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Abstract

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Keywords

job lock

THE STATE CHILDREN'S HEALTH INSURANCE PROGRAM AND JOB MOBILITY: IDENTIFYING JOB LOCK AMONG WORKING PARENTS IN NEAR-POOR HOUSEHOLDS

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To assess whether near-poor parents' job mobility is reduced due to the non-portability of employer-provided health insurance—an effect termed job lock—the authors examine data from the Survey of Income and Program Participation for 1996 and 2001, years bracketing the introduction of the State Children's Health Insurance Program (SCHIP). Among the working fathers whose children met the SCHIP eligibility criteria, those whose wives did not have their own employer-provided insurance were 5–6% more likely to separate from their current employer in the year of the later survey than in the year of the earlier survey, whereas those whose wives were insured exhibited no comparable change in mobility. These results confirm the presence of job lock: for men whose wives were uninsured, but not for those whose wives were insured, the authors argue, the SCHIP program presented a new opportunity to switch jobs without losing health insurance.

Working adults with children in the United States are likely to seek employment opportunities that provide health insurance benefits. For a number of reasons, procuring health insurance through an employer is often less expensive than purchasing insurance individually. Unlike the earned income needed to privately purchase coverage, the value of employer-provided coverage is not taxed. In addition, large employers can purchase insurance at a rate per beneficiary that is considerably lower than that faced by individual households on the open market.

These cost advantages suggest that the value that many individuals place on employer-provided health insurance benefits exceeds the concurrent marginal cost to employers. This differential may be exacerbated if alternative employers do not offer health benefits, or refuse to provide coverage for pre-existing medical conditions for up to one year as allowed under the provisions of the Health Insurance Portability and Accountability Act (HIPAA), or impose length-of-service requirements before they provide any benefits. Consequently, some workers may bypass alternative employment with higher pay and superior non-monetary job attributes. Parents in particular, whose children are likely to use health insurance benefits intensively, may find themselves "locked"

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A data appendix with additional results, and copies of the computer programs used to generate the results presented in the paper, are available from the first author at Department of Economics, St. Lawrence University, 23 Romoda Drive, Canton, NY 13617; cbansak@stlawu.edu.

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into particular jobs by the need to maintain health coverage for their children.

The recent expansion of eligibility for public health insurance through the State Children's Health Insurance Program (SCHIP) provides a novel opportunity to assess the degree to which the job mobility of working parents is reduced by the need to maintain health insurance for their children. SCHIP expanded the pool of children eligible for public health insurance benefits from roughly 30% in 1997 (under Medicaid eligibility rules) to roughly 50% in 2001 (both Medicaid and SCHIP combined), effectively de-linking health insurance coverage from parents' employment status for an increasing proportion of children. In this paper, we exploit the introduction of SCHIP as a natural experiment that can help reveal the extent to which near-poor parents are subject to job lock.

Using data from the 1996 and 2001 Survey of Income and Program Participation (SIPP), we first identify working fathers whose children met the SCHIP eligibility criteria. We then evaluate a simple quasi-experiment defining a treatment group and comparison group among the fathers of SCHIP-eligible households based on a characteristic that would be expected, *a priori,* to either magnify or reduce the extent of job lock. Specifically, we compare the pre-post change in one-year separation rates for married men whose spouses had employer-provided coverage in their own names to the corresponding change for married men whose spouses did not have such benefits. For the former group, the introduction of SCHIP is unlikely to have relieved job lock, since they already had a viable alternative source of coverage. For the latter group, however, SCHIP provided an alternative source of coverage where one previously did not exist.

Identifying Job Lock

Most Americans obtain health coverage through the group plans of their employers.[1] However, not all employers offer such benefits. Thus, it is quite natural to ask whether health benefits tie workers to their employers. Even if a worker receives an alternative job offer with health benefits, changing employers and health plans may create several transaction costs and involve spells of being uninsured. For example, those who switch employers may have to switch primary care physicians; the new employer may require initial physical exams and exclude coverage for the treatment of pre-existing conditions for up to a year;[2] the new employer may require some minimum length of service before extending health benefits to new employees; or, if the worker changes a job mid-year, he or she may lose credit toward deductibles, or contributions made toward a pre-tax health care reimbursement account.

If household valuations of these transaction costs are substantial, workers may bypass preferred employment opportunities that arise. Furthermore, one might expect that those individuals who place a particularly high value on their health insurance benefits will be less likely to initiate on-the-job search in the first place.

In cross-sectional data, workers with employer-provided health insurance are considerably less likely to separate from their current employers than are uninsured workers. For example, tabulations of the 1996 SIPP show that roughly 18% of workers with employer-provided health benefits separate from their employers within a year, compared with 41% of uninsured workers. The key identification problem is that there is a high likelihood of correlation between having employer-provided health insurance and both job and worker characteristics that are also likely to be related to mobility. For example, jobs that

[1]Our tabulations of the 2001 SIPP indicate that 84% of working adults in that year were covered by a health insurance plan whereby either their employer or the employer of someone else in the household paid part of the cost. Roughly 70% of children had private health insurance coverage, and among these, the overwhelming majority were covered through the employer group plans of a parent or guardian.

[2]Under the Health Insurance Portability and Accountability Act of 1996, employers must offer coverage for the treatment of pre-existing conditions after one year of service.

offer health benefits probably offer other fringe benefits as well, such as a pension or vacation time. Moreover, employees with health benefits are likely to be more skilled on average, and perhaps more stable, than are those without health benefits.

In their review of job lock research, Gruber and Madrian (2002) identified two principal identification strategies for measuring job lock. The first strategy exploits variation in whether a given worker has access to health insurance coverage through a source other than his or her employer. The second strategy exploits the fact that the personal valuation of one's health insurance benefits will vary with one's personal circumstances, such as the number and composition of dependents or health conditions.

With regard to the first strategy, several studies have sought to determine whether the job mobility effects of employer-provided health benefits are lower for workers with an alternative source of health insurance than for those without. For example, Madrian (1994) investigated whether married men's mobility was affected by whether their spouses were independently insured through their own employers. Among men with insured wives, the wife's health benefits (and possibly the husband's as well) were not dependent on the husband's current employment. For such men, concerns over losing health benefits were unlikely to constrain mobility, and thus any mobility difference between those with and without employer-provided benefits was likely attributable to observable and unobservable job and worker attributes.

On the other hand, for a man without an independently insured spouse, both his coverage and the coverage of his dependents are tied to his current job. Subject to some assumptions, one can attribute to job lock the differential effect of employer-provided health benefits among these men relative to men with insured spouses. Several studies have pursued this strategy, including Madrian (1994), Buchmueller and Valletta (1996), Holtz-Eaken (1994), Anderson (1997), Gilleskie and Lutz (2000), and Adams (2004). With the exception of Holtz-Eaken (1994), all of these studies have found evidence of job lock among workers with no alternative source of health coverage, with the estimates suggesting a 25–50% reduction in job mobility.

The second identification strategy investigates whether the mobility effects of employer-provided health insurance are mediated by the worker's valuation of health benefits. Madrian (1994) compared the mobility of two groups of married men whose employers provided them with health insurance: those with pregnant wives, and those without. She estimated 25% lower voluntary mobility for the former group than for the latter. Kapur (1998) compared the effects of benefits on mobility for workers with and without chronic health conditions. Similar approaches were taken by Brunetti et al. (2000) and Stroupe et al. (2001).

Research focusing on workers' valuation of health benefits has yielded mixed evidence on job lock. However, the bulk of the studies have examined how chronic poor health affects the rate of voluntary turnover among workers in jobs with employer-provided health benefits, and most of these studies have found little evidence of job lock.

An identification strategy that has not been extensively pursued is to exploit variation in policy pertaining to either the availability of alternative sources of health coverage or the portability of existing coverage. Gruber and Madrian (1994) provide the sole exception. The authors explored the effect of state variation in continuous coverage mandates on the likelihood that workers separated from their current employers. Before passage of the Consolidated Omnibus Reconciliation Act of 1985 (COBRA), state regulation governed the length of time that employers were required to allow former employees to buy into their group plans (usually at the average cost per beneficiary to the employer). After 1986, the federal law mandated that employees may retain their health insurance for 18 months after leaving a job. If the state and federal statutes are at odds, firms must abide by the law that provides for more generous coverage. The authors found positive and statistically significant effects of the extension of continuous coverage protection on quarterly job separation rates during the 1980s.

Below, we outline an identification strategy that exploits both differential access to health coverage through sources other than one's employer (following Madrian [1994] and others) and exogenous policy-induced variation in access to public health care.

Using the Expansion of SCHIP to Identify Job Lock among Working Parents in Near-Poor Households

In 1997, Congress created the State Children's Health Insurance Program (SCHIP), providing $40 billion in federal matching funds through fiscal year 2007 for state-designed and -operated public health insurance programs. SCHIP targets children in low-income families with incomes too high to qualify for Medicaid benefits. For the most part, children in families with income less than 200% of the poverty line that are ineligible for Medicaid benefits are eligible for SCHIP,[3] though some states extend coverage to households with income up to 350% of the poverty line. Unlike Medicaid, SCHIP benefits are not an entitlement. States are allotted funds based on a matching formula, and each state is allowed to define the "targeted" group of low-income children to receive health insurance through the SCHIP program.[4]

The introduction of SCHIP greatly expanded the proportion of children eligible for public health insurance. In 1997, 34% of U.S. children were eligible for public health insurance, all through the Medicaid program. By 2001, the percentage of children eligible for public health insurance had climbed to 51%, with 19% eligible for SCHIP benefits and 32% eligible for Medicaid (Bansak and Raphael 2007). Restricting the focus to uninsured children, as of 2001, roughly half were eligible for Medicaid benefits while a quarter were eligible for SCHIP benefits.

Our identification strategy for measuring job lock compares the pre-post SCHIP changes in mobility rates among employed fathers of SCHIP-eligible children after we stratify these parents into groups who are likely to differ in terms of the degree to which they were locked to their current employer through health benefits. Specifically, for the period surrounding the introduction of SCHIP, we compare the pre-post change in mobility rates for married fathers whose wives were independently insured through their own employment to the corresponding change for married fathers whose wives did not have such benefits. For men in the former group, the introduction of SCHIP provided a second alternative source of health insurance, since household dependents were likely to be eligible for benefits through the spouse's employer. Thus, for this group, the introduction of SCHIP affected a non-binding (or, perhaps, less-binding) constraint on job mobility.[5]

On the other hand, for parents without independently insured spouses, the SCHIP program provided the first alternative source of health insurance for their dependent children. For these parents, the program directly affected a binding constraint on job mobility. To the extent that job lock is important among SCHIP-eligible parents, one should observe an increase in the job mobility rates of parents without an insured spouse relative to those with an insured spouse.

Using these defined treatment and comparison groups, we tabulate difference-in-difference estimates of the program's introduction on job mobility rates. We calculate the before-after change in one-year separation

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[3]While SCHIP is aimed at low-income children, some groups of low-income children are not eligible. For example, children eligible for Medicaid and children who are members of families currently eligible for state employee insurance are not eligible to receive coverage. Children who live in an "institution for mental diseases" are also ineligible. For a fuller discussion of eligibility criteria, see Bansak and Raphael (2007).

[4]Each state has a fixed allotment of SCHIP funds that are distributed as a federal match with an enhanced matching rate, ranging from 65% to 85% (Green Book 2004). State allotments are determined through a formula that takes into account both the "number of children" and a "state cost factor" that reflects the cost of health care in a given state (Bansak and Raphael 2007).

[5]To be sure, SCHIP does not as a rule relieve job lock due to concern about one's own health benefits, as the SCHIP program extends benefits to parents in only a handful of states. Thus, while the program partially unlinks health insurance coverage and the employer for dependents, it does not sever the relationships entirely.

Table 1. Proportion of Parents of SCHIP-Eligible Children Whose Children Had Publicly Provided Health Insurance, 1996 and 2001.

(Standard Errors in Parentheses)

Description	*1996*	*2001*	*Δ (2001–1996)*
Panel A: Married Men			
Insured Spouse	0.035 (0.011)	0.078 (0.008)	0.042 (0.019)**
No Insured Spouse	0.077 (0.008)	0.145 (0.012)	0.068 (0.014)***
Diff.-in-Diff.	—	—	0.026 (0.028)
Panel B: Married Men with Employer-Provided Health Insurance			
Insured Spouse	0.049 (0.018)	0.100 (0.028)	0.050 (0.032)
No Insured Spouse	0.052 (0.008)	0.080 (0.011)	0.027 (0.013)**
Diff.-in-Diff.	—	—	–0.023 (0.033)

**Difference is statistically significant at the .05 level; ***at the .01 level.

rates for men with independently insured spouses and for men without, then test to determine whether the two groups differ with respect to this change.

Our difference-in-difference estimator rests on several important assumptions that merit discussion. First, we are assuming that our stratification actually identifies variation in the degree to which parents are locked to their employer due to concerns about health insurance coverage for their children. Second, we are assuming that the mobility trends for fathers in our comparison group provide adequate counterfactuals for those we are identifying as being treated by the SCHIP program.

There are several reasons to question the first assumption. For one thing, in households with two employed and independently insured parents, one might surmise that the parents would be likely to enroll their dependent children in the group plan that offers the best benefits at the lowest cost. In other words, the benefits offered by the wife's employer may not perfectly substitute for those offered through the husband's own employer. Moreover, having an independently insured spouse does not necessarily imply that one's children are eligible for benefits through the spouse's health plan.[6] Thus, even for parents with an insured spouse, health insurance concerns may bind them to their current employers.

Despite these concerns, we do observe some evidence of a relatively larger take-up rate among the SCHIP-eligible children of married men without insured spouses, although this pattern is not consistent across the samples that we use for our analysis. Table 1 presents these comparisons.[7] For the years 1996 (before SCHIP's introduction) and 2001 (several years into the program's life), the table presents the proportion of working fathers in SCHIP-eligible households with children who were covered by publicly provided health insurance. Panel A presents figures for all married men, while Panel B presents tabulations for married men who had employer-provided health insurance benefits. Among all fathers with insured spouses, the proportion with children receiving public health benefits rose from 0.035 in 1996 to 0.078 in 2001, a statistically significant increase of 4.2 percentage points.[8] In

[6]The questions in the SIPP pertaining to health insurance allow determination of whether a spouse is covered by a plan that is paid for in part by the spouse's employer. These questions do not, however, reveal whether the spouse's benefits would cover minor dependents in the household.

[7]These tabulations are calculated using data from the 1996 and 2001 SIPP. We discuss the data in detail below.

[8]Presumably, the proportion of SCHIP-eligible children covered by public health insurance in 1996 was zero, as such children were ineligible for Medicaid and SCHIP was introduced in 1997. The positive proportion receiving benefits pre-SCHIP reflects error in our imputation of the SCHIP-eligible population. Previous studies of take-up in the Medicaid and SCHIP programs have encountered similar problems with respect to observed take-up among presumably ineligible households (for example, see Cutler and Gruber 1996; LaSasso and Buchmueller 2002; Bansak and Raphael 2007).

Table 2. Mean Characteristics of Married Men with and without Spouses Who Carried Their Own Employer-Provided Health Insurance.
(Standard Errors in Parentheses)

Variable	*Insured Spouse*	*No Insured Spouse*	*Difference*
Black	0.167 (0.016)	0.082 (0.006)	0.084 (0.014)***
Asian	0.045 (0.009)	0.039 (0.004)	0.006 (0.009)
America Indian	0.008 (0.003)	0.008 (0.002)	0.000 (0.004)
Hispanic	0.131 (0.015)	0.174 (0.009)	–0.043 (0.018)**
Age	37.920 (0.335)	37.861 (0.175)	0.059 (0.379)
Enrolled in School	0.056 (0.009)	0.052 (0.005)	0.004 (0.011)
Veteran	0.215 (0.017)	0.157 (0.008)	0.058 (0.018)***
High School Dropout	0.146 (0.015)	0.139 (0.008)	0.007 (0.17)
High School Graduate	0.443 (0.021)	0.361 (0.010)	0.082 (0.013)***
Some College, No Degree	0.174 (0.016)	0.197 (0.009)	–0.023 (0.019)
Associate Degree	0.119 (0.014)	0.124 (0.007)	–0.004 (0.016)
Bachelors	0.093 (0.013)	0.126 (0.007)	–0.034 (0.015)
Masters or Higher	0.024 (0.006)	0.052 (0.005)	–0.028 (0.010)
Has Employer-Provided Health Benefits	0.484 (0.022)	0.781 (0.009)	–0.297 (0.021)***
Union	0.159 (0.015)	0.234 (0.009)	–0.076 (0.020)***
Log Wages	2.225 (0.021)	2.487 (0.009)	–0.261 (0.022)***
N	532	1,937	—

**Difference is statistically significant at the .05 level; ***at the .01 level.

contrast, the proportion of working fathers without insured spouses whose children were covered by public health insurance increased from 0.077 in 1996 to 0.145 in 2001, a slightly larger increase of approximately 7 percentage points. When the sample is restricted to married men with employer-provided health insurance, the relative change flips signs, with the children of men with insured spouses experiencing a larger increase in public coverage relative to the children of men without insured spouses. However, the change for men with insured spouses is poorly measured and is statistically insignificant, as is the relative change across the two groups.

Whether our proposed comparison group provides an adequate gauge of the counterfactual path that the mobility rates of our treatment group would have followed had SCHIP not been introduced is, of course, impossible to assess. However, we can compare the two groups in terms of pre-intervention average characteristics and compare pre-intervention values for our key dependent variable to assess the degree of similarity between our treatment and comparison groups. Table 2 presents average values for a host of observable characteristics of the married fathers of SCHIP-eligible children with and without independently insured spouses. While there are a few notable differences in basic demographic and human capital characteristics (men with insured spouses were more likely to be black and a veteran and less likely to be Hispanic), age, the overall proportion minority, and the variables measuring educational attainment are fairly similar. However, there are large differences in average wages and the fraction union, both factors that have been shown in previous research to be strongly correlated with the likelihood of a job separation (Bansak and Raphael 2006). Moreover, the proportion of these men who had employer-provided health benefits was nearly 30 percentage points higher for those with uninsured spouses than for those without.

To account for these differences in observable covariates, we take several steps. To begin, in addition to presenting difference-in-difference estimates based on the sample used to tabulate the figures in Tables 2, we present estimates based on a somewhat more restrictive subsample. Specifically, we present separate difference-in-difference estimates using (1) the sample of all married fathers of

SCHIP-eligible children and (2) the sample of all married fathers of SCHIP-eligible children with employer-provided health benefits.

The two samples we use carry various costs and benefits. With regard to the more inclusive sample, our treatment and comparison groups look less similar *a priori*. Moreover, among the treated men in the sample were some who had no employer-provided health insurance—men who, by definition, could not be job locked. However, using this larger sample obviates a concern that the program itself might have affected the composition of those without employer-provided health insurance. In particular, if the introduction of SCHIP caused some fathers to move from jobs with insurance benefits to jobs without, and if these men on the margin were less stable than other men with health benefits who did not respond in this manner, the post-program sample of men with benefits will be more negatively selected (in terms of separation probabilities) than the pre-program sample. The effect of this selection on the difference-in-difference estimate will depend on the relative responsiveness along this margin of men with independently insured spouses as compared to men without. Our unrestricted sample bypasses this potentially thorny selection problem. Moreover, the unrestricted sample affords a somewhat larger sample size.[9]

A clear advantage of further restricting the sample to married men with employer-provided health benefits is improvement of the balance between the treatment and comparison groups along observable dimensions. In particular, conditioning the sample on having employer-provided health benefits eliminates the disparity in health coverage, narrows the average wage differential slightly, and eliminates the difference in the proportion union. Moreover, as is evident in Table 3, pre-program separation rates were considerably more alike in the restricted sample than in the more inclusive sample. However, this additional restriction reduces our sample size and raises concerns regarding sample selection. Given the relative benefits and problems associated with these two samples, we present results for both through the remainder of the paper.[10]

In addition to restricting the analysis samples in this manner, we also present regression-adjusted difference-in-difference estimates that account for any remaining differences in the observable characteristics listed in Table 2. To illustrate this adjustment, let $NoSpouse_i$ be an indicator variable equal to one if parent i does not have an independently insured spouse and $Y2001_i$ be an indicator variable for observations from the post-SCHIP sample. The regression-adjusted difference-in-difference model comes from estimating the linear probability model

$$(1) \quad Separation_i = \alpha_0 + \alpha_1 NoSpouse_i + \alpha_2 Y2001_i + \alpha_3 NoSpouse_i * Y2001_i + \beta' X_i + \varepsilon_i,$$

where $Separation_i$ is a dummy variable indicating that parent i separated from his employer within the year, X_i is a vector of the observable characteristics, ε_i is a disturbance term, and α_0, α_1, α_2, α_3, and β are parameters to be estimated. The parameter α_3 provides the difference-in-difference estimate after

[9]While we present estimates for this more inclusive sample to explore whether our findings from the sample restricted to married men with insurance benefits are robust to the potential selection bias noted above, these estimates can also be interpreted as a policy evaluation of an imperfectly targeted intervention to relieve job lock. Specifically, if the intent of SCHIP was to relieve job lock among low-earning workers, the "intent-to-treat" effect of such an intervention would be dulled by the existence of newly eligible working adults without health insurance benefits at their current job.

[10]In addition to the above-mentioned costs and benefits of these two samples, it is worth noting that the treatment and control groups in the first sample—all married fathers of SCHIP-eligible children (the larger sample)—may be comprised of different subgroups who could be affected differentially by SCHIP. For example, the treatment group—men without spousal coverage—is comprised of some men with EPHI and some without. The latter group is less likely to be job-locked and less likely to be affected by SCHIP. Furthermore, the control group—men with spousal coverage—also is comprised of some men with EPHI and some without. While we do not think there will be much of a SCHIP effect for these subgroups of the control group, some increase in mobility is possible. Parents in the dual coverage group could become self-employed or open their own business; meanwhile, in the group with spousal coverage and no EPHI, the spouse could become more mobile and hence increase the mobility of the married father.

adjusting for the variables included in the vector X_i. We discuss the exact specifications of these models below with the presentation of the results.

Data Description

The data for this project come from the public release files of the Survey of Income and Program Participation (SIPP), a monthly longitudinal household survey. Every four months for several years, each SIPP respondent is interviewed to obtain detailed retrospective information pertaining to demographics, employment, income receipt, and other variables. Thus wave 1 of the 1996 SIPP corresponds to the first four months of the panel, wave 2 corresponds to months 5 through 8, and so on. For each of the 1996 and 2001 panels, we merge data from wave 1 and wave 4. In this section, we detail our method for imputing SCHIP eligibility, the manner in which we gauge the labor market mobility of the parents of eligible children, our characterization of the benefits available to parents and their spouses through employer-provided group plans, and the additional sample restrictions that we place on the analysis sample.

Identifying Employed Parents of SCHIP-Eligible Children

Using the 1996 and 2001 Panels, we first identify all children 18 years of age and under and impute SCHIP eligibility based on family income and composition. We identify children who were eligible for SCHIP benefits in 2001 as well as children who would have been eligible in 1996 (under 2001 income criteria) had the program been in existence. Identifying children in the SIPP who were eligible for public health insurance benefits requires (a) information on family income net of allowable disregards and (b) state-level information on Medicaid and SCHIP eligibility criteria. The income eligibility criteria for both Medicaid and SCHIP are based on family net income relative to the federal poverty line.

To gauge income, we first adjust household income for allowable childcare and work-related expenses. We deduct $2,500 in child-care expenses from annual household income for each child in the household and an additional $1,080 for work-related expenses. We then divide the remaining household income by the federal poverty line corresponding to the state of residence,[11] household size, and year.

To determine which children in the 1996 panel (waves 1 and 4) were hypothetically eligible for SCHIP benefits, we identify those who met the SCHIP income criteria for 2001 but did not meet the 1996 Medicaid criteria. Note that since SCHIP did not exist in 1996, this group of children essentially represents the SCHIP target group prior to the program's implementation (see Bansak and Raphael [2007] for a detailed discussion of this imputation).

For the 2001 panel, we apply the 1996 Medicaid criteria to identify Medicaid-eligible children and the 2001 SCHIP income criteria in conjunction with Medicaid income and age limits to identify the SCHIP-eligible population. Note that this schema attributes all expansions in coverage between 1996 and 2001 to the introduction of SCHIP.[12]

Inspection of pre-post program changes in health insurance status around the SCHIP cutoff reveals large and statistically significant increases in the proportion of children covered by public health insurance up to approximately the state SCHIP income cutoff plus one-half of the poverty line. Beyond this point, observed changes in the fraction covered by public health insurance are small and statistically insignificant. These patterns suggest that our basic imputation underestimates the SCHIP-eligible population. To account for this under-estimation, we designate all children with income above the Medicaid cutoff but below the SCHIP eligibility cutoff plus one half of the poverty line as SCHIP-eligible.

Once we have identified children who met the income eligibility requirements for

[11]The federal poverty line varies by household size and is the same for all states except Hawaii and Alaska.

[12]Note that several states provide SCHIP benefits through an expansion of their existing Medicaid programs, and thus in these states Medicaid eligibility criteria are currently more generous than they were in 1997.

SCHIP in each year, we then identify the mothers and fathers of these children (either both parents or only one, depending on who was present in the household), using the mother and father identification codes provided in the SIPP's records of the children and the personal identification codes for the parents. At this point, we restrict the adult sample to married fathers of SCHIP-eligible children who were employed in the first month of each panel.

Measuring Job Mobility

To measure job mobility, we construct an indicator variable for each employed father of a SCHIP-eligible child that is equal to one if the father separated from his employer over the course of one year. We identify job separators from a series of employer identification codes constructed from the interview control cards used by the SIPP surveyors.

In the first-wave interview, the SIPP interviewers recorded the identity of the respondent's primary and secondary employers on an interview control card that was used in all subsequent interviews. Each employer was assigned a consecutively numbered employer identification value. In subsequent interviews, if the respondent's primary or secondary employers matched either the primary or secondary employers recorded in previous interviews, the employer identification variables remained unchanged. When the worker changed employers, the new employer name was recorded on the control card and the next available employer identification number was assigned. If the worker was unemployed or had left the labor force, the employer identification code was set to missing (not in universe).

We define job separations relative to the respondent's primary employer as of the first month of the panel. To do so, we compare the employer I.D. of the primary employer in month 1 of the panel to the employer I.D.'s of the individual's primary and (if relevant) secondary employers in month 13 of the panel. If the month 1 I.D. does not equal the I.D. number for either the primary or secondary employer in month 13 (either because the worker had switched employers or because he was not employed in month 13), then we code the parent as having separated from his initial primary employer.

We explored a number of alternative methods for constructing the job separation variable. For example, we merged waves 1 through 4 of each panel and defined a separation as any break in the sequence of employer I.D.'s over the 13-month period. We also computed separation rates that required any break from the primary employer to persist for at least 6 months. All three methods produced nearly identical one-year separation rates (with approximately 25% of the sample separating within one year in each panel). However, constructing the separation rate by matching wave 1 to wave 4 yielded the largest sample size, since this approach requires a completed interview in only two rather than four separate waves.[13]

Characterizing the Insurance Status of Fathers and Their Spouses

For all parents of SCHIP-eligible children, we determine whether the parent had employer-provided health insurance by making use of two questions in the SIPP. First, all respondents were asked whether they were covered by a health insurance plan in their own name or in someone else's name. Respondents indicating either that they were covered in their own name or that they were covered by both a plan in their own name and someone else's plan are further asked, regarding the plan in their own name, whether their employer or union covered all or part of the costs of this plan. We code those adults indicating that they had coverage in their own name and that either their employer or union bore part of the cost as having employer-provided health insurance.

[13]We also explored using information from the topical modules accompanying each core wave of the SIPP pertaining to the reason behind a separation—for example, voluntary versus involuntary job separations. However, this variable pertains only to job separations that occurred within waves. A detailed analysis of these data (Bansak and Raphgael 2006) shows that the majority of separations were recorded between waves (a common "seaming" problem that has been noted previously in the SIPP), and thus information on the reason for the switch is missing for the majority of observations.

Table 3. One-Year Separation Rates for Parents of SCHIP-Eligible Children, by Insurance Status of the Spouse.
(Standard Errors in Parentheses)

Description	*1996*	*2001*	*Δ (2001–1996)*
Panel A: Married Men			
Insured Spouse	0.346 (0.028)	0.241 (0.028)	–0.105 (0.039)***
No Insured Spouse	0.208 (0.012)	0.264 (0.015)	0.056 (0.019)***
Diff.-in-Diff.	—	—	0.162 (0.042)***
Panel B: Married Men with Employer-Provided Health Insurance			
Insured Spouse	0.208 (0.034)	0.179 (0.036)	–0.029 (0.049)
No Insured Spouse	0.160 (0.013)	0.221 (0.016)	0.061 (0.020)***
Diff.-in-Diff.	—	—	0.090 (0.053)*

*Difference is statistically significant at the .10 level; ***at the .01 level.

For each identified parent of an SCHIP-eligible child, we match the parent to his or her spouse (if the spouse was residing in the household) using the spouse I.D. codes provided in the SIPP. We obtain spouses from the unrestricted sample of adults, in order to make sure we are capturing households with remarried parents where children may not have been living with both biological parents. Using the same two insurance questions discussed above, we then code whether the spouse had employer-provided health insurance benefits.

Additional Sample Restrictions

Throughout our analysis, the sample is restricted to the married fathers of SCHIP-eligible children. For each year, we also restrict the sample to fathers who were employed in the first month of the panel. We impose several additional restrictions on the samples drawn from the 1996 and 2001 panels. We eliminate family workers and parents who were members of the armed forces. We also restrict the sample to parents between the ages of 18 to 65 years as of the beginning of each panel. Finally, we discard all observations with incomplete interviews in either wave 1 or wave 4 of each panel.

Empirical Results: Job Lock and Labor Mobility

Table 3 presents a simple preliminary difference-in-difference analysis in job separation rates that does not adjust for observable covariates. The table shows estimates of the proportion of employed fathers who separated from their employers within one year, with figures presented for 1996 (one year before the introduction of SCHIP) and 2001 (four years after), and by whether the worker had an independently insured spouse. Recall that the analysis sample is restricted to parents of SCHIP-eligible children in 2001 and parents in the target income range of the SCHIP program in 1996. We present results for the sample of all married men and for the sample of married men with employer-provided health insurance.

Beginning with the results for all married men, we find a large, sizable, and statistically significant decline in the separation rates of men with insured spouses of 10.5 percentage points. Among married men without insured spouses, the one-year separation rate increased by 5.6 percentage points, a change that is significant at the 1% level. The relative increase in separation rates is fairly large (16.2 percentage points) and is statistically significant ($p < .01$).

Panel B presents results restricting the sample to married men with employer-provided insurance. The overall separation rates for such workers were considerably lower than the overall separation rates for all fathers of SCHIP-eligible children presented in Panel A. Notably, the pre-program separation rates were also considerably closer to one another than the comparable separation rates in Panel A. Among those with insured spouses,

Table 4. Regression-Adjusted Estimates of the First Difference (2001 minus 1996) and Difference-in-Difference in the One-Year Separation Rate: Relative Comparisons Based on the Insurance Status of the Spouse.

(Standard Errors in Parentheses)

Specification	$\Delta_{1996\ to\ 2001}$ *with an Insured Spouse*	$\Delta_{1996\ to\ 2001}$ *without an Insured Spouse*	*Difference-in-Difference*
Panel A: Married Men			
Specification (1)	–0.105 (0.040)***	0.056 (0.019)***	0.162 (0.042)***
Specification (2)	–0.100 (0.041)**	0.045 (0.019)**	0.152 (0.041)***
Specification (3)	–0.102 (0.044)**	0.044 (0.019)***	0.153 (0.042)***
Specification (4)	–0.098 (0.045)**	0.058 (0.019)***	0.152 (0.043)***
Panel B: Married Men with Employer-Provided Health Insurance			
Specification (1)	–0.029 (0.049)	0.061 (0.020)***	0.090 (0.053)*
Specification (2)	–0.036 (0.055)	0.056 (0.020)***	0.090 (0.053)*
Specification (3)	–0.054 (0.063)	0.058 (0.021)***	0.087 (0.054)
Specification (4)	–0.052 (0.065)	0.069 (0.021)***	0.092 (0.054)*

Notes: Specification (1) is the raw difference with no controls. Specification (2) adds all of the control variables listed in Table 4 with the exception of wages, plus twelve industry dummies and six occupation dummies. Age is entered as a third-order polynomial. Specification (3) adds a full set of state fixed effects and a complete set of dummy variables for income relative to the poverty line (measured in 25 percentage point increments). Specification (4) adds log wages. For the difference-in-difference estimates, all specifications include a dummy variable for not having an insured spouse along with an interaction term between this variable and the 2001 year dummy. In this model, the effects of the explanatory variables are constrained to be constant across the two groups.

*Difference is statistically significant at the .10 level; **at the .05 level; ***at the .01 level.

the separation rate declined slightly over the study period (a statistically insignificant decline of 2.9 percentage points). For those without insured spouses, the separation rates increased by 6.1 percentage points (p-value < .01). The unadjusted difference-in-difference calculation yields a job lock estimate of 9 percentage points (p-value = 0.0904).

To be sure, the estimates using the larger, more inclusive sample in Panel A are implausible, with the large point estimate of 16.2 percentage points being driven by the large observed decline in separation rates among our quasi-experimental control group. Nonetheless, we do indeed observe a statistically significant increase in separation rates among fathers without insured spouses. At a minimum, we can certainly conclude that the raw changes indicate a strongly significant relative increase in the separation rate among those whom we believe to have been affected by the introduction of SCHIP. The relatively modest estimates in Panel B for the smaller sample are clearly more plausible. In fact, these results using the restricted sample (along with the regression-adjusted results presented in Tables 4 and 5) are our preferred estimates. Granted, given the smaller sample size on which they are based, these results are considerably less precise than those in Panel A, and may be subject to the selection bias discussed above. Nonetheless, in combination, these basic results do suggest that job lock was an important factor for fathers whose labor market earnings were in this portion of the earnings distribution.

The unadjusted results in Table 3 reveal an increase in separation rates among those fathers whom we, *a priori,* designated as more likely to be locked to their current employer. However, the patterns we observe may be driven by changes in either observable or unobservable factors that influence separation rates. To probe the robustness of our results, we turn to our difference-in-difference job lock estimates that control for observable characteristics. Table 4 presents a series of regression-adjusted estimates of the before-after change in separation rates for those married men with insured spouses, the comparable change for those with uninsured spouses, and the difference-

Table 5. Difference-in-Difference Estimates of Job Lock Effects for Married Fathers of SCHIP-Eligible Children and Married Fathers in Households with Incomes above the SCHIP Eligibility Cutoffs.

(Standard Errors in Parentheses)

Specification	Δ^2, *Fathers of SCHIP-Eligible Children*	Δ^2, *Higher Income Fathers*	$\Delta^3 = \Delta^2$*SCHIP-Eligible* – Δ^2*Higher Income*
Panel A: Married Men			
Specification (1)	0.162 (0.042)***	0.022 (0.028)	0.139 (0.050)***
Specification (2)	0.152 (0.041)***	0.028 (0.028)	0.122 (0.049)***
Specification (3)	0.153 (0.042)***	0.034 (0.028)	0.118 (0.049)**
Specification (4)	0.152 (0.043)***	0.035 (0.028)	0.118 (0.050)**
Panel B: Married Men with Employer-Provided Health Insurance			
Specification (1)	0.090 (0.053)*	0.019 (0.033)	0.070 (0.061)
Specification (2)	0.090 (0.053)*	0.019 (0.033)	0.066 (0.061)
Specification (3)	0.087 (0.054)	0.025 (0.033)	0.070 (0.061)
Specification (4)	0.092 (0.054)*	0.022 (0.033)	0.078 (0.062)

Notes: See notes to Table 4.

in-difference in these changes in separation rates. We present estimates for each of the two alternative samples used in Table 3 using four different model specifications. The first specification presents the unadjusted estimates reproduced from Table 3. Specification (2) adds twelve industry dummies, six occupation dummies, age squared, age cubed, and all of the covariates listed in Table 2 with the exception of wages. Specification (3) adds a full set of state fixed effects, as well as a set of dummy variables gauging household income relative to the poverty line in 25-percentage-point increments.[14] Finally, specification (4) adds log-wages to all of the variables in specification (3). Here we only present the adjusted first-differences and the adjusted difference-in-difference estimates.

Beginning with the first-difference results for men with insured spouses, separation rates declined for all married men (Panel A) in all specifications, with statistically significant declines in all models. Among married men with employer-provided health insurance (Panel B), all of the point estimates are negative but none are statistically significant.

For those parents without insured spouses, the change in the one-year separation rate was positive and statistically significant (at either the 1% or 5% level) in all models in both panels. For all married men, the change in separation rates ranged from 4.4 to 5.8 percentage points. Among married men with employer-provided health insurance, estimates of the increase in job separation rates range from 5.8 to 6.9 percentage points.

The difference-in-difference estimates all show a relative increase in separation rates among fathers with uninsured spouses. Moreover, in all models, the difference-in-difference estimate exceeds the first-difference estimate for parents without insured spouses. All of these estimates are sizable and significant at the 1% level for the sample of all married men. For married men with employer-provided health insurance, three of the four difference-in-difference estimates are significant at the 9% level, and one estimate (the regression using specification 3) is not statistically significant. Interestingly, for both samples the relative change in separation rates survives controlling for household income as well as the father's current wage level.

A Falsification Check: Do We See Similar Patterns for the Higher-Income Fathers of Ineligible Children?

In this section, we conduct a simple falsi-

[14]That is to say, we include a set of dummies indicating whether a household has income that is between 100% and 125% of the poverty line, or between 125% to 150%, etc., covering the full support of this variable among SCHIP-eligible households.

fication check of our results. Specifically, we repeat the analysis described above, but focusing on fathers in households with incomes *above* the imputed eligibility cutoffs. To the extent that the relative changes observed in Tables 3 and 4 are being driven by the introduction of SCHIP alone, one would not expect to see similar relative changes among the working parents of children who were not eligible for SCHIP benefits.

Table 5 presents the results from this exercise. The first column of figures reproduces the unadjusted and adjusted difference-in-difference estimates for the fathers of SCHIP-eligible children presented in Table 4. The next column presents comparable estimates for fathers in households with income ranging from our imposed SCHIP cutoff to the SCHIP cutoff plus twice the poverty line.[15] Beginning with the results for all married men in Panel A, the relative change in separation rates for men without insured spouses ranges from 2.2 to 3.5 percentage points. None of the estimates are statistically significant, and all are substantially smaller than the difference-in-difference estimates for fathers of SCHIP-eligible children. For married men with employer-provided health insurance, the difference-in-difference estimates for higher income fathers range from 1.9 to 2.5 percentages points. Again, none of these estimates are statistically significant and all are less than the corresponding estimates for SCHIP-eligible fathers.

Finally, shown in the last column of Table 5 is a series of triple-difference estimates, equal to the difference-in-difference estimate for fathers of the SCHIP-eligible minus the difference-in-difference for the higher-income fathers of ineligible children.[16] For each calculation, we note whether the positive triple-difference estimate is statistically significant (a test of whether the higher difference-in-difference estimate for the SCHIP-eligible is statistically distinguishable from the smaller estimates for the ineligible). For all married men, all of the triple-difference estimates of a job lock effect are positive, sizable, and statistically significant at either the 1% or 5% level. These point estimates range from 11.8 to 13.9 percentage points and are slightly lower than the difference-in-difference estimates using this sample.

For married men with employer-provided health insurance, all of the triple difference estimates are again positive. However, here none of the estimates are statistically significant, as the two difference-in-difference estimates used to produce this estimator are poorly measured.

Conclusion

Our analysis of data from two SIPP surveys, one conducted before the start of the State Children's Health Insurance Program (SCHIP) and one conducted after the program had been operating for a few years, shows that married men who were fathers of SCHIP-eligible children and whose wives did not possess health insurance through their own employers were 5–6% more likely to separate from their current employer after the introduction of the SCHIP program, whereas otherwise similar men whose wives did possess independent coverage exhibited no comparable change. The difference

[15]We chose the width of the higher income range to match approximately the width of the income range corresponding to SCHIP-eligible households (a range roughly 1.8 times the poverty line). We experimented with smaller and larger widths for this upper income range and found results quite similar to those presented in Table 5.

[16]We calculate the triple-difference estimates using a linear probability model in which the dependent variable is the separation rate and the key explanatory variables are a dummy for an SCHIP eligible household, a dummy for the year 2001, a dummy for those with uninsured spouses, a complete set of two-way interactions between these three dummies, and a triple interaction term between these three dummies. The triple difference estimate comes from the coefficient on the three-way interaction term and is interpreted as the degree to which the differential increase for men with uninsured spouses is greater for the fathers of SCHIP-eligible children. For the triple-difference estimates that adjust for observable covariates, the triple difference may not be exactly equal to the difference in the difference-in-difference estimators presented. This is because the difference-in-difference estimates in the first two columns allow the effects of the included covariates to vary across income groups (due to the estimation of two separate models), while the figures in the third column constrain the effects of these covariates to be constant.

between the two groups' change in likelihood of separation from their employers was substantial and statistically significant. This relative increase in the likelihood of separation affected only fathers of SCHIP-eligible children, not fathers in the next highest income band. We argue *a priori* that the separation rates for men with independently insured spouses should be less sensitive to the introduction of SCHIP, given that these men already have a likely alternative source of insurance coverage for their children, than should the separation rates for men whose spouses are not independently insured. Thus, the basic patterns suggest that the program's introduction did indeed relieve job lock among near-poor working parents.

These relative changes in separation rates indicate that job lock is indeed a significant factor among near-poor families. But how large are these effects? And how do our estimates of job lock compare to previous estimates?

To be sure, there is an inherent difference between the implicit model underlying our experiment and models underlying previous research on job lock. In previous research, having employer-provided health insurance has served as a proxy for the differential valuation of health benefits by employees (relative to employers' valuation of the cost of providing such benefits) as well as for the transaction costs associated with switching plans when moving between alternative employment opportunities. That is to say, the corpus of existing research focuses on estimating the partial correlation coefficient on a dummy for possession of employer-provided health insurance in a model of employment mobility. In these models, having health insurance ties one's own coverage as well as the coverage of one's dependents to one's current employment situation. Thus, an employer-provided benefits dummy serves as a proxy for being tethered to one's current job by the need to maintain health insurance for everyone in the household.

In the present exercise, the expansion of public health insurance to near-poor families loosens job lock or relaxes the constraint associated with having one's children's health benefits tied to one's current employment status. In general, the program does not relax the constraint with respect to one's own health benefits, since only a small group of states extend benefits to the parents of SCHIP-eligible children.[17] Thus, the first-difference and difference-in-difference estimates are essentially estimating the effect of relaxing the constraint with respect to an employee's dependents (at least some of the employee's dependents) but not with respect to the employee. If public health insurance benefits were extended to all members of the household, one might expect to find larger effects on mobility than we have found for the SCHIP program.

Several of the calculations presented above serve as alternative estimates of the job lock effect. The first-difference increases in separation rates for fathers without insured spouses, for example, would serve as a lower bound. Among all married fathers, we observe an increase in the separation rate of approximately 5.5 percentage points. For married fathers with insurance benefits, the comparable increase is 6 percentage points. Given the baseline separation rates for these groups reported in Table 3, these first difference estimates suggest job lock effects ranging from 29% to 37%.

Our difference-in-difference estimates are larger in all comparisons, with an unadjusted estimate of 15 percentage points for all married men and 9 percentage points for married men with health benefits. Again based on the baseline separation rates presented in Table 3, these point estimates provide job lock effects of 72% and 56%, respectively. A similar exercise with the triple difference estimates provides corresponding job lock effects of 56% and 44%. Given that the large difference-in-difference estimates for the all-married-men sample is driven by the large decline in separation rates for our comparison group, we favor the more conservative yet less precisely measured estimates using the sample of married men with employer-provided health insurance.

[17]In 2001, only Minnesota, New Jersey, Rhode Island, and Wisconsin extended benefits to the parents of SCHIP-eligible children.

The existing research has found job lock effects ranging from zero to roughly 40%. Most of the studies identifying job lock from the interaction between spousal insurance and employer-provided health insurance have found effects ranging from 20% to 40%. For example, Madrian (1994) found that mobility rates were reduced by roughly 25–30%. Buchmueller and Valleta (1996) found job lock effects of roughly 25–30%, and Anderson (1998) found effects of 20–40%. Our estimates range from 29% to 72%, with our preferred estimate range being from 29% to 56%. Thus, the range of our results largely overlaps with the range of estimates from previous studies.

The findings in this study do indeed indicate that job lock is a significant factor in the labor supply and mobility decisions of parents in near-poor households. Moreover, in addition to affecting mobility rates, job lock may have an impact on the average quality of job matches in the economy, with respect to both pecuniary and non-pecuniary working conditions of employees. Future research should focus on how wage levels and growth, as well as non-wage job attributes, are affected by job lock. For example, one might observe a relative increase in the proportion of the treatment group working standard hours rather than non-standard hours. Such parents may move toward safer jobs or jobs that offer a better mix of other non-wage benefits such as pension, sick time, or vacation benefits. While these questions are unanswerable using the data we assembled for this study, some of them could be explored with various monthly supplements of the Current Population Survey. Given the size of the mobility effects documented here, this is a potentially fertile area for future research.

In addition, our findings demonstrate that exogenous government policy–related variation in the factors that tend to tie workers to their jobs can be exploited to identify and gauge job lock. To our knowledge, only the present paper and Gruber and Madrian (1994) have examined the effects of policy-induced variation in job lock across those with and without alternative sources of coverage. Among studies that exploit variation in health conditions or other predictors of individual valuation of health benefits, none have made use of policy variation. However, the proscriptions against the long-term exclusion of coverage for pre-existing conditions introduced in the 1996 Health Insurance Portability and Accountability Act may provide an opportunity to improve on studies exploiting this latter identification strategy.

REFERENCES

Adams, Scott. 2004. "Employer-Provided Health Insurance and Job Change." *Contemporary Economic Policy,* Vol. 22, No. 3 (July), pp. 357–69.

Anderson, Patricia M. 1998. "The Effect of Employer-Provided Health Insurance on Job Mobility: Job-Lock or Job-Push?" Unpublished manuscript, Dartmouth College.

Bansak, Cynthia, and Steven Raphael. 2006. "Have Employment Relationships in the United States Become Less Stable?" *International Advances in Economic Research,* Vol. 12, No. 3 (August), pp. 342–57.

____. 2007. "The Effect of State Policy Design Features on Take Up and Crowd Out Rates for the State Children's Health Insurance Program." *Journal of Policy Analysis and Management,* Vol. 26, No. 1 (Winter), pp. 149–75.

Brunetti, Michael J., Kamran Nayeri, Carlos Dobkin, and Henry E. Brady. 2000. "Health Status, Health Insurance, and Worker Mobility: A Study of Job Lock in California." Unpublished manuscript, University of California, Berkeley.

Buchmueller, Thomas G., and Robert G. Valletta. 1996. "The Effects of Employer-Provided Health Insurance on Worker Mobility." *Industrial and Labor Relations Review,* Vol. 49, No. 3 (April), pp. 439–55.

Cutler, David M., and Jonathan Gruber. 1996. "Does Public Insurance Crowd Out Private Insurance?" *Quarterly Journal of Economics*, Vol. 111, No. 1 (February), pp. 381–460.

Gilleskie, Donna B., and Byron F. Lutz. 2002. "The Impact of Employer-Provided Health Insurance on Dynamic Employment Transitions." *Journal of Human Resources,* Vol. 37, No. 1 (Winter), pp. 129–62.

Gruber, Jonathan, and Brigitte C. Madrian. 1994. "Health Insurance and Job Mobility: The Effects of Public Policy on Job-Lock." *Industrial and Labor Relations Review,* Vol. 48, No. 1 (October), pp. 86–102.

____. 1997. "Employment Separation and Health Insurance Coverage." *Journal of Public Economics,* Vol. 66, No. 3 (December), pp. 349–82.

____. 2002. "Health Insurance, Labor Supply, and Job Mobility: A Critical Review of the Literature." National Bureau of Economic Research Working Paper #8817.

Holtz-Eakin, Douglas. 1993. "Health Insurance Provision and Labor Market Efficiency in the United States." National Bureau of Economic Research Working Paper #4388.

Kapur, Kanika. 1998. "The Impact of Health on Job Mobility: A Measure of Job Lock." *Industrial and Labor Relations Review,* Vol. 51, No. 2 (January), pp. 282–98.

Los Angeles Times. 1998. "Work and Careers on the Job: Fear of Losing Health-Care Benefits Has Many Workers Locked into Jobs." August 30, p. D5.

LoSasso, Anthony T., and Thomas C. Buchmueller. 2002. "The Effect of the State Children's Health Insurance Program on Health Insurance Coverage." National Bureau of Economic Research Working Paper #9405.

Madrian, Brigitte C. 1994. "Employment-Based Health Insurance and Job Mobility: Is There Evidence of Job-Lock?" *Quarterly Journal of Economics,* Vol. 109, No. 1 (February), pp. 27–54.

New York Times. 1991. "Health Benefits Found to Deter Switches in Jobs." September 26, p. A1.

Stroupe, Kevin T., Eleanor D. Kinney, and Thomas J. Kniesner. 2001. "Chronic Illness and Health Insurance–Related Job Lock." *Journal of Policy Analysis and Management,* Vol. 20, No. 3 (Summer), pp. 525–44.

U.S. House of Representatives, Committee on Ways and Means. 2004. *Green Book.* Washington, D.C.: U.S. Government Printing Office.