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ABSTRACT

The Omnibus Budget Reconciliation Act of 1990 introduced a refundable tax credit for low-income working families who purchased health insurance coverage for their children. This health insurance tax credit (HITC) existed during tax years 1991, 1992, and 1993, and was then rescinded. We use Current Population Survey data and a difference-in-differences approach to estimate the HITC's effect on private health insurance coverage of low-earning single mothers. The findings suggest that during 1991–1993, the health insurance coverage of single mothers was about 6 percentage points higher than it would have been in the absence of the HITC.

JEL Classification Codes: H2, H51, I18, J32

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1. Introduction

A longstanding proposal for increasing the health insurance coverage of low-earning workers is to subsidize their coverage through a refundable tax credit (Pauly 1999; Cogan, Hubbard, and Kessler 2005; Antos and Miller 2010). Such a Health Insurance Tax Credit (HITC) would grant a tax credit up to a specified maximum—for example, $2,000 for an individual and $5,000 for a family—on a tax return where the filer purchased a private health insurance policy (either employer-provided or in the market). The HITC was included in several of President George W. Bush's budget proposals (McClellan and Baicker 2002; Office of Management and Budget 2005) and is a centerpiece of Wisconsin Representative Paul D. Ryan's "Roadmap for America's Future" (Ryan 2010). By reducing the price of health insurance to eligible taxpayers and extending the tax-favored treatment of health insurance to workers who lack access to employer-provided health insurance, the HITC would be expected to increase the percentage of low-income individuals and families covered by private health insurance.

But controversy has surrounded the extent to which workers would take up an HITC and hence whether an HITC would reduce the number of uninsured individuals. Pauly and Herring (2001, 2002), Pauly, Song, and Herring (2001), and Wozniak and Emmons (2000) simulated a variety of HITC policies and found that a "reasonably generous" credit could reduce the number of uninsured individuals by roughly 50 percent. However, simulations by Gruber (2000a,b) suggested that the HITC might reduce the number of uninsured by only about 10 percent.[1]

[1] Pauly, Song, and Herring (2001) also note that health insurance tax credits could lead to broader changes in health insurance markets—including greater price competition among insurers—that are not accounted for in simulation models. Holtzblatt (2008) and Furman (2008) compare the HITC with other tax-based approaches to reforming health insurance.

A number of careful empirical studies have estimated workers' responsiveness to health insurance subsidies and changes in premiums, but the range of estimates produced by these studies is disturbingly wide. Estimated price elasticities of health insurance take-up range from close to 0 (Chernew, Frick, and McLaughlin 1997; Blumberg, Nichols, and Banthin 2001) to approximately –1.2 (Gruber and Poterba 1994; Marquis et al. 2005). We review these studies more fully in the conclusion, but the wide range of estimates suggests a useful role for further research on how workers would respond to health insurance subsidies.

In this paper, our goal is to obtain direct evidence on the effect of refundable tax credits on the health insurance coverage of low-earning workers. We do this by examining a supplemental credit that Congress added to the Earned Income Tax Credit (EITC) during 1991, 1992, and 1993. This policy provided a refundable tax credit of up to $428 in 1991 ($451 in 1992, and $465 in 1993) to EITC-eligible households that bought health insurance for a qualifying child. We treat this supplemental credit as a natural experiment and estimate its impact on the health insurance coverage of working single mothers using a difference-in-differences approach applied to Current Population Survey data.

We first describe the tax credit (section 2), the approach to estimation (section 3), and the data we use (section 4). We then present the main findings (section 5), which suggest that the HITC increased private health insurance coverage of low-education working single mothers by about 6 percentage points (from a base of slightly less than 16 percent). Section 5 also reports several falsification (or "placebo") tests. Section 6 reports a range of sensitivity tests, including checks for whether the HITC affected labor supply, and whether the findings could be biased by Medicaid crowd-out, state-level economic conditions, state AFDC waivers that preceded federal welfare reform, and state-level EITCs. Section 7 concludes by examining whether the implications of the findings make sense in light of the number of HITCs that were actually claimed, and by comparing the implications of the HITC findings with existing estimates of the price elasticity of health insurance take-up.

2. Background: The Health Insurance Tax Credit, 1991–1993

When Congress passed the Omnibus Budget Reconciliation Act (OBRA) of 1990, it added a supplemental credit for health insurance purchases to the basic Earned Income Tax Credit (EITC) program (U.S. Government Accountability Office 1991, 1993). This HITC was a refundable tax credit for low-income workers with one or more children who bought health insurance—either employer-provided or private nongroup—covering the child or children. The credit offset only the cost of health insurance and did not cover co-payments, deductibles, or out-of-pocket health expenses. To encourage participation, the credit was refundable, so taxpayers with no federal income tax liability could still receive a payment from the Internal Revenue Service. The HITC was repealed effective December 31, 1993, so it was available only during tax years 1991, 1992, and 1993.[2]

The HITC had the same eligibility criteria as the EITC: To receive a credit, a household needed earnings and a qualifying child. To qualify, a child had to meet three requirements: (1) be a child, stepchild, grandchild, or foster or adopted child of the taxpayer; (2) have the same place of residence as the taxpayer for more than half the tax

[2] See U.S. Government Accountability Office (1991, 1993, 1994) for discussions of why Congress eliminated the credit. We touch on these reasons later in this section and in the conclusion.

year; and (3) be under age 19 (or 24 if a full-time student) or be permanently disabled. Unlike the basic EITC, the HITC remained the same regardless of the number of qualifying children in the family.

The HITC schedule closely followed the EITC's.[3] In 1991, a taxpayer with earnings and a qualifying child could receive a credit up to $428 if he or she bought private health insurance that covered the child. For households with earned incomes of $1 to $7,140, the credit was 6 percent of earned income. For households with earnings between $7,140 and $11,250, the credit was $428 (6 percent of $7,140). For households with earnings between $11,250 and $21,250, the credit phased out at a rate of 4.28 percent per marginal dollar earned and fell to $0 for earnings at or above $21,250 (see Figure A1). In 1991, the maximum HITC of $428 was 36 percent of the maximum EITC of $1,192. Like the EITC schedule, the HITC schedule was indexed to inflation.

In 1991, the HITC's first year, the average credit was $233, or 23 percent of the reported average annual health insurance premium of $1,029. Also in 1991, 2.3 million taxpayers received health insurance credits of $496 million (U.S. Government Accountability Office 1991). A U.S. Government Accountability Office (GAO) study (1994) estimated that the take-up rate for the HITC in 1991 (its first year) was in the range of 19 to 26 percent.[4] In contrast, the take-up rate for the regular EITC was 80 to 86 percent. The GAO attributed the relatively low HITC take-up rate to two factors. First,

[3] Table A1 in the Appendix gives details of the HITC schedules and of the EITC schedules in the years before, during, and after the HITC existed. Figure A1 in the Appendix is a graphical representation of the HITC and EITC schedules in 1991.

[4] This take-up rate is based on GAO's (1994) estimate from the March 1992 CPS that about 8.8 million families were eligible for the HITC in 1991; however, our tabulations suggest that this is an estimate not of the number of eligible *families* but eligible *individuals*. Our tabulations suggest that 5.2 families were eligible for the HITC in 1991, which implies an HITC take-up rate of about 44 percent. This latter estimate seems reasonable in light of Scholz's (1994) estimate that the 1990 EITC take-up rate was 80 to 86 percent. We thank Paul Fronstin for advice on this point.

interviews with taxpayers at IRS service sites in six cities suggested that fewer than 30 percent of EITC-eligible taxpayers knew the credit existed. Second, the GAO suggested the credit was too modest to induce low-income workers to buy health insurance.

The GAO's findings appear to have played a role in persuading Congress to eliminate the HITC in 1993 (effective 1994). However, as the evidence below suggests, the implicit conclusion that the HITC was ineffective may have been premature. Given the attention paid to tax credits for health insurance in the intervening years, it is curious that no analysis of the HITC's impact on health insurance coverage appears to exist.

3. Approach to Estimation

We treat the HITC as a natural experiment and adopt a difference-in-differences approach to estimating its effects on the health insurance coverage of working single mothers with less than a high school education—hereafter, *low-education working single mothers*. The approach follows a large literature on the labor supply effects of the EITC including Eissa and Leibman (1996), Meyer and Rosenbaum (2000), Eissa and Hoynes (2004), and Hotz, Mullin, and Scholz (2006).

The population potentially affected by the HITC consisted of low-income working families with children. If the HITC had any effect on private health insurance coverage, then the coverage of low-income working families with children would have been greater than otherwise during 1991, 1992, and 1993. For three reasons, we focus on the HITC's possible effect on private health insurance of low-education working single mothers. First, working single mothers were roughly 44 percent of all EITC-eligible households in 1990, making them the largest group of taxpayers eligible for the EITC and hence for the HITC (Liebman 2000). Second, for households headed by a single woman, we can plausibly abstract from decisions made jointly with other family members (Eissa and Liebman 1996). Third, by focusing on high school dropouts, we can estimate the effect of the HITC on a group that is likely to be eligible (because it is likely to have low earnings) without conditioning explicitly on income or earnings. Conditioning on income or earnings is ruled out because the EITC creates incentives for earners to change their hours of work so as to qualify for the credit (Eissa and Hoynes 2006). Accordingly, our main “treatment” group is low-education working single mothers.

A convincing difference-in-differences analysis requires a comparison or “control” group that is as similar as possible to the treatment group without being eligible for the HITC. In particular, we want a control group whose trend in health insurance coverage would have been the same as low-education working single mothers’ trend if the latter had not been affected by the HITC (Angrist and Krueger 1999; Blundell and MaCurdy 1999). Following Eissa and Liebman (1996), we choose *working single women without children and with less than high school* as the comparison group for low-education working single mothers because, although they do not have children (and hence are ineligible for the EITC and HITC), they face essentially similar labor markets and economic conditions. As Figure 1 shows, the two groups had similar trends in health insurance coverage in the years preceding the HITC.

Two concerns invariably arise with the difference-in-differences approach. First, if the treatment and control groups do differ in their characteristics, each may be affected differently by contemporaneous shocks (other than the HITC). In this case, the difference-in-differences approach may still be valid if we can control convincingly for observables that capture characteristics of individuals that are correlated with health insurance coverage. Second, the difference-in-differences estimator may be contaminated if the compositions of the treatment and control groups change over time. In the present case, changes in tax and welfare programs increased the work incentives of single mothers between 1984 and 1996 (Meyer and Rosenbaum 2000, 2001). If single mothers who entered the labor force in later years were more likely to work part-time and hence less likely to have employer-provided health insurance, then changes in the characteristics of single mothers could have blunted any rise in private health insurance that may have occurred as a result of the HITC. We can again mitigate this problem by controlling for observable characteristics using regression, but it will also be important to examine the samples carefully to see whether and how much they did in fact change over time. We do this below in section 6.1.

In the case of the HITC, a third issue arises—the treatment is in fact a combination of the HITC and increases in the EITC. As Appendix Table A1 shows, in the years preceding adoption of the HITC (1988–1990), the maximum credit under the EITC increased by 3, 4, and 5 percent per year ($23, $36, and $43). In contrast, during the HITC years (1991–1993) the maximum EITC increased by 25, 8, and 11 percent per year ($239, $89, and $110). Clearly, the increased generosity of the EITC may have contributed to any effect we attribute to the HITC—the specific subsidy of $428 to $465 targeted on private health insurance—and we have no ready method for disentangling the contributions of each. Accordingly, when we refer to an HITC effect below, it should be understood that we are referring to an affect of the HITC combined with the above increases in the EITC.

The model we estimate can be written:

$$\Pr(covered_i = 1|\bullet) = F[\beta_0 + \beta_1\, treatgroup_i + \beta_2\, HITC_t + \beta_3\, treatgroup_i \times HITC_t + X_i\, \beta] \quad (1)$$

where i indexes individuals and t indexes years; $covered_i$ is a binary indicator of private health insurance coverage; $treatgroup_i$ equals one for a working single woman with a dependent child and less than high school, and zero otherwise; $HITC_t$ equals one for 1991, 1992, and 1993 (years during which the HITC was in effect), and zero for 1988, 1989, and 1990 (years before the HITC was in effect); and $treatgroup_i \times HITC_t$ captures the change in coverage rates for working single mothers, relative to working single women without children, after the HITC took effect.

In the basic specification we estimate, the vector of controls X_i includes age, indicators of race (white, black, and other), indicators of number of children (under age 6, aged 6–18, and aged 19–24 *and* a full-time student), indicators of work status (full-time/full-year, full-time/part-year, part-time/full-year, and part-time/part-year),[5] earned income,[6] and unearned income.[7] For some of the specification tests reported in Table 8, we include additional controls, as described later. We let F denote the standard normal cumulative density and estimate equation (1) as a probit.

4. Data

We estimate equation (1) using data from the March 1989–1994 Annual Demographic Supplements to the Current Population Survey (CPS), which provide

[5] Full-year work implies at least 50 weeks of work in the previous year. Full-time work implies usual weekly hours of 35 or more in the previous year.
[6] Earned income includes income from wages, salaries, and self-employment.
[7] Unearned income includes income from unemployment compensation, worker's compensation, social security or railroad retirement, supplemental security, public assistance or welfare, veteran payments, survivor benefits, disability, retirement funds, interest, dividends, rent, educational assistance, child support, alimony, contributions, financial assistance from friends, and other nonearnings.

information for tax years 1988 through 1993. Respondents to the March 1989, 1990, and 1991 CPS constitute a before-HITC sample (tax years 1988, 1989, and 1990). Respondents to the March 1992, 1993, and 1994 CPS constitute a during-HITC sample (tax years 1991, 1992, and 1993). The relevant unit of observation is the tax-filing unit, which in the CPS implies allocating primary families and subfamilies to separate tax-filing units.

The sample includes women aged 19 to 44[8] who worked (had annual hours greater than zero), were single (never married, widowed, or divorced), and had less than a high school education. We exclude women who reported negative earnings, those in school full-time, those who were separated from their spouses, and those who reported being ill or disabled. The resulting sample, after pooling all six years, includes 3,661 observations.

We allocate working single women with at least one dependent child to the treatment group, and working single women without a child to the control group. We consider any child in the tax-filing unit who was under age 19 (or under age 24 if a full-time student) to be a dependent child for tax purposes. Consistent with the EITC literature, we do not try to impose the support or residency test for HITC eligibility.[9]

For two reasons, we limit the analysis to 1988 through 1993 and do not attempt to estimate the effect of the HITC's repeal. First, when Congress passed the OBRA of 1993, which repealed the HITC, it enacted the largest expansion of the EITC in the credit's

[8] We restrict the sample to women aged 19 to 44 because women in this age range are most likely to have at least one dependent child (a child under age 19 or under age 24 and a full-time student who lives at home). When women aged 45 to 54 are added to the sample, the findings are essentially similar (available on request).
[9] This is mainly due to limitations of the CPS. However, using data from the SIPP and IRS, Scholz (1994) shows that the support test does not greatly change estimates of EITC eligibility.

history (see Appendix Table A1; Baughman and Dickert-Conlin 2003). In 1993, a mother of one child with earnings up to $7,750 could receive a credit of 18.5 percent of earned income, resulting in a maximum credit of $1,434. In 1994, the credit rate rose to 26.3 percent of earned income, resulting in a maximum credit of $2,038. Also beginning in 1994, eligibility for the credit was expanded to include families with no children, for whom the credit was 7.65 percent of earnings up to $4,000 (for a maximum 1994 credit of $306). As a result, the effects of repealing the HITC are confounded with the EITC expansion of 1994–1996.

The second reason for restricting the analysis to 1988 through 1993 is that the CPS remained unchanged throughout this period. In March 1988, the Bureau of Labor Statistics modified the CPS health insurance questions to capture more accurately the insurance coverage of dependents. The next important revisions to the CPS health insurance questions occurred in March 1995, when BLS introduced a more detailed set of health insurance questions. This led to an increase in the number of persons reporting employer-provided coverage.[10]

During the years we examine, the CPS health insurance questions read as follows:

> 75A. Other than government sponsored policies, health insurance can be obtained privately or through a current or former employer or union. Was anyone in this household covered by health insurance of this type at any time during 19xx [last year]?
>
> 75B. Who was that?
>
> 75C. Was …'s health insurance coverage from a plan in …'s own name?

[10] In particular, previous surveys asked about employer coverage as a subset of private coverage, but beginning in March 1995 the survey asked separate questions about employer-provided and other types of private health insurance. Appendix T of Unicon (2005) is a good discussion of changes in the CPS health insurance questions.

75F. What other persons were covered by this health insurance policy? Possible answers are Spouse, Children in household, Children not in the household, Other, and No one.

These questions allow us to define three alternative measures of health insurance:

1. **private insurance coverage**, defined broadly to include coverage by a privately purchased or employer-provided health insurance plan, whether or not in the respondent's own name (that is, positive responses to questions 75A and 75B)
2. private insurance **in the respondent's own name** (a subset of the first definition because it implies a positive response to question 75C)
3. private insurance in the respondent's own name that **covers children in household** (a subset of the second definition because it implies a "children in household" response to question 75F)

Table 1 shows average private health insurance coverage rates for working single mothers and working single women without children during 1988–1993, using the measures of coverage defined above. For single mothers, the rate of insurance coverage in the respondent's own name fell by 5.4 percentage points (from 32.3 to 27 percent); however, only about one-fifth of this decrease occurred after 1990. For single women without children, the coverage rate also fell somewhat between 1988 and 1990, but fell sharply after 1990 (from 37.8 to 20.9 percent). A likely explanation for the continued drop after 1990 is the recession of 1991, which would have reduced both employment and access to employer-provided health insurance of single women.

In the following analysis, the outcome of interest is coverage by private health insurance, defined as *whether a working single woman has private insurance in her own name that covers her child or children*. We focus on this outcome because the HITC could be used only to purchase a health insurance policy—either in the market or through an employer or union—covering a qualifying child. Figure 1 graphs the coverage rates for single mothers and single women without children.[11]

Table 2 displays mean characteristics of working single mothers and working single women without children pooled for 1988–1993. The two groups differ in some important ways. Relative to single women without children, single mothers are more likely to be black (33.6 versus 14.7 percent) and less likely to work full-time, full-year (37.8 versus 49.1 percent). Also, single mothers have lower average earnings than single women without children ($7,257 versus $8,686).

Table 3 displays summary statistics for both single mothers and single women without children in the before-HITC and during-HITC periods. With one exception, the characteristics of both single mothers and single women without children appear fairly stable. The exception is that the percentage of women (both single mothers and single women without children) working full-time, full-year dropped by 2.5 percentage points during the HITC period, reflecting the recession of 1991–1992.

5. Empirical Findings

5.1. Main findings—single women with less than high school

Table 4 displays the average private health insurance coverage rates for single mothers and single women without children in the years before and during the HITC. The first row shows that health insurance coverage for single mothers fell by 2.4 percentage

[11] Note that the private health insurance variable is constructed as "private insurance in own name that covers children" for single mothers and as "private insurance in own name" for single women without children. In an earlier version of this paper, we defined the private insurance variable as "private insurance in own name" for both single mothers and single women without children. The estimates we obtained (available on request) are essentially similar to those we report in this section and the next.

points between 1988–1990 and 1991–1993. The second row shows that, over the same period, coverage fell for single women without children by 9 percentage points. The implication is that, after netting out the declining trend in insurance coverage, the private health insurance coverage of single mothers was higher by 6.5 percentage points than it would have been without the HITC. The bootstrap standard error of this point estimate is 0.031.

Table 5 displays estimates of the key coefficients in equation (1).[12] Specification 1 includes no control variables. Specification 2 controls for age, race (white, black, and other), number of children under age 6, number of children aged 6–18, and number of children aged 19–24 *and* a full-time student. Specification 3 controls in addition for work status (full-time/full-year, full-time/part-year, part-time/full-year, and part-time/part-year), earned income, and unearned income.

In specification 1, the coefficient on *treatgroup* is –0.145 and statistically significant (p-value = 0.000). With the addition of demographic characteristics in specification 2, it falls to –0.080 (p-value = 0.006). In specification 3, when we control for work status and income along with demographic characteristics, it changes slightly from –0.080 to –0.084 (p-value = 0.000). That the coefficient on *treatgroup* falls as controls are added to the model (between specifications 1 and 2) suggests that observable characteristics other than the presence of children are important in explaining the difference between single mothers and single women without children in private health insurance coverage.

[12] As described in the appendix, we follow DeLeire (2004) and obtain the probit DD estimator as the discrete double difference of the standard normal cumulative distribution function, averaged over all observations in the sample.

In specification 1, the coefficient on *HITC* is negative (–0.090) and statistically significant (*p*-value = 0.000). This is consistent with the declining trend in average health insurance coverage for both single mothers and single women without children. Including additional controls (specifications 2 and 3) leaves this estimate essentially unchanged.

The estimate of main interest is the coefficient on the interaction term. This is essentially similar in size and statistical significance across the three specifications. The estimate from specification 3 estimate is 0.063 (*p*-value = 0.019), which suggests that private health insurance coverage of working single mothers with less than high school was higher by 6.3 percentage points than it would have been without the HITC.[13]

5.2. Findings for single women with more education

We would expect the estimated effect of the HITC on single mothers who completed high school (or who have some college or completed college) to be smaller than its effect on single mothers with less than high school for a simple reason: Single mothers with more education are less likely to be low earners and hence less likely to be eligible for the EITC and HITC. Accordingly, single mothers with more education offer a useful set of falsification (or "placebo") tests of the estimates in Table 5: Finding that single mothers with more education experienced coverage increases similar to those of single mothers with less than high school would cast doubt on the findings in Table 5.

Table 6 displays tests of the hypothesis that single mothers with more education were less likely than single mothers with less than high school to experience an HITC-

[13] Eligibility for the HITC depended on the presence of a dependent child, so the credit could have affected women's fertility decisions. (Hotz and Scholz [2003] discuss fertility and the EITC.) To test whether the findings are sensitive to the assumption that the presence of children is exogenous, we reestimated the basic specification (Table 5, specification 3) after dropping single mothers with children younger than age 3 from the treatment group. The findings are robust to this deletion.

induced increase in health insurance coverage. We do this by estimating equation (1) using samples of women with successively more educational attainment. The specifications are the same as specification 3 of Table 5; that is, they include demographic characteristics, indicators of work status, earned income, and unearned income as controls. We refer to this specification as the "basic specification."

The first row of Table 6 repeats the main finding from comparing single mothers and single women with *less* than high school (specification 3 of Table 5). The second row of Table 6 shows the results of comparing single mothers and single women *with* a high school education. The estimate on the interaction term is 0.043 with a p-value of 0.002, which suggests that the private health insurance coverage rate of single mothers with high school was 4.3 percentage points higher than it would have been without the HITC. This accords with the expectation that the HITC effect on single mothers should be less as educational attainment increases.

The third and fourth rows of Table 6 show findings from comparisons for single mothers with some college and with a college degree or more. As expected, these comparisons produce still smaller estimated effects of the HITC on health insurance coverage. For single mothers with some college, the estimated HITC effect on health insurance coverage is 2.5 percentage points, but the estimate is imprecise (row 3 of Table 6). The estimated effect of the HITC on health insurance coverage of single women with college or more is –0.006 percentage points—essentially zero. That these falsification tests behave as expected, with negligible responses to the HITC among single mothers who are increasing unlikely to be eligible for the HITC, tends to increase our confidence that Table 5 is identifying an HITC effect for low-education single women.

5.3. Findings by work status

As another falsification test, we explore whether the effect of the HITC differed by work status. Because non-workers were not eligible for the HITC, we would not expect to observe HITC effects on the health insurance coverage of non-working single mothers. To test this hypothesis, we estimated equation (1) using a sample that includes 3,136 non-working single mothers with less than high school in addition to the 3,661 working single mothers used to obtain Table 5's estimates. We allowed the HITC to have a differential effect on the health insurance coverage of working single mothers and non-working single mothers. The estimated HITC effect for non-working single mothers (not reported in a table) is –0.004 percentage points (*p*-value = 0.718), compared with an estimated effect for working single mothers of 6.5 percentage points (*p*-value = 0.018). Accordingly, the HITC regime appears to have had no effect on a group of HITC-ineligible women, many of whom would likely have been eligible had they worked. (We explore the possible effect of the HITC on labor supply below.)

5.4. Findings disaggregated by year

The estimates in Tables 5 and 6 are restrictive because they aggregate the three before-HITC years and the three during-HITC years. We would like to know whether it is reasonable to restrict the estimated effect of the HITC to be the same in the three during-HITC years (1991, 1992, and 1993). Accordingly, we estimate a model that includes a set of year dummies (for 1988, 1989, 1991, 1992, and 1993), a treatment group dummy, and interactions of the treatment and year dummies. We estimate this model for the main group of interest—low-education working single women—and report estimates from a specification analogous to the basic specification. The estimates are year-specific difference-in-differences estimates of the HITC effect on private health insurance coverage rates.

The estimates, reported in Table 7, suggest that the 1991, 1992, and 1993 coverage rates of single mothers with less than high school were higher by 6.4, 7.2, and 11 percentage points than they would have been in the absence of the HITC (although only the estimate for 1993 has a *p*-value less than 0.05). The pattern of these estimates—increasing over time—has at least two possible interpretations. First, it could be that the impact of the HITC increased as the existence of the program became known and its implications better understood. This would be consistent with the pattern of participation in several social programs (Madrian and Shea 2001; Remler and Glied 2003; Currie 2006). Alternatively, it could be that we have omitted one or more variables that affected the health insurance coverage of low-education single mothers (but not other low-education single women). We investigate the latter possibility next.

6. Sensitivity Tests

This section describes findings from four sets of sensitivity tests. We begin by exploring possible effects of the HITC on labor supply. We then attempt to control for factors affecting private health insurance in the early 1990s other than the HITC: (1) expansion of the Medicaid program, (2) state-level economic conditions and other state-specific effects, and (3) welfare reform and the introduction of state-level EITCs.

6.1. The HITC and labor supply

Several empirical studies have found that the EITC substantially increases labor force participation of eligible groups (see the reviews by Hotz and Scholz 2003 and Eissa and Hoynes 2006). Because the HITC's structure was essentially similar to the EITC's, the HITC may also have drawn women into the labor force. If so, then the composition of working single mothers during the HITC years may have differed from the composition before the HITC. Such compositional change, if substantial, could contaminate the difference-in-differences estimator.

To examine whether the HITC substantially affected labor force participation we estimate a model along the lines of equation (1), but using *employment* (a dummy variable equal to 1 if a woman reported working at least one hour during the year, 0 otherwise) as the dependent variable. The specification is the same as specification 1 of Table 5. The estimates, reported in row 2 of Table 8, suggest that during the HITC years, labor force participation of low-education single mothers was higher by 2.1 percentage points; however, the estimated HITC effect is insignificant at conventional levels (*p*-value = 0.40). It follows that, if the HITC did induce a change in the composition of working single mothers, it was likely small and hence unlikely to overturn the main finding from Table 5.

6.2. Medicaid crowd-out

During the years we are examining, eligibility for Medicaid expanded substantially to include low-income pregnant women and children with no ties to AFDC (see Appendix Table A2). Because Medicaid and private health insurance are potential substitutes—they offer similar health coverage, and Medicaid is much less costly—Medicaid expansion may have drawn some low-income single mothers out of private health insurance and into Medicaid. Cutler and Gruber (1996) first referred to such substitution as "crowding out," and to the extent it exists, the estimates in Tables 5, 6, and

7 could understate the HITC's effect on private health insurance coverage. Alternatively, the availability of the HITC could conceivably have drawn some low-income single mothers out of Medicaid and into private health insurance, in which case the estimates in Table 5 would overstate the net impact of the HITC on health insurance coverage.

To address these concerns, we estimate variants of equation (1) that use two alternative dependent variables: *medicaid,* a dummy indicator of whether a woman had Medicaid coverage, and *insured*, a dummy indicator of whether a woman had coverage from either Medicaid or private health insurance. In the first case, the question addressed is whether Medicaid coverage of low-education single mothers changed (relative to low-education single women without children) during the HITC years. In the second case, the question addressed is whether overall health insurance coverage of low-education single mothers changed (relative to low-education single women without children) during the HITC years.

Row 3 of Table 8 displays findings from the model in which *medicaid* is the dependent variable. The estimates offer only marginal evidence that relative Medicaid coverage of low-education single mothers was higher during the HITC period than in the prior years. (The point estimate is 1.4 percentage points, but the estimate is imprecise.) We interpret this finding as evidence that, if the Medicaid expansions did crowd out private health insurance during the HITC years, crowd-out was slight.

Row 4 of Table 8 displays findings from the model in which *insured* (by either private health insurance or Medicaid) is the dependent variable. The estimates suggest that during the HITC period, relative net health insurance coverage of low-education single mothers was higher by 8.4 percentage points (p-value = 0.007) than in the preceding years. Estimates from the basic specification in row 1 (from Table 5, specification 3) suggest that the HITC was the main factor in this net increase in health insurance coverage, and that Medicaid expansion played a relatively minor role. Together, rows 1, 3, and 4 suggest that 6.3 of the 8.4 percentage points were due to an increase in private coverage, and 1.4 percentage points were due to an increase in Medicaid coverage (although the latter is not statistically significant). A relatively small residual (0.7 percentage points) cannot be explained by either the HITC or changes in Medicaid.

6.3. State-level economic conditions and state fixed effects

It is also possible that changes in private health insurance coverage of low-education single mothers were related to state-level economic conditions or to state fixed effects. To account for these possibilities, rows 5, 6, and 7 of Table 8 report estimates from models that successively add to the basic specification a set of state dummy variables (to control for time-invariant state-level effects on health coverage), the year-specific contemporaneous state unemployment rate (to control for cyclical labor market influences on health insurance coverage), and both state fixed effects and the year-specific state unemployment rate.

The findings reported in rows 5, 6, and 7 of Table 8 suggest that including these state-level controls produces estimates that are essentially similar to the basic specification—an increase in single mothers' private health insurance coverage of roughly 6.5 percentage points (with *p*-values less than 0.02).

6.4. Welfare reform and state EITCs

Although the findings in Table 5 are consistent with the HITC increasing private health insurance coverage of low-education single mothers, an alternative explanation for this increase is the welfare reforms adopted by some states after 1990. California, Michigan, New Jersey, Oregon, and Utah all implemented welfare waivers in 1993 (DeLeire, Levine, and Levy 2006). These waivers changed the nature of AFDC by imposing time limits on AFDC participation and introducing work requirements for AFDC participants (Meyer and Rosenbaum 2000). If these changes moved single mothers into the labor market and onto private (employer-provided) health insurance, the estimates in Tables 5, 6, and 7 could overstate the HITC's effect on private health insurance coverage.

To address this possibility, we restrict the sample to low-education single women in states that did not have welfare waivers. Row 8 of Table 8 shows that restricting the sample in this way leaves the estimated HITC impact on single mothers essentially unchanged—the private health insurance coverage of single mothers increases by 6.3 percentage points (*p*-value = 0.035).

In addition to implementing welfare reforms, several states made changes in their EITCs during the period we examine. By 1993, six states had their own EITCs: Minnesota, Vermont, and Wisconsin had refundable EITCs, while Iowa, Maryland, and Rhode Island had non-refundable EITCs (Baughman 2005). When we restrict the sample to single women in states that did not have state-level EITCs, we again see no major change in the estimated effect of the HITC. The estimates reported in row 9 of Table 8 show a relative increase in the coverage rate of single mothers of 6.5 percentage points

(p-value = 0.017), suggesting that changes in state EITCs were not responsible for the HITC effect we observed earlier.

Row 10 of Table 8 shows the findings when we restrict the sample to women in states with neither welfare reforms nor state-level EITCs. Again, the main findings are essentially unchanged.

Overall, the findings in Table 8 suggest that Table 5's estimated HITC effect is robust to several alternative specifications, sample restrictions, and other tests.

7. Discussion and Conclusion

A refundable tax credit for health insurance has been frequently proposed as a policy to expand health insurance coverage in the United States, but the extent to which an HITC would be effective has been a matter of debate. The Health Insurance Tax Credit of 1991 through 1993 offers a natural experiment that we have used to examine the effectiveness of credits. The main findings in Table 5 suggest that, as a result of the HITC, health insurance coverage among low-education working single mothers was higher than it would otherwise have been by roughly 6 percentage points—that is, without the HITC, only 15.7 percent of these women would have had health insurance covering their children during 1991–1993, whereas the actual (estimated) coverage rate was 22 percent (Tables 4 and 5). This main finding appears to stand up to a variety of falsification and sensitivity tests (see Tables 6 and 8 and the accompanying discussions).

It makes sense to perform two further checks on the main findings. First, we explore whether enough households received an HITC to yield an effect of the size we have estimated. Second, we compare our findings with previous evidence on the price elasticity of health insurance.

Did enough households claim the HITC to account for the estimated effect? In order for the estimated HITC effects reported in rows 1 and 2 of Table 6 to make sense, 36,500 single mothers with less than high school would need to have been newly covered by private health insurance, and 80,500 single mothers with only high school would need to have been newly covered.[14]

In light of the number of HITCs actually claimed, are such increases reasonable? Internal Revenue Service records tabulated by the GAO (1994) show that about 2.3 million tax returns claimed the HITC for tax year 1991. (Congress did not request GAO reports on the HITC in subsequent years.) Because IRS records do not include data on education, we need to impute the number of single mothers with less than high school (and with only high school) who claimed the HITC. We do this by applying population percentages from the CPS to the number of HITCs actually claimed. Tabulations of the March 1992 CPS suggest that, in 1991, about 5.2 million families were HITC-eligible, of whom 11.7 percent (608,000) were headed by single mothers with less than high school, and 36.1 percent (1,877,000) were headed by single mothers with only high school. Applying the population proportion of HITC-eligibles who were working single mothers with less than high school (0.117) to the number of HITCs received (2.3 million) suggests that about 269,100 of these women received a credit in 1991. By similar reasoning, about

[14] The CPS estimate of the number of HITC-eligible single mothers with less than high school in 1991 is 608,000. Roughly 21 percent of these, or 127,000, had private health insurance. The estimated HITC effect of 6 percentage points in Table 6, row 1, would imply 36,500 newly covered families as a result of the HITC. We obtain the estimate for single mothers with only high school in a similar way.

830,300 (0.361 of 2.3 million) working single mothers with only high school would have received a credit in 1991.

Table 9 summarizes these imputations, which suggest that the estimated HITC effects for single mothers with less than high school (Table 6, row 1) are reasonable if 13.6 percent of the HITCs that went to those women resulted in new private health insurance coverage. Similarly, the estimated effects for single mothers with only high school (Table 6, row 2) are reasonable if 9.7 percent of the HITCs claimed by those women resulted in new coverage. This is clearly an exercise in judgment and heuristics; however, we suggest that the estimates in Table 6 are plausible in light of the number of HITCs claimed and the estimated number of HITC-eligible households.

Are the estimates we obtain from the HITC natural experiment consistent with existing estimates of the responsiveness of health insurance to premium changes? The main estimate we report implies a price elasticity of health insurance take-up of –1.27, and a semi-elasticity (the percentage point change in coverage resulting from a one percent increase in price[15]) of –0.24. We calculate these elasticities as follows: First, the HITC we are examining covered 23 percent of the reported average annual health insurance premium (U.S. Government Accountability Office 1991), which implies a 26 percent decrease in the insurance premium (taking a midpoints approach, 23/88.5 = 0.260). Also, the estimated 6.3 percentage-point increase in private health insurance coverage of single mothers during 1991, 1992, and 1993 represents a 33 percent increase in coverage (because from Table 4, 22 percent of single mothers were covered during the HITC years, which implies a counterfactual coverage rate of 15.7 percent). Accordingly,

[15] See Gruber and Poterba (1994) and Congressional Budget Office (2005) for useful discussions of elasticity concepts in this literature.

a 26 percent reduction in the price of health insurance led to a 33 percent increase in health insurance coverage, which implies an elasticity of –1.27. For the semi-elasticity, a 26 percent price reduction led to a 6.3 percentage point increase in coverage, giving a semi-elasticity of –0.24.

As Table 10 shows, the price elasticity and semi-elasticity estimated from the HITC natural experiment are at the elastic end of the range of existing estimates that might be seen as comparable.[16] The take-up elasticity we obtain from the HITC is similar to Gruber and Poterba's (1994) for single, self-employed workers, and at the elastic end of the range found by Marquis et al. (2004).[17] The semi-elasticities we obtain are at the elastic end of the range estimated by Chernew, Frick and McLaughlin (1997) and Blumberg, Nichols, and Banthin (2001), and are similar to those obtained by Marquis et al. (2004) and the Congressional Budget Office (2005). The last two sets of estimates are obtained from samples of employed individuals who are near poverty—workers who should be similar to the low-education working single mothers we have examined.

Table 5's findings suggest that enough eligible low-earning workers with children would take up an HITC to significantly increase the health insurance coverage of such workers and their children. In that sense, they appear favorable to the HITC. Nevertheless, during the existence of the HITC, reports emerged that some insurers offered policies of little real value that happened to cost the same amount as the credit. Indeed, a Congressional investigation found that some insurers sold policies for children

[16] In assembling Table 10, we gave priority to estimates based on lower-income, single workers in order to show estimates for workers who are similar to low-education working single mothers. The empirical literature has consistently estimated larger price elasticities for such workers than for higher-income and married workers [for discussions, see Gruber and Poterba (1994) and Marquis et al. (2004)]. We have included estimates both for those with and those without access to group health insurance, although estimated elasticities for the latter are generally smaller than for the former (Marquis et al. 2004).
[17] The take-up elasticity estimated from the HITC is larger than most others in Table 10 in part because the base level of coverage for the sample we examine is so low.

covering only “cancer, heart attacks, strokes, and other diseases that few children have” (Solomon 2007). Senator Lloyd Bentsen, the HITC’s original sponsor, considered these reported abuses serious enough that he led efforts to repeal the HITC. It is an open question whether such concerns could be mitigated by a well-developed private health insurance market or appropriate regulation.

Also, our rough-and-ready attempt to calculate how many low-education working single mothers received the HITC (above and Table 9) suggests that the HITC went to many who would have purchased health insurance even without the credit. This lack of target efficiency may not be a major issue, given that the credits went to low-income households. Still, it raises questions, discussed in some depth by Gruber (2008) about whether the HITC is the most efficient way to expand health insurance coverage among low-income households.

Finally, although the results presented here are consistent with the HITC increasing health insurance coverage substantially, the analysis is hampered by small sample sizes. It seems possible that further research using different data sources—panel data in particular—could give further evidence by allowing observation of the same individuals before and after adoption of the HITC.

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Table 1
Health insurance coverage rates for low-education working single mothers and low-education working single women without children

	1988	1989	1990	1991	1992	1993
Single mothers						
Private insurance	0.350	0.353	0.297	0.297	0.279	0.308
Private insurance in own name	0.323	0.325	0.276	0.278	0.255	0.270
Private insurance in own name that covers children	0.221	0.249	0.218	0.209	0.197	0.202
Single women without children						
Private insurance	0.466	0.409	0.414	0.367	0.320	0.272
Private insurance in own name	0.398	0.385	0.378	0.341	0.283	0.209

Notes: The data are from the March 1989–1994 Annual Demographic Supplements to the Current Population Survey (CPS). The sample contains working single women with less than a high school education. We define "working" as having positive hours and positive earnings during the year. We exclude women who are in school full-time, those who are separated from their spouse, and those who report being ill or disabled. Means are tabulated using CPS March supplement weights. Sample sizes are 2,228 (single mothers) and 1,433 (single women without children).

Table 2
Summary statistics: low-education working single mothers and low-education working single women without children

Variable	Single mothers	Single women without children
Age (years)	30.5	29.7
White (%)	63.8	81.0
Black (%)	33.6	14.7
Other race (%)	2.6	4.4
Has children under age 6 (%)	48.6	0.0
Has children aged 6–18 (%)	69.9	0.0
Has children aged 19–24 and a full-time student (%)	2.0	0.0
Full-time, full-year (%)	37.8	49.1
Part-time, full-year (%)	8.9	9.2
Full-time, part-year (%)	31.7	29.0
Part-time, part-year (%)	21.6	12.8
Earned income ($)	7,257	8,686
Unearned income ($)	1,833	464
Number of observations	2,228	1,433

Notes: See Table 1. Dollar amounts are converted to 1993 dollars using the Consumer Price Index, All Urban Consumers (CPI-U).

Table 3
Summary statistics for low-education working single mothers and low-education working single women without children, before and during HITC

	Single mothers		Single women without children	
Variable	Before HITC (1988–90)	During HITC (1991–93)	Before HITC (1988–90)	During HITC (1991–93)
Age (years)	30.3	30.7	29.6	29.8
White (%)	64.6	63.0	81.8	80.1
Black (%)	32.9	34.4	15.5	13.7
Other race (%)	2.5	2.7	2.6	6.2
Has children under age 6 (%)	48.6	48.7	0.0	0.0
Has children aged 6–18 (%)	70.5	69.3	0.0	0.0
Has children aged 19–24 and a full-time student (%)	1.5	2.4	0.0	0.0
Full-time, full-year (%)	39.0	36.5	50.4	47.6
Part-time, full-year (%)	7.2	10.7	6.8	11.6
Full-time, part-year (%)	31.6	31.7	31.7	26.0
Part-time, part-year (%)	22.2	21.1	11.1	14.7
Earned income ($)	6,770	7,764	8,003	9,419
Unearned income ($)	1,599	2,077	472	455
Number of observations	1,153	1,075	741	692

Notes: See Table 1. Dollar amounts are converted to 1993 dollars using the Consumer Price Index, All Urban Consumers (CPI-U).

Table 4
Private health insurance coverage rates for low-education working single mothers and low-education working single women without children

	Before HITC (1988–1990)	During HITC (1991–1993)	Difference
Single mothers	0.244 (0.013) [1,153]	0.220 (0.013) [1,075]	–0.024 (0.018)
Single women without children	0.389 (0.018) [741]	0.299 (0.017) [692]	–0.090 (0.025)
Difference	–0.145 (0.022)	–0.080 (0.022)	—
Difference-in-differences	—	—	0.065 (0.031)

Notes: See Table 1. Figures are average rates of private health insurance coverage in own name and covering a child. Robust standard errors in parentheses. Sample sizes in brackets.

Table 5

Difference-in-differences estimates of the effect of the HITC on private health insurance coverage of low-education working single mothers from probit estimates of equation (1)

Dependent variable: *covered* by private health insurance	Specification		
	(1)	(2)	(3)
treatgroup	–0.145	–0.080	–0.084
	(0.020)	(0.029)	(0.026)
HITC	–0.090	–0.087	–0.104
	(0.023)	(0.022)	(0.021)
treatgroup×*HITC*	0.065	0.060	0.063
	(0.030)	(0.028)	(0.028)
Number of observations	3,661	3,661	3,661
Pseudo R-squared	0.020	0.061	0.256

Notes: See Table 1. Figures are estimated changes (marginal effects) in the probability of private health insurance coverage in own name and covering a child from probit models. Bootstrap standard errors are in parentheses. Specification 1 includes no control variables. Specification 2 includes age, indicators of race (white, black, and other), number of children under age 6, number of children aged 6–18, and number of children aged 19–24 and a full-time student. Specification 3 adds indicators of work status (full-time/full-year, full-time/part-year, part-time/full-year, and part-time/part-year), earned income, and unearned income to the controls in specification 2. Estimated marginal effects for all regressors are reported in Appendix Table A3.

Table 6
Difference-in-differences estimates from alternative treatment and comparison groups

Dependent variable: *covered* by private health insurance	*treatgroup*	*HITC*	*treatgroup*×*HITC*
1. Treatment group:	–0.084	–0.104	0.063
Single mothers with less than high school [*N*=2,228]	(0.026)	(0.021)	(0.028)
Control group:			
Single women with less than high school [*N* =1,433]			
2. Treatment group:	–0.166	–0.107	0.043
Single mothers with high school [*N*=6,794]	(0.016)	(0.011)	(0.016)
Control group:			
Single women with high school [*N* =6,608]			
3. Treatment group:	–0.173	–0.098	0.025
Single mothers with some college [*N* =3,630]	(0.020)	(0.012)	(0.021)
Control group:			
Single women with some college [*N* =5,060]			
4. Treatment group:	–0.184	–0.048	–0.006
Single mothers with college [*N* =1,179]	(0.034)	(0.009)	(0.024)
Control group:			
Single women with college [*N* =5,736]			

Notes: Estimates in row 1 come from specification 3 of Table 5. Estimates in rows 2, 3, and 4 come from applying the same model to different samples of women, as indicated.

Table 7
Difference-in-differences estimates of the change in private health insurance coverage relative to 1990, low-education working single mothers

Dependent variable: *covered* by private health insurance	
1988×*treatgroup*	0.025
	(0.042)
1989×*treatgroup*	0.036
	(0.044)
1990×*treatgroup*	—
1991×*treatgroup*	0.064
	(0.046)
1992×*treatgroup*	0.072
	(0.048)
1993×*treatgroup*	0.110
	(0.046)
Number of observations	3,661
Pseudo R-squared	0.264

Notes: See Table 1. Figures are estimated changes in the probability of private health insurance coverage from a probit model. Bootstrap standard errors are in parentheses. The specification includes age, race, number of children, work status, earned income, unearned income, and a set of year indicators. The marginal effects of the complete set of regressors are reported in Appendix Table A4.

Table 8
Sensitivity tests for HITC effects on low-education working single mothers

Dependent variable: *covered* by private health insurance (except rows 2, 3, and 4)		*treatgroup*	*HITC*	*treatgroup*×*HITC*	Number of observations	Pseudo *R*-squared
1.	Basic specification (Table 5, specification 3)	–0.084 (0.026)	–0.104 (0.021)	0.063 (0.028)	3,661	0.256
2.	Basic specification, but dependent variable is whether employed	–0.263 (0.021)	–0.063 (0.020)	0.021 (0.404)	6,797	0.034
3.	Basic specification, but dependent variable is whether covered by Medicaid	0.129 (0.022)	0.059 (0.020)	0.014 (0.025)	3,661	0.132
4.	Basic specification, but dependent variable is covered by either private insurance or Medicaid	–0.018 (0.027)	–0.074 (0.027)	0.084 (0.030)	3,661	0.067
5.	Add state dummies to the basic specification	–0.090 (0.021)	–0.106 (0.020)	0.068 (0.028)	3,661	0.291
6.	Add state unemployment rate to the basic specification	–0.085 (0.021)	–0.079 (0.021)	0.063 (0.029)	3,661	0.259
7.	Add state dummies and state unemployment rate to the basic specification	–0.089 (0.021)	–0.136 (0.023)	0.067 (0.028)	3,661	0.292
8.	Restrict sample to single women in states without welfare waivers in 1993	–0.089 (0.030)	–0.120 (0.022)	0.063 (0.033)	2,845	0.268
9.	Restrict sample to single women in states without state EITCs	–0.093 (0.027)	–0.106 (0.020)	0.065 (0.023)	3,484	0.258
10.	Restrict sample to single women in states without welfare waivers in 1993 and without state EITCs	–0.102 (0.028)	–0.125 (0.024)	0.067 (0.031)	2,668	0.270

Notes: See Table 5. Estimates in row 1 (basic specification) come from specification 3 of Table 5. The basic specification (row 1) includes age, race, number of children, work status, earned income, and unearned income as controls. Bootstrap standard errors in parentheses.

Table 9
Implications of the estimates: Proportions of HITCs resulting in new health insurance coverage

	Working single mothers with less than high school	Working single mothers with only high school
1. Covered by private health insurance (from CPS)	127,000	953,200
2. Number newly covered implied by estimated effects in Table 6	36,500	80,500
3. HITC-eligible in 1991 (from CPS)	608,000	1,876,500
4. Received HITC in 1991 (imputed from sample proportions)	172,500 (= 2,300,000 × 0.117)	453,100 (= 2,300,000 × 0.361)
5. Row 2/Row 4 (proportion of HITCs resulting in new coverage	0.136	0.097

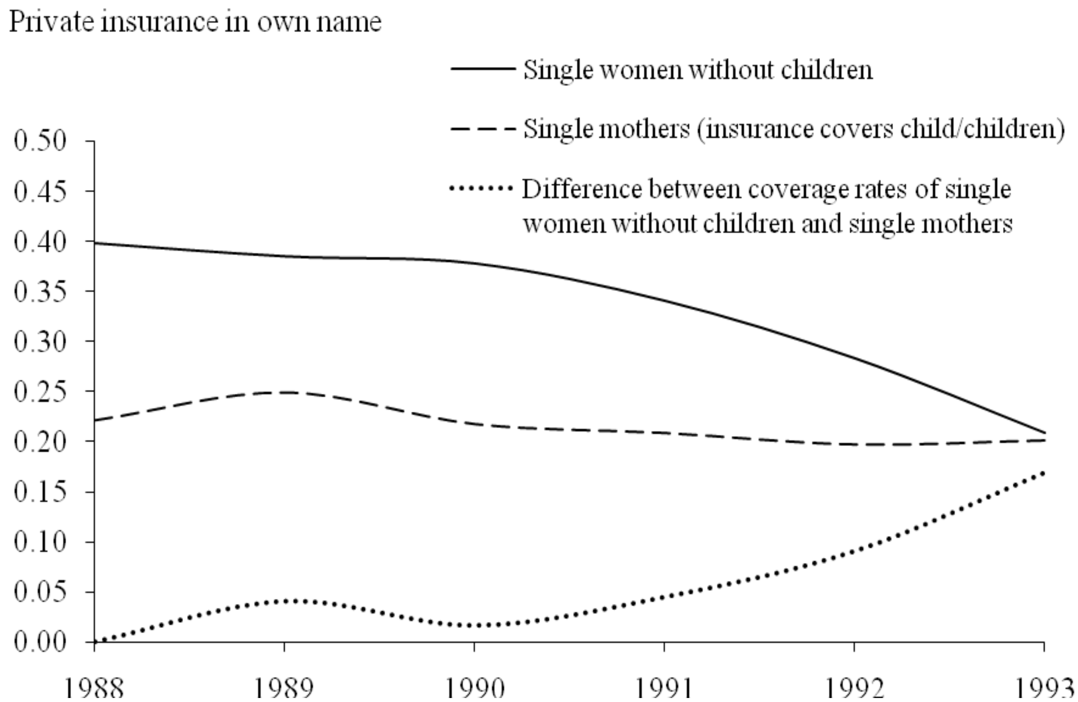
Note: Rows 1 and 3 report authors' tabulations of the March 1992 Current Population Survey. Row 2 reports the number of working single mothers newly covered by private health insurance implied by the estimated HITC effects in Table 6, rows 1 and 2. Row 4 reports imputations based on IRS records of the number of HITCs claimed (from GAO 1994) and CPS sample proportions of HITC-eligible working single mothers with less than high school and only high school. See text for discussion.

Table 10
Estimated price elasticities of health insurance (HI) take-up, selected studies

Study	Source of price variation	Sample	Elasticity of take-up	Semi-elasticity of take-up
Gruber and Poterba (1994)	1986 Tax Reform Act	Single, self-employed workers	–1.23	–1.78
Marquis and Long (1995)	HI premium imputed to family	Working families without access to group HI		
		below 200% poverty	–0.64 to –0.31	–0.09 to –0.07
		above 200% poverty	–0.45 to –0.27	–0.10 to –0.08
Chernew, Frick, and McLaughlin (1997)	Observed HI premium or employee contribution	Single workers below 200% poverty	–0.27 to –0.03	–0.19 to –0.03
Blumberg, Nichols, and Banthin (2001)	Observed or imputed HI premium	Workers with children (or married) aged 18 to 64 and below 200% poverty	–0.30 to –0.06	–0.28 to –0.05
Long and Marquis (2002)	Public subsidy schedule	Low-income adults in Washington State	–0.7 to –0.3	–0.06 to –0.04
Marquis, Buntin, Escarce, Kapur, and Yegian (2004)	HI premium imputed to individual	California families without access to group HI		
		below 200% poverty, age < 35, self-employed	–1.21 to –0.71	–0.14 to –0.10
		below 200% poverty, age < 35, not self-employed	–1.03 to –0.52	–0.24 to –0.21
Congressional Budget Office (2005)	Premium differences from state-level regulations	Individual workers without access to group health plan		
		full sample	–0.57	–0.09
		below 200% poverty	–0.84	–0.19
Cebi and Woodbury (2010)	Introduction of Health Insurance Tax Credit	Working single mothers with less than high school	–1.27	–0.24

Sources: Reported estimates and (in some cases) authors' calculations from the papers cited.

Figure 1
Health insurance coverage rates for low-education working single mothers and low-education working single women without children



Notes: See Table 1.

Appendix

Table A1
Earned Income Tax Credit parameters

Basic Tax Credit parameters, 1987–1996, nominal dollars

Tax year	Phase-in rate (%)	Phase-in range ($)	Maximum credit ($)	Phase-out rate (%)	Phase-out range ($)
1987	14.00	0–6,080	851	10.00	6,920–15,432
1988	14.00	0–6,240	874	10.00	9,840–18,576
1989	14.00	0–6,500	910	10.00	10,240–19,340
1990	14.00	0–6,810	953	10.00	10,730–20,264
1991					
One child	16.70	0–7,140	1,192	11.93	11,250–21,250
Two children	17.30	0–7,140	1,235	12.36	11,250–21,250
1992					
One child	17.60	0–7,520	1,324	12.57	11,840–22,370
Two children	18.40	0–7,520	1,384	13.14	11,840–22,370
1993					
One child	18.50	0–7,750	1,434	13.21	12,200–23,050
Two children	19.50	0–7,750	1,511	13.93	12,200–23,050
1994					
No child	7.65	0–4,000	306	7.65	5,000–9,000
One child	26.30	0–7,750	2,038	15.98	11,000–23,755
Two children	30.00	0–8,245	2,528	17.98	11,000–25,296
1995					
No child	7.65	0–4,100	314	7.65	5,130–9,500
One child	34.00	0–6,160	2,094	15.98	11,290–24,396
Two children	36.00	0–8,640	3,110	20.22	11,290–26,673
1996					
No child	7.65	0–4,220	323	7.65	5,280–9,500
One child	34.00	0–6,330	2,152	15.98	11,610–25,078
Two children	40.00	0–8,890	3,556	21.06	11,610–28,495

Health Insurance Tax Credit parameters, 1991–1993, nominal dollars

Tax year	Phase-in rate (%)	Phase-in range ($)	Maximum credit ($)	Phase-out rate (%)	Phase-out range ($)
1991	6.00	0–7,140	428	4.28	11,250–21,250
1992	6.00	0–7,520	451	4.28	11,840–22,370
1993	6.00	0–7,750	465	4.28	12,200–23,050

Source: U.S. Government Accountability Office (1991, 1994); U.S. House of Representatives (1998).

Table A2
Major legislative changes affecting low-income women, 1987–1994

Tax year	*Earned Income Tax Credit*
1988	The beginning and end of the phase-out range increased by about $3,000.
1991	The credit rates rose by 2 percentage points. Additional credits established for families with two or more children. The new increment to the maximum credit for a second child was $43. Supplemental credits added for child health insurance premiums and children less than one year of age.
1992	The credit rates rose by one percentage point. The increment to the maximum credit for a second child rose to $60.
1993	The credit rates rose by one percentage point. The increment to the maximum credit for a second child rose to $77.
1994	The credit rates rose by about 10 percentage points. The increment to the maximum credit for a second child rose to $490. Supplemental credits for health insurance and children less than one year of age repealed. Small credits established for taxpayers without children and between the ages of 25 and 65. The IRS began to notify eligible taxpayers of the advance payment option, which had been available since 1979.

Tax year	*Medicaid*
April 1987	States permitted to extend Medicaid coverage to children under age two in families below 100 percent of the poverty line.
July 1988	States permitted to extend Medicaid coverage to children under age five in families below 100 percent of the poverty line.
October 1988	States permitted to extend Medicaid coverage to children under age eight in families below 100 percent of the poverty line, and to children under age one in families below 185 percent of the poverty line.
July 1989	States required to extend Medicaid coverage to children under age one in families below 75 percent of the poverty line.
April 1990	States required to extend Medicaid coverage to children under age six in families below 133 percent of the poverty line.
July 1991	States required to extend Medicaid coverage to children under age 19 in families below 100 percent of the poverty line.

Source: U.S. General Accountability Office (1993); U.S. House of Representatives (1998); Meyer and Rosenbaum (2000).

Table A3
Estimated marginal effects from probit models underlying Table 5 estimates

Dependent variable: *covered* by private health insurance	(1)	(2)	(3)
treatgroup	–0.145	–0.080	–0.084
	(0.020)	(0.029)	(0.026)
HITC	–0.090	–0.087	–0.104
	(0.023)	(0.022)	(0.021)
treatgroup×*HITC*	0.065	0.060	0.063
	(0.030)	(0.028)	(0.028)
Age	—	0.011	0.005
		(0.001)	(0.001)
White	—	–0.012	–0.009
		(0.035)	(0.031)
Black	—	–0.057	–0.033
		(0.036)	(0.033)
Number of children under age 6	—	–0.046	–0.004
		(0.015)	(0.014)
Number of children aged 6-18	—	–0.030	–0.019
		(0.010)	(0.009)
Number of children aged 19-24 and a full-time student	—	–0.034	–0.029
		(0.053)	(0.047)
Full-time, full-year	—	—	0.259
			(0.034)
Part-time, full-year	—	—	0.077
			(0.035)
Full-time, part-year	—	—	0.138
			(0.027)
Part-time, part-year	—	—	—
Earned income	—	—	0.018
			(0.002)
Unearned income	—	—	0.005
			(0.003)
Number of observations	3,661	3,661	3,661
Pseudo R-squared	0.020	0.061	0.256

Notes: Figures are estimated changes in the probability of private health insurance coverage from probit models. Bootstrap standard errors are in parentheses.

Table A4
Estimated marginal effects from probit model underlying Table 7 estimates

Dependent variable: *covered* by private health insurance	
treatgroup	–0.104 (0.032)
1988	0.036 (0.036)
1989	0.012 (0.039)
1990	—
1991	–0.037 (0.034)
1992	–0.094 (0.034)
1993	–0.144 (0.034)
1988×*treatgroup*	0.025 (0.042)
1989×*treatgroup*	0.036 (0.044)
1990×*treatgroup*	—
1991×*treatgroup*	0.064 (0.046)
1992×*treatgroup*	0.072 (0.048)
1993×*treatgroup*	0.110 (0.046)
Age	0.005 (0.001)
White	–0.008 (0.030)
Black	–0.031 (0.034)
Number of children under age 6	–0.004 (0.013)
Number of children aged 6–18	–0.019 (0.009)
Number of children aged 19–24 and a full-time student	–0.027 (0.055)

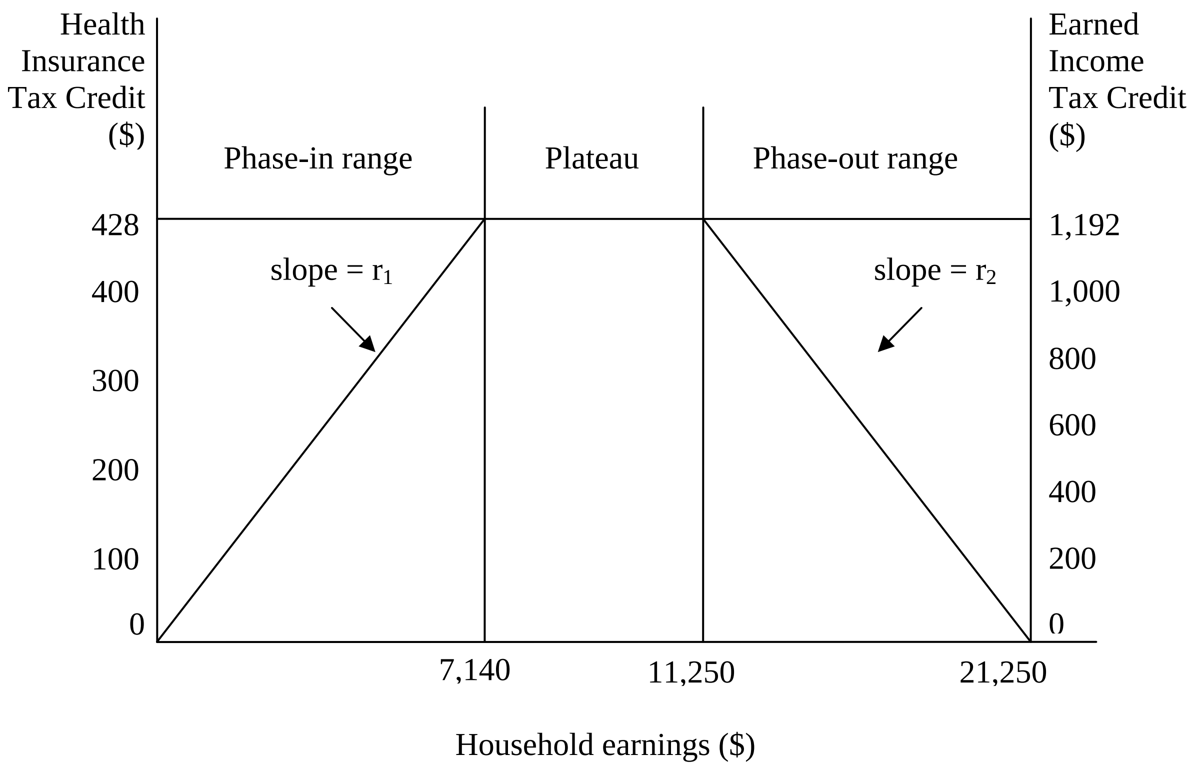
(continued)

Table A4 (continued)

Dependent variable: *covered* by private health insurance	
Full-time, full-year	0.257
	(0.034)
Part-time, full-year	0.077
	(0.039)
Full-time, part-year	0.138
	(0.028)
Part-time, part-year	—
Earned income	0.018
	(0.002)
Unearned income	0.006
	(0.003)
Number of observations	3,661
Pseudo R-squared	0.264

Notes: Figures are estimated changes in the probability of private health insurance coverage from a probit model. Bootstrap standard errors are in parentheses.

Figure A1
Earned Income Tax Credit and Health Insurance Tax Credit schedules, 1991



Difference-in-differences estimates in linear probability and probit models

In a linear probability model, calculating a difference-in-differences (DD) estimate is straightforward. In the simplest case, we have two time periods (pre and post) and two groups (control and treatment). Letting y be the binary outcome of interest, the linear probability model can be written:

$$\Pr(y = 1|\bullet) = \beta_0 + \beta_1\, treatment + \beta_2\, post + \beta_3\, treatment \times post + X\beta \qquad \text{(A.1)}$$

where X denotes a vector of explanatory variables. In this setting, the DD estimator is the OLS estimator of β_3 (the coefficient on the interaction between *treatment* and *post*).

If (A.1) is estimated as a probit model, the DD estimator is no longer the coefficient on the interaction term. To illustrate, consider the following:

$$\Pr(y = 1|\bullet) = \Phi(\beta_0 + \beta_1\, treatment + \beta_2\, post + \beta_3\, treatment \times post + X\beta) \qquad \text{(A.2)}$$

where Φ is the standard normal cumulative distribution function. The marginal effect of the interaction term (the analog to β_3 in the linear probability model) is:

$$\frac{\Delta \Pr(y = 1 | \bullet)}{\Delta treatment \times post} = \begin{bmatrix} \Phi(\beta_0 + \beta_1 treatment + \beta_2 post + \beta_3 + X\beta) - \\ \Phi(\beta_0 + \beta_1 treatment + \beta_2 post + X\beta) \end{bmatrix} \qquad \text{(A.3)}$$

But the DD estimator is the marginal effect of changes in both *treatment* and *post*:

$$\frac{\Delta \Pr(y = 1 | \bullet)}{\Delta treatment \quad \Delta post} = \frac{\Delta[\Phi(\beta_0 + \beta_1 + \beta_2 post + \beta_3 post + X\beta) - \Phi(\beta_0 + \beta_2 post + X\beta)]}{\Delta post}$$

$$= \begin{Bmatrix} [\Phi(\beta_0 + \beta_1 + \beta_2 + \beta_3 + X\beta) - \Phi(\beta_0 + \beta_1 + X\beta)] \\ - [\Phi(\beta_0 + \beta_2 + X\beta) - \Phi(\beta_0 + X\beta)] \end{Bmatrix} \qquad \text{(A.4)}$$

As discussed by Ai and Norton (2003), Norton, Wang, and Ai (2004), and DeLeire (2004), many authors incorrectly interpret equation (A.3) as a DD estimator, but equation (A.4) makes clear that, in the probit model, the marginal effect of a change in the interaction term is not the difference in differences.

Following DeLeire (2004), we obtain DD estimates from probit models by taking the discrete double difference of the standard normal cumulative distribution function. In particular, we estimate the probit analog to equation (1):

$\Pr(covered_i = 1|\bullet) = \Phi[\beta_0 + \beta_1 treatgroup_i + \beta_2 HITC_t + \beta_3 treatgroup_i \times HITC_t + X_i\,\beta]$

then predict four counterfactual probabilities for each observation in the sample:

1. The predicted probability of private health insurance coverage for the treatment group in the post-HITC period: $\Phi\left(\hat{\beta}_0 + \hat{\beta}_1 + \hat{\beta}_2 + \hat{\beta}_3 + X_i\hat{\beta}\right)$
2. The predicted probability of private health insurance coverage for the treatment group in the pre-HITC period: $\Phi\left(\hat{\beta}_0 + \hat{\beta}_1 + X_i\hat{\beta}\right)$
3. The predicted probability of private health insurance coverage for the control group in the post-HITC period: $\Phi\left(\hat{\beta}_0 + \hat{\beta}_2 + X_i\hat{\beta}\right)$
4. The predicted probability of private health insurance coverage for the control group in the pre-HITC period: $\Phi\left(\hat{\beta}_0 + X_i\hat{\beta}\right)$

Using these predicted probabilities, we calculate the DD for each observation as

$$\left\{\begin{matrix}\left[\Phi\left(\hat{\beta}_0 + \hat{\beta}_1 + \hat{\beta}_2 + \hat{\beta}_3 + X_i\hat{\beta}\right) - \Phi\left(\hat{\beta}_0 + \hat{\beta}_1 + X_i\hat{\beta}\right)\right] - \\ \left[\Phi\left(\hat{\beta}_0 + \hat{\beta}_2 + X_i\hat{\beta}\right) - \Phi\left(\hat{\beta}_0 + X_i\hat{\beta}\right)\right]\end{matrix}\right\}$$

The probit DD estimator is the average of this double difference over all observations. Obtaining standard errors by the bootstrap method is straightforward.