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Health Insurance Tax Credits and Health Insurance Coverage of Low-Earning Single Mothers

Upjohn Institute Staff Working Paper 09-158

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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 This health insurance tax credit (HITC) existed during tax years 1991, 1992, and 1993, and was then rescinded.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: H2; H51, I18, J32

Keywords: Health insurance; Low-wage workers; Tax credits and subsidies

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1. Introduction

A longstanding suggestion 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; Furman 2008). Such a Health Insurance Tax Credit (HITC) would grant a tax credit up to a specified maximum—for example, \$1,000 for an individual or \$2,000 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 would reduce the price of health insurance to eligible taxpayers and extend the tax-favored treatment of health insurance to workers who lack access to employer-provided health insurance. Accordingly, 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) and Gruber and Levitt (2000) suggested that the HITC might reduce the number of uninsured by only about 10 percent. Emmons, Madly, and Woodbury (2005) replicated Gruber’s simulation model and found (as is often true) that relatively minor changes in assumptions could result in substantial changes in simulated impacts of the HITC. Their conclusions echoed those of Pauly, Song, and Herring (2001): Simulations of the impact of health insurance tax credits are

uncertain because little empirical evidence exists to guide modelers in choosing appropriate behavioral assumptions.¹

Only a few empirical studies have estimated the responsiveness of workers to health insurance subsidies and changes in premiums. Gruber and Poterba (1994) examined the tax subsidy for self-employed workers' health insurance introduced under the 1986 Tax Reform Act. Their findings suggest that the self-employed responded strongly to the new subsidy and imply an elasticity of demand for health insurance of about 1.8. Other studies estimate much smaller elasticities. Chernew, Frick, and McLaughlin (1997) and Blumberg, Nichols, and Banthin (2001) examined matched employer-employee data, and both estimated the elasticity of insurance take-up with respect to premiums to be less than -0.1 . Gruber and Washington (2005) examined a change in the tax treatment of federal employees' health insurance premiums in which the employee's share, previously paid with after-tax earnings, became payable with pre-tax earnings. Their main finding is that federal employees' take-up of health insurance was minimally responsive to the change in subsidy. We return to these studies in the conclusion.

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

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.

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, the approach to estimation, and the data we use. We then present the main findings, followed by several sensitivity tests and a discussion of possible alternative explanations of the findings.

2. 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 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. (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.) 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.³ 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 We have not been able to replicate this take-up rate, which 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. Our tabulations suggest that 5.2 families were eligible for the HITC in 1991. See the conclusion for further discussion.

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 between 1991 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 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 we will see, the two groups did have 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.

The model we estimate can be written:

$$\Pr(\text{covered}_i=1|\bullet) = F[\beta_0 + \beta_1 \text{treatgroup}_i + \beta_2 \text{HITC}_t + \beta_3 \text{treatgroup}_i \times \text{HITC}_t + X_i \beta] \quad (1)$$

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 $\text{treatgroup}_i \times \text{HITC}_t$ 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),⁴ earned income,⁵ and unearned income.⁶ 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 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 44⁷ 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

⁴ 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.

⁵ Earned income includes income from wages, salaries, and self-employment.

⁶ 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.

⁷ 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).

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.⁸

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 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 choosing the period between 1988 and 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

⁸ 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.

insurance coverage of dependents.⁹ 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. 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. This change led to an increase in the number of persons reporting employer-provided coverage.

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?

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)

⁹ For more detailed information, see Appendix T of Unicon (2005), which describes changes in the health insurance questions.

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. Clearly, the private health insurance coverage rates of single mothers fell between 1988 and 1993. For example, 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. The coverage rate for single women without children also fell between 1988 and 1990, but fell sharply even after 1990 (from 37.8 to 20.9 percent). A likely explanation for the drop after 1990 is the recession of 1991, which would have reduced both employment and access to employer-provided health insurance of single women. Figure 1 graphs the coverage rates for single mothers (private insurance in own name that covers children) and single women without children (private insurance in own name).

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

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.

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 points between 1988–1990 and 1991–1993. The second row shows that, over the same time 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).¹⁰ 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.

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

¹⁰ 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.

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.¹¹

5.2. Findings for single women with more education

We would expect the estimated effect of the HITC on low-education single mothers to be greater than its effect on single mothers with more education for a simple reason: Low-education single mothers are more likely to be low earners and hence eligible for the EITC and HITC. For single mothers with more education, we would expect a smaller estimated HITC effect because fewer are eligible.

We view this as a useful falsification test because it may suggest whether the HITC estimates in Table 5 are convincing. We know of no *a priori* reason why single mothers with high school or more (and hence higher average earnings) should have experienced improved health insurance coverage from 1988–1990 to 1991–1993. Accordingly, finding that single mothers with high school or more 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 higher levels of education were less likely to experience an HITC-induced improvement in health insurance coverage than single mothers with low education. We do this by estimating equation (1) using samples of women with successively more educational attainment. The

¹¹ 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.

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 not statistically significant (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, which is essentially zero. That we observe negligible responses to the HITC among single mothers with more than high school—women who are unlikely to be eligible for the HITC—tends to increase our confidence that we are identifying an HITC effect for low-education single women in Table 5.

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 our main group of interest—low-education working single women, and report estimates from a specification that is analogous to that in 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 several sensitivity tests that attempt to control for influences on private health insurance in the early 1990s other than the HITC. These include (1) expansion of the Medicaid program, (2) state-level economic conditions and state fixed effects, and (3) welfare reform and the introduction of state-level EITCs. We also explore possible effects of the HITC on labor supply.

6.1. 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 the

AFDC Program (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 of the HITC’s impact on private health insurance coverage in Tables 5, 6, and 7 may be downward-biased. 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 2 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 standard error is large.)

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 3 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 (which repeats the main findings 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. Specifically, rows 1, 2, and 3 suggest that 6.3 percentage points 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.

6.2. 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 4, 5, and 6 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 4, 5, and 6 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).

6.3. 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 impact 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 7 of Table 8 shows that restricting the sample in this way leaves the estimated HITC impact on single mothers essentially unchanged—an increase in the private health insurance coverage of single mothers of 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 tax EITCs, while Iowa, Maryland,

and Rhode Island had non-refundable EITCs (Baughman 2005). When we restrict our 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 8 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 9 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.

6.4. 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 (not reported in a table) 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 not enough to overturn the main finding from Table 5.

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

7. Discussion and Conclusion

The ongoing debate over how best to increase health insurance coverage in the United States has led to a range of proposals for reform. One incremental proposal is the HITC, but the extent to which an HITC could be expected to increase the health insurance coverage of low-earning workers has been a matter of debate. Although Gruber and Poterba's (1994) research suggests workers are highly responsive to tax subsidies for health insurance, other research (Chernew, Frick, and McLaughlin 1997; Blumberg, Nichols, and Banthin 2001; Gruber and Washington 2005) seems to offer little evidence that an HITC would be effective.

The Health Insurance Tax Credit of 1991 through 1993 offers a natural experiment that we have used in this paper to examine the effectiveness of credits. The main findings reported 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 14 percent of these women would have had health insurance covering their children during

1991–1993, whereas the actual (measured) coverage rate was about 20 percent. This main finding appears to hold up to a variety of falsification and sensitivity tests (see Tables 6 and 8 and the accompanying discussions). The estimate implies an elasticity of health insurance take-up with respect to tax credits close to one.¹²

In view of evidence from several studies that the elasticity of health insurance coverage with respect to price of health insurance is close to zero, the findings in Table 5 may seem surprising. The obvious question is whether the findings we report are plausible. We address this question in two ways. First, we explore whether enough households received an HITC to lead to an effect of the size we have estimated. Second, we consider how applicable previous evidence on the price elasticity of health insurance is to the group we examine—working single mothers.

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.¹³

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

¹² The HITC we are examining covered 23 percent of the reported average annual health insurance premium in 1991 (U.S. Government Accountability Office 1991). The estimated 6.3 percentage-point increase in health insurance coverage of single mothers during 1991, 1992, and 1993 implies a 25.8 percent increase in private coverage (because 24.4 percent of single mothers had private health insurance coverage during the before-HITC years). Accordingly, a 23 percent reduction in the price of health insurance led to a 25.8 percent increase in health insurance coverage, which implies an elasticity of 1.1.

¹³ 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 1, 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.

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.¹⁴ 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 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.

¹⁴ As we noted earlier, GAO (1994) reported an estimate of 8.8 million HITC-eligible families, which is substantially larger than our estimate of 5.2 million. Our tabulations of the CPS suggest that 8.8 million is an estimate of the number of *individuals* in families that were eligible for the HITC (not the number of eligible families). The HITC take-up rate based on the GAO estimate of 8.8 million families is about 26 percent. The HITC take-up rate based on our estimate of 5.2 million eligible families is about 44 percent, which seems reasonable in light of Scholz's (1994) estimate that the 1990 EITC take-up rate was 80 to 86 percent.

How applicable are previous research findings to the HITC? The studies cited in the introduction differ from this one in two main ways. First, they estimate the effect of subsidies or changes in premiums on health insurance take-up, rather than estimating the effect of a tax credit that reimburses an individual dollar-for-dollar for health insurance premium payments. Second, the worker populations studied are quite heterogeneous; that is, the studies focus on workers whose earnings range widely, rather than on low-earning workers. For example, Chernew, Frick, and McLaughlin (1997) and Blumberg, Nichols, and Banthin (2001) use variation in premiums to estimate workers' demand for health insurance. The former study examines a sample of 295 workers in small firms in seven cities, and the latter examines a nationally representative sample of workers who are older, have more education, and are more male than the workers we study. Gruber and Washington (2005) examine a tax subsidy to a group of relatively high-wage government workers, most of whom were already covered by health insurance. Accordingly, it is unclear whether we should draw strong inferences about the effects of an HITC on health insurance coverage from these studies.

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. In that sense, they appear more favorable to the HITC than prior research. 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 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 could be addressed by appropriate regulation of health insurance.

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 the question whether the HITC is the most efficient way of expanding health insurance coverage to 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 us to observe the same individuals before and after the 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

Variable	Single mothers		Single women without children	
	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: <i>covered</i> by private health insurance	Specification		
	(1)	(2)	(3)
<i>treatgroup</i>	−0.145 (0.020)	−0.080 (0.029)	−0.084 (0.026)
<i>HITC</i>	−0.090 (0.023)	−0.087 (0.022)	−0.104 (0.021)
<i>treatgroup</i> × <i>HITC</i>	0.065 (0.030)	0.060 (0.028)	0.063 (0.028)
Number of observations	3,661	3,661	3,661
Pseudo <i>R</i> -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: <i>covered</i> by private health insurance		<i>treatgroup</i>	<i>HITC</i>	<i>treatgroup</i> × <i>HITC</i>
1.	Treatment group: Single mothers with less than high school [<i>N</i> =2,228]	−0.084 (0.026)	−0.104 (0.021)	0.063 (0.028)
	Control group: Single women with less than high school [<i>N</i> =1,433]			
2.	Treatment group: Single mothers with high school [<i>N</i> =6,794]	−0.166 (0.016)	−0.107 (0.011)	0.043 (0.016)
	Control group: Single women with high school [<i>N</i> =6,608]			
3.	Treatment group: Single mothers with some college [<i>N</i> =3,630]	−0.173 (0.020)	−0.098 (0.012)	0.025 (0.021)
	Control group: Single women with some college [<i>N</i> =5,060]			
4.	Treatment group: Single mothers with college [<i>N</i> =1,179]	−0.184 (0.034)	−0.048 (0.009)	−0.006 (0.024)
	Control group: Single women with college [<i>N</i> =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: <i>covered</i> by private health insurance	
<i>treatgroup</i>	−0.104 (0.032)
1988× <i>treatgroup</i>	0.025 (0.042)
1989× <i>treatgroup</i>	0.036 (0.044)
1990× <i>treatgroup</i>	—
1991× <i>treatgroup</i>	0.064 (0.046)
1992× <i>treatgroup</i>	0.064 (0.046)
1993× <i>treatgroup</i>	0.110 (0.046)
Number of observations	3,661
Pseudo <i>R</i> -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: <i>covered</i> by private health insurance (except rows 2 and 3)		<i>treatgroup</i>	<i>HITC</i>	<i>treatgroup</i> × <i>HITC</i>	Number of observations	Pseudo <i>R</i> -squared
1.	Basic specification	−0.084 (0.026)	−0.104 (0.021)	0.063 (0.028)	3,661	0.256
2.	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
3.	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
4.	Add state dummies to the basic specification	−0.090 (0.021)	−0.106 (0.020)	0.068 (0.028)	3,661	0.291
5.	Add state unemployment rate to the basic specification	−0.085 (0.021)	−0.079 (0.021)	0.063 (0.029)	3,661	0.259
6.	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
7.	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
8.	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
9.	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 are in parentheses.

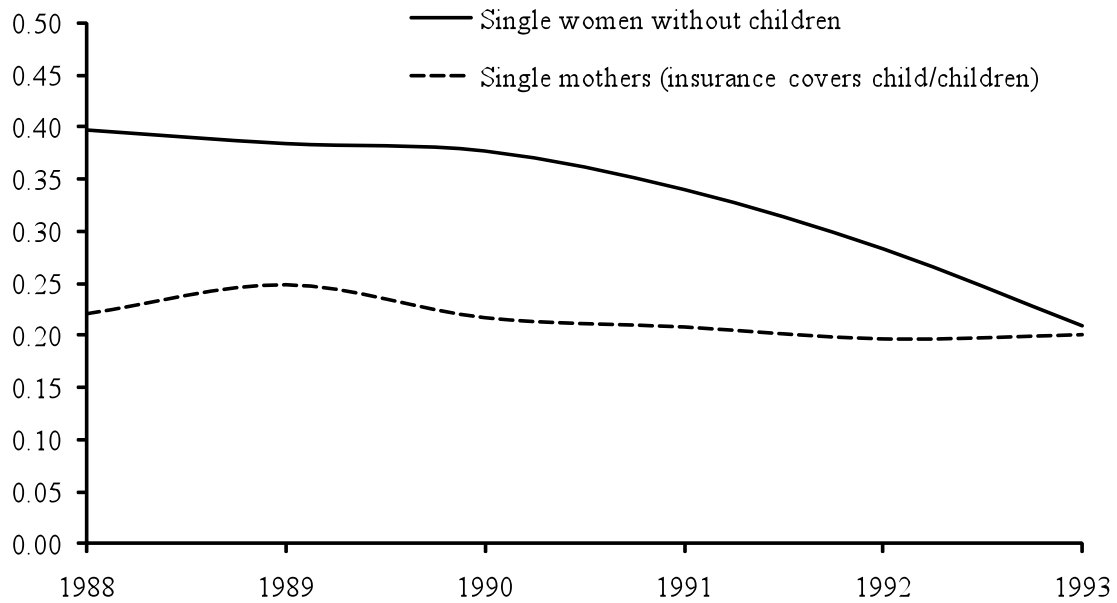
Table 9
Are the estimates reasonable?

	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	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.

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

Private insurance in own name



Notes: See Table 1.

Appendix

Table A1
Earned Income Tax Credit parameters

<i>Basic Tax Credit parameters, 1987–1996, nominal dollars</i>					
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
<i>Health Insurance Tax Credit parameters, 1991–1993, nominal dollars</i>					
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

<i>Tax year</i>	<i>Earned Income Tax Credit</i>
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.
<i>Tax year</i>	<i>Medicaid</i>
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: <i>covered by private health insurance</i>	(1)	(2)	(3)
<i>treatgroup</i>	−0.145 (0.020)	−0.080 (0.029)	−0.084 (0.026)
<i>HITC</i>	−0.090 (0.023)	−0.087 (0.022)	−0.104 (0.021)
<i>treatgroup</i> × <i>HITC</i>	0.065 (0.030)	0.060 (0.028)	0.063 (0.028)
Age	—	0.011 (0.001)	0.005 (0.001)
White	—	−0.012 (0.035)	−0.009 (0.031)
Black	—	−0.057 (0.036)	−0.033 (0.033)
Number of children under age 6	—	−0.046 (0.015)	−0.004 (0.014)
Number of children aged 6-18	—	−0.030 (0.010)	−0.019 (0.009)
Number of children aged 19-24 and a full-time student	—	−0.034 (0.053)	−0.029 (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 <i>R</i> -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: <i>covered</i> by private health insurance	
<hr/>	
<i>treatgroup</i>	−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 × <i>treatgroup</i>	0.025 (0.042)
1989 × <i>treatgroup</i>	0.036 (0.044)
1990 × <i>treatgroup</i>	—
1991 × <i>treatgroup</i>	0.064 (0.046)
1992 × <i>treatgroup</i>	0.072 (0.049)
1993 × <i>treatgroup</i>	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)

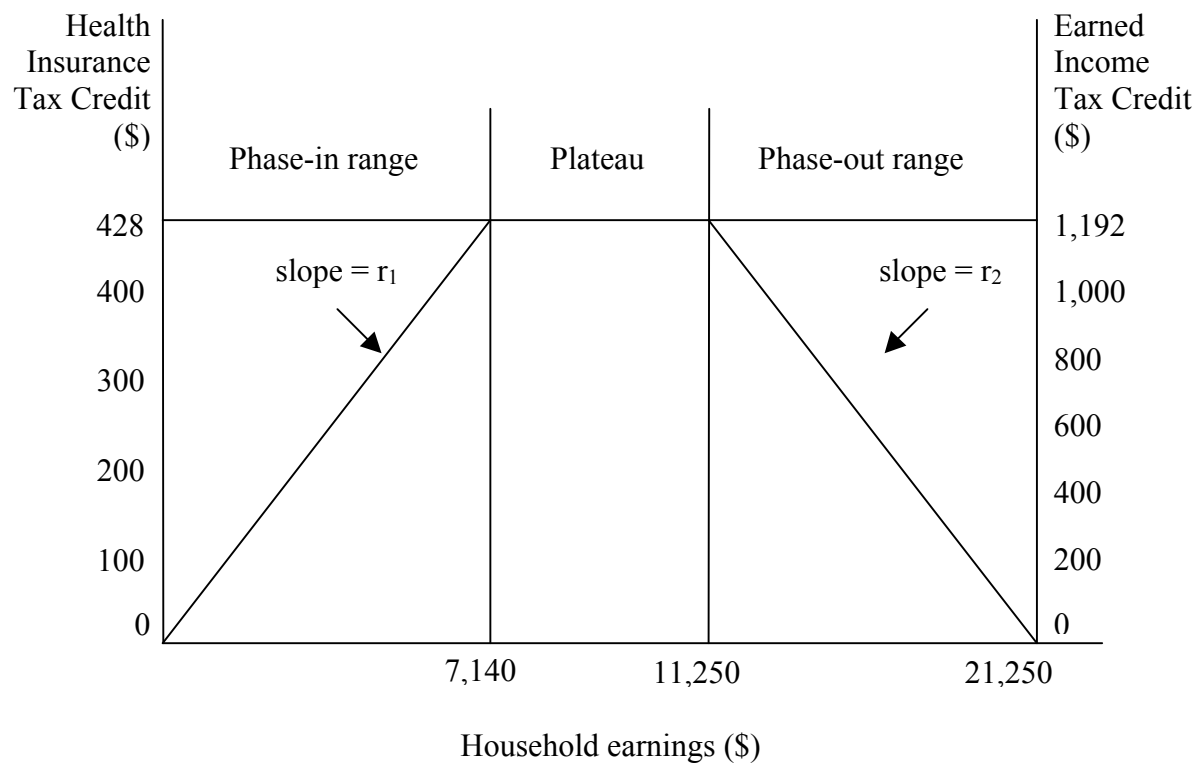
(continued)

Table A4 (continued)

Dependent variable: <i>covered</i> by private health insurance	
<hr/>	
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 <i>R</i> -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:

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

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:

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

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:

$$\frac{\Delta \Pr(y = 1 | \bullet)}{\Delta \text{treatment} \times \text{post}} = \left[\frac{\Phi(\beta_0 + \beta_1 \text{treatment} + \beta_2 \text{post} + \beta_3 + X\beta) - \Phi(\beta_0 + \beta_1 \text{treatment} + \beta_2 \text{post} + X\beta)}{\Delta \text{treatment} \times \text{post}} \right] \quad (\text{A.3})$$

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

$$\begin{aligned} \frac{\Delta \Pr(y = 1 | \bullet)}{\Delta \text{treatment} \Delta \text{post}} &= \frac{\Delta [\Phi(\beta_0 + \beta_1 + \beta_2 \text{post} + \beta_3 \text{post} + X\beta) - \Phi(\beta_0 + \beta_2 \text{post} + X\beta)]}{\Delta \text{post}} \\ &= \left\{ \frac{[\Phi(\beta_0 + \beta_1 + \beta_2 + \beta_3 + X\beta) - \Phi(\beta_0 + \beta_1 + X\beta)]}{-[\Phi(\beta_0 + \beta_2 + X\beta) - \Phi(\beta_0 + X\beta)]} \right\} \quad (\text{A.4}) \end{aligned}$$

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):

$$\Pr(\text{covered}_i=1|\bullet) = \Phi[\beta_0 + \beta_1 \text{treatgroup}_i + \beta_2 \text{HITC}_t + \beta_3 \text{treatgroup}_i \times \text{HITC}_t + X_i \beta]$$

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

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

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

The probit DD estimator is the average of this double difference over all observations.

Obtaining standard errors by the bootstrap method is straightforward.