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Cancer Survivorship, Health Insurance, and Employment Transitions among Older Workers

This study examined the effect of job-related health insurance on employment transitions (labor force exits, reductions in hours, and job changes) of older working cancer survivors. Using multivariate models, we compared longitudinal data for the period 1997–2002 from the Penn State Cancer Survivor Study to similar data for workers with no cancer history in the Health and Retirement Study, who were also ages 55 to 64 at follow-up. The interaction of cancer survivorship with health insurance at diagnosis was negative and significant in predicting labor force exits, job changes, and transitions to part-time employment for both genders. The differential effect of job-related health insurance on the labor market dynamics of cancer survivors represents an additional component of the economic and psychosocial burden of cancer on survivors.

Job-related health insurance is particularly valuable to cancer survivors. Cancer survivors, like other people with expensive or potentially expensive chronic illnesses, are limited in their ability to purchase individual health insurance in the nongroup market. The regulations governing the nongroup market vary by state, but survivors are likely to encounter limited coverage of cancer-related expenses because of pre-existing condition exclusions, high premiums, and refusals by some insurers to sell to them at all (Pollitz, Richard, and Thomas 2001; Hewitt, Greenfield, and Stovall 2005). A March 2007 story in *The New York Times* described the health insurance predicament of one survivor, who was diagnosed with breast cancer after leaving a job with health insurance benefits to become a self-employed real estate agent. Although her cancer treatment was successfully completed, the only individual policy that she was offered cost more than $27,000 per year (Pear 2007). Even if cancer survivors are able to buy individual insurance, they are likely to find that the coverage is not as comprehensive as that offered by most employers, a particular hardship for survivors with ongoing medical problems resulting from their cancer and its treatment.

Under these circumstances, one would expect cancer survivors with job-related health insurance to give considerable weight to maintaining access to group insurance when contemplating changes in employment. Because part-time workers are often ineligible for health insurance benefits available to full-

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time workers (Kaiser Family Foundation and Health Education Research Trust 2007), even cutting back from full-time to part-time work without changing employers may threaten continued coverage. Some unlucky survivors may find themselves in a Catch-22 situation where their ability to work has been compromised by their cancer and treatment, yet because of health insurance considerations they are unable to adapt by reducing hours, changing jobs, or exiting the labor force.

Two federal laws, the Consolidated Omnibus Budget Reconciliation Act of 1985 (COBRA) and the Health Insurance Portability and Accountability Act of 1996 (HIPAA), offer limited protection to workers who leave jobs with health insurance benefits (Hewitt, Greenfield, and Stovall 2005; Hoffman 2005). COBRA requires employers with 20 or more employees to extend job-related coverage for 18 months after an employee or dependent would otherwise leave the plan, but only by charging the total premium (employer *and* employee shares) with a 2% add-on for administrative costs. Many states extend this requirement to smaller employers (Kaiser Family Foundation 2007).

HIPAA makes coverage more seamless for workers who move from one employer plan to another, by prohibiting pre-existing condition exclusions for workers with prior group coverage (within the preceding 63 days) and by prohibiting the exclusion of individuals from group plans because of health status. HIPAA also guarantees conversion to individual insurance for workers who leave job-related health insurance, but only after exercising and exhausting COBRA. Because these federal laws guarantee access to health insurance without regulating premiums or guaranteeing affordability, cancer survivors who leave jobs with health insurance still run the risk of finding themselves with no other source of affordable coverage.

Economists have demonstrated that health insurance considerations affect labor supply and job choices in a variety of contexts (Gruber and Madrian 2004). Empirical studies have shown that retiree health insurance encourages retirement (Gruber and Madrian 1996; Rogowski and Karoly 2000; Blau and Gilleskie 2001), that women are more likely to work in jobs with health insurance when their husbands do not have insurance benefits (Buchmueller and Valletta 1999), and—somewhat less definitively–that job mobility in the United States is reduced by the reluctance of workers to leave jobs with health insurance benefits (a phenomenon commonly called "job lock"). Identifying and measuring the effects of insurance on labor market dynamics is an empirical challenge because demonstrating that workers with health insurance are more attached to their jobs than workers without health insurance is not sufficient to demonstrate a causal connection to insurance. Other unmeasured differences in the quality of jobs, or the quality and professionalism of workers in jobs with better fringe benefits, also could explain the greater attachment of insured workers to their jobs.

One generally accepted solution is to use a difference-in-differences analysis to identify and measure the specific effects of health insurance. This test for the effect of insurance on employment transitions involves showing that the difference in transition rates between workers with and without health insurance is larger in a subgroup where health insurance is particularly valuable, compared to another subgroup where it is less valuable. Most often, this test has been formulated by comparing differences between policyholders who do and do not have access to employer health insurance through a spouse (Madrian 1994; Holtz-Eakin 1994; Buchmueller and Valletta 1996; Berger, Black, and Scott 2004; Adams 2004; Bradley et al. 2006). However, subgroups with differing valuations of job-related insurance also have been constructed on the basis of family size (Madrian 1994; Adams 2004), pregnancy (Madrian 1994), and health status (Kapur 1998; Stroupe, Kinney, and Kniesner 2001; Pelkowski and Berger 2003; Berger, Black, and Scott 2004; Rashad and Sarpong 2006). As an alternative to employing a difference-in-differences analysis in identifying job lock, studies also have examined the effects of continuation-of-coverage laws on job mobility (Madrian 1994; Gruber and Madrian 1996) or considered variation in the predicted probability of obtaining coverage on a new job (Cooper and Monheit 1993).

Studies aimed at determining whether workers with more health problems or a greater demand for medical care are "stuck" in jobs because of health insurance mostly have found (Stroupe, Kinney, and Kniesner 2001; Pelkowski and Berger 2003; Rashad and Sarpong 2006), but not always (Kapur 1998), that less healthy workers are more likely to remain in jobs with health insurance. In a recent study that examined the effects of employment-contingent health insurance on the labor supply of married breast cancer survivors at six, 12, and 18 months after diagnosis, Bradley et al. (2006) found that reductions in the labor supply shortly after a cancer diagnosis were larger for working women insured at diagnosis by their spouse's job-related insurance compared to women with their own job-related health insurance.

The average five-year survival rate for cancer has improved to 65% (Jemal et al. 2004), and there are now 10.8 million cancer survivors in the United States (Office of Cancer Survivorship 2008). These trends have given rise to a new conception of cancer as a chronic illness, encouraging a new emphasis on the quality of life of patients who survive cancer (Centers for Disease Control and Prevention and Lance Armstrong Foundation 2004; The President's Cancer Panel 2004; Hewitt, Greenfield, and Stovall 2005). Both the Institute of Medicine's Committee on Cancer Survivorship (Hewitt, Greenfield, and Stovall 2005) and the President's Cancer Panel (2004) have expressed concern that health insurance issues may limit the employment choices and opportunities available to cancer survivors. In the telephone interviews conducted for our study, 27% of insured survivors reported that the fear of losing their insurance had kept them from looking for another job.

In this article, we use a difference-in-differences approach to measure the effect of health insurance on employment transitions for cancer survivors. Comparing workers with and without job-related health insurance, we test the hypothesis that differences in transition rates are greater in a sample of cancer survivors than in a comparison group without cancer. Our analyses compare longitudinal data from the Penn State Cancer Survivor Study (PSCSS)—for survivors who were working when diagnosed from 1997 through 1999 and who were 55 to 64 years old in 2002—to longitudinal data for a noncancer sample of the same ages from the Health and Retirement Study (HRS), an ongoing national cohort study of Americans born from 1931 to 1947.

As already noted, the term "job lock" is used to describe the reluctance of workers to move from one job to another out of a desire to retain health insurance from their current job. In this study, we go beyond job lock to consider the broader effects of job-related health insurance on three types of employment transitions: changing jobs (job lock), reducing hours from full time to part time, and quitting work altogether. Depending on the alternatives available to a particular worker, making any of these changes could result in losing employer-sponsored health insurance, a type of insurance generally preferable to individually purchased insurance because of favorable tax treatment, lower premiums, more comprehensive coverage, and absence of individual risk rating. By making change more costly for some workers, we expect the presence of employer-sponsored insurance to be associated with lower average rates of change in all three of these domains. Furthermore, because the loss of employer-sponsored insurance is particularly costly for cancer survivors, we expect the average rate of change to be even lower among cancer survivors with employer-sponsored insurance.

Our study builds in several ways on previous work by Bradley and colleagues (2006), which demonstrated insurance-related inertia in the employment choices of breast cancer survivors. Our sample includes survivors of both genders with all types of cancer, except for superficial skin cancers. While the Bradley study focused on insurance-related differences *within* a sample of cancer survivors, we implement a difference-in-differences analysis that compares insurance-related differences in our cancer sample to insurance-related differences in a noncancer sample. We examine employment changes from diagnosis to a later point in survivorship, an average of nearly four years post-diagnosis, and consider job lock in addition to changes in employment status and hours. We also control for two additional job characteristics, pensions and job tenure, that could otherwise inflate the apparent attachment of cancer survivors to jobs with health insurance.

Data

Cancer Sample

The Penn State Cancer Survivor Study was a longitudinal survey that tracked the employment of a cohort of cancer survivors diagnosed from 1997 through 1999 at three large hospitals in central and northeastern Pennsylvania (Milton S. Hershey Medical Center, Geisinger Medical Center, and Lehigh Valley Hospital) and The Johns Hopkins Hospital in Baltimore, Md. After the sample was identified from the hospitals' tumor registries, hospital employees in each site recruited subjects and obtained informed consent. The research protocol was approved by institutional review boards at Penn State's University Park campus and the four hospitals.

Because the main research goal was to study the effects of cancer survivorship on employment, eligibility was limited to adults ages 25 to 62 years old at diagnosis. Virtually all types of cancer, except for superficial skin cancers, were included. Most cases diagnosed at Stage 4 were excluded because these people were unlikely to survive to the end of a multiyear study; the exceptions were some Stage 4 leukemias, lymphomas, and plasma cell cancers with relatively favorable prognoses. Because the registry at John Hopkins lacked administrative control over urological cancers, they were omitted from that facility's sample. Survivors who could not be interviewed in English were also excluded.

Approximately 5,000 cases (2,000 from Johns Hopkins and 3,000 from the Pennsylvania registries) met the eligibility criteria. Forty-three percent of the eligible cases gave consent (N=2,076). Of these, 75 subjects were discovered to be ineligible when they were contacted for an interview. Eighty-eight percent of the consenting and eligible cases (N=1,763) were interviewed (Short, Vasey, and Tunceli 2005). Eighty-nine percent of surviving respondents in the first interview participated in the second interview, which was conducted in 2002 (N=1,511). For these analyses, we restricted the 2002 sample to the age group 55 to 64 years old, the age range common to the PSCSS and HRS (N=603). Additional exclusions included subjects who were still in initial treatment for active cancer in 2002 (N = 12), who were not working (N = 123) or were self-employed (N = 75) at diagnosis, or who were involuntarily separated from their jobs from diagnosis to follow-up because of business closures or layoffs (N = 17). The resulting cancer sample consisted of 376 survivors who were disease-free (N= 302), who had new cancers (N = 69), or who had missing data with respect to new cancers (N = 5).

Considering the low rate of recruitment into the study, we conducted extensive analyses using de-identified data for nonparticipants to identify correlates of nonparticipation and to test for biases (Short and Mallonee 2006). Although there were significant differences in participation by gender, race, cancer site, and facility among the roughly 5,000 eligible individuals, further tests did not find evidence of residual biases in multivariate models similar to those estimated here. We also assessed attrition after the first interview and found a few significant differences related to age, nonwhite race, any college education, poverty, and treatment at the facility with the lowest initial participation rate.

The first of four computer-assisted telephone interviews was conducted from October 2000 to December 2001. Many of the questions were taken from the HRS to facilitate comparisons between the surveys. The first interview ascertained current employment (at the time of interview) and, retrospectively, employment at the time of diagnosis. In this article, we compared employment data from the second PSCSS interview in 2002 with employment data from the 2002 HRS interview. At the second PSCSS interview, 71% of the cancer sample had survived from three to five years post-diagnosis. Survival time from diagnosis ranged from 26 to 70 months (mean = 45.9 months, standard deviation = 10.3).

Comparison Group

A comparison group with no history of cancer was drawn from the Health and Retirement Study. The HRS is a national longitudinal cohort study funded by the National Institute on Aging and conducted biennially by the Institute for Social Research (2006) at the University of Michigan. The sample interviewed in the first wave of the HRS was nationally representative of people ages 51 to 61 in 1992 (people who were born from 1931 to 1941), as well as their spouses regardless of age (Gustman, Mitchell, and Steinmeier 1995). In 1998, another cohort born during the period 1942 to 1947 was added. We used RAND's longitudinal compilation file (RAND 2004) where possible, and otherwise employed the HRS public use files for each wave.

We characterized the employment status of HRS subjects at a baseline comparable to the date of diagnosis in the cancer sample by following a procedure similar to one used by Bradley and colleagues (2005). Baseline dates were randomly assigned to the HRS sample from dates of diagnosis in the cancer sample, separately by gender, with replacement. We excluded HRS subjects with any history of cancer, other than superficial skin cancers, and those who were not working or self-employed at baseline, or who were involuntarily separated at the time of the follow-up. After also restricting the HRS sample to the age range common to both surveys, there were 2,772 cases for analysis.

While we might have implemented these comparisons solely with data from HRS, which include self-reporting of cancer, we see the Penn State Cancer Survivor Survey as offering several advantages as the source of our cancer sample. First, the number of cancer cases meeting the requirements of our analysis (incident cases diagnosed in adults 62 or younger with three to five years of follow-up) is considerably smaller in the HRS. Second, we expect that the nature and timing of cancer incidence (incorporating the exclusion of common skin cancers) is measured more reliably from cancer registries than from comparing and reconciling respondents' answers to the question: "Has a doctor ever told you that you have cancer or a malignant tumor, excluding minor skin cancer?"

Empirical Strategy

We adopted a standard difference-in-differences (DD) estimation strategy for identifying the extent to which employer-sponsored health insurance (ESI) inhibits employment transitions (job-to-job, full time to part time, and labor force exits) among cancer survivors. For expositional convenience in discussing this methodology, we use the term "job lock" in a broad sense to refer to insurance-induced reductions in any of these transition probabilities.

As we subsequently explain, the DD estimation strategy addresses first-order issues of data comparability associated with using different data sources for the cancer and noncancer comparison groups. It also accounts for first-order differences in job and worker quality associated with job-related health insurance. The basic relationship to be estimated takes the form,

$$\begin{aligned} y_i = \alpha_o + \alpha_1\, CANCER_i + \alpha_2\, ESI_i \\ + \alpha_3\, CANCER_i * ESI_i + \varepsilon_i, \end{aligned} \tag{1}$$

where the DD estimator is α_3, the difference in the effect of ESI on employment transitions for workers with a history of cancer compared to workers with no history of cancer. Here, ε_i is the error term that denotes the effect of unobservable factors affecting the propensity for employment changes.

Following Meyer (1995), the statistical prerequisites for obtaining an unbiased estimate of α_3 can be identified by taking expected values of the differences implied by equation 1 for workers in the four subgroups defined by the cross-classification of cancer and health insurance. The expected differences between individuals with and without insurance are described by the following expressions, shown separately for the cancer and noncancer comparison groups. The superscripts c and o identify cancer survivors and other adults, respectively.

$$\begin{aligned} E[\Delta y^c] = \alpha_2 + \alpha_3 + E[\varepsilon_i | CANCER, ESI] \\ - E[\varepsilon_i | CANCER, No\ ESI] \end{aligned} \tag{2a}$$

$$E[\Delta y^o] = \alpha_2 + E[\varepsilon_i|No\ CANCER, ESI] - E[\varepsilon_i|No\ CANCER, No\ ESI] \tag{2b}$$

The expected value of the difference in these differences is the expected value of the DD estimator itself:

$$E[\Delta y^c - \Delta y^o] = \alpha_3 + \{(E[\varepsilon_i|CANCER, ESI] - E[\varepsilon_i|CANCER, No\ ESI]) - (E[\varepsilon_i|No\ CANCER, ESI] - E[\varepsilon_i|No\ CANCER, No\ ESI])\} \tag{3}$$

The difference-in-differences estimator is unbiased if the expression in curly brackets in equation 3 is zero. The expression in curly brackets corresponds to the difference between the following two expressions,

$$E[\varepsilon_i|CANCER, ESI] - E[\varepsilon_i|CANCER, No\ ESI]) \tag{4}$$

$$(E[\varepsilon_i|No\ CANCER, ESI] - E[\varepsilon_i|No\ CANCER, No\ ESI]) \tag{5}$$

Each of these last two expressions is the difference in the average unobserved propensity of workers with and without health insurance to make employment transitions, for the cancer and noncancer groups, respectively.

As equation 3 demonstrates, the validity of the DD estimator does not require comparability of the unobservable labor-market tendencies of the two insurance groups (i.e., zero values for both expressions 4 and 5) because differencing between the cancer and noncancer groups (i.e., subtracting expression 5 from expression 4) "nets out" any insurance-related differences. Similarly, by rearranging the terms in equation 3 to difference out any cancer-related (or survey-specific) differences between the insurance groups, one can show that the unobservable labor market tendencies in the cancer and noncancer samples need not be identical either. The netting out of unobservable differences is the great advantage of the DD estimator over a single-difference estimator that contrasts cancer survivors with and without insurance but has no noncancer group (where the correlation between unobservables and health insurance must be zero), or a single-difference design that contrasts insured cancer survivors with other insured adults but has no uninsured group (where the correlation between unobservables and cancer must be zero).

What the DD estimator does require is that any insurance-related differences in the unobservable tendencies affecting employment transitions must be similar in the cancer and noncancer groups (or, equivalently, that any cancer-related differences must be similar in the two insurance groups). This condition could be violated in our application if working cancer survivors value jobs with ESI more highly than other workers for reasons other than the availability of health insurance. For example, it could be that jobs with ESI provide other workplace amenities that are differentially valued by those with a history of cancer; examples might include more flexible work schedules, more sick days, or an ability to work from home. Because these amenities generally will not be as valuable to healthy survivors as they are to those with ongoing health problems, a natural specification check is to re-estimate the baseline DD model using only healthy survivors. This is not a perfect test, however, since healthy survivors may not value health insurance as much as less healthy survivors either.

To support the assumptions underlying our DD estimator, we put considerable weight on comparing the two samples over the same calendar time period. This consideration, along with the availability of the requisite data items, largely dictated our decision to select the noncancer group from the HRS instead of another longitudinal national survey, although our use of the HRS necessarily limited the analysis to older workers with and without cancer.[1] Otherwise, for example, had there been differences in macroeconomic conditions at the times the samples were interviewed, and had these differences exerted a varying influence on employment transitions for jobs with and without health insurance, the identifying assumption underlying our DD estimator would be violated. Also, because the cancer survey questionnaire relied heavily on the HRS questionnaire, drawing the noncancer sample from the HRS enhanced data comparability.

Given that our cancer sample was regional rather than national, we sought to control for regional differences in labor market conditions by including indicator variables for the census divisions in our multivariate DD models. To account for regional differences in a less parametric way, we also re-estimated the models with the comparison group restricted to the same census divisions as the cancer sample, thereby (implicitly) interacting census division with all other coefficients in the model (including the interaction of health insurance and cancer survivorship that identified job lock).

Initially, we compared separately for the cancer and noncancer samples unadjusted differences in employment transitions associated with having job-related health insurance from one's own employer. Job-related health insurance was measured at baseline (1997, 1998, or 1999), just prior to diagnosis for the cancer survivors. Beginning with everyone employed at baseline, three kinds of transitions from baseline to follow-up in 2002 were examined: a change in employers; a change in employment status (working vs. not working); and, for full-time workers at baseline, a change from full-time to part-time employment (defined as working fewer than 35 hours per week).

These simple DD estimates were supplemented with a regression-based approach that was conditioned on a set of job and worker characteristics likely to influence the degree of job attachment and/or the likelihood of having ESI.[2] In particular, we controlled for the following job characteristics, measured at baseline: pension type (none, defined benefit, defined contribution, both defined benefit and defined contribution, and "unknown"); occupation (managerial, professional, or technical versus all others); and job tenure (measured in months). In addition, an indicator for being employed full time was included in the models of labor force exits and job change[3] to account for the degree of labor market attachment in the baseline period. We also controlled for two spousal job characteristics at baseline (spousal employment and spousal health insurance) that may influence job attachment by providing access to an alternative source of income, or in the case of spousal health insurance, an alternative source of group health coverage.

In addition to the baseline employment variables described previously, we also controlled for a number of sociodemographic variables. These include age (dummy variables for each year), race (white versus nonwhite), marital status (married/partnered versus not married), the presence of children under age 18 in the household, educational attainment (less than high school, high school, some college, college, and post college), and census region and division (Mid-Atlantic, including Pennsylvania; South Atlantic, including Maryland; New England; rest of South; Midwest; West). All models included an indicator for the presence of any of five common chronic conditions (diabetes, chronic lung disease, heart disease, stroke, and arthritis). These were measured at follow-up, since comparable baseline measures were not available for the cancer sample.

We did not use the HRS weights in our analysis, since we were not attempting to make estimates for the U.S. population from the HRS. The HRS weights are post-stratified by birth year, gender, and race/ethnicity to control totals for the U.S. population. Our regressions controlled for differences between the HRS and cancer samples along these dimensions.

Results

Table 1 describes the characteristics of the male and female PSCSS ("cancer") and HRS samples by insurance status. Overall, the cancer survivors were more likely to work in managerial/professional/technical jobs and, in the group without ESI, to have pensions (especially defined benefit pensions) and longer tenure in the baseline job. These employment differences were consistent with sociodemographic differences between the two samples, where the cancer sample was younger (except for males without ESI), more often white, and more likely to have a college degree. The cancer sample was less likely to

Table 1. Characteristics of males and females working at baseline/diagnosis, by insurance status and sample

	Males (%)				Females (%)			
	Own ESI		Not own ESI		Own ESI		Not own ESI	
Characteristic	PSCSS	HRS	PSCSS	HRS	PSCSS	HRS	PSCSS	HRS
Any new cancers	18.7	0.0	21.3	0.0	14.2	0.0	21.7	0.0
Chronic conditions								
Diabetes	16.5	15.7	19.4	16.1	13.8	12.0	8.4	12.9
Chronic lung disease	9.9	6.4	8.1	4.8	13.8***	7.1	11.2	6.8
Heart disease	17.6	15.5	30.7***	14.8	20.7***	12.3	15.9*	10.3
Stroke	3.3	3.1	1.6	2.3	4.3	4.1	.9	2.5
Arthritis	38.5	41.1	35.5	38.2	48.3	54.5	47.7	54.1
Pension (baseline)								
None	7.7	11.9	22.6***	58.4	8.6	11.9	40.2***	58.9
Defined benefit	48.4***	31.4	45.2***	12.6	43.1*	34.8	25.2***	13.3
Defined contribution	30.8	32.0	27.4	20.3	40.5**	30.8	25.2*	18.1
Both	6.6***	20.3	3.2	4.8	1.7***	17.4	3.7	4.3
Unknown	6.6	4.4	1.6	3.9	6.0	5.1	5.6	5.5
Job tenure in years (mean)	18.0	16.0	17.2***	8.2	15.2	13.9	11.4***	8.7
Occupation categories								
Managerial, professional, technical	53.9***	37.2	45.2***	26.8	50.9***	36.4	42.1***	23.8
Others	46.2***	62.3	54.8***	71.3	49.1***	62.3	57.9***	73.9
Unknown	0.0	.5	0.0	1.9	0.0	1.3	0.0	2.3
Census division								
Mid-Atlantic	76.9***	12.6	66.1***	12.0	53.5***	13.5	61.7***	13.1
Midwest	0.0	26.0	0.0	22.1	0.0	24.7	0.0	28.5
South Atlantic	23.1	25.3	33.9	30.8	46.6***	27.6	38.3**	27.3
Rest of South	0.0	16.4	0.0	16.2	0.0	14.7	0.0	11.7
West	0.0	16.6	0.0	15.9	0.0	15.5	0.0	15.8
Age								
55–59	51.7*	42.5	33.9	32.6	62.1**	50.5	58.9**	47.8
60–61	22.0	24.5	16.1**	31.3	20.7	23.2	20.6	21.3
62–64	26.4	33.1	50.0**	36.1	17.2**	26.3	20.6**	31.0
Nonwhite	5.5***	16.1	6.5**	17.9	7.8***	24.1	4.7***	17.5
Married/partner (baseline)	90.1	85.8	91.9	91.3	65.5	65.0	84.1	78.5
Spouse has employer insurance (baseline)	14.3***	28.1	35.5*	48.1	17.2***	30.3	63.6*	54.4
Children < 18	9.9	8.7	4.8	6.8	4.3	.2	5.6	3.9
Education								
Less than high school	6.6***	19.4	3.2***	22.6	4.3***	15.2	4.7***	22.2
High school	34.1**	46.5	32.3**	47.1	37.9***	54.9	31.8***	56.9
Some college	17.6***	4.1	21.0***	6.8	22.4***	5.2	25.2***	5.1
College	11.0	15.2	17.7	14.5	10.3	12.6	18.7***	9.0
Post college	30.8***	14.7	25.8***	8.4	25.0***	12.0	18.7***	6.8
Spouse employment								
Spouse working	60.4	55.3	59.7	62.9	48.3	45.9	65.4	59.0
Spouse not working	29.7	27.1	30.7	25.5	11.2	14.9	15.0	17.8
Without a spouse	9.9	14.2	8.1	8.7	34.5	35.0	15.9	21.5
Unknown	0.0*	3.5	1.6	2.9	6.0	4.3	3.7	1.8
Number	91	1,013	62	310	116	936	107	513

Notes: PSCSS= Penn State Cancer Survivor Study; HRS= Health and Retirement Study; ESI= employer-sponsored health insurance.
* $p<.1$, ** $p<.05$, *** $p<.01$: significant difference between cancer survivors and HRS respondents.

have a spouse with ESI, except for females without ESI. The samples differed by region of the country because the PSCSS sample was regional (central and northeastern Pennsylvania and Baltimore, Md.) while the HRS sample was national.

In Tables 2, 3, and 4 we present simple (unadjusted) DD estimates of the extent to

Table 2. Exit rates of employees at follow-up, by cancer and insurance status

	Male exits		Female exits	
	Yes ESI	No ESI	Yes ESI	No ESI
Cancer survivor				
Yes				
Percent	17.9	46.8	15.6	38.3
Number	91	62	116	107
No				
Percent	23.2	23.2	19.6	25.1
Number	1,013	310	936	513
Row difference	− 5.6 (4.2)	23.5 (6.8)***	− 4.0 (3.6)	13.2 (5.1)**
Difference-in-differences column difference	−29.2 (8.0)***		−17.2 (6.2)***	

Note: Standard errors are in parentheses.

which having ESI through one's employer inhibited employment transitions by male and female cancer survivors. These estimates are similar in magnitude to the regression-adjusted estimates subsequently shown in Table 5, but provide a more transparent picture of the influences of ESI on employment transitions.

Focusing first on exits from the labor force (Table 2), we find that among individuals *without* ESI, cancer survivors left work at a greater rate than their counterparts with no history of cancer (men: +23.5 percentage points, $p<.01$; women: +13.2 percentage points, $p<.05$). In contrast, for individuals *with* ESI, the exit rate for cancer survivors was much smaller and was not statistically different from the exit rate for those with no history of cancer.[4] Differencing these differences yields DD estimates for the effect of ESI on exit rates among cancer survivors of −29.2 percentage points ($p<.01$), and −17.2 percentage points ($p<.01$), for men and women, respectively, suggesting that having health insurance through one's employer substantially reduced transitions out of the labor force for both male and female cancer survivors.

For those working at follow-up, job-to-job transition rates varied substantially across workers with and without a history of cancer and by insurance status (Table 3). Among individuals without employer coverage, cancer survivors changed jobs at a *higher* rate than other workers (men: +27.4 percentage points, $p<.01$; women: +16.7 percentage points, $p<.05$). Among workers with ESI, job transition rates were significantly *lower* for female cancer survivors than for other female workers (−9.7 percentage points, $p<.05$), while the difference for males (−5.9 percentage points) was also negative but not

Table 3. Rates of employee job change (changing job vs. keeping same job) at follow-up, by cancer and insurance status

	Male change rates		Female change rates	
	Yes ESI	No ESI	Yes ESI	No ESI
Cancer survivor				
Yes				
Percent	13.3	60.6	8.2	48.5
Number	75	33	98	66
No				
Percent	19.3	33.2	17.8	31.8
Number	778	238	752	384
Row difference	−5.9 (4.1)	27.4 (9.1)***	−9.7 (3.1)***	16.7 (6.6)**
Difference-in-differences column difference	−33.4 (10.0)***		−26.4 (7.3)***	

Note: Standard errors are in parentheses.
*$p<.1$; ** $p<.05$; *** $p<.01$.

Table 4. Part-time job rates at follow-up, by cancer and insurance status among employees who worked full time at baseline

	Males working part time		Females working part time	
	Yes ESI	No ESI	Yes ESI	No ESI
Cancer survivor				
Yes				
Percent	12.9	42.3	6.7	41.9
Number	70	26	89	43
No				
Percent	6.9	11.4	9.9	22.7
Number	35	185	668	225
Row difference	5.9 (4.1)	30.9 (10.0)***	−3.1 (2.9)	19.2 (8.0)**
Difference-in-differences column difference	−25.0 (10.8)**		−22.3 (8.5)***	

Note: Standard errors are in parentheses.
*$p<.1$; ** $p<.05$; *** $p<.01$.

statistically significant. The DD estimates for job-to-job transitions were large and statistically significant: −33.4 percentage points ($p<.01$) for males and −26.4 percentage points ($p<.01$) for females.

Finally, among those working full time at baseline, we examined variations in transition rates into part-time work by cancer history and insurance status (Table 4). We found a large and statistically significant difference between cancer survivors and other workers in the propensity to move to part-time work for those *without* ESI (men: +30.9 percentage points, $p<.01$; women: +19.2 percentage points, $p<.05$), but not for workers *with* ESI. Taken together, these changes yield DD estimates of −25.0 percentage points, ($p<.05$) for men and −22.3 percentage points ($p<.01$) for women, pointing to a large role for ESI in impeding survivors' transitions into part-time employment.

In Table 5, we replicate our simple DD estimates using linear probability models that allowed us to condition on the set of job and worker characteristics described in the previous section. Our regression-adjusted job-lock estimates, measured by the coefficients on the interaction terms involving cancer survivorship and ESI, are shown in the third row of Table 5. Consistent with the simple DD estimates, we found large negative effects of ESI on the employment transitions of cancer survivors of both genders, with magnitudes that were similar to the unadjusted DD estimates presented in Tables 2 to 4. Estimates were significant at the .05 level or better for all three employment transitions.[5]

Turning briefly to other covariates, the independent effect of cancer on employment was evident for both genders, while the independent effect of ESI was significant only among women. Defined-benefit pension coverage had a predictable effect on labor supply in our sample of people ages 55 to 64, increasing the likelihood of retirement and decreasing the probability of changing employers. Spousal employment and spousal health insurance coverage were significant predictors of employment mobility in some specifications, although their magnitude and statistical significance varied by gender. Female workers with employed spouses were more likely to work and less likely to change jobs at follow-up compared to females who were unmarried or married to nonworking spouses. In addition, female workers who were married to husbands with health insurance coverage were less likely to work at follow-up.

In Table 6, we report the results of three sensitivity checks. For comparison purposes, our baseline estimates from Table 5 are displayed in Panel A. In Panel B, we re-estimated the DD parameter using a probit specification in conjunction with a procedure developed by Ai and Norton (2003) for calculating interaction effects and their associated standard errors in logit and probit models (Norton, Wang, and Ai 2004). There is little difference between the average interaction effects derived from the probit models

Table 5. Linear probability models by gender

Variables	Male: Exit (n=1,454)	Male: Job change (n=1,107)	Male: Part time (n=1,009)	Female: Exit (n=1,651)	Female: Job change (n=1,288)	Female: Part time (n=1,022)
Cancer survivor	.234*** (.073)	.438*** (.098)	.262*** (.094)	.175*** (.054)	.173** (.068)	.251*** (.081)
Own employer-sponsored insurance	−.013 (.032)	−.036 (.039)	−.030 (.025)	−.050* (.028)	−.068** (.032)	−.101*** (.032)
Interaction term for cancer survivor and own employer-sponsored insurance	−.240*** (.081)	−.402*** (.106)	−.214** (.102)	−.154** (.063)	−.271*** (.072)	−.233*** (.085)
Months from diagnosis	.001 (.001)	.005*** (.001)	−.001 (.001)	−.001 (.001)	.001 (.001)	.001 (.001)
Presence of any other conditions	.031 (.023)	.020 (.026)	.007 (.019)	.061*** (.021)	.018 (.026)	.010 (.023)
Full-time worker	.022 (.049)	.001 (.061)		−.054* (.028)	.005 (.032)	
Job tenure in months	−.001 (.001)	−.006*** (.001)	.002* (.001)	.002* (.001)	−.009*** (.001)	−.001 (.001)
Managerial, professional, technical occupation	.003 (.030)	−.034 (.034)	−.019 (.024)	−.009 (.026)	.030 (.030)	.019 (.026)
Defined benefit pension plan	.096*** (.036)	−.096** (.043)	−.024 (.032)	.062* (.033)	−.099*** (.038)	−.039 (.036)
Defined contribution pension plan	.001 (.033)	−.061 (.042)	−.025 (.029)	.026 (.030)	−.100*** (.036)	−.101*** (.034)
Defined benefit *and* defined contribution	.063 (.041)	−.118** (.048)	−.044 (.033)	.113*** (.041)	−.081* (.047)	−.100*** (.037)
Age 56	.052 (.040)	−.007 (.051)	.001 (.034)	.001 (.037)	.020 (.046)	.017 (.034)
Age 57	.075 (.043)	−.017 (.052)	−.016 (.030)	.062 (.043)	.014 (.049)	−.011 (.035)
Age 58	.140*** (.046)	.107* (.061)	.024 (.038)	.052 (.041)	−.030 (.046)	.011 (.036)
Age 59	.070 (.039)	−.009 (.049)	−.048* (.028)	.083** (.039)	.101** (.047)	.090** (.039)
Age 60	.100*** (.038)	.012 (.047)	.013 (.031)	.027 (.038)	−.022 (.044)	.044 (.035)
Age 61	.128*** (.039)	.019 (.048)	.028 (.034)	.064 (.041)	.015 (.046)	.075* (.039)
Age 62	.260*** (.046)	−.006 (.055)	.091** (.043)	.126*** (.044)	.028 (.053)	.163*** (.049)
Age 63	.259*** (.046)	.018 (.053)	.151*** (.049)	.197*** (.047)	.042 (.054)	.217*** (.058)
Age 64	.297*** (.045)	.032 (.055)	.104** (.045)	.117** (.046)	−.055 (.049)	.114** (.051)
Nonwhite	−.035 (.031)	.039 (.037)	−.011 (.026)	−.018 (.026)	.003 (.030)	.048* (.028)
Any children under 18	−.049 (.032)	.019 (.040)	.001 (.029)	−.083* (.050)	.004 (.064)	−.034 (.048)
Single	.037 (.039)	−.047 (.043)	.005 (.032)	−.105*** (.035)	.009 (.038)	−.010 (.034)
Spouse has employer-sponsored insurance	−.004 (.026)	−.005 (.029)	.009 (.022)	.057** (.026)	.012 (.029)	.000 (.027)
Spouse working	−.034 (.027)	−.014 (.031)	−.011 (.022)	−.129*** (.033)	−.067* (.035)	.014 (.033)
High school	−.021 (.032)	.033 (.037)	−.009 (.025)	−.068** (.032)	.068** (.033)	.029 (.032)
Some college	.009 (.058)	.032 (.068)	.058 (.053)	−.039 (.049)	.153** (.057)	.015 (.051)
College	−.056 (.043)	.046 (.051)	.031 (.034)	−.098** (.044)	.113** (.050)	.002 (.044)
Post college	−.120** (.046)	.063 (.054)	.027 (.038)	−.067* (.047)	.055 (.051)	.030 (.047)
Mid-Atlantic	.000 (.074)	−.128 (.079)	.096*** (.030)	.008 (.055)	−.001 (.063)	−.066 (.074)
Midwest	.043 (.072)	−.016 (.078)	.098*** (.025)	.035 (.053)	.001 (.062)	−.022 (.073)
South Atlantic	−.037 (.071)	−.027 (.078)	.102*** (.024)	.017 (.052)	.019 (.061)	−.088 (.071)
Rest of South	.070 (.075)	−.058 (.081)	.107*** (.029)	.028 (.056)	.045 (.066)	−.040 (.074)
West	−.020 (.073)	−.010 (.080)	.094*** (.027)	.067 (.056)	.009 (.065)	−.051 (.074)

Note: Standard errors are in parentheses.
*$p<.10$; ** $p<.05$; *** $p<.01$.

and those obtained from our baseline linear probability models.

In Panel C, we re-estimated our baseline DD models using only HRS respondents who lived in the same census divisions as participants in the Penn State Cancer Survivor Study (the Mid-Atlantic and South Atlantic regions). These geographic indicators crudely control for local labor market conditions, within the constraints of the geographic identifiers available on the HRS public use files. The fact that our original estimates remained virtually unchanged after removing seven of the nine census divisions from our

Table 6. Specification checks

Variables	Male			Female		
	Exit	Job change	Part time	Exit	Job change	Part time
Panel A: Baseline estimates from Table 5						
N	1,454	1,107	1,009	1,651	1,288	1,022
Cancer survivor	.234*** (.073)	.438*** (.098)	.262*** (.094)	.174*** (.054)	.173** (.068)	.251*** (.081)
Own employer-sponsored insurance	−.013 (.032)	−.036 (.039)	−.030 (.025)	−.050* (.028)	−.068** (.032)	−.101*** (.032)
Interaction term for cancer survivor and own employer-sponsored insurance	−.240*** (.081)	−.402*** (.106)	−.214** (.102)	−.154** (.063)	−.271*** (.072)	−.233*** (.085)
Panel B: Ai-Norton[a] interaction effects (probit models)						
N	1,454	1,107	1,009	1,651	1,288	1,022
Interaction term for cancer survivor and own employer-sponsored insurance	−.236*** (.088)	−.422*** (.115)	−.166 (.100)	−.169** (.069)	−.280*** (.085)	−.283*** (.099)
Panel C: Mid-Atlantic and South Atlantic census regions only						
N	663	522	482	807	630	502
Cancer survivor	.234*** (.079)	.469*** (.110)	.280*** (.100)	.162*** (.057)	.140* (.073)	.278*** (.083)
Own employer-sponsored insurance	−.006 (.047)	−.003 (.055)	−.001 (.041)	−.087** (.042)	−.152*** (.049)	−.067 (.043)
Interaction term for cancer survivor and own employer-sponsored insurance	−.250*** (.087)	−.408*** (.119)	−.229** (.109)	−.139** (.069)	−.191** (.081)	−.244*** (.089)
Panel D: Healthy survivors only (no recurrence)						
N	1,424	1,086	992	1,612	1,263	1,004
Cancer survivor	.262*** (.078)	.373*** (.116)	.195* (.103)	.151*** (.057)	.187** (.073)	.253*** (.088)
Own employer-sponsored insurance	−.011 (.032)	−.033 (.038)	−.025 (.025)	−.051* (.028)	−.064** (.032)	−.103*** (.032)
Interaction term for cancer survivor and own employer-sponsored insurance	−.321*** (.087)	−.347*** (.123)	−.169 (.111)	−.156** (.067)	−.274*** (.078)	−.264*** (.092)

Note: Standard errors are in parentheses.
[a] Ai and Norton (2003); Norton, Wang, and Ai (2004).
*$p<.1$; ** $p<.05$; *** $p<.01$.

comparison group suggests that geographic differences in labor market dynamics did not play a major role in our results.

Finally, in Panel D, we re-estimated our baseline DD models restricting attention to "healthy survivors" (those with no recurrence of their cancer). The similarity between these estimates and those from the full sample provides some evidence that our original estimates were not confounded by cancer survivors differentially valuing jobs with ESI for reasons beyond the availability of health insurance.

Discussion

When we compared changes in the employment of older cancer survivors to changes in the employment of otherwise similar older workers without cancer, we found significant differences only for survivors in jobs without health insurance at diagnosis (compared to other workers in jobs without insurance). These survivors, who had no job-related health insurance to lose by changing employment, were more likely to stop working, switch jobs, and cut back from full-time to part-time work than their counterparts without cancer. However, survivors in jobs that did have health insurance at diagnosis—who had something to lose—were no more likely to make changes in their employment after being treated for cancer than insured workers without cancer. These findings, which relate to employment changes over an average of four years for older male and female survivors treated for a variety of cancers, are consistent with the findings of Bradley and her colleagues (2005) for breast cancer survivors over a shorter time frame of six to 18 months post-diagnosis.

Despite the barriers that cancer survivors confront in purchasing health insurance in the individual market, national statistics show that their enrollment in private insurance is comparable to the rate of private enrollment for other Americans (Sabatino et al. 2006). If anything, adult cancer survivors are somewhat more likely to be insured than other adults under age 65 because of increased enrollment in Medicare and Medicaid (public programs that cover severely disabled adults under age 65 who are unable to work). The propensity of survivors to remain in jobs with health insurance benefits, as documented in this paper, may help to explain why the rate of private insurance for cancer survivors is similar to the national average.

Although our findings imply that cancer survivors succeed in staying insured by holding onto jobs with insurance benefits, these findings also imply that the employment opportunities available to survivors are more constrained by health insurance considerations than the opportunities available to other workers. Not only are survivors handicapped in advancing their careers or pursuing leisure interests by their need for health insurance, but those who have continuing health problems as a result of cancer and treatment may be less able to accommodate changes in their health or functional status by changing jobs or cutting back on work. In these ways, survivors pay a higher "price" for health insurance that affects their quality of life and adds to the economic burden of cancer.

In making comparisons to the HRS, which follows a national sample of older Americans into retirement, we focused on employment transitions in an older group of working cancer survivors (ages 55 to 64 in 2002). Because cancer is a disease associated with aging, older workers are particularly affected by cancer. Nearly half of new cancers diagnosed in the age group from 20 to 64 are diagnosed after age 54 (Ries et al. 2006). Nevertheless, because our estimates of insurance-related differences in employment transitions do not necessarily generalize to younger workers, the age group that we examined is an important limitation of our study.

The extended recall of employment at diagnosis and the representativeness of the cancer survivor sample are also potential limitations. Employment status and job characteristics at diagnosis were ascertained in the first interview of our cancer survivor survey, one to five years post-diagnosis. Not only did the long recall increase the likelihood of reporting errors, but we were unable to ask about key variables at the time of diagnosis (including wage rates and health status).

The generalizability of employment patterns in our cancer sample to the employment of cancer survivors nationally is also a potential concern. Compared to cancer survivors nationally (Ibrahim, Short, and Tunceli 2004; Short, Vasey, and Tunceli 2005), our cancer sample—drawn, as already noted, from the tumor registries of four medical centers in Pennsylvania and Maryland—was relatively advantaged in terms of socioeconomic status, employment rates, and enrollment in private insurance. Although we used data stripped of personal identifiers from the sampling frame to find correlates of participation and controlled for them in our employment analyses, nonparticipation bias is another issue potentially affecting the representativeness of the cancer sample.

Our difference-in-differences approach should net out most unobserved differences in job attachment between the cancer and noncancer samples, or between workers with and without job-related health insurance. However, we acknowledge that our DD estimator will be biased if cancer-related differences in unobserved job attachment vary with insurance (or, equivalently, insurance-related differences vary between the cancer and noncancer samples). Nevertheless, the fact that our DD estimates are large and highly significant, consistent across three different employment outcomes for both genders, and robust to several specification checks heightens confidence that our estimates primarily reflect insurance-induced reductions in employment transitions, rather than confounding from unobservables.

Cancer survivors live with a variety of health and financial risks. In addition to the possibility of recurrence, their health risks include a higher rate of unrelated second cancers and the possibility of continuing or delayed effects from treatment (Hewitt, Greenfield, and Stovall 2005). Their financial risks include the possibility of large out-of-pocket health care expenditures and a considerable loss of earnings in the event of severe disability or premature death. Even if they successfully negotiate these other risks, cancer survivors also confront the risk of becoming uninsured without job-related health insurance. Our findings suggest that survivors effectively "insure" against this risk by remaining in jobs with health insurance in situations where exiting the labor force, reducing hours, or changing employers would otherwise be the more attractive alternative. Public policies that provide affordable alternatives to job-related health insurance would spread this risk more broadly and relieve cancer survivors (and others with chronic illnesses) of the costs associated with protecting their access to health insurance.

Notes

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1 For example, the Survey of Income and Program Participation now fields only one panel at a time, and none of the panels spanned the same calendar years as our cancer survivor data. The Panel Study of Income Dynamics identifies respondents with employer-sponsored insurance, but does not distinguish policyholders from dependents.

2 For simplicity, we used linear probability models to implement the DD estimator. Estimates from a probit specification, with the interaction effect calculated using a procedure derived by Ai and Norton (2003) and Norton, Wang, and Ai (2004), were virtually identical.

3 Models of the full-time / part-time transition apply only to individuals working in the follow-up period.

4 Note that for men who did not have cancer, exit rates were identical for those with and without ESI (23.2%), while there was a large differential for men with a history of cancer: 17.9 percentage points for those with ESI versus 46.8 percentage points for those without ESI. The pattern was similar for women.

5 We also estimated a DDD model in which our DD estimates were differenced again using the existence of spousal health insurance as the third margin. Estimates were mostly in the same range as our DD estimates, but were not precise and were never statistically significant. Given the small number of individuals who simultaneously had ESI, had a spouse with health insurance, and survived cancer (13 males and 20 females, respectively), we do not believe that such an approach can be used to reliably identify job lock in our sample.

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