The EITC over the Great Recession: Who benefited?☆

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Abstract

In this paper, I examine the impact of the Great Recession on Earned Income Tax Credit (EITC) eligibility. Because the EITC is structurally tied to earnings, the direction of this impact is not obvious. Families who experience complete job loss for an entire tax year lose eligibility, while those experiencing underemployment (part-year employment, a reduction in hours, or spousal unemployment in married households) may become eligible. Determining the direction and magnitude of the impact is important for a number of reasons. The EITC has become the largest cash-transfer program in the U.S., and many low-earning families rely on it as a means of support in tough times. The program has been considered a replacement for welfare, enticing former welfare recipients into the labor force. Recent research has begun to focus on the effectiveness of the EITC during a time when jobs are scarce; however, due to data limitations, the mechanism of EITC eligibility loss has not been addressed. To answer this question, I use the 2006 Current Population Survey Annual Social and Economic Supplement matched to tax data from 2005 through 2011 to examine changes in eligibility experienced by individuals over time. In assessing three competing causes of eligibility loss, I find that less-educated, unmarried women experienced a greater hazard of loss due a yearlong lack of earnings compared with other labor-market groups. Meanwhile, this same group was less likely than other groups to gain eligibility through underemployment. Not only did many families headed by unmarried, low-skilled women lose all earnings income during the recession, but they also lost credit dollars that averaged approximately $2000 per tax year. I discuss the implications of these findings on the view of the EITC as a safety-net program.

Keywords: EITC; Unemployment; Great Recession; Safety Net

JEL: J64, I38, 32, H24

☆The views expressed in this paper are those of the authors and do not necessarily represent the views of the U.S. Census Bureau.I would like to thank Steven Martin and Hilary Hoynes for helpful comments on an earlier draft.

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1. INTRODUCTION

After the welfare reform era of the 1990s, the Earned Income Tax Credit (EITC) emerged as the largest cash-transfer program in the United States (Bitler and Hoynes, 2010). It is a tax credit that is paid to filers as part of a tax refund, and its receipt is dependent on three key prerequisites: a recipient must have earned income in the tax year; he or she must file a federal income tax return; and he or she must specifically file for the credit. While there are other eligibility requirements that come into play, these three rules must be met before receipt occurs. The first prerequisite—earnings—is strictly outlined in the tax code. Those who file for and receive the EITC are required to have income that was earned in the tax year in question, such as wage and salary earnings from an employer or self-employment earnings.

Thus, both a strength and a weakness of the EITC is its tie to work. While the program has been shown to encourage work among those who formerly made up welfare rolls (Hotz and Scholz, 2006; Meyer and Rosenbaum, 2001), its usefulness as the key component of the social safety net has been questioned due to its focus on labor force participants (Williams and Maag, 2008; Moffitt, 2013). For single earners incapable of finding work over a tax year, the program provides zero assistance. On the other hand, in cases of underemployment for both married and unmarried earners—or for families where one spouse has become unemployed but the other remains working—the credit may provide crucial financial help.

Recent research has looked into the question of unemployment and the EITC (for example, Bitler et al., 2014). These studies have relied on caseload data, and have used the state unemployment rate as a source of identification. However, while point-in-time estimates of the correlation between unemployment and caseloads are useful, they are not able to identify the direction of the effect of unemployment on eligibility loss or gain. This is a direct result of the EITC's program parameters, which require that earnings be greater than zero but less than a threshold. Panel data on individuals is necessary to isolate the causes of eligibility loss and compare the incidence of these causes across labor market groups.

In the current work, I examine the dynamics of employment and EITC eligibility over the Great Recession using a panel data set of linked Current Population Survey Annual Social and Economic Supplement (CPS ASEC) and Internal Revenue Service (IRS) data. The main research question I answer is: for which labor market groups were unemployment and loss of EITC eligibility co-determined. My hypothesis is that groups who have traditionally had a more tenuous attachment to the labor market experienced a greater incidence of eligibility loss through a complete lack of employment earnings than did groups with a more stable attachment. I use a definition of labor market groups that is well established in the literature, and look separately at groups defined by marriage, sex, and educational attainment.

I focus on the population that is determined to be eligible based on a methodology developed to create the IRS's official EITC take-up rates. This methodology uses the CPS ASEC linked at the individual level to data from W-2s and 1040s supplied by the IRS. Compared with caseload data, this population captures labor-market group characteristics, such as education, sex, and marriage, that are strongly correlated with labor market attachment.

In what follows, I identify three risks of eligibility loss based on the program parameters: yearlong lack of earnings; earnings or income above the program's thresholds; and family changes. The competing risks analysis captures what happened to eligible filers over time, identifying how eligibility spells ended for filers first observed in the 2006 CPS ASEC. I find that unmarried women with low educational attainment (a high school diploma or less) were more likely to lose eligibility due to a yearlong lack of earnings than were unmarried women with high educational attainment. Tests of equivalence of the effect of education indicate that low education was a predictor of eligibility loss through zero earnings only for this group. A symmetric analysis examines eligibility gain over the same period. During this time, women with low educational attainment were less likely than other labor market groups to become eligible through underemployment, and showed no greater propensity to gain eligibility through labor market entry or family change.

These results bring into question the effectiveness of the EITC as a substitute for welfare, as has been emphasized by both policy experts and economists. The current structure of the EITC was established during the late 1990s as part of welfare reform, with the intent of drawing welfare leavers into the job market. During the Great Recession, the policy appears to have missed this main target population—low-skilled single mothers who would otherwise be welfare users. This is not to suggest that the EITC does not "work" as a policy. In fact, the EITC appears to work exactly as intended, and has been instrumental in drawing welfare leavers into the workforce when jobs are available. However, the results have serious implications for the existing focus in policymaking on tying income supports to work. Economic analyses of the EITC have consistently presupposed that jobs are available to welfare leavers; the current work demonstrates what happens when this assumption is not met.

The paper proceeds as follows. Section 2 provides information on the EITC and previous literature. Section 3 goes over the data used in the analysis, describing the sources for the data and the processes by which data sets are linked. Section 4 describes the methods used. Section 5 presents the results and describes some implications of the results. Section 6 provides some sensitivity analyses, and section 7 concludes.

2. BACKGROUND AND LITERATURE

The EITC is a refundable tax credit that arrives as a lump sum in an earner's tax return.[1] The main original intent of the EITC was to reimburse payroll taxes for low-income

[1] The Advance Earned Income Tax Credit was eliminated beginning in tax year 2011. Only about 3% of yearly EITC recipients used the program (http://www.gao.gov/highlights/d071110high.pdf).

earners, for whom these taxes represent a disproportionately high percentage of earnings (Hoffman and Seldman, 2003). While the credit is modest for earners without children, families with children can receive credits as high as 40 percent of their wage and salary earnings. The EITC has been credited with expanding the labor-market participation of single mothers—in effect "making work pay" (Meyer and Rosenbaum, 2001). Use of the EITC as a "safety net" program changed the nature of government assistance to low-income families, from "out-of-work aid (welfare) to in-work aid (EITC)" (Bitler et al., 2014). For those who retained eligibility over the Great Recession, the EITC provided valuable income support (Larrimore et al., 2013).

Figure 1 shows the program parameters for the EITC in 2011, the latest year of data in this analysis. According to the program parameters,[2] a modeled tax filer becomes eligible when his earnings become greater than zero and less than the maximum allowable income (defined as the maximum of earnings or adjusted gross income (AGI)). He loses eligibility when earnings drop to zero or increase beyond the maximum. A married labor-force participant becomes eligible when she and her spouse earn more than zero and less than the maximum allowable income. It is easy to see from the figure that, for a given person, a change from eligible to ineligible from one year to another may occur due to earnings falling below zero or surpassing the threshold. Family change may also play a part; holding earnings constant for an individual, the aging of children may drive a previously eligible person into an ineligible earnings range. The same holds true for divorce, since the earnings thresholds are approximately $5000 higher for married tax filers. The nature of the program parameters makes it difficult to assign causality to eligibility loss when the researcher is unable to measure an individual's earnings and family changes from year to year.

It is also clear from the figure that the program provides zero assistance to those unable

[2]The description that follows is simplified. Earnings is the key eligibility requirement, but there are other rules governing eligibility, such as a limit on investment income.

to find work over a tax year. A large body of research has demonstrated that the EITC was instrumental in drawing single mothers into the workforce; however, research that assessed the impact of the EITC on labor-force participation of welfare leavers did so over a period when the economy was expanding and unemployment rates were low (Meyer and Rosenbaum, 2001; Grogger, 2003). Even during this period, evidence indicated that many female-headed households fell through the cracks, receiving neither welfare nor wages. For unmarried mothers who did make the transition, employment was often tenuous and unlikely to be covered by unemployment insurance. Thus, although the transition from welfare to work has been well-documented for most low-income women, a non-trivial proportion were unable to make the transition (Turner et al., 2006).

In cases when one earner becomes full-year unemployed, a married family may still be eligible if the other spouse has earnings. Therefore, due to the recession's differential effects on low-skilled earners and men, plus the interaction of marriage and education, it is likely that the direction of eligibility change is different between groups. This belief is supported by research into the effect of recessions on different skill, sex, and race groups. Elsby and Hobijn (2010) found that—similar to earlier recessions—young, male, and less-educated workers and those from ethnic minorities were more strongly affected by the economic downturn than other groups. In examining an earlier time period (1979-1992), Hoynes (1999) found that the labor market outcome of low-skilled workers exhibit greater cyclicality (wider swings between employment and unemployment, for example) in response to economic downturns compared with higher-skilled workers (an update to this paper, Hoynes et al., 2012, found a similar pattern for the Great Recession). In tandem with labor market outcomes, earners are known to cycle in and out of EITC eligibility as their incomes fluctuate (Horowitz, 2002).

Moreover, the recent recession was marked by two characteristics of importance to EITC receipt. One was the low rate of exit from unemployment to jobs, indicating a higher rate of all-year job loss compared with earlier recessions. The other was the persistence in the reduction of hours for workers who managed to be employed over the period, which would lead to lower household income (Elsby and Hobijn, 2010). Together, these effects may conspire to cancel out population-level changes in EITC eligibility. Eligibility by group, however, may change depending on which labor market outcome was predominant.

A further issue is the dynamics of eligibility and marriage. While married men are more likely to participate in the labor force than are single men, married women have lower participation rates than single women, and these differences are themselves affected by skill and presence of children (Juhn and Potter, 2006). For married tax filers, "marriage insurance" or an "added worker effect" may exist—the extent to which wives (or husbands) change their participation in the labor force to cushion the shock of a spouse's unemployment (for example, Stephens, 2002 and Juhn and Potter, 2007).

Thus, EITC eligibility and employment may be simultaneously predicted by race, gender, education, marriage, and childbearing. Jones (2013), for example, found that men, joint filers, and families with children experienced differentially greater increases in EITC eligibility over the Great Recession, while low-skilled workers experienced decreases.

The cyclical nature of both income and transfer programs has also been the subject of attention (for example, Blank, 2001 and Ziliak et al., 2000). Unsurprisingly, previous research has shown that caseloads for public assistance rise when the economy turns down, illustrating the countercyclical nature of transfer programs. There is precedent for using a competing risks analysis to model the loss of program eligibility or take-up: Blank and Ruggles (1996) used just such a framework to examine the cyclical nature of eligibility and take-up in the Aid to Families with Dependent Children (AFDC) and Food Stamp programs, finding that most program spells were short-lived and were likely to end with an increase in income beyond the program thresholds.

3. DATA

Beginning with tax year 2005, the U.S. Census Bureau has estimated the take-up of the EITC for the IRS using linked survey and tax records. Each CPS ASEC, administered annually in March, includes questions regarding family structure and earnings that can be used to estimate tax filings for the preceding year. The use of nationally representative survey data is necessary for the denominator of the take-up rate, since not all those who are eligible actually file taxes. Take-up for EITC is lowest for those who are estimated as eligible and have earnings in the phase-in region of the credit; thus, using eligibility rather than take-up better captures people with more tenuous connections to the labor market. Survey data also provide demographic measures not available in tax data or caseload data, such as sex, education, race, and Hispanic origin. In this paper, I use the CPS ASEC 2006 linked with tax records from 2005 to 2011.

Records were linked using a process whereby individuals in each data set were given a unique key, called a Protected Identification Key (PIK). The Center for Administrative Records Research and Applications (CARRA) assigned these unique identifiers via the Person Identification Validation System (PVS), which employs probability record linkage techniques (see Wagner and Layne (2014) for more information).[3] CARRA uses personally identifiable information (PII) such as name, date of birth, and address to assign a PIK. CARRA then removes the PII from the data file to anonymize the data and preserve confidentiality so it can be used for statistical purposes and research. Only those observations that received the unique key are used in the analysis. Furthermore, a match is only used if CPS earnings were not imputed or allocated.[4]

[3]Over the time period included, the PVS system was altered to include 1040 observations with Individual Taxpayer Identification Numbers (ITINs). The inclusion of ITINs changes the sample slightly, with more non-citizen tax filers being identified. These observations would, however, only be EITC eligible if they were married to a 1040 filer with an SSN. To check whether this had any influence on my results, I ran all analyses on citizens. The results were unaffected.

[4]There is some bias introduced in who receives a PIK and who does not; see (Bond et al., 2014). To check whether this bias affects my estimates, I ran weighted models where the CPS sample weights had been recalculated based on the probability of receiving a PIK. Estimates were unchanged.

The Census Bureau estimates EITC eligibility first using the survey responses to questions that reflect the EITC program parameters. These responses allow for the creation of tax filing units, assignment of dependents, and determination of filing status. Individuals are identified as potential tax filers, and their EITC eligibility is determined, based on these values. Then, these values are superseded by the "true" values from the tax data whenever possible, and eligibility is recalculated. Superseded variables include earnings, adjusted gross income, investment income, filing status, and claimed dependents. Uncertainties in determining eligibility tend to arise around household formation, since the presence of children is a prerequisite for the higher EITC amounts and the wider range of eligible earnings.

For this paper, eligibility for 2006 CPS ASEC respondents was estimated using further years of 1040 and W-2 data. I consider household structure to be fixed at the characteristics reported in the survey, and I then age-out children from qualifying status using the ages reported for them in 2006. The number of children reported for a household is replaced by the number of children claimed on 1040 data in future years. Similarly, I assume that an individual's marital status remains at 2006 reports unless he or she reports otherwise in later 1040 data. I limit the sample to those who are present in 1040 or W-2 data in 2005.

Using W-2 earnings, 1040 earnings, 1040 AGI, 1040 interest income, marriage in 2006 superseded by filing status in later years, and children reported in 2006 superseded by children claimed in later years, I run the same eligibility estimation procedure described above for each subsequent tax year. I retain all observations in the file who experienced a spell of eligibility between 2005 and 2011.

There are some concerns in using this data as described, some of which can be partially addressed. The first issue is that some respondents to the 2006 CPS ASEC may have left the sample, and their attrition is due to something other than lack of labor-force participation and 1040 filing. Some 2006 survey respondents may have died. Others may have left the labor force due to disability or retirement, although we would expect these respondents to have filed taxes in later years due to sources of income such as Social Security benefits. To address this issue to some extent, I remove any survey respondents if I observe them in a given year, but never in a subsequent year. For example, if I have a W-2 record for a survey respondent in 2009, but I do not have a W-2 or 1040 record for that person in 2010 or 2011, I drop the person from the sample. The loss of a such a respondent means that the long-term unemployed will be underrepresented in the final sample. Because the relationship between unemployment and EITC eligibility is the focus of the analysis, any coefficients reporting this relationship will be underestimated. I examine the impact of this sample restriction in section 6.

A second concern is education, as many EITC-eligible respondents may continue their education over the time period. There is no method by which these changes can be captured. As a partial solution to this problem, I restrict the sample to respondents who were 25 or older at the time of the CPS ASEC survey, thus retaining those who have more likely finished their education by tax year 2005. I also restrict the sample to earners who had at least one qualifying child in at least one tax year.

A person is considered as becoming at risk for losing eligibility when he or she is eligible in any time period, including the first. A person is considered as at risk for becoming eligible from the start of the time period, excluding those who are already eligible at the start. For those who filed a joint return at the start of the period, each filer's eligibility is considered separately, since they may file a single return in a later year. For each modeled tax filer, I look only at the first episode of eligibility. Meanwhile, eligibility that does not end by 2011 is censored.

Table 1 shows the sample that results, displaying EITC eligibility status by year. Panel A of Table 1 shows the pattern of eligibility spells for tax filers with children in the 2006 CPS ASEC, with each cell capturing the number who started a spell in the year listed in the row and ending in the year listed in the column. As reported in the table, a total of

20,759 CPS ASEC 2006 filers were EITC eligible at some point between 2006 and 2011. Of these, 10,745 started the period in an eligible state. Of those who were eligible in the first period, 3,156 (30 percent) retained eligibility for the entire period (reported in Panel B). For those who became eligible after tax year 2005 (12,045 respondents), 3,826 retained eligibility to the end (37 percent). Most spell lengths for those eligible in the first period were seven years, followed by one year. For those starting in period 2 or later, most eligibility spells were one year. As with any duration data, it is impossible to know what eligibility spells would look like if we had everyone's lifetime history. However, it does appear as though EITC eligibility, at least over the recession, was persistent for much of the sample.

The next step in the data construction was to assign to filers the different ways a person could lose or gain EITC eligibility. Eligibility loss (or, in the terms of survival analysis, "failure") can be pooled into three categories: total loss of earnings, earnings or AGI above the maximum, and family change (including loss of qualifying children and marriage or divorce). To the extent that eligibility determination is based on EITC rules, these risk categories are exhaustive but not mutually exclusive without making some important assumptions. A person without earnings is, by definition, ineligible. However, a person who chooses not to earn a wage because of high total income would not have been eligible regardless of whether he or she wanted to earn a wage. Thus I categorize the initial risk as due to loss of earnings, but replace the risk category as due to high income if a respondent without earnings fits this definition. Income maximums for eligibility are, however, dependent on family structure, with higher income levels allowed for married couples and those with more children. Thus I supersede the high-AGI category with the family-change category for cases in which the respondent reported earnings, had income above the maximum, and in the same period experienced a change in family status.

It is possible that members of a group that cycle out of eligibility through a given risk may include members who cycle in due to a countervailing risk. To address this question, I also examine the risks of eligibility gain. These are symmetric with the eligibility loss risks. A person becomes eligible upon his or her earnings going above zero and staying below a threshold; by income decreasing below the threshold; or by a change in family structure that brings the person into eligibility regardless of a change in earnings or income.

Table 2 shows, by year, the numbers of CPS ASEC 2006 tax filers with children who went from EITC eligible to ineligible, by risk. For each column, the number who do not experience a failure include those who have not yet become EITC eligible. In each year, eligibility loss due to lack of earnings is approximately 20 percent, with an increase in 2009 to 30 percent, the year when the national unemployment rate increased to 10 percent. There is a slight increase in later years in the number failing due to income increasing beyond the EITC threshold, and a slight decrease in later years in the percent failing due to family change.These changes in percent may simply reflect the effects of time on earnings growth and the ages of children for the modeled filers in the 2006 CPS ASEC. For comparison, Table 2 also shows the numbers entering eligibility by risk.

4. EMPIRICAL METHODS

Using panel data, my key question is how eligibility changed for individuals over time in response to changes in economic circumstances. To answer this question, I use competing risks models that examine how eligibility spells for the EITC end. To check for a counterbalancing movement from ineligibility to eligibility, I also examine how spells begin, but the main change of interest is eligibility loss.

As mentioned, the three risk categories are mutually exclusive; they are competing in the sense that an EITC-eligible individual, by definition, may lose eligibility in the next period due to one failure risk, but not another. The main risk I am interested in is lack of earnings, which can be viewed as the risk that implies the worst outcome induced by unemployment: a loss of earnings simultaneous with a loss of a tax credit. Each of the three risks is assigned a numeric category, while those who do not fail in the period are assigned a zero. For entry into eligibility, risk 1 is entry due to partial loss of earnings; risk 2 is entry due to earnings increasing above zero; and risk 3 is family change.

Following previous literature, I look at the experiences of eight labor market groups defined on sex, education (low and high), and marital status. I first examine the risk profile for the entire sample of tax filers from the 2006 CPS ASEC. Then, rather than break up groups by all three labor market characteristics (marital status, sex, and education), I run separate models on four groups—unmarried women and men and married women and men—and use education as the main independent variable of interest. This allows me to test whether the effect of skill differs across the four groups and across risks. Alternatively, I could have split the groups into eight categories, or split them by marriage and education and examined differences by sex. The resulting risk ratios on education in each model simply express differences in likelihood that a low-education person will experience the risk compared with a high-education person of the same group.

The modeling technique I use is that of Fine and Gray (1999) (hereafter, FG). In contrast to a Cox regression analysis, an FG competing risks model allows a researcher to assess events that compete with failure from the event of interest, rather than having competing events treated as though they were censored. Rather than a survivor function, competing risks models consider a failure function—the cumulative incidence function (CIF), which describes the probability that an event will take place before a certain time ($P(T \leq t$ and event type k, out of a set of events $\epsilon \in (1, ..., K)$). Considering $\epsilon = 1$ and covariates x, F&G define the "subdistribution hazard":

$$\lambda_1(t; \mathbf{X}) = \lim_{\delta\mathbf{t} \to \mathbf{0}} \frac{\mathbf{1}}{\delta\mathbf{t}} \mathbf{Pr}[\mathbf{t} \leq \mathbf{T} \leq \mathbf{t} + \delta\mathbf{t}, \epsilon = \mathbf{1} | \mathbf{T} \geq \mathbf{t} \cap \epsilon \neq \mathbf{1}, \mathbf{X}] \tag{1}$$

where λ_1 is the hazard of interest. Those who fail due to other hazards remain at risk. The cumulative hazard function can be easily calculated from the subdistribution hazard. The baseline subhazard is subsumed in the model, while the effects of covariates, as in a

Cox regression, are proportional:

$$\lambda_1(t; \mathbf{X}) = \lambda_{\mathbf{1,0}}(\mathbf{t}) \exp[\mathbf{X}^{\mathbf{T}}(\mathbf{t})\beta] \tag{2}$$

In the model as outlined above, for an event k happening in a small interval $(t, t + \delta t)$, the risk set includes those who are alive at t and those who failed before t from events other than the event in question. Each event is considered separately. Similar to a Cox proportional hazards model, time-varying coefficients can be included to test the proportionality assumption. In other words, for a given covariate x_1, the proportion becomes

$$\lambda_1(t; X) = \lambda_{1,0}(t) \exp[(\beta_1 + \gamma_1 g(t))x_1] \tag{3}$$

where $g(t)$ is a linear function of time. Thus, γ_1 gives the magnitude of the deviations from the main effect, β_1, over time. Because proportional subhazards assume that β_1 is time-invariant, I include time-varying coefficients in the model for each control variable to ensure β_1 reflects the baseline hazard for the risk.

The FG model takes the last value of a time-dependent covariate for an observation that fails because of a competing risk. The inclusion of a time-dependent covariate thus may bias the results because later values are not captured. I use only time-invariant covariates or those measured at baseline (tax year 2005). These include the two education levels (defined as a high school education or less and more than a high school education), race, Hispanic origin, age, number of qualifying children, and the state unemployment rate and state EITC. When unemployment rate is included as a time-varying measure, the risk ratios on the variable of interest—education—do not change substantially, likely because local unemployment rates experience some serial correlation.

The focus of this analysis is on the intersection of sex, education, and marriage in the joint determination of labor-market and eligibility outcomes. While the unemployment rate is not the independent variable of interest, full-year unemployment, expressed by the total loss of earnings, is included as the risk of interest among the possible competing risks, while educational attainment becomes the independent variable that allows me to assess the differential incidence of this risk for each group. To check for a countervailing entry by groups *into* eligibility, I compare estimates from the same models, but using the risks of entry defined previously.

5. RESULTS

5.1. Competing risks results

Table 5 shows the results of the competing risks analysis when the full panel is used. Hazard ratios are reported for each of the risks in separate columns. Each hazard ratio indicates the main effect of the variable on the probability that an eligibility spell ended through the given risk. The main effect for each time-invariant covariate is shown.[5]

The reported ratios express the additional likelihood that a modeled filer with the given characteristic (or higher value) failed for the reason specified. Results show that women were less likely than men to lose eligibility by any means, and that the hazard ratio was fairly similar across risks. In general, women had about a 30 percent lower risk of any type of eligibility loss compared with men. Compared with the more highly educated, those with a high school education or less had a 35 percent lower risk of loss due to high income, and a 21 percent lower risk of loss due to family change. Those who were unmarried at baseline were about 40 percent less likely than those who were married to lose eligibility due to total earnings loss or high income, and 19 percent more likely to lose eligibility due to family change.

The control variables give a picture of eligibility loss, indicating that Hispanic and non-White modeled filers tended to retain eligibility, with the exception of those reporting Asian alone, who experienced a higher likelihood of losing eligibility due to high earnings. Older earners tended to have a higher risk of eligibility loss due to any cause,

[5]I do not show time-varying coefficients, but they are described in the footnote to each table.

while those with a greater number of children experienced a lower risk of eligibility loss. Finally, the baseline unemployment rate in the state was associated with an increased risk of eligibility loss due to high income (8 percent). This last finding is surprising, although it could be capturing earnings growth over the period for those who manage to stay in the workforce in high-unemployment areas. This earnings growth may reflect a greater number of hours worked for those who retained a job during this period as employers limited hiring.

Table 6 shows the competing risks analysis for women who were unmarried at baseline and men who were unmarried at baseline (separate models). Compared with more highly educated unmarried women, unmarried women with low education were 22 percent more likely to experience eligibility loss due to lack of earnings, and about half as likely to lose eligibility due to high income. In contrast, low-education unmarried men were no more likely to lose eligibility due to lack of earnings than their more highly educated peers. They were, however, about 30 percent less likely to lose eligibility due to high income. Unmarried women who reported their race as Black did not differ in their risk profile compared with the full sample, with the exception that the hazard ratio on loss due to lack of earnings is not precisely estimated. Black unmarried men did not differ from White unmarried men except for the risk of loss due to high income, with Black men 61 percent less likely to lose eligibility due to this risk (the same pattern was seen for women who reported their race as "other"). Women who reported their race as Asian alone were 2.7 times as likely than White alone, unmarried women to experience eligibility loss due to family change. These results may be driven by a small number of observations, since there are few Asian women who were unmarried at baseline but have children at some point during the sample period. For unmarried women, having more children at baseline translated into a lower risk of eligibility loss due to high income or family change. An increase in one qualifying child was associated with a 37 percent decrease in loss due to high income and a 34 percent decrease in the risk of loss due to family change. For men, a one-child increase was associated with a 41 percent lower risk of loss due to high earnings.

Table 7 shows the same analysis for women and men who were married at baseline. Because I am looking at individual trajectories, it seems important to point out that the subsamples shown in these tables include observations who were married to one another at baseline. If many of these couples remained married, their failure from a particular risk would have been concurrent. Thus, it is not surprising that the risk profile for married women and men look more similar to one another than the profiles for unmarried women and men. Both men and women with low educational attainment were less likely than their better-educated counterparts to experience loss due to high income (38 percent less for women and 32 percent for men) and family change (29 percent less for women and 26 percent for men). Married women who reported their race as Black alone were less likely than married White women to lose eligibility due to lack of earnings or family change. Both men and women who reported Asian alone were less likely than married whites to lose eligibility due to family change. Hispanic men were less likely than non-Hispanic men to lose eligibility due to high earnings. For married men and women, a one-year increase in age at baseline was associated with a 1 to 2 percent increase in eligibility loss for any reason; however, the ratios were not statistically significant for the lack of earnings risk for women and the family change risk for men. Baseline number of children and unemployment rate were, for married men, associated with an increase in eligibility loss due to high income.

In sum, the results confirm the main hypothesis of this research: that for more vulnerable labor market participants—unmarried, low skilled women with children—there was a greater risk of year-to-year EITC eligibility loss due to full-year unemployment and lack of earnings.

To check whether there was a countervailing entry by low-education women into eligibility, I used the same competing risks framework and examined entry risk. The results of this analysis appear in Tables 6 to 8. Low-education unmarried women were half as likely as those with a higher education to enter eligibility due to a decrease in income (in other words, from having earnings or income above the threshold in t and below the threshold in $t + 1$). Meanwhile, the low-education group was no more likely to enter eligibility due to joining the labor market after a period of zero earnings, or through a change in family status. For the sake of completeness, the results for the full sample and for the married sample are shown, in Tables 6 and 8. As one would expect, women are more likely than men to enter eligibility by having their earnings increase from zero and through family change, and less likely to enter through decreased income. Those who were unmarried in 2005 were less likely to become eligible by any risk compared with married filers.

5.2. *Impact of eligibility loss*

The preceding analysis uncovered the greater propensity for unmarried women with low educational attainment to lose eligibility due to a lack of earnings in a tax year. In examining this risk alone, panel A of Table 9 shows the unconditional probability, by year, that a person of a given labor market group will lose eligibility through zero earnings. The table gives a sense of how this probability changed over the recession for the eight different labor market groups.

In 2006, low-education, unmarried women had a probability of zero-earnings eligibility loss that was average. Beginning in 2007, this group consistently had the highest unconditional probability of exit through zero earnings; in each year except 2011, the probability is statistically higher than the overall average for all groups. Unmarried men with low educational attainment also showed high rates, with statistically higher-than-average rates from 2007 to 2009.

To give an idea of the scale of the impact of this risk of loss on low-education single women, Panel B of Table 9 uses the CPS ASEC population weight to show the flows of this group into and out of eligibility over time, with the flows defined only for the first spell of eligibility. As shown in the table, the key years of the recession (2008 and 2009) are associated with an uptick in net flows out of eligibility. Meanwhile, the number losing eligibility from one year to the next through the risk of zero earnings ranged from a high of 370 thousand in 2009 to a low of 168 thousand in 2011. The average EITC for those who are eligible in 2005 and exit in a later period is approximately $2000 (in 2011 dollars). This evidence paints a dire picture for families headed by low-skilled and unmarried women, a large number of which lost all employment earnings and their EITC in the same tax year.

6. SENSITIVITY ANALYSES

The creation and analysis of the panel data required some assumptions that may have affected the results just discussed. The first issue is attrition: observations for whom we do not have a further record in the tax data after a given year. Among the many reasons why we would not have data for these observations is withdrawal from the labor market. Dropping these observations may significantly underestimate the effect of the Great Recession on eligibility, since one of the distinguishing features of the latest recession was long-term unemployment and the discouragement of workers (Elsby and Hobijn, 2010). To test the effect of attrition, I looked at models that retained the missing observations and coded the failure type for these observations as due to a lack of earnings. The results are reported in Table 8.

The table shows the results for unmarried and married respondents when those who attrit from the data are retained and their risk is coded as "no earnings." For unmarried women, the risk of eligibility loss due to lack of earnings is intensified (36 percent versus 22 percent), which fits the hypothesis that workers with less skill are more likely to withdraw from the labor market due to discouragement. The corresponding risk ratios for the other groups are not different. For loss due to lack of earnings, age now appears to have a positive relationship for all groups, with older ages associated with a higher risk. This positive effect may reflect loss that has been coded as lack of earnings but is not due to discouragement, but rather a withdrawal from the workforce due to death, disability, or retirement (although we would expect this last group to still file 1040s to reflect retirement income).

A second issue of possible concern is the use of the Fine and Gray model. The justification for using the model depends on how the question is framed. The Fine and Gray analysis presented hazard ratios when the comparison groups were retained as still at risk rather than censored. Using data duplication (such as the technique outlined in Putter et al., 2007, among others), similar models can be run using Cox regressions where the hazards for each risk are modeled simultaneously.

Table 9 shows the results of running Cox models. The results are similar to the main models, with certain hazards slightly greater and others diminished. Specifically, the risk that unmarried women lose eligibility due to a lack of earnings is again 18 percent greater if she has low educational attainment (compared with 22 percent in the Fine and Gray model). Also of interest are the hazard ratios on unemployment rate, none of which are statistically significant in the Cox models.

Using Cox models also allows for tests that the effect of education on the risk of loss due to lack of earnings is statistically different between groups that share an important characteristic. For example, we might want to know whether the effect of education, which is significant only in the models restricted to unmarried women, truly differs between married and unmarried women (or unmarried women and unmarried men). Such an analysis is possible with a Cox model that has been stratified by the variable of interest to estimate the separate baseline hazard rates. Currently, there is no similar test available with the Fine and Gray model (Putter et al., 2007).

Each Cox model is run with the variable of interest—education—interacted with each independent variable and then used to stratify the model. When looking only at women, the coefficient on married times education provides a test of whether the effect of education is the same for unmarried and married women. The Chi-squared value for this test was 5.7, allowing for a rejection of equality. Similarly, when looking only at those who were unmarried at baseline, the coefficient on sex times education provides a test of whether the effect of education is the same for men and women. The Chi-squared value in this case was 5.5, again allowing for a rejection of equality. These tests reinforce the interpretation that unmarried women whose educational attainment was low faced a greater risk of eligibility loss due to zero earnings than did other groups who shared a common characteristic (female or unmarried).

7. CONCLUSION

Considering the EITC's importance in bolstering the income of low-wage earners, any evaluation of the program needs to take into account what happens when earnings are threatened by an economic downturn. Recent research using caseloads and looking at the effect of state unemployment on eligibility indicate that married earners saw increases in eligibility over the recession, while eligibility was flat for single earners. The only explanation for this in light of almost-universal negative job market experiences is that the presence of spousal earnings retained eligibility, or drove families into eligibility when their previous earnings had been too high. Clearly, marriage has a protective effect when it comes to eligibility for the EITC. Similar research using administrative records also shows that eligibility retention differs by education, sex, and race groups.

The results of a competing-risks analysis support and extend this research. Unmarried women experienced a higher risk of loss due to zero earnings when their educational attainment was low. This result is troublesome, since it indicates that, at least in a downturn, the EITC fails to reach the target population for whom it was specifically expanded during the welfare reform era. Looking at the outcomes for individuals over time indicates that marriage, gender, and skill were each important factors in how individuals transitioned out of eligibility. Population estimates indicate that during the recession, hundreds of thousands of families lost both earnings income and distributions from the key cash-transfer program in the U.S.

The effectiveness of the EITC in meeting its policy goals is not the subject of this analysis. During times when employment opportunities are widely available, there is no doubt that the EITC is strong enticement for labor force participation. There is also no doubt that it provides a substantial boost in income for those who can find work, especially married families where one spouse remains employed. However, all of the excellent aspects of the policy become irrelevant for single earners when no jobs are available. Between the difficulty of enrolling in TANF or remaining in the program, a lack of coverage of unemployment insurance and curtailment of benefits, and cuts to the Supplemental Nutrition Assistance Program, many female-headed households may soon find themselves without a safety net program to assist them in times of need.

Compliance with Ethical Standards: The author declares that she has no conflict of interest.

Bitler, M., Hoynes, H., Kuka, E., 2014. Do In-Work Tax Credits Serve as a Safety Net? Tech. rep., National Bureau of Economic Research.

Bitler, M. P., Hoynes, H. W., 2010. The State of the Social Safety Net in the Post–Welfare Reform Era. Brookings Papers on Economic Activity.

Blank, R., Ruggles, P., 1996. When do women use Aid to Families with Dependent Children? Journal of Human Resources 31 (1), 57–89.

Blank, R. M., 2001. What Causes Public Assistance Caseloads to Grow? Journal of Human Resources 36 (1), 85–118.

Bond, B., Brown, J. D., Luque, A., O'Hara, A., 2014. The Nature of the Bias When Studying Only Linkable Person Records: Evidence from the American Community Survey. Tech. rep., U.S. Census Bureau, Center for Administrative Records Research and Applications.

Elsby, M. W., Hobijn, B., 2010. The labor market in the great recession. Brookings Papers on Economic Activity.

Fine, J. P., Gray, R. J., 1999. A proportional hazards model for the subdistribution of a competing risk. Journal of the American Statistical Association 94 (446), 496–509.

Grogger, J., 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.

Hoffman, S. D., Seldman, L. S., 2003. Helping Working Families: The Earned Income Tax Credit, w Edition. W.E. Upjohn Institute for Employment Research, Kalamazoo, MI.

Horowitz, J. B., 2002. Income mobility and the earned income tax credit. Economic Inquiry 40 (3), 334–347.

Hotz, V. J., Scholz, J. K., 2006. Examining the effect of the Earned Income Tax Credit on the labor market participation of families on welfare. Tech. rep., National Bureau of Economic Research.

Hoynes, H., 1999. The employment, earnings, and income of less skilled workers over the business cycle. Tech. rep., National Bureau of Economic Research.

Hoynes, H., Miller, D. L., Schaller, J., 2012. Who Suffers During Recessions? The Journal of Economic Perspectives 26 (3), 27–47.

Jones, M. R., 2013. Changes in EITC eligibility and participation: 2005–2009. IRS Research Bulletin.

Juhn, C., Potter, S., 2006. Changes in labor force participation in the united states. The Journal of Economic Perspectives 20 (3), 27–46.

Juhn, C., Potter, S., 2007. Is there still an added-worker effect? FRB of New York Staff Report (310).

Larrimore, J., Burkhauser, R. V., Armour, P., 2013. Accounting for income changes over the great recession (2007-2010) relative to previous recessions: the importance of taxes and transfers. Tech. rep., National Bureau of Economic Research.

Meyer, B. D., Rosenbaum, D. T., 2001. Welfare, the Earned Income Tax Credit, and the labor supply of single mothers. Quarterly Journal of Economics 116 (3), 1063 – 1114.

Moffitt, R. A., 2013. The Great Recession and the Social Safety Net. The ANNALS of the American Academy of Political and Social Science 650 (1), 143–166.

Putter, H., Fiocco, M., Geskus, R., 2007. Tutorial in biostatistics: competing risks and multi-state models. Statistics in medicine 26 (11), 2389–2430.

Stephens, M., 2002. Worker Displacement and the Added Worker Effect. Journal of Labor Economics 20 (3), 504–537.

Turner, L. J., Danziger, S., Seefeldt, K. S., 2006. Failing the Transition from Welfare to Work: Women Chronically Disconnected from Employment and Cash Welfare. Social Science Quarterly 87 (2), 227–249.

Wagner, D., Layne, M., 2014. The Person Identification Validation System (PVS). Tech. rep., U.S. Census Bureau, Center for Administrative Records Research and Applications.

Williams, R., Maag, E., 2008. The recession and the Earned Income Tax Credit. Tech. rep., The Urban Institute.

Ziliak, J. P., Figlio, D. N., Davis, E. E., Connolly, L. S., 2000. Accounting for the decline in AFDC caseloads: Welfare reform or the economy? Journal of Human Resources, 570–586.

Table 1: Patterns of EITC eligibility gain and loss for the 2006 CPS ASEC panel, 2006 to 2011

	Exit year							
Panel A	2006	2007	2008	2009	2010	2011	Never	Total
Start year: 2005	3,372	1,449	999	558	524	401	3,156	10,459
2006		1,229	711	295	245	169	997	3,646
2007			707	222	145	116	387	1,577
2008				526	268	152	461	1,407
2009					823	324	604	1,751
2010						542	545	1,087
2011							832	832
Total	3,372	2,678	2,417	1,601	2,005	1,704	6,982	20,759

Panel B	EITC eligibility	
Starting in	Period 1	Period 2 or later
Number	10,459	10,300
percent	50.38	49.62
Number right censored	3,156	3,826
percent	30.17	37.15
Spell length (years)		
1	32.24	42.23
2	13.85	20.10
3	9.55	11.61
4	5.34	7.98
5	5.01	5.40
6	3.83	9.68
7	30.17	-

Source: CPS ASEC 2006 linked with 1040 and W-2 data from 2005–2011. Panel A reports the number of observations fitting into cells defined by the year eligibility began (rows) and the year it ended (columns). Panel B reports the number and proportion of observations experiencing the start time and length of eligibility spell.

Table 2: Number and percent of CPS ASEC 2006 tax filers who transition from EITC eligible to ineligible, or ineligible to eligible, by year and risk type

Panel A. Exit year	2006	2007	2008	2009	2010	2011
No earnings	856	460	455	476	444	298
percent of failure	0.25	0.17	0.19	0.30	0.22	0.17
Income¿max	1,152	928	727	582	893	755
percent of failure	0.34	0.35	0.30	0.36	0.45	0.44
Family change	1,364	1,290	1,235	543	668	651
percent of failure	0.40	0.48	0.51	0.34	0.33	0.38
No failure	17,387	14,709	12,292	10,691	8,686	6,982
Panel B. Entry year	2006	2007	2008	2009	2010	2011
Decreased income	745	490	540	912	494	347
percent of entry	3.65	2.4	2.65	4.47	2.42	1.7
Earnings¿0	637	154	114	117	111	92
percent of entry	3.12	0.75	0.56	0.57	0.54	0.45
Family change	2,197	883	713	603	425	366
percent of entry	10.77	4.33	3.5	2.96	2.08	1.79

Source: CPS ASEC 2006 linked with 1040 and W-2 data from 2005–2011. Panel A reports the numbers and percentages that transitioned from eligible to ineligible due to the risk listed in column 1. The number who do not experience a failure in a given year includes those who are eligible and do not fail in that year and those who are not yet eligible. Total observations: 20,759. Panel B shows those who became eligible over the time period, excluding the 51 percent of the panel that was eligible in 2005.

Table 3: Competing risks models estimating exit from EITC eligibility, full sample of CPS ASEC 2006 filers with children

	Cause of eligibility loss		
	No earnings	Income>max	Family change
Labor market variables			
Female	0.68***	0.73***	0.65***
	(0.06)	(0.05)	(0.04)
Low education	0.95	0.65***	0.79***
	(0.09)	(0.04)	(0.05)
Unmarried in 2005	1.00	0.57***	1.19***
	(0.04)	(0.02)	(0.03)
Control variables			
Black alone	0.77*	0.52***	0.60***
	(0.09)	(0.05)	(0.05)
Asian alone	1.05	1.19*	0.69***
	(0.10)	(0.07)	(0.05)
Other race	0.71	0.73	0.74*
	(0.14)	(0.12)	(0.11)
Hispanic	0.85	0.77***	0.79**
	(0.09)	(0.07)	(0.06)
Age	1.02***	1.02***	1.01*
	(0.01)	(0.00)	(0.00)
Number of children	1.03	0.95***	0.90***
	(0.02)	(0.01)	(0.01)
Unemployment rate	0.94	1.08*	0.94
	(0.04)	(0.04)	(0.03)
State EITC	0.79*	1.07	0.88
	(0.08)	(0.08)	(0.06)
Percent failing	14.40	24.26	27.70
Observations		20,759	

Source: CPS ASEC 2006 linked with 1040 and W-2 data from 2005–2011. $*p < 0.05, **p < 0.01, ***p < 0.001$. Reported are the risk ratios from Fine and Gray competing risks model in which each risk is modeled separately. For example, leaving eligibility due to no earnings is compared to leaving eligibility either for earnings/AGI above the maximum or family change. All variables reflect status in tax year 2005. Standard errors, clustered on the individual, appear in parentheses. Not reported are time-varying coefficients for sex, education, Black alone, other race, Hispanic, age, and baseline unemployment and EITC. Of the sample, 33.6 percent never lost eligibility after gaining it.

Table 4: Competing risks models estimating exit from EITC eligibility, women and men unmarried at baseline

	Women			Men		
	No earnings	Income>max	Family change	No earnings	Income>max	Family change
Labor market variable						
Low education	1.22**	0.50***	0.93	1.02	0.69***	0.93
	(0.08)	(0.03)	(0.04)	(0.11)	(0.06)	(0.07)
Control variables						
Black alone	0.92	0.57**	0.48***	0.63	0.39**	0.81
	(0.18)	(0.11)	(0.07)	(0.21)	(0.13)	(0.16)
Asian alone	3.57	2.91	2.71**	1.10	1.04	1.25
	(2.42)	(1.70)	(1.04)	(0.84)	(0.54)	(0.94)
Other race	0.85	0.48*	0.93	0.58	0.49	0.68
	(0.32)	(0.17)	(0.24)	(0.31)	(0.27)	(0.20)
Hispanic	1.02	1.00	0.81***	1.17	0.96	0.67***
	(0.09)	(0.09)	(0.05)	(0.15)	(0.11)	(0.07)
Age	1.00	1.02*	1.01	1.03*	1.03	1.00
	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
Number of children	0.99	0.63***	0.66***	0.81	0.59***	0.83
	(0.10)	(0.06)	(0.05)	(0.13)	(0.09)	(0.09)
Unemployment rate	0.91	1.06	0.91	1.00	0.94	0.90
	(0.08)	(0.08)	(0.05)	(0.14)	(0.11)	(0.07)
State EITC	0.82	1.07	0.95	0.84	0.94	0.95
	(0.16)	(0.18)	(0.11)	(0.25)	(0.24)	(0.18)
Percent failing	14.81	15.50	30.76	15.17	21.84	33.88
Observations		5,781			2,267	

Source: CPS ASEC 2006 linked with 1040 and W-2 data from 2005–2011. $*p < 0.05, **p < 0.01, ***p < 0.001$. Reported are the risk ratios from Fine and Gray competing risks model in which each risk is modeled separately. For example, leaving eligibility due to no earnings is compared to leaving eligibility either for earnings/AGI above the maximum or family change. All variables reflect status in tax year 2005. Standard errors, clustered on the individual, appear in parentheses. Not reported are time-varying coefficients for sex, education, Black alone, other race, Hispanic, age, and baseline unemployment and EITC. Of the sample, 33.6 percent never lost eligibility after gaining it.

Table 5: Competing risks models estimating exit from EITC eligibility, women and men married at baseline

	Women			Men		
	No earnings	Income>max	Family change	No earnings	Income>max	Family change
Labor market variable						
Low education	1.07	0.62***	0.71**	1.05	0.68***	0.74*
	(0.17)	(0.07)	(0.09)	(0.17)	(0.08)	(0.10)
Control variables						
Black alone	0.60*	0.84	0.64*	1.27	0.83	0.88
	(0.15)	(0.19)	(0.12)	(0.31)	(0.18)	(0.17)
Asian alone	1.10	1.16	0.68**	1.00	1.17	0.59***
	(0.17)	(0.11)	(0.09)	(0.17)	(0.11)	(0.09)
Other race	0.88	0.77	1.01	0.67	1.19	0.53
	(0.33)	(0.22)	(0.31)	(0.23)	(0.36)	(0.18)
Hispanic	0.76	0.79	0.79	0.88	0.66**	0.91
	(0.13)	(0.12)	(0.12)	(0.16)	(0.09)	(0.13)
Age	1.01	1.02**	1.02*	1.02*	1.02*	1.01
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Number of children	0.99	1.02	1.06	1.04	1.10	1.02
	(0.09)	(0.07)	(0.07)	(0.09)	(0.07)	(0.07)
Unemployment rate	0.99	1.11	1.00	0.90	1.15*	1.03
	(0.08)	(0.07)	(0.06)	(0.07)	(0.07)	(0.07)
State EITC	0.78	1.13	0.90	0.75	1.08	0.88
	(0.14)	(0.14)	(0.12)	(0.13)	(0.13)	(0.12)
Percent failing	13.52	27.38	24.74	14.67	30.10	25.73
Observations		6,625			6,086	

Source: CPS ASEC 2006 linked with 1040 and W-2 data from 2005–2011. $*p < 0.05, **p < 0.01, ***p < 0.001$. Reported are the risk ratios from Fine and Gray competing risks model in which each risk is modeled separately. For example, leaving eligibility due to no earnings is compared to leaving eligibility either for earnings/AGI above the maximum or family change. All variables reflect status in tax year 2005. Standard errors, clustered on the individual, appear in parentheses. Not reported are time-varying coefficients for sex, education, Black alone, other race, Hispanic, age, and baseline unemployment and EITC. Of the sample, 33.6 percent never lost eligibility after gaining it.

Table 6: Competing risks models estimating entry into EITC eligibility, full sample of CPS ASEC 2006 modeled filers with children

	Cause of eligibility gain		
	Decreased income	Earnings>0	Family change
Labor market variables			
Female	0.40***	4.35***	1.19**
	(0.04)	(0.61)	(0.08)
Low education	0.91	1.00	1.31***
	(0.09)	(0.13)	(0.08)
Unmarried in 2005	0.54***	0.70***	0.65***
	(0.02)	(0.05)	(0.02)
Control variables			
Black alone	0.96	0.52***	3.35***
	(0.16)	(0.10)	(0.28)
Asian alone	0.86	1.20	0.95
	(0.07)	(0.16)	(0.07)
Other race	1.08	0.68	1.39*
	(0.32)	(0.23)	(0.21)
Hispanic	0.75*	0.79	2.25***
	(0.10)	(0.14)	(0.18)
Age	1.00	0.99	1.00
	(0.01)	(0.01)	(0.00)
Number of children	1.00	1.04	0.63***
	(0.02)	(0.03)	(0.01)
Unemployment rate	1.05	0.95	1.03
	(0.05)	(0.06)	(0.03)
State EITC	1.04	1.21	0.90
	(0.11)	(0.17)	(0.06)
Percent entering	14.40	24.26	27.70
Observations		20,759	

Source: CPS ASEC 2006 linked with 1040 and W-2 data from 2005–2011. $*p < 0.05, **p < 0.01, ***p < 0.001$. Reported are the risk ratios from Fine and Gray competing risks model in which each risk is modeled separately. For example, leaving eligibility due to no earnings is compared to leaving eligibility either for earnings/AGI above the maximum or family change. All variables reflect status in tax year 2005. Standard errors, clustered on the individual, appear in parentheses. Not reported are time-varying coefficients for sex, education, Black alone, other race, Hispanic, age, and baseline unemployment and EITC. Of the sample, 33.6 percent never lost eligibility after gaining it.

Table 7: Competing risks models estimating entry into EITC eligibility, women and men unmarried at baseline

	Women			Men		
	Decreased income	Earnings>0	Family change	Decreased income	Earnings>0	Family change
Labor market variable						
Low education	0.52***	1.07	1.01	0.82	0.77	1.15*
	(0.05)	(0.14)	(0.06)	(0.13)	(0.14)	(0.08)
Control variables						
Black alone	0.51**	1.34	1.92***	1.22	3.20	2.56***
	(0.13)	(0.44)	(0.28)	(0.71)	(2.04)	(0.47)
Asian alone	1.71	0.93	2.91*	0.21	4.22	1.56
	(0.88)	(0.77)	(1.26)	(0.33)	(5.06)	(0.65)
Other race	1.17	0.47	1.11	0.35	0.96	1.14
	(0.69)	(0.31)	(0.30)	(0.30)	(1.23)	(0.35)
Hispanic	0.67**	1.04	1.25**	0.85	1.01	1.19*
	(0.10)	(0.17)	(0.09)	(0.16)	(0.23)	(0.09)
Age	1.03*	0.95**	1.02***	0.99	0.96	1.03***
	(0.01)	(0.02)	(0.01)	(0.02)	(0.02)	(0.01)
Number of children 1.04	1.13	0.20***	1.27	1.24	0.28***	
	(0.13)	(0.17)	(0.03)	(0.32)	(0.34)	(0.07)
Unemployment rate 1.25*	0.89	0.98	0.96	0.90	1.10	
	(0.14)	(0.12)	(0.06)	(0.21)	(0.22)	(0.08)
State EITC	0.97	1.19	0.96	1.37	0.57	0.93
	(0.22)	(0.36)	(0.13)	(0.62)	(0.30)	(0.15)
Percent entering	18.14	8.98	29.98	25.97	4.33	22.43
Observations		5,781			2,267	

Source: CPS ASEC 2006 linked with 1040 and W-2 data from 2005–2011. $*p < 0.05, **\ p < 0.01, ***p < 0.001$. Reported are the risk ratios from Fine and Gray competing risks model in which each risk is modeled separately. For example, entering eligibility due a loss of earnings is compared to leaving eligibility either for an increase of earnings/AGI above the maximum or family change. All variables reflect status in tax year 2005. Standard errors, clustered on the individual, appear in parentheses. Not reported are time-varying coefficients for sex, education, Black alone, other race, Hispanic, age, and baseline unemployment and EITC. Of the sample, 44.99 percent do not change from ineligible to eligible during the time period.

Table 8: Competing risks models estimating entry into EITC eligibility, women and men married at baseline

	Women			Men		
	Decreased income	Earnings>0	Family change	Decreased income	Earnings>0	Family change
Labor market variable						
Low education	0.76	0.85	1.30*	1.14	0.94	1.17
	(0.17)	(0.15)	(0.14)	(0.16)	(0.29)	(0.16)
Control variables						
Black alone	0.81	0.23***	4.78***	0.52*	0.28	3.27***
	(0.36)	(0.09)	(0.70)	(0.15)	(0.19)	(0.64)
Asian alone	0.84	1.02	0.89	0.78*	1.25	0.95
	(0.11)	(0.18)	(0.10)	(0.10)	(0.32)	(0.13)
Other race	1.39	0.98	1.60	0.82	1.17	1.56
	(0.96)	(0.53)	(0.45)	(0.38)	(0.91)	(0.50)
Hispanic	0.88	0.72	2.39***	0.83	0.78	2.47***
	(0.27)	(0.18)	(0.31)	(0.15)	(0.33)	(0.39)
Age	1.00	1.00	0.98**	1.00	1.01	0.98*
	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)
Number of children 0.70**	0.93	0.76***	0.92	0.82	0.74**	
	(0.08)	(0.08)	(0.06)	(0.06)	(0.11)	(0.08)
Unemployment rate 0.96	1.02	1.02	0.99	0.92	1.02	
	(0.10)	(0.09)	(0.05)	(0.07)	(0.13)	(0.07)
State EITC	0.94	1.42	0.84	1.12	1.25	0.92
	(0.22)	(0.28)	(0.10)	(0.17)	(0.42)	(0.14)
Percent entering	8.00	2.45	18.11	16.91	2.80	38.75
Observations		6,625			6,086	

Source: CPS ASEC 2006 linked with 1040 and W-2 data from 2005–2011. $*p < 0.05, **\ p < 0.01, ***p < 0.001$. Reported are the risk ratios from Fine and Gray competing risks model in which each risk is modeled separately. For example, leaving eligibility due to no earnings is compared to leaving eligibility either for earnings/AGI above the maximum or family change. All variables reflect status in tax year 2005. Standard errors, clustered on the individual, appear in parentheses. Not reported are time-varying coefficients for sex, education, Black alone, other race, Hispanic, age, and baseline unemployment and EITC. Of the sample, 33.6 percent never lost eligibility after gaining it.

Table 9: Patterns of eligibility loss through zero earnings and financial impact

Panel A. Unconditional probability of eligibility loss through zero earnings

	2006	2007	2008	2009	2010	2011
Unmarried, low-education women	0.06	0.05*	0.05*	0.06*	0.04*	0.03
Unmarried, high-education women	0.04*	0.04	0.03	0.04	0.03	0.03
Unmarried low-education men	0.05	0.04*	0.04*	0.06*	0.03	0.03
Unmarried high-education men	0.04	0.04*	0.04	0.05	0.03	0.03
Married, low-education women	0.06*	0.03	0.03	0.03*	0.03	0.03
Married, high-education women	0.04*	0.02*	0.02*	0.02*	0.02*	0.03
Married low-education men	0.08*	0.03	0.03	0.03	0.03	0.03
Married high-education men	0.05	0.02	0.03	0.03*	0.03	0.02

Panel B. Flows into and out of eligibility for low-education, unmarried women in the 2006 CPS ASEC, population counts

	2005	2006	2007	2008	2009	2010	2011
Became eligible in year	3,778,613	1,026,180	312,380	257,564	169,283	108,286	102,091
Ineligible in year+1		756,126	667,187	737,104	622,629	468,727	402,498
Flow		270,054	-354,807	-479,540	-453,346	-360,441	-300,407
Total eligible in year	3,778,613	4,048,667	3,928,920	3,683,572	3,535,220	3,305,665	3,179,383
Total losing eligibility due to earnings=0		347,653	266,599	287,230	370,551	282,087	169,492
Per person dollar loss based on 2005 start		$1,472	$1,706	$1,802	$2,017	$2,091	$2,221

Source: CPS ASEC 2006 linked with 1040 and W-2 data from 2005–2011. Panel A gives the unconditional probability that an EITC-eligible filer will exit eligibility through zero earnings in the period. An * notes probabilities that are statistically different from the average for all groups. Panel B reports the weighted numbers of women who have a high-school education or less in the CPS ASEC, file as single or head of household, and claim children at some time between 2005 and 2011. Dollars reflect the average EITC in 2005 for those in the "ineligible in year+1" category and are adjusted to 2011 dollars using the Bureau of Labor Statistics' Consumer Price Index.

Table 10: Competing risks models, unmarried and married, full sample with those who attrit retained

	Unmarried						Married					
		Women			Men			Women			Men	
	No earnings	Income >max	Family change	No earnings	Income >max	Family change	No earnings	Income >max	Family change	No earnings	Income >max	Family change
Main												
Low education	1.36***	0.47***	0.88**	1.09	0.68***	0.93	0.93	0.58***	0.73*	0.91	0.62***	0.72*
	(0.07)	(0.03)	(0.04)	(0.09)	(0.06)	(0.06)	(0.12)	(0.07)	(0.09)	(0.12)	(0.07)	(0.09)
Black alone	1.00	0.57**	0.47***	0.48**	0.37**	0.78	0.61*	0.81	0.68*	1.30	0.72	0.88
	(0.13)	(0.10)	(0.06)	(0.12)	(0.12)	(0.15)	(0.12)	(0.19)	(0.12)	(0.22)	(0.16)	(0.17)
Asian alone	2.70	3.25*	2.67*	0.87	0.92	1.19	1.14	1.13	0.67**	1.11	1.12	0.60***
	(1.52)	(1.90)	(1.04)	(0.43)	(0.48)	(0.75)	(0.14)	(0.11)	(0.09)	(0.14)	(0.11)	(0.09)
Other race	1.08	0.53	0.85	0.96	0.51	0.69	1.19	0.66	1.02	0.87	1.00	0.55
	(0.27)	(0.18)	(0.22)	(0.36)	(0.27)	(0.21)	(0.34)	(0.19)	(0.31)	(0.21)	(0.33)	(0.17)
Hispanic	0.99	1.01	0.81***	1.07	0.97	0.71***	0.86	0.72*	0.77	0.88	0.64**	0.87
	(0.07)	(0.09)	(0.05)	(0.10)	(0.11)	(0.07)	(0.13)	(0.11)	(0.12)	(0.13)	(0.09)	(0.13)
Age	1.02**	1.02	1.01	1.05***	1.02	1.00	1.02**	1.02**	1.02**	1.02**	1.01*	1.01
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Number of children	1.12	0.63***	0.66***	0.90	0.57***	0.83	1.12	1.02	1.07	1.07	1.11	1.03
	(0.08)	(0.06)	(0.05)	(0.11)	(0.08)	(0.09)	(0.08)	(0.06)	(0.07)	(0.07)	(0.07)	(0.07)
Unemployment rate	0.95	1.06	0.90	1.00	0.92	0.94	1.01	1.10	0.98	0.95	1.16**	1.00
	(0.06)	(0.08)	(0.05)	(0.10)	(0.11)	(0.07)	(0.06)	(0.07)	(0.06)	(0.06)	(0.07)	(0.07)
State EITC	0.93	1.10	0.94	1.04	0.93	0.95	0.79	1.12	0.95	0.85	1.07	0.90
	(0.12)	(0.19)	(0.11)	(0.21)	(0.23)	(0.17)	(0.11)	(0.14)	(0.12)	(0.11)	(0.13)	(0.12)
Percent failing	24.26	13.87	28.21	24.27	19.25	31.60	18.34	25.75	24.52	21.02	27.95	24.90
Observations		6,689			2,649			7,161			6,755	

Source: CPS ASEC 2006 linked with 1040 and W-2 data from 2005–2011. $*p < 0.05, **p < 0.01, ***p < 0.001$. Reported are the risk ratios from Fine and Gray competing risks model in which each risk is modeled separately. For example, leaving eligibility due to no earnings is compared to leaving eligibility either for earnings/AGI above the maximum or family change. All variables reflect status in tax year 2005. Standard errors, clustered on the individual, appear in parentheses. Not reported are time-varying coefficients for sex, education, Black alone, other race, Hispanic, age, and baseline unemployment and EITC. For the subsamples, 33.7 percent of unmarried women, 24.9 percent of unmarried men, 31.4 percent of married women, and 26.1 percent of married men never lost eligibility after gaining it Sample includes modeled tax filers from the CPS ASEC 2006 who are not found in further tax data after a given year (see text for description).

Table 11: Cox proportional hazard models: Unmarried and married, full sample

	Unmarried						Married					
	Women			Men			Women			Men		
	No earnings	Income >max	Family change	No earnings	Income >max	Family change	No earnings	Income >max	Family change	No earnings	Income >max	Family change
Main												
Low education	1.18*	0.50***	0.92*	0.61	0.47**	0.77	1.06	0.65***	0.94	1.05	0.69***	0.94
	(0.08)	(0.03)	(0.04)	(0.18)	(0.12)	(0.14)	(0.07)	(0.03)	(0.04)	(0.07)	(0.03)	(0.04)
Black alone	0.98	0.57**	0.46***	0.58	0.35**	0.74	1.05	0.60***	0.92	1.32**	0.73***	1.02
	(0.19)	(0.11)	(0.06)	(0.19)	(0.11)	(0.15)	(0.11)	(0.05)	(0.06)	(0.13)	(0.06)	(0.08)
Asian alone	3.73	3.06	2.92**	0.92	1.29	0.54**	1.08	1.10	0.71**	1.00	1.09	0.65**
	(2.59)	(1.88)	(1.20)	(0.27)	(0.24)	(0.13)	(0.17)	(0.10)	(0.08)	(0.16)	(0.10)	(0.09)
Other race	1.29	0.79	0.93	0.55	0.47	0.70	0.88	0.58	0.87	0.66	1.08	0.50*
	(0.17)	(0.12)	(0.09)	(0.30)	(0.26)	(0.22)	(0.33)	(0.17)	(0.28)	(0.23)	(0.36)	(0.17)
Hispanic	1.03	0.99	0.84**	1.17	0.97	0.70***	0.76	0.59***	0.65**	0.82	0.53***	0.79
	(0.09)	(0.08)	(0.05)	(0.14)	(0.10)	(0.06)	(0.13)	(0.09)	(0.10)	(0.14)	(0.07)	(0.12)
Age	1.00	1.02*	1.02**	1.04*	1.03	1.00	1.01	1.03***	1.02**	1.02**	1.02**	1.01
	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Number of children	0.94	0.56***	0.59***	0.72	0.47***	0.72**	1.09	1.06	1.15	1.10	1.16*	1.07
	(0.11)	(0.06)	(0.05)	(0.13)	(0.08)	(0.09)	(0.12)	(0.08)	(0.10)	(0.11)	(0.09)	(0.09)
Unemployment rate	0.92	1.07	0.92	0.99	0.93	0.90	0.95	1.08	0.95	0.89	1.12	0.97
	(0.08)	(0.08)	(0.06)	(0.14)	(0.11)	(0.07)	(0.08)	(0.07)	(0.06)	(0.07)	(0.07)	(0.07)
State EITC	0.82	1.08	0.98	0.83	0.93	0.90	0.77	1.12	0.99	0.75	1.08	0.95
	(0.16)	(0.19)	(0.12)	(0.25)	(0.24)	(0.18)	(0.14)	(0.15)	(0.14)	(0.13)	(0.14)	(0.14)
Percent failing	24.26	13.87	28.21	24.27	19.25	31.60	18.34	25.75	24.52	21.02	27.95	24.90
Observations		6,689			2,649			7,161			6,755	

Source: CPS ASEC 2006 linked with 1040 and W-2 data from 2005–2011. $*p < 0.05, **p < 0.01, ***p < 0.001$. Reported are the risk ratio from a Cox competing risks model in which the data are duplicated (see text for details). Each risk is examined separately while other risks are treated as censored. Standard errors appear in parentheses. For the subsamples, 33.7 percent of unmarried women, 24.9 percent of unmarried men, 31.4 percent of married women, and 26.1 percent of married men never lost eligibility after gaining it.

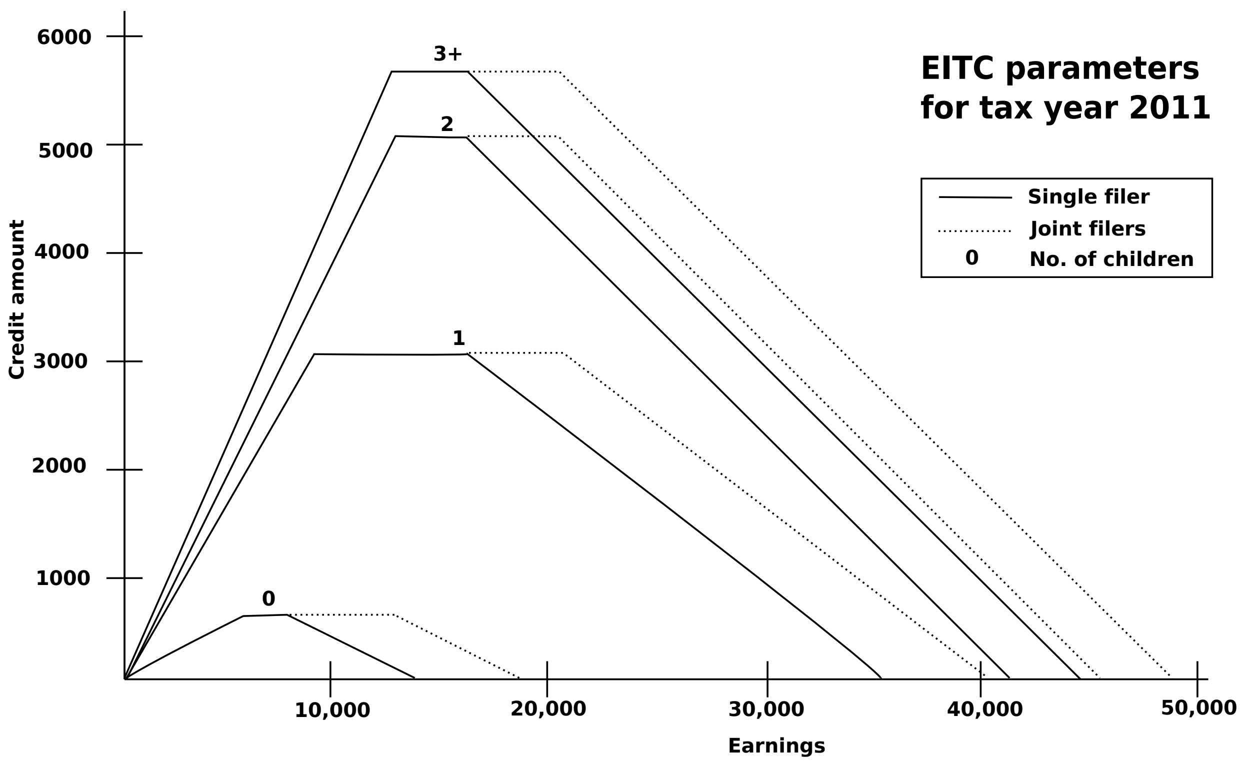

Figure 1: EITC parameters for tax year 2011. The graph gives a clear picture of what may cause a transition from eligibility to ineligibility from one year to another. First, earnings may fall to 0 due to full-year unemployment. Second, earnings or income may increase beyond the threshold. Third, earnings and income for a tax filer may remain constant, but the filer loses dependents or changes marital status.