The Effect of Expanded Mental Health Benefits on Treatment Initiation and Specialist Utilization

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Objective. To measure the effects of a mental health benefit design change on treatment initiation for psychiatric disorders of employees of a large U.S.-based company.

Data Sources. Mental health treatment administrative claims data plus eligibility information provided by the company for the years 1995–1998.

Study Design. We measure the effect of a change in mental health benefits consisting of three major elements: a company-wide effort to destigmatize mental illness; reduced copayments for mental health treatment; and an effort to increase access to specialty mental health providers.

Data Extraction Methods. We identified the subsample of employees that were continuously enrolled in the company's health plan over the period 1995–1998, were between the ages of 18 and 65, and were actively employed.

Principal Findings. Our results suggest that the combined effect of destigmatization and reduced copayments led to an 18 percent increase ($p < .01$) in the probability of initiating mental health treatment. The results suggest that the effort to increase access to specialty providers was effective, but only for nonphysician providers: initiation at nonphysician mental health providers increased nearly 90 percent ($p < .01$) relative to nonspecialty providers, while use of psychiatrists declined by nearly 40 percent ($p < .01$).

Conclusions. Our results suggest that the benefit change increased initiation for mental health treatment overall and encouraged the use of nonphysician specialty mental health providers.

Key Words. Behavioral health, corporate health benefits, provider distance, employer sponsored health insurance

Mental health disorders represent one of the most common problems facing adults in the labor force. It is estimated that within a 12-month-period nearly 30 percent of the U.S. population experiences some diagnosable mental health or addictive disorder (Kessler et al. 1994). Some prior studies have found that mental disorders can substantially inhibit an employee's ability to work by increasing absenteeism and reducing on-the-job productivity (Kessler and

Frank 1997; Burton and Conti 1999; Burton et al. 1999; Kessler et al. 1999). Early initiation into treatment is frequently mentioned as an important means of controlling the severity of an episode of mental illness and associated costs such as absenteeism and impaired productivity (Goldberg and Steury 2001). Moreover, treatment of mental illness by mental health professionals can represent an improvement in quality of care relative to treatment by generalists. In this study, we assess the impact of a company-wide initiative to broaden mental health benefits on mental health treatment initiation at a large, nationally known communications and electronics company. The benefit policy change included three major components: (1) reduced copayments for mental health treatment, (2) increase modes of access to specialty mental health providers, and (3) an effort to destigmatize the treatment of mental disorders.

Prior research has shed light on different aspects of the company's benefit change. Results from the RAND Health Insurance Experiment (HIE) described the effects of cost sharing on the demand for mental health services. Keeler, Manning, and Wells (1988) examined how the demand for episodes of mental health services is affected by changing coinsurance rates and found that persons in the 95 percent coinsurance plan initiated mental health treatment at 24 percent of the rate of those in the free plan, while those on the 25 percent coinsurance plan initiated mental health services at 70 percent of the rate of those in the free plan.

Prior research has documented the effect of restricting coverage of mental health treatment within a corporate setting. Using data from a large U.S. manufacturing firm, Rosenheck et al. (1999) demonstrated the consequences of employer cost containment efforts affecting mental health treatment coverage. The restrictions on mental health coverage came in the form of increased deductibles, prior authorization for inpatient mental health care, and utilization review for both inpatient and outpatient care. Findings indicated that over a 3-year period the restrictions resulted in decreased mental health treatment costs that were completely offset by increased nonmental health costs. In addition, sick-days increased disproportionately among users

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of mental health services over time. Altogether, the results suggested that restrictions on mental health treatment might in fact lead to an overall cost increase when factoring in other aspects of the employers' bottom line. It is clear from the HIE and other studies (e.g., Goldman, McCulloch, and Sturm 1998) that benefit design can have a strong effect on behavioral health treatment use.

Mental health provider choice is an important issue for patient care; the debate regarding the merits of specialty mental health care versus non-specialty care is not new (see Frank and Kamlett 1989; Sturm and Wells 1995; Sturm, Meredith, and Wells 1996). While there is evidence that patient sickness has the strongest effect on the use of specialty care, other factors such as income, patient education, and ethnicity also predict the use of specialty care (Sturm, Meredith, and Wells 1996). In addition, other research has suggested that care initiated with nonphysician specialty providers is the most economical (Holmes and Deb 1998).

The stigma associated with mental illness has potentially important implications in the workplace. A recent study of employers found that when presented with hypothetical job applicants who were identical except for a diagnosis of either depression or diabetes, employers were less likely to hire the person diagnosed with depression (Glozier 1998). More recently, the stigma associated with mental illness has also been shown to decrease adherence to depression treatment (Sirey et al. 2001). To our knowledge, the present study is the first to focus on a program with a major stated emphasis to destigmatize mental health treatment.

INSTITUTIONAL SETTING

The setting for our study is the U.S.-based employees of a multi-national high-technology company specializing in communications and embedded electronic solutions. During the time period of our study the company employed between 70,000 and 80,000 people in the U.S. The company had offices and production facilities throughout the U.S., although a plurality of employees was located in the Midwest. The remainder of the U.S.-based employees was split roughly evenly between the southern and western states. Relatively few employees worked in the northeast.

The company had long provided a mental health and substance abuse benefit although in the early- to mid-1990s the company began to notice that expenses for mental health treatment services were increasing at a striking

rate. Beginning in January 1996, several important changes in the behavioral health benefits were implemented at the company. The changes are summarized in Table 1. Note that the out-of-pocket costs required to access (in-network) services between the pre- and postperiods were reduced. Coinsurance rates on inpatient care were changed from 80 to 90 percent, while 80 percent coinsurance for outpatient services was replaced with a flat \$10 co-payment.¹ The company also implemented selective contracting with mental health providers in place of a traditional fee-for-service indemnity benefit. The stated goal was to encourage treatment of mental illness at the "least intensive locus of care." Employees retained the option of accessing providers outside the network, but had substantially lower out-of-pocket costs if they used in-network providers.

As part of the change in benefits, the company allowed employees to self-refer through either the Internet or a telephone referral service. With the aid of computerized kiosks, employees were able to identify providers. In addition, the company started a campaign to eliminate the stigmatization associated with seeking mental health treatment. They distributed brochures and web-based information that demonstrated the recognition that the daily stress of life can sometimes take its toll.

In most parts of the country the company offered one or more fully insured health maintenance organizations (HMOs) as competing plan options alongside the company's self-insured health plan. In general, the company experienced an influx of enrollees into the self-insured plan and an outflow from the HMOs: the proportion of employees enrolled in the HMOs decreased from roughly 35 percent in 1996 to 16 percent in 1999. As discussed later, to minimize the possibility of selection bias from those in need of mental health treatment disproportionately enrolling in the self-insured plan after the benefit expansion, we focus our analyses only on those persons who were continuously enrolled in the self-insured plan over the period from 1995 to 1998.

DATA AND METHODS

Claims data files were acquired from the company through their claims data manager, Medstat. In addition, enrollment information for the period between 1995 and 1998 for persons enrolled in the company's health insurance plan was also acquired. The enrollment data were updated on a quarterly basis and contained date of birth, family status, gender, employment status, and home zip code. We alternatively estimate our models based on an unrestricted

Table 1: Description of Mental Health Benefits Pre- and Post-1996

	Post-1996		
	Pre-1996	In Network	Out of Network
Hospital inpatient	80% of semi-private room	90% of negotiated fees for semi-private room	50% of Negotiated Network fee for semi-private room
Hospital inpatient services	80% of miscellaneous charges	90% of negotiated fees	Maximum of 10 days per year 50% of Negotiated Network fee
Physician, inpatient hospital	80% of charges	90% of negotiated fees	Maximum of 10 days per year 50% of Negotiated Network fee
Physician's office	80% of charges	\$10 copayment per visit	Maximum of 10 days per year 50% of (maximum) Negotiated Network fee for licensed Ph.D./M.D.
Access to care	Contact EAP representative Contact provider directly No utilization review for IP or OP treatment		Maximum of 20 visits per year Contact CallCARE (referral service) Contact EAP representative Self-refer using kiosk or WWW
			Utilization review for IP admits; no utilization review for OP treatment

sample of all individuals that were ever enrolled for at least 1 complete year in the company's self-insured health plan and a sample that is restricted to employees that were continuously enrolled in the company's health plan over the 4-year period. In the unrestricted sample, we have 72,991 unique individuals and 202,801 person-year observations, and in the restricted model we have 29,883 individuals and 119,532 person-year observations. The fraction of employees enrolled in the company's health insurance plan grew over the 4-year period: in 1995 the 29,883 individuals represented about 67 percent of the individuals who were enrolled for the full year, but by 1998 the continuously enrolled 29,883 persons represented 55 percent of the total full-year enrollees. Because of the possibility of bias resulting from the benefit change drawing in a different group of mental health services users over time, our preferred estimates are derived from the restricted sample.

We identified mental health treatment based on the presence of an ICD-9 diagnosis code of 290–314, 780.1, 783.0, V11.3, V61.41, or V79.1 or a CPT code of 90801–90899. We define treatment initiation as the occurrence of a mental health claim after a period of no mental health claims. Following the previous literature, we define a new episode of mental health treatment as one occurring after an 8-week (56 days) period in which no mental health treatment was observed (Keeler, Manning, and Wells 1988; Huskamp 1999). In order to consistently measure initiation without censoring throughout all 4 years of our study, we exclude initiations occurring within the first 2 months of the year. Initiation at a mental health specialist is defined by whether the initial provider was a mental health professional.²

Table 2 shows the percentage of individuals that initiate treatment over the 4-year period using the alternative sample formulations. In 1995, just over 6 percent of employees initiated mental health treatment regardless of how the

Table 2: Mental Health Treatment Initiation Rates over Time

<i>Year</i>	<i>Restricted Sample</i>		<i>Unrestricted Sample</i>	
	<i>Observations</i>	<i>% Initiation</i>	<i>Observations</i>	<i>% Initiation</i>
1995	29,883	6.07	44,368	6.18
1996	29,883	6.72	51,493	6.98
1997	29,883	7.13	52,905	7.27
1998	29,883	7.22	54,035	7.24

Note: Restricted sample includes employees consistently enrolled in the company plan during the period 1995–1998. Unrestricted sample includes employees that were enrolled for at least 1 complete calendar year during the period 1995–1998.

sample was defined. In 1996, the rate of mental health treatment initiation increased 0.65 percent point in the restricted sample and 0.80 percent point in the unrestricted sample. The rate of initiation continued to increase in subsequent years. The regression models presented below will control for other factors that might explain the increase.

Using the claims data, we extracted the set of unique provider identifiers and zip codes associated with these mental health providers. For our purposes, this set of providers represents the universe of potential providers available to individuals in the analytic sample. Distance to the nearest mental health provider in our analytic sample was then calculated by computing the shortest distance from each home zip code (centroid) to a provider zip code (centroid). We use all providers observed in the period from 1995 to 1998 to calculate the distance. However, for the years 1996–1998, when the provider network was in place, distance was calculated using only those providers observed during those 3 years. Using the claims data set we found that employees had access to 6,022 mental health providers in 1995, of which 2,143 were nonphysician specialists and 391 were psychiatrists. Beginning in 1996, employees could access 4,625 providers, of which 1,892 were nonphysician specialists and 303 were psychiatrists.

Table 3 displays pre/postperiod descriptive characteristics for the restricted and unrestricted sample of employees. Because of the addition of the provider network, distance to mental health providers should generally increase between the pre- and postperiods. This holds for all cases with the exception of the unrestricted sample for nonphysician specialists. Note that because the composition of the unrestricted sample can change between the pre- and postperiod, it is possible that the new enrollees could be nearer to providers than the previous enrollees. Nonspecialist physicians are the nearest to employees, while psychiatrists are farthest. There are several noteworthy differences between the characteristics of the restricted and unrestricted samples. The unrestricted sample tends to have a greater fraction of nonwhites, women, persons with single coverage, and younger persons.

Following earlier research (Sturm, Meredith, and Wells 1996; Holmes and Deb 1998), we define three types of providers for the second part of our analysis examining the type of provider with whom patients initiated. Using provider type codes available in the claims data, we identified (1) nonspecialist physician providers, (2) psychiatrist providers, and (3) nonphysician specialist mental health providers. Table 4 displays pre/postdescriptive statistics for people initiating with the three different provider types. For convenience we

Table 3: Descriptive Characteristics for the Restricted and Unrestricted Sample of Employees

	<i>Restricted Sample</i>		<i>Unrestricted Sample</i>	
	<i>1995</i>	<i>1996–1998</i>	<i>1995</i>	<i>1996–1998</i>
Distance (miles)				
... to physician	2.14	2.63	1.68	1.82
... to nonphys specialist	3.21	3.43	2.80	2.72
... to psychiatrist	8.16	9.72	7.65	7.75
Race				
White	77.84%	77.84%	74.43%	74.15%
Nonwhite	22.16	22.16	25.57	25.85
Gender				
Female	30.06	30.06	32.66	32.39
Male	69.94	69.94	67.34	67.61
Coverage status				
Single	39.61	35.63	46.67	40.38
Family	60.39	64.37	53.33	59.62
Age category				
18–29	18.14	12.43	22.36	20.53
30–39	35.82	34.31	36.41	35.03
40–49	29.78	31.43	26.47	26.86
50–64	16.26	21.83	14.75	17.58
Region				
Northeast	1.99	2.29	2.42	2.99
Central	42.72	42.35	41.82	40.91
South	29.84	29.91	31.41	30.29
West	25.45	25.44	24.36	25.82
Sample size	29,883	89,649	44,368	158,433

Note: Restricted sample includes employees consistently enrolled in the company plan during the period 1995–1998. Unrestricted sample includes employees that were enrolled for at least 1 complete calendar year during the period 1995–1998.

only display sample statistics for the restricted sample; statistics for the unrestricted sample are similar. We observe that the percentage of treatment episodes initiated with nonspecialist physicians or psychiatrists dropped between the pre- and postperiods, while the percentage initiated with a nonphysician specialists increased sharply. The average distances tend to suggest that individuals preferred the closer of the three provider types, with the exception of psychiatrists, who always tended to be farther away than the alternatives.

Table 4 also contains a pre/postsummary of several severity indicator variables that were modeled on work by Ettner et al. (1998) for the purposes of

Table 4: Descriptive Statistics of Pre/Postprovider Choice Conditional upon Initiation, Sample Restricted to Employees Continuously Enrolled

	<i>Preperiod, 1995</i>			<i>Postperiod, 1996–1998</i>		
	<i>Nonspecialist Physician</i>	<i>Nonphysician Specialist</i>	<i>Psychiatrist</i>	<i>Nonspecialist Physician</i>	<i>Nonphysician Specialist</i>	<i>Psychiatrist</i>
Percentage	51.60	34.07	14.33	42.59	49.90	7.51
Distance (miles)						
To physician	2.57	1.78	1.16	3.13	2.25	1.54
To specialist	3.89	2.07	2.54	4.01	2.80	2.27
To psychiatrist	9.78	6.27	5.52	11.16	8.98	5.50
Race						
White	85.15%	85.92%	86.54%	84.01%	82.56%	85.84%
Nonwhite	14.85	14.08	13.46	15.99	17.44	14.16
Gender						
Female	40.81	43.37	43.46	41.63	41.49	41.65
Male	59.19	56.63	56.54	58.37	58.51	58.35
Coverage status						
Single	36.00	37.54	41.54	36.34	34.68	33.83
Family	64.00	62.46	58.46	63.66	65.32	66.17
Age category						
18–29	10.90	12.14	17.69	9.02	10.28	11.42
30–39	34.62	37.86	40.77	32.80	40.06	40.59
40–49	36.22	36.25	29.23	37.35	33.85	31.50
50–64	18.27	13.75	12.31	20.83	15.81	16.49
Region						
Northeast	1.82	1.46	2.31	2.20	1.40	1.48
Central	36.97	21.04	59.23	36.75	32.01	65.33
South	31.09	36.73	18.85	32.46	32.20	20.93
West	30.13	40.78	19.62	28.59	34.39	12.26
Severity						
Major depression	24.47	16.83	25.38	34.10	17.85	30.02
Manic disorder	5.45	3.56	5.38	6.82	1.75	1.69
Anxiety disorder	22.44	16.18	16.15	19.64	14.44	12.47
Personality disorder	1.28	2.27	1.54	1.6	1.43	1.90
Substance use	9.29	5.02	7.31	7.49	5.22	7.19
Conduct disorder	0.32	1.46	0.38	0.41	0.76	0.21
Sample size		1,814			6,299	

Note: Restricted sample includes employees consistently enrolled in the company plan during the period 1995–1998.

risk adjusting mental health claims data. One concern in examining results conditional upon initiation is that there may have been changes in the composition of the treatment initiators between the pre- and postperiods. For example, if the destigmatization effort was successful we would anticipate that people with less severe disorders would initiate in the postperiod. We observe

that the most notable change was the overall increase in the proportion of major depression diagnoses, the bulk of which were treated by generalists. It is interesting to note that despite the dramatic increase in the use of non-physician specialists between the pre- and postperiods, the overall severity composition of the treatment initiators did not change dramatically.

A potential concern with our data is that we observe a pre/postchange from a single employer, without the use of an untreated control group. Thus, if treatment initiation exhibited a secular trend over our analysis period, 1995–1998, we might wrongly attribute this broader trend to the company's benefit change. Using two national surveys, the National Household Survey on Drug Abuse and the Medical Expenditure Panel Survey, we computed measures of mental health treatment initiation for subsamples in each data set that were as comparable as possible to our study database. We found no significant secular trend over the analysis period.

To estimate the impact of the benefit change on mental health treatment initiation, we specify the following model:

$$\Pr(Y_{it} = 1) = F(\alpha Post_{it} + X'_{it}\beta) \quad (1)$$

where Y represents treatment initiation for individual i at time t ; $Post$ represents an indicator variable for whether the observation occurs during the 1996–1998 benefit change period, with α measuring the total effect of the benefit expansion; X is a matrix including indicators for gender, race, family status, age, and region; also included are distance measures to the nearest provider of each of the three types. Thus, the distance measures represent the relative “cost” to access each provider type. As such, we expect that individuals will be less likely to access mental health treatment if the distance to the nearest provider is lengthy, but the effect may differ by provider type. β is the vector of coefficients associated with X . Under the assumption that the function $F(\cdot)$ is logistically distributed the model can be estimated as a logistic regression. We use a Huber correction to adjust the standard errors for multiple observations on the same individual.³

To examine the provider choice decision, we condition on treatment initiation. In doing so, we do not have a full panel with which to estimate the provider choice model because employees do not initiate every year. Hence, we opt for a multinomial choice model to estimate the conditional probability of treatment initiation with each provider type. As above, we use a Huber correction to adjust the standard errors and thus to control for the correlation within individuals.

The specification of the provider type regression relies on several assumptions. First, we assume that the decision-making process is sequential. In other words, a patient decides whether or not to seek help (i.e., to initiate) and then, conditional upon deciding to initiate, chooses a provider type. Second, the patient makes the decision by comparing the expected utilities of visiting a nonspecialist physician, a psychiatrist, and a nonphysician specialist and then selecting the option with the greatest expected utility. The independent variables are the same as in (1) above. Formally, we posit:

$$\Pr(\text{Gen}_{it} = 1 | Y_{it} = 1) = F(\theta^{\text{Gen}} \text{Post}_{it} + X'_{it} \beta^{\text{Gen}}) \quad (2)$$

$$\Pr(\text{Psych}_{it} = 1 | Y_{it} = 1) = F(\theta^{\text{Psych}} \text{Post}_{it} + X'_{it} \beta^{\text{Psych}}) \quad (3)$$

$$\Pr(\text{NonPhys}_{it} = 1 | Y_{it} = 1) = F(\theta^{\text{NonPhys}} \text{Post}_{it} + X'_{it} \beta^{\text{NonPhys}}) \quad (4)$$

where each of the left-hand side variables represent the decision to initiate treatment with each of the three different provider types, Y_{it} is the decision to initiate treatment at all, and the θ terms represent the effect of the change in benefits on the conditional probability to initiate treatment at each provider type. If $F(\cdot)$ follow a logistic distribution then a multinomial logit model can be used to estimate the model in (2)–(4).

RESULTS

In Table 5, we present logit regression results for the factors affecting treatment initiation for our sample using two alternative sample definitions. We observe in both samples that coincident with the implementation of the mental health benefit change, treatment initiation increased. The coefficients for Post in the first row show that individuals were roughly 18 percent more likely to initiate treatment for a mental disorder in the postperiod relative to the prebenefit change period. The slightly (although trivially) larger initiation result for the unrestricted sample is potentially the result of adverse selection in the form of individuals anticipating greater mental health utilization needs entering the company's self-insured health plan. Consequently, our preferred initiation estimates are for the sample restricted to employees who were enrolled in the company plan during the entire 4-year period.

The results for the distance variables suggest that lengthier distances to mental health providers are associated with a lower likelihood of treatment initiation, although relationship is only statistically significant in the

Table 5: Logit Regressions for Mental Health Treatment Initiation, Restricted and Unrestricted Sample of Employees

	<i>Restricted Sample</i>		<i>Unrestricted Sample</i>	
	<i>Coefficient (Standard Error)</i>	<i>Odds Ratio</i>	<i>Coefficient (Standard Error)</i>	<i>Odds Ratio</i>
Post	0.164*** (0.025)	1.178	0.165*** (0.021)	1.180
Log distance, physician	− 0.033 (0.025)	0.968	− 0.069*** (0.019)	0.933
Log distance, specialist	− 0.027 (0.024)	0.973	− 0.037** (0.017)	0.964
Log distance, psychiatrist	− 0.007 (0.019)	0.993	0.010 (0.014)	1.010
Single	− 0.099*** (0.033)	0.906	− 0.109*** (0.024)	0.897
Relative to family				
Male	− 0.622*** (0.032)	0.537	− 0.645*** (0.023)	0.525
Ages 30–39	0.295*** (0.050)	1.343	0.298*** (0.031)	1.347
Relative to 18–29				
Ages 40–49	0.295*** (0.053)	1.343	0.309*** (0.034)	1.362
Relative to 18–29				
Ages 50–64	− 0.064 (0.059)	0.938	− 0.075* (0.040)	0.928
Relative to 18–29				
Nonwhite	− 0.527*** (0.041)	0.590	− 0.496*** (0.027)	0.609
Central	− 0.062 (0.121)	0.940	0.029 (0.078)	1.029
Relative to Northeast				
South	0.238** (0.119)	1.268	0.307*** (0.077)	1.360
Relative to Northeast				
West	0.330*** (0.122)	1.392	0.376*** (0.078)	1.456
Relative to Northeast				
Constant	− 2.489*** (0.137)		− 2.531*** (0.088)	
<i>N</i> (observations)		119,532		202,801
<i>N</i> (individuals)		29,883		72,991

*Coefficient is significant at 10%.

**Coefficient is significant at 5%.

***Coefficient is significant at 1%.

unrestricted sample for nonspecialist physicians and nonphysician specialists. Overall initiation appears to be least sensitive to the distance to psychiatrists, although this result could be because initiating with a psychiatrist is the least common occurrence. The other independent variables indicate that men, younger persons, persons with single coverage, and nonwhite individuals

are in general less likely to initiate mental health treatment, controlling for distance to provider.

In Table 6, we present the results of the multinomial logit regression of provider type conditional on initiation.⁴ As in the previous table, the regression includes measures of the distance to each provider type. We use the natural logarithm of distance to correct for the skewed nature of the distance measures. The regression results indicate that in the postperiod treatment initiators were significantly more likely to initiate treatment with a nonphysician specialist than with a nonspecialist physician and significantly less likely to initiate treatment with a psychiatrist than with a nonspecialist physician. The regression results indicate that conditional upon entry into mental health treatment, employees initiated treatment with a nonphysician specialist versus a nonspecialist physician at 1.9 times the rate. At the same time, employees initiated treatment with a psychiatrist versus a nonspecialist physician at roughly 0.6 times the rate. The relationship between initiation and distance to each type of provider confirmed our expectations, although only in the case of psychiatrist initiation was the relationship statistically significant: the greater the distance to psychiatrists, the less likely people were to initiate with them.

Table 6: Multinomial Logit Regression for Provider Type, Relative to Initiation at Nonspecialist Physician, Restricted Sample of Employees

	<i>Nonphysician Specialist</i>	<i>Psychiatrist</i>
Post	0.618*** (0.056) [1.855]	- 0.485*** (0.088) [0.616]
Log distance, nonspecialist physician	- 0.057 (0.047) [0.945]	- 0.019 (0.086) [0.981]
Log distance, nonphysician specialist	- 0.046 (0.045) [0.955]	0.116 (0.081) [1.123]
Log Distance, psychiatrist	- 0.023 (0.035) [0.978]	- 0.316*** (0.058) [0.729]

Sample size = 8,113. Regression also includes age brackets, gender, single versus family coverage indicator, nonwhite indicator, region indicators, and a constant.

*Coefficient is significant at 10%.

**Coefficient is significant at 5%.

***Coefficient is significant at 1%.

Odds ratios are in square brackets.

One concern with the design of the study is that employees may have been aware in 1995 of the impending benefit expansion. This could have led them to delay the onset of treatment until the postperiod. If this were the case, we would expect to see a one-time jump in initiation in 1996 and a subsequent reversion to preperiod levels (Sturm, Goldman, and McCullugh 1998). To test for this possibility, we estimated the same regression model as in (2) but with year dummies and the results indicated that initiation continued to rise over the 3-year period.

DISCUSSION

The results reveal that the combination of destigmatization and lower copayments appeared to be associated with a statistically significant and clinically meaningful increase in the probability of initiating mental health treatment. The probability of initiating with nonphysician mental health specialists increased dramatically, controlling for distance to provider. The provider network set up by the company intentionally included a large number of nonphysician specialists, as indicated on the provider lists made available through multiple means by the company. Because we are focusing on initiation, we believe that the provider choice effect is not likely because of the selectivity of the network. Rather, because the company publicized the list of network providers they helped induce individuals to use mental health specialists, rather than generalists, for the treatment of their disorders. We would expect a similar increase in initiation at nonphysician specialists if the company had eliminated financial incentives to visit within network providers, while still publishing the same list of providers. A change in the severity of illness is a possible explanation for the shift towards nonphysician specialists because the severity of the new initiators may be lower because of destigmatization. However, our analysis of diagnosis type does not reveal an increase in less-severe cases.

Additional work is needed to examine whether this change in company benefits was successful in reducing the company's mental health care costs and in increasing quality of care for persons with mental illness disorders (which were also the goals of their benefit change). However, our results do show that a combination of destigmatization and lowered copayments significantly increased the likelihood of treatment initiation. Furthermore, by including a large number of nonphysician specialists in the provider network, the company dramatically increased the probability of initiation at nonphysician specialists.

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NOTES

1. It should be noted that switching from coinsurance to copayments entails a relatively complex change in the behavioral response as measured by a change in the budget constraint. However, it is beyond the scope of this paper to model these complexities. Although throughout we refer to the reduction copayments, we actually are referring to the reduction in average cost-sharing required by patients.
2. We also attempted to define a referral measure by identifying whether individuals who initiated with a nonspecialist are observed to have a mental health specialist claim within 30 days of the start of the treatment episode. However, the rate of apparent referral was very low, and the analyses were not particularly informative.
3. We also estimated the model using a random effects logit and found that the results did not differ significantly from the simple logit with cluster-correction.
4. We use the standard Hausman test for the validity of the independence of irrelevant alternatives assumptions. The data do not reject the assumption.

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