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6 Health Insurance Provision and Labor Market Efficiency in the United States and Germany

Douglas Holtz-Eakin

> Worker mobility is one of the most important values in an entrepreneurial society, where most jobs are created by small businesses. The present health care system is a big brake on that.
>
> President Bill Clinton

Health insurance and health care provision have claimed a prominent place on the policy agenda in the United States. Critics argue that the *status quo* has led to spiraling health care costs, an inequitable distribution of quality medical care, and a failure to provide care to many individuals. Further, there is a perception that the dominant form of providing health insurance in the United States—private provision as part of employee compensation—interferes with smooth functioning of the labor market. The statement (above) by President Bill Clinton[1] is characteristic of claims that individuals are locked into jobs because of their fear of large changes in their health insurance status if they change jobs. Clearly, to the extent that this is true, a U.S.-style system of providing this fundamental part of the social safety net interferes with the efficient matching of employers and employees. Other things equal, one would prefer to avoid such a labor market inefficiency when providing health insurance.

Do individuals forgo changing jobs on the basis of health insurance? In a recent *CBS/New York Times* survey, roughly 30 percent of respondents indicated that they had stayed in a job to retain their current health insurance cover-

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1. Quotation from Greider, W., P. J. O'Rourke, H. S. Thompson, and J. S. Wenner (1992), The Rolling Stone interview: Bill Clinton, *Rolling Stone*, 17 September, 44.

age.[2] In other suggestive evidence, a KPMG Peat Marwick survey cited in the *Wall Street Journal* indicated that job switching may be impeded by the absence of coverage for preexisting conditions, a situation affecting over two-thirds of employees.[3] In addition, researchers have begun to analyze the job lock issue. The results in Madrian (1992), the first careful econometric analysis of the issue, suggest the presence of job lock from health insurance in the United States.

In contrast to the U.S. system, virtually all German citizens are guaranteed health insurance as part of a privately operated but compulsory health insurance system. As a result, the insurance is apparently portable, and one would expect that it would have no impact on job arrangements. However, one feature of the German system is that individuals may pay different—perhaps very different—premiums for essentially the same coverage. Moreover, the cost of coverage will depend upon the insurance company, or sickness fund, chosen by each employer. Accordingly, health insurance may be portable, but the *price* is not. In some circumstances, however, individuals—not their employers—choose the sickness fund that provides for their health insurance, thereby providing portability of both coverage and price. *Ceteris paribus*, one would anticipate a greater propensity to move among those individuals whose insurance does not change in price across jobs. To the extent that this is an important issue, changes in the price of health insurance when switching jobs offer a potential impediment to the smooth functioning of the German labor market.

The basic goal of this paper is to assess the empirical magnitude of health insurance–related impediments to job mobility in the United States and Germany. It is important to stress at the outset that the impact on labor market efficiency is not the sole, or even the best, means by which to gauge the performance of a health insurance system. One might, for example, wish to assess the efficiency of the insurance market itself.[4] Still, a comparison of the U.S. and German experiences may provide input in assessing the relative strengths of the alternative systems.

To anticipate the basic results, I find little evidence that health insurance provision interferes with job mobility in either the United States or Germany. The outline for the remainder of the paper is as follows. Section 6.1 looks at health insurance and labor market mobility in the United States, focusing on the job lock hypothesis. In section 6.2, I examine the role of health insurance financing in labor market mobility in Germany. The final section contains a summary and assesses the implications of the results.

2. *New York Times* (1991), 26 September, 1. The question asked was: "Have you or anyone else in your household ever decided to stay in a job you wanted to leave mainly because you didn't want to lose health coverage?"

3. *Wall Street Journal* (1991), 31 December, A1.

4. Fuchs (1991) provides an overview of the issues.

6.1 Health Insurance and the Labor Market in the United States

As noted at the outset, policy toward health insurance is currently much debated in the United States. In part, this reflects the increasing share of national resources devoted to health care expenditures; the share of such spending in net national product rose from 4.8 percent to 13.7 percent between 1950 and 1990 (Aaron 1991, table 3-1). In addition, there is widespread concern that the large number of uninsured families leaves a critical part of the population, especially children, needlessly vulnerable to health problems. Finally, there is the notion that the tradition of packaging health insurance along with other job-related benefits may have detrimental effects on labor market efficiency.

Some suggestive evidence on the latter issue—the job lock conjecture—in the United States is displayed in table 6.1. As detailed below, the 1984 wave of the Panel Study of Income Dynamics (PSID) identifies individuals receiving employer-provided health insurance. The table shows the frequency with which those who do and do not receive such health insurance change employers in subsequent years. The job transition rates are for the one-year period from 1984 to 1985 and the three-year span between 1984 and 1987.[5] To the extent that job lock is a significant feature of labor market dynamics, transitions between employers should be relatively higher among the uninsured and relatively lower among those who receive health insurance. That is, in terms of the table, the proportion of the entries in the first row of each category, under "Job Change," should be greater than the proportion of the entries in the second row. Looking first at the transitions for married individuals, one finds exactly the predicted pattern for both one-year and three-year transitions.[6] Also reported below each transition category is a Chi-square test statistic. In each case, the transition rates of the insured are significantly different from those of uninsured. Turning to single individuals, one finds the same pattern: transition rates are lower for the insured. Once again, the differences are statistically different.

To be sure, transition tables focus on a single variable—insurance here—and thus ignore many aspects of the decision to change jobs. For this reason, they are not a powerful test of the job lock hypothesis. This section is devoted to developing a more complete assessment of the role of job lock. Before doing so, however, I begin with a brief review of private and public health insurance institutions in the United States. In the subsequent subsection, I discuss information available regarding health insurance in the PSID, the source of individual-level data for the United States in this study. Section 6.1.3 discusses

5. The sample consists of full-time employed individuals aged twenty-five to fifty-five. I follow the recommendations of Brown and Light (1992) in identifying job changes.

6. At each step in the analysis that follows, results for two-year transitions are quite similar to those for the three-year horizon. To conserve space, these are not reported.

Table 6.1 **Job Transition in the United States**

	One-Year Transition		Three-Year Transition	
	No Change	Job Change	No Change	Job Change
Married individuals				
No insurance	1,707	399	1,283	757
	(0.834)	(0.166)	(0.629)	(0.371)
Insurance	1,789	234	1,499	524
	(0.884)	(0.116)	(0.741)	(0.259)
χ^2	21.4**		59.1**	
Single individuals				
No insurance	500	152	367	258
	(0.767)	(0.233)	(0.563)	(0.437)
Insurance	429	96	341	184
	(0.817)	(0.183)	(0.650)	(0.350)
χ^2	4.41*		9.11**	

Notes: Numbers in parentheses = column entry ÷ (no job change + job change). χ^2 = Chi-square test statistic for the null hypothesis that those with and without insurance have the same transition rates.

*Significant at the 5 percent level.

**Significant at the 1 percent level.

the empirical facts regarding the uninsured as revealed by these data.[7] In the final section, I undertake some simple multivariate analyses designed to quantify the importance of health insurance in reducing labor market mobility.

6.1.1 The Provision of Health Insurance

For working-age individuals and their dependents, employer-provided private coverage is the dominant source of health insurance in the United States. Two out of every three Americans under the age of 65, constituting roughly 75 percent of employees, are covered by employer-provided private insurance (Aaron 1991, 54). The precise terms of this coverage, however, vary widely. In part due to the existence of alternative health care providers such as health maintenance organizations (HMOs) and in part because many employers support more than a single plan, the variance in actual coverage and its cost is enormous.

There is widespread government involvement in the provision of health insurance as well. Perhaps the most important policy is the exclusion of premiums for employer-provided insurance from taxable income under the U.S. individual income tax. The value of this exclusion is nearly $80 billion[8] and

7. As discussed below, the PSID does not ask directly about insurance coverage. Instead, individuals are classified as uninsured if they do not receive employer-provided health insurance.

8. Calculated from Congressional Budget Office (1992, 258). The taxation of employer-paid insurance would generate $230 billion in income tax revenues and $160 billion in payroll tax revenues over the period 1993–97. Converting the total ($390 billion) into an annual average yields the number in the text.

provides a clear incentive to add health insurance as a fringe benefit.[9] In addition, Medicare and Medicaid are large-scale programs that provide health insurance. Retired individuals who worked in employment covered by Social Security or railroad retirement are eligible for Medicare, which began in 1966. The program is divided into two parts: Part A covers the costs of hospitalization and some nursing home care; Part B covers physicians' charges. Part B is limited to covering 80 percent of allowable charges above a $100 deductible, so there is an incentive to purchase private, "medigap" insurance.[10]

Medicaid is a health insurance program for low-income individuals that covers roughly 9 percent of the population (Wolfe 1992). In practice, it consists of individual programs in each of the states, operating under general guidelines from the federal government. The federal government finances part of the cost by offering matching grants to those states that offer specified services to target populations. In particular, states are required to offer benefits to recipients of Aid to Families with Dependent Children (AFDC) and Supplemental Security Income (SSI).

6.1.2 Analyzing Health Insurance, Using the PSID

The empirical analysis presented below uses the Panel Study of Income Dynamics. The PSID offers many advantages, in particular a wealth of longitudinal data on labor market performance and a structure comparable to that of the Socioeconomic Panel of Germany (GSOEP) data used in the empirical analysis of Germany.[11] Unfortunately, there are drawbacks as well, the most important of which is the relative paucity of information on the health insurance status of individuals. Ideally, one would like to have annual information on the type, cost, and coverage from all sources of health insurance. Instead, even the relatively circumscribed information on health insurance coverage is limited to a single year.

In the 1984 wave of the PSID, individuals were asked the question: "Does your employer pay for any medical, surgical, or hospital insurance that covers any illness or injury that might happen to you when you are not at work?" For married couples, there is an identical question regarding the payment of health insurance by the spouse's employer. In what follows, those individuals who answered "yes" are classified as having employer-provided health insurance, and those who answered "no" will be referred to as "uninsured."[12] Clearly, this is a great departure from the ideal. In particular, some of those categorized as

9. See, for example, Sloan and Adamache (1986) or Hamermesh and Woodbury (1990) for a discussion of the relationship between tax policy and the provision of fringe benefits.

10. In 1990, Congress placed limits on the medigap insurers, especially limiting their ability to exclude those with preexisting conditions, mandating experience rating, and requiring a minimum ratio of benefit payments to premium income. See Rovner (1990).

11. For a description of each data source, see the appendix.

12. Those that answered either "don't know" or "not applicable" were eliminated from the sample.

uninsured may have purchased private insurance or obtained access in some other way. Further, there is no information regarding the extent or *cost* of coverage, especially the degree to which spouses are covered by any plan. In what follows, I will interpret the results, using the assumption that individuals are eligible for coverage under their spouse's plan, if present, but a degree of caution is clearly warranted. Lastly, the focus below is on private health insurance. To the extent that the absence of private insurance is offset by Medicare or Medicaid, the labor market behavior of individuals will be altered.[13]

6.1.3 Who Are the Uninsured?

Estimates indicate that there are thirty to forty million uninsured individuals in the United States, including a substantial fraction that are uninsured throughout each year (Wolfe 1992). In addition, there is concern that the fraction of the population that is uninsured has risen over the past decade.[14] As noted above, the PSID does not contain information regarding health insurance provision for multiple years. Thus, it is not possible to address many interesting questions regarding the dynamics of insurance, such as the extent to which uninsured status is transitory, the extent to which the rise in the number of uninsured reflects changes in family structure, and so forth. It is possible, however, to take a "snapshot" view of the uninsured population.

In doing so, I focus on employed individuals ages twenty-five to fifty-five, the same group for which I test the job lock hypothesis. Thus, this section is best viewed as providing a glimpse at the characteristics of the working uninsured. The basic facts are laid out in table 6.2. Part A focuses on married individuals. In the top table the upper-left entry indicates for 49 percent of the married individuals who did not have insurance, neither did their spouses. In contrast, the upper-right entry indicates that the remaining 51 percent of the uninsured were married to individuals whose employers provided health insurance. The second row of the matrix gives corresponding information for those with insurance. Notice that the probability of having a spouse with insurance is higher for the uninsured than for the insured. Put differently, the probability of having insurance is negatively correlated among spouses. The Chi-square statistic (16.4) indicates that this correlation is statistically significant at the 1 percent level. Thus, spousal insurance tends weakly to offset the lack of insurance for a married individual.

From here, I proceed in two routes. The data on the left investigate the relationship between the individual reporting employer-provided insurance (Indi-

13. In practice, two pieces of evidence suggest that the existence of government-provided insurance does not have a significant impact on the results. First, there is no significant difference in behavior between those belonging to the low-income subsample and those in the remainder of the PSID. Second, eliminating all individuals who report assistance from any income assistance program (Aid to Families with Dependent Children, Supplemental Security Income, unemployment insurance, or Medicaid) does not affect the basic nature of the results.

14. Most counts of the uninsured are based on the Current Population Survey, and the wording of the questions in this survey has changed over time. As a result, there is some ambiguity in interpreting changes in the number of uninsured (see Swartz 1989).

Table 6.2 Insurance Relationships: Characteristics of Insured and Uninsured Individuals, Married and Single

A. Married Individuals

	Spouse Uninsured	Spouse Insured	χ^2
Uninsured	0.494	0.506	16.4**
Insured	0.557	0.443	

	Individual			Individual or Spouse		
Characteristic	No Insurance	In-surance	χ^2	No Insurance	In-surance	χ^2
Gender						
Male	0.352	0.646	350**	0.496	0.500	0.0520
Female	0.648	0.354		0.504	0.500	
Race						
Nonwhite	0.257	0.251	0.209	0.297	0.240	13.2*
White	0.743	0.749		0.703	0.760	
Health						
Good	0.971	0.995	34.0**	0.957	0.991	49.6**
Poor	0.029	0.005		0.043	0.009	
Union member						
Nonunion	0.967	0.727	454**	0.961	0.810	134**
Union	0.033	0.273		0.039	0.190	

B. Single Individuals

Characteristic	No Insurance	Insurance	χ^2
Gender			
Male	0.333	0.404	6.33*
Female	0.667	0.596	
Race			
Nonwhite	0.661	0.410	74.3**
White	0.339	0.590	
Health			
Good	0.929	0.983	18.6**
Poor	0.071	0.017	
Union membership			
Nonunion	0.952	0.762	91.8**
Union	0.048	0.238	

Notes: Each entry is the number of individuals, expressed as a decimal fraction of the total (e.g., male plus female) in that category (e.g., no insurance). χ^2 denotes the Chi-square test statistic for independence of the rows and columns.

*Significant at the 5 percent level.

**Significant at the 1 percent level.

vidual") and the characteristics of such individuals. In contrast, in the data on the right, we undertake the same comparisons, using instead as the measure of insurance whether either the individual or the spouse (or both) has employer-provided insurance ("Individual or Spouse").

Consider first the results for those that report receiving employer-provided coverage. A glance down the column reveals (statistically) significant differences along three major dimensions. First, the uninsured are more likely to be female than are the insured. Second, the uninsured are more likely to report that they are in poor health (in 1984, 2.9 percent in poor health, compared with 0.5 percent of the insured). Third, and not very surprising, the insured are much more likely to be union workers than are the uninsured population. The latter suggests that large changes in the status of unions may be related to changes in the uninsured population in the United States. Interestingly, the uninsured have roughly the same racial composition as the insured.

When one looks at the data on the right, a slightly different picture emerges. Recall that an individual is classified as insured here if either the individual or the individual's spouse (or both) report receiving employer-provided medical insurance. Using this definition, the gender composition of the uninsured and the insured populations is essentially the same. In contrast, we now find differences along racial lines. The uninsured population has a greater fraction of nonwhite individuals, and the difference is statistically significant at the 1 percent level. The remaining two relationships are unchanged: the uninsured are more likely to be in poor health and to be nonunion workers.

Part B of table 6.2 undertakes a comparable analysis for single individuals. The same patterns emerge from the data. Just as with married individuals, the uninsured tend to be female, nonwhite, in poor health, and nonunion workers.

Of course, it is desirable to examine these relationships simultaneously rather than in a sequence of bivariate comparisons. The probit analysis in table 6.3 is designed for this purpose. The table shows the results of estimating a probit model in which the dependent variable is equal to one if the individual has employer-provided insurance, and zero otherwise.[15] It is important to note that these probit estimates are designed to be descriptive; there is no putative causal relationship. Instead, they serve to summarize the empirical relationship between uninsured status and a myriad of other (endogenously determined) variables.

There are several interesting results. First, the coefficients on age, education, and number of children are insignificant.[16] The last in particular indicates no differential propensity for those with children to be at risk of not having health insurance. Similarly, the lack of correlation with race persists in this analysis; neither of the coefficients for white and black individuals is individually significant. (Also, one cannot reject the null that they are equal.)

There are a wide variety of variables that do enter significantly. The dummy variables to control for occupation and industry capture statistical differences

15. A more expansive definition of coverage would include those whose spouses have employer-provided insurance. Estimating a probit using this definition yields results broadly similar to those in table 6.3. To the extent that they differ, the probability of coverage rises with age, declines for nonwhites, and is unrelated to sex and tenure.

16. One cannot reject the joint hypothesis that both age and age squared have coefficients equal to zero.

Table 6.3 Probit Analysis: Probability of Employer-Provided Insurance*

Variable	Coefficient (standard error)	Variable	Coefficient (standard error)
Age	−0.0387	Children	−0.00842
	(0.0275)		(0.0207)
Age squared × 10^{-3}	0.312	Wages × 10^{3}	0.00134
	(0.350)		(0.00132)
Education	0.0193	Assets × 10^{-6}	−0.764
	(0.0117)		(0.156)
White	0.157	Tenure	0.0204
	(0.162)		(0.00513)
Black	0.0645	Union member	0.565
	(0.165)		(0.145)
Female	−0.134	Union job	0.273
	(0.0565)		(0.133)
Poor health	−0.255	Spouse insured	0.0124
	(0.215)		(0.0556)
Married	−0.0988	Spouse wages × 10^{-3}	0.00259
	(0.0651)		(0.00198)
Average unemployment	−0.0160	Public sector	1.95
	(0.00578)		(0.147)
Professional	0.474	Retail sales	1.55
	(0.0807)		(0.129)
Sales	0.492	Real estate	1.90
	(0.0811)		(0.151)
Blue-collar	0.426	Business services	1.33
	(0.0846)		(0.158)
Agriculture, fisheries, forestry	0.942	Personal services	1.34
	(0.202)		(0.156)
Mining	2.51	Entertainment	1.34
	(0.325)		(0.285)
Construction	1.17	Professional services	1.78
	(0.156)		(0.123)
Manufacturing	2.20	Public administration	2.14
	(0.133)		(0.144)

*Based on 5,037 observations. The dependent variable is equal to one if the individual has employer-provided insurance, and zero otherwise.

in coverage across different sectors of the economy and job levels within firms. Similarly, the more detailed characteristics of the job also help to predict insurance coverage. The probability of medical insurance rises with wage earnings by the individual, tenure on the job, union membership, and whether the individual's job is covered by a collective bargaining agreement ("union job"). Thus, the presence of medical insurance reflects good jobs—that is, jobs with good wages, stability, and other benefits.

Individual characteristics matter as well. Females are less likely to be insured, although separate probits for married and single individuals (not reported) reveal that this finding stems from married females. (This also explains the negative but statistically insignificant effect on the variable for married individuals.) The probability of insurance falls with net assets, likely reflecting the ability of wealthy individuals to purchase insurance directly or forgo it altogether. An individual's work history also enters the likelihood of having insurance in a significant fashion; the probability of being insured falls when the average number of weeks of unemployment between 1981 and 1983 ("average unemployment") rises.

Finally, in contrast to the raw correlation in table 6.2, the probit analysis indicates no significant correlation between the presence of a spouse who has employer-provided insurance and the probability of having health insurance. A difficulty in interpreting such a reduced form, however, stems from the fact that some spouses may elect to decline insurance coverage. In such circumstances, it is not obvious how the individual will answer the survey question on insurance coverage.

6.1.4 Health Insurance and Job Mobility

Should employment-based health insurance reduce job mobility? At a first pass, the likely answer seems to be no. Health insurance is only part of the overall compensation package for a worker. Thus, one might anticipate that wages or other noninsurance aspects of the total compensation package would vary so as to offset differences in the cost of providing health insurance. Indeed, Gruber (1992) finds that changes in the costs of insuring workers for maternity benefits are reflected by almost identical offsets in wages. Thus, if workers differ in terms of the cost of health insurance, the result could still be a compensation package that matches the productivity of each worker.

Changes in health status, however, complicate this simple story. Consider an individual who experiences a significant decline in health. As part of his or her current group plan, say, he or she may be relatively inexpensive to insure. As a result of experience rating, however, this cost would rise if the individual moved to another firm, and the individual would be a less attractive candidate for other jobs unless he or she was willing to accept lower wages in the new job. Thus, in the current firm, the individual could receive health insurance of the same value and a larger wage income than in another firm. In this way, employment-based insurance may act as a "tax" on labor mobility by driving a wedge between the cost of insurance in a new firm versus the current employer. Notice, however, that even this effect may be short lived. To the extent that insurance companies raise premiums, even the current employer may move to a mix of greater health insurance and less wage income.

A similar scenario may follow from clauses precluding coverage for preexisting conditions. A 1987 survey indicated that 57 percent of employers had clauses limiting or excluding coverage for expenses stemming from preexisting conditions in their insurance arrangements. For smaller firms these are even more prevalent, with 64 percent of small employers (less than five hundred employees) having such clauses.[17] These features of the insurance market have led to many calls for reforms that move away from employer-based insurance (see, e.g., Mitchell 1990) on the grounds that the features trap workers in their current jobs. Still, firms have the option of paying more to cover preexisting conditions, so the presence of job lock behavior and its efficiency consequences is ultimately an empirical issue.

While there is a relatively large literature on the relationship between fringe benefits and job mobility (especially pensions), there is a paucity of studies examining the role of health insurance alone.[18] The major exception is Madrian (1992), who finds some evidence of job lock, using the 1987 National Medical Expenditure Survey. Madrian focuses on three comparisons to find evidence of job lock: (1) the job mobility of insured men whose spouses have health insurance coverage versus those whose spouses do not, (2) the behavior of men with large versus small families, and (3) the behavior of men with pregnant spouses versus those who are not expecting a child. In each case, mobility is lower in those situations where the current insurance coverage is more valuable. Indeed the empirical magnitudes are quite striking. Madrian estimates that voluntary mobility differentials due to job lock range from 25 percent (estimated using the first parameter) to 50 percent (using the third). The existence of mobility differentials of this magnitude suggests that an insurance system divorced from employment status (see, e.g., Mitchell 1990) would enhance efficiency.

To see the nature of the test for job lock, consider the following simplified model. Let the probability of changing employers be given by

$$p(\text{change}) = \phi(z) + \alpha_1 d_1 + \alpha_2 d_2 + \alpha_3 d_3 + \alpha_4 d_4 , \tag{1}$$

where $\phi(\cdot)$ captures non-insurance-related aspects of the job change decision. The variable d_1 is a dichotomous variable equal to one if the individual is the only person in the household to have insurance and equal to zero otherwise; d_2 is defined similarly and indicates that only the individual's spouse has insurance; d_3 equals one if both the individual and the spouse are insured, and d_4 indicates that neither has insurance. (It is assumed that if the spouse has insurance, the individual can be covered by it.) If the lack of portable insurance impedes job transitions, it should be apparent only when employment and in-

17. The survey of two thousand employers offering health insurance was conducted by Foster Higgens, an employee benefits consulting firm (see Cotton 1991).

18. Mitchell (1982, 1983) explores the link between fringe benefits and job mobility. There is also a substantial literature examining the degree to which pension plans (and their vesting rules) produce additional job attachment (see, e.g., Allen, Clark, and McDermed 1991). Gustman and Steinmeier (1990), however, find that jobs with pensions also contain a wage premium that dominates the financial effects of pension provision; they conclude that the wage premium is the source of lower propensities to leave these jobs. As was noted above, the probability that a job provides health insurance is also positively correlated with the wages received by the individual.

surance are tied. In terms of equation (1), this corresponds to having $d_1 = 1$. For all others, the access to insurance is independent of employment. For example, if $d_2 = 1$, then the individual is covered by the spouse's policy, which is unaffected by a change to a new employer. Or if $d_4 = 1$, the individual has no insurance to lose, and it is therefore not a factor when changing jobs.

It turns out to be useful to express this in a slightly different form:

$$\text{(2)} \qquad p(\text{change}) = \phi(z) + \beta_0 + \beta_1 \mathit{Self} + \beta_2 \mathit{Spouse} + \beta_3 \mathit{Both}\,,$$

where *Self* indicates that the individual has insurance, *Spouse* indicates that the spouse has insurance, and *Both* is the interaction (product) of these two variables. A bit of algebra reveals the correspondence between equations (1) and (2):

$$\text{(3)} \qquad \begin{aligned} p(\text{change}) = \phi(z) &+ (\beta_0 + \beta_1)d_1 + (\beta_0 + \beta_2)d_2 + \\ &(\beta_0 + \beta_1 + \beta_2 + \beta_3)d_3 + \beta_0 d_4\,. \end{aligned}$$

Consider now equation (1). As argued above, if individuals are locked into their current jobs by health insurance, it should be apparent only when $d_1 = 1$. Note that the other states are equivalent from the perspective of health insurance—the individual loses no insurance if he or she leaves a job. Hence, one would expect $\alpha_2 = \alpha_3 = \alpha_4$. Thus, the notion that employer-based insurance affects job transitions amounts to testing the null hypothesis that $\alpha_1 = \alpha_2 (= \alpha_3 = \alpha_4)$. This has several implications for the parameters in equation (3): (1) $\alpha_2 = \alpha_3$ implies that $\beta_1 + \beta_3 = 0$; (2) $\alpha_3 = \alpha_4$ implies that $\beta_1 + \beta_2 + \beta_3 = 0$, so that these together require that $\beta_2 = 0$; and (3) $\alpha_1 = \alpha_2$ implies that $\beta_1 = \beta_2$, so that β_1 must also be zero. Collecting results, this requires that β_3, the coefficient on the interaction variables (*Both*), be zero. Thus, this line of reasoning suggests the (not surprising) result that one should test whether all of the coefficients on the insurance variables in equation (2) are equal to zero. If they are, this is consistent with the notion that health insurance has no effect on transitions among jobs. Rejecting this null hypothesis, however, suggests the presence of job lock due to health insurance.

However, one could argue that the presence of an employer-provided insurance plan is really serving as an indicator of whether the individual has a "good job"; the probit analysis in table 6.3 leads directly to this notion. If so, all that such an exercise establishes is that people are less likely to leave good jobs than bad jobs. In terms of equation (1), the "good jobs" argument essentially says that $\phi(\cdot)$ does not control completely for attributes of the job that are correlated with the presence of insurance. It is likely, then, that the coefficients on d_1 and d_3 are contaminated by these job-related attributes. A similar argument may be put forward with regard to spouses; that is, the coefficients on d_2 and d_3 reflect unobserved attributes of the spouse. Thus, one may rewrite equation (1) as:

$$(1') \qquad p(\text{change}) = \phi(z) + (\alpha_1 + j)d_1 + (\alpha_2 + s)d_2 + (\alpha_3 + j + s)d_3 + \alpha_4 d_4,$$

or

$$(4) \qquad p(\text{change}) = \phi(z) + \gamma_1 d_1 + \gamma_2 d_2 + \gamma_3 d_3 + \gamma_4 d_4,$$

where j is the contamination due to job effects and s is the corresponding contamination due to spouse effects. Because of the presence of s and j, it is not possible to test the relevant hypothesis regarding the coefficients in equation (1). Indeed, one cannot even learn about health insurance effects by looking at $(\gamma_2 - \gamma_1)$, $(\gamma_3 - \gamma_1)$, or $(\gamma_4 - \gamma_1)$, because each contains either s or j. However, bringing equation (3) into play allows one to compare the differences of the differences, thereby eliminating the unobserved attributes and isolating the effect on job changes. Algebraically, $(\gamma_3 - \gamma_2) - (\gamma_1 - \gamma_4) = (\alpha_3 - \alpha_2) - (\alpha_1 - \alpha_4)$, which does not depend on s or j. Under the null hypothesis, this should equal zero. Returning to equation (3), it is straightforward to verify that $(\alpha_3 - \alpha_2) - (\alpha_1 - \alpha_4) = \beta_3$. Thus, testing the null hypothesis in the presence of job effects and spouse effects involves testing whether the coefficient on the interaction variable differs from zero. Intuitively, in the absence of health insurance effects on transitions, the impact of having an employer-provided plan should not depend on whether the worker can be covered by a spouse's plan.

Table 6.4 looks at the propensity to change employers for married individuals who have employer-provided medical insurance. Within each gender the

Table 6.4 **Job Transitions in the United States for Married Individuals with Employer-Provided Health Insurance**

	One-Year Transitions		Three-Year Transitions	
	No Change	Job Change	No Change	Job Change
Men				
Uninsured spouse	789	87	655	221
	(0.901)	(0.099)	(0.748)	(0.252)
Insured spouse	376	55	322	109
	(0.872)	(0.128)	(0.747)	(0.253)
χ^2	2.39		0.001	
Women				
Uninsured spouse	212	39	176	75
	(0.845)	(0.155)	(0.701)	(0.299)
Insured spouse	412	53	346	119
	(0.886)	(0.114)	(0.744)	(0.256)
χ^2	2.49		0.218	

Notes: Numbers in parentheses = column entry ÷ (no job change + job change). χ^2 = Chi-square test statistic for the null hypothesis that those with and without insurance have the same transition rates.

rows display the mobility rates for those whose spouses do not have insurance versus those whose spouses have employer-provided insurance. Consider the results for married males. For one-year transitions, one finds that 12.8 percent of males with an insured spouse undertook a job transition, and this exceeds the 9.9 percent rate for those whose spouses do not receive insurance. The greater propensity to change jobs is suggestive of the job lock phenomenon. As indicated by the Chi-square test statistic, however, the difference in behavior is not statistically significant.[19] The lack of an effect from spousal insurance is even clearer in the three-year transition matrix. Here the transition rates are virtually identical in each row.

Table 6.4 displays analogous transition data for married women. In both time spans, the propensity to change employers is greater, not smaller, for those whose spouses do not have insurance. Again, however, the differences are not statistically significant. Thus, the raw data reveal little linkage between insurance status and job mobility.

Intuitively, one would also expect job lock (if any) to become more important as insurance became more valuable to the individual. Using the information from the PSID, one may focus on several indicators of the value of insurance. First, the PSID contains measures of health status for 1984 and 1986 and of the change in health status between 1982 and 1984. Thus, for example, one might expect that health insurance would be more valuable to those with poor health in 1984.[20] In a probit equation for job mobility, then, this would lead one to anticipate that the interaction between poor health and the provision of health insurance would tend to decrease mobility; that is, the sign of the coefficient on such an interaction variable should be negative. In contrast, poor health should raise the value of access to medical insurance via one's spouse. I investigate these interactions in the analysis of the PSID.

The use of 1984 health status places the emphasis on contemporaneous relationships. Alternatively, it may be that individuals anticipate the need to address developing health conditions, and they value their current insurance more highly as a result. To gain a feel for this aspect of the data, I perform an analogous examination of the effects of interacting future health status (that reported for 1986) with the provision of health insurance in 1984. As before, one would anticipate that the coefficient on such an interaction variable should be (if anything) negative.

Of course, the discussion of preexisting conditions indicates that job lock may hinge on the *change* in health status as much as or more than on the state of the individual's health. The 1984 wave of the PSID also contains survey

19. The difference between the mobility rate of *uninsured* males whose spouses do and do not have insurance is 0.037. Thus, the differences-in-differences point estimate of the effect of insurance is negative: $-0.008 = 0.029 - 0.037$.

20. I focus on the "poor health" response in what follows. Attempts at first distinctions in health status provided no additional insights.

information on health status in 1984 versus 1982. I focus on those who report that their health is *worse* in 1984 than in 1982.[21] The logic of preexisting conditions indicates that the worsening of health condition should make health insurance more valuable and thus lower mobility. As before, I check the degree to which the data are consistent with this hypothesis by examining the coefficients on an interaction between the variable indicating provision of health insurance and that indicating a worsening of health status over the two years prior to 1984.

As a last check on the interaction between the value of insurance to the individual and job mobility, I focus on interactions with the age of the individual. Here, the basic notion is simple: as one ages, the expected cost of medical care rises, *ceteris paribus.* Thus, as before, the interaction between age and employer-provided health insurance should serve to reduce mobility.

Thus far, the discussion has focused on the health status of the individual. However, the health status of others living with (and covered by the policy for) the individual may have just as important an effect on the perceived value of the benefit. I use the number of children living in the family to proxy for the expected value of the insurance policy. If correct, one would expect the coefficient on such an interaction variable (number of children and employer-provided insurance) to be negative.

No empirical strategy is entirely without pitfalls, and the approach used herein is no exception. Thus far, for example, the discussion has treated health insurance status as exogenous. To the extent that this is not the case, two related problems arise. First, it is difficult to understand the notion of job lock when the health insurance package is (in part) self-inflicted. Second, the right-side variables in the probit equations below will be endogenous, and more-refined statistical techniques will be required to identify residual evidence of job lock. A conceptually similar issue arises in the literature on pensions. Allen, Clark, and McDermed (1991), for example, model the endogenous determination of pension status. In addition, when health insurance status is endogenous, there may be some gain to explicitly modeling the distribution of insurance among spouses. These extensions are beyond the scope of this paper.

Another caveat is that the absence of information on the benefits package for health insurance makes the interpretation of the dummy variables more tenuous. While the interaction effects discussed above are designed to reveal the differential value of insurance across individuals, they are unlikely to capture fully such variations in the net benefits of insurance coverage.

Last, two minor footnotes on legal institutions are in order. First, the Consolidated Omnibus Reconciliation Act of 1986 (COBRA) contains provisions (effective 1 July 1986) guaranteeing access to health insurance for up to eighteen months after separating from a job. (For a full discussion of COBRA, see

21. Again, attempts to make finer distinctions yielded little additional insight.

Flynn 1992.) Thus, while I examine job changes as late as 1987 below, those that occur after 1985 are subject to the caveat that COBRA may make job transitions easier. Similarly, even prior to COBRA many states had state-specific laws regarding access to employer medical insurance after leaving a job.[22]

6.1.5 Probit Analysis

The results of testing for health insurance–related job lock in the PSID are summarized in table 6.5 and table 6.6. Begin by looking at table 6.5, which reports the coefficient estimates for three variables indicating health insurance status. The first is a dichotomous variable equal to one for those that have employer-provided health insurance. The second is an analogously defined variable indicating that the spouse has health insurance. The final variable is the interaction between the first two variables and indicates that both the individual and the spouse have insurance. For single individuals, the last two clearly are not appropriate.

Part A of the table reports the coefficient estimates for one-year transitions, while Part B is devoted to three-year transitions. Within each part, two sets of estimates are reported for each sample. The "No Controls" estimates are obtained by estimating a probit equation for job transitions in which only the insurance variables (and a constant) appear on the right-hand side. The "Controls" estimates contain a rich set of noninsurance control variables and are discussed below. Examination of the "No Controls" estimates suggests two broad conclusions. First, for the probits analyzing married men and women, the coefficient on the interaction variable "Both insurance" is always on the wrong sign and statistically insignificant. Recall from the earlier discussion that this variable is the natural point of focus when one suspects that the specification does not fully capture attributes of jobs or spouses that are correlated with the provision of employer health insurance. The "No Controls" estimates in table 6.5 are an extreme case because no other controls are included in the equation. Second, from a job lock perspective, even the econometric performance of the "Own insurance" variable is somewhat uneven; only for married men is it uniformly negative and statistically significant.

The columns labeled "Controls" in table 6.5 contain the results for the insurance variables of fully specified mobility equations for one-year and three-year transitions, respectively.[23] Before discussing the variables related to the job lock hypothesis, consider the variables included in the probit to control for other aspects of job mobility. Each probit equation contains dummy variables for occupation and industry, reflecting differential conditions in these markets. In an attempt to control, albeit incompletely, for non–health insurance attributes of individuals' jobs, I include reported tenure on the job, dummy vari-

22. A statistical test of the importance of dummy variables for each state did not reject the null hypothesis of no significant differences across states.

23. Complete results are available from the author.

Table 6.5 **Probit Analysis of Job Transitions (standard errors in parentheses)**

A. One-Year Transitions

	Married Men		Married Women		Single Men		Single Women	
	No Controls (1)	Controls (2)	No Controls (3)	Controls (4)	No Controls (5)	Controls (6)	No Controls (7)	Controls (8)
Own insurance	−0.251**	−0.143	0.00519	−0.0341	−0.223	0.291	−0.175	−0.050
	(0.0901)	(0.134)	(0.117)	(0.141)	(0.133)	(0.236)	(0.111)	(0.217)
Spouse insurance	0.160	0.247	0.0618	−0.0568				
	(0.120)	(0.140)	(0.0854)	(0.129)				
Both insurance	−0.00209	0.0860	−0.270	−0.233				
	(0.155)	(0.164)	(0.150)	(0.163)				

B. Three-Year Transitions

	Married Men		Married Women		Single Men		Single Women	
	No Controls (1)	Controls (2)	No Controls (3)	Controls (4)	No Controls (5)	Controls (6)	No Controls (7)	Controls (8)
Own insurance	−0.175**	−0.0912	−0.0671	−0.156	−0.291*	0.304	−0.204	−0.0451
	(0.0748)	(0.110)	(0.0366)	(0.120)	(0.123)	(0.224)	(0.0960)	(0.180)
Spouse insurance	0.174	0.164	0.0344	−0.0692				
	(0.105)	(0.120)	(0.0268)	(0.108)				
Both insurance	−0.181	−0.0518	−0.0833	−0.104				
	(0.132)	(0.140)	(0.0458)	(0.136)				
Observations	2007	2007	2024	2024	420	420	738	738

*Significant at the 5 percent level.
**Significant at the 1 percent level.

ables for each of the following: whether the job is covered by a collective bargaining agreement, whether the job provides dental insurance, whether it provides life insurance, and whether there is a pension plan. To further refine the pension measures, two additional variables are included. The first identifies those who have been vested in the pension plan, while the remaining variable indicates those who participate in a defined contribution pension plan.

Also present in the equation are individual attributes such as age, education, race, an indicator variable for health-related work limitations, an indicator variable for poor health (in 1984), the number of children in the household, and union membership status.

Finally, the equations control for resources and prices associated with the mobility decision by including wage earnings in the current job and net assets of the individual. For married individuals, the wages of the spouse are included as well. (Descriptive statistics for key variables are shown in appendix table 6A.1.)

Consider now the variables related to medical insurance. For married individuals, as discussed earlier, the interaction variable is the focus of attention. For single individuals, such an interaction is not available. However, one may be able to discern evidence of job lock by looking at the interaction between insurance and health status or other variables. I return to these tests below.

Column 2 of part A in table 6.5 shows the estimates for married men when focusing on job changes over the one-year period 1984–85. The coefficient for having insurance is -0.14 but not statistically significant. As argued above, however, neither this coefficient nor the coefficient on spouse insurance (0.25) is the proper center of attention. Instead, the coefficient on the interaction variable is likely to be the most reliable indicator of job lock. The estimated coefficient is positive (0.09), which is suggestive of job lock but not statistically significant at conventional levels. Thus, in contrast to Madrian (1992), these data do not provide evidence of health insurance–related job lock.

The results for married women present a slightly different picture, as the coefficient on "Own insurance" is positive and imprecisely estimated. However, the interaction variable is of the "wrong" sign from the job lock perspective and, as in the case of married men, statistically insignificant.

The basic thrust of these results is reinforced by those reported in part B, which shows the estimates for three-year transitions. For both married men and married women, the point estimate of the interaction variable is negative and has a large standard error. In sum, using the presence of spouse insurance to test for job lock gives little support to the proposition. Thus, these estimates do not favor identifying health insurance as a major culprit in job market inefficiencies.

Recall from the discussion surrounding equation (4) that the coefficient on the own insurance variable may be contaminated by correlation with unobserved attributes of the job. If one could be confident that such unobserved heterogeneity was quantitatively small, then it would be possible to test for job lock by direct examination of the own insurance variable, a testing procedure that would permit examination of the results for single individuals (columns 6 and 8 of parts A and B as well. In this context, a comparison of the No Controls and Controls columns is relevant. The estimated coefficients for the insurance variables appear somewhat sensitive to the inclusion of controls for differences among individuals and jobs. This argues against focusing on the own insurance variable. Moreover, as noted earlier, the pattern of estimated coefficients for the own insurance variable in the No Controls columns does not support the job lock notion. This conclusion is amplified by the results in the Controls columns.

A simple indicator variable may not adequately capture differences across individuals in the value of insurance. Hence, as noted earlier, using interactions between the insurance variable (both insurance) and indicators of the value of insurance such as poor health, or worsened health, may provide better insight into the significance of job lock. Moreover, to the extent that these variables are significant in the mobility equations for single individuals, they permit one to detect job lock where the use of the spouse insurance interaction was not feasible.

Table 6.6 is devoted to summarizing the results of such an exercise. The table contains t-statistics to test the null hypothesis that the coefficient on the interaction between the insurance variable and the variable shown in each row is zero. Thus, for example, consider the entry in the first row of column 1 of the table. To generate the test statistic, I estimate a variant of the basic mobility probit, which also includes the interaction between the both insurance variable and the dummy variable for poor health in 1984. The test statistic, 1.52, indicates that one cannot reject the null hypothesis that the coefficient for the interaction is zero. This procedure is repeated for each of the variables shown in the table. (For single individuals, the own insurance variable is used for these interactions.)

Table 6.6 **t-test Statistics for Interactions with Insurance Variables**

	One-Year Transitions	Three-Year Transitions
Married men		
Poor health, 1984	1.52	1.18
Poor health, 1986	0.239	3.02**
Worse health, 1982–84	1.66	0.400
Children	0.007	−1.36
Age	0.585	0.709
Married women		
Poor health, 1984	−0.039	1.22
Poor health, 1986	1.17	1.43
Worse health, 1982–84	0.672	0.773
Children	−0.584	−0.406
Age	−1.36	−0.973
Single men		
Poor health, 1984	1.07	1.02
Poor health, 1986	1.83	1.34
Worse health, 1982–84	3.04**	3.17**
Children	−0.248	−0.552
Age	1.79	1.77
Single women		
Poor health, 1984	1.71	1.15
Poor health, 1986	−0.161	0.644
Worse health, 1982–84	2.12*	1.28
Children	−0.854	1.07
Age	−0.176	−0.704

Note: Test statistics for the null hypothesis that the interaction between the "Both Insurance" variable (for single individuals, "Own Insurance") and the row variable is zero.

*Significant at the 5 percent level.

**Significant at the 1 percent level.

What results emerge? There is little in the table to suggest an important or pervasive effect of job lock. Few of the interactions are statistically significant, and one that is—the interaction of worsening health status with insurance among single men—is of the wrong (positive) sign. In sum, using a variety of indicators of the value of health insurance to the individual does not provide evidence of job lock in the United States.

6.2 Health Insurance and the Labor Market in Germany

The German systems of health care provision and finance have attracted rising attention in the United States. As in the United States, the majority of care in Germany is provided by private sector doctors operating in private hospitals and financed by health insurance provided by private companies ("sickness funds," see below). Also in the United States, these private markets are subject to large-scale government intervention to ensure satisfactory health and budgetary outcomes. In 1987, however, health expenditures were only 8.1 percent of gross domestic product in West Germany, compared with 11.2 percent in the United States (Aaron 1991, 80, table 4-1). Moreover, mandatory insurance coverage for virtually all Germans precludes the possibility of large numbers of uninsured people. For these reasons, some analysts have pointed to the German system as a model for U.S. reforms.[24]

6.2.1 Health Insurance Provision in Germany

The German system of social health insurance[25] was introduced by Bismarck in 1883 and has gradually expanded to cover roughly 90 percent of the population.[26] The core of the financing system is provided by the roughly 1,150 private sickness funds, or insurance companies. Regional associations of sickness funds, in turn, bargain with regional associations of physicians to determine the rates charged for specific services.[27] Similarly, there are negotiations with each hospital for specific in-patient rates. All these negotiations are undertaken within the guidelines for rate increases established by the national committee (Concerted Action) set up in the 1977 health care reform to control the growth rate of health costs.

In some broad sense, then, the fund system is decentralized and self-governed by autonomous administrations. That is, it resembles a private system in that there are no explicit government agencies. German law requires, how-

24. In this context, it is somewhat ironic that the West German health care and finance system underwent major reforms in 1977, 1982, 1983, and 1989, in part to address dissatisfaction over the inability to contain costs.

25. This brief overview draws upon excellent surveys by Glaser (1991), GAO (1991), Henke (1990), and Reinhardt (1990).

26. The system dates from the Social Insurance Code of 15 June 1883 (Commission of the European Communities 1990, 50).

27. German physicians do not directly bill the sickness fund for services rendered. Instead, the regional association of physicians pays its members out of premium income collected from the sickness funds.

ever, that all persons with incomes below a cutoff (in 1989, DM 54,900, or roughly $27,300 in 1990 U.S. dollars) receive mandatory health insurance coverage. Those people with incomes above the cutoff may voluntarily join the mandatory system, purchase private health insurance, or remain uninsured.[28]

In large part, individuals receive insurance from the sickness fund chosen by their employer. (I will return to the exceptions below.) The retired are covered by the sickness fund of their former employer, and the unemployed receive insurance from the sickness fund of their previous employer. Self-employed individuals must enroll in one of the sickness funds. The health insurance premiums are financed from a variety of sources. The bulk of contributions take the form of a payroll tax rate, which is statutorily split equally between the employer and the employee. Government subsidies contribute toward the cost of covering the retired, the unemployed, and full-time students.

The payroll tax base consists mainly of wage and salary income.[29] Health insurance premiums, and thus payroll tax rates, are based on the average cost of insurance within each sickness fund. To calculate the rate, an insurance fund effectively divides the expected insurance costs by the total payroll tax base of its members. The result is a single payroll tax rate that is applied to the earnings of each individual (and his or her employer) in the fund. In this way, the German system embodies a form of "community rating" in which insurance rates are independent of the medical risks of individuals and their dependents.[30] Rates do depend, however, upon the sickness fund of the insured, with the result that West German rates ranged from 8 percent to 16 percent in 1990 (Schneider 1991).

Such a system provides clear incentives to migrate from high-cost insurance funds. Freedom of choice, however, is carefully circumscribed in the German insurance system. As noted above, employers choose the sickness fund that covers most of their employees. The most common sickness funds are the General Local Sickness Funds, or local funds organized on a regional basis. In addition, companies may organize their own establishment sickness funds, or company funds, to provide insurance to their employees. White-collar workers, however, also have the option of joining alternative private funds known as substitute funds, which are organized on a national basis.[31] In addition, there are several sickness funds for specific occupations (guild funds) and for miners, farmers, and mariners.[32] In the end, approximately 50 percent of individu-

28. Prior to the 1989 reforms, individuals whose incomes fluctuated first above, then below the income cutoff simply rejoined the mandatory system, an option that is no longer available.

29. Recently, some pension income has been included in the payroll tax base.

30. It is illegal to charge rates that discriminate by age, but in practice the system moves beyond even this restriction.

31. As a result of the 1989 reforms, the options of blue-collar workers are now comparable to those of white-collar workers.

32. The breakdown is as follows: 266 local sickness funds, 691 company sickness funds, 152 guild sickness funds, 19 agricultural sickness funds, 1 seamen's fund, 1 miners' fund, and 15 substitute sickness funds.

als may choose their own health insurance fund. Note that most of the variation across funds occurs in the cost of coverage, not benefits received (HCFR 1989, 94).

6.2.2 Health Insurance Information in the GSOEP

With the PSID as the standard for comparison, the German Socioeconomic Panel (GSOEP) provides relatively good information regarding the insurance status of each individual. Each year, information is collected regarding whether the individual has insurance and, if so, from which type of sickness fund. In addition, there is information about whether the individual is a voluntary or compulsory member of the sickness fund. Like the PSID, however, the GSOEP does not contain information on the health insurance premium paid by each individual.[33]

6.2.3 Health Insurance and Job Turnover

Although the German system is designed to provide universal coverage, its features potentially generate a job lock phenomenon. In this regard, the key aspect is the degree to which the price of insurance is portable across jobs. Several cases are straightforward. Individuals with private health insurance, for example, have insurance that is portable in both access and price. Similarly, members of guild or substitute sickness funds (each of whom is, by definition, a voluntary member) will be equally unencumbered by their insurance status.

For the remainder, the possibility of a nonportable price arises. Members of company funds—voluntary or compulsory—face a change in the cost of health insurance if they change employers, and the empirical averages indicate that it will be a higher cost. Schulenburg (1989) reports that the company funds have the lowest average payroll tax rate among types of sickness funds.[34] Similarly, both voluntary and compulsory members of local sickness funds may have to change funds if the employment switch takes them outside of the current area or if their new employer chooses not to use the local sickness fund.

To deal with the ambiguity, I create two variables identifying candidates for job lock on the basis of health insurance. The first, "Insurance Lock 1," consists of company fund members, local fund members, and compulsory members of the "other" category.[35] If all employment changes involve interregional moves, this variable appropriately identifies candidates for job lock. The second variable, "Insurance Lock 2," excludes members of local funds and is appropriate

33. There have been attempts to impute premiums to individuals, but the imputation scheme does not reflect differences in types of sickness funds. Experimentation with these imputed payroll tax rates did not prove fruitful, and the analysis presented below does not rely on these measures.

34. See table 6.6. Schulenburg reports that in 1988 the average payroll tax rate required for health insurance was highest (13.5 percent) in local sickness funds and lowest (11.5 percent) in company funds.

35. The latter assignment was done on the basis of testing whether the coefficient on such a variable was different from that on the health lock variable. I could not reject this hypothesis.

only for intraregional moves. While only a conjecture, it would seem likely that the large number of local funds and relatively small land area would combine to make the first measure more appropriate on the whole.[36]

As with the analysis of the PSID, I begin with a simple look at the data in the GSOEP. Tables 6.7 and 6.8 contain transition data computed for those with and those without "Insurance Lock" status. Consider table 6.7, which shows the transition rates of for individuals with and without the Insurance Lock 1 status. The table first shows a comparison of job mobility rates for married individuals. For comparability with the analysis of the U.S. data, I restrict the sample to individuals ages twenty-five to fifty-five who are full-time employees. The combination of excluding those in the very early stages of their labor market experience and the general nature of the German labor market contributes to the low overall rate of job transition that characterizes both tables.

Looking at the transition rates in table 6.7, one finds that they are in accord with the simple hypothesis for both one-year transitions (between 1984 and 1985) and three-year transitions (between 1984 and 1987). Only the former differences, however, are statistically significant. For single individuals, job transition rates for the insurance locked are lower over both the one-year and the three-year horizon; indeed, they are less than one-half the rates for the control group. Moreover, as shown by the Chi-square statistics, these differences are significant.

Table 6.8 repeats the analysis using Insurance Lock 2, the more circumscribed definition, as the indicator of insurance lock-in. Here the results are qualitatively similar—transition rates are uniformly lower for married and single individuals in the insurance lock category—but the details of statistical significance differ. As with the transition data from the PSID, such a simple test is hardly conclusive, so I turn now to a multivariate analysis intended to shed additional light on the issue.

6.2.4 Probit Analysis

The results of the probit analysis for job lock in the GSOEP, presented in table 6.9, are organized in a fashion parallel to those for the United States.[37] As before, I begin by analyzing transition equations that include only the insurance lock variables. Because of the low overall rate of job transitions, the use of such a parsimonious specification avoids the pitfall of overfitting the equations. Further, in a fashion analogous to the discussion earlier, comparison of these estimates with those from a richer specification sheds light on the degree to which the insurance variables are correlated with unobserved heterogeneity in

36. The public use sample of the GSOEP does not contain geographic identifiers, thus precluding direct examination of this issue.

37. The sample size for married women who are full-time employed was too small to obtain satisfactory estimates, so I restrict my attention to married men, single men, and single women. Full results are available upon request.

Table 6.7 **Job Transitions for Germany: Insurance Lock 1**

	One-Year Transitions		Three-Year Transitions	
	No Change	Job Change	No Change	Job Change
Married individuals				
Insurance Lock 1	703	9	670	42
	(0.987)	(0.013)	(0.941)	(0.059)
No lock	717	20	692	45
	(0.973)	(0.027)	(0.939)	(0.061)
χ^2	3.88*		0.027	
Single individuals				
Insurance Lock 1	971	24	945	50
	(0.976)	(0.024)	(0.950)	(0.050)
No lock	331	21	309	43
	(0.940)	(0.060)	(0.878)	(0.122)
χ^2	10.2**		20.9**	

Notes: Numbers in parentheses = column entry ÷ (no job change + job change). χ^2 = Chi-square test statistic for the null hypothesis that those with and without insurance have the same transition rates.

*Significant at the 5 percent level.

**Significant at the 1 percent level.

Table 6.8 **Job Transitions for Germany: Insurance Lock 2**

	One-Year Transitions		Three-Year Transitions	
	No Change	Job Change	No Change	Job Change
Married individuals				
Insurance Lock 2	212	2	209	5
	(0.991)	(0.009)	(0.977)	(0.023)
No lock	1,208	27	1,153	82
	(0.978)	(0.022)	(0.934)	(0.066)
χ^2	1.46*		5.98*	
Single individuals				
Insurance Lock 2	230	4	227	7
	(0.983)	(0.017)	(0.870)	(0.030)
No lock	1,072	41	1,027	86
	(0.963)	(0.037)	(0.923)	(0.077)
χ^2	2.33		6.75**	

Notes: Numbers in parentheses = column entry ÷ (no job change + job change). χ^2 = Chi-square test statistic for the null hypothesis that those with and without insurance have the same transition rates.

*Significant at the 5 percent level.

**Significant at the 1 percent level.

Table 6.9 **Probit Analysis of Job Transitions (standard errors in parentheses)**

	Married Men		Single Men		Single Women	
A. One-Year Transitions	No Controls	Controls	No Controls	Controls	No Controls	Controls
Insurance Lock 1	−0.261	−0.501*	−0.446**	−0.381	−0.459	−0.583
	(0.168)	(0.227)	(0.140)	(0.262)	(0.244)	(0.530)
Insurance Lock 2	−0.342	−0.348	−0.373	−0.107	−0.242	0.317
	(0.279)	(0.315)	(0.246)	(0.285)	(0.419)	(0.538)
B. Three-Year Transitions	Married Men		Single Men		Single Women	
	No Controls	Controls	No Controls	Controls	No Controls	Controls
Insurance Lock 1	0.179	−0.153	−0.447**	−0.409*	−0.691**	−0.952*
	(0.114)	(0.153)	(0.138)	(0.207)	(0.199)	(0.396)
Insurance Lock 2	−0.477*	−0.509*	−0.542**	−0.384	−0.273	0.102
	(0.199)	(0.213)	(0.204)	(0.226)	(0.329)	(0.377)
Observations	1,173	1,173	879	879	453	453

*Significant at the 5 percent level.
**Significant at the 1 percent level.

job attributes for the individuals. The results for separate probits using each insurance lock measure are shown in the No Controls columns of table 6.9.

With a single exception—three-year transitions for married men—each of the estimated coefficients is negative, which is consistent with the job lock notion. In terms of statistical significance, the effect appears to be centered among the single individuals, with the more expansive definition of insurance lock performing the best.

As in the case of the PSID, it is useful to see whether this pattern survives in the context of a multivariate analysis. I begin by noting the noninsurance variables entered in the probit equation to control for other aspects of job changes. Each probit in the Controls columns includes controls for the type of employment (blue-collar, civil servant).[38] The equations also include individual attributes such as age, indicator variables for alternative educational backgrounds, an indicator of health-related work limitations, an indicator of poor health (in 1984), and the number of children in the household.

Finally, the equations control for resources and prices associated with the mobility decision by including wage earnings in the current job, household income (for married men) and the capital income (dividends and interest) of the individual.[39] Descriptive statistics for key variables are shown in the appendix table 6A.2.

38. Experiments with including industry dummy variables yielded very large standard errors but had only a small effect on the point estimates for the insurance lock variables.

39. In contrast to the PSID, these equations do not have a control for job tenure (which is not available) or pension characteristics.

Returning to table 6.9, part A looks at one-year job transitions between 1984 and 1985. For married men, the coefficient on the Insurance Lock 1 variable is negative and exceeds its standard error by roughly 1.9. Thus, especially if one employs a one-tailed test for a negative value, there is suggestive statistical evidence that job mobility is lower among those individuals who belong to insurance funds for which the price of insurance is not portable. The result for Insurance Lock 2 is qualitatively similar but even less precisely estimated. The pattern is reversed for three-year transitions among married men; both coefficients are negative, but that for Insurance Lock 2 is weakly significant.

Recall as well from table 6.7 that the results for one-year transitions were somewhat stronger for single individuals than for married men. In the probit analysis, however, this is not the case. Looking at the estimated coefficients for single men and women in part A, one finds that they are typically of the anticipated sign but are not statistically significant at conventional levels in either one- or two-tailed tests.

Are these results special to one-year transitions? Consider the estimates for three-year transitions between 1984 and 1987. Again, there is no strong pattern of reduced mobility for those having insurance lock status. For both single men and single women, the coefficient on Insurance Lock 1 is negative but is significant only at the 5 percent level. Moreover, the results from using the Insurance Lock 2 variable instead are even weaker.

Taken as a whole, the parameter estimates in table 6.9 raise the possibility that the low rate of job mobility is reduced further by the institutions of the German health insurance system. At the same time, the statistical link is not sufficiently firm to warrant a strong position on the basis on this evidence alone.[40] Rather, the picture that emerges is one that does not support any conjecture of widespread labor market interference as a result of the health insurance system in Germany.

6.3 Summary

The potential for employer-provided insurance to interfere with the smooth working of the labor market has attracted considerable attention in the United States, and some analysts have pointed to Germany as a model for a system that avoids impediments to labor market mobility. As in the United States, however, the provision and cost of health insurance in Germany are in part determined by individuals' employers. Although it has not attracted comparable attention, the German system also generates the potential for insurance-related job lock.

40. I experimented with interactions of the health insurance lock variable and age, number of children, and health status. The anecdotal evidence suggests that the benefits package does not differ widely across sickness funds. If so, one would not expect these variables—which proxy for differences in the value of benefits across individuals—to be significant. They were not.

To date, much of discussion of job lock has been restricted to the use of anecdotal evidence. As a step toward filling the research void, this paper has been devoted to gauging the empirical magnitude of the job lock phenomenon. On the whole, these initial results suggest no evidence of job lock in either country. For the United States, analysis of the PSID suggests little in the way of correlation between insurance variables and the probability of changing employers. Some suggestive correlations are present when the health insurance variables are analyzed in isolation. However, in the presence of a rich set of noninsurance variables to control for other aspects of the incentives to change employers, their apparent importance disappears. This suggests that access to richer data for each individual and employer may explain the apparent difference between these findings and those in, for example, Madrian (1992).

The results of analyzing the GSOEP data for West Germany have the same flavor. When viewed in isolation, membership in a sickness fund for which the price of insurance is not portable across jobs is correlated with lower mobility. When analyzed simultaneously with a larger set of socioeconomic variables, however, the link becomes more tenuous. A difficulty unique to this analysis is the low overall rate of mobility among employers reported in the GSOEP.

From a slightly different perspective, the results of the empirical analysis in each country suggest a very important result. The health insurance systems in these countries should not be judged by their secondary effects on labor mobility, as these effects are small at best. Instead, they should be judged by their primary effects: access to health care and the efficiency of the provision of health insurance.

Appendix
Data Sources

The empirical analyses use data from two longitudinal data sets: for the United States, the Panel Study of Income Dynamics (PSID); for Germany, the Public Use Version of the Socioeconomic Panel of Germany (GSOEP). Since 1968, the PSID has interviewed annually a representative sample of some five thousand families. At least one member of each family was either part of the original families interviewed in 1968 or born to a member of one of these families. (See Survey Research Center 1984 for a complete discussion.)

The GSOEP is a more recent longitudinal data set developed at the Universities of Frankfurt and Mannheim in cooperation with the Deutsches Institut für Wirtschaftsforschung, Berlin (DIW) and initially financed by the German National Science Foundation. In 1990 the DIW assumed control of the panel with funding through 1995 from the Bund-Länder-Kommission für Forschungsförderung. The panel started in the spring of 1984. It comprises about six thousand families. Nine yearly waves have been conducted (1984–92), and six waves (1984–89) are available, providing information on calendar year 1983 through 1988. (In 1990 the GSOEP was expanded to include a representative sample of East Germans.) The data are representative of the German population, including "guest workers." Wagner, Burkhauser. and Behringer (1993) contains a detailed discussion of these data.

Table 6.A.1 Descriptive Statistics for Key Variables: PSID

	Married Men	Married Women	Single Men	Single Women
Job change 1984–85	0.1266	0.1512	0.2548	0.1789
	(0.3326)	(0.3583)	(0.4363)	(0.3835)
Job change 1984–87	0.2795	0.3463	0.4381	0.3726
	(0.4489)	(0.4759)	(0.4967)	(0.4838)
Own insurance	0.6457	0.3498	0.5000	0.4173
	(0.4784)	(0.4770)	(0.5006)	(0.4935)
Spouse insurance	0.3189	0.6275		
	(0.4662)	(0.4836)		
Both insurance	0.2113	0.2263		
	(0.4083)	(0.4185)		
Age	37.64	37.18	34.76	38.00
	(8.301)	(8.565)	(7.566)	(9.155)
Education (years)	12.66	12.60	12.51	12.04
	(2.626)	(2.210)	(2.702)	(2.350)
White	0.7514	0.7451	0.5429	0.4052
	(0.4323)	(0.4359)	(0.4988)	(0.4913)
Black	0.2327	0.2288	0.4333	0.5867
	(0.4227)	(0.4201)	(0.4961)	(0.4928)
Children	1.593	1.501	0.3643	1.098
	(1.214)	(1.205)	(0.8892)	(1.285)
Own wages. 1984	21.707	7.920	15.870	9.437
	(17.632)	(8.391)	(14.115)	(8.790)
Spouse wages. 1984	7.390	21.647		
	(8.207)	(20.986)		
Tenure	10.70	7.563	7.833	8.696
	(7.995)	(5.843)	(7.027)	(7.283)
Union member	0.2148	0.08646	0.1429	0.1179
	(0.4108)	(0.2811)	(0.3503)	(0.3227)
Poor health. 1984	0.01545	0.01976	0.02619	0.05556
	(0.1234)	(0.1392)	(0.1599)	(0.2292)
Poor health. 1986	0.02691	0.02223	0.02857	0.05149
	(0.1619)	(0.1475)	(0.1668)	(0.2212)
Worse health. 1982–84	0.09866	0.9289	0.1214	0.1626
	(0.2983)	(0.2903)	(0.3270)	(0.3693)

Table 6A.2 Descriptive Statistics for Key Variables: GSOEP

	Married Men	Married Women	Single Men	Single Women
Job change, 1984–85	0.02090	0.01103	0.03600	0.02838
	(0.1470)	(0.1046)	(0.1864)	(0.1663)
Job change, 1984–87	0.06372	0.04412	0.07537	0.05677
	(0.2444)	(0.2057)	(0.2641)	(0.2317)
Local fund, 1984	0.3339	0.3897	0.5928	0.5109
	(0.4718)	(0.4886)	(0.4916)	(0.5004)
Company fund, 1984	0.1453	0.06985	0.1766	0.1245
	(0.3525)	(0.2554)	(0.3816)	(0.3305)
Guild fund, 1984	0.06627	0.02574	0.04387	0.01965
	(0.2489)	(0.1586)	(0.2049)	(0.1390)
Substitute fund, 1984	0.2702	0.4044	0.08661	0.2795
	(0.4442)	(0.4917)	(0.2814)	(0.4492)
Other fund, 1984	0.04344	0.02214	0.02931	0.006550
	(0.2039)	(0.1474)	(0.1688)	(0.08076)
Private insurance, 1984	0.1325	0.09927	0.06637	0.06550
	(0.3392)	(0.2996)	(0.2491)	(0.2479)
No insurance, 1984	0.01956	0.003690	0.01237	0.002183
	(0.1385)	(0.06075)	(0.1106)	(0.04673)
Age	40.97	38.77	38.93	38.16
	(8.351)	(9.145)	(8.429)	(8.781)
Blue-collar	0.4435	0.2500	0.7840	0.5480
	(0.4970)	(0.4338)	(0.4117)	(0.4982)
Civil servant	0.1487	0.06985	0.05287	0.03712
	(0.3559)	(0.2554)	(0.2239)	(0.1893)
Chronic illness	0.2574	0.2427	0.1856	0.2576
	(0.4374)	(0.4295)	(0.3890)	(0.4378)
Health limitation	0.05438	0.09191	0.06412	0.09170
	(0.2269)	(0.2894)	(0.2451)	(0.2889)
Wages	44,998	32,098	36.140	28,231
	(19,814)	(28,577)	(22,317)	(11,355)
Capital income	306.8	301.9	190.2	131.5
	(2052)	(1301)	(2099)	(1581)
Children	1.037	0.5441	1.051	0.7249
	(0.9518)	(0.8138)	(1.239)	(1.022)

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