percentile		0.004	
		(0.031)	
Two or more children*Post* 75th–87th Percentile		–0.002	
		(0.033)	
Two or more children*Post*Annual rental costs (in $thousands)			–0.005
			(0.005)
Below 25th Percentile		–0.034	
		(0.098)	
25th–50th Percentile		0.050	
		(0.085)	
50th–75th Percentile		0.032	
		(0.072)	
75th–87th Percentile		0.008	
		(0.050)	
Annual rental costs (in $thousands)	–0.014***	–0.012	–0.011***
	(0.003)	(0.012)	(0.004)
Observations	36,944	36,944	36,944

Note: Authors' calculations from the 1991–1995 CPS. The dependent variable equals one if the respondent reported working last week. Other covariates include age, age squared, nonwhite, MSA unemployment rate, and educational attainment fixed effects. Reported coefficient estimates represent the mean marginal effects. All regressions are weighted with the CPS household weight. Standard errors, clustered by MSA, are reported in parentheses. Statistical significance is as follows: * 10%, ** 5%, and *** 1%.