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THE ECONOMICS OF GENDER AND MENTAL ILLNESS

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INTRODUCTION

Dave E. Marcotte and Virginia Wilcox-Gök

The past quarter-century has seen research on the economic impacts of mental illness flourish. Innovations in measurement and the release of several community-based and often nationally representative data sets containing valid and reliable diagnostic information have enabled researchers to make substantial advances in understanding the myriad ways that mental illness impacts the economic lives of the ill and their families. Among the most interesting and persistent findings in this literature is that mental illness affects women and men differently. Not only do women and men have very different rates of prevalence for various diseases, but mental illness is also commonly found to have different effects in their economic lives.

Research into the ways in which mental illness shapes the economic lives of the ill has identified important gender differences throughout the life cycle. Beginning during adolescence, mental illness among family members appears to have different effects on boys' and girls' educational attainment. During adulthood, women often withdraw from the labor market more quickly than do men when afflicted with mental illness. Among those who are employed, earnings losses due to mental illness are often found to vary by gender. Further, women and men undertake treatment for mental illness at different rates, and this may affect their ability to maintain productive lives as they adapt to their illnesses.

The dimensions along which the experiences of men and women with mental illness differ are numerous, as are the data, methods and economists who have studied the topic. While gender has so often been found to be an important determinant of prevalence and outcomes of mental illness, economists have rarely focused on gender differences as a central element of their analyses. In this volume, our aim is to direct the focus of research in the economics of mental health more squarely on

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the topic of gender. Each paper in this volume each provides insight into the ways in which women and men are afflicted and affected by mental illness in the labor market; highlighting both differences and similarities. It is our intention that by compiling a set of first-rate papers that focus on gender, we can facilitate a more complete understanding of the patterns and economic consequences of mental illness over the life course.

We hope that the research compiled herein, as individual papers and as a collection, will provide the reader with a richer understanding of the prevalence of mental disorders, the educational, employment and earnings impacts of psychiatric disease, and prospects for treating and providing access to health care for the mentally ill. Accordingly, we have organized this volume along just these lines: prevalence, effects of disease, and treatment and access to care.

The volume begins with a study by Kessler that reports on gender differences in patterns of mental illness in the United States from the National Comorbidity Survey (NCS). The NCS is the first nationally representative data set that allows estimation of 12-month and lifetime rates of prevalence of mental illness as well as the assessment of demographic, social, and economic correlates of prevalence. The NCS employed a modified version of the World Health Organization's Composite International Diagnostic Interview (CIDI) to make diagnoses using the criteria defined by the Diagnostic and Statistical Manual, 3rd edition, revised (DSM-III-R). The structured diagnostic interview is used by non-clinicians and has been shown to be a reliable and valid instrument for measuring mental illness. While Kessler finds small differences in the overall rates of prevalence of mental illness, he reports substantial differences in prevalence for nearly all specific illnesses. Most notably, women are substantially more likely to suffer from mood disorders and anxiety disorders, while men were more likely to have substance abuse and conduct and antisocial behavior disorders. Kessler reports that among the ill, women are much more likely than men to report seeking professional help for their psychiatric problems. He presents some evidence that this is because women can more readily perceive need for such help. Finally, Kessler notes that it is troubling that women have equally persistent and severe histories of mental illness, despite their higher rate of treatment. This leaves open important questions about gender differences in therapeutic efficacy.

The next paper focuses on understanding economic effects of mental illness at an early stage of the life cycle. Wilcox-Gök, Marcotte, Farahati and Borkoski point out that while we have come to understand the employment and earnings effects of mental illness fairly well, little is known about the impact of adolescent mental illness on educational attainment. Using the NCS data, Wilcox-Gök et al. estimate the effects of depression on the likelihood of dropping out of high school. They do so because high school dropout is known to be associated with substantial, and growing, earnings losses during a person's working life, and claim that this focus provides an indication of heretofore little examined *indirect* labor market costs of depression. To identify the effects of adolescent depression on dropout, Wilcox-Gök et al. focus on the sub-group with onset before the age of 16 (the age to which most states make attendance compulsory), dropping from the analysis those with onset between 16 and 18. Comparing these respondents to a reference group without early onset mental illness, the authors find that depression among boys is associated with a substantially elevated risk for high school dropout, but the same is not true for girls. While Wilcox-Gök et al. find that early onset of alcohol disorders increases the risk of high school dropout for both sexes, they report evidence that this cannot be attributed to the comorbidity between depression and alcohol disorders.

The paper by Alexandre, Fede and Mullings analyzes the effects of mental illness and labor supply. Making use of data from the first-ever wave of the National Household Survey on Drug Abuse to include measures of serious mental illness, the authors estimate effects of mental illness on unemployment and the number of workdays skipped among the employed. Alexandre et al. attempt to identify the labor supply effects of mental illness by exploiting measures of religiosity as instruments for serious mental illness. They find that serious mental illness substantially increases the risk of unemployment among men, but not women. The authors suggest this pattern is likely the result of women dropping out of the labor force altogether when afflicted with mental illness, not to the lack of an employment effect. This would be entirely consistent with previous research. Equally suggestive of a negative effect on labor supply, Alexandre et al. find that among those who remain employed, the seriously mentally ill miss more work days each month, likely because of the disabilities imposed by illness.

The next paper provides a unique look at the relationship between psychiatric illness and employment and earnings, as well as the likelihood of marriage (or co-habitation). Westergaard-Nielsen, Agerbo, Eriksson and Mortenson use data from the Danish Longitudinal Labor Market Register, which is a compilation of various administrative data containing information on income and unemployment, among other things, for a 5% sample of the Danish adult population. The data provide observations on more than 200,000 persons, and tracks them for nearly 2 decades, from 1976 to 1993. The authors have merged these data with an administrative data set covering all admissions into psychiatric inpatient hospitals in Denmark; providing information on length of stay and diagnosis. Westergaard-Nielsen et al. make use of a matched control group design to examine patterns of employment and marriage before and after a psychiatric hospital admission compared to pattern for observationally similar but healthy peers. The authors find that the mentally ill suffer employment and earnings losses many years prior to admission. They find no substantial differences between men and women in the likelihood of maintaining employment, though they find that men suffer relatively large earnings effects. The authors suggest this may be due to differences in the extent to which jobs accommodate or are impeded by mental illness. Other possibilities include gender differences in disease or severity, or access to treatment. Westergaard-Nielsen et al. also find that differences between cases and controls grow most rapidly just prior to admission, but then stabilize in the years following admission. A hopeful interpretation of this last finding is that treatment prevents further decay in the economic and social lives of the ill. On the other hand, the ill never recover their pre-admission levels of employment and earnings vis-à-vis controls.

The next paper helps provide some context on employment transitions and mental health that, among other things, is useful for interpreting the Alexandre et al. and Westergaard-Nielsen et al. papers. In this paper, Gresenz and Sturm make use of the Healthcare for Communities data to analyze the relationship between mental health and transitions between employment, unemployment and being out of the labor force. The authors exploit panel features of the data to examine employment transitions between waves and their relationship with recent mental health episodes as measured at the initial interview. Gresenz and Sturm find interesting and important differences between the employment transitions of women and men subsequent to an episode of mental illness. Most notably, they find that women recently suffering an episode associated with depression or an anxiety disorder are more likely to drop out of the labor force altogether. Interestingly they are also more likely to rejoin the labor force if they were out of the labor force. For men, however, there is not this pattern of exiting the labor force. Nonetheless, Gresenz and Sturm find that men with anxiety or depressive disorders are less likely to find employment.

In the final paper in the volume to examine employment consequences, Salkever, Slade and Karakus examine gender differences in employment patterns and transitions among persons with schizophrenia. They analyze data from a large scale survey conducted at six sites around the United States: the Schizophrenia Care and Assessment Program (SCAP). SCAP is a longitudinal study, so participants were interviewed at six month intervals over a period of up to three years. Not surprisingly, employment rates and employment stability are quite low among those with schizophrenia compared to the general population. Importantly, Salkever et al. find that among those with schizophrenia, women are more likely to have long-term periods of non-employment, and are more likely to lose a job if employed. The authors find that among schizophrenics, education has a larger positive effect on the likelihood of employment for women. But, women's employment chances are also more negatively affected by negative symptoms of disease. Salkever and his colleagues speculate that this may be related to gender differences in the types of jobs those with schizophrenia hold: perhaps women hold jobs that require interpersonal rather than manual skills, and hence negative symptoms impose more important limitations on their job performance, on average. While the authors do not have the data to test this hypothesis, their results serve to both answer and raise important questions about the ways in which women and men suffering from debilitating mental illness experience the labor market differently.

Each of these papers on labor market impacts of mental illness confirm some important findings of previous research, and each raises important questions, too. A particular value of these papers is that they each utilize new or novel data to examine this relationship, thereby contributing important new perspective into the existing body of knowledge.

The next three papers also make use of unique and excellent data, and provide unusual insight into the relevance of gender in mental health services utilization, re-employment after disability leave, and also access to health insurance among poor, depressed women. The first of these papers makes use of data from a multinational U.S. communications and electronics company, employing around 75,000 workers in the U.S. that implemented a change in mental health benefits. In this paper, LoSasso, Lindrooth and Lurie describe a change in health insurance benefits at this company that reduced co-payments for mental health treatment, provided employees access to a network of mental health providers which included non-physician specialists, and adopted efforts to de-stigmatize mental health treatment. The authors found significant increases in the likelihood that employees initiated mental health treatment subsequent to the benefit change, and most of this was the result of increases in the use of non-physician mental health specialists. LoSasso et al. found no differences in the way women and men responded to the benefit change. While women were more likely to initiate mental health treatment prior to and after the benefit change, both men and women increased their rates of mental health care utilization subsequent to the change in the company's mental health benefit package and de-stigmatization effort. Taken as a whole the authors' findings provide cause for optimism that both working men and women can be encouraged to seek mental health treatment.

In the next paper, Dewa, Hoch and Goering examine gender differences in outcomes from short-term depression-related disability leave. The authors make use of administrative records from three major financial companies in Canada, together employing some 63,000 worker (about 12% of Canadian workers in the industry). Dewa et al. compiled data from each company's short-term disability claims, prescription drug claims and occupational health department records. The authors used the data to examine how work experience, occupation, depression symptoms, recommended clinical guideline use of antidepressants along with gender affected two key outcomes: use of anti-depressant medication during the disability episode, and; return to work at the end of the disability episode. They find that while female employees were much more likely to incur at least one depression-related short-term disability episode in a year, men remained on disability longer, and were less likely to return to work In addition, men were significantly more likely to use anti-depressant pharmacotherapy than were women. While men were found to go on short-term depression-related disability less often than women, Dewa and her colleagues conclude that they comprise an important group, because they were more likely to be managers and more experienced. This translates into relatively large losses of human capital and friction costs for employers.

In the final paper, Roberts examines the relationships between access to health insurance and income and depression for women. Her study is motivated by the welfare reform of 1996 that resulted in a substantial decline in the welfare roles and raised concern that many women would lose health insurance coverage. Roberts makes use of data from the 1999 National Health Interview Survey (NHIS) and an adult supplement that includes the CIDI – Short Form. Because the NHIS data are so rich, she is able to estimate a series of multinomial probit models to estimate the effect of major depressive disorder and low income on the likelihood a woman has no health insurance, compared to public or privately financed insurance, controlling for a host of other attributes. Roberts finds that like having low income, depression increases the risk a woman has no insurance coverage. Though here data do not permit her to sort out the causal pathways here, she finds that depression among low income women does not pose any greater risk of lacking insurance than does depression among more affluent women.

In this introduction, we have been able to provide only the briefest of reviews of the very rich findings contained in this volume's papers. Collectively, these papers advance our understanding of the importance of gender in shaping the prevalence of mental illness; it's consequences in the labor market; patterns of use of mental health services; disability leave; and, the importance of depression in insurance coverage for women. Not only is the centrality of gender in each of these papers an important contribution, but these papers employ data sets that make new and unique contributions to the field of mental health economics. It is our hope that as a volume, these papers can help researchers and policy analysts develop a more integrated and complete understanding of the many ways that mental illness affects the economic lives of women and men.

GENDER DIFFERENCES IN MENTAL DISORDERS IN THE U.S. NATIONAL COMORBIDITY SURVEY

Ronald C. Kessler

This chapter presents a broad overview of data from the U.S. National Comorbidity Survey (NCS) (Kessler et al., 1994) on gender differences in the prevalences and correlates of commonly occurring mental disorders, including mood disorders, anxiety disorders, substance use disorders, and antisocial personality disorder. The NCS is the first nationally representative survey in the United States to administer a structured psychiatric interview to a large sample of the general population, providing a unique opportunity to investigate these gender differences.

Previous epidemiological surveys have consistently found that women are more likely to have anxiety and mood disorders than men, while men are more likely to have substance use disorders and antisocial personality disorder than women. These basic patterns have been found consistently in community epidemiological studies, using a variety of diagnostic schemes and interview methods (Bebbington, 1988; Nolen-Hoeksema, 1987; Weissman & Klerman, 1992) in the United States (Robins & Regier, 1991b) as well as in such other countries as New Zealand (Wells, Bushnell et al., 1989) Canada (Bland, Newman et al., 1988; Bland, Orn et al., 1988) Taiwan (Hwu et al., 1989) and Puerto Rico (Canino et al., 1987). We consequently expected similar patterns in the NCS.

Previous psychiatric epidemiological surveys have been less concerned with the consequences of mental disorders, although interest in consequences has grown in the past decade. Mental disorders have been shown in these recent studies to have substantial personal and social costs (Kouzis & Eaton, 1994) with impairments as

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great as those associated with serious chronic physical illnesses (Wells, Stewart et al., 1989). Mental disorders have also been linked to substantially reduced quality of life (Wohlfarth et al., 1993) and impaired work role functioning (Kessler & Frank, 1997). The analysis of gender differences has not played a prominent role in these investigations, though, although there is reason to believe that the implications of mental disorders could be quite different for women than men.

Prior to the NCS, the most recent evidence of gender differences in treatment of mental disorders in the U.S. came from the Epidemiologic Catchment Area (ECA) Study (Robins & Regier, 1991b). The ECA surveyed more than 20,000 people in five U.S. communities between 1980 and 1985 and estimated the prevalence of specific DSM-III disorders and past-year use of services for these disorders. The ECA data suggested that women with recent mental disorders were significantly more likely than comparable men to obtain treatment (Robins & Regier, 1991a), which is consistent with virtually all previous U.S. evidence on gender differences in mental health service use (Kessler et al., 1981; Leaf & Bruce, 1987). This was not due to women having more psychiatric problems than men, however, as the ECA investigators and others before them found that the gender difference in use persisted after adjusting for differential need for services. Nor is it plausible to argue that women have more ready access to services than men, as available evidence clearly shows that both financial and non-financial barriers to care are greater for women than men. Although it is not possible to evaluate all plausible alternative possibilities here, some data relevant to reasons for the gender differences in mental health service use are reviewed in this chapter.

METHODS

The Survey

The NCS was based on multistage area probability sample of people age 15–54 in the non-institutionalized civilian population of the coterminous U.S., with a supplemental sample of students living in campus group housing (Kessler et al., 1994). The survey was carried out face-to-face in the homes of respondents by professional interviewers employed by the Survey Research Center (SRC) at the University of Michigan. The survey was carried out between September 1990 and February 1992. The response rate was 82.6%, with a total of 8,098 respondents participating in the survey. A supplemental non-response survey was carried out to adjust for non-response bias, with a random sample of initial non-respondents offered a financial incentive to complete a short form of the diagnostic interview. A non-response adjustment weight was constructed for the main survey data to compensate for elevated rates of disorders found among the initial non-respondents in this non-response survey. Significance tests were made using design-based methods because of this weighting and clustering of the data (Kish & Frankel, 1970; Koch & Lemeshow, 1972; Woodruff & Causey, 1976). More details about design and weighting procedures are reported elsewhere (Kessler, Little et al., 1995).

Diagnostic Assessment

Diagnoses were made in the NCS using the definitions and criteria of the American Psychiatric Association's Diagnostic and Statistical Manuel of Mental Disorders, Third Edition, Revised (DSM-III-R). The interview used to generate these diagnoses was a modified version of the World Health Organization's Composite International Diagnostic Interview (CIDI) (World Health Organization, 1990) a state-of-the-art structured diagnostic interview designed to be used by trained interviewers who are not clinicians (Robins et al., 1988). A number of modifications were introduced into the CIDI in the NCS in order to improve accuracy of response. These modifications are discussed elsewhere (Kessler et al., 1999).

The DSM-III-R diagnoses assessed in the NCS included mood disorders (major depression, mania, dysthymia), anxiety disorders (panic disorder, generalized anxiety disorder, simple phobia, social phobia, agoraphobia, post-traumatic stress disorder), substance use disorders (alcohol abuse, alcohol dependence, drug abuse, drug dependence), conduct disorder, adult antisocial behavior, antisocial personality disorder (the conjunction of conduct disorder and adult antisocial behavior), and non-affective psychosis (NAP; a summary category made up of schizophrenia, schizophreniform disorder, schizoaffective disorder, delusional disorder, and atypical psychosis).

As described in more detail elsewhere, clinical reappraisal interviews with NCS respondents showed that all but one of the CIDI diagnoses were assessed with acceptable reliability and validity, with no evidence of gender differences in reliability or validity (Kessler et al., 1998).The exception was NAP, which was not accurately assessed in the CIDI (Kendler et al., 1996). Because of this inaccuracy, no results regarding NAP are reported in the current chapter. This failure to assess NAP accurately is consistent with similar difficulties in other community epidemiological surveys, which have consistently involved systematic over-estimation of prevalence due to a large proportion of the population endorsing questions that were designed to assess delusions and hallucinations.

RESULTS

Prevalence

As shown in Table 1, 48.5% of the women in the NCS and 51.2% of the men reported a history of at least one of the lifetime (LT) DSM-III-R disorders assessed in the survey ($z = 1.0$, $p = 0.294$), while 32.3% of women and 29.4% of men

Table 1. Lifetime and 12-Month Prevalences of DSM-III-R Disorders in National Comorbidity Survey by Gender.[a]

Disorders[a]	Women				Men			
	Lifetime		12-Month		Lifetime		12-Month	
	%	se[b]	%	se[b]	%	se[b]	%	se[b]
I. Mood disorders								
Major depressive episode	21.3[c]	0.9	12.9[c]	0.8	12.7	0.9	7.7	0.8
Manic episode	1.7	0.3	1.3	0.3	1.6	0.3	1.4	0.3
Dysthymia	8.0[c]	0.6	3.0	0.4	4.8	0.4	2.1	0.3
Any mood disorder	23.9[c]	0.9	14.1[c]	0.9	14.7	0.8	8.5	0.8
II. Anxiety disorders								
Panic disorder	5.0[c]	1.4	3.2[c]	0.4	2.0	0.3	1.3	0.3
Agoraphobia without panic	7.0[c]	0.6	3.8[c]	0.4	3.5	0.4	1.7	0.3
Social phobia	15.5[c]	1.0	9.1[c]	0.7	11.1	0.8	6.6	0.4
Simple phobia	15.7[c]	1.1	13.2[c]	0.9	6.7	0.5	4.4	0.5
Generalized anxiety disorder	6.6[c]	0.5	4.3[c]	0.4	3.6	0.5	2.0	0.3
Posttraumatic stress disorder	10.1[c]	0.8	5.4[c]	0.7	4.8	0.6	2.3	0.3
Any anxiety disorder	34.3[c]	1.8	24.7[c]	1.5	22.6	1.2	13.4	0.7
III. Addictive disorders								
Alcohol abuse	6.4	0.6	1.6	0.2	12.5[d]	0.8	3.4[d]	0.4
Alcohol dependence	8.2	0.7	3.7	0.4	20.1[d]	1.0	10.7[d]	0.9
Drug abuse	3.5	0.4	0.3	0.1	5.4[d]	0.5	1.3[d]	0.2
Drug dependence	5.9	0.5	1.9	0.3	9.2[d]	0.7	3.8[d]	0.4
Any substance disorder	17.9	1.1	6.6	0.4	35.4[d]	1.2	16.1[d]	0.7
IV. Other disorders								
Antisocial personality disorder	1.0	0.2	–	–[e]	4.8[d]	0.5	–	–[e]
Nonaffective psychosis	0.7	0.2	0.4	0.1	0.3	0.1	0.2	0.1
V. Any NCS disorder	48.5	2.0	32.3	1.6	51.2	1.6	29.4	1.0

[a] Disorders are defined without DSM-III-R hierarchy rules.
[b] se = standard error.
[c] Significantly higher prevalence among women than men at the 0.05 level (two tailed test).
[d] Significantly higher prevalence among men than women at the 0.05 level (two tailed test).
[e] Antisocial personality was only assessed on a lifetime basis.

reported at least one 12-month disorder ($z = 1.5$, $p = 0.124$). While gender differences in these overall proportions are not large, there are significant differences in the prevalence of almost all specific disorders. Consistent with previous studies, women were much more likely than men to have mood and anxiety disorders, while men were much more likely than women to have substance use disorders, conduct disorder, and adult antisocial behavior.

There is some indication from other studies that gender differences in mental disorder have declined in recent years (Kessler & McRae, 1981). We examined this possibility by asking respondents when they first experienced each disorder and by comparing these reports across sub-samples defined by age at the time of interview. By focusing on the same recalled ages, we were able to produce retrospective age of onset curves for each sample cohort to see whether there has been change over time and then comparing female vs. male patterns.

Figure 1 presents an example of these cohort-specific curves. This example focuses on the cumulative LT prevalence of a major depressive episode among women in the NCS. Lifetime prevalence is higher at all ages in successively younger cohorts. The pattern is roughly similar to those in Fig. 1 for most other NCS disorders, suggesting that mental disorders have become more common over the half century spanned by the lives of NCS respondents. It is important to note, though, that age-related biases of various sorts could also explain the pattern (Giuffra & Risch, 1994).

The more important question for this chapter is whether this increase is associated with a change in the Female:Male (F:M) LT prevalence ratio. The answer is presented in Table 2 for major depressive episode. As shown there, an elevated F:M ratio is observed by age 14 in the youngest cohort and by age 9 in the next two older cohorts, but not until age 24 in the oldest cohort. The F:M ratio does not differ greatly across cohorts by age 24, the oldest age at which we can compare all four NCS cohorts. The ratio tends to be larger at ages older than 24 in successively older cohorts, suggesting that risk of depression onset during the middle years of life has become smaller in more recent cohorts. The similarity across cohorts is particularly striking in light of the suggestion in Fig. 1 that the LT prevalence of depression increased roughly five-fold across these cohorts.

Summary results are presented in Table 3 of parallel analyses for the other NCS/DSM-III-R disorders. A tendency for the F:M ratios to be fairly stable over the age range of respondents is clear in most cases. However, cohort differences in the ratios are apparent for three male-dominated disorders: alcohol dependence, drug dependence, and antisocial behavior. These differences are all associated with a decrease in the F:M ratio in recent cohorts, which means women are catching up to men in the prevalence of these disorders. An inspection of unconditional cumulative prevalence by cohort (results not shown) shows that

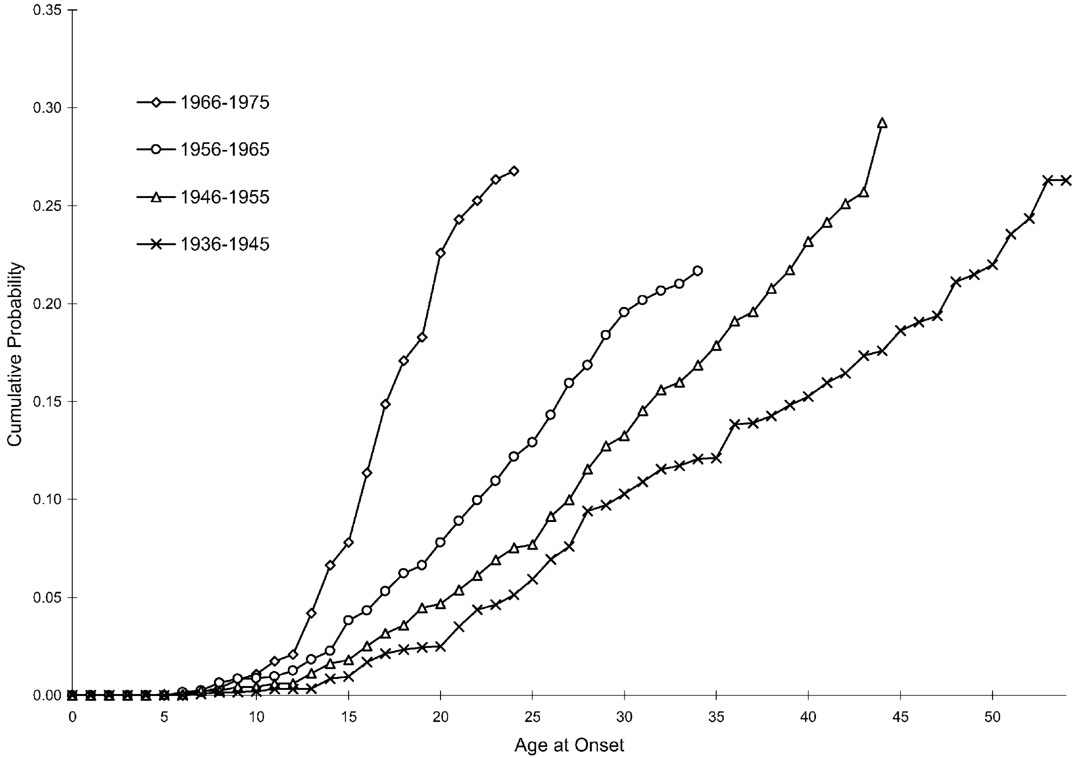

Fig. 1. Cohort Differences in the Cummulative Age of Onset of Major Depressive Episode in Women.

Table 2. Female: Male Rate Ratios for Lifetime Major Depressive Episode in National Comorbidity Survey by Age of Onset.

Age of Onset	Cohort							
	1966–1975		1956–1965		1946–1955		1936–1945	
	OR	(95% CI)	OR	(95% CI)	OR	(95% CI)	OR	(95% CI)
9	0.7	(0.2, 1.9)	1.9	(0.4, 7.9)	1.7	[a]	0.2	[a]
14	2.0[b]	(1.0, 3.8)	2.4[b]	(1.3, 4.3)	1.5	(0.8, 3.2)	0.6	(0.1, 2.8)
19	2.0[b]	(1.3, 3.2)	1.6[b]	(1.2, 2.3)	1.5	(0.9, 2.6)	0.8	(0.3, 2.2)
24	1.9[b]	(1.3, 2.9)	1.8[b]	(1.4, 2.3)	1.6[b]	(1.1, 2.4)	1.6	(0.7, 3.4)
29			1.6[b]	(1.3, 2.1)	1.7[b]	(1.2, 2.4)	2.6[b]	(1.5, 4.5)
34			1.5[b]	(1.2, 2.0)	1.9[b]	(1.4, 2.5)	2.8[b]	(1.7, 4.5)
39					1.8[b]	(1.4, 2.3)	2.0[b]	(1.3, 3.0)
44					1.7[b]	(1.4, 2.2)	1.8[b]	(1.2, 2.7)
49							1.9[b]	(1.2, 2.8)
54							2.0[b]	(1.3, 3.0)

[a] Low precision.
[b] Significant at the 0.05 level (two-tailed test).

this narrowing is due to alcoholism becoming more prevalent in recent cohorts of women rather than less prevalent among men. In fact, the cohort-specific cumulative prevalence curves show that alcoholism has increased among men as well in recent cohorts, but that the increase has been much greater among women, accounting for the increase in the F:M ratio.

A similar pattern was found for antisocial behavior, where the significant cohort effect is due to a greater proportional increase among women than men in recent cohorts. The change in the F:M ratio for drug dependence, however, is more complex. The NCS cohort data show that drug dependence was much more common among women than men before the widespread use of recreational drugs among youth in the early 1970s. Most drug dependence prior to this time was associated with use of prescription medications (e.g. valium) and was largely a problem of middle-age women. The situation changed radically in the 1970s, however, with the introduction of marijuana into youth culture. Drug use began to be associated with youthful experimentation and, like alcohol use in earlier generations, became much more of a male than female behavior. As a result, the F:M ratio for drug dependence changed quite dramatically from a strong female preponderance (3.0) among people in their late 20s in the 1936–1945 NCS cohort to a strong male preponderance (0.4) in the 1946–1955 cohort. The ratio has remained below 1.0 since that time. We also see a secondary trend in recent cohorts similar to the trend described above for alcohol dependence – young women are beginning to catch up to men in their rates of

Table 3. Change in the Relationships Between Gender and Lifetime DSM-III-R Disorders in the National Comorbidity Survey Across the Life Span and Over Cohorts.[a]

Disorders	Age Effects[b]		Cohort Effects
	OR	95% CI	X_3^2
Mood disorders			
Major depressive episode	0.9	0.8–1.1	1.0
Manic episode	0.7	0.5–1.1	5.6
Dysthymia	1.1	0.9–1.3	0.4
Anxiety disorders			
Panic disorder	1.0	0.8–1.4	3.7
Agoraphobia without panic	1.1	0.9–1.4	6.3
Social phobia	0.9	0.7–1.2	0.2
Simple phobia	0.9	0.7–1.1	2.5
General anxiety disorder	1.1	0.8–1.3	2.6
Posttraumatic stress disorder	0.7[c]	0.6–0.9	5.7
Substance use disorders			
Alcohol abuse	1.1	0.8–1.3	5.3
Alcohol dependence	1.0	0.9–1.2	10.2[c]
Drug abuse	0.9	0.7–1.2	3.1
Drug dependence	1.3[c]	1.1–1.6	8.1[c]
Other disorders[d]			
Conduct disorder	–		3.2
Adult antisocial behavior	–		11.5[c]

[a] The results are based on a series of discrete-time survival models in which sex, age (person-year), cohort, and their interactions were used to predict onset of each disorder. The ORs associated with the age effect are interactions between age and sex in predicting the outcomes. The x^2 values associated with the cohort effect are x^2 differences associated with the introduction of interactions between sex and three dummy variables (to define the four decades of birth in the NCS cohorts from 1936–1945 through 1966–1975) in predicting the outcomes.
[b] Age effects are reported as the relative-odds associated with a decade of life.
[c] Significant at the 0.05 level (two-tailed test).
[d] Conduct disorder is the childhood component of antisocial personality disorder, while adult antisocial behavior is the adult component. A significant cohort effect was found only for the second of these two components. Age effects were not assessed because these disorders were evaluated only on a lifetime basis in the NCS. We did not include nonaffective psychosis in this analysis because the small number of cases of this disorder made it impossible to estimate the interaction effects.

illicit drug dependence. Comparing the three ten-year NCS cohorts born after World War II during their late teens, we see that an initial strong preponderance of drug dependence among men (0.2) in the 1946–1955 cohort decreased meaningfully in the 1956–1965 cohort (0.3), and even more in the 1966–1976 cohort (0.6).

Course

As shown in Table 1, men and women differ not only in whether they have ever been depressed or anxious or alcoholic, but also in their likelihood of suffering from these conditions recently. Between one-third and one-half the people who had ever experienced these disorders still had them during the year prior to the interview. These high percentages could be an artifact, as recency of LT disorder may be positively related to probability of recall. However, if not artifactual, they could reflect either higher rates of chronicity, high rates of recurrence, or both.

An investigation of chronicity is important because some theories of gender differences in anxiety and depression emphasize the importance of differences in chronicity and recurrence. Sex role theory, for example, suggests that the chronic stresses and lack of access to effective coping resources associated with traditional female roles lead to the higher rates of chronic anxiety and depression found among women (Barnett & Baruch, 1987; Mirowsky & Ross, 1989), while rumination theory suggests that women are more likely than men to dwell on problems and, because of this, to let transient symptoms of dysphoria grow into clinically significant episodes of depression (Nolen-Hoeksema, 1987). The gross associations concerning the ratios of 12-month to LT disorders reported in Table 1 do not support this view, as either greater chronicity or greater recurrence would create a significant association between gender and 12-month prevalence. No systematic evidence of such an association exists in the data.

The possibility of gender differences in course should not be crossed off too quickly, though, as the results mentioned in the last paragraph fail to adjust for the possibility of gender differences in either age of onset or time since onset. A more formal analysis is presented in Table 4, where we examine the F:M ratio of 12-month prevalence among people with an LT disorder, controlling for age of onset and years since onset. No evidence for a gender difference in the course of mood disorders was found. The situation is different, however, for anxiety disorders and substance use disorders, with 12-month anxiety disorders appearing to be more prevalent among women (F:M ratios greater than 1.0) and substance use disorders more prevalent among men with LT disorders. These results have to be considered tentative, however, as the data are cross-sectional.

Life Course Consequences

The effects of early-onset mental disorders were examined in predicting three important subsequent life course role transitions: educational attainment, teenage childbearing, and marriage. This was done by using retrospectively reported age of

Table 4. Gender Differences (Female:Male) in Persistence[a] of Disorders in the National Comorbidity Survey.

Disorders	OR	95% CI
Mood disorders		
Major depression episode	1.0	0.7–1.4
Manic episode	0.5	0.2–1.3
Dysthymia	0.8	0.5–1.3
Anxiety disorders		
Panic disorder	1.0	0.5–2.3
Agoraphobia without panic	1.3	0.7–2.5
Social phobia	1.0	0.8–1.4
Simple phobia	3.0[b]	1.9–4.6
Generalized anxiety disorder	1.6[b]	1.0–2.5
Posttraumatic stress disorder	1.1	0.6–1.9
Substance use disorders		
Alcohol abuse	1.0	0.6–1.5
Alcohol dependence	0.6[b]	0.4–0.9
Drug abuse	0.4[b]	0.2–0.7
Drug dependence	0.6[b]	0.3–1.1

[a] Persistence is defined as 12-month prevalence among respondents with a lifetime history of the disorder and age of onset more than one year prior to the time of interview. The ORs are the effects of gender (female:male) in a multivariate logistic regression equation controlling for age of onset and number of years since onset.
[b] Significant at the 0.05 level (two-tailed test).

onset reports to classify respondents according to the existence of mental disorders prior to the age of the relevant transition and then to use this information to predict the transitions. Control variables were included for socio-demographics and a variety of childhood adversities that could predispose both to early onset of mental disorders and the role of transitions. More details on the analysis methods are reported elsewhere (Kessler et al., 1997; Kessler & Forthofer, 1999; Kessler, Foster et al., 1995). Early-onset disorders were found to be powerful predictors of failure to complete high school, failure to go on to college among high school graduates, and failure to graduate from college among college entrants (Kessler, Foster et al., 1995). However, no significant gender differences in these effects were found. The analysis of teenage childbearing (for women) and fathering (for men) was carried out in a similar way. As reported in more detail elsewhere (Kessler et al., 1997), early-onset mental disorders were found significantly to predict these outcomes, with only one significant gender differences (a greater effect of conduct disorder among men than women). The analysis of marital timing, finally, documented that people with child and adolescent mental disorders are more likely than others to marry by age 18 and that mental disorders are a handicap in the marriage market among people who do not marry by the age of 18. As in analyses of the other role transitions, no consistent pattern of gender differences was found.

Treatment

A significantly higher proportion of women (15.5%) than men (10.8%) reported use of services for mental health problems in the 12 months before the NCS interview ($z = 2.9$, $p = 0.004$), a pattern that held for each of the different types of professional treatment considered in the survey. Women and men did not differ, in comparison, in their use of self-help groups (3.2% of women and 3.2% of men). We also looked at the average number of visits among respondents in treatment for emotional problems in each service sector by type of disorder. The average person in ambulatory care reported making 16 visits for treatment of emotional problems during the 12 months prior to the interview. Average number of visits differed substantially across service sectors and was much higher in the sub-sample of people who met criteria for one of the NCS/DSM-III-R disorders than among those who did not. However, only one case out of the many examined showed a statistically significant gender difference in average number of visits: men with no disorder who participated in self-help groups attended a significantly larger number of visits, on average, than comparable women (42.5 v 16.2, $z = 2.5$, $p = 0.012$). Further analysis showed that this reflects a much higher use of Alcoholics Anonymous and other self-help groups among men than women with a history of substance problems.

Previous research has suggested that gender differences in rates of seeking treatment might be due to women being more likely than men to perceive themselves as needing help even when they have the same level of measured need as men (Kessler et al., 1981). We investigated this issue by creating a “perceived need” score that combined the people who reportedly sought treatment because of perceived need (as opposed to treatment sought under pressure from loved ones or legal authorities) with the people who reportedly felt that they needed help at some time during the past year but did not get it. A strong monotonic relationship was found between a summary measure of objective disorder severity and this measure of perceived need among both women and men. Consistent with the hypothesis, we found that women had a significantly higher level of perceived need than men at each level of severity.

Importantly, no significant gender differences were found in patterns of help-seeking for emotional problems among women and men with the same objective disorder severity who perceived themselves as needing treatment. This finding suggests that perceived need mediates the relationship between gender and help-seeking. It is important to note that the measure of perceived need was not simply a close proxy for service use, as only 41.2% of the women with perceived need and 44.4% of the men with perceived need obtained treatment. Perceived need was a critical mediator, however, as only very small proportions of women (1.2%) and men (1.1%) who did not think of themselves as needing treatment obtained help. (All of these reported that they sought treatment involuntarily either because of court order or because of a close loved one pressuring them to see a professional about their emotional problems.) Based on these results, it appears that the greater perception of need among women than men explains the higher rate of female treatment for emotional problems.

DISCUSSION

Prevalence and Course

As noted in the introduction, many previous epidemiological surveys have found that women are more likely to experience mood and anxiety disorders and that men are more likely to experience substance use and antisocial disorders. The NCS replication of these results was not unexpected. It was somewhat more interesting to find that these differences did not change greatly over the age range of the sample, especially in light of the fact that gender differences in recall bias about mental disorders has been reported in longitudinal research (Ernst & Angst, 1992; Wilhelm & Parker, 1994).

Differential recall bias might also be implicated in the NCS results regarding gender differences in the course of mental disorders. However, it is unlikely that any such bias is so severe that true gender differences in the course of these disorders are more pronounced than differences in risk of initial onset. The significant cohort effects documented for most of these disorders separately among women and men could be due to recall bias; indeed, simulations suggest that this is quite plausible (Giuffra & Risch, 1994). To the extent that evidence of cohort effects is real, however, it strongly suggests that environmental factors are responsible and presumably account for the gender differences. We also know from twin and adoption studies that there are powerful genetic effects on most of these disorders (Cadoret, 1991; Kendler et al., 1995), suggesting that some combination of environmental and genetic influences is at work, including genetic influences on sensitivity to environmental effects.

Other Disorders

While the presentation in this chapter focused on anxiety, mood, substance, and antisocial personality disorders, at least three other important broad classes of mental disorders are important to mention: non-affective psychoses, dementias (most notably Alzheimer's disease), and impulse-control disorders (e.g. intermittent explosive disorder, attention-deficit hyperactivity disorder, pathological gambling disorder).

As noted in the section on diagnostic assessment, little is known about the general population prevalence or correlates of non-affective psychosis due to the fact that the CIDI and other fully structured assessment measures are unable to measure NAP with good accuracy. As a result, our knowledge about the epidemiology of NAP is based largely on studies carried out in clinical samples (Bromet et al., 1992). These studies suggest that there is no significant gender difference in either lifetime risk or course of NAP, but that the average age of onset of NAP is five years earlier in men (beginning in adolescence) than women (beginning in early adulthood) (Hafner, 2003). The age-of-onset difference appears to be associated with gender differences in normal age-related maturational changes (Spauwen et al., 2003). Estrogen seems to be implicated here, as indicated both by changes in illness severity among post-menopausal women and by positive effects of estrogen therapy on course of illness (Riecher-Rossler, 2002). Gender differences that vary with menopausal status have also been documented in response to novel anti-psychotic medications (Goldstein et al., 2002). There has been much less research on whether socio-cultural factors play a part in gender differences in the course of psychosis, although a number of provocative ideas about such effects have been proposed (Nasser et al., 2002).

Dementia was omitted from the NCS diagnostic assessment because Alzheimer's disease (AD) and related dementias are very rare in the age range of the NCS sample. Epidemiological surveys of older populations consistently show that AD is more prevalent among women than men (Hauser & Amatniek, 1998). As women live longer than men, though, it is important to evaluate age-related incidence as well as prevalence. There is controversy about incidence in the AD literature, with some reviews suggesting that incidence becomes higher among women than men at older ages and others failing to find a gender difference in incidence (Jorm & Jolley, 1998). There is also controversy regarding gender differences in the severity of AD, with some studies suggesting that the cognitive deficits caused by AD are more pronounced among women than men (Buckwalter et al., 1996) and others finding no such gender differences (Bayles et al., 1999).

These inconsistencies appear to be due, at least in part, to gender-related selection bias into studies of AD (Bonsignore et al., 2002).

Epidemiological research on impulse control disorders is scant. We know that patients with these disorders, despite quite variable presentations (e.g. pyromania, kleptomania, pathological gambling, intermittent explosive disorder, attention-deficit, hyperactivity disorder), typically have an early age of onset, a chronic course, and serious impairments in role functioning (Hollander & Rosen, 2000). We also have good indirect evidence to suggest that some impulse control disorders, such as intermittent explosive disorder (Coccaro, 2000; Olvera, 2002) and adult attention-deficit hyperactivity disorder (Wender et al., 2001), are fairly common in the general population, with cumulative lifetime prevalence across these disorders likely to be as high as 10–15%. Furthermore, the available evidence suggests that the most frequently occurring impulse control disorders are more prevalent among men than women (Arnold, 1996; Coccaro, 2000; Hollander & Rosen, 2000). However, there is some evidence from treatment studies to suggest that impulse control disorders are often more persistent and severe among women than men (Arcia & Conners, 1998; Tavares et al., 2001).

Social Consequences

The mental disorders assessed in the NCS are significantly associated with adverse outcomes in educational attainment, age at first childbearing, probability of ever marrying, and marital timing. Although the NCS cross-sectional research design does not permit these associations to be interpreted unequivocally as causal, such an interpretation is plausible. Although we found no systematic gender differences in these associations, some outcomes are more eventful for women than men. This is especially true for teenage childbearing. Given the profile of effects documented in the NCS, in fact, we would expect early-onset mental disorders to be significantly associated with welfare dependency among women. A number of recent longitudinal welfare-to-work studies have shown that this is the case and that mental disorders significantly interfere with the transition from welfare to work (Dooley & Prause, 2002; Montoya et al., 2002).

Treatment

The NCS data document that women are significantly more likely than men to seek professional help for their psychiatric problems. This is quite striking in light of the fact that women have a relative disadvantage compared to men in access to health insurance. It appears that there is a lower severity threshold for seeking help among women than men. A similar gender difference in the threshold for help-seeking has been documented for physical illness as well (Davis, 1981; Kessler, 1986). One plausible explanation for this difference is that women might be more sensitive to bodily symptoms than men and thereby more likely to perceive need and seek help (Mechanic, 1979). However, the empirical evidence fails to document a strong difference of this sort (Kessler, 1986). A second plausible explanation is that women are socialized to more readily adopt the sick role than men. Numerous beliefs and attitudes have been proposed along these lines (Mechanic, 1976; Nathanson, 1977). Even though not all have been examined empirically, the available evidence is consistent with this hypothesis. For example, there is evidence that women have a greater tendency than men to consult physicians for hypothetical physical health problems (Cleary et al., 1982) and to believe in the efficacy of medical intervention (Depner, 1981).

It is sobering that women do not have a less persistent or severe course of mental illness than men despite their higher treatment rate. One possible explanation for this is that treatment is less effective for women than men, although research on treatment of mental disorders has shown little evidence for such a difference (Thase et al., 1998). A more plausible possibility is that women receive less adequate treatment than men due to the fact that they are more likely than men to receive their treatment from primary care doctors rather than from mental health specialists. It is very difficult to evaluate this possibility with non-experimental data due to a significant association between illness severity and treatment being provided in the specialty sector. However, this possibility is consistent with the finding that many people who are treated for their emotional problems receive therapies of undocumented effectiveness even though therapies of proven effectiveness exist (Katz et al., 1998; Wells et al., 1994). Because of this finding, future efforts to increase appropriateness of treatment are as important as efforts to increase the proportion of mentally ill women and men who obtain treatment.

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EARLY ONSET DEPRESSION AND HIGH SCHOOL DROPOUT

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1. INTRODUCTION

Mental illness, in its various forms, is common in the United States. Tens of millions of Americans are afflicted by an episode of mental illness every year. Estimates of the 12-month prevalence of mental disorders in the U.S. (including alcohol and substance abuse or dependence) indicate that 22–30 persons per 100 in the adult population are afflicted each year.[1] An episode of a psychiatric disorder, like a physical disorder, is debilitating – often disrupting the ability of the afflicted to carry on normal personal, social, and work activities. Mental illness also commonly results in large medical expenses. In addition, a number of recent papers have found that mental illness imposes large labor market losses on the ill, decreasing the likelihood of employment and limiting earnings for the employed.[2] In particular, research by two of the authors indicates that depressive disorders cause significant reductions in the labor force participation of women and the earnings of both men and women.[3]

Studies of the costs of mental illness among adults typically focus on the current costs to the affected individuals and ignore the possible impact of mental illness on school attainment. Treating schooling attainment as exogenous is a potentially serious problem. The onset of many mental illnesses occurs during adolescence, leading to a relatively high prevalence among high school students. For example, studies of high school students indicate that 8–10% of students between the ages

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of 15 and 18 suffer from clinical depression.[4] Evidence from the educational psychology literature indicates that school counselors and administrators consider emotional and behavior problems symptomatic of mental illness as serious risks for dropout.[5]

If the high prevalence of mental illness during schooling years limits educational attainment, it will have long-term economic costs. Studies in the human capital literature have established that schooling attainment is directly related to adult labor force outcomes.[6] Thus, any negative impacts of mental illness on learning and performance in school would have a lifelong, detrimental effect on labor force participation, employment, and earnings due to the reduced investment in human capital during schooling years.

Despite the obvious importance of this economic cost, the effects of mental illness on schooling are not well understood. Several empirical studies have established the linkage between health and human capital formation, but little is known about the impact of mental illness.[7] In this research, we address this shortcoming: Using an empirical model that controls for comorbid alcohol disorders, we estimate the impact of depressive disorders on the probability of an individual's failure to complete high school.

The relatively high prevalence of psychiatric disorders, especially depression, among the high school age population is of particular policy importance because many of the negative effects of these disorders on schooling attainment and adult life may be largely avoidable. For example, if depression significantly increases the probability of high school dropout, attention to the problem at the levels of both practice and policy could substantially mitigate these consequences. Adolescent depression is readily diagnosable, even by non-clinicians. Moreover, the disease is highly treatable through a variety of means. If this disease greatly increases dropout risk, the results would heighten the importance of efforts by administrators, counselors, and teachers to identify and refer ill students for treatment. Moreover, as a matter of policy, such results would support the need for a campaign to bring attention to the problem of depression among students and provide schools with the resources and information needed to combat it. More resources for school-based efforts to identify, counsel, and refer ill students for treatment would have positive effects that would last a lifetime, suggesting that policies designed to mitigate the effects of mental illness on schooling attainment would be cost-effective.[8]

Below, we first describe relevant aspects of what is known about depression among children and adolescents. We pay particular attention to studies of alcohol use and abuse because depression and alcohol disorders are highly comorbid. We briefly summarize the literatures from education, psychology, public health, and economics that bear on our understanding of mental illness and dropout. In the sections that follow, we describe our empirical model and our estimates of the impact of depressive disorders on high school dropout. We conclude with a discussion of our findings.

2. BACKGROUND

There is abundant empirical evidence in the labor economics literature of the importance of education and high school completion in determining labor market outcomes.[9] While the long-term importance of completing high school is well established, the evidence on how health and mental illness affect the chances of completing high school is sketchy. Many psychological studies report that adolescent depression may produce functional academic impairment (Judd et al., 1996). In addition, it has been found that adolescent depression is highly associated with other disorders such as anxiety, conduct, eating, and substance abuse disorders (Rice & Leffort, 1997). Despite the evidence from psychological studies of the negative impact of psychiatric disorders on schooling, few economic studies have incorporated psychiatric disorders into models of human capital formation. Those studies that have examined psychiatric disorders have largely focused solely on the effects of alcohol and substance abuse.[10]

Mullahy and Sindelar (1989), in a study using data from the New Haven site of the Epidemiologic Catchment Area (ECA) surveys, found that alcoholism at an early age significantly reduces educational achievement for young men. Using National Longitudinal Survey of Youth data (NLSY79), Cook and Moore (1993) investigate the effect of youthful drinking on years of schooling and successful completion of college. The results indicate that since higher state beer taxes and higher minimum ages reduce teenagers' consumption of alcohol, students who attend high school in states with relatively high alcohol taxes and high minimum legal drinking ages have a higher probability of graduating from college.

Using data from a longitudinal survey of 1392 students in a southeastern U.S public school system, Bray et al. (2000) investigate the relationship between the initiation of marijuana use and the probability of dropping out of high school. They find that marijuana users are 2.3 times more likely to drop out of high school than non-users, though they recognize the limitations of their regional sample.

Two studies use data from the National Comorbidity Survey to investigate the impact of psychiatric disorders on educational outcomes. In the first, Kessler et al. (1995) used 10-year birth cohorts from the National Comorbidity Survey to examine how the effects of early onset of psychiatric disorders on educational outcomes have changed over time. The authors performed survival analysis to study changes in the impact of anxiety disorders, mood disorders, substance use disorders, and conduct disorders. One of the transitions examined is failure to complete high school. They find that the proportion of high school dropouts with psychiatric disorders has increased in recent years to 14.2%. The authors find that among four major types of psychiatric disorders, those having the most deleterious impact on schooling attainment are conduct disorders (for men) and anxiety disorders (for women).

Jayakody et al. (1998) use a sample of men between the ages of 25 and 54 from the National Comorbidity Survey to investigate the impact of the early onset of affective disorders, anxiety disorders, substance use/abuse disorders and conduct disorders on schooling attainment, marital status, employment, and current mental illness. The authors find that the early onset of psychiatric disorders (before age of 16) reduces a man's educational attainment. Looking at specific types of disorders, they find that men with an early onset of conduct disorders are three times more likely to drop out of high school than healthy men.

Like Kessler et al. (1995) and Jayakody et al. (1998), our research uses the National Comorbidity Survey to examine the effect of psychiatric disorders on schooling attainment. However, this research differs from these two studies in important ways: (1) we focus solely on the impact of depression and alcohol disorders on high school completion; (2) because we focus solely on high school completion, we use a younger sample that Jayakody et al. (1998), who limit their sample to white men between the ages of 25 and 54; (3) unlike Jayakody et al. (1998), which excluded observations of women, we consider the effects of early onset psychiatric disorders on the probability of high school dropout for both men and women; (4) we include factors in our estimating model omitted from these earlier studies. Unlike Kessler et al. (1995), we include measures of parental psychiatric problems in our analysis. These measures, extensions of our earlier research on high school completion using the NCS, differ from those used in Jayakody et al. (1998).[11]

3. THEORETICAL MODEL

Schooling attainment depends on many factors. These include the student's ability, student and parental preferences, and individual characteristics. School age psychiatric disorders are a component of individual characteristics that may decrease schooling attainment. To examine the impact of depression and alcohol disorders on schooling attainment, we use a model of human capital investment similar to those typically used in schooling attainment studies.[12] We posit optimal schooling as a function of individual and family characteristics and factors assumed to be exogenous:

$$S^* = s(C, F, E), \tag{1}$$

where S^* represents optimal schooling attainment, C represents individual characteristics, F represents family characteristics, and E represents exogenous

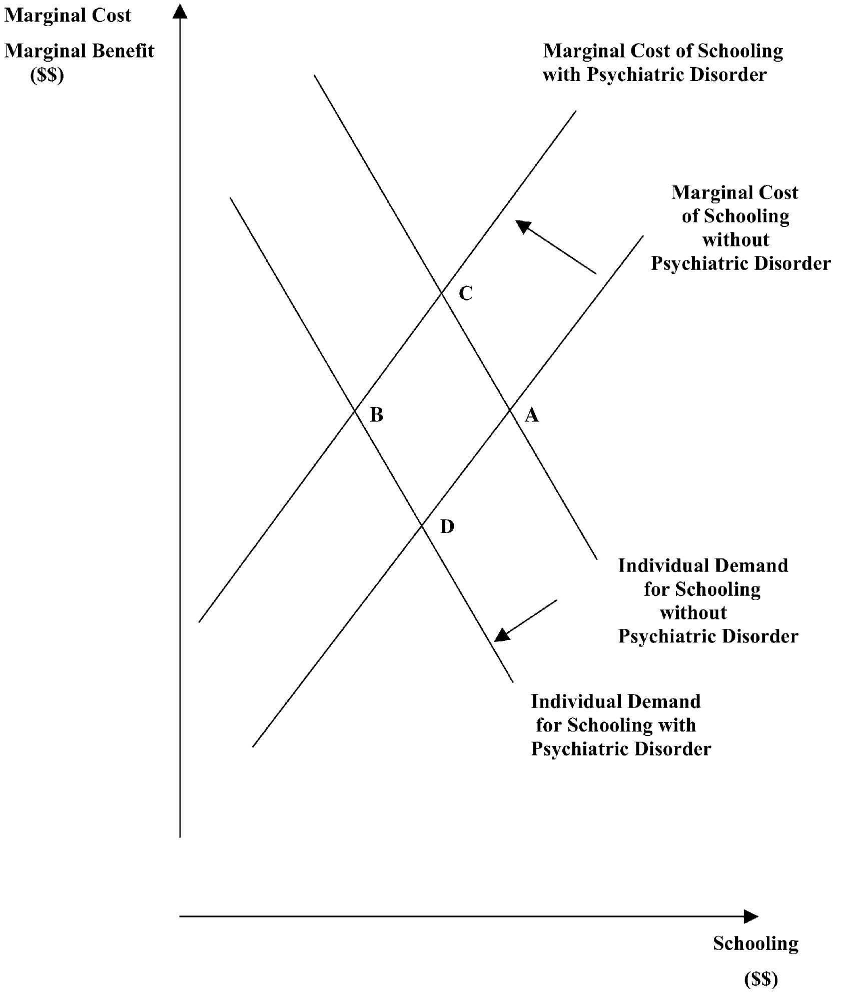

Fig. 1. Impact of Psychiatric Disorders on Schooling.

factors. We add the effects of early onset of depression (D) and alcohol disorders (A) to the model. The expanded function for optimal schooling attainment is

$$S^* = s(C, F, E, D, A). \tag{2}$$

We hypothesize that the onset of these disorders during schooling years will cause a decrease in the optimal level of schooling attainment, other things equal. Our primary interest is in estimating the impact of the early onset of depression on schooling attainment. The hypothesis is $H_1 : \partial S^*/\partial D < 0$. We also expect the early onset of alcohol disorders to have a negative impact on schooling: $H_2 : \partial S^*/\partial A < 0$. Thus, an empirical estimate indicating that these disorders increase the probability of high school dropout (decrease the probability of high school completion) will support our hypotheses. We also will examine the possibility of an interactive effect of depression and alcohol disorders by estimating a model including $D{\cdot}A$ (in addition to D and A).

The hypothesized effects are illustrated in Fig. 1. If depression and or alcohol disorders reduce the individual's ability to benefit from schooling, the demand for schooling shifts down to the left. Alternatively, if they increase the marginal cost of schooling, the supply of schooling shifts up to the left. In either case, the optimal level of schooling decreases.[13] To test our hypothesis, we estimate the effect of depression during schooling years on high school dropout, holding other factors constant. In the following section, we describe our study sample and the empirical model used for the hypothesis test.

4. EMPIRICAL METHODS

The data used in this study are drawn from the National Comorbidity Survey (NCS). The NCS was conducted between September 1990 and February 1992. The NCS is the first survey to administer a structured psychiatric interview based on a stratified, multistage nationally representative sample design. The NCS interviewed individuals between the ages 15 and 54 in the non-institutionalized civilian population (including students living in-group housing) in the 48 coterminous states of the United States. The interview data contain two parts: Part I describes the core interview administered to all respondents, part II includes a series of diagnostic questions administered to a sub-sample of respondents. We constructed our study sample from respondents in part II of the interview data and restricting the age range to include respondents old enough to have completed high school (19–54).

The NCS data is unique in the richness of detail describing individuals' psychiatric disorders. The data permit us to estimate empirical models that include the usual sociodemographic variables found in the schooling attainment literature, as well as variables representing characteristics of the individual's psychiatric disorder(s) and the parents' history of psychiatric problems. The NCS data provide information describing: (1) the respondent's characteristics, such as race, gender, age, and mental and physical health; and (2) family characteristics, such as income, education, parents' history of mental illness, family structure, language spoken in the home, number of siblings, rural residency, and number of times the person moved during childhood. The detail available in the data allows us to control for many of the confounding factors that may influence schooling attainment, so that we are able to identify the effects of individual's mental illness on the survey respondent's schooling attainment.

In our theoretical model, we specified a function for schooling in which the optimal level of schooling depends, among other things, upon individuals' mental illness. Our measure of schooling attainment is a dummy variable indicating whether the individual dropped out of high school. High school dropout is defined as completing less than 12 years of schooling.[14] We use high school dropout as the dependent variable in multivariate logistic analyses testing our hypothesis that early onset of mental illness is associated with a higher probability of dropout.

Our theoretical hypothesis assumes that we control for other factors influencing schooling attainment. To control for other factors, we include in our estimating model many of the explanatory variables common to models in the schooling literature as well as variables indicating the presence or absence of a psychiatric disorder during the first 16 years of his or her life.

The simple logistic function for DROPOUT for person is as follows:

$$\text{DROPOUT} = f(C, F, E, D, A), \tag{3}$$

where DROPOUT $= 1$ if the individual failed to complete high school and
$= 0$ if the individual completed high school;

C is a vector of individual characteristics;
F is a vector of family characteristics;
E is a vector of variables representing other exogenous factors;
D is a variable representing the individual's depressive disorder; and
A is a variable representing the individual's alcohol disorder.

The null hypothesis is that the marginal effect of the depression indicator variable will be positive. Alternatively, H_1 implies that this coefficient estimate will be negative.

Although we start with a model estimated for the pooled sample of men and women, we proceed to analyze the dropout behavior of young men and women separately. There are three reasons for doing so. First, the prevalence of depressive disorders is very different for young men and women. Second, women historically have lower high school dropout rates than men. This has been typically attributed to the fewer opportunities available to women in the labor market. With fewer labor market opportunities, the opportunity cost of being in school is lower. Finally, it may be that the effect of depression on dropout risk is different for girls than boys, because of differences in symptoms, treatment or learning. Below, we discuss the explanatory variables used in the empirical analyses. Definitions of the variables discussed below are presented in Table 1.

4.1. Depression (D) and Alcohol Disorders (A)

In our schooling model, we hypothesized that episodes of mental illness during schooling years have a negative impact on schooling attainment, other things equal. The individual's mental illness taxes personal resources leaving less time and effort for schooling. We represent a depressive disorder in our analyses with a dummy variable equal to one if the survey respondent reports the onset during schooling years of either major depression or dysthymia.[15] We represent an alcohol disorder with a dummy variable equal to one if the survey respondent reports the onset during schooling years of alcohol abuse or dependence.

We separately examine the effects of onset of depression and alcohol disorders before age 16 and age 19. While defining "early onset" to include the onset of a disorder before age 19 increases the number of respondents with reported disorders, this use of respondents with higher age of onset may introduce endogeneity between high school dropout and psychiatric disorders. That is, there may be respondents whose depression and/or alcohol disorders are at least partially caused by their having failed to complete high school. Endogeneity will bias our findings. Alternatively, defining "early onset" to include the onset of a disorder before age 16 eliminates the possibility that high school dropout is the cause of the reported disorders because dropping out of high school before age 16 is highly unlikely (and unlawful in most states). We examine the exogeneity of depression and alcohol disorders by performing tests of exogeneity using both definitions.

Given the definition of "early onset," we exclude from the study sample all respondents with early onset of psychiatric disorders other than depression and alcohol disorders.[16] This permits us to establish a reference group of respondents who were free from psychiatric disorders during schooling years. We estimate the effects of depression and alcohol disorders relative to this reference group.

Table 1. Definitions of Variables.

Variables	Definition
Dependent	
DROPOUT	1 if respondent has less than 12 years of schooling
Psychiatric disorder indicators	
DEPRESS	1 if respondent had major depression or dysthymia before age 16
ALCOHOL	1 if respondent reports alcohol dependence/abuse before age 16
Control variables	
AGE	Age of respondent
AGE2	Age squared
GOOD HEALTH	1 if respondent reports good or excellent physical health
BLACK	1 if respondent reports race as black
HISPANIC	1 if respondent reports race as Hispanic
OTHER RACE	1 if respondent reports race as not black, Hispanic, or white
ENGLISH	1 if language other than English was spoken at home
INTACT FAMILY	1 if respondent was raised by both natural parents
PARENT EDUC	Years of education of adult providing child’s primary financial support
INCOME BETTER	1 if respondent’s family was financially better off than average
INCOME WORSE	1 if respondent’s family was financially worse off than average
SIBLINGS	Number of respondent sisters and brothers
MOVES	Number of times respondent changed neighborhood of residence in childhood
RURAL	1 if respondent was raised in rural area
NORTHEAST	1 if respondent was raised in Northeast area
MIDWEST	1 if respondent was raised in Midwest area
WEST	1 if respondent was raised in West area
UNEMPLOY	Maximum unemployment rate when respondent was 16–18 years of age
VIETNAM	1 if respondent was in high school between 1965 and 1974
FATH HIST 1	1 if father suffered from psychiatric disorder(s) for at least two weeks
FATH HIST 2	1 if father’s psychiatric disorder(s) interfered with life
FATH HIST 3	1 if father had a history of treatment for psychiatric disorder(s)
MOTH HIST 1	1 if mother suffered from psychiatric disorder(s) at least two weeks
MOTH HIST 2	1 if mother’s psychiatric disorder(s) interfered with life
MOTH HIST 3	1 if mother had a history of treatment for psychiatric disorder(s)

4.2. Parents’ Mental Illness (P)

Having a parent with a psychiatric disorder disrupts the home environment, leaving less time, effort, ability, and financial resources for parents to invest in their children’s schooling. Thus, we expect parental mental illness to be negatively associated with the respondent’s schooling. The NCS data provide information describing several psychiatric disorders in parents: major depression, anxiety disorder, alcohol abuse, and drug dependence and/or abuse. In this research, we consider the relationship between the severity of a parent's disorder and the child's schooling attainment. We create three sets of dummy variables for the mother and father representing the increasing severity of a parent's psychiatric disorder. The first two dummy variables indicate whether a parent has a history of psychiatric disorders (FATH HIST 1, MOTH HIST 1). The second and third pairs of dummy variables indicate whether a psychiatric disorder interfered with life (FATH HIST 2, MOTH HIST 2) and whether the parent was treated for a psychiatric disorder (FATH HIST 3, MOTH HIST 3).[17]

4.3. Individual Characteristics (C)

A number of socioeconomic variables are included to control for gender, age, health status, race, and native language. We include a quadratic of the age variable to capture any nonlinear effects of age (AGE, AGE2). The respondents in our study samples range in age from 19 to 54 born between 1936 and 1975. The average years of schooling has increased over the last several decades in the United States. The age variables will capture this trend. Given that schooling attainment has increased over the last several decades, so that older people are less likely to have completed high school, we expect to find a positive association between age and high school dropout.

In the survey, each respondent was asked to categorize his or her health as excellent, very good, good or poor. We use a dummy variable that has a value of one if the respondent indicates that his or her health is excellent, very good, or good and a value of zero if he or she reports poor health (GOOD HEALTH). We expect better health to be associated with a lower probability of high school dropout, other things equal.

Several variables are included to control for differences in preferences due to racial and cultural factors. Typically, studies have found less high school completion among nonwhite groups. This is partially attributed to the lower return on schooling among nonwhites. We include several variables to represent different races and ethnicity in our specifications (HISPANIC, BLACK, OTH RACE). Furthermore, we added a variable to identify if the respondent spoke a language other than English at home while a child (ENGLISH). There is some evidence in the literature indicating that limited English proficiency is associated with high school dropout.[18] Including this variable implies that the variable indicating Hispanic ethnicity represents cultural differences not related to language.

4.4. Family Characteristics (F)

Several family characteristics are included as regressors in the analysis. Family income is an important determinant of children's schooling attainment. In the survey, respondents were asked to indicate whether their family income during childhood was above average, below average, or average compared to other families in their community. Using this information, we include two dummy variables in our analyses indicating if the respondent reported that his or her family was financially better off (INCOME BETTER) or worse off (INCOME WORSE) than the average family. The reference group for these variables includes respondents who indicated that their families were neither better off nor worse off than the average family. Other things equal, we expect that the probability of dropout will be lower among children living in families with above average income and higher among families with below average income. Unfortunately, because the variable represents family income relative to other families in the same community, it is unlikely to be a reliable measure of actual family income or family income relative to the national average.

While family income represents the availability of financial support for a child's schooling, it does not capture the availability of the parents' time and effort for a child's schooling. To represent parental time and effort, we include two variables in our analysis. The first indicates whether the respondent lived in a household with two parents present until at least age 15 (INTACT). We assume that having both parents present increased the time and effort devoted to the child's schooling. We therefore expect living with both natural parents to be negatively associated with high school dropout. Manski et al. (1992) found that growing up with both parents present increases the probability that a child will graduate from high school.

The second variable is the number of siblings in the respondent's childhood family (SIBLINGS). We assume that the greater the number of siblings, the less time and effort the parents have, on average, to invest in each, and therefore the lower the probability of completing high school, other things equal.[19] Behrman and Taubman (1989) found that the number of siblings is negatively related to years of schooling.

Parents' educational attainment is often included in schooling studies to reflect the parents' taste for schooling or their efficiency in the production of the household component of schooling. Evidence in the literature supports these views. Many studies have found that parents' educational attainment is a highly significant factor influencing children's schooling attainment (see Behrman & Taubman, 1989). In our analysis, we include a variable to represent the years of schooling of the adult who was the child's primary source of financial support (PARENT EDUC). In our larger sample, 78% of the respondents reported that the primary financial support came from the father, 13% reported that the primary financial support came from the mother, 3% reported that it came from both, and 6% reported that it came from other individuals. We expect PARENT EDUC to be negatively associated with dropout rates, other things equal.

Schooling is a production process that can be derailed if a child's course of studies is interrupted frequently. Lack of long-term stability is often measured in schooling studies by residential mobility. To control for residential mobility, we include in our analysis a variable measuring the number of times the respondent's family moved during schooling years (MOVES). We predict that higher mobility will be positively associated with dropping out of high school. Studies by Haveman and Wolfe (1994) and Astone and Mclanahan (1994) found that residential mobility decreases schooling success.[20]

Finally, we included several variables to control for the location in which the respondent grew up. The first is a dummy variable equal to one if the respondent lived in a rural area (RURAL). Particularly for older cohorts, we expect rural location to have a positive effect on high school dropout rates. For older cohorts, a rural location may imply that staying in school involves a high opportunity cost in terms of foregone farm labor. Furthermore, for younger cohorts, availability of employment in urban areas may imply a higher high school dropout rate caused by migration to urban areas.

It is important to control for cultural and/or economic differences across regions that may be associated with differences in dropout rates and are not explicitly controlled for by other variables in the analysis. The ideal variables would represent the state or region of residence during the respondent's schooling years. These are not available in the NCS data. However, an examination of the study sample indicates that 49% of respondents have lived in the same state for their whole life and 72% of respondents live within 200 miles of the place where they were raised during most of their childhood. Based upon this information, we use dummy variables indicating the region of residence at the time of the survey (NORTHEAST, MIDWEST, WEST) to proxy for the region of residence during schooling years. The reference region is the south.

4.5. Exogenous Factors (E)

Many empirical studies have reported a decrease in the rate of return to education in the 1970s followed by an increase in the 1980s.[21] Several factors depressed the rate of return between 1968 and 1973. The entry into the labor force of the post-World War II baby boom cohorts was an important demographic phenomenon.

Given a relatively stable demand for labor, the increased supply of labor resulted in higher unemployment and lower labor earnings. Higher unemployment and lower labor earnings may have significantly lowered the benefits of leaving high school and joining the full-time labor force. During other periods, a tighter labor market increased the benefits of high school dropout. Empirical evidence supports this hypothesis. For example, Hill (1979) found that a higher demand for teenage labor increased the probability of dropping out of high school for both men and women.

To control for variations in the unemployment rate over time, we include the maximum unemployment rates for teenage men and women during the period when the survey respondent was 16, 17, and 18 years old (UNEMPLOY). Consistent with Hill's evidence, we expect to observe that a higher unemployment rate has a negative effect on high school dropout rates, other things equal.[22]

During the years of the Vietnam War military draft, young men seeking to avoid the draft had an incentive to continue schooling longer than they would have otherwise. To capture the effect of the military draft on the optimal level of schooling attainment, we include a dummy variable (VIETNAM) indicating if the survey respondent was in high school during the Vietnam War military draft (1965–1974). We expect a negative association between high school dropout rates and this dummy variable.

5. RESULTS

Table 1 contains definitions of the variables used in the analysis. Table 2 contains the gender-specific means and standard deviations of these variables. When we define "early onset" as onset of a disorder before age 16, the study sample contains 1119 observations of women and 925 observations of men. We find that 6.5% of women and 9% of men in our study sample report not completing 12 years of schooling. This is consistent with the generally observed gender difference in dropout rates in the United States population.

In our study sample, 5.9% of women and 3% of men report having depression before 16 years of age. The higher percentage observed for women is consistent with the higher prevalence of depression found among women in the general population. We also observe that 4% of men and 1.5% of women report early onset of alcohol disorders. These observations are consistent with Kessler's (1994) findings that of American with psychiatric disorders, men are more likely to suffer from substance use and antisocial personality disorders, while women are more likely to suffer from affective and anxiety disorders.

Table 2 reveals interesting differences between women and men in the extent to which men and women report that during their childhood their parents suffered

Table 2. Means and Standard Deviations of Variables by Gender.[a]

Variables	Women		Men	
	Mean	S.D.	Mean	S.D.
DROPOUT[b]	0.065	0.26	0.09	0.33
DEPRESS[b]	0.059	0.25	0.03	0.20
ALCOHOL[b]	0.015	0.13	0.04	0.24
AGE	35.47	10.05	35.98	10.48
GOOD HEALTH	0.94	0.25	0.94	0.27
BLACK[b]	0.12	0.34	0.10	0.34
HISPANIC[b]	0.05	0.23	0.09	0.33
OTHER RACE	0.03	0.19	0.03	0.21
ENGLISH[b]	0.11	0.33	0.17	0.43
INTACT FAMILY	0.85	0.37	0.86	0.41
PARENT EDUC[c]	11.54	3.77	11.01	4.30
INCOME BETTER	0.21	0.43	0.21	0.47
INCOME WORSE	0.08	0.28	0.09	0.34
SIBLINGS[b]	3.34	2.73	3.00	2.62
MOVES[b]	1.77	2.78	1.45	2.90
RURAL[b]	0.21	0.43	0.25	0.50
NORTHEAST	0.22	0.43	0.23	0.49
MIDWEST[c]	0.29	0.47	0.25	0.50
WEST	0.17	0.40	0.18	0.44
UNEMPLOY[b]	16.96	3.01	17.41	3.63
VIETNAM	0.31	0.49	0.31	0.54
FATH HIST 1	0.08	0.28	0.07	0.31
FATH HIST 2	0.03	0.18	0.02	0.18
FATH HIST 3	0.23	0.44	0.22	0.48
MOTH HIST 1	0.12	0.35	0.12	0.38
MOTH HIST 2[b]	0.05	0.22	0.03	0.20
MOTH HIST 3[b]	0.16	0.39	0.14	0.40

[a] Means and standard deviations calculated for weighted observations of 1119 women and 925 men.
[b] The difference between the variable means for women and men is significant for $a = 0.05$.
[c] The difference between the variable means for women and men is significant for $a = 0.10$.

from psychiatric disorders. There are only small (statistically insignificant) differences between women and men in the means of the variables representing the father's reported history of psychiatric illness. However, when we consider the means of the variables representing the mother's reported history of psychiatric illness, the gap widens and is statistically significant for the two more severe indicators. It may be that women have over-reported or that men have under-reported their parents' mental illnesses. Kendler et al. (1997) found that individuals who have psychiatric disorders are more sensitive to and more likely to report that disorder in their relatives compared to other individuals. This suggests that the

Table 3. Percentage of Respondents Who Report Dropping Out of High School by Disorder.[a]

	All	Women	Men
Depression[b]	14.80	10.80	24.06
Alcohol abuse/use	22.17	24.21	21.50
Neither depression nor alcohol abuse/use	6.84	6.06	7.60

[a] Percentages calculated for weighted observations of 1119 women and 925 men.
[b] The patterns of percentages differ significantly between men and women for alpha = 0.05.

transmission of gender-linked psychiatric disorders from mother to daughter would make daughters more likely to report their mothers' psychiatric disorders. Alternatively, men may be less aware of the outward symptoms of some psychiatric disorders or simply less willing to report disorders suffered by their parents.

Table 3 reports the percentages of women and men who dropped out of high school by disorder status. The percentages of both men and women with early onset depressive disorders who dropped out of high school are significantly higher than the percentages for respondents who report neither type of disorder. Although more women report early onset of depressive disorders (see Table 2), the percentage of men with early onset depressive disorders dropping out of high school is significantly higher than the percentage of women. Among men, the onset of a depressive disorder before 16 years of age more than triples the probability of dropping out of high school. Among women, early onset of depression increases the likelihood of dropping out of high school by almost 80%.

Similarly, compared to respondents who report neither type of disorder, we see that there are significantly higher percentages of both men and women with early onset alcohol disorders who dropped out of high school. Although more men report early onset of alcohol disorders than women (see Table 2), the *difference* in the percentage of each group that drops out of high school does not differ significantly from zero. Among women, early onset of an alcohol disorder almost quadruples the likelihood of dropping out of high school. Among men, early onset of an alcohol disorder almost triples the likelihood of dropping out of high school.

The simple findings of Table 3 suggest that early onset of depressive disorders and alcohol disorders is strongly related to the probability of dropping out of high school. To test our hypotheses regarding the impact of these disorders, we report in Table 4 the marginal effects calculated from our weighted logistic analyses of high school dropout.[23] Prior to performing these analyses, we tested the exogeneity of depression and alcohol by performing Hausman tests of exogeneity. Reported at the bottom of Table 4, the test results indicate that the exogeneity of depression and alcohol disorders cannot be rejected.

Table 4. Marginal Effects of Independent Variables on Likelihood of Dropping Out of High School (Age of Onset < 16 Years).[a]

Variables	All		Women		Men	
	Marginal Effect	Prob $> X^2$	Marginal Effect	Prob $> X^2$	Marginal Effect	Prob $> X^2$
DEPRESS	0.03	0.02	0.005	0.67	0.06	0.001
ALC	0.04	0.002	0.03	0.03	0.03	0.03
FEMALE	−0.01	0.08				
AGE	−0.004	0.17	−0.0002	0.94	−0.01	0.04
AGE2	0.00004	0.35	−0.00000	0.97	0.0001	0.14
GOOD HEALTH	−0.02	0.08	−0.01	0.20	−0.03	0.04
BLACK	−0.03	0.007	−0.04	0.003	−0.01	0.47
HISPANIC	0.01	0.38	0.01	0.26	0.003	0.83
OTHER RACE	−0.04	0.09	−0.005	0.85	−0.06	0.02
ENGLISH	0.02	0.09	−0.005	0.65	0.03	0.01
INTACT	−0.01	0.24	−0.01	0.20	0.01	0.64
PARENT EDUC	−0.01	0.0001	−0.005	0.0001	−0.01	0.0001
INCOME BETTER	−0.01	0.15	−0.01	0.40	−0.007	0.55
INCOME WORSE	−0.03	0.49	−0.01	0.52	−0.01	0.40
SIBLINGS	0.005	0.0001	0.003	0.002	0.006	0.0002
MOVES	0.003	0.004	0.002	0.05	0.002	0.30
RURAL	0.01	0.31	−0.0005	0.95	0.01	0.33
NORTHEAST	−0.02	0.10	0.001	0.91	−0.03	0.05
MIDWEST	−0.01	0.15	0.003	0.71	−0.03	0.004
WEST	−0.03	0.002	−0.02	0.11	−0.04	0.008
UNEMPLOY	−0.001	0.45	0.0002	0.94	−0.004	0.14
VIETNAM	−0.03	0.02	−0.03	0.01	−0.02	0.18
FATH HIST 1	−0.04	0.01	−0.02	0.10	−0.04	0.05
FATH HIST 2	−0.08	0.11	−0.33	0.99	−0.02	0.60
FATH HIST 3	−0.003	0.70	0.002	0.76	−0.002	0.84
MOTH HIST 1	0.0005	0.96	0.01	0.48	−0.004	0.78
MOTH HIST 2	0.02	0.31	−0.0004	0.98	0.03	0.15
MOTH HIST 3	0.02	0.07	0.02	0.01	−0.004	0.76
CONSTANT	0.14	0.01	0.01	0.90	0.26	0.001
Sample Size	2044		1119		925	
−2 Log L	999.24		480.79		469.07	
Test of exogeneity						
DEPRESS |t| statistic (prob>|t|)	0.90 (0.37)		0.87 (0.39)		1.38 (0.17)	
ALC |t| statistic (prob > |t|)	0.10 (0.33)		0.78 (0.44)		0.60 (0.55)	

[a] Marginal effects calculated from coefficients of a weighted logistic regression.

The first two columns of Table 4 contain the marginal effects and coefficient *p*-values estimated for the pooled sample of 2044 respondents. The results indicate that early onset of depression significantly increases the probability of dropout by 0.03. While this is smaller than the increase associated with depression observed in Table 3, it is quite large relative to the average dropout rate for all men of 0.08. Early onset of alcohol disorders has a comparable impact (0.04) on the probability of dropout in the pooled sample.

As expected, the sex of the respondent has a significant effect on the probability of dropout. If the effects of DEPRESS and ALC differ by sex, then the marginal effects reported in the first column are not representative of the actual effects for either men or women. Consequently, in the second and third set of results in Table 4, we report separate estimates of the marginal effects of depression and alcohol disorders on dropout for women and men. The results for women indicate that early onset of depression does not have a statistically significant effect on high school completion. In comparison, early onset of depression has a large and statistically significant effect on dropout among men. Jayakody et al. (1998) report a small, positive effect of depression on high school dropout for men that is statistically insignificant. In comparison, the magnitude of the positive effect for men reported in Table 4 is much larger than that in Jayakody et al. (1998) and statistically significant.

Among women, we find that early onset of an alcohol disorder significantly increases the probability of dropout. The magnitude of the marginal effect is large relative to the average dropout rate among women (0.065) reported in Table 2. For men, early onset of alcohol disorders has a statistically significant effect on dropout, but the magnitude of the effect is only half that of depression.

The lack of effect among young women suggests important gender differences in the way depression is manifested or treated. One possibility is that alcohol abuse and dependence, which is often comorbid with depression, may be masking the negative effect of depression on the likelihood of finishing high school. To examine this possibility, we performed the same analysis after dropping ALC from the specification, but find that the marginal effects and *p*-values are virtually unchanged by dropping ALC. Thus, we find no evidence supporting this explanation.

The outcomes for many of the control variables have the expected effect on the probability of high school dropout. The probability of high school dropout decreases with years of age among men. For both sexes, higher educational attainment of a parent (or other adult providing the primary financial support for the respondent during his or her adolescent years) lowers the probability that the respondent dropped out of high school.[24] For both men and women, reporting good or excellent health as an adult is associated with a lower probability of dropout.[25]

The probability of dropout is significantly lower for black women than for white women. While there is a small negative marginal effect of BLACK on the probability of high school dropout rates among men, this effect is not statistically significant. However, for men who identified themselves as neither black nor white, there is a large reduction in the probability of having dropped out of high school. Hispanic ethnicity does not have a significant effect on high school dropout for either gender.

Family income is not significantly related to the probability of dropout. The number of siblings in the family has a small but positive effect on the probability that a respondent of either gender drops out of high school. The number of moves made by a child during their childhood has a small positive marginal effect on the probability of dropout (statistically significant for women only).

Compared to residence in the South, living in the Northeast, Midwest, or West reduces the probability of dropout among men. The gender-specific unemployment rate during the respondent's high school years does not affect the dropout rate for either sex, but the military draft during the Vietnam War years reduces the probability of dropout among women.

Prior work by three of the authors investigated the impact of parental psychiatric history on educational attainment (Farahati et al., 2000). In that research, we found that simple dummy variables indicating the presence of any parental psychiatric disorder during the respondent's youth was not significantly related to the probability of high school completion.[26] The initial specification examined in this research found a similar lack of significance between dummy variables indicating any psychiatric disorder present in the mother and father during the respondent's youth and the probability of high school dropout. However, extending the simple specification to include dummy variables representing increasing levels of severity yields significant marginal effects. We find that the probability of dropout for both men and women is affected by parental psychiatric disorders, although the effects differ for men and women. The probability of dropout for both sexes significantly declines if the respondent reports that his or her father had a psychiatric disorder (but not one that interfered with life nor for which the father received treatment). In contrast, a woman's probability of dropout is increased by 0.02 if she reports that her mother received treatment for a psychiatric disorder.

In Table 5 we report the estimated marginal effects when the age of onset of either depression or alcohol disorders was raised to 18. The test statistics (at the bottom of Table 5) indicate that at a significance level of 0.05, we must accept the exogeneity of depression and alcohol disorders. However, when the significance level is dropped to 0.10, we must reject the exogeneity of depression among men. The weak endogeneity of depression in Table 5 is reflected in the decrease in the estimated marginal effect and loss of statistical significance of depression

Table 5. Marginal Effects of Independent Variables on Likelihood of Dropping Out of High School (Age of Onset < 19 Years).[a]

Variables	All: Marginal Effect	All: Prob > X^2	Women: Marginal Effect	Women: Prob > X^2	Men: Marginal Effect	Men: Prob > X^2
DEPRESS	0.02	0.17	0.01	0.56	0.03	0.13
ALC	0.04	0.0001	0.03	0.02	0.05	0.0001
FEMALE	−0.01	0.12				
AGE	−0.003	0.42	−0.003	0.54	−0.01	0.23
AGE2	0.00002	0.64	0.00004	0.3659	0.0001	0.42
GOOD HEALTH	−0.03	0.005	−0.02	0.13	−0.05	0.002
BLACK	−0.04	0.006	−0.06	0.002	−0.009	0.61
HISPANIC	0.02	0.21	0.01	0.67	0.02	0.28
OTHER RACE	−0.05	0.04	−0.01	0.86	−0.08	0.01
ENGLISH	0.02	0.13	0.01	0.73	0.03	0.06
INTACT	−0.01	0.44	−0.004	0.74	−0.002	0.87
PARENT EDUC	−0.01	0.0001	−0.01	0.0001	−0.01	0.0001
INCOME BETTER	−0.01	0.14	−0.02	0.24	−0.005	0.70
INCOME WORSE	−0.02	0.08	−0.01	0.38	−0.03	0.03
SIBLINGS	0.01	0.0001	0.01	0.001	0.01	0.0001
MOVES	0.002	0.05	0.002	0.08	0.0002	0.90
RURAL	−0.002	0.80	−0.01	0.52	−0.002	0.85
NORTHEAST	−0.02	0.02	−0.01	0.54	−0.04	0.01
MIDWEST	−0.01	0.15	0.005	0.66	−0.03	0.01
WEST	−0.03	0.01	−0.03	0.07	−0.02	0.13
UNEMPLOY	−0.002	0.32	0.003	0.43	−0.005	0.07
VIETNAM	−0.03	0.02	−0.03	0.06	−0.03	0.08
FATH HIST 1	−0.02	0.08	−0.02	0.30	−0.02	0.30
FATH HIST 2	−0.02	0.48	−0.06	0.19	0.03	0.26
FATH HIST 3	−0.0003	0.97	0.005	0.64	−0.004	0.70
MOTH HIST 1	0.002	0.88	0.01	0.62	−0.001	0.96
MOTH HIST 2	−0.001	0.94	−0.02	0.54	0.01	0.54
MOTH HIST 3	0.02	0.07	0.03	0.02	0.00002	0.99
CONSTANT	0.12	0.05	−0.004	0.96	0.25	0.004
Sample Size	2608		1290		1318	
−2 Log L	1284.98		559.99		675.32	
Tests of exogeneity						
DEPRESS |t| statistic	0.10		0.64		1.74	
(prob>|t|)	(0.92)		(0.52)		(0.08)	
ALC |t| statistic	0.15		0.25		0.86	
(prob>|t|)	(0.88)		(0.81)		(0.39)	

[a] Marginal effects calculated from coefficients of a weighted logistic regression.

compared with the values in Table 4. This suggests that depression cannot reliably be treated as exogenous in this case.

We also estimated models in which the measure of psychiatric illness was broadened to either a single dichotomous measure of mental illness or to two measures representing substance abuse disorders and non-substance abuse psychiatric disorders. In both of these cases, exogeneity must be rejected. This suggests that models in which nondepressive psychiatric disorders are treated exogenously, such as Jayakody et al. (1998), are likely to yield biased estimates of the disorders' effects. We estimated instrumental variable models for these broader categories of psychiatric disorders, but were unable to obtain consistent estimates of the effects of non-depressive psychiatric illnesses on the probability of dropout because our instruments do not perform well.[27] The instruments used in these analyses were the same as those used successfully in instrumental variables estimation of the effects of adult mental illness on labor market earnings in the authors' prior research (2000a). This suggests that over time, adult children learn to compensate for the negative productivity effects of parental characteristics.

6. DISCUSSION

In this chapter we have examined the role of early onset depression and alcohol disorders in determining risk for high school dropout. Using data on a sample of 925 men and 1119 women, ages 19 through 54, who responded to the National Comorbidity Survey, we have found that onset of depression prior to the age at which school attendance is no longer compulsory significantly increases the probability of high school dropout among men, but not among women. Onset of alcohol abuse or dependence at this age significantly increases the probability of high school dropout for both men and women.

Our analyses are based on estimating multivariate logistic regression models in which we are able to include control variables typically used in studies of educational attainment. Our estimates for these control variables are similar to results found in prior studies, suggesting that this is an appropriate sample for investigating the impact of mental illness on high school dropout. Moreover, because we are able to identify age at onset, we are able to exclude from our sample those who first experience depression between the age of 16 and 18, for whom identifying the causal relationship between depression and completion of high school is difficult. By focusing on those with early onset depression – during years when dropout was not legally possible – we feel more confident that depression can plausibly be viewed as exogenous.

We believe this insight into the role of depression in shaping economic prospects and outcomes for the ill. While much has been done during the past few decades to advance our understanding of the costs imposed by mental illness in the labor market, little is known about the role of adolescent mental illness in limiting human capital accumulation. Our work here suggests that for men, early onset of depression imposes non-trivial indirect costs by increasing the risk of high school dropout, in addition to the direct costs of illness in the labor market.

The fact that we find no such increased risk among girls suggests important gender differences in the way depression is manifested or treated. One explanation for these differences is that girls may be more likely to receive treatment for depression than boys. There is evidence that among depressed adults, women are more likely to undertake treatment than men. This pattern may be true of adolescents as well, both because of girls' own preferences, and because parents might be less willing to seek treatment for their sons if the stigma effect of a mood disorder is larger for boys. Unfortunately, there is no way to test this explanation of the gender differences observed here.

We hope that the research reported in this chapter serves as a platform for further research on the impacts of psychiatric disorders on human capital formation. In further research, several extensions of this research seem fruitful. For example, to capture more comprehensively the effects of psychiatric disorders, it would be useful to include additional mental disorders and employ richer measures of disorder, including frequency of episodes. Another potentially valuable extension would be an examination of the role of treatment, and the age of first treatment in mitigating costs. Finally, the connections between early onset of psychiatric disorders, high school dropout, and teen parenting (a frequent outcome among teenagers with psychiatric disorders) need to be carefully explored.

Despite the fact that this initial analysis leaves many questions unanswered, the results have important policy implications. At the very least, the results suggest that policy initiatives designed to recognize and treat adolescent depression and to provide support to people suffering from depression may have important long term indirect benefits by promoting higher levels of schooling attainment and socioeconomic success.

NOTES

1. See Regier et al. (1993) for estimates based on the Epidemiologic Catchment Area Study and Kessler et al. (1994) for estimates from the National Comorbidity Survey.
2. See Ettner et al. (1997), Ettner (2000), Marcotte et al. (2000a, b) and Slade and Albers (2000).
3. Marcotte et al. (2000a, b).
4. Another 15–20% suffer from less severe or moderate depression. See Sullivan and Engin (1986), Ehrenberg et al. (1990), and Connelly et al. (1993).
5. See Bull et al. (1991), Egyed et al. (1998), and Ekstrom et al. (1986).

6. See Becker (1993), Nerdrum (1999), Card (1999), Blackmore and Low (1984), Stern et al. (1989), and Bedi and Gaston (1999).

7. Grossman (1972a, b), Hamilton et al. (1997), Hunt-McCool and Bishop (1998) examine health and human capital formation.

8. Indeed, empirical research suggests that the estimated costs of treatment for many types of disorders are substantially lower than the productivity losses due to these disorders (e.g. Lave et al., 1996).

9. See Card and Krueger (2000) and Ashenfelter et al. (1999) for reviews of this literature.

10. Bray et al. (2000), Cook and Moore (1993).

11. Farahati et al. (2003).

12. Our model, available on request from the authors, is similar to those of Rosen (1973) and Griliches (1977).

13. A decrease in the demand for schooling due to mental illness will decrease the rate of return to schooling, while an increase in the cost of schooling will increase the rate of return to schooling. We do not attempt to ascertain the underlying structural relationships and the impact of I on the rate of return to schooling.

14. The NCS data do not indicate whether an individual actually graduated from high school or obtained a graduate equivalency degree.

15. Major depression typically involves episodes of severe depression, while dysthymia is a less severe form of chronic depression. We exclude manic disorders, such as bipolar disorders, because manic episodes may have productive effects and because these disorders are not treated as successfully as depressive disorders.

16. Other disorders include anxiety disorders, conduct disorders, and drug disorders.

17. In an earlier study (Farahati et al., 2003), the authors reported an investigation of the impacts of the various types of parental psychiatric disorders on high school dropout.

18. The United States Department of Education (1990).

19. Becker and Lewis (1973) and Becker and Tomes (1976) developed the theory describing the tradeoff between the quantity and quality of children.

20. Note, however, that Long (1992) reported that family dissolution is significantly associated with residential mobility, so that at least part of the mobility effect may be due to family dissolution. Astone and Mclanahan (1994) found that 30% of the difference in the probability of high school dropout between children from stepfamilies and children from intact families could be explained by the differences of their residential mobility.

21. See, for example, Mincer (1991).

22. There were also exogenous shocks to the demand for labor, such as oil crises, that may have contributed to the lower rate of productivity growth in the 1970s. All of these shocks will be reflected in the unemployment rates for teenage men and women.

23. In an alternative specification, we also included the interaction of DEPRESS and ALC, but this effect was not statistically significant for any of the samples examined in Tables 4 and 5.

24. Jayakody et al. (1998) found similar effects for age and parental schooling in their study of male respondents to the NCS.

25. The data does not include a general health indicator for respondents during school age years. Although GOOD HEALTH is included to proxy for adolescent health, in fact it measures adult health. Because better adult health may be an outcome of more schooling, it is not clear whether this effect represents the effect of health on dropout or the effect

of dropout on adult health. Dropping GOOD HEALTH has no impact on the coefficient estimates.

26. However, Farahati et al. (2003) identifies several specific parental mental disorders that are significantly related to the child's probability of dropout.

27. Overidentification tests reject the appropriateness of our instruments.

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GENDER DIFFERENCES IN THE LABOR MARKET EFFECTS OF SERIOUS MENTAL ILLNESS

Pierre Kébreau Alexandre, Joseph Yvard Fede and Marsha Mullings

1. INTRODUCTION

Mental disorders collectively account for 4 of the 10 leading causes of disability and represent more than 15% of the overall burden of disease in the United States (SAMHSA, 1999). The first Surgeon General's Report on Mental Health reported that in 1999 nearly 20 million American adults (9.5% of the population) were clinically depressed and that, at any one time, 1 in every 20 employees is experiencing depression (SAMHSA, 1999). The indirect costs of mental disorders to the American economy amounted to an estimated $79 billion in 1990, with loss of productivity because of illness accounting for about 80% of these costs ($63 billion) (Rice & Miller, 1996). Additionally, significant costs may accrue from decreased productivity due to symptoms that sap energy, affect work habits, and cause problems with concentration, memory, and decision-making (SAMHSA, 1999).

Mental disorders often affect individuals during their most productive years and may represent one of the most important factors influencing labor market decisions (Mitchell, 1990; Mitchell & Butler, 1986). Since the 1970s, economists have directed more attention toward the relationship between mental health and labor

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supply, but lack of data sets with detailed information on mental health and labor market information imposed significant constraints to rigorous research. Bartel and Taubman (1979) conducted one of the first studies and found that individuals diagnosed as either psychotic or neurotic had a lower probability of being in the labor force, received lower wages, and worked fewer weekly hours. Other researchers have linked mental illness to a variety of labor market related negative functional outcomes. For example, Conti and Burton (1994) indicate that, with regard to absenteeism, mental disorders tend to surpass other common chronic medical conditions such as heart disease and lower back pain in terms of the average length of the disability period. Hays et al. (1995) and Wells et al. (1996) reported that both chronic and episodic depressions had similar or greater functional limitations than those attributed to other chronic medical illnesses. Focusing on the associations between work outcomes and treatment of mental disorders, Mintz et al. (1992) found that work outcomes were consistently better for patients whose mental illness had remitted and that they improved with duration of treatment. Moreover, in a study based on clinical trials, Berndt et al. (1998) found that the level of work performance was negatively related to the severity of depressive status and that a reduction in depressive severity improved the patient's perceived work performance.

Based on large community-based samples, the Epidemiologic Catchment Area (ECA) study, one of the first main sources of data on the prevalence of psychiatric disorders in the U.S., has enabled researchers to examine the mental illness/labor supply relationships for diverse cohorts of the U.S. population. Conducted in the 1980s, the ECA was based on the Diagnostic Interview Survey, a self-assessed, non-clinician research diagnostic interview and covered five communities: Baltimore, MD; Los Angeles, CA; New Haven, CT; the Piedmont area of North Carolina, and Saint Louis, MO (Robins & Regier, 1991). Mitchell and Anderson (1989) used the ECA to explore the relationship between mental health and the labor force status of older American workers (aged 50–64) in Baltimore, the Piedmont area, and Los Angeles. They found that mental illness negatively affected the labor force participation of older men, but had no significant effect on the labor force participation of older women. Mullahy and Sindelar (1990) used Wave I of the New Haven site and estimated four alternative specifications with successive incremental refinements of standard economic variables, self-reported mental health variables, diagnosed mental health variables, and self-reported physical health variables. The authors found that mental health was a significant predictor of labor force participation for both men and women, but that the effects varied by gender. Frank and Gertler (1991) examined ECA data from the Baltimore site. Defining mental distress as impairment plus the presence of a diagnosis, they indicated that mental distress at baseline contributed to a 21% reduction in the annual income of male workers.

Other attempts to estimate the labor market impacts of mental illness using community-based samples include a study by French and Zarkin (1998), who used data from a large manufacturing worksite to explore the relationship between symptoms of emotional and psychological problems and employee absenteeism and earnings. They found that workers who reported symptoms of emotional and psychological problems had higher absenteeism rate and lower earnings than otherwise similar coworkers. Alexandre and French (2001) used data collected between 1996 and 1997 in crime-ridden and low-income neighborhoods of Miami-Dade County, Florida to examine the relationship between depression and labor supply while examining the robustness of the model estimates to the co-morbid effects of substance abuse. They indicated that depression significantly decreased the probability of being employed as well as the number of weeks worked during the last 12 months. The authors also showed that co-morbid substance abuse significantly contributed to the estimated effects of depression.

Utilization-based measures are dependent on whether the individuals were treated for mental illness and are likely to bias the estimates (Frank & Gertler, 1991). Community-based measures, on the other hand, assess mental illness from the responses to psychological questions and the community-based variables are independent of whether respondents were treated for a mental health problem. However, results from community-based samples do not generalize to the general population. Most recently, in an attempt to bring new information and more robust estimates of the effects of mental illness on labor market outcomes for the general population, a few researchers have used the National Comorbidity Survey (NCS). In one of these studies, Ettner et al. (1997) used the NCS to examine the impact of psychiatric disorders on employment, conditional work hours, and income for 2,225 men and 2,401 women. They found that psychiatric disorders reduced employment and conditional earnings among both men and women, but adversely affected conditional work hours for men only. Marcotte et al. (2000) also used the NCS and indicated that affective disorders had resulted in significant earnings and employment losses, with the most pronounced impacts on depressive women.

The present study contributes to the growing literature in many respects. First, it utilizes the 2001 National Household Survey on Drug Abuse (NHSDA), a national representative survey of U.S. households, which for the first time included a series of questions designed to assess mental illness among adults aged 18 and older (SAMHSA, 2002). Second, this study incorporates a number of methodological strategies that address empirical problems typically associated with estimating mental illness/labor supply relationships. Specifically, the existing literature suggests that mental illness may be an endogenous variable, that is, missing or unobservable determinants of both mental illness and labor supply outcomes may be correlated. While major efforts have been made to generate representative data sets at the national level, researchers are still confronting challenging estimation issues in examining the mental illness/labor supply relationships (Alexandre & French, 2001; Marcotte & Wilcox-Gök, 2001). Finally, previous research demonstrated that males and females differed considerably in their mental illnesses and labor market behavior (Ettner et al., 1997; Kaestner, 1994; Mullahy & Sindelar, 1990, 1991). While economists have spent a great deal examining gender differences in labor market behavior (see, for example, Blundell & Meghir, 1986), mental health is an important aspect of both female and male labor supply that has received little attention. Thus, the present analysis separated males and females into sub-samples to conduct a richer examination of gender differences.

2. EMPIRICAL MODELS AND ESTIMATION ISSUES

To study the relationship between serious mental illness (thereafter referred to simply as mental illness) and labor supply, we used the linear econometric specification below (see, for example, Alexandre & French, 2001; French & Zarkin, 1998; Miller & Kelman, 1992; Mitchell & Anderson, 1989; Mullahy & Sindelar, 1990, 1991):

$$L = \beta_1 M + \beta_2 X + \varepsilon \tag{1}$$

where L is a measure of labor supply; M represents mental illness; X is a vector of exogenous socio-demographic and environmental characteristics; the β'*s* are parameters to be estimated; and ε is an error term containing, among other things, unobserved characteristics and preferences.

We examined two labor supply variables: "unemployment" and "workdays skipped" during the past 30 days prior to the interview date. These measures are consistent with the Bureau of Labor Statistics (BLS, 2000) definitions, but it is worth noting that individuals who may have secured a job shortly before the interview would be designated as "employed," while those who may have continuously worked for several months and stopped working within 30 days of the interview would be designated as "unemployed." Similar issues may exist for the measurement of "workdays skipped" since some individuals may have not missed work for several months, but skipped days of work during the past 30 days prior to the interview. While these measurement concerns probably applied only to a small percentage of survey respondents, it is important to recognize that changes in variable definitions and in the timing of survey administration could have an impact on the empirical findings.

The analysis defined mental illness (M) as having a diagnosable mental, behavioral, or emotional disorder that met criteria in the 4th edition of the Diagnostic and Statistical Manual of Mental Disorders (DSM-IV) and that resulted in functional impairment that substantially interfered with or limited one or more major life activities (American Psychiatric Association, 1994; SAMHSA, 2002). We expect mental illness to negatively affect labor supply. Thus, the probability of unemployment would be higher, and conditional on being employed, individuals with mental health problems would devote less time to work activities. But, estimation of Eq. (1) with standard techniques such as OLS or probit would generate biased results if M is significantly correlated with ε (Manning et al., 1987; Mullahy, 1998; Mullahy & Sindelar, 1996; Staiger & Stock, 1997). Specifically, the effect of M on L can be written as $\mathrm{d}L/\mathrm{d}M = LM^{*} + LM\mathrm{d}\varepsilon/\mathrm{d}M$, which will not be equal to L_M if $\mathrm{d}\varepsilon/\mathrm{d}M$ is non-zero, that, is if M is endogenously determined in Eq. (1).

The endogenous relationship may work in many ways. First, stresses associated with unemployment may be important unobserved factors distinguishing the mentally ill and the non-mentally ill. Second, reduced income associated with employment gaps and other labor supply problems may lead to increased mental distress (Hamilton et al., 1997). Third, employed individuals may have more disposable income to spend on mental health services as well as increased access to mental health care through employer-provided insurance coverage. To the extent that mental health and physical health are positively correlated, even insurance coverage that solely covers physical health might also improve mental health. Finally, mentally ill and non-mentally ill individuals may differ in unobserved ways that impact skills and ability and, thus, are likely to affect labor market outcomes. For example, mental disorders may make skill acquisition more costly and thereby result in lower skill levels and ability and, thus, are likely to affect employment.

An efficient or consistent estimate of $\mathrm{d}L/\mathrm{d}M$ can be obtained by addressing the endogeneity of M in the labor market equation through variations of a two-stage IV technique (Alexandre & French, 2001; Angrist & Krueger, 2001; Evans et al., 1999; French et al., 2001; Marcotte et al., 2000; Mullahy & Sindelar, 1996; Norton et al., 1998). Davidson and MacKinnon (1993) note that the two-stage IV method can be used to obtain consistent estimates regardless of how that correlation may have arisen.[1]

In the two-stage IV procedure, we specified a first-stage probit specification for mental illness (M) with the observable explanatory variables noted earlier (X) and a composite measure of religiosity (Z):

$$\Pr(M = 1) = \Phi(\gamma_1 X + \gamma_2 Z) \tag{2}$$

where the γ'*s* are parameters to estimate, and Z is a dichotomous variable equal to 1 if the respondent agreed or strongly agreed with *each* of the following statements: "religious beliefs are important to me," "religious beliefs influence my decisions," and "it is important that my friends share my religious beliefs." This composite measure for religiosity is intended to identify the strongly religious individuals in the sample. The expectation is that M is inversely related to L as cross-sectional and longitudinal studies have consistently found significant associations between religious beliefs and mental and physical health (Levin, 1994). Moreover, religious involvement was shown to have significant protective effects for the emotional well being of individuals in crisis (Bradley, 1995; Ellison, 1991; Gorsuch, 1995). To complete the two-stage IV procedure, the predicted values of mental illness ($\hat{M}$) from Eq. (2) replaced the mental illness variable (M) in Eq. (1). We corrected the estimated standard errors for the use of $\hat{M}$ in Eq. (1) (Murphy & Topel, 1985).

The first response variable of interest, employment, is a binary measure. A second-stage probit estimated the probability of being employed with the corresponding explanatory variables (X) and the predicted values of mental illness ($\hat{M}$) (Angrist & Krueger, 2001; Maddala, 1983). The second measure of labor supply represented the number of workdays the employed individuals skipped in the past 30 days prior to the interview. These count data displayed a large number of zeros. We conducted a likelihood ratio test for overdispersion (Stata, 2003) and the Vuong test (Vuong, 1989), which suggested the use of the zero-inflated negative binomial (ZINB) as the appropriate estimation approach for both the male and the female samples. The two-stage IV procedure in this instance corresponded to the two-stage quasi-maximum likelihood (2SQML) technique, suggested by Mullahy (1997). Specifically, second-stage ZINB estimated the weekdays skipped specifications with the corresponding X and $\hat{M}$.

Single equation models usually have a smaller mean squared error, implying a practical tradeoff between bias and variance (Davidson & MacKinnon, 1993). Because the models are non-linear, we used the method suggested by Smith and Blundell (1986) to test for the potential endogeneity of mental illness. The null hypothesis that mental illness was exogenous involved a χ^2-test of the explanatory power of the residuals from the first-stage equation for mental illness when added to Eq. (1). Similar to the Hausman-Wu test (Hausman, 1983; Wu, 1973), the Smith-Blundell test is influenced to a large degree by the reliability of the instruments (Bollen et al., 1995; Nelson & Startz, 1990; Norton et al., 1998; Staiger & Stock, 1997). We used χ^2-tests for the significance of the instrument in explaining mental illness to test instrument reliability (Bollen et al., 1995; Norton et al., 1998).[2]

3. SAMPLE AND DATA

The empirical analysis used data from the 2001 National Household Survey on Drug Abuse (NHSDA), which is the twenty-first survey in a series that has been conducted since 1971. Starting in 1999, the survey has conducted about 70,000 interviews each year using a computer-assisted interviewing methodology. The sampling design was a nationally stratified multistage area probability sample of the non-institutionalized household population in the 50 United States who were 12 years of age and older. The NHSDA is one of the largest surveys ever undertaken in the U.S. A major change in NHSDA questionnaire design was initiated in 1994, which placed more emphasis on health status and health care, access to care, and mental health (SAMHSA, 2002). Of particular interest for the present study is that the 2001 NHSDA included a new series of questions designed to assess serious mental illness (SMI), defined on the basis of criteria specified in the Diagnostic and Statistical Manual of Mental Disorders, 4th edition (DSM-IV) (American Psychiatric Association, 1994). A scale consisted of six NHSDA questions were used to assess SMI. These questions asked how frequently a respondent experienced symptoms of psychological distress during one month in the past year when the individual was at his/her worst emotionally (SAMHSA, 2002). Based on clinical assessments done on survey respondents, methodological research has determined that the scale was a good predictor of SMI (Kessler et al., in press).

The NHSDA also has limitations that can affect our analyses (SAMHSA, 2002). The data are self-reported, which raises questions regarding validity and reliability. NHSDA procedures are designed to maximize honesty and recall, but ultimately the value of the data depends on respondents' truthfulness and memory. A few studies have examined the validity of self-reported information in this context and have found the measures were quite good (Harrison & Hughes, 1997; Preston et al., 1997; Turner et al., 1992). Furthermore, a small segment of the U.S. population (slightly less than 2%) was excluded from the sampling frame because they were not part of the target population. The excluded subpopulations were members of the active duty military and persons in institutional group quarters (e.g. hospitals, prisons, nursing homes, treatment centers). For additional details on sample design, response rates, major findings, and other technical details of the 2001 NHSDA, refer to SAMHSA (2002).

Table 1 reports the variable definitions of all of the variables included in the empirical models. The core demographic variables were the standard economic variables used in labor supply analyses and included age, gender, race, marital status, and education. Dummy variables were created to identify whether there were children living in the household, whether the individual was born in the U.S., received unearned income during the past 12 months, has moved during the past 12 months; or lived in a Metropolitan Statistical Area (MSA) during the past 12 months. Individuals under the age of 25 or over the age of 64 were excluded from our analyses given their unique employment choices (e.g. full-time students, new labor market participants, occasional workers, and retirees).

Table 1. Variable Definitions.

Variables	Definitions
Labor supply	
Unemployed	Unemployed = no job in the past 30 days
Work days skipped	Number of days skipped work in past 30 days if employed
Socio-demographics	
Age 26–34	Equal 1 if individuals aged 26–34; 0 otherwise
Age 35–49	Equal 1 if individuals aged 35–49; 0 otherwise
Age 50–64	Equal 1 if individuals aged 50–64; 0 otherwise
African-American	Equal 1 if African-American; 0 otherwise
Hispanic	Equal 1 if Hispanic; 0 otherwise
White	Equal 1 if non-Hispanic White; 0 otherwise
Other ethnic groups	Equal 1 if Native American; Asian, or Pacific Islander; 0 otherwise
Married or living as married	Equal 1 if married or living as Married; 0 otherwise
Less than high school	Equal 1 if less than a high school diploma or GED; 0 otherwise
High school	Equal 1 if received a high school diploma or GED; 0 otherwise
Some college	Equal 1 if attended college or technical school; 0 otherwise
Graduate school	Equal 1 if attended graduate school school; 0 otherwise
Any kids in household	Equal 1 if children under 18 live in household; 0 otherwise
Born in U.S.	Equal 1 if born in the U.S.; 0 otherwise
Any unearned income in past 12 months	Equal 1 if received income from sources other than a job; 0 otherwise
Had moved in past 12 months	Equal 1 if had moved in the past 12 months; 0 otherwise
MSA residency	Equal 1 if living in a metropolitan statistical area; 0 otherwise
Instrumental variable	
Strongly religious	Equal to 1 if the respondent agreed or strongly agreed with *each* of the following statements: "religious beliefs are important to me," "religious beliefs influence my decisions," and "it is important that my friends share my religious beliefs."

Table 2 presents the mean values of all variables used in the empirical analysis. Mean values are reported separately for males and females, and by mental illness status. We used *t*-tests of differences for independent samples to determine if there were significant differences across the mental illness categories. Unemployment was statistically significantly different across mental illness status for males, but not significant for females. The workdays skipped variable was statistically significantly different across mental illness status in both samples. The following section reports the findings of the multivariate analysis when the labor supply variables are estimated with univariate models and two-stage IV techniques.

Table 2. Gender-Specific Variable Means, by Mental Illness Status.

Variable	Males		Females	
	SMI ($N = 348$)	NSMI ($N = 7{,}474$)	SMI ($N = 690$)	NSMI ($N = 6{,}897$)
Labor supply				
Unemployed (%)[a]	8.05 (27.24)	3.21 (17.63)	7.25 (25.94)	5.63 (23.04)
Work days skipped[a,b]	1.35 (4.10)	0.58 (2.54)	1.36 (3.80)	0.80 (3.14)
Socio-demographics (%)				
Age 26–34[a,b]	44.83	34.40	42.90	33.54
Age 35–49	46.55	50.55	48.41	49.24
Age 50–64[a,b]	8.62	15.05	8.70	17.22
African-American	8.62	8.35	9.42	11.58
Hispanic[a]	8.05	11.65	7.68	9.28
White[b]	77.59	74.53	77.68	73.86
Other ethnic groups	5.75	5.47	5.22	5.28
Married or living as married[a,b]	48.27	68.66	52.60	64.33
Less than high school	14.65	12.14	9.85	8.75
High school	33.04	31.48	32.89	31.52
Some college[b]	27.59	24.62	31.45	27.32
Graduate school[a]	24.71	31.75	25.80	32.41
Any kids in household[a]	36.78	46.38	51.59	49.20
Born in U.S.[b]	87.93	85.32	91.30	87.99
Any unearned income in past 12 months[a]	13.21	9.78	12.89	10.64
Had moved in past 12 months[a,b]	37.68	20.85	27.82	18.45
MSA residency	69.82	72.94	75.36	72.55
Instrumental variable (%)				
Strongly religious[a,b]	4.02	7.02	7.10	8.62

Note: SMI = Seriously Mentally Ill; NSMI = Non-Seriously Mentally Ill. Standard deviations are in parentheses.

[a] Significant differences between SMI and NSMI, male sample, $p \leq 0.05$.

[b] Significant differences between SMI and NSMI, female sample, $p \leq 0.05$.

4. ESTIMATION RESULTS

Our analysis estimated a variety of empirical models to determine the relationships between mental illness and labor supply. The labor supply variables included unemployment in the past 30 days and workdays skipped in the past 30 days prior to the interview.[3] We estimated separate equations for males and females. Tests of endogeneity of mental illness and reliability of the instruments were executed. For illustration purposes, we estimated the labor supply models with and without

Table 3. Coefficient Estimates for Univariate Probit of SMI.

Variable	Males	Females
Age 35–49	0.0839 (0.0557)	−0.1178*** (0.0452)
Age 50–64	−0.2409** (0.0952)	−0.4460*** (0.0752)
African-American	−0.0514 (0.0951)	−0.2653*** (0.0717)
Hispanic	−0.2499** (0.1103)	−0.1636* (0.0863)
Other ethnic groups	0.0023 (0.1200)	0.0129 (0.0998)
Married	−0.2948*** (0.0615)	−0.2236*** (0.0443)
High school	−0.1156 (0.0845)	−0.0889 (0.0779)
Some college	−0.0933 (0.0874)	−0.0614 (0.0788)
Graduate school	−0.2368*** (0.0892)	−0.2656*** (0.0805)
Any kids in household	−0.0344 (0.0627)	0.0117 (0.0449)
Born in U.S.	−0.0007 (0.0961)	0.1732** (0.0814)
Any unearned income in past 12 months	0.1411* (0.0789)	0.1465** (0.0640)
Had moved in past 12 months	0.2927*** (0.0583)	0.1980*** (0.0502)
MSA residency	−0.0743 (0.0582)	0.1068** (0.0483)
Strongly religious	−0.1823*** (0.0657)	−0.0942** (0.0472)
Constant	−1.2827*** (0.1381)	−1.1934*** (0.1213)

Note: SMI = Serious Mental Illness. Standard errors in parentheses.
*Statistically significant, $p \leq 0.10$.
**Statistically significant, $p \leq 0.05$.
***Statistically significant, $p \leq 0.01$.

a correction for potential endogeneity of mental illness, but the focus on the coefficient estimates presented is dictated by the results of the endogeneity tests (Bollen et al., 1995). Results of the multivariate analyses are presented in Tables 3–5, including probit estimates of the mental illness specifications. The most important results for the unemployment and workdays skipped specifications correspond to coefficient estimates for the mental illness variable. We present post estimation of the marginal effects of mental illness in squared brackets. The marginal effects, evaluated at the mean values for the other variables, represent the difference in expected probabilities of the labor supply variable between the mentally ill and the non-mentally ill (Greene, 2000).

4.1. The Determinants of Mental Illness

We report the first-stage probit results for mental illness in Table 3. With a few exceptions, the results were quite similar for men and women. Men and women who were between the ages of 50 and 64 were less likely to suffer from mental illness relative to those who were between 25 and 34. Regardless of gender,

Table 4. Estimation Results for Unemployment: Males and Females.

Variable	Males		Females	
	Univariate Probit	IV Estimation	Univariate Probit	IV Estimation
SMI	0.3509*** (0.1052) [0.0304]	0.5654 (2.6374)	0.1089 (0.0787)	−0.8608 (2.1455)
Age 35–49	0.1243 (0.0744)	0.1299* (0.0694)	0.0021 (0.0558)	−0.0182 (0.0721)
Age 50–64	0.1168 (0.0964)	0.1203 (0.1133)	0.1143 (0.0767)	0.0517 (0.1581)
African-American	0.0956 (0.0982)	0.0963 (0.1014)	0.0441 (0.0760)	−0.0009 (0.1252)
Hispanic	0.1286 (0.1064)	0.1343 (0.1270)	−0.0642 (0.0948)	−0.0934 (0.1143)
Other ethnic groups	0.2630** (0.1207)	0.2650** (0.1207)	−0.0642 (0.1185)	−0.0659 (0.1186)
Married	−0.3116*** (0.0667)	−0.3008*** (0.1068)	0.0113 (0.0521)	−0.0262 (0.0982)
High school	−0.1482* (0.0849)	−0.1484 (0.0918)	−0.3234*** (0.0746)	−0.3401*** (0.0825)
Some college	−0.4171*** (0.0962)	−0.4128*** (0.0998)	−0.6083*** (0.0813)	−0.6207*** (0.0850)
Graduate school	−0.2558*** (0.0909)	−0.2520** (0.1120)	−0.7104*** (0.0826)	−0.7527*** (0.1246)
Any kids in household	−0.1351* (0.0696)	−0.1407** (0.0700)	−0.1153** (0.0536)	−0.1126** (0.0538)
Born in U.S.	0.0700 (0.0987)	0.0738 (0.0988)	−0.2039** (0.0837)	−0.1770* (0.1023)
Any unearned income in past 12 months	0.4028*** (0.0773)	0.3981*** (0.0882)	0.1943*** (0.0700)	0.2184** (0.0875)
Had moved in past 12 months	0.2255*** (0.0652)	0.2216** (0.1112)	0.2570*** (0.0582)	0.2920*** (0.0979)
MSA residency	−0.0589 (0.0654)	−0.0577 (0.0678)	−0.0792 (0.0541)	−0.0600 (0.0672)
Constant	−1.6475*** (0.1442)	−1.6654*** (0.2751)	−0.9514*** (0.1265)	−0.8367*** (0.2830)

Note: SMI = Serious Mental Illness.
Estimates for marginal effects of SMI, estimated for the statistically significant coefficients only, are in squared brackets.
Standard errors reported in parentheses.
* Statistically significant, $p \leq 0.10$.
** Statistically significant, $p \leq 0.05$.
*** Statistically significant, $p \leq 0.01$.

Table 5. Regression Results for Workdays Skipped: Males and Females.

Variable	Males		Females	
	ZINB	2SQML	ZINB	2SQML
SMI	0.7585*** (0.1951) [0.5967]	0.2707 (0.5170)	0.5110*** (0.1244) [0.5086]	1.0174 (0.8608)
Age 35–49	−0.0128 (0.0918)	0.0185 (0.1034)	−0.1969** (0.0832)	−0.1080 (0.1356)
Age 50+	−0.0498 (0.1395)	0.0086 (0.1906)	−0.4133*** (0.1234)	0.0180 (0.4072)
African-American	0.1707 (0.1485)	0.1466 (0.1551)	0.0035 (0.2206)	0.2392 (0.3283)
Hispanic	−0.1490 (0.1567)	−0.1363 (0.2229)	0.1887 (0.1419)	0.3339* (0.2002)
Married	0.1809 (0.1975)	0.1277 (0.1973)	0.3434* (0.1954)	0.3455* (0.1961)
Other ethnic groups	−0.0840 (0.1030)	−0.0528 (0.1965)	0.1146 (0.0803)	0.3132 (0.2097)
High school	−0.1948 (0.1377)	−0.1595 (0.1470)	−0.3369** (0.1397)	−0.2591 (0.1619)
Some college	−0.3333** (0.1408)	−0.3031** (0.1480)	−0.4910*** (0.1410)	−0.4292*** (0.1552)
Graduate school	−0.8981*** (0.1409)	−0.8382*** (0.1806)	−0.7466*** (0.1418)	−0.5065* (0.2726)
Any kids in household	−0.2399** (0.0994)	−0.2310** (0.1005)	−0.1748** (0.0801)	−0.1731** (0.0803)
Born in U.S.	0.4496*** (0.1469)	0.3855*** (0.1461)	−0.0286 (0.1356)	−0.1968 (0.2108)
Any unearned income in past 12 months	0.6308*** (0.1333)	0.6212*** (0.1567)	0.4233*** (0.1157)	0.2812* (0.1667)
Had moved in past 12 months	−0.0162 (0.1046)	−0.0712 (0.1842)	−0.0836 (0.0964)	−0.2534 (0.1979)
MSA residency	0.1307 (0.0957)	0.1116 (0.1026)	0.0609 (0.0849)	−0.0353 (0.1312)
Constant	−0.5759*** (0.2228)	−0.0511 (0.7194)	0.2758 (0.2113)	1.6011 (1.0732)

Note: SMI = Serious Mental Illness.
Estimates for marginal effects of SMI, calculated for the statistically significant coefficients only, are in squared brackets.
Standard errors reported in parentheses.
*Statistically significant, $p \leq 0.10$.
**Statistically significant, $p \leq 0.05$.
***Statistically significant, $p \leq 0.01$.

being Hispanic, being married, and having received unearned income or attended graduate school were significantly and negatively related to mental illness while having changed residency in the last 12 months was significantly and positively related to mental illness. No other coefficient was significant for the male sample. For the female sample, similar to women aged 50–64, those who were between 35 and 49 were less likely to be mentally ill relative to women of ages 25–34. African-American women were less likely to be mentally ill relative to their white female counterparts. The results for the female sample also indicate that having been born in the U.S., or lived in a MSA was positively related to mental illness. The χ^2-statistics for the significance of the composite religiosity instrument in the first-stage equations were statistically significant for both samples ($p < 0.05$), indicating that the instrument was a good predictor of mental illness for both males and females.

4.2. Unemployment Equations

The estimation results for the unemployment equations are presented in Table 4, for both males and females. First, the Smith-Blundell test of exogeneity failed to reject the null hypothesis that mental illness was exogenous in the unemployment equations for both the male sample ($\chi^2(1) = 0.01$; $p = 0.92$) and the female sample ($\chi^2(1) = 2.12$; $p = 0.14$). This implies that univariate probit model can consistently estimate the effects of mental illness on employment. The results from the univariate model, our focus, are presented in the next paragraph.

The univariate probit findings indicate that men who were mentally ill had a significantly higher probability of being unemployed compared to men who were not mentally ill ($p < 0.01$). Specifically, the marginal effects find that mental illness increased the probability of being unemployed by more than 3 percentage points for the male sample. There was no significant relationship between mental illness and unemployment for the female sample. Numerous other coefficients were significant in the unemployment equations. For example, men or women who had more than a high school education as well as individuals who had children in their households were significantly less likely to be unemployed. Men or women who had received unearned income or moved in the past twelve months were more likely to be unemployed. Moreover, men from other ethnicities as a group comprising native Americans, non-Hispanic Asians and Pacific Islanders were more likely to be unemployed compared to White men; married men were less likely to be unemployed relative to non-married men. For the female sample, we found that women born in the U.S. were less likely to be unemployed compared to their female counterparts.

4.3. Workdays Skipped Equations

Analyzed for the employed only, the results should be interpreted as the impacts of mental illness on the number of workdays skipped in the past 30 days by the respondents who had a job. Coefficient estimates from the ZINB and the 2SQML models for the workdays skipped equations are presented in Table 5 for both males and females. As indicated earlier, we tested for the endogeneity of mental illness in the workdays skipped equation using the Smith-Blundell test of exogeneity. The null hypothesis that mental illness was exogenous in the workdays skipped equation was not rejected for either males ($\chi^2(1) = 0.006$; $p = 0.94$) or females ($\chi^2(1) = 0.55$; $p = 0.46$), hence we present the results from the ZINB specifications (columns 1 and 3 of Table 5) in the next paragraph.

The results from the standard ZINB model indicate that employees who suffered from mental illness skipped significantly more workdays during the last 30 days prior to the interview, relative to the workers who had no mental health problems. Marginal effects, reported in square brackets, suggest that mentally ill men and women skipped approximately 0.59 more workdays than the non-mentally ill during the past 30 days. In addition, we found that men and women who attended college skipped work significantly less relative to those with less than a high school education; individuals with kids in the household skipped significantly more workdays. In addition, men who were older than 34 years old skipped more workdays compared to those who were 25–34 years old; Hispanic men skipped significantly fewer workdays than White men. For the female sample, we found that African-Americans and those from other ethnicities (Native Americans, Pacific Islanders, and Asians as a group) skipped significantly more workdays compared to White women.

5. CONCLUSIONS AND SUGGESTIONS FOR FURTHER RESEARCH

The primary objective of this study was to examine the relationships between mental illness and labor supply, while conducting separate analyses for males and females for a richer examination of the mental illness/labor supply nexus. We used the 2001 National Household Survey on Drug Abuse, which included a new series of questions that allowed distinguishing individuals who were mentally ill based on DSM-IV criteria. The main findings of our analysis were that mentally ill men were more likely to be unemployed relative to non-mentally ill men, but no significant relationship between mental illness and unemployment was found for the female sample. For the workdays skipped specifications, we found that, regardless of gender, mentally ill individuals skipped significantly more workdays in the last 30 days prior to the interview compared to non-mentally ill individuals.

The findings of the present study are consistent with other studies on the labor market effects of substance use in the sense that mental illness had general negative effects on labor supply, but that men and women often differed in the significance or the magnitude of the effects (see Alexandre & French, 2001; Ettner et al., 1997; Mitchell & Anderson, 1989; Mullahy & Sindelar, 1990). While the study suggests that mental illness causes men and women to miss work if employed, in terms of employment transitions it appears to lead to job loss and unemployment for men only, perhaps because women would drop out of the market.

The study offers support for the expansion of mental health services as a means to improve quality-of-life and promote economic benefits. Programs that prevent mental illness or improve mental health may sustain or even enhance work force productivity through avoiding the negative consequences of mental illness on employment and work time. Previous studies suggest that these programs were generally cost-effective, could substantially improve mental health, and increased employment and job retention (Metsch et al., 1995, 1999, 2001; Schoenbaum et al., 2001; Wells et al., 2000). More importantly, treatment programs that closely meet gender-specific needs have been associated with superior treatment outcomes (Nelson-Zlupko et al., 1995; Schliebner, 1994). Although public health interventions may not lead to overall increases in employment, they may be justified on social welfare grounds. The negative effect of mental illness on workdays skipped is particularly important to employers who oversee and finance workplace intervention programs, including employee assistance programs. With the increase in workloads during recent years and the associated increase in the number of employees experiencing psychological problems related to occupational stress (Murphy, 1996), it may be economically efficient for employers to allocate more resources for EAP referrals, mental health insurance coverage, as well as organizational development and job design programs.

NOTES

1. Regarding the two-stage model, full-information maximum likelihood (FIML) also yields consistent estimates as long as the errors are normally distributed (Davidson & MacKinnon, 1993).

2. Since the number of excluded instruments (religiosity) is equal to the number of potentially endogenous variables (mental illness) in the present study, the model is exactly identified. Thus, tests for overidentifying restrictions of the instruments do not apply (Davidson & MacKinnon, 1993). But, bias in the coefficient estimates is greatly reduced for the exactly identified model (Angrist & Krueger, 2001).

3. The "unemployed" also includes individuals who were not in the labor force (not actively looking job), which represented only 0.8% in the male sample and 3% in the female sample. We estimated the unemployment equations while excluding those who were out of the labor force; the main results did not change.

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MENTAL ILLNESS: GENDER DIFFERENCES WITH RESPECT TO MARITAL STATUS AND LABOUR MARKET OUTCOMES

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1. INTRODUCTION

In a number of recent studies, it has been demonstrated that mental illness imposes real and large costs over and above the direct expenses of care and treatment. Each year in the U.S., 5–6 million workers between 16 and 54 years of age lose, fail to seek or cannot find employment as a consequence of mental illness. Among those who do work, it is estimated that mental illness decreases annual income by an amount between USD 3500 and USD 6000 (Marcotte & Wilcox-Gök, 2001). Similar results have been shown in a number of studies (Ettner et al., 1997).

The seriousness of these costs is highlighted by findings in the United States that 48% of adults have had a substance or non-substance abuse disorder over their lifetime and close to 30% experienced a disorder within the previous 12-month period. (Kessler et al., 1994). More than half of all lifetime disorders occurred in the 14% of the population who had a history of three or more comorbid disorders. Less than 40% of all with a lifetime disorder had ever received professional treatment. The most common disorders were major depression and alcohol dependence (Kessler et al., 1994).

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A compounding problem is that since mental illness appears to hit relatively young persons, it has a negative effect on investments in human capital (Berndt et al., 1998; Marcotte, 2000). This will in turn mean lower future earnings. These and a number of similar results are all based on surveys covering 5–8000 persons in the U.S. (see Ettner et al., 1997) and the information on mental illness is based on diagnoses established using self assessment of symptoms.

We take these findings based on survey data of the U.S. population as a point of departure for this study. These studies illustrate that the onset of a psychiatric disease may reduce the productivity of the ill. Depending on the severity of the disease, the availability of alternative means of support in the welfare system and the ability of the wage system to accommodate a drop in productivity, this may result in a loss of employment for some and in lower earnings for those still working.

However, our study differs from previous survey based research in at least four aspects: First, it is based on a representative sample of 5% of the entire adult population of Denmark, covering about 200,000 persons. This means, that there is no sample bias in the Danish data and it also means that no person can be too ill to be included, thus it is overcoming a potential bias in mail or telephone surveys as recognized by Ettner (2000). Second, the information on illness comes from hospital records and the information on social background, education and economic variables all come from administrative records. The reliance on hospital records means that we will have a much lower proportion of cases than the comorbidity survey and similar self-reported surveys. The administrative records contain the psychiatric diagnoses that resulted in hospital admission, admission and exit dates. About 12% of the Danish population experience an admission to a psychiatric hospital or ward before the age of 70. These 12% can be compared to 48% of the U.S. population with a diagnosable disorder over their lifetime from the National Comorbidity Survey (NCS) (Kessler et al., 1994). Of the NCS sample, 40% say that they were treated, suggesting about 19% of the U.S. population have been treated for a mental illness. Treatment in the NCS question includes both outpatient therapy and pharmacotherapy, as well as hospital admissions. Since the data from Denmark include only admissions to mental hospitals, these prevalence rates can be plausibly reconciled. Third, diagnoses are made by doctors and not by self-reporting. This should increase the quality of the diagnosis. Furthermore, all background information and information on earnings and employment come from reliable registers. Fourth, all individuals can be followed over time both before and after admission to hospital. This permits us to observe how social and economic connections for the ill change over a number of years and illustrates that mental illness is not a static concept. Instead, it is likely that the kind of measures used (earnings and employment) will pick up long-run effects of the disorder early on.

Though these differences make it more difficult to compare with the survey-based results, this type of investigation points at other important features of psychiatric disorders.

2. PREVIOUS RESEARCH

Most studies show that mental illness leads to reductions in labour earnings. Bartel and Taubman (1979) estimated an earnings loss of about 20% associated with poor mental health. Similarly, Benham and Benham (1981) estimated a 30% loss due to an indicator of psychosis. In both cases, data were not from self-assessments but from diagnostics records. Ettner et al. (1997) present an overview of similar results. There seems to be agreement on the sign of the impact but some disagreement on the effect of different diagnoses. The substance abuse literature contains reports of negative as well as positive effects of alcoholism on the employment rate and on income.[1] The same diversity of results is found for drug abuse.[2] Though there are good theoretical reasons to expect an adverse effect on employment, some studies show that this hypothesis is not supported unanimously, see Ettner et al. (1997) for an overview.

More recently, a number of papers using the U.S. NCS (National Comorbidity Study) have addressed the impact on employment and earnings. Ettner et al. (1997) estimate that having any mental illness reduces the likelihood of working with about 11 percentage points for women and somewhat lower for men. Furthermore, they estimate that earnings for those who work are reduced by 20–50%. Similarly, Marcotte et al. (2000) look at the effects of different diagnoses of major mental affective disorders. They find large but in general insignificant effects due to the relatively small sample. Depression is estimated to reduce earnings with about USD 6000. These findings are generally supported by other studies using other survey data.

One of the crucial topics in this study and in most of the literature is the direction of causation. One possibility is that the stress and anxiety associated with income losses and decreases in material living standards may trigger mental health problems. More recent research in the fields of social medicine and psychology indicates that the unemployment experience and the stress associated with it may change an individual's self-consciousness with ensuing identity crises and loss of time perspective. These strains may in turn weaken a person's resistance to somatic as well as psychic illnesses. Furthermore, workers are less prone to depression and other ills because of their opportunities for establishing social networks, achieving role satisfaction and gaining power. However, such beneficial effects may be offset by the detrimental impact of occupational stress. Likewise, it has been argued that the impact of employment on physical health is negative because of occupational hazards or the physiological effects of the work overload and performance pressure associated with paid employment. But there may also be an indirect effect through income and health insurance coverage, where lower income means a lower budget spent on health (Ettner, 1997).

The most straightforward hypothesis is that the disease itself lowers productivity, resulting in lower earnings and poorer employment prospects. In this scenario it is likely that health problems create problems in the labor market. Indeed these problems can then exacerbate one another, leading to a downward spiral that ends in hospitalization. This pattern makes it very difficult to disentangle the causation, because it all goes back to what triggered the first push and what is it that enables some people to overcome these traits and others to give in.

Some earlier attempts to test for causation have produced indirect evidence from longitudinal data sets by looking at the change in mental health upon re-employment. There are very few studies in which the level of mental health prior to becoming unemployed is known. However, it is difficult to disentangle the direction of causation by observation. One reason for this is that we are not directly observing the early stages of mental disorders. As a result, researchers use indirect methods such as instrumental variables (IV). In the above-mentioned studies by Ettner et al. and Marcotte et al., an instrumental variable method (IV) is applied using own past history and mental illness of parents as instruments.

In Ettner (2000) the predictive power of health measures on labour market outcomes is tested. She found that health has greater ability to predict employment than job characteristics conditional on employment. Several health measures were tested but functional limitations were the most important in predicting labour outcomes, indicating that better health leads to better jobs. Furthermore, the analysis suggests that the effects on labour market outcomes were not particularly sensitive to reverse causality.

Ettner (2000) is aware that the analysis is subject to limitations because separate identification of the effects of multiple health measures is not plausible and because the use of telephone interviews and written questionnaires may disproportionately exclude the unemployed and hence bias the estimated effects of health downwards.

Similarly, Ettner et al. (1997) estimated IV models, where all psychiatric illnesses were aggregated into one group. As instruments they use the number of psychiatric disorders exhibited by the respondent's parents and the number of psychiatric disorders experienced by the respondent before the age of 18. In all cases but one, the effect of having a psychiatric disorder became larger after instrumenting. The crucial assumption for the use of the proposed instruments is that these variables are not correlated with labour market outcome. This seems questionable, since psychiatric disorders may have a long-stretched impact on labour market behaviour.

Marcotte et al. (2000) also use family history as an instrument in their study of labour market effects of affective disorders. They find that the IV estimation changes the sign to the effect on income of depression (from positive to negative) and sharpens the effects on employment and earnings. So on one side there is no doubt that properly done IV estimation on cross section data is superior to OLS. However, it all rests on the ability to find good instruments for the onset of the disease which are not correlated with earnings and participation.[3] On the other side, the literature shows little indication that there is a serious causal effect going from labour market conditions to psychiatric disorders. In this paper, we will be circumspect about the true direction of causation, and will not try to disentangle the question. Instead we will throw light on the development of the effects for those hit by mental illness using the longitudinal structure of our data.

An important aspect of a psychiatric disorder is how it develops over time. Most studies date the onset of a mental disease to be relatively close to the effect on the labour market situation. In this study, we are able to investigate how employment and wages are affected several years before and after the disease actually develops into a diagnosed disorder requiring hospitalization. This is not only important in an assessment of the total costs to society of mental disorders, but is also useful as regards attempts to discover and treat diseases at an early stage as argued in Marcotte and Wilcox-Gök (2001). This is relevant information for policy makers, insurance companies, health plan administrators, private and public employers and all others who are engaged in some sort of human resource management.

Ettner et al. (1997) discuss this issue and conclude that "a portion of the wage effect may be mediated through work experience." And they continue: "Unless treatment is provided at the onset of the disorder, subsequent treatment cannot compensate for the loss of prior work experience that reduces current productivity". In this study, we are able to demonstrate that most mental disorders develop over a number of years before they get so serious that they are treated.

3. DATA

The data are constructed by merging two longitudinal data sets using a common identifier, the CPR-number.[4] The first data set is The Longitudinal Labour Market Register which consists of records from various Danish administrative registers on income, unemployment, education, work history along with basic demographic information. The register contains a 5% random sample of the adult Danish population (15–74 years of age) or 206,784 persons who are followed over the years

1976–1993. The sample is representative for the population between 15 and 74 years of age. The sample is not balanced, which means that we are observing most persons for the whole period, but a small number of persons disappears from the sample due to attrition (death and emigration) and a similar small number is added to the sample. The sample is supplemented with people turning 15 or immigrants, so that the representative nature of the sample is maintained. The data have been used extensively over the last 20 years and have proven to give a reliable picture of flows in the labour market, see Westergaard-Nielsen (1989). One reason for the reliability is that the underlying administrative data are updated almost every time a person approaches public authorities. The sample is updated annually, and people are followed for as long as they are alive, fulfilling the age requirement and still living in the country.

These data were merged with the Danish Psychiatric Case Register[5] using the CPR-number as the identifying key. The Danish Psychiatric Case Register has been described in detail by Munk-Joergensen and Mortensen (1997), and in short it covers all psychiatric inpatient facilities in Denmark. The register includes all admission and discharge dates and diagnoses according to the World Health Organization ICD8 and ICD10 of all psychiatric inpatient facilities in Denmark since 1969. There are no private psychiatric hospitals in Denmark, and all treatment is free of charge. Registration of outpatient treatment was established in 1995. The resulting sample consists of 5% of the population with psychiatric records on all those in the sample who have ever been admitted to a psychiatric hospital, which also means that we can distinguish between those who have been admitted (cases) and those who have not been admitted at the point of observation (controls). About 5% (10,400) of the total sample of 206,784 persons appear to have been admitted to a hospital due to a psychiatric diagnosis during the period 1976–1993.

We use three outcome variables, marital status, employment status and earnings. They all come from the registers. Marital status, defined as cohabitating or not, is determined by Statistics Denmark based on their comprehensive housing register. This register allows Statistics Denmark to identify if two persons of the opposite sex with a moderate age differential live in the same dwelling (apartment or one-family house). We could also have used the more conventional marital status, but that the prevalence of legal marriages is cohort specific in the period we observe. Cohabitation tends to give a picture of how easy it is to live together with the person who eventually gets ill.

In this study, we have defined employment according to the labour earnings of a person, i.e. if the person earns more than a threshold set to the basic tax allowance of about DKK 30,000 (about 4000 USD) in 1994-kroner then he or she is defined as employed.[6] If the person earns less we define him or her as not

Table 1. Summary Table of Participation in 1993.

	Labour Market Participation in % of All in Cohort					
	Never Admitted	One Admission	Admitted More Than Once	Total	Number	Of which Admitted
Women	0.57	0.36	0.22	0.55	103201	0.05
Men	0.63	0.43	0.25	0.62	103583	0.04
Total	0.6	0.39	0.23	0.59	206784	0.05
Age <25	0.47	0.37	0.23	0.47	36712	0.01
25 <Age <34	0.77	0.53	0.29	0.76	42313	0.03
34 <Age <44	0.8	0.57	0.32	0.78	39093	0.06
44 <Age <54	0.75	0.5	0.3	0.72	38773	0.07
55 <Age <64	0.46	0.23	0.15	0.45	26809	0.06
65 <Age	0.06	0.03	0.02	0.06	23084	0.06

employed. Earnings below the basic tax allowance means that no tax was paid. This definition is chosen to capture economic activity on the labour market instead of relying on a formal definition of attachment to the labour market that could cover various tax arrangements exploiting the basic tax allowance.

Data on earnings come from the tax register. This is generally believed to give a true picture of the earnings of individuals because it is based on self-reporting together with a comprehensive employer reporting system, where employers can only claim expenditures to labor if they declare the individual earnings of their labor.

The summary table (Table 1) shows that 5% have been admitted out of the total of 206,784 persons covered. The overall labour force participation ratio was 59% in 1993 (55% for women and 62% for men), 39% among people who at that time had been admitted once and 23% among people who at that time had been admitted more than once.

Employment is much lower among patients than among people who have not been admitted to a hospital and more than one admission reduces the labor force participation even more, and these patterns are similar for women and men. The second part of the summary table contains two time dimensions: a calendar time and age. The calendar dimension picks up changes to the practice of admitting people to hospital. It is known that first-time admission has decreased over the investigated period, as it has in most countries. It is also known from time-series studies that increases in the aggregate rate of unemployment is followed by more admissions, see Eriksson et al. (1998). The age dimension contains all age patterns in the admission probability.

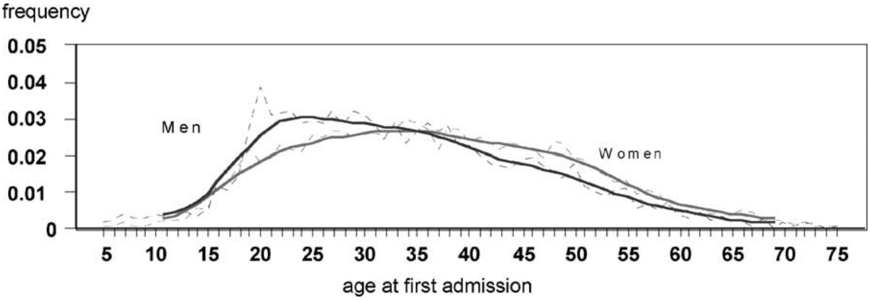

Fig. 1. Admission Rates by Age Group. *Note:* The dotted lines are actual age specific admission rates, and the thicker lines are 6 periods smoothed kernels using the technique in Silverman, 1986.

Figure 1 shows that first admission is age dependent with its highest frequency at 45 years of age. Peaks are found for men at the age of 24 and for women at the age of 35.

Diagnoses are also clearly related to age as shown in Table 2. Affective disorders and alcoholism are related to hospital admissions of people from the late 20s to the late 40s with median values about 38, while the other diseases are more prevalent among younger people. For all diseases, it is clear that women are hospitalized later in life than men.

Our main interest is, however, on patterns of employment and cohabitation before and after hospital admission for the ill. To understand these patterns we project all cases on the same time axis so that the first hospital admission is located in $t = 0$. The years previous to admission are located on the negative part of the axis and the years after on the positive axis. Even though the observation period covers 17 years, the position of the first hospital admission in this span of years

Table 2. Age at Maximum Frequency (Modus), and Median Age at First Admission for Groups of Diagnoses. Age at First Admission.

	One Admission	Admitted More Than Once	Total	Number
Affective	0.35	0.25	0.29	1086
Alcohol	0.37	0.25	0.3	4807
Non-organic and non psychotic	0.44	0.28	0.36	297
No main diagnosis	0.47	0.45	0.46	1006
Schizofrenia	0.27	0.13	0.16	274
Substance	0.24	0.18	0.21	585
Organic	0.23	0.11	0.16	197136
Never admitted	0.23	1	0.6	0

determines how many years before and after we may be able to follow in each case. A person admitted in 1993 could be followed in The Longitudinal Labour Market Register from 1980 to 1993, which is 13 years before admission. Similarly, a person being admitted in 1980 could be followed 13 years after admission. Only cases with first admissions are included. A first admission is defined as the date, where the individual appears in the Danish Psychiatric Central Register for the first time conditioned on that the person has not been admitted before the beginning of the register in 1969.[7]

To develop a group against which the employment patterns pre- and post-admission could be compared, we matched the ill with comparable persons who had not previously been admitted to a psychiatric hospital. For each point in time where we observe that a person has been admitted we randomly choose 10 control persons among those who have not been admitted, up to this point in time (including childhood and the period prior to 1980) and who belong to the same cohort and gender as the case person. Controls and cases will have the same characteristics with respect to the number of years we observe them before and after the selection. Therefore, we will expect that on average controls have an employment rate, a wage level and a marital status corresponding to that of the cases, had they not been sick. The selection of controls means that we can always compare persons who have the same time distance to admission (cases) or selection (controls).

For each point on this new time axis, we can calculate mean values of employment (and later of wage levels and cohabitation status). Asymptotic pointwise 95% confidence intervals for the labour market participation are calculated using the standard large-sample asymptotic distributions (see Agresti, 1990). The sample size suggests that these confidence intervals have good coverage properties.

The virtue of this method is that it is simple, highly intuitive and justified by the abundance of observations. The drawback is that we are not able to disentangle multiple causes. Another point is that our control group will contain people who will be admitted to hospital or who have the same symptoms as those admitted. This presents a potential downward bias of our results because our control group may itself have lower earnings, etc. due to hidden or not yet recognized illness.

In the following, we will present a number of graphs of the sampled cases and the comparable controls calculated as described above. Because disorders have different age and gender characteristics, we create different control groups for different disorders, gender and age groups.

4. FREQUENCY OF EMPLOYMENT

Figure 2 presents the rate of employment of cases and controls. A number of features are revealed in this graph. First, the standard errors are rather small, which is

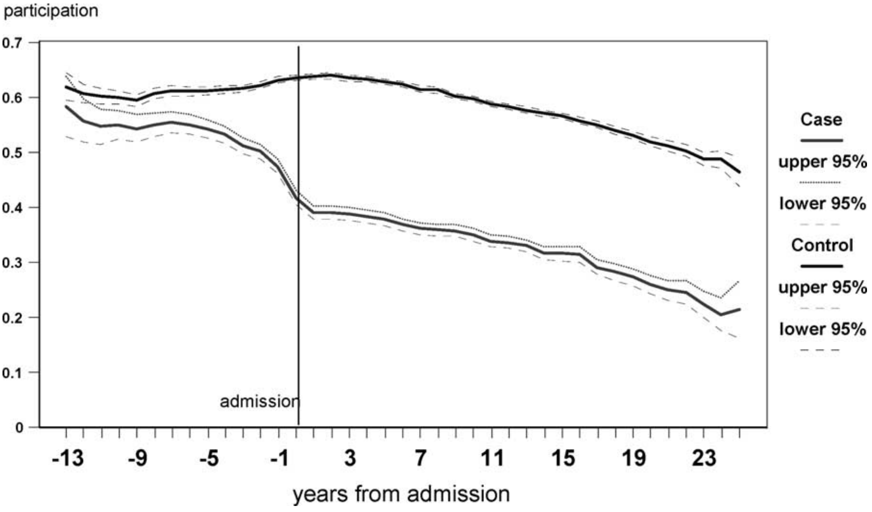

Fig. 2. Employment Rates of Patients and Controls. *Note:* The dotted lines are approximate pointwise 95% confidence limits around the estimated labour market participation.

to be expected from such a large sample. Naturally, this results in highly significant differences between the cases and controls. The significance is somewhat less in the two ends of the x-axis, where there are fewer observations due to the construction of the sample. Second, cases have a lower level of employment from the very beginning. Third, people already start becoming affected by their psychiatric disorder about 5 years before admission. Fourth, the decline of average employment during the last two years previous to admission is substantial. Fifth, after admission the employment rate seems to stabilize at an average level of about 2/3 of the comparable population. Sixth, this difference in the employment rate seems to hold for the rest of the observation period indicating that cases continuing in the labour market retire in the same way as the general population. The general decline of the participation rate is mainly due to age related retirement. Denmark has for all of the period under investigation had a non-health related early retirement plan from the age of 60 like many other European countries. Beside that there is also a health related plan with no lower age limit.

5. GENDER DIFFERENCES

In Fig. 3, we present a similar graph for men and women, separately. The controls show the traditional differences between gender, where men have an employment

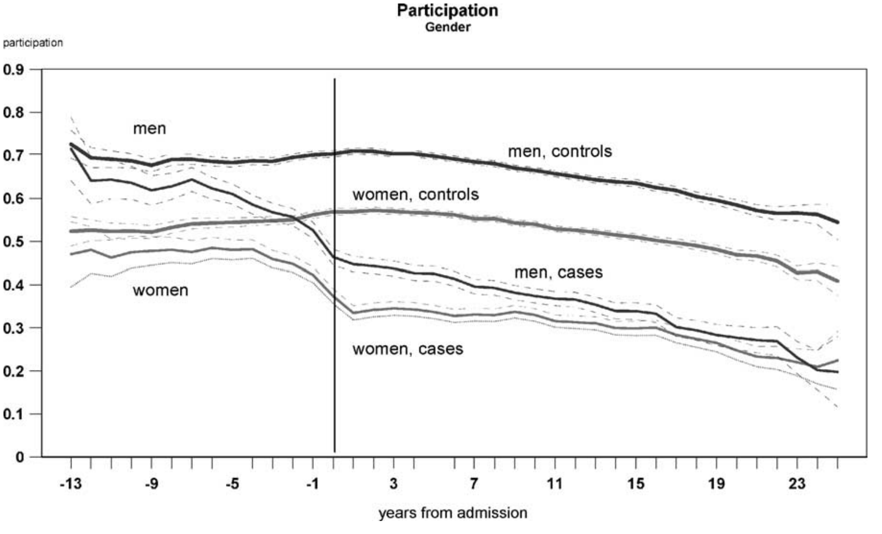

Fig. 3. Frequency of Employment for Cases and Controls for Men and Women. See the Note of Fig. 2.

rate about 15 percentage points higher than women. The graph for cases shows that both genders are almost equally affected by mental illness. In addition, the relative differences in Table 3 show that the relative reduction in employment is almost the same for men and women. It starts with a reduction of about 5–7 percentage

Table 3. Frequency of Employment for Cases and Controls, All, Females and Males.

	Years from Admission						
	−10	−5	−1	0	1	5	10
All							
Controls	0.60	0.61	0.63	0.63	0.64	0.63	0.60
Cases	0.55	0.54	0.47	0.42	0.42	0.38	0.35
Females							
Controls	0.52	0.54	0.56	0.57	0.57	0.57	0.54
Cases	0.48	0.48	0.42	0.37	0.37	0.34	0.33
Relative diff	0.09	0.12	0.25	0.34	0.34	0.40	0.39
Males							
Controls	0.69	0.68	0.70	0.70	0.71	0.70	0.66
Cases	0.64	0.61	0.53	0.46	0.46	0.43	0.37
Relative diff	0.07	0.11	0.25	0.34	0.35	0.39	0.44

points. One year previous to admission the difference is 25 percentage points; after admission the difference increases to 34 percentage points during the first year and it is 39–40 percentage points after 9–10 years. These relative differences are not different for gender.

6. THE AGE PATTERN

The age pattern of cases and controls reflects the expected life cycle participation for all three age groups in Fig. 4. The total effect of disorders is found to be almost the same for different age groups. Looking more closely at the youngest group of cases, we can see that this group clearly deviates from the controls because their participation rates do not increase as do those of the controls. The reason is most likely that an early onset of the disease has prevented many from acquiring normal levels of education and training, as concluded in Berndt et al. (1998) and Jayakody et al. (1996) both using U.S. samples. After having been admitted, their employment rate increases slightly after which time it stabilizes at about 50%. This is high compared with the older groups. One explanation for this pattern is that persons with early onset have sorted into jobs which better accomodate their disorders, while persons who become sick later in their lives are more likely to have started careers they are unable to continue, Berndt et al. (1998).

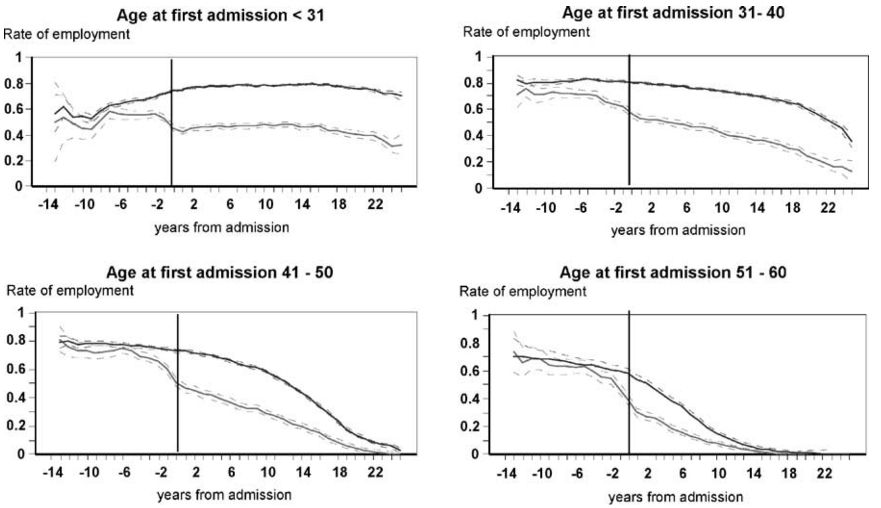

Fig. 4. Age Patterns in Participation.

Finally, the same graphs are shown for the older age group. It is remarkable that the immediate impact for the group older than 50 years of age is similar to the younger age groups. For all age groups, there is a drop in participation rates of about 25 percentage points. Of course the total cost to society is lower for the older groups because the drop in participation lasts for fewer years.

7. WAGES OF CASES AND CONTROLS

We will now turn to the development of the wages of the admitted persons who maintain employment. For this analysis, we will use the same case-control method as above. This means that for every time we observe a patient, we randomly pick 10 control persons, who have identical characteristics with respect to cohort and gender. For the years where these persons have a wage above the basic tax allowance, we calculate the mean for cases as well as for controls. For each case and control, we observe wages before and after the selection point. All wages are transformed into 1980 terms. At this point of the analysis, we neglect any differences due to differences in education and experience. By controlling for these factors, we could have achieved a higher precision. However, because controls are randomly selected from a large number of observations and the generally small standard deviations of the mean values for cases and controls indicate that we do not have to be overly concerned about this caveat (Fig. 5).

The controls clearly show the expected upward but decreasing slope due to the normal life cycle trajectory. The cases show lower earnings from the very first observations 13 years previous to admission. Here, the difference is about DKK 4,000 (1980-level) per year. However, the differences are only statistically significant from about 10 years before and up to 23 years after the first admission. While the average wage income increases for the controls, it decreases slightly for cases, and then drops precipitously at the time of admission and the subsequent year. At that time, the difference is about DKK 15,000. This difference is preserved almost for the rest of the observed time, though there is a slight catch-up effect, which could be due to non-random exiting behaviour among those who stop working among the cases.

Figure 6 presents similar graphs by gender. The basic pattern is the same for men and women, though the differences between cases and controls are much less for women than for men. So while the maximum reduction for women is about DKK 10,000 (15%) it is more than DKK 20,000 (22%) for men. Another difference is that women seem more able to catch up with their controls than men. Part of the explanation is undoubtedly that a somewhat smaller proportion of women are employed, and it is probably the most productive among those who continue

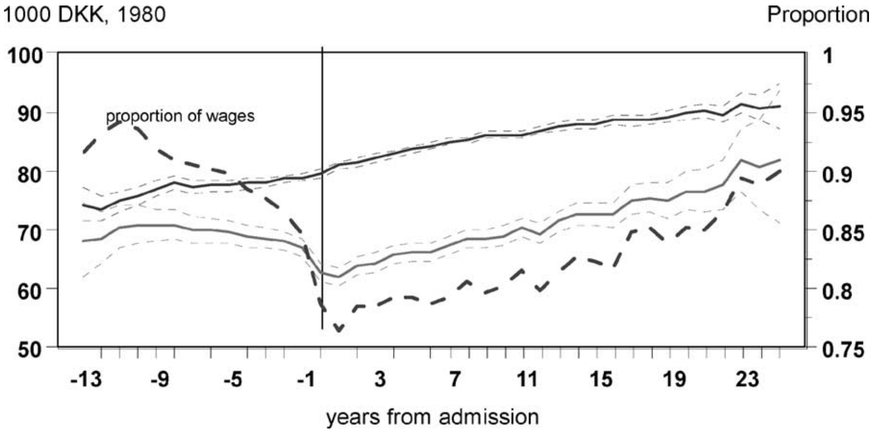

Fig. 5

Average Wages for Cases and Controls, and the Wages of Cases as a Proportion of the Wage of the Controls (Right-Hand Axis). *Note:* The dotted lines are describing the 95% confidence intervals. Pointwise 95% confidence limits for the wage income are based on the logarithm transformed wage.

working after having been admitted. Another reason could be that in general men have more human capital and the illness periods may therefore hurt them more than women. A third reason could be the differences in the gender composition of diagnoses and their different associations to income losses.

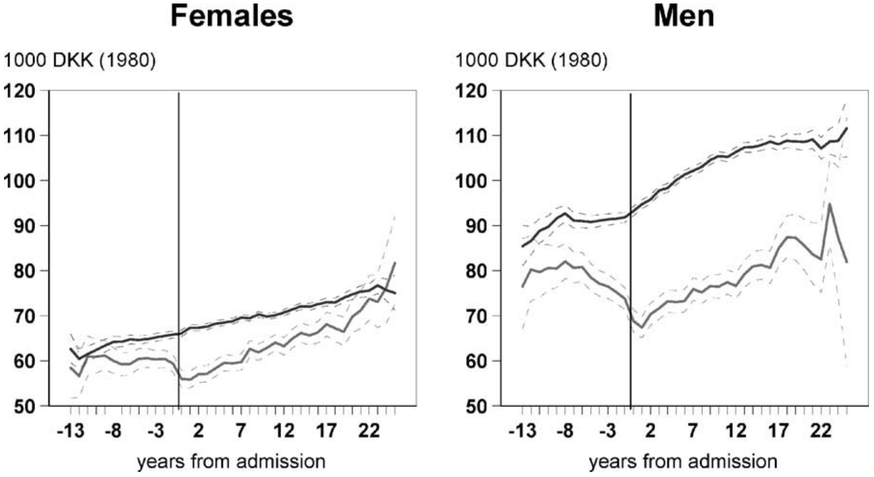

Fig. 6. Wage Income for Cases and Controls for Women and Men. See Note in Fig. 5.

A third feature of the gender differences is that the early differences between cases and controls are much larger for men than for women. This could be due to the earlier onset for men than for women, but could also be because of the gender composition of diagnoses.

The loss in wage for alcoholics is about DKK 25,000 and highly significant while it is about DKK 20,000 for organic diagnoses but with a higher variance. The loss for substance abusers is also high, but here the high variance among the cases indicates that some individuals in this group earn almost as much as the controls after treatment. The diagnosis group with the lowest difference between cases and controls is that of affective diseases.

8. COHABITATION

The last dimension we examine is the pattern of cohabitation over the life course for cases and controls, to better understand how illness might affect the formation and maintenance of families. First, Fig. 7 shows that cases actually have a different behavior than controls for years before admission. Cases have a significantly lower cohabitation rate 12 years prior to their eventual admission to hospital, and this rate is reduced dramatically beginning about 5 years before admission. After an initial dip at the time of admission, there is indeed an increase in cohabitation until about 3–4 years after. From then on, cohabitation rates among the ill follow the pattern of controls.

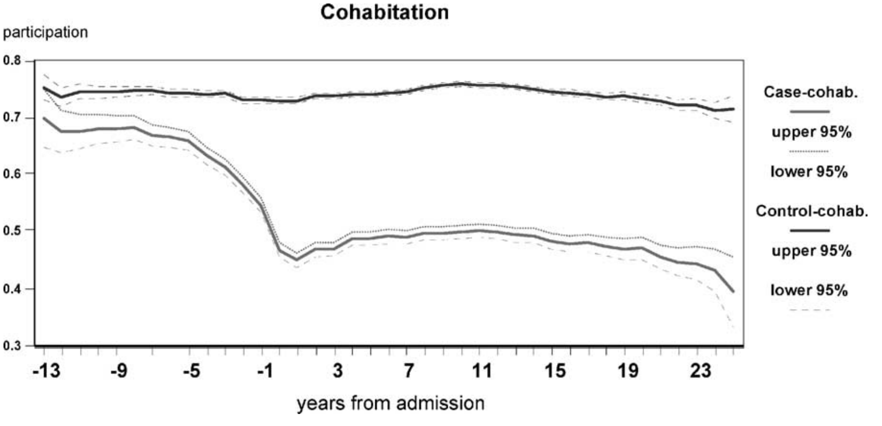

Fig. 7. Cohabitation for Cases and Controls.

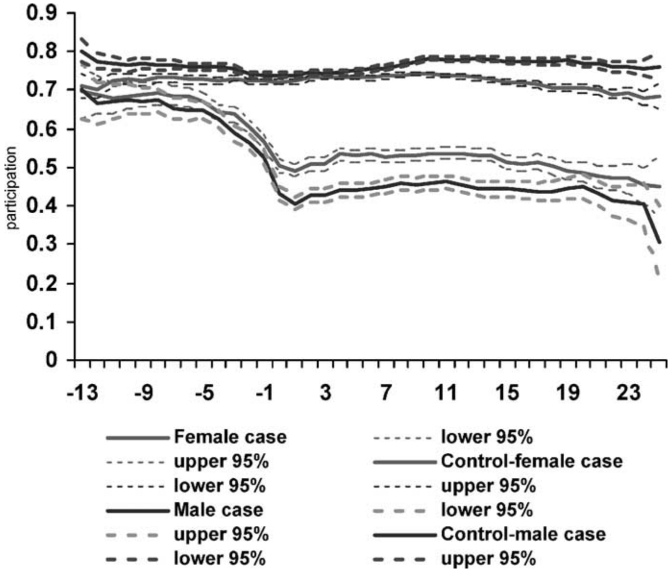

Fig. 8. Cohabitation for Gender for Cases and Controls.

This pattern is, however, somewhat different between men and women, as is demonstrated in Figs. 8 and 9. First, men who ultimately are admitted have a lower cohabitation rate from the start. For women this difference becomes significant at a later point of time (from about 8 years before admission). Furthermore, the

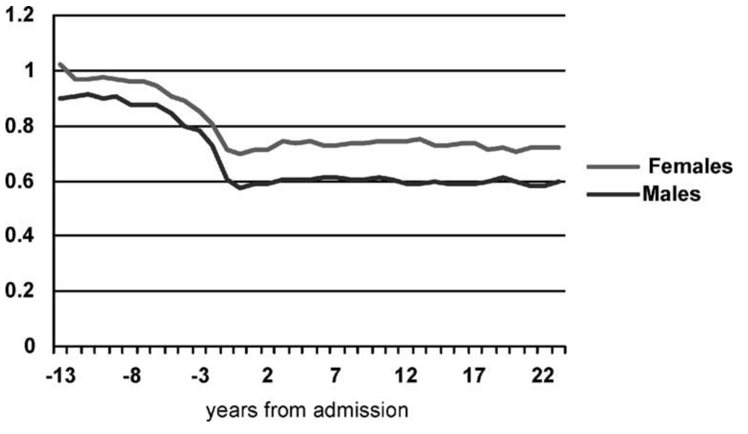

Fig. 9. Cohabitation Among Cases Compared to Controls for Men and Women.

reduction in cohabitation seems to be larger for men than for women. The dip and the subsequent rise in cohabitation are of similar magnitudes for men and women.

This relationship is clearly seen in Fig. 9, where cases are compared to controls. The loss for females subsequent to a hospital admission is about 25–30%, compared to about 40% for men. Furthermore, this loss appears to be permanent for men, while there is some improvement for women from about year 2 to 15 after admission for the first time.

9. DISCUSSION AND CONCLUSION

In this paper, we find that on average persons who are eventually admitted to a psychiatric hospital start behaving differently from their control groups with respect to employment well before they are actually admitted to a psychiatric hospital or ward for the first time. Differences are first observed as early as 10 years previous to admission with a slightly lower average employment rate, then a more substantial difference beginning about 5–7 years before admission. Among those who work, those later admitted to hospital earn about 90% of controls beginning 10 years before admission. This may be because some of the cases have lower educational levels than the controls because of an early onset of the disorder as found in Kessler et al. (1995) and in Jayakody et al. (1996). However, the distribution of income between cases and controls does not differ very much at this stage. Unlike Mullahy and Sindelar (1993) we find roughly the same effects on employment for men and women, but we find that earnings of women seem to be less vulnerable to mental diseases than men. One reason might be that on average women may hold jobs where one can function with a minor psychiatric disease.

In the years immediately preceding admission, both the rate of employment and the average wage fall more and more. The process seems to speed up from about 3 years previous to admission. In the year of admission, the employment frequency is reduced by 34% compared to controls, and earnings are reduced by 21% for those still working. From then on the situation stabilizes. This dip is probably related to the period spent in the hospital. Five years after first admission, the employment rate is reduced by 39% and earnings are still 21% lower than for the controls.

These results reflect a lower ability to maintain a job together with a lower productivity for a vast number of those admitted to a psychiatric hospital, even though about half of all the cases succeed in maintaining their earnings position. Our findings are, however, also influenced by the availability and the level of disability pensions and the praxis of the governmental bodies determining when people are eligible. If the disability pension is high, more people will probably try to leave the labour market compared to a situation with a low disability pension.

Since a Government board determines eligiblity, this board has an impact on how many will actually give up a working career. Their decisions are dependent on diagnoses. One example may clarify this latter point: Alcoholism is not found to have the most severe consequences for employment, but it is seen to have the most severe effects on earnings. This may be a result of the fact that the boards may be more reluctant to give pensions to people with severe alcoholism than other disorders.

It is hard to find comparable results in the literature, which are both nationally representative and using treatment-based definitions of disorders. The National Comorbidity Study (Kessler et al., 1995), seems to be an obvious candidate for benchmarking. There are, however, a number of differences that have to be taken into account. First, we can observe wage income directly and therefore disregard income from pensions etc. that might be related to the mental disorder in the NCS. We are using hospital admission data, where NCS uses diagnostic interview shedules. This has the effect that the NCS will include a broad range of disorders, with a broad range of severity, where we will only see the most severe cases which involve a hospital admission. On the other hand, the NCS exclude the most severe cases and all persons who are currently hospitalized.

Because of these features and differences, we will expect our results to show a more grave picture of the consequences of mental disorders compared to the NCS. Table 4 compares the results which generally support our priors.

Table 4 shows our results measured 5 years before, at the year of hospitalization and 5 years after, compared to NCS for men and women. This allows for a different degree of severity in the condition in the Danish data in order to make it more

Table 4. Comparison Between our Results and Results from the National Comorbidity Study. Reductions of Employment and Wages for Cases Compared to Controls. (The First Row Presents the Values for Controls).

	Reduction for People With Psychiatric Disorders									
	Employment			NCS		Wage income			NCS	
	−5	0	+5	Men	Women	−5	0	+5	Men	Women
No disorder/ controls	0.61		0.63	93.30	81.80	77566	79601	83643	33245	18586
All cases	0.11	0.34	0.40			0.10	0.21	0.21		
Men	0.11	0.34	0.39	0.17		0.14	0.26	0.27	0.10	
Women	0.20	0.34	0.40		0.17	0.06	0.13	0.13		0.29

Source: The NCS numbers are from Ettner et al. (1997).

comparable with the NCS, though it cannot overcome the selectivity problem. This is done for employment and for earnings.

First, it seems reasonable that our results 5 years before admission are comparable with the results from NCS. At later points on the time scale, we get much stronger effects than those found in NCS. Second, five years before we find relatively large differences between men and women with respect to reductions in participation, later on we find no differences in accordance with NCS. Third, we find that the earnings of men suffer much more than earnings for women at all points of measurement. The opposite result is found in the NCS. It is possible that the difference can be explained by different criteria for being included in the sample of NCS, or by differences in the U.S. and Danish labor markets. The larger differences in earnings for men in our sample are furthermore supported by the larger loss in cohabitation for men and indicates that the onset of mental diseases have larger effects for men. Nonetheless, more remains to be done to explain the difference in earnings loss between NCS and the Danish data.

For many cases, treatment is found to stabilize both the wage income for those working and their participation level, compared to a group of controls. Thus, we find that there are actually quite a number of persons who live and work with their mental disorder. Some diagnoses seem to be more difficult to combine with working than others.[8] Some of this "element of revival" is probably related to treatment. Though we do not go into detail here, it is clear that treatment is associated with a worsening of the situation and a subsequent stabilization of the effects. Undoubtedly, some of this effect is directly related to treatment because treatment at hospital will almost always prevent people from working. However, the stabilization is a real phenomenon. Though a number of persons are seen to overcome the disease, the general picture still is that the psychiatric disorder has marked many for the rest of their labour market career. This is also the case for cohabitation status, where loss rates are found to be between 25 and 30% for women and 40% for men. In a system as the Danish with good access to disability pension, this loss in personal relations can be seen as the most severe, since it is harder to substitute.

The loss to society can be measured as the sum of those who stop working due to their disease and the loss in income for those who continue working.[9] The income loss could be measured as the gap between the graphs of cases and controls in Fig. 5 times the number of admitted persons in the current year with different diagnoses. Similarly, the loss of income for those who continue working can be calculated as the mean loss multiplied with the number of persons. However, this calculation appears to neglect the fact that a number of persons experience almost the same symptoms but are never admitted to hospitals. These will actually weigh the control group downwards and thus contribute indirectly to a lower gap and to the negligence of a loss occurring before the first admission.

Though we have not contributed directly to the literature evaluating the direction of causation between mental disorders and labour market outcome, we have demonstrated that individuals eventually receiving treatment start behaving differently on the labour market 5–7 years before admission to hospital. And personal relations measured by cohabitation seem to start breaking down a couple of years before that. Men who eventually end up with an admission to a psychiatic hospital have a lower cohabitation level even 13 years before admission. Similar effects are observed for women, but not as many as 13 years prior to admission. These observations are clearly in accordance with the main view on causation presented in Ettner (2000) and Marcotte et al. (1999, 2000). But our findings could also be supportive of reverse causation. We have shown that the disorder develops over a period, and at some point it becomes so serious that employment and earnings suffer even before some of these persons later receive treatment.

These findings are of course important for the whole concept of onset of a mental disorder. There are very good reasons to believe that there is not a well-defined point in time where one can say that the disorder starts. This makes estimating the relationship between mental illness and labour market outcomes using cross-sectional data a very difficult task, indeed. In particular it complicates the job of identifying valid instruments for mental illness. For example, past history of "onset," an instrument commonly used in previous research, could clearly be problematic. The problem is that at the time persons with a mental disorder are identified in the cross-section, we cannot be certain that they are different from their observationally equivalent peers only in their current state.

NOTES

1. Thus, a negative impact on earnings and employment is found by Mullahy and Sinclair (1993), while Berger and Leigh (1988) find that drinkers earn more. Part of the diversity is probably because of differences with respect to the severity of alcoholism reflected in the measure.

2. Kaestner (1991) finds a positive effect on earnings while Gill and Michaels (1992) find a negative effect.

3. It seems to be even more demanding to find good instruments in panel data passing the overidentification test.

4. The CPR-number can be compared to a social security number though it is used more widely.

5. The research facilities of Statistics Denmark make it possible for us to use hitherto confidential data. The CPR-number is a ten-digit number that can be logically checked for errors. Each individual born in or immigrating to Denmark is assigned a number which is used uniformly across all registration systems in Denmark, including the population registers and registers of cause of death.

6. This threshold is so low compared to basic living costs that it is not sustainable without additional support from the welfare system.
7. Only first admissions are accounted for.
8. Westergaard-Nielsen et al. (2003) report large differences between diagnoses of schizophrenia and depression.
9. This cost concept is not in accordance with the so-called friction cost approach as described in Koopmanschap and van Ineveld (1992).

ACKNOWLEDGMENTS

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MENTAL HEALTH AND EMPLOYMENT TRANSITIONS

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1. INTRODUCTION

It is well known that mental health disorders cause substantial functional limitations and disability (Surgeon General, 1999). Less well known is the central role that mental health plays in economic disparities. The prevalence of depressive disorders is almost 3 times as high among individuals in the bottom 20% than among individuals in the top 20% of the income distribution, a much steeper gradient than for hypertension, heart disease, arthritis, chronic pain, or the number of medical problems (Sturm & Gresenz, 2002). In addition, individuals with mental disorders are less likely to have savings than individuals with physical health problems and the disparity widens with advancing age (Gresenz & Sturm, 2000).

This study analyzes the relationship between mental health and employment transitions among different employment states (employed, unemployed and out of the labor force) among adults in the U.S. While we know from numerous cross-sectional analyses that individuals with mental illness are less likely to be employed, only longitudinal data can provide insight into processes underlying this inverse relationship between employment and mental health. Does mental health affect job retention among the employed? Decrease labor force attachment? Reduce the probability of gaining new employment among the unemployed? How do labor market transitions differ between severe and more common disorders? We explore these questions in this paper. In keeping with a long tradition among labor economists of separately analyzing labor market outcomes for men and

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women, we allow for differences in the associations between mental health and labor market transitions with gender-specific analyses.

Research on the relationship between employment and mental health status extends back more than 60 years (Eisenberg & Lazarsfeld, 1938) and there is universal agreement that labor market status and mental health are closely related. Economists have generally focused on the adverse effects of mental health on employment or labor force participation (Alexandre & French, 2001; Ettner et al., 1997; Marcotte et al., 2000; Mitchell & Anderson, 1989; Ruhm, 1992), whereas psychologists and sociologists emphasize the mental health consequences of low quality jobs or unemployment (Dooley et al., 1996, 2000; Fergusson et al., 2001; Montgomery et al., 1999; Theodossiou, 1998; Viinamaki et al., 1993; Wilson & Walker, 1993). In line with traditional economics research, our focus is on the effect of mental health on labor market transitions. Data limitations – namely a lack of detailed information on the timing of changes in mental health and labor market status – prevent us from also exploring the reverse question (effect of labor market transitions on mental health).

Previous research has been handicapped by the absence of nationally representative longitudinal data. The studies with the largest sample sizes are cross-sectional and do not follow the same individuals over time, whereas other data collections focus either on health or on economic issues with limited measures in the other area. The most extensive longitudinal economic surveys contain no diagnostic mental health measures and at most some global psychological scale (e.g. Panel Study of Income Dynamics, PSID; National Longitudinal Survey of Youth, NLSY; Survey of Income and Program Participation, SIPP; Health and Retirement Survey, HRS). Most health surveys contain little information about an individual's labor market history or wealth (e.g. National Health Interview Survey, NHIS; Epidemiologic Catchment Area Study, ECA; Medical Outcomes Study, MOS; and National Comorbidity Study, NCS).

Because usually only one observation is available for an individual that measures mental health and labor market outcomes at the same time, economists have tried to measure causal directions using instrumental variables (Alexandre & French, 2001; Ettner, 1996; Ettner et al., 1997; Hamilton et al., 1997). It is difficult, if not impossible, to find unambiguously exogenous instrumental variables in this context and instrumental variable methods tend to be very sensitive to model specification. Longitudinal data do not automatically solve the issue of joint determination, but repeated observations on the same individuals help us to understand the temporal patterns of labor market outcomes and mental health. Dooley et al. (1994) estimate a two period model based on the ECA and find employment status affects depression, but no statistically significant effect in the other direction. Slade and Albers (2000) also use ECA data (from Baltimore) in a study of the effect of mental disorders on job loss. In contrast to the Dooley et al. study, Slade and Albers report a higher probability of job loss among individuals with symptoms of depression, but no similar effects for individuals with symptoms of panic, agoraphobia, or social phobia. Other longitudinal studies include Viinamaki et al. (1993), which tracks the mental health of workers following a factory closure (as well as a group of matched controls), and Dooley et al. (2000) and Fergusson et al. (2001), both of which analyze longitudinal data on young adults but with a focus on the causal effect of employment status on mental health.

By now, the data from the ECA are about 20 years old and predate dramatic changes in mental health care as well as significant changes in job requirements. Treatment rates for depression have increased with the increased availability of new medications, especially SSRIs for depression that came on the market in the late 1980s and atypical antipsychotics in the 1990s. These developments should have made it easier for individuals with mental illness to remain in the labor market. On the other hand, employment options have shifted from manual "blue collar" labor (which declined by roughly 20% in the past 2 decades) to administrative "white collar" jobs. While many of these new "white collar" jobs may not require high skill levels, they are likely to be more cognitively demanding than manual labor jobs and therefore may not be particularly well suited for individuals with mental illness. However, these occupational trends are also likely to have improved the opportunities for women relative to men. While there are no nationally representative data that would allow us to study how labor market outcomes for individuals with mental illness have changed over the past 20 or 30 years, this paper describes the situation at the beginning of the 21st century and provides a benchmark to evaluate future trends over time.

2. DATA

We analyze the first two waves of Healthcare for Communities (HCC), a national study to track the effects of the changing health care system for individuals at risk for alcohol, drug abuse, or mental health disorders. HCC links primary data collected from households, employers, and public agency administrators with secondary data sources (Sturm, Gresenz et al., 1999). The HCC household sample for Wave 1 was selected from a random sample of 30,375 adults who participated in the main component of the Robert Wood Johnson Foundation Tracking Initiative, the Community Tracking Study (CTS) household survey administered 15 months previously. The design of the CTS is described in more detail in Kemper et al. (1996) and at www.hschange.com. 14,985 respondents were selected for an expected completion of 10,000 interviews. The first wave obtained 9,585 eligible responses (a response rate of 64%), most interviews were conducted in 1998.

HCC Wave 2 re-interviewed all participants in HCC1 approximately 31 months after their HCC Wave 1 survey. HCC Wave 2 also added a new cross-section by interviewing CTS Wave 2 respondents not in HCC1. About half the interviews were conducted in 2000 and the other half in 2001. We use only the longitudinal component of the data – that is, respondents who are in both HCC 1 and 2. For the longitudinal component, 6,659 interviews were completed out of 9,585 potential respondents from HCC Wave 1, for a response rate of 65.9%. Because we are interested in labor market transitions, we subset the data to individuals between the ages of 19 and 60 at the time of the HCC Wave 1 interview (who therefore would be before typical retirement age in Wave 2) with complete information about labor market status in each time period ($n = 5211$). We separately analyze individuals with severe mental illness ($n = 144$), such as bipolar depressive or psychotic disorders, because these disorders are qualitatively different from the much more common unipolar depressive and anxiety disorders. Severe mental illness often starts in late adolescence or early adulthood, whereas depression and anxiety tends to peak in early middle age. Mechanic et al. (2002) have recently provided employment statistics for severely mentally ill from 3 surveys. The sample size of individuals with severe mental illness is too small for a detailed analysis of employment transitions. We therefore present only a descriptive table of employment transitions for these individuals.

The HCC household survey covers demographic information, health and daily activities, mental health, alcohol and drug use, medications, use and quality of behavioral health care, general health insurance, insurance coverage for behavioral health care and prescription medication, labor market status, income, wealth, and life difficulties. Mental health is measured with the MHI-5 (Wells et al., 1996) as well as with screening questions for particular disorders. Questions to assess risk for generalized anxiety disorder (GAD) and depression are based on the CIDI-SF (Kessler et al., 1998). Panic, dysthymia, mania, and substance abuse screens are derived from the stem items of the World Health Organization's (WHO) Composite International Diagnostic Interview (CIDI, WHO, 1990, 1997). HCC also had a brief screen for psychotic disorders. While these screening items are more specific to clinical disorders (and often also more sensitive) than global scales of psychological distress, they are not the same as a full diagnostic assessment. Therefore, whenever we refer to a specific disorder, it is best interpreted as "high likelihood of having" this disorder. A consequence of the possible misclassification of individuals based on these screening items is that the contrast of the groups with and without illness is not as clear as it would be with clinical diagnoses and the estimated effects will be smaller than the "true" effects, at least in descriptive statistics.

The economic data collected include the current employment status of the respondent. Respondents were allowed to describe their labor market status with multiple categories (e.g. student and working or homemaker and working). In assigning respondents to one of three states (employed, unemployed or out of the labor force), priority was given to any indication that an individual was doing any work for pay or wanted to do work for pay. Thus, only those homemakers or students who reported neither working for pay nor looking for work for pay were considered out of the labor force.

The HCC trades off some level of clinical detail in order to be able to measure a broad set of economic outcomes. For example, HCC does not collect information about the history of respondents' or parental mental health (including onset and number of previous episodes of mental health disorders). Similarly, economic outcomes are less detailed than in purely economic studies. For example, there is no employment history battery. Compared to clinical studies, where follow-up periods are typically measured in weeks or a few months, the HCC spans multiple years, which is similar to other social science surveys. The average time between Wave 1 and 2 interviews was 31 months, with a range from 21 to 47 months. The time span of the HCC is useful for studying economic phenomena such as job loss or job gain that are likely to change more slowly compared to outcomes such as functionality, productivity or absenteeism.

Despite the relatively large sample size of the HCC, power considerations are a factor in this analysis. Analyzing broader social outcomes, such as health care costs or labor market, income, or wealth changes, requires sample sizes that are much larger than sample sizes for clinical trials (Sturm et al., 1999). While meaningful differences in clinical symptoms or quality of care can often be detectable with sample sizes of a few dozens, such sample sizes are not sufficient for social outcomes for three reasons. First, the distributions of several measures, especially income and wealth, do not have a ceiling like many psychometric measures, and are very skewed. Second, the heterogeneity of the samples in real life studies will have a stronger effect on social outcome measures than on the clinical measures that an intervention focuses on, increasing both the variance of outcome measures and possibly also reducing the effect sizes that should be expected. Finally, the interpretation of what constitutes a meaningful change differs. For interventions, an absolute change of 20 percentage points on a measure of quality of care may be considered a moderate effect, yet for a change in labor market participation at the population level, such a difference would be dramatic. Unfortunately, it is unlikely that longitudinal data collections much larger than HCC would be funded.

3. METHODS

The potential ways in which health affects employment decisions have been described before (e.g. Mullahy & Sindelar, 1990) and include that health may affect the wage rate that an individual earns, individuals' valuation or utility of time spent in leisure or work, the accumulated assets or wealth that an individual has, and the rate at which individuals discount future time periods.

Our primary dependent variable is a measure of individuals' labor market status change between the two waves. We distinguish three employment states – employed, unemployed, or out of the labor force (OLF) – and we analyze each of the nine potential transitions (from each of the three employment possibilities at Wave 1 to each of the three at Wave 2) for men and women separately.

For multivariate analyses, we use a multinomial logit regression, subset by gender and baseline labor market status (working, unemployed, out-of-the labor force). The probability of an employment transition can be written:

$$\Pr(\mathrm{ES}_t|\mathrm{ES}_{t-1}) = \alpha + \gamma\mathrm{MH}_{t-1} + \beta X + \varepsilon$$

where ES is employment status, MH is mental health status, and X is a vector of demographic and other variables that affect individuals' decisions about whether or not to begin, end or otherwise change employment. Each regression includes controls for age, race, education, marital status, number of dependents, citizenship and physical health. Physical health is measured with variables indicating whether an individual has one chronic condition (out of a possible list of 17 such conditions), two chronic conditions, or three or more conditions (versus no chronic conditions).

The primary explanatory variable is the presence of a likely mental health disorder. Mental health disorders, like physical health conditions, span a wide range and the relationship between employment and specific conditions is likely to vary. But, specificity in disorders in analysis must be balanced against statistical power considerations. The balance we strike is to analyze the relationship between a broader group of disorders and specific labor market outcomes. We distinguish common depressive and anxiety disorders (including generalized anxiety disorder, panic, dysthymia, and major depressive disorder) from more severe disorders (psychotic and bipolar disorders).

The presence or absence of a disorder is measured with screening questions about symptoms during the 12 months prior to the interview date. Individuals are classified as having a disorder in a particular wave if they have had symptoms during the year prior to interview. Mental health disorder at Wave 1 should not be affected by job transitions that occur between Waves 1 and 2 (unless the job transition is anticipated and there is an anticipatory effect on mental health). Thus, coefficient on the indicator of mental health disorder at Wave 1 should largely capture the causal effect of mental health on future labor market transitions.

4. RESULTS

Table 1 provides the percentage of men and women with no mental health disorder, anxiety or depressive disorder, or bipolar or psychotic disorder (severe mental illness) who are employed, unemployed or out of the labor force (OLF). Compared to women with no mental health disorder, women with an anxiety or depressive disorder have lower rates of employment (difference = 10.6 percentage points, $p < 0.001$), and higher rates of being unemployed or out of the labor force (difference = 4.1 percentage points, $p < 0.06$; difference = 5.7 percentage points, $p < 0.02$, respectively). Similarly, men with an anxiety or depressive disorder are less likely to be employed (difference = 12.5 percentage points, $p < 0.0001$) and more likely to be out of the labor force (difference = 11.2 percentage points, $p < 0.0001$). However, rates of being unemployed were not statistically different between men with/without an anxiety or depressive disorder. For both men and women, those with severe mental illness are much more likely to be unemployed or out of the labor force and much less likely to be employed.

Table 2 presents the transition probabilities between the three employment states (employed, unemployed or OLF) for women with a depressive or anxiety

Table 1. Employment Status Among Individuals with No Mental Health Disorder vs. Anxiety or Depressive Disorder or Severe Mental Illness (Bipolar or Psychotic Disorder).

	No Mental Health Disorder (%)	Anxiety or Depressive Disorder (%)	Severe Mental Illness (%)	All (%)
Women				
Employed (*n*)	79.0 (1910)	68.3 (461)	49.5 (46)	75.8 (2417)
Unemployed (*n*)	6.7 (165)	11.5 (78)	24.7 (19)	8.2 (262)
OLF (*n*)	14.4 (338)	20.2 (151)	25.8 (20)	16.0 (509)
Total	100 (2413)	100 (690)	100.0 (85)	100 (3188)
Men				
Employed (*n*)	90.9 (1486)	78.5 (237)	42.1 (31)	33.7 (1754)
Unemployed (*n*)	4.1 (69)	5.3 (18)	27.2 (10)	1.9 (97)
OLF (*n*)	5.0 (98)	16.2 (56)	30.8 (18)	3.3 (172)
Total	100 (1653)	100 (311)	100.0 (59)	39 (2023)

Table 2. Employment Transitions Among Individuals With Anxiety or Depressive Disorder vs. No Mental Health Disorder.

Wave 1: Employment Status	N	Wave 1: Mental Health	Wave 2: Employment Status			Total
			Employed	Unemployed	OLF	
Women						
Employed	1910	No disorder	90.1	4.3	5.6	100.0
	461	Anx/dep disorder	85.3	4.5	10.2	100.0
		Difference	4.9**	−0.2	−4.7**	
Unemployed	165	No disorder	47.6	32.7	19.7	100.0
	78	Anx/dep disorder	29.7	28.0	42.4	100.0
		Difference	17.9**	4.8	−22.7**	
OLF	338	No disorder	23.6	10.3	66.2	100.0
	151	Anx/dep disorder	23.5	26.2	50.3	100.0
		Difference	0.1	−15.9***	15.8**	
Men						
Employed	1486	No disorder	94.1	3.3	2.6	100.0
	237	Anx/dep disorder	92.4	3.1	4.4	100.0
		Difference	1.7	0.2	−1.9	
Unemployed	69	No disorder	71.5	12.3	16.2	100.0
	18	Anx/dep disorder	41.1	28.7	30.1	100.0
		Difference	30.4**	−16.4	−13.9	
OLF	98	No disorder	36.8	5.1	58.0	100.0
	56	Anx/dep disorder	19.8	12.0	68.2	100.0
		Difference	17.1	−6.9	−10.2	

**$p < 0.05$.
***$p < 0.01$.

disorder and for those with no probable mental health disorder. "Row" percentages are reported which indicate the percentage of individuals in a given employment status at Wave 1 who transition to each employment status in Wave 2. For example, among women with no mental health disorder who are employed at Wave 1, 90.1% remain employed at Wave 2.

There are notable differences in transitions between women who have or are free of anxiety or depressive disorders. Women with an anxiety or depressive disorder who are employed in Wave 1 are less likely to continue to be employed (difference = 4.9 percentage points, $p < 0.05$). Most women with anxiety or depressive disorders who transition out of employment leave the labor force entirely, rather than become unemployed: Transitions to OLF from employment among women with an anxiety or depressive disorder occur at nearly twice the rate among women with no mental health disorder (10.2 vs. 5.6%).

Among women who are unemployed at baseline, the proportion remaining unemployed at Wave 2 is statistically similar among those with/without an anxiety or depressive disorder (32.7 vs. 28.0%). But, the transition path out of unemployment varies dramatically depending on mental health status: Women with an anxiety or depressive disorder are more likely to stop job search (move from unemployed to OLF, 42.4 vs. 19.7%) while those with no mental health disorder are more likely to become employed (47.6 vs. 29.7%).

Transitions also vary among women out of the labor force. Women with an anxiety or depressive disorder who are OLF are more likely to re-enter the labor force (become either employed or unemployed). Among those with a disorder, 49.7% rejoined the labor force (50.3% remained OLF), compared to only 33.8% of those with no disorder who transitioned either to work or to seeking work (with 66.2% remaining OLF). Compared to those with no disorder, transitions were more common to unemployment among those with an anxiety or depressive disorder (26.2 vs. 10.3%).

Thus, we observe substantially more churning in labor force participation among women with an anxiety or depressive disorder. Transitions out of the labor force from employment or unemployment and transitions back into the labor force among those out of the labor force are more common among women with a disorder compared to among those without a disorder. Women with an anxiety or depressive disorder who are in the labor force thus exhibit weaker labor force attachment compared to others, but those who are out of the labor force show stronger resiliency in labor force participation compared to others.

Transitions among employed men vary little by mental health status, but there is significant variation among men who are OLF or unemployed. As with women, unemployed men with an anxiety or depressive disorder are much less likely to find employment compared to men with no mental health disorder (41.1 vs. 71.5%). Unemployed men with an anxiety or depressive disorder appear more likely to remain unemployed or to leave the labor market compared to men with no mental health disorder, but while the differences are large in size, they are not statistically significant with only 18 men with an anxiety or depressive disorder who are unemployed at baseline. Transitions from OLF to employment appear less likely and transitions from OLF to unemployment more likely among men with an anxiety or depressive disorder, though again, despite their relatively large magnitude, the differences are not statistically significant.

For men, the power to detect meaningful differences in transitions is much lower compared to women. There are fewer than half as many men as women with anxiety or depressive disorder in the sample and *ex post* power calculations show power of less than 0.5 for all transitions and less than 0.2 for several transitions. Thus, for men, the data are not sufficient for consistently capturing differences in

Table 3. Effect of Anxiety/Depressive Disorder on Labor Market Transitions Multinomial Logistic Regression Results (Women Only).

Regression	Sample Included	Employed RRR (95% CI)	Unemployed RRR (95% CI)	OLF RRR (95% CI)
(1)	Employed	Omitted	0.95 (0.52–1.76)	1.92** (1.14–3.21)
(2)	Unemployed	0.81 (0.27–2.38)	Omitted	2.38* (0.86–6.60)
(3)	OLF	1.89 (0.90–3.95)	2.74** (1.20–6.26)	Omitted

Note: All regressions control for age, race, marital status, education, number of dependents, citizenship, and physical health.
*$p < 0.10$.
**$p < 0.05$.

these specific transitions between those with and without an anxiety or depressive disorder. Given the limited power for men, we restrict our regression analysis of the more refined transitions to women.

The results of multinomial logistic regressions for more specific employment transitions among women appear in Table 3. These regressions were performed separately for individuals grouped by employment status at Wave 1. The omitted category in each case was no change in employment status. Relative risk ratios and 95% confidence intervals around those ratios are reported (results for other control variables are not shown). The regression analysis confirms the finding from the descriptive statistics: The fragile labor market attachment of women with mental health disorders who are in the labor force, and the resiliency of labor market attachment among those who are out of the labor force, is not a consequence of other sociodemographic factors, but most likely a consequence of mental health.

The results show that the separation of job loss into transitions from employment to unemployment and transitions from employment to OLF is important for women (regression 1). Women who have an anxiety or depressive disorder at Wave 1 have nearly twice the relative risk of switching from employment to labor force non-participation. On the other hand, women with and without anxiety or depressive disorders have statistically similar risks of transitioning from employment to unemployment. Regressions (2) and (3) show that mental health disorders substantially increase the relative risk of leaving the labor market among the unemployed, or becoming unemployed for women initially out of the labor market. Among those OLF, women with an anxiety or depressive disorder are more likely to try to rejoin the labor force (become employed or unemployed; results not shown, relative risk ratio $= 2.3$, $p = 0.010$).

Finally, Table 4 shows labor market transitions for the small group of individuals with probable bipolar depressive or psychotic disorders. This table has the same

Table 4. Employment Transitions Among Individuals with Severe Mental Illness vs. No Mental Health Disorder.

Wave 1: Employment Status	*N*	Wave 1: Mental Health	Wave 2: Employment Status			Total (%)
			Employed (%)	Unemployed (%)	OLF (%)	
Women						
Employed	1910	No disorder	90.1	4.3	5.6	100.0
	46	SMI	77.9	14.1	7.9	100.0
		Difference	12.2	−9.8	−2.4	
Unemployed	165	No disorder	47.6	32.7	19.7	100.0
	19	SMI	15.3	53.3	31.5	100.0
		Difference	32.3***	−20.5	−11.8	
OLF	338	No disorder	23.6	10.3	66.2	100.0
	20	SMI	29.4	2.7	67.8	100.0
		Difference	−5.9	7.5**	−1.7	
Men						
Employed	1486	No disorder	94.1	3.3	2.6	100.0
	31	SMI	89.5	4.6	5.9	100.0
		Difference	4.6	−1.3	−3.3	
Unemployed	69	No disorder	71.5	12.3	16.2	100.0
	10	SMI	72.8	22.0	5.1	100.0
		Difference	−1.3	−9.7	11.0	
OLF	98	No disorder	36.8	5.1	58.0	100.0
	18	SMI	10.5	16.9	72.6	100.0
		Difference	26.3**	−11.8	−14.5	

**$p < 0.05$.
***$p < 0.01$.

structure as Table 2, which focused on individuals who were likely to have had unipolar depressive and anxiety disorders. Results are very imprecise with such small samples so the results should not be over-interpreted. Nevertheless, it appears that severe disorders reduce the probability of retaining employment among employed women (not statistically significant), substantially lower the probability of transitioning to employment for unemployed women ($p < 0.01$), and decrease the probability of beginning job search (transitioning to unemployment) among women initially out of the labor force ($p < 0.05$). Among men initially out of the labor force, those with who are severely mentally ill are less likely to transition to employment (10.5 vs. 36.8%, $p < 0.05$) and appear to be more likely to begin the search for employment (not statistically significant). Unemployed men with severe mental illness appear more likely to remain unemployed (22.0 vs. 12.3% not statistically significant).

5. DISCUSSION

Many studies have shown that individuals with mental illness are less likely to be employed. This study adds to previous research by analyzing relatively new national data that provide insight into the associations between mental illness and employment transitions. We distinguish three employment states (employed, unemployed, and out of the labor force) and find different patterns for men and women.

There is a clear and statistically significant relationship between an episode of an anxiety or depressive disorder and later employment transitions among women. In particular, among those with an anxiety or depressive disorder, we observe substantially more churning in and out of labor force participation. Women who experience an episode of an anxiety or depressive disorder and are in the labor force, either employed or unemployed, are more prone to leave the labor force compared to others. Among women who are out of the labor force, those with an anxiety or depressive disorder are more likely to later try to rejoin the labor force. The differences in transitions between individuals with no mental health disorder and with an anxiety or depressive disorder are not accounted for by socio-demographic differences.

Most transitions for men also vary by mental health status, although our sample sizes are too small to provide precise estimates. Nevertheless, it appears that there is little difference by mental health status in transitions among employed men (in contrast to women). On the other hand, unemployed men with anxiety or depressive disorders are less likely to find a job. Men with an anxiety or depressive disorder who are OLF appear less likely to transition to employment and more likely to transition to unemployment compared to men with no mental health disorder; but, while the magnitude of these differences is large, the sample size is too low to achieve statistical significance.

There are many questions that we cannot address with this data. Arguably the most salient would be an analysis of whether employment transitions predict the onset of an episode of depression or anxiety. Other important questions for future research include whether the effect of mental illness on transitions varies with demographic factors such as individual's age or race.

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GENDER-SPECIFIC PATTERNS OF EMPLOYMENT AND EMPLOYMENT TRANSITIONS FOR PERSONS WITH SCHIZOPHRENIA: EVIDENCE FROM THE SCHIZOPHRENIA CARE AND ASSESSMENT PROGRAM (SCAP)

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1. INTRODUCTION

Schizophrenia is a profoundly disabling chronic mental disorder with an estimated annual prevalence rate of about 1.3% for the U.S. population age 18 to 54 (USDHHS, 1999). One reason it is so disabling is that onset typically occurs in early adulthood, impacting on a range of life experiences that influence later employment, such as completion of schooling, and early-career experiences at work. For most people, successful navigation of these experiences creates a solid foundation for later career advancement through the development of work skills and social supports.

Though the onset of schizophrenia often results in disruption of these influential experiences for both men and women, evidence of gender differences in age of disease onset (later for women), symptom expression, and response to treatment suggests less adverse effects on employment outcomes among women.[1] Goldstein

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(1988) reviewed a large number of studies that followed up patients after their first admission for schizophrenia, and reported a consensus that "women experienced fewer rehospitalizations, shorter hospital stays, better social and work functioning, better response to neuroleptics, lower relapse rates, and less severe psychopathological outcomes . . ." (p. 684). Goldstein also reported the results of her own analysis of a 10-year follow-up of 90 patients enrolled after a first or second hospitalization. These confirmed the earlier studies, showing significantly fewer hospitalizations and days in the hospital for women over both 5-year and 10-year follow-up periods. More recently, a 10-country comparative study (Jablensky et al., 1992) reported significantly more favorable outcomes for women along four of the seven outcome dimensions that were studied: pattern or course of the disorder, percent of follow-up period in psychotic episodes, percent of follow-up in complete remission, and percent of follow-up with unimpaired social functioning.

The quantitative impact of these gender-related differences on employment outcomes is not entirely clear. Those few studies that have examined gender differences in employment for persons with schizophrenia have found that women are less likely to be employed. Although this result parallels a lower rate of employment among women in the general population, the usual explanation for this differential, i.e. differences in domestic responsibilities, may be less applicable for persons with schizophrenia. These persons typically report lower rates of marriage and parenting, presumably resulting from the early age of disease and disability onset. In the only study that directly compared gender effects for persons with serious mental illness and those with no mental illness (Mechanic et al., 2002), the researchers found a much stronger negative effect of gender on the odds of being employed for females with *no* mental illness than for females with serious mental illness.

In the present study, we undertake a more detailed analysis of gender differences in employment patterns and employment transitions for persons with schizophrenia. The database for our analysis is the Schizophrenia Care and Assessment Program (SCAP), a recently completed large-scale survey and data collection effort from six geographic locales in the U.S. We begin by presenting background information on the nature of schizophrenia and reported employment rates for persons with the disorder. We then describe the SCAP database and our study population. The next two sections describe gender-specific employment patterns and present a cross-sectional analysis of similarities and differences by gender in the determinants of employment status for this population. This is followed by two sections that present descriptive data and analysis on gender differences in employment stability over time and on transitions into and out of employment. The final two sections of the paper discuss our results, conclusions, and directions for future research.

2. BACKGROUND

Schizophrenia is characterized by a wide variety of associated symptoms.[2] One group, referred to as positive symptoms, includes unpredictable and bizarre social behavior, delusions, hallucinations, difficulty in carrying out goal-directed behavior, catatonic behavior, and disorganized speech and thought. A second group, referred to as negative symptoms, includes apparent disinterest in goal-directed behaviors, lessening of fluency and productivity of speech, and flattened affect. Cognitive difficulties are also often observed in persons with schizophrenia, including problems in information processing, abstract categorization, attention, memory, visual processing, and cognitive flexibility. It is common for symptom levels to vary over time, particularly for the positive symptoms; negative symptoms fluctuate less but are more persistent. Over the long-term course of the illness, most people do not ever return to their pre-morbid level of mental functioning and about 10% remain severely ill for extended periods of time. The majority of patients, however, do experience significant improvement in functioning over time relative to their condition during the initial acute phase of the disorder.

Persons with schizophrenia face many serious challenges in getting and/or keeping a job. Deficits in cognitive processing, judgment, attention, and decision-making substantially impair abilities to carry out and complete even simple job tasks. Difficulties in social interactions and communication with co-workers and supervisors also adversely affect job performance. While some effects of the disorder can often be ameliorated by treatment interventions (such as antipsychotic medications and supportive psychotherapy), recurrence of acute symptomatic episodes is common. Moreover, side effects of efficacious drug therapies may also impair productivity. Finally, the early age of onset of schizophrenia magnifies these negative effects. Many patients have not had time to accumulate advanced educational degrees or substantial job experience before the occurrence of their first acute episode.

In view of these challenges, it is not surprising that many studies have reported very low rates of employment for persons with schizophrenia. Steinwachs, Lehman et al. (1997) interviewed a sample of 440 subjects recruited from community-based treatment providers in Georgia and Ohio, and reported that only 14.2% were currently working at a paid job. In this same sample, 19.8% reported receiving any wages or salary from a job in the past month. In a survey of 1,041 consumer members of the National Alliance for the Mentally Ill, of whom 58.7% reported a current diagnosis of schizophrenia, 37.7% reported receiving wages or salary from a job in the past month (Skinner et al., 1992). Slade and Salkever (2001) analyzed interview data for 1,643 adults with a schizophrenia diagnosis recruited from systems of care in six geographic areas across the U.S. and reported that 21.8%

had any kind of employment in the four weeks preceding the interview. Using National Health Interview Survey (NHIS) data from 1989 and from 1994/1995, Mechanic et al. (2002) report employment rates for persons with schizophrenia of 22% in both surveys.

Available evidence on job or employment retention for persons with schizophrenia is fragmentary but also indicates large differentials vis-à-vis the general population. Salkever et al. (2003) studied 159 persons with schizophrenia who were employed at a baseline time period in community jobs without formal job supports (such as a job coach or special supervisor). They observed that more than 30% of these individuals were no longer employed 6 months later. We computed the corresponding figure for the general population ages 18 to 62 using data from the 1996 Survey of Income and Program Participation (SIPP), and found that less than 5% were no longer employed after 6 months. A separate analysis of SIPP data for 1993 and 1996 showed that among employed persons with reported work limitations those with limitations due to mental disorders were significantly more likely at follow-up points 15 months, 23 months and 27 months later to report they had lost their employment (Salkever, 2003).

3. STUDY POPULATION AND DATA

In this study we use data from the U.S. Schizophrenia Care and Assessment Program (SCAP) to explore gender differences in employment. The SCAP is an ongoing longitudinal study of 2,327 adults with schizophrenia recruited from systems of behavioral health care providers in six areas of the U.S. These systems include community mental health centers, public psychiatric hospitals, Veterans Administration mental health providers, and university outpatient clinics. Recruitment began in July 1997 and was completed in January 2001. The six geographic areas were San Diego, Baltimore, New Haven, and selected counties in central Florida, Colorado (including Denver), and North Carolina (including Durham). All participants were over age 18 and had a schizophrenia, schizophreniform, or schizoaffective diagnosis prior to recruitment. Data were obtained from interviews, clinical assessments, and medical records reviews. (Recent studies providing additional information on the SCAP are Mark et al., 2002; Slade & Salkever, 2001; Swanson et al., in press.)

Study subjects were interviewed at baseline and were re-interviewed at 6-month intervals over a three-year follow-up period. The analyses reported in this paper are based on data available as of July 31, 2002 for 2,169 persons age 62 or less at baseline. Three-year follow-up data were available for 755 subjects, two-year follow-up data for 1,358 subjects, and one-year follow-up data for 1,966 subjects.

Employment information was obtained at each interview from subject self-report. Subjects were asked whether they had worked at a job for pay in the past four weeks. Those responding affirmatively were then asked whether their job was “in a sheltered workshop”. In the analyses reported in this paper, work in a sheltered workshop is excluded from our definition of employment. Subjects who worked for pay were also asked the number of days they worked for pay in the four weeks preceding the interview, and the number of hours worked on the days when they did work. We used this information to classify employed subjects into categories by hours of work in the past four weeks (<40, 40 to 79, 80 to 119, and 120+). We interpret greater hours worked as an indication of greater work attachment.

Table 1 presents the characteristics of study subjects by gender. Comparison of sociodemographic variables shows similar racial, ethnic, and educational characteristics for males and females. Their distribution across living arrangements is also fairly similar, although women are slightly more likely to be living with

Table 1. Baseline Characteristics of the SCAP Study Population.

Variable Names and Definitions	Males		Females	
	Mean or %	*N*[a]	Mean or %	*N*[a]
AGE = Age in Years (x10^{-1})	3.959	1345	4.247	824
YRSEDUC = Education in Years	11.69	1331	11.84	815
AGEONS = Age at Onset of Mental Health Problems-18 yrs.[b]	1.46	1254	2.68	757
AFR_AMER = African-American[c]	34.35	1345	38.83	824
HISPAN = Hispanic[c]	11.15	1345	8.37	824
EVERMARR = Ever Married[c]	26.91	1334	55.27	816
ALONE = Living Alone****	22.34	1343	22.90	821
SUPHOUS = Living in Supported Housing[c]	24.57	1343	19.24	821
INST = Living in Institution[c]	9.46	1343	8.53	821
Living in Shelter/Homeless (included in ALONE)	1.27	1343	0.73	821
Living with Spouse, Friends or Family (Default)	42.37	1343	48.60	821
PANSSPOS = Positive Symptom Score	16.50	1345	15.59	824
PANSSNEG = Negative Symptom Score	18.85	1345	17.20	824
MADRS = Montgomery-Asberg Depression Score	13.57	1345	14.71	824
SASCA = Simpson-Angus Scale	3.87	1345	3.81	824
GAF = Global Assessment of Functioning	41.21	1337	43.47	813
LYHOSP = Any Psych. Hospitalization in Past Year[c]	37.69	1332	37.39	813

[a] *N* = number of subjects with valid data for each item.
[b] AGEONS is set to zero if onset is before age 18.
[c] Denotes 0–1 binary variable.

friends or family while men are slightly more likely to be in a supported housing placement with access to a mental health professional. Female subjects are slightly older, were slightly older at the first onset of mental health problems,[3] and are more than twice as likely to have ever been married.

Comparison of clinical characteristics shows no evidence of large gender differences. Men have slightly higher average positive and negative symptom scores and slightly lower depression scores, but Parkinsonian-like side effects of medication (measured by the Simpson-Angus Scale) showed almost no gender difference. The Global Assessment of Functioning, which is an overall rating assigned by the clinical assessor that includes social functioning as well as symptom levels and side effects (higher scores indicate better functioning), was slightly higher for women.

4. EMPLOYMENT AND HOURS OF WORK

Table 2 presents information on employment in the four weeks prior to the baseline interview date for our SCAP study subjects along with comparative data for the

Table 2. Percent Employed in Prior Month and Personal Characteristics: Comparison of Persons with Schizophrenia to the General Population.

	SCAP		1996 SIPP	
	% Emp.[a]	*N*	% Emp[b]	*N*
All males	18.82	1344	84.19	14171
<HS Educ	16.94	791	64.77	1759
HS Educ	17.36	144	83.17	4498
>HS Educ	23.13	402	88.76	7914
Age 18–34	21.64	379	82.33	5157
Age 35–49	18.13	761	89.19	5928
Age 50+	16.18	204	77.72	3086
All females	12.06	821	70.54	16254
<HS Educ	8.94	470	40.82	2050
HS Educ	11.63	43	67.34	5295
>HS Educ	16.78	304	78.37	8909
Age 18–34	13.81	181	69.70	5888
Age 35–49	12.79	430	76.11	6810
Age 50+	9.05	210	61.73	3556

[a] Employment is defined to exclude working in sheltered workshops.
[b] 0SIPP percentages are based on weighted data.

general U.S. population from the 1996 Survey of Income and Program Participation (SIPP), conducted by the U.S. Census Bureau.[4] Individuals are classified by their reported hours of work in the past four weeks. In the SIPP, we computed hours worked in the past month by multiplying the respondent's reported usual hours worked per week by their reported weeks worked in the last month. Responses indicating 5 weeks worked in the last month were recoded to 4.

In our SCAP study data, the overall employment rates are very low, both for males (18.82%) and for females (12.06%). The male-female differential, though positive, is not statistically significant ($P = 0.224$ based on a chi-squared test). The comparative data for persons age 18–62 in the 1996 SIPP pertains to employment in the first month of that survey. Employment rates are much higher for both genders and the male-female differential is almost 14% (84.19 vs. 70.54), but that differential is small relative to the overall employment rate for the general population.

Comparison of age-employment and education-employment gradients in the SCAP with those in the SIPP show several interesting differences. In the SCAP, employment is positively correlated with education and negatively correlated with age. In the SIPP, there is also a positive education gradient and a negative age gradient, but the decline in employment rates by age begins later (i.e. after ages 35–49) than in the SCAP. Also, the education gradient in the SIPP is much larger for females than it is for males.

While the employment rate differences between persons with schizophrenia and the general public are very large, there are also large differences in hours of work (Table 3).[5] Employed persons with schizophrenia are much less likely to be working full time and much more likely to be working at casual, temporary,

Table 3. Percent Distribution of Employed Persons by Hours Worked in a Four-Week Period of First Survey Month.

	SCAP (%)	1996 SIPP (%)
Males		
<40 Hrs.	46.23	0.98
40–79 Hrs.	22.11	1.68
80–119 Hrs.	15.08	3.48
120+ Hrs.	16.58	93.86
Females		
<40 Hrs.	46.67	2.76
40–79 Hrs.	18.89	5.54
80–119 Hrs.	14.44	10.29
120+ Hrs.	20.00	81.41

or other very part time jobs. For both males and females, the SCAP data show that almost half of all employed respondents averaged less than 10 hours of work per week, and only about one-third report working more than 30 hours per week on average. This contrasts sharply with the data for the general public in the SIPP, which show that 93.86% of working males and 81.41% of working females averaged more than 30 hours per week. The other interesting comparison pertains to gender differences. Among the general public, employed males are more likely to work longer hours; among the SCAP respondents, there is no clear pattern of gender differences.

5. EXPLORING GENDER DIFFERENCES IN RATES OF EMPLOYMENT

The large *relative* difference by gender in employment rates for persons with schizophrenia raises the question of the underlying reasons for this difference. We begin to examine this question by modeling, via maximum-likelihood probit regression analysis, the probability of being employed at baseline among SCAP respondents. In this analysis, we view the individual subject's employment status as reflecting the combined effects of human capital that determines their market productivity, market opportunities, household characteristics that influence the productivity of non-market time, and preferences for market work vs. other activities that may reflect unobservable individual taste factors as well as group-level socio-cultural differences.

As Altonji and Blank (1999) have noted, the effects of these factors may also be expected to vary by gender. Given the relatively large number of observations in our employment status regressions, and the gender differences in employment status observed in Tables 2 and 3, we allow for these gender differences by estimating separate regression models by gender.

Within the general approach just described, the rationales for including specific characteristics in our models are fairly straightforward. Baseline characteristics relating to mental health status and symptoms (PANSSPOS, PANSSNEG, MADRS, SASCA) are presumed to influence employment status primarily through their impact on market productivity of the individual. Other variables related to the need at baseline for support services (SUPHOUS, INST) and to a recent history of hospitalization (LYHOSP) are also presumed to be positively related to severity of disability, and thus negatively related to market productivity. Educational attainment (YRSEDUC) is included as measure of general human capital accumulation. Age at onset (AGEONS) is included to control for potential human capital appreciation through work experience (since we also control for

AGE and YRSDEDUC).[6] An additional binary variable for age of onset <18 (AGEONLT18) was included on the assumption that work experience prior to age 18 was a negligible source of human capital. A binary indicator for missing age of onset was also included to avoid dropping the substantial number of study subjects for whom AGEONS was not ascertained. AGE is included to capture depreciation in human capital. Based on the descriptive data in Table 2 showing a monotonic relationship between age and employment for SCAP subjects, the effect of AGE is presumed to be linear. EVERMARR and ALONE are included as indicators (albeit crude ones) of non-market work opportunities (e.g. childcare, housekeeping), which should be negatively related to human capital accumulation and to employment. In the context of gender comparisons, Altonji and Blank refer to these factors as differences in "comparative advantage." These latter two variables may also proxy for other family income sources besides the individual's own earnings. The indicators of minority race-ethnicity (AFR_AMER and HISPAN) are intended to control for socio-cultural differences in labor-supply behavior; however, they may also capture differences in employment opportunities due to labor-market discrimination and socio-cultural differences in the response to serious mental disorders. Finally, we assume that labor-market opportunities vary across but not within study locations. Thus, inclusion of separate intercepts for each of our six study locations controls for the effects of market opportunities.

5.1. Results

The results of the analysis are reported in Table 4. Columns (1) through (6) present results when the explanatory variables included all the respondent characteristics included in Table 1 (except for the GAF score which itself depends on employment status). Columns (7) through (12) replicate the analyses deleting the two regressors that had p-values >0.3 for *both* males and females. Values reported as $\mathrm{d}F/\mathrm{d}x$ represent the marginal predicted change in proportion employed resulting from a one-unit change in the value of a continuous variable or a change in the value of an indicator variable from 0 to 1. Several significant findings are consistent across genders. A higher negative symptom score (PANSSNEG) is strongly associated with reduced probability of employment. This result has been previously noted in several studies (Pearlson et al., 1984; Slade & Salkever, 2001; Tsang et al., 2000). Living in housing with supportive mental health services available (SUPHOUS = 1) is associated with a reduction in employment probability ($\mathrm{d}F/\mathrm{d}x$) of about $+0.05$ (see columns 8 and 11); this is consistent with the expectation of more severe psychopathology and greater need for support services among persons in these

Table 4. Probit Regression on Baseline Employment Dummy for Respondents Age 18–62 at Baseline.

	Females			Males			Females			Males		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	Coef.	d*F*/d*x*	*P*>|*z*|	Coef.	d*F*/d*x*	*P*>|*z*|	Coef.	d*F*/d*x*	*P*>|*z*|	Coef.	d*F*/d*x*	*P*>|*z*|
EVERMARR*	0.214	0.035	0.132	−0.010	−0.003	0.920	0.194	0.032	0.168	−0.002	−0.001	0.981
ALONE*	−0.032	−0.005	0.834	−0.208	−0.051	0.059	−0.051	−0.008	0.738	−0.190	−0.047	0.081
SUPHOUS*	−0.264	−0.039	0.174	−0.206	−0.051	0.054	−0.316	−0.046	0.093	−0.222	−0.054	0.036
INST*	−0.207	−0.030	0.421	−0.403	−0.088	0.019	−0.260	−0.037	0.301	−0.427	−0.093	0.011
LYHOSP*	−0.067	−0.011	0.633	−0.088	−0.023	0.346						
PANSSPOS	−0.012	−0.002	0.372	0.006	0.002	0.460						
PANSSNEG	−0.040	−0.007	0.001	−0.025	−0.006	0.001	−0.044	−0.007	0.000	−0.023	−0.006	0.001
MADRS	−0.004	−0.001	0.527	−0.006	−0.002	0.214	−0.007	−0.001	0.335	−0.006	−0.002	0.181
SASCA	−0.041	−0.007	0.035	−0.015	−0.004	0.187	−0.044	−0.007	0.025	−0.015	−0.004	0.201
HISPAN*	−0.557	−0.066	0.073	−0.055	−0.014	0.689	−0.474	−0.060	0.100	−0.059	−0.015	0.666
AFR_AMER*	0.094	0.016	0.546	−0.166	−0.042	0.101	0.078	0.013	0.615	−0.167	−0.042	0.097
AGEBL/10	−0.161	−0.026	0.032	−0.117	−0.030	0.021	−0.162	−0.027	0.028	−0.105	−0.027	0.033
YRSEDUC	0.067	0.011	0.011	0.025	0.006	0.179	0.068	0.011	0.010	0.025	0.006	0.176
AGEONS-18	0.005	0.001	0.644	0.016	0.004	0.081	0.005	0.001	0.681	0.015	0.004	0.098
AGEONSM*	0.138	0.025	0.603	−0.058	−0.015	0.780	0.100	0.018	0.703	−0.062	−0.016	0.761
AGEONLT18*	−0.130	−0.021	0.434	0.146	0.039	0.164	−0.163	−0.026	0.321	0.158	0.042	0.131

Notes: All regressions include separate intercepts for each of the six study areas. Variable are defined in Table 1 (above) except for two additional 0–1 binary variables: AGEONSM = 1 age of onset is missing, and AGEONLT18 = 1 if age at onset <18. Reference categories are never married, living with friends and/or family, and non-Hispanic non-African-American. d*F*/d*x* = change in employment probability for a one unit change in continuous variables or a 0–1 change in a binary variable.

Table 5. Ordered Probit Estimates of Employment at Baseline (0 = Unemployed, 1 = Less Than 80 hrs, 2 = Greater Than 79 hrs) for Respondents Age 18–62 at Baseline.

	Females		Males		Females				Males			
	Coef.	$P>\|z\|$	Coef.	$P>\|z\|$	Coef.	Marg. Effects		$P>\|z\|$	Coef.	Marg. Effects		$P>\|z\|$
						EMP = 1	EMP = 2			EMP = 1	EMP = 2	
EVRMARR[a]	0.116	0.350	−0.072	0.446	0.111	0.016	0.009	0.369	−0.072	−0.013	−0.010	0.446
ALONE[a]	−0.059	0.668	−0.181	0.068	−0.069	−0.010	−0.005	0.613	−0.182	−0.032	−0.024	0.068
SUPHOUS[a]	−0.169	0.305	−0.175	0.071	−0.182	−0.026	−0.014	0.268	−0.175	−0.030	−0.023	0.070
INST[a]	−0.198	0.372	−0.364	0.016	−0.224	−0.032	−0.018	0.311	−0.364	−0.063	−0.049	0.016
LYHOSP[a]	−0.172	0.176	−0.149	0.078	−0.183	−0.026	−0.014	0.147	−0.151	−0.026	−0.020	0.073
PANSSPOS	−0.011	0.350	−0.001	0.876								
PANSSNEG	−0.039	0.000	−0.015	0.023	−0.042	−0.006	−0.003	0.000	−0.015	−0.002	−0.002	0.017
MADRS	−0.010	0.108	−0.007	0.095	−0.011	−0.002	−0.001	0.062	−0.007	−0.001	−0.001	0.079
SASCA	−0.039	0.022	−0.015	0.131	−0.039	−0.006	−0.003	0.020	−0.016	−0.003	−0.002	0.130
HISPAN[a]	−0.382	0.127	−0.074	0.556	−0.381	−0.055	−0.030	0.129	−0.075	−0.011	−0.009	0.554
AFR_AMER[a]	0.004	0.976	−0.176	0.055	0.001	0.000	0.000	0.993	−0.176	−0.030	−0.023	0.055
AGEBL/10	−0.128	0.051	−0.152	0.001	−0.130	−0.019	−0.011	0.047	−0.152	−0.025	−0.019	0.001
YRSEDUC	0.079	0.001	0.024	0.147	0.078	0.011	0.006	0.001	0.024	0.005	0.004	0.149
AGEONS-18	0.011	0.281	0.034	0.000	0.011	0.002	0.001	0.255	0.034	0.006	0.005	0.000
AGEONSM[a]	−0.038	0.880	−0.090	0.633	−0.057	−0.008	−0.005	0.821	−0.092	−0.015	−0.012	0.624
AGEONLT18[a]	0.045	0.757	0.189	0.049	0.040	0.005	0.003	0.783	0.189	0.035	0.027	0.049

Notes: All regressions include separate intercepts for each of the six study areas. Variable definitions are given in Tables 1 and 4 above.
[a] Denotes a 0–1 Binary variable.

living situations. An additional year of age at baseline has a significant negative effect on employment probability of -0.0027 ($= 0.1$ times dF/dx for AGE).

Other results differ across genders. The negative coefficient for living alone is larger and more significant for males; the same is true for living in an institution. The positive effect of education is larger and more significant for females, while Parkinsonian symptoms (SASCA) have somewhat larger negative effects for females. Older age of onset is positively related to employment for men but not for women. This is consistent with a work-experience effect if men are more likely to accumulate work experience in the years between age 18 and the onset of the disorder. Negative employment differentials are also observed for Hispanic women and African-American men.

To examine possible gender differences in employment patterns when differentials in hours of work are also accounted for, we re-estimated the regressions in Table 4 using ordered probit with three outcomes: non-employed, worked less than 80 hours, and worked 80 hours or more. These results are reported in Table 5.

Some of the results indicate findings different from those for employment vs. non-employment. Coefficients for all three living situation dummies become negative and strongly significant for males, but they are not significant for females. Results for variables relating to symptoms and disease course are also more pronounced. Negative symptoms again have very strong negative effects on employment status, with their marginal effects being somewhat larger for females than for males. The depression rating (MADRS) has a moderately significant negative impact on employment status. As in the simple probit, Parkinsonian symptoms (SASCA) have a much stronger negative effect for women than for men.

Results for the socio-demographic variables are quite similar to the simple probit findings. Age has a strong negative effect while age at onset is significantly positive for males but not for females. (The strongly positive coefficient for males for the binary indicator of onset before age 18 is, however, counterintuitive.) Education has a very strong positive effect on employment status for women but not for men, and African-American men are less likely to be employed.

6. EMPLOYMENT STABILITY AND INSTABILITY

Since we have up to 7 different observations (called "assessments") on employment status for each respondent at 6-month intervals, it is possible to examine gender differences in employment stability over a period as long as three years. Basic descriptive data are provided in Table 6. We group respondents by the number of assessments at which data on employment status were reported; this

Table 6. Comparison by Gender of Employment Stability.

Panel A

No. of Assessments w. Reported Employment Status	Males		Females		Prob. (χ^2)
	N	% w. No Employment	*N*	% w. No Employment	
2	135	68.89	72	77.78	0.175
3	173	61.85	92	71.74	0.107
4	175	55.43	101	75.25	0.001
5	200	64.50	121	67.77	0.55
6	278	54.32	168	66.07	0.015
7	274	49.64	206	69.90	$<$0.001

Panel B

	Males w. Any Employment		Females w. Any Employment		Prob. (χ^2)
	N	% Always Employed	*N*	% Always Employed	
2	42	47.62	16	31.25	0.261
3	66	22.73	26	15.38	0.433
4	78	15.38	25	16	0.941
5	71	12.68	39	17.95	0.453
6	127	15.75	57	8.77	0.202
7	138	7.97	62	9.68	0.689

number ranges from 2 to 7. (Respondents with data from only one assessment were uninformative about employment stability and were therefore excluded.) The table reports two tests for gender differences. First (in Panel A), we test for gender differences in the percent of subjects who did not report being employed at *any* assessment. We perform this test for each subgroup of subjects defined by the number of available assessments (since random factors would cause the probability of no employment to fall as the number of assessments increases). Second (in Panel B), we perform analogous tests of gender differences on the percentage of subjects who reported being employed at *every* assessment. Note that in this second case we restrict our comparisons to subjects who reported employment in at least one assessment.

Panel A of the table indicates that a substantial fraction of all respondents report no employment over the entire period of observation. This percentage is always lower for males and in 3 of 6 instances this difference is highly significant. We also note that the percentage falls with the number of assessments for males much more than for females, suggesting that the lack of employment is a less permanent status for males.

Panel B shows that among persons with any work, the percent working at every assessment was not significantly different between males and females. For both genders, most persons who work report an intermittent work history while a relatively small fraction report persistent attachment to employment. We also compared hours of work in the baseline job for persons with persistent attachment and found that for both genders about 29% of these persons had full-time jobs (working 30+ hours per week). Thus the majority of persons reporting employment at every assessment point (excluding missing values) were working in part-time positions.

7. EMPLOYMENT STATUS TRANSITIONS

An alternative approach to studying possible gender differences in employment stability is to examine the transitions processes from employed to not employed and vice versa. We noted in the introduction that short employment tenure appeared to be a serious problem for persons with schizophrenia who were working relative to other employed persons. Thus, in comparing SIPP data to our SCAP respondents we find that while 95.5% of working persons in the SIPP were employed 6 months later, this was true for only 59.2% of working persons in the SCAP survey. In view of the greater prevalence of casual and part-time work among SCAP respondents, this disparity is not surprising. Table 7 documents, however, that this disparity persists even when we control for the extent of work. Examining the data by gender and hours of work, we see that within every grouping, the percentage still retaining employment is lower for SCAP respondents by 12% or more. What is also interesting is that this differential is considerably greater for females with the exception of those working more than 120 hours in the four-week baseline. For males, the size of the differential only ranges from 15.48% (for those working 120+ hours) to 32.64% (for those working 40–79 hours). For females, the differential exceeds 40% for three of the four groups.

A final point of interest from Table 7 is comparison of gender differences within each of the surveys. The SIPP data indicate that for the general employed population there is virtually no difference in employment retention rates by gender once we account for hours of work. In contrast, for SCAP respondents, retention rates within every category of hours of work are lower for females than for males except for persons working 120+ hours. (The differentials are statistically significant for persons working $<$40 hours and those working 80–119 hours; corresponding P-values are 0.046 and 0.013.) The implication is that schizophrenia does impact

Table 7. Employment Retention at 6 Months for Persons Age 18–62 Working at Baseline by Hours Worked in Four Weeks Pre-Baseline.

	1996 SIPP[a]		SCAP	
	N	(%)	*N*	(%)
All Males	11955	96.70	199	63.32
<40 Hrs	117	79.26	92	57.61
40–79 Hrs.	201	87.19	44	54.55
80–119 Hrs.	416	90.75	30	73.33
120+ Hrs.	11221	97.30	33	81.82
All Females	11335	94.16	90	50.00
< 40 Hrs.	313	79.63	42	38.10
40–79 Hrs.	628	87.98	17	47.06
80–119 Hrs.	1166	91.04	13	46.15
120+ Hrs.	9228	95.51	18	83.33

[a] SIPP percentages are based on weighted data.

employment retention rates differentially by gender and that this is particularly true for persons employed less than full time (i.e. less than 30 hours per week).

7.1. Probit Estimates of Employment-Retention

To test the significance of the gender difference within the SCAP data, we estimated probit regressions on a binary indicator of employment at 6 months post-baseline. Explanatory variables were a gender dummy (MALE), and dummies for working 40–79 hours (W4079), 80–119 hours (W80119) or 120+ hours (W120) at baseline. Results are reported as Model 1 in Table 8 below. We find a highly significant

Table 8. Probit Regressions on Employment Retention at 6 Months ($N = 284$).

	Model 1			Model 2		
	Coefficient	d*F*/d*x*	Prob.	Coefficient	d*F*/d*x*	Prob.
MALE	0.381	0.149	0.020	0.435	0.169	0.013
W4079	0.011	0.004	0.953			
W80119	0.355	0.131	0.117	0.417	0.151	0.078
W120	0.914	0.306	<0.000	0.726	0.250	0.002
YRSEDUC				0.111	0.042	0.001
LYHOSP				−0.590	−0.229	0.001
FAM_KID				−0.747	−0.291	0.006
Constant	−0.225		0.153	−1.313		0.005

gender difference that implies the probability of employment retention for males is 0.15 greater than for females. This probability also increases with hours of work at the baseline job, and is significantly higher for persons working full time at baseline.

Inclusion of additional explanatory variables allowed us to test for other factors that may influence the job retention probability. These included the baseline clinical ratings of GAF, PANSSPOS, PANSSNEG, MADRS, and SASCA, and three other variables used in our employment status regressions in Tables 4 and 5 (YRSEDUC, AGEBL/10, and LYHOSP). We also included a dummy for respondents who are living with friends, family or spouse, *and* who reported that housekeeping or childcare is their main activity (FAM_KID). This dummy is used as an indicator of domestic demands in the home such as childcare. (A more direct measure is not possible since the survey did not specifically ask about the number of children living in the respondent's home.)

Only three of these additional variables were significant. Results for these variables are reported as Model 2 in Table 8. (Also note that W4079 was dropped from Model 2 since its estimated coefficient was very close to zero and clearly insignificant.) Years of education (YRSEDUC) was strongly predictive of employment retention while persons with a psychiatric hospitalization in the pre-baseline year (LYHOSP) were much less likely to retain employment. This may arise from an association between LYHOSP and recurrence of acute episodes in the post-baseline period. Domestic responsibilities were also strongly and negatively related to employment retention. Also note that inclusion of these factors did not diminish the size or significance of the gender differential (as indicated by the coefficient of MALE).[7]

7.2. Hazard Regression Analysis of Employment Duration

We also extended our analysis to the *length* of employment spells beyond the 6-month follow-up. More specifically, we estimated a hazard regression on data for these same 284 persons employed at baseline including data for consecutive follow-up assessments for which employment status was not missing. The hazard function describes the probability at each assessment (from assessment 2 through 7) of ceasing to be employed. Explanatory variables in the hazard regression included the same variables used in the probit regressions. Observations on respondents for whom employment status is missing at any assessment are truncated. Failure is defined as reporting no employment at a particular assessment. Assessments subsequent to a reported failure are dropped from the analysis.

Rather than specify a parametric form for the hazard function we adopt the Cox semi-parametric approach. An advantage of this method, relative to fully

Table 9. Proportional Hazard Regressions on Length of Employment ($N = 284$).

	Model 1		Model 2	
	Hazard Ratio	Prob.	Hazard Ratio	Prob.
MALE	0.716	0.079	0.724	0.099
W4079	0.992	0.972		
W80119	0.722	0.255	0.697	0.227
W120	0.359	0.004	0.448	0.024
YRSEDUC			0.918	0.026
LYHOSP			1.573	0.021
FAM_KID			1.679	0.051

parametric approaches, is that it avoids possible inconsistency of parameter estimates (due to incorrect specification of the form of duration dependence) with relatively little efficiency loss (Bernell, 2000). All explanatory variables are assumed to impact the hazard function proportionally across all of the six time periods (Assessments 2 through 7) in the analysis. Hazard Ratios and *p*-values are reported in Table 9. Hazard ratios represent a proportional shift in the baseline hazard for employment loss up or down depending on whether the ratio is greater than 1 or less than 1, respectively.

Once again we find a gender difference in the risk of becoming not employed, albeit with a reduced significance level.[8] Similarly, higher levels of hours of work at baseline are again predictive of a lower hazard for non-employment (greater employment retention) but with slightly reduced significance levels. Results for duration, prior hospitalization, and domestic responsibilities also parallel those in the probit regression, but in this case their inclusion tends to weaken the evidence of a pure gender difference. Results when some or all of the clinical variables were included (available from the authors) did not yield any significant results but further diluted the gender coefficient, raising it to about 0.8, and raising its *p*-value to about 0.15–0.2. Thus, we conclude that evidence of a pure gender difference in length of employment spells is somewhat weaker than evidence of a gender difference in short-term employment retention. Finally, as in the probit case (see note 5 above), we tested for gender interactions with other explanatory variables, and could not reject the null hypothesis of no interactions.

7.3. Hazards for Transitions from Non-Employment

To complete our analysis, we also estimated hazard models for transition to employment of respondents who were not employed at baseline. The variables used

Table 10. Proportional Hazard Regressions on Length of Non-Employment ($N = 6{,}358$).

	Model 1		Model 2		Model 3	
	Hazard Ratio	Prob.	Hazard Ratio	Prob.	Hazard Ratio	Prob.
MALE	1.449	0.002	1.443	0.002	10.818	<0.000
AGEBL	0.971	<0.000	0.971	<0.000	0.972	<0.000
YRSEDUC	1.043	0.077	1.043	0.075	1.167	<0.000
LYHOSP	1.006	0.962				
FAM_KID	0.659	0.035	0.655	0.032	0.66	0.035
GAF	1.013	0.006	1.014	0.003	1.013	0.005
PANSSPOS	0.973	0.118	0.971	0.063	0.976	0.124
PANSSNEG	1.005	0.723				
MADRS	0.999	0.949				
SASCA	0.959	0.147			0.958	0.138
MALE-ED					0.847	0.001

in this analysis are the same as those used in our probit and survival models of employment retention, with the exception of the three work hours dummies (which would be zero for all respondents not employed at baseline).

Results are reported in Table 10. In this case, the gender difference is highly significant, with men having 44% greater odds of moving from non-employment to employment in each assessment period. Age is strongly and negatively related to the odds of transitioning to employment while the effects for clinically assessed functional status (GAF) and education are positive. Respondents with greater involvement in domestic responsibilities (FAM_KID = 1), and those with higher levels of positive symptoms, are significantly less likely to move into employment. Finally, tests for pooling rejected the null hypothesis of equal coefficients for men and women, but only indicated a substantial difference for the education coefficient. Thus we re-estimated the model with a gender-education interaction. The results, shown in the last two columns of Table 10, indicate a strong positive education effect for women and essentially no effect for men.[9]

8. DISCUSSION

Our analysis of a new and substantial database of persons with schizophrenia confirms previous reports of the disorder's dramatic negative effects on employment rates, especially on rates of full-time employment (i.e. working 30+ hours per week). Examination of employment data by gender reveals several interesting departures from patterns observed in the general population. While overall employment levels are clearly lower for women with schizophrenia, the distribution of hours of work among those with jobs does not vary by gender, even though in the general population women work fewer hours on average. The factors affecting employment probabilities and hours of work for persons with schizophrenia also vary somewhat by gender. Education has a stronger positive effect for women while negative symptoms and Parkinsonian side effects have a stronger negative effect for women. We speculate that these results may be related to gender differences in the types of employment opportunities; unfortunately our data set does not provide information on job titles or characteristics that would allow us to test this hypothesis.

The results with respect to living situation and marital status are somewhat surprising. Living situation is a much more important factor for men, while marital status does not have strong effects for either gender. Although prior research on gender differences for the general population indicates that being married and being in a family living situation should impact negatively on employment for women (Altonji & Blank, 1999), these results are not observed in our data.

Data on employment stability over time document dramatically lower rates of employment retention among persons with schizophrenia relative to the general population. They also show greater risk of transition to non-employment for employed women, relative to men, controlling for hours of work; this pattern is not observed in the general population. Analysis of the other predictors of this risk (besides hours of work) identified education as a significant protective factor, and recent hospitalization and responsibilities for housekeeping and child care as contributors to this risk. Only the latter, however, is strongly correlated with gender; controlling for this factor does not eliminate or substantially diminish the significant gender difference in risk. Furthermore, we observe a significant gender difference in the probability of transitioning *to* employment among those who are not employed. The size of this gender difference appears to be inversely related to the level of education.

Finally, we compared employment patterns over time by gender for schizophrenia patients and observed: (1) a greater tendency for females to be persistently non-employed; and (2) similar persistence of employment across gender for those reporting any employment. We also found that among those who were persistently employed, the percentage working full time is only about 29% for both men and women.

First, women with schizophrenia are more likely to be persistently non-employed, other factors held constant. Second, women who are employed less than 30 hours per week are at a greater risk of becoming non-employed, other factors held constant.

A variety of factors not captured in our analyses could explain these differences. Persistent non-employment could be a consequence of factors such as lack of job experience and vocational skills. It might be also due in part to factors such as lack of motivation to work. In any event, it seems reasonable to consider whether evidence-based vocational services programs, such as supported employment, effectively reach women with schizophrenia.

Higher risk of employment loss could be related to gender differences in types of jobs. For example, if women are more likely to be in jobs that involve interpersonal skills, negative symptoms (which interfere with these skills) may be a greater impediment to job retention. There may also be differences in stability of the jobs themselves. Women could be more likely to fill casual or temporary part-time jobs whereas men might find more stable part-time situations. To sort out the truth in these speculations, further collection and analysis of detailed data about the nature of jobs and the reasons for termination or quitting will obviously be required.

9. CONCLUDING REMARKS

Evidence from the mental health literature suggests that while the disabling consequences of schizophrenia are severe for both men and women, gender differences in age of onset, response to treatment, and long-term course of the illness may produce more positive long-term outcomes for women. In spite of this evidence, we observe higher rates of long-term non-employment for women and higher risks of employment loss for women who are employed. These findings are also somewhat surprising in view of the fact that gender differences in domestic responsibilities thought to account for gender differences in employment in the general population are less pronounced for persons with schizophrenia, and in view of the fact that no corresponding gender differences in risk of employment loss are observed in the general population.

Thus, we need to look beyond clinical factors and factors that produce gender differences for the general population in order to understand the reasons for our results. We have suggested several avenues for further inquiry that may be promising in this regard. One is to obtain and examine more detailed evidence on the nature of the jobs that women with schizophrenia actually hold and the factors that result in termination or quits from these jobs. A second is to examine the extent to which effective vocational interventions, such as supported employment programs, reach women as well as men within the population of persons with schizophrenia. We are optimistic that these further investigations will lead to improvements in our strategies for supporting greater workforce involvement by women with schizophrenia.

NOTES

1. See Goldstein (1988) and the references cited there.

2. The descriptive material in this paragraph and the next is based on the Surgeon General's Report on Mental Health (USDHHS, 1999).

3. Note that onset of mental health problems does not specifically mean first acute schizophrenic episode. Recent work by Agerbo et al. (forthcoming) indicates that psychopathology and social dysfunction may in fact precede the first acute episode by a number of years.

4. The 1996 SIPP cohort was used for this comparison since it comes closest in calendar time to the start date of the SCAP project (which was July 1997). The only other SIPP cohort with any time overlap with the SCAP is the 2000 SIPP.

5. In the SIPP, we computed hours worked in Table 3 by multiplying the respondent's reported usual hours worked per week by their reported weeks worked in the last month. (Responses indicating 5 weeks worked in the last month were recoded to 4.)

6. Charles (2003) has argued that age of onset may also be relevant in determining post-onset human capital investments, since such investment will yield returns over a longer period of time for those with earlier onset ages. While he developed this argument for analyses of a broad range of disabilities with widely varying onset ages, it is relevant to note that in our data the age of onset is relatively young and not as variable as with many other types of disabilities.

7. Because our sample for these regressions was only 284 respondents, estimation of separate models for males and females produced generally insignificant results. Likelihood ratio tests of these separate regressions vs. pooled models with separate intercepts by gender could not reject the null hypothesis that other coefficients did not vary by gender.

8. Note that hazard ratios in Table 9 refer to the hazard of becoming not employed. Thus a positive (negative) coefficient in Table 8 is comparable to a hazard ratio less than (greater than) 1.0 in Table 9.

9. The effect of an additional year of education for men is the product of the coefficients for YRSEDUC and MALE-ED; this is 1.167×0.847, or 0.988, which is essentially 1.0, implying no change in the hazard ratio. Also note that while the reported coefficient for MALE is much larger when the interaction variable MALE-ED is included, the net effect of gender is not substantially altered. For example, based on the estimated coefficients for both MALE and MALE-ED, the odds ratio for a man with 12 years of education relative to a woman with 12 years of education is 1.434.

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THE ROLE OF GENDER IN A COMPANY-WIDE EFFORT TO EXPAND AND DESTIGMATIZE MENTAL HEALTH TREATMENT

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INTRODUCTION

The role of gender in psychiatric disorders is not well understood, but several broad trends are known: while men and women experience psychiatric symptoms at roughly the same rate, women are more likely to experience depressive symptoms (Kessler et al., 1993) and men are more likely to experience substance use disorders (DHHS, 1999). However, women are more likely to use primary care services for mental health care than are men (Wells et al., 1986). Equally controversial and not well understood has been the differential responses to treatment interventions by gender (Kornstein, 1997). One recent study found that a depression intervention was more cost-effective for women than for men (Pyne et al., in press). Indeed, the study found that the intervention was essentially cost *and* outcome neutral for men, while women were found to have a cost-effectiveness ratio of over $5000 for each QALY saved.

In this study we test whether differential gender responses exist after a company-wide benefit policy change to broaden mental health benefits on mental

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health treatment initiation at a nationally known communications and electronics company. The benefit policy change included three major components: (1) reduced copayments for mental health treatment, (2) a newly created network of select mental health providers, and (3) efforts to destigmatize the treatment of mental disorders. Though the benefit policy change was gender neutral, we test whether there was a differential response by gender.

Each of these benefit changes may have important effects on mental health treatment decisions. First, results from the RAND Health Insurance Experiment (HIE) showed how cost sharing tempered the demand for mental health services (Keeler et al., 1988). Nevertheless, it is not clear *a priori* whether cost sharing would differentially affect the actions of women vs. men. Second, using the same data as in this study, Lindrooth et al. (2003) showed that the selective contracting network was not binding in terms of initiation, though it did appear to encourage the use of non-MD mental health specialists. Because the selective contracting network was not binding, we do not expect to see differences in initiation by gender. Finally, the stated emphasis of the program was destigmatization of mental health treatment. The stigma associated with mental illness has been shown to decrease adherence to depression treatment (Sirey et al., 2001). The important contribution of this study is to test whether there are gender differences in initiation in a program where the principal emphasis is on destigmatizing 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 U.S. based workers. The company had offices and production facilities throughout the US, though 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; as discussed later, the northeast-based employees were dropped from the analysis because of their relatively small number.

The company had long provided a mental health and substance abuse benefit though in the early- to mid-1990s it became apparent that expenses for mental health treatment services were increasing at a striking rate. Beginning in 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 post-periods were reduced. Coinsurance rates on inpatient care were changed from 80 to

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

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

90%, while 80% coinsurance for outpatient services was replaced with a flat $10 copayment. The company also implemented a 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 did retain 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. Moreover, the company started a campaign to eliminate the stigma associated with seeking mental health treatment. They disseminated 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 HMO products 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% in 1996 to 16% in 1999. As discussed later, to avoid 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.

The company had more than twice as many men than women enrolled in their self-insured health insurance plan (70% vs. 30%), which generally reflected the overall composition of the workforce. The gender composition of the company's workforce reflects the company's large engineering and manufacturing component to its business-line. The company did not specifically target gender in its benefit changes, thus the results we observe are likely to be a reflection of the "real world" gender impact of a benefit expansion of this kind.

DATA

Analytical File Construction

We acquired claims data files for the company from their claims data manager, Medstat. We had complete set of claims data (outpatient, inpatient, and pharmaceutical) for employees diagnosed with a mental health or substance abuse disorder at any time during the period 1995–1998 for persons enrolled in the company's self-insured health insurance plan. As mentioned earlier, employees could opt annually to join an HMO instead of the company's health insurance plan. We did not have access to utilization information for employees enrolled in an HMO. Because of the potential for sample selection effects associated with persons more likely to use mental health treatments to increasingly enroll in the company plan after implementation of the benefit change, we focus only on those employees who were continuously enrolled in the company plan over the entire 4-year period. While this approach runs the risk of introducing sample selection bias because the sample is comprised of employees who possibly have longer tenure with the firm than average, we believe this potential source of bias is the lesser of the two evils.

We also had access to enrollment information for the period between 1995 and 1998 for persons enrolled in company's health insurance plan. The enrollment data were updated on a quarterly basis and contained date of birth, family status, gender, employment status, and home zip code. Over the 4-year period, 104,646 employees had at some point been part of the company's self-insured health plan. We restricted our sample, however, to employees that were continuously enrolled in the company's health plan over the 4-year period, leaving us with 33,983 individuals. During the 4-year analysis period, the number of employees enrolled in the company's health insurance plan grew somewhat. In 1995, those same 33,983 individuals represented about 53% of the individuals who were enrolled at any time during the year. By 1998, the percentage of total enrollment represented by our continuously enrolled subsample was 45%.

A separate administrative data set contained information on employee race, but only for individuals who were employed in 1999. Merging with this administrative data set reduced our sample to 32,345 individuals. To allow for greater ease in defining distance to nearest provider, we focus only on individuals who did not change zip code of residence over the 4-year period, yielding a sample of 26,458 individuals. To focus on individuals who were exposed to the company's "stigma-free" initiative, we further limited our sample to individuals that were between 18 and 65 years of age over the 4-year period and were neither COBRA-enrolled nor retirees. These exclusions pared our sample size to 24,802. Finally, we excluded employees who lived in the Eastern region because there were too few to accurately identify the provider network in this area. This led to our final sample size of 24,297.

Measures

By merging mental health claims with the enrollment data we created a continuously enrolled sub-sample with which to examine mental health treatment decisions. Using inpatient and outpatient claims for the individuals, 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. The diagnosis codes represent all recognized behavioral health disorders including, but not limited to, depression, anxiety disorders, personality disorders, and substance abuse disorders. The percentage of people in our sample who had at least one mental health claim in each year was 8.35, 8.90, 9.70, and 9.99%, in 1995, 1996, 1997 and 1998 respectively.

Because we did not have access to an external data source containing all possible mental health providers available to individuals, we use the claims of those who sought treatment to calculate the distance to the nearest provider for both users and non-users of mental health treatment services. This introduces a potential for biased distance measures if, as we expect, treatment initiation and distance are correlated due to the introduction of the selective contracting network. To estimate the distance to the closest potential provider, and to ensure that this distance estimate is as independent as possible from measures of treatment initiation, we used the set of claims data for those individuals *not* included in our analytic sample of 24,297 persons. Thus, the mental health provider data were derived from the 80,359 persons who either were not continuously enrolled in the employer's self-insured health plan or were omitted from the analytic sample for some other reason.

Using the claims data for this different group of employees, we identified all mental health treatment claims (defined as a claim presenting one of the aforementioned diagnoses or CPT codes). We then extracted the set of provider IDs and zip codes associated with these mental health claims. For our purposes, this set of providers represents the "universe" of potential providers available to individuals in the analytic sample. Importantly, this set of providers is not correlated with the utilization of mental health care services among the members of our sample. Distance to any 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). In 1995, prior to the implementation of the selective contracting network, patients could access any provider. Thus, we included all providers observed in the period from 1995 through 1998. For the years 1996–1998, when selective contracting was in place, distance was calculated using only those providers observed during those 3 years.

Our prior work has demonstrated that the selective contracting network restrictions were rarely binding for enrollees (Lindrooth et al., 2003). That is, the network was defined broadly enough that few enrollees were made more distant from their nearest provider as a result of the selective contracting network. For example, of those individuals who had a mental health provider in the same zip code as their residential zip code during 1995 (two-thirds of the sample), only 4% no longer had a provider in the same zip code during 1997. Given the limited number of enrollees for whom selective contracting was binding, it is not surprising that no effect of selective contracting was found. However, distance clearly was an important factor affecting treatment utilization, thus had the provider network been more narrowly defined, treatment initiation would have been affected.

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 day) period in which no mental health treatment was observed (Huskamp, 1999; Keeler et al., 1988). We exclude initiations occurring within the first 2 months of the year to consistently measure initiation without censoring throughout all 4 years of our study. Following earlier research (Holmes & Deb, 1998; Sturm et al., 1996), 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) generalist MD providers; (2) MD psychiatrist

Table 2. Descriptive Statistics for Outcome Measures, By Gender and Pre/Post Benefit Change, 1995–1998.

	Pre-Period (1995)		Post-Period (1996–1998)	
	Female	Male	Female	Male
Observations	7,381	16,916	22,143	50,748
Initiation				
Any initiation percentage	8.31%	4.94%	9.34%	5.71%
Number of initiations	613	835	2069	2899
Initiation type				
General	50.24%	53.77%	43.69%	42.64%
Non-MD	35.73%	32.22%	48.96%	49.95%
MD Psych	14.03%	14.01%	7.35%	7.42%
Outpatient visits				
All visits	14.19	10.20	17.97	13.95
Mental health visits	6.61	5.59	8.23	7.33
Severity				
Major depression	27.57%	19.04%	30.64%	22.35%
Manic depressive	4.24%	4.91%	3.43%	3.79%
Anxiety	20.72%	18.68%	16.92%	16.90%
Personality disorder	4.57%	8.86%	3.53%	8.00%
Alcohol or drug abuse	0.98%	0.48%	0.10%	0.79%
Conduct disorder	1.63%	1.68%	1.21%	1.76%

providers; and (3) non-MD mental health providers. Finally, in order to examine the intensity of outpatient treatment services utilization, we compute the total number of outpatient visits for any reason and for mental health visits, as identified via ICD-9 diagnosis and procedure codes.

Table 2 shows descriptive statistics in the pre- and post-periods for the outcome measures in our study. Note that the percentage of continuously enrolled persons initiating mental health treatment increased by similar amounts across gender. For women the percentage initiating increased from 8.3 to 9.3%, while for men it increased from 4.9 to 5.7%. For both women and men the fraction of initiations occurring with non-MD mental health specialists increased from 35.7 to 49% and 32.2 to 50%, respectively. By contrast, initiations occurring with generalists and psychiatrists dropped by similar amounts for both women and men. Lastly, both any outpatient visits and outpatient mental health visits (measured annually) increased nearly identical amounts for men and women.

Table 3 displays descriptive statistics for our sample. Because the sample is continuously enrolled for the 4-year period, we present means taken from the 1995 data. The table indicates that women were more likely to be non-white relative to men: about 70% of women were white and 80% of men were white. Women were more likely to have single coverage relative to men: about 50% of women had

Table 3. Descriptive Statistics, Continuously Enrolled Persons from 1995–1998.

	Female	Male	Total
Sample size	7,381	16,916	24,297
Race			
White	71.09%	81.16%	78.1%
Non-White	28.91%	18.84%	21.9%
Coverage status			
Single	49.45%	32.35%	37.54%
Family	50.55%	67.65%	62.46%
Age category			
18–29	14.39%	15.69%	15.29%
30–39	32.37%	36.84%	35.48%
40–49	33.33%	30.69%	31.49%
50–64	19.92%	16.78%	17.73%
Region			
Central	44.32%	43.00%	43.40%
South	28.76%	30.98%	30.31%
West	26.92%	26.02%	26.29%

single coverage while only a third of men had single coverage. This likely reflects the fact that men traditionally are the source of health insurance for most families. The age distribution by gender was roughly comparable, though it appears that women tended to populate the higher age brackets relative to men. Women and men were comparably distributed across regions of the country. Most employees were located in the central region.

METHODS

Our econometric model accounts for individual heterogeneity over our 4-year time period given the following random utility framework:

$$\mathrm{U}_{it}^{*} = \alpha^{G}\mathrm{Post}_{it} + X_{it}'\beta + \nu_{it}, \tag{1}$$

where U_{it}^{*} is the excess utility from initiating treatment for individual i at time t; Post represents an indicator variable for whether the observation is made during the 1996–1998 benefit expansion period, with α^{G} measuring the total effect of the benefit expansion; the superscript G indicates gender; X is a demographic matrix including gender, race dummy, family status dummy, age dummy, and regional dummies; b is the vector of coefficients associated with X; and ν_{it} is the error term. We assume that the error term ν_{it} is constructed from an individual-specific unobserved effect and an idiosyncratic component, $\nu_{it} = A_i + \varepsilon_{it}$, where A_i and ε_{it} are assumed to be logistically distributed. This framework allows us to control for any within-person correlation over time of the decision to initiate treatment. We do not observe U_{it}^{*}, but only the decision to seek treatment, Y_{it}. Thus, we define $Y_{it} = 1$ if $\mathrm{U}_{it}^{*} > 0$ and zero otherwise. We estimate the following model:

$$\Pr(Y_{it} = 1) = F(\alpha^{G}\mathrm{Post}_{it} + X_{it}'\beta | A_i), \tag{2}$$

which can be estimated with a random effects logit model. We estimate stratified regressions by gender as well as a pooled model with a gender times post-period interaction term to identify differential gender effect.

A concern with the random effects specification arises when unobserved attributes of individuals are correlated with the observed attributes; that is, $\mathrm{E}(\nu_{it}|X_{it}, \mathrm{Post}_{it}) \neq 0$. In this circumstance, the random effects model is inconsistent. For example if stigma, an unobserved attribute affecting individual willingness to seek treatment, is affected by the benefit change as embodied in the post-period dummy variable, then the estimated coefficients of the model are potentially biased. As an alternative specification to test the robustness of our random effects results, we will estimate a fixed effects logit model. As a matter of practice the fixed effects logit model is estimable conditional upon the binary dependent variable changing over time for a given individual (Baltagi, 1995). In our context, a person must have some years with treatment and some years without treatment. The implication of this conditioning restriction is that the vast majority of observations contribute nothing to the (conditional) likelihood function and, therefore, nothing to the coefficient estimates. Nonetheless, the fixed effects model does provide a means of validating the results in the random effects model.

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 not everyone initiates each year. Hence, we opt for a multinomial choice model to estimate the conditional probability of initiation of treatment with each provider type. In cases where there are repeated observations for the same individual over time, 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 generalist, a psychiatrist, and a non-MD mental health professional and then selecting the option with the greatest expected utility. The independent variables are the same as above, except that we include separate distance measures to each of the three provider 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 a particular provider type if the distance to the nearest provider of that type is far. Formally, we posit:

$$E(\mathrm{U}_i^{\mathrm{Gen}}|Y_i = 1) = \alpha^{\mathrm{Gen}}\mathrm{Gender}_i + \theta^{\mathrm{Gen}}\mathrm{Post}_i + \delta^{\mathrm{Gen}}\mathrm{Gender}_i \times \mathrm{Post}_i + X_i'\beta^{\mathrm{Gen}} + \nu_i^{\mathrm{Gen}} \tag{4}$$

$$E(\mathrm{U}_i^{\mathrm{Psych}}|Y_i = 1) = \alpha^{\mathrm{Psych}}\mathrm{Gender}_i + \theta^{\mathrm{Psych}}\mathrm{Post}_i + \delta^{\mathrm{Psych}}\mathrm{Gender}_i \times \mathrm{Post}_i + X_i'\beta^{\mathrm{Psych}} + \nu_i^{\mathrm{Psych}} \tag{5}$$

$$E(\mathrm{U}_i^{\mathrm{nonMD}}|Y_i = 1) = \alpha^{\mathrm{nonMD}}\mathrm{Gender}_i + \theta^{\mathrm{nonMD}}\mathrm{Post}_i + \delta^{\mathrm{nonMD}}\mathrm{Gender}_i \times \mathrm{Post}_i + X_i'\beta^{\mathrm{nonMD}} + \nu_i^{\mathrm{nonMD}} \tag{6}$$

where U_i is the utility associated with the decision to initiate treatment with each of the three different provider types – Gen for generalist MD mental health providers, Psych for MD psychiatrist providers, and nonMD for non-MD mental health professionals. Y_i represents the decision to initiate treatment, which here equals unity because we are conditioning on treatment initiation, and the θ terms represent the effect of the change in benefits on the conditional probability to initiate treatment at each provider type. A positive θ does not necessarily imply that individuals get more utility in the post-period; rather, the destigmatization campaign and the associated literature on providers increased the expected utility of going to that particular provider type. The δ terms represent the gender-specific impact of the benefit expansion on the willingness to use each provider type. If the ν terms follow an extreme value distribution then a multinomial logit model can be used to estimate the model embodied in (4)–(6).

We will use ordinary least squares (OLS) regressions to examine the differential impact of the benefit change on the number of outpatient visits, conditional on initiation. For these models we will use annualized measures of outpatient visits as the dependent variable.

A potential limitation in our analysis is that we observe a pre/post change 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. We only have one year of pre-benefit change data, thus it is not feasible to include a time trend in addition to a post-period indicator because the trend term would not allow us to differentiate between actual secular trend in treatment and a lagged implementation of the benefit change effect. Using two national surveys, we computed measures of mental health treatment utilization and initiation over our time period. The National Household Survey on Drug Abuse (NHSDA) measures any mental health treatment utilization consistently for a period coincident with our analysis period, 1995–1998. Because the NHSDA contains a self-reported measure of any mental health treatment for the previous 12 months, the values are not directly comparable to our claims-based measure of treatment initiation. However, there was no discernable trend over our analysis period in self-reported mental health treatment utilization. Similarly, because of the concern that the NHSDA does not produce a measure comparable to our claims-based initiation measure, we used the Medical Expenditure Panel Survey (MEPS), which includes claims data for a sub-sample of respondents, to construct a measure of mental health treatment initiation more comparable to our own measure. Unfortunately, the MEPS data begin in 1996, and therefore its period does not directly overlap with our analysis period. However, the MEPS initiation measure also indicates that there was no discernable trend in initiation throughout the late 1990s. Thus, given that there are no apparent trends in the rate of mental health treatment initiation over the late

1990s, we believe that our data allow us to accurately identify the impact of the benefit change at this employer.

RESULTS

Results for Treatment Initiation

In Table 4, we present pooled and gender-stratified random effects logit regression results for the factors affecting treatment initiation for our sample. We observe that coincident with the implementation of the mental health benefit change, continuously enrolled individuals were significantly more likely to initiate mental health treatment. The odds ratio indicates that the overall rate of initiation was about 13% higher in the post-period. Interestingly, while men were significantly less likely to initiate mental health treatment overall, men and women did not experience differential rates of initiation in response to the benefit change. Men and women both were more likely to initiate mental health treatment in the response to the benefit change. The small difference between the gender effects was not significant (*Post* × *Male*).

Several differential relationships between the demographic variables and initiation are apparent when examining the stratified results. The natural log of distance to the nearest mental health provider is negative and significant for women but not for men, suggesting that distance was a greater apparent barrier to treatment for women than it was for men. This result could be indicative of greater transportation issues for women or greater time burdens for women (e.g. child care). The age profile of treatment initiation differed somewhat by gender. Note that women aged 50–64 were less likely to initiate treatment than were women in their 20s. Women's initiation of mental health treatment appeared to peak at 30–39 years of age. By contrast, men 30 and older were more likely to initiate mental health treatment than men in their 20s. Gender differences were not apparent by single vs. family coverage, by region of the country, or by race.

Because of the aforementioned concern regarding the random effects specification, Table 5 displays results from fixed effects conditional logit regressions. Note that any time-invariant independent variables are eliminated by the fixed effect regression because they are subsumed in the fixed effect term. As observed in the random effects specification, men and women both experienced increases in the likelihood of initiating mental health treatment: the overall rate of initiation was 16% higher in the post-period with the women's initiation rate 16% higher and the men's nearly 20% higher. However, the difference between the genders was not statistically significant.

Table 4. Random Effect Logit Regression for Mental Health Treatment Initiation.

	Pooled	Female	Male
Post	0.124***	0.145***	0.128***
	(0.038)	(0.039)	(0.031)
	[1.131]	[1.156]	[1.136]
Male	−0.583***	–	–
	(0.050)		
	[0.558]		
Post × Male	0.017	–	–
	(0.048)		
	[1.017]		
Log of distance to nearest provider	−0.048**	−0.093***	−0.031
	(0.019)	(0.035)	(0.024)
	[0.953]	[0.911]	[0.969]
Ages 30–39 relative to 18–29	0.263***	0.238***	0.301***
	(0.051)	(0.080)	(0.068)
	[1.301]	[1.269]	[1.351]
Ages 40–49 relative to 18–29	0.231***	0.026	0.378***
	(0.054)	(0.083)	(0.072)
	[1.260]	[1.026]	[1.459]
Ages 50–64 relative to 18–29	−0.029	−0.271***	0.149*
	(0.058)	(0.090)	(0.078)
	[0.971]	[0.762]	[1.161]
South relative to Central	0.265***	0.326***	0.242***
	(0.037)	(0.063)	(0.047)
	[1.303]	[1.386]	[1.274]
West relative to Central	0.366***	0.452***	0.316***
	(0.038)	(0.061)	(0.048)
	[1.443]	[1.571]	[1.371]
Single coverage relative to family	−0.107***	−0.083*	−0.099**
	(0.032)	(0.049)	(0.044)
	[0.899]	[0.920]	[0.905]
Non-White	−0.462***	−0.589***	−0.343***
	(0.040)	(0.061)	(0.054)
	[0.630]	[0.555]	[0.710]
Constant	−2.643***	−2.542***	−3.316***
	(0.068)	(0.093)	(0.082)
N (observations)	97,188	29,524	67,664
N (Individuals)	24,297	7,381	16,916

Note: Odds ratios in square brackets.
*Indicates the coefficient is significant at 10%.
**Indicates the coefficient is significant at 5%.
***Indicates the coefficient is significant at 1%.

Table 5. Fixed Effect Conditional Logit Regression for Mental Health Treatment Initiation.

	Pooled	Female	Male
Post	0.148**	0.151**	0.180***
	(0.058)	(0.061)	(0.051)
	[1.159]	[1.163]	[1.197]
Post × Male	0.035	–	–
	(0.074)		
	[1.036]		
Log Distance	0.174	0.657**	−0.069
	(0.170)	(0.317)	(0.199)
	[1.190]	[1.929]	[0.933]
Ages 30–39 relative to 18–29	0.180	0.163	0.199
	(0.131)	(0.191)	(0.179)
	[1.197]	[1.177]	[1.220]
Ages 40–49 relative to 18–29	0.164	–0.060	0.319
	(0.166)	(0.251)	(0.223)
	[1.178]	[0.942]	[1.376]
Ages 50–64 relative to 18–29	0.166	−0.154	0.391
	(0.207)	(0.318)	(0.275)
	[1.181]	[0.857]	[1.478]
Single coverage relative to family	−0.455***	−0.563***	−0.354**
	(0.127)	(0.184)	(0.176)
	[0.634]	[0.569]	[0.702]
N	16148	6508	9640

Note: Sample sizes reflect only those observations that experienced a change from use to non-use or non-use to use during the period 1995–1998.

Odds ratios in square brackets.

** Indicates the coefficient is significant at 5%.

*** Indicates the coefficient is significant at 1%.

Results Conditional Upon Treatment Initiation

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 post-periods. For example, if the destigmatization effort was successful we would anticipate that people with less severe disorders would initiate in the post-period. We use the severity indicators defined in Ettner et al. (1998) for the purposes of risk adjusting mental health claims data. Similarly, the selective contracting effort, while not observed to have a strong impact on the overall sample, might have had a differential gender impact. We descriptively explore these issues in Table 6.

Table 6. Distance to Nearest Mental Health Provider and Severity Indicators, Pre and Post Implementation of Benefit Change, by Gender.

	Pre-Period		Post-Period	
	Female	Male	Female	Male
Sample size	7,381	16,916	22,143	50,748
Distance to closest provider				
Zero miles (same zip code)	65.21%	66.83%	62.43%	64.26%
0–2 miles	19.43%	16.55%	19.94%	17.07%
2–5 miles	7.89%	8.25%	9.57%	9.69%
5+ miles	7.48%	8.36%	8.06%	8.97%
Initiations	613	835	2069	2899
Severity among initiates				
Major depression	27.57%	19.04%	30.64%	22.35%
Manic depressive	4.24%	4.91%	3.43%	3.79%
Anxiety disorder	20.72%	18.68%	16.92%	16.90%
Personality disorder	4.57%	8.86%	3.53%	8.00%
Alcohol or drug abuse	0.98%	0.48%	0.10%	0.79%
Conduct disorder	1.63%	1.68%	1.21%	1.76%

Table 6 indicates that, as expected based on the imposition of selective contracting, men and women were generally farther from their nearest mental health provider in the post-period relative to the pre-period. A smaller fraction of people had a mental health provider in their residential zip code (zero miles) in the post-period; the people who previously had a provider in their residential zip code were distributed across the other three more distant categories. However, no differential impact of the policy change is apparent by gender. When attention is restricted to those who initiated treatment, we observe that the fraction of initiates that exhibited major depression or personality disorder diagnoses increased for both men and women; the fraction of initiates that exhibited manic depressive or anxiety disorder symptoms decreased for both men and women. Substance abuse symptoms and conduct disorders were more prevalent in the post-period for men but less so for women. However, these last two diagnoses were quite rare in the data set.

In Table 7, we present the results of the multinomial logit regression of provider type conditional on initiation. 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 post-period treatment initiators were significantly more likely to initiate treatment with a non-MD mental health specialist than with a generalist and significantly

Table 7. Multinomial Logit Regression for Provider Type, Relative to Initiation to a Generalist Provider, Pooled and Stratified by Gender.

	Pooled		Female		Male	
	Non-MD	Psychiatrist	Non-MD	Psychiatrist	Non-MD	Psychiatrist
Post	0.519***	−0.606***	0.505***	−0.600***	0.736***	−0.445***
	(0.099)	(0.151)	(0.098)	(0.152)	(0.087)	(0.135)
	[1.680]	[0.545]	[1.656]	[0.549]	[2.088]	[0.641]
Male	−0.197	−0.136	–	–	–	–
	(0.122)	(0.170)				
	[0.821]	[0.873]				
Post × Male	0.203	0.190	–	–	–	–
	(0.131)	(0.203)				
	[1.225]	[1.209]				
Log of distance to generalist	−0.006	−0.011	0.112	−0.056	−0.071	0.001
	(0.054)	(0.102)	(0.089)	(0.179)	(0.070)	(0.124)
	[0.994]	[0.989]	[1.119]	[0.945]	[0.932]	[1.001]
Log of distance to non-MD specialist	−0.108**	0.090	−0.146*	0.120	−0.106	0.068
	(0.051)	(0.091)	(0.081)	(0.146)	(0.067)	(0.115)
	[0.897]	[1.094]	[0.864]	[1.127]	[0.900]	[1.070]
Log of distance to MD Psychiatrist	−0.032	−0.320***	−0.069	−0.374***	−0.002	−0.283***
	(0.039)	(0.066)	(0.061)	(0.099)	(0.051)	(0.090)
	[0.968]	[0.726]	[0.933]	[0.688]	[0.998]	[0.754]
Ages 30–39 relative to 18–29	0.046	0.031	−0.190	−0.280	0.233	0.295
	(0.123)	(0.187)	(0.180)	(0.260)	(0.170)	(0.272)
	[1.047]	[1.031]	[0.827]	[0.756]	[1.262]	[1.344]
Ages 40–49 relative to 18–29	−0.206	−0.398**	−0.418**	−0.898***	−0.051	−0.040
	(0.127)	(0.198)	(0.185)	(0.283)	(0.176)	(0.284)
	[0.814]	[0.672]	[0.658]	[0.407]	[0.950]	[0.961]
Ages 50–64 relative to 18–29	−0.435***	−0.392*	−0.764***	−0.647**	−0.231	−0.206
	(0.137)	(0.222)	(0.199)	(0.308)	(0.195)	(0.318)
	[0.648]	[0.675]	[0.466]	[0.524]	[0.794]	[0.814]
South relative to Central	0.316***	−0.878***	0.169	−0.792***	0.395***	−0.932***
	(0.082)	(0.142)	(0.129)	(0.229)	(0.106)	(0.183)
	[1.372]	[0.416]	[1.184]	[0.453]	[1.484]	[0.394]
West relative to Central	0.411***	−1.247***	0.315**	−1.110***	0.447***	−1.369***
	(0.082)	(0.153)	(0.126)	(0.238)	(0.110)	(0.195)
	[1.508]	[0.287]	[1.371]	[0.330]	[1.563]	[0.254]
Single	−0.035	−0.137	0.093	0.017	−0.113	−0.269
	(0.072)	(0.119)	(0.102)	(0.176)	(0.105)	(0.165)
	[0.965]	[0.872]	[1.097]	[1.017]	[0.893]	[0.764]
Non-White	0.010	−0.137	0.241*	−0.186	−0.175	−0.072
	(0.088)	(0.157)	(0.125)	(0.246)	(0.124)	(0.205)
	[1.010]	[0.872]	[1.273]	[0.830]	[0.839]	[0.930]
Major depression	−0.859***	−0.280**	−0.760***	−0.380**	−0.962***	−0.185
	(0.075)	(0.122)	(0.107)	(0.185)	(0.105)	(0.161)
	[0.424]	[0.755]	[0.468]	[0.684]	[0.382]	[0.831]
Manic depressive	−1.065***	−0.931***	−0.712***	−0.484	−1.295***	−1.276***
	(0.194)	(0.328)	(0.268)	(0.502)	(0.270)	(0.407)
	[0.345]	[0.394]	[0.491]	[0.616]	[0.274]	[0.279]

Table 7. (*Continued*)

	Pooled		Female		Male	
	Non-MD	Psychiatrist	Non-MD	Psychiatrist	Non-MD	Psychiatrist
Anxiety	−0.555***	−0.590***	−0.422***	−0.672***	−0.665***	−0.536***
	(0.085)	(0.145)	(0.128)	(0.240)	(0.115)	(0.185)
	[0.574]	[0.554]	[0.655]	[0.510]	[0.515]	[0.585]
Personality disorder	0.222	0.469	0.348	0.100	0.118	0.629
	(0.251)	(0.373)	(0.374)	(0.642)	(0.341)	(0.448)
	[1.249]	[1.599]	[1.416]	[1.105]	[1.125]	[1.875]
Alcohol or drug abuse	−0.630***	−0.249	−0.442*	−0.625	−0.681***	−0.073
	(0.124)	(0.209)	(0.231)	(0.446)	(0.148)	(0.239)
	[0.532]	[0.779]	[0.642]	[0.536]	[0.506]	[0.930]
Conduct disorder	0.942**	0.049	–[a]	–[a]	0.975*	0.493
	(0.435)	(0.794)			(0.516)	(0.827)
	[2.566]	[1.050]			[2.651]	[1.637]
Constant	0.044	0.169	0.218	0.494	−0.289	−0.238
	(0.161)	(0.249)	(0.211)	(0.325)	(0.199)	(0.306)
N	6416		2682		3734	

Note: [a] Because of small cell sizes Conduct Disorder had to be dropped from the female model to achieve convergence.
Robust standard errors in parentheses.
Odds ratios in square brackets.
*Indicates the coefficient is significant at 10%.
**Indicates the coefficient is significant at 5%.
***Indicates the coefficient is significant at 1%.

less likely to initiate treatment with an MD psychiatrist than with a generalist, even when controlling for the severity indictors. The post-period coefficients indicate that conditional upon entry into to mental health treatment, women and men were roughly 65% and 110% more likely to initiate with a non-MD mental health professional than with a generalist, respectively. At the same time, women and men were 45% and 35% less likely to initiate treatment with an MD psychiatrist than with a generalist, respectively. However, in both cases the gender differential was insignificant as indicated by the interaction terms in the pooled model.

The relationship of initiation and distance to each type of provider confirmed our expectations: the greater the travel distance to non-MD specialists, the less likely people were to initiate with them; the same held true for psychiatrists as well. However, as observed in the initiation results, we again see that men appear generally less sensitive to distance than did women. Among the other coefficients in the regression, older women appeared much more likely to initiate treatment with generalists, while no age profile was evident for men. This suggests that the policy to encourage the use of non-MD specialists was more successful among younger individuals, particularly women. No other noteworthy gender differences were apparent in the regression results.

One concern in interpreting the multinomial logit results is that they describe the relative shift in provider type, but do not necessarily indicate the absolute change in provider type utilization. In order to calculate the unconditional effects of the benefit change on provider type, we estimate a sequential or two-part model of provider type. Hence for each of the three provider types we calculate:

$$\begin{aligned}\Pr(\text{Provider} = Z) = {} & \Pr(Y = 0) \cdot \Pr(\text{Provider} = Z|Y = 0) \\ & + \Pr(Y = 1) \cdot \Pr(\text{Provider} = Z|Y = 1),\end{aligned}$$

which because $(\text{Provider} = Z|Y = 0) = 0$ yields:

$$\Pr(\text{Provider} = Z) = \Pr(Y = 1) \cdot \Pr(\text{Provider} = Z|Y = 1),$$

where provider Z represents generalist, non-MD specialist, and psychiatrist and Y represents the decision to initiate treatment. The initiation probabilities are estimated with a pooled logit model and the provider type probabilities are calculated from the multinomial logit model. The unconditional estimates are provided in Table 8. Note that the unconditional predicted probability estimates confirm the pattern of results observed in the multinomial logit results: the probability of using non-MD specialists increased and the probability of using psychiatrists decreased, while generalist use decreased slightly.

Table 9 displays pooled and gender stratified results for total outpatient visits and total outpatient mental health visits. Note once again that for both outpatient measures, very similar post-period coefficients are observed for men and women. Both men and women have on average nearly 4 more outpatient visits in the post-period and nearly 2 more mental health outpatient visits in the post-period. Despite the generally lower baseline number of outpatient visits for men, the benefit change appears to result in similar increases across gender. In contrast to the earlier regressions, men appear more sensitive to distance when it comes to the intensity of service use, conditional on initiation, than did women. Other coefficients in the model are generally consistent across gender.

Table 8. Unconditional Provider Type Probability Estimates.

Regression	General		Non-MD		MD Psych	
	Pre (%)	Post (%)	Pre (%)	Post (%)	Pre (%)	Post (%)
Pooled						
Female	4.60	4.37	3.41	5.09	1.22	0.70
Male	2.85	2.55	1.77	3.11	0.69	0.42
Stratified						
Female	4.66	4.47	3.44	5.18	1.26	0.72
Male	2.85	2.54	1.77	3.10	0.69	0.42

Table 9. Regressions of Total Outpatient Visits and Total Outpatient Mental Health Visits Per Year Conditional on Initiation.

	Total Outpatient Visits			Mental Health Visits		
	Pooled	Female	Male	Pooled	Female	Male
Post	3.789***	3.805***	3.667***	1.743***	1.776***	1.743***
	(0.507)	(0.508)	(0.394)	(0.294)	(0.293)	(0.238)
Male	−3.415***	–	–	−0.606*	–	–
	(0.593)			(0.327)		
Post × Male	−0.107	–	–	0.007	–	–
	(0.638)			(0.377)		
Log of distance	−0.445**	−0.368	−0.487**	−0.372***	−0.038	−0.549***
	(0.213)	(0.404)	(0.246)	(0.117)	(0.227)	(0.129)
Ages 30–39 relative to 18–29	1.959***	1.591*	2.387***	1.364***	1.008*	1.736***
	(0.592)	(0.880)	(0.818)	(0.370)	(0.566)	(0.492)
Ages 40–49 relative to 18–29	2.882***	2.293**	3.525***	0.903**	0.394	1.441***
	(0.623)	(0.924)	(0.871)	(0.377)	(0.579)	(0.508)
Ages 50–64 relative to 18–29	4.598***	3.409***	5.615***	−0.131	−0.980	0.604
	(0.728)	(1.072)	(1.015)	(0.426)	(0.646)	(0.589)
South relative to Central	0.140	0.546	−0.133	0.448*	0.379	0.501
	(0.453)	(0.789)	(0.542)	(0.264)	(0.444)	(0.328)
West relative to Central	0.431	0.418	0.477	0.250	0.381	0.167
	(0.455)	(0.742)	(0.568)	(0.258)	(0.434)	(0.315)
Single relative to family	1.481***	1.362**	1.727***	0.823***	0.837**	0.925***
	(0.430)	(0.621)	(0.609)	(0.247)	(0.355)	(0.354)
Non-White	−0.551	−1.180	0.154	−0.714**	−1.037***	−0.331
	(0.489)	(0.752)	(0.636)	(0.294)	(0.381)	(0.451)
Major depression	4.073***	5.327***	2.929***	3.510***	4.224***	2.866***
	(0.461)	(0.724)	(0.587)	(0.284)	(0.428)	(0.378)
Manic depressive	4.226***	7.536***	2.191**	4.274***	6.296***	3.009***
	(1.025)	(1.960)	(1.061)	(0.697)	(1.308)	(0.740)
Anxiety	2.426***	2.195***	2.594***	1.751***	1.530***	1.903***
	(0.482)	(0.729)	(0.638)	(0.306)	(0.425)	(0.434)
Personality disorder	5.145***	7.967***	3.522*	4.955***	7.517***	3.460***
	(1.488)	(2.491)	(1.811)	(0.971)	(1.759)	(1.115)
Alcohol or drug abuse	−0.259	0.802	−0.650	0.973**	2.406**	0.501
	(0.579)	(1.327)	(0.644)	(0.449)	(1.106)	(0.471)
Conduct disorder	4.664**	3.010	4.778*	4.897***	4.435	4.687**
	(2.185)	(3.414)	(2.640)	(1.809)	(2.862)	(2.186)
Constant	9.094***	9.157***	5.391***	3.751***	3.793***	2.900***
	(0.763)	(0.989)	(0.896)	(0.469)	(0.610)	(0.586)
N	6416	2682	3734	6416	2682	3734

Note: Robust standard errors in parentheses.
*Indicates the coefficient is significant at 10%.
**Indicates the coefficient is significant at 5%.
***Indicates the coefficient is significant at 1%.

DISCUSSION

Our results indicate that while the mental health benefit policy change at this company had significant effects on treatment initiation, provider type choice, and treatment service use, no significant differences in the benefit impact were apparent by gender. This finding runs counter to previous research indicating differential gender impacts of mental health interventions. Thus, while the benefit change did not appear to narrow the gap in treatment seeking behavior among men and women, it did not enlarge the gap.

One concern that emerged from the stratified regression results for initiation is that distance to providers was more of an apparent barrier to treatment for women than it was for men. While the importance of geographic accessibility of providers has been noted previously (e.g. Fortney et al., 1999), a gender differential in the impact of distance has not been noted. This finding has implications for network design in other corporate or Medicaid settings. Although in this case the selective contracting network was not defined narrowly enough to represent a barrier to a large number of employees, in other settings benefit designers may define narrow networks and thus impair women's access to mental health care.

This study has a number of limitations that can make generalizing to other settings problematic. First, data are only from one company and therefore company-specific factors could confound relationships observed in the data. Second, as a consequence of using a pre/post design and a single company, it is difficult to say for certain that the relationships observed are not the result of secular trends. However, an examination of employed individuals from the MEPS and the NHSDA did not reveal general trends in initiation or outpatient utilization between 1995 and 1998, thus we believe that the secular trend was relatively flat during our analysis period and that the differences observed in this study represent the impact of the benefit change. Third, in an effort to control for the possibility of selection effects we used data on employees continuously enrolled in the company's self-insured health plan. In doing so, there is the potential for bias in our estimates because we only examine people who have relatively stable employment and they may have atypical mental health treatment patterns.

Additional work is needed to examine whether the change in company benefits was successful in improving quality of care for persons with mental illness disorders for all employees. However, our results do show that a combination of destigmatization and lowered copayments significantly increased the likelihood of treatment initiation regardless of gender. Intensity of outpatient treatment increased similar amounts for men and women. In addition, the benefit change was associated with a large increase in the probability of initiating with a non-MD mental health professional for both men and women. This result is perhaps surprising because the mid- to late-1990s was a period in which highly effective pharmacotherapy became common in treatment of mental disorders. However, given that most non-MD mental health specialists had prescribing arrangements with MDs, the jump in non-MD specialist use was not inconsistent with the rising trend toward pharmaceutical use. We believe that the observed impact of the benefit change in provider type was consistent with the company's desire to encourage treatment from the "least intensive locus of care." This goal was furthered by including a large number of non-MD mental health specialists in the selective contracting network.

The specific reason for the lack of a difference by gender when previous studies tend to find differences by gender is unclear. It does not appear that the company made a special outreach effort to encourage men or women to take-up treatment to a greater extent than they otherwise might. Of note is that prior to the program, more women than men sought treatment for their disorders. One possible explanation for the lack of gender difference in the effect of the program is that we are studying changes rather than levels. In other words, in levels there are gender differences in the rate and probability of initiation pre and post, but the effect of the program on the rate of initiation is not gender-specific. It is also possible that composition of the firm mattered: it employs many engineers and scientists, a group that is more likely to be comprised of men. Many of the women employed by the company are in either manufacturing or administrative jobs, which tend to require less education. To the extent that lower educated individuals may be less likely to initiate treatment, the difference in the composition of the employees at this company could explain the pattern of the results.

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GETTING MEN AND WOMEN RECEIVING DEPRESSION-RELATED SHORT-TERM DISABILITY BENEFITS BACK TO WORK: WHERE DO WE BEGIN?

Carolyn S. Dewa, Jeffrey S. Hoch and Paula Goering

INTRODUCTION

During the past two decades, there has been a growing awareness of the impact of mental illness on the population (Regier et al., 1988; U.S. Department of Health and Human Services, 1999; World Health Organization, 2001). However, only recently have issues surrounding its effect on the labor force been raised (Berndt et al., 2000; Dewa & Lin, 2000; Kessler et al., 1999; Marcotte et al., 1999; Stewart et al., 2003).

Part of this new focus is based on the growing realization that the number of disability claims for mental and nervous disorders is increasing. The Health Insurance Association of America (HIAA) (1995) reports that between 1989 and 1994, mental and nervous disorder claims doubled. In addition, in their 1994 survey, HIAA found respondent companies spent between $360 and $540 million on disability claims related to this group of disorders.

The effects of mental illness on the labor market are a valid and pressing concern. Yet, little is known about the working population disabled by mental and nervous

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disorders. Absence of this basic type of information makes it virtually impossible to plan occupational health programs to either help prevent the occurrence of disability related to nervous and mental disorders or promote return to work.

In a previous study, we began to address this gap in knowledge by examining the degree to which the general population's experience with depression is similar to the working population's disabled by depression. Population-based studies characterizing the general population and working population suffering from psychiatric disorders indicate the prevalence of depression is greater among women than men (Kessler et al., 1994; Marcotte et al., 1999; Offord et al., 1996). This prevalence pattern appears to be consistent among the working population using short-term disability benefits (Dewa et al., 2002). Specifically, we observed that about 2.5% of all employees in the companies had at least one short-term disability episode related depression. There was a statistically significant difference in the prevalence rates among males and females; 1% of male employees and 3% of female employees incurred at least one depression-related short-term disability episode during a 12-month period.

We also observed that compared to women, men were significantly less likely to return to work at the end of the disability episode (77.3% vs. 69.4%, respectively) (Dewa et al., 2002). This finding raised a number of questions including what factors are associated with disability outcomes for men and women and do they have implications for planning workplace interventions?

The purpose of this study is to examine gender differences among factors that are associated with disability outcomes of workers on depression-related short-term disability. To this end, we address two questions. First, do men and women with depression-related short-term disability claims differ in terms of age, work experience and antidepressant use? Second, to what extent are there differences in the association between these different factors and the return-to-work of men and women?

BACKGROUND

Much of the literature on labor market disability focuses on the impact of workplace factors on productivity. Two main areas of concentration have included the relationship between stress and job performance (Van der Heck & Plomp, 1997) and the role of workplace support systems on disability outcomes (Akabas, 1995). Less attention has been paid to the potential contribution of pharmaceutical interventions particularly the newer antidepressants that were introduced during the early 1990s.

The introduction of fluoxetine in the late 1980s ushered in a new era of antidepressant therapy. It was the first of a new group of antidepressants known collectively as selective serotonin reuptake inhibitors (SSRIs). By 1994, there were four different SSRIs on the market. During the past decade, this family of drugs experienced a phenomenal growth (Olfson et al., 2002; Simon et al., 1993). In 1999, three of the SSRIs were in the top 10 most frequently used drugs in the U.S. (Cauchon, 1999). In addition, they were one of the major contributors to the growth in psychotropic expenditures during the past five years (Dewa & Goering, 2001; Foote & Etheredge, 2000; NIHCM, 2000). Thus, it appears that antidepressants are being widely used. But, to what benefit?

Only a handful of studies have examined the relationship between antidepressant use and outcomes in the workplace. For example, using data from a clinical trial, Berndt and his colleagues (1998) found evidence of a positive relationship between worker self-perceived low productivity and severity of depression. They also observed the use of the antidepressants, sertraline and imipramine, had a significant impact on the severity of depression. Through inductive reasoning, one might conclude there is an association between antidepressant treatment and workplace functioning. However, Berndt et al.'s study does not directly test the impact of antidepressant treatment and workplace functioning, thereby stopping short of examining the direct relationship between antidepressant use and productivity.

In their analysis, Mintz and colleagues (1992) pooled data from 10 studies and used the Social Adjustment Scale in an attempt to measure the direct impact of treatment on productivity. They found their productivity measure was positively associated with treatment. In addition, they identified symptom remission and longer duration of treatment (i.e. four to six months) as the most important predictors of work impairment. But, their measure of productivity is more difficult to translate into policy recommendations than specific measures of work performance would be.

Using administrative data to examine the relationship between absenteeism and treatment, Claxton and colleagues (1999) observed differences among various antidepressants in terms of mean lost work days. Comparing two types of antidepressants (tricyclic amines (TCAs) SSRIs), they found a lower average number of days absent for the group using SSRIs. These results offer an important first step toward understanding the impact of antidepressant treatment on absenteeism. However, they did not look at or control for other factors that could also be associated with absenteeism such as age, sex and patterns of antidepressant use.

In addition, none of these studies specifically focused on whether there were differences between how men and women used antidepressants and whether these differences were associated with workplace productivity. Yet, given that it is well documented that men and women differ in their use of mental health-related services in general (Katz et al., 1997; Lin et al., 1996; Rhodes et al., 2002), this may be an important line of inquiry.

METHODS

Data Sources

This study was conducted using administrative data from three major Canadian financial/insurance sector employers. At the time of the project, the companies had a combined workforce of approximately 63,000 employees nationwide, representing approximately 12% of their sector's workforce (Statistics Canada, 1996).

The primary information sources were company short-term disability claims, prescription drug claims, and occupational health department records. Because of its relatively smaller size, claims from one company were taken for short-term disability episodes beginning between January 1996 and December 1998. For the remaining two, data were abstracted from claims beginning in 1997 or 1998.

Company Disability Management

The three participating companies self-insured for short-term disability. This arrangement is representative of many mid- to large-size employers. For example, in their survey, Watson Wyatt (1997) found 53% of the firms they surveyed self-administered their short-term disability benefits while 45% depended on third-party administration (e.g. insurance carriers) with the remainder covered by government programs.

In each company, short-term disability benefits were managed by the occupational health department staff that included occupational health nurses and case managers who reported to the corporate Medical Director/Advisor and the Manager of the Occupational Health Department. All followed similar claims management processes. As a result, all had similar data sources. Employees applying for short-term disability benefits were required to provide the occupational health department with medical evidence of disability. In addition, they received a short-term disability benefits application package containing:

- application form requesting basic demographic and occupational data,
- medical information release form, and
- *Attending Physician's Report* to be completed by the employee's physician. It requested information about the specific reason(s) for the employee's disability application, restrictions to his/her abilities and the physician's treatment plan.

For extended absences, an *Attending Physician's Report* form or a supplementary form was sent to the employee approximately every month to be completed and returned to the occupational health department by his/her health care provider. Between reports, a case manager was responsible for maintaining telephone contact with the employee and his/her supervisor to monitor and record the employee's progress. Paper records of these communications were stored in the employee's file in the occupational health department.

All the occupational health departments also maintained a database containing electronic records for each employee collecting short-term disability benefits. These records contained basic information including episode initiation date, diagnosis, treatment plan and status at different points throughout the leave. Case managers used these records to monitor the employee's progress.

Data Abstraction from Occupational Health Records

A list of eligible disability claims from company electronic databases formed the basis for identifying the occupational health records from which information was abstracted. Several procedures were instituted to protect employee confidentiality. Six trained nurse-abstractors were bound by a signed confidentiality agreement and worked in collaboration with the designated staff at each company's occupational health department. A predefined set of data elements were extracted from each subject's company occupational health record for the disability episode using a customized, secure, computerized data entry form. Personal identifiers (e.g. names, addresses or identification numbers) were not recorded during the abstraction process.

Study Population

Subjects included in our analysis met two criteria. First, they met the company-defined criteria for collecting short-term disability benefits. This meant the subjects in our study had depression-related absences from work for at least 10 consecutive work days prior to their disability leave (starting sample, $n = 1{,}521$).

Because the prescription benefits were provided using pay direct plans, the burden of filing a drug claim was taken from the employee. Instead, at the time of a prescription fill employees presented a type of credit card to be swiped by the pharmacist and an invoice was sent directly to the insurer. By decreasing the inconvenience associated with filing a claim, the probability of the benefit being used was increased. In addition, the drug benefit formularies were very generous and covered most available prescription drugs.

This led to the second criterion. Subjects had to have used their prescription drug benefit at least once during the study period for any type of prescription. This assumed that most people used their drug benefit for some type of ailment including minor complaints such as colds. If an individual did not fill a prescription (e.g. for at least one antibiotic) we assumed s/he probably did not use his/her drug benefit. This excluded sixty subjects.

Dependent Variables

Two outcome variables were examined. The first was a dummy variable to identify whether the subject filled an antidepressant prescription at any time during his/her short-term disability episode.

The second outcome variable was a dummy variable to indicate whether the subject returned to work at the end of the disability episode. After a total of six months on short-term disability, employees in all three participating companies were required to either return-to-work, transition to long-term disability or terminate their employment. Employees who either claimed long-term disability benefits or terminated their employment were categorized as not having returned to work.

Independent Variables

Four categories of independent variables were used: (1) demographic (i.e. age and sex); (2) work experience (i.e. length of employment and occupational status); (3) depression symptom and complexity; and (4) indicators of recommended clinical guideline use of antidepressants.

Demographic Variables

Two demographic variables were created. One indicated whether the subject was female or male. The second variable was age calculated as the difference between birth date and the date of the start of the short-term disability episode.

Work Experience Variables

Two work experience variables were created. The first variable indicated whether the subject was in a management position (i.e. supervisor or manager). The second variable contained the length of time the subject was with the company. It was calculated as the difference between the subject's hire date and date of the start of the short-term disability episode.

Depression Symptom and Complexity Indicators

To reflect the number of symptoms reported by the subjects, we created a count of the number of depression-related disability symptoms reported on the short-term disability application form. Information was abstracted from occupational health records using a checklist covering the major DSM-IV depressive symptoms categories (APA, 1994).

An examination of the reported depression-related symptoms revealed an emphasis on the more salient symptoms rather than providing a thorough clinical description. However, those with a depression-related nervous/mental disorder-related episode had significantly higher average numbers of depression symptoms compared to those whose episode was not identified as depression-related. This suggested that though average number of depression-related symptoms could not be used to develop a DSM-IV diagnosis, it is a crude indicator of the relative depression-related severity.

We also created a variable to indicate whether the short-term disability was attributed to depression only or depression comorbid with either another mental or physical problem. Finally, assuming greater severity was associated with a recurring episode, we developed a variable to indicate whether a worker had a prior short-term disability episode during the past 12-months as another proxy for severity.

In addition, in previous work we observed that despite concordance with guideline recommended first-line agents and use within recommended timeframes, there are a group of users who experience a complex course of antidepressant use. These complex patterns have been reported by Claxton et al. (1999) and Thompson et al. (1996). There is evidence that this complexity is associated with a greater need for high intensity health services. This in turn, may be linked to the severity of the episode. For example, Thompson et al. (1996) observed that those who switched and augmented their antidepressant use had more inpatient hospital use. These findings were corroborated by Dobrez et al. (2000) who found these groups of users also use more healthcare services overall. Dewa et al. (2003) also observed certain patterns suggesting a greater severity of illness and resistance to treatment. For example, those who switched and those who augmented use on average reported a higher number of symptoms than those who either had one antidepressant fill or used one antidepressant exclusively. This suggests that the former two groups may have a relatively greater severity leading to a more difficult course of treatment.

Based on the literature and previous findings (Dewa et al., 2003), we created four pattern variables to capture the complexity of antidepressant use. The indicators were:

(1) *One fill only* indicates the subject had only one prescription fill for antidepressants during the short-term disability episode.
(2) *One exclusively* indicates the subject filled >1 prescription for an antidepressant and did not change antidepressants during the short-term disability episode.
(3) *Switched* indicates that >1 prescription was filled and antidepressants were changed at least once during the short-term disability episode.
(4) *Augmented* indicates that >1 prescription was filled and two prescriptions for different antidepressants were filled on same day during the short-term disability episode.

Defining Recommended Antidepressant Treatment

Recommended antidepressant treatment was based on the Canadian Network for Mood and Anxiety Treatment (CANMAT) guidelines (1999). CANMAT is a national network of Canadian healthcare professionals in research, academic and clinical centers that seeks to improve the treatment of individuals with mood and anxiety disorders. Their guidelines are written for physicians practicing in general medical settings.

Based on patterns of drug use during the 200 days following the initiation of the short-term disability episode, we developed two variables to characterize different aspects of drug utilization. These included:

(1) *Use of recommended first-line antidepressant* indicates whether one of the CANMAT first choice antidepressants was the first drug used during the short-term disability episode. These include the antidepressants – fluoxetine, fluvoxamine, paroxetine, sertraline, bupropion, moclobemide, nefazedone, or venlafaxine.

(2) *Antidepressant was received within 30-days of the initiation of short-term disability*. This indicator variable captures whether the antidepressant prescription was filled either within the 30-day period prior to or following the start of the short-term disability episode.

Company Fixed Effects

Because there may exist non-random company-specific factors associated with either antidepressant use or return to work, company-specific fixed effects were included in both models. Under ideal conditions, these non-random factors would be controlled for by the inclusion of variables that are correlated with outcomes and vary between companies. However, given the limitations inherent in the data, it was not possible to explicitly measure company factors and their contribution to the outcomes. As a result, company fixed effects allowed us to adjust our estimates for unobserved company-related heterogeneity.

Analysis

Differences in demographic, work experience, depression severity and complexity and pharmacologic treatment experience characteristics of men and women in this study population were examined. The chi-square test was employed to examine the statistical significance of the difference between discrete variables and sex. Two-sided *t*-tests were used to test the associations between continuous variables and sex.

The regression analysis plan was selected to help describe factors associated with decisions to: (1) Fill an antidepressant prescription; and (2) Return to work. To allow for the potential that important characteristics might differ for men vis-à-vis women, the regression analyses were stratified by sex. The two main research questions are both binary and patient driven. With regard to the second point, it is assumed that subjects will only fill an antidepressant prescription or return to work if they believe this will maximize their utility. The utility of each decision may be influenced by a variety of factors; however, the subjects' choices reveal their preferences.

For this reason, a random utility model was employed (Greene, 1993). For (1), let

$$U^{\text{FILL}} = \text{the utility of filling an antidepressant prescription and}$$
$$U^{\text{noFILL}} = \text{the utility of not filling an antidepressant prescription.}$$

Similarly, for (2), let

$$U^{\text{WORK}} = \text{the utility of returning to work and}$$
$$U^{\text{noWORK}} = \text{the utility of not returning to work.}$$

The observed decisions reveal which options provide greater utility; the actual utilities themselves, however, are unobserved (Greene, 1993). That is, when

$$\Delta U^{\text{FILL}} = U^{\text{FILL}} - U^{\text{noFILL}} > 0,$$

it is expected that a subject will fill an antidepressant prescription. Likewise, for (2) when

$$\Delta U^{\text{WORK}} = U^{\text{WORK}} - U^{\text{noWORK}} > 0,$$

it is expected that a subject will return to work. As per the random utility model, the differences in utility are specified as unobserved variables such that

$$\Delta U^{\text{FILL}} = \boldsymbol{\alpha}'\boldsymbol{x}_{\text{Fill}} + \delta,$$
$$\Delta U^{\text{WORK}} = \boldsymbol{\beta}'\boldsymbol{x}_{\text{Work}} + \varepsilon, \text{ and } \delta \text{ and } \varepsilon \text{ have standard normal distributions.}$$

The $\boldsymbol{x}$ matrices contains data from the administrative database (e.g. the four categories of variables: (1) demographic; (2) work experience; (3) depression symptom and complexity; and (4) indicators for recommended clinical guideline use of antidepressants), and the coefficient vectors are $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$.

The following data are observed

$$\text{FILL} = 1 \text{ (a filled antidepressant prescription) when } \Delta U^{\text{FILL}} > 0 \text{ and}$$
$$\text{FILL} = 0 \text{ (an unfilled antidepressant prescription) when } \Delta U^{\text{FILL}} < 0.$$

Also it is possible to observe

$$\text{WORK} = 1\ \text{(returned to work) when}\ \Delta U^{\text{WORK}} > 0\ \text{and}$$
$$\text{WORK} = 0\ \text{(did not return to work) when}\ \Delta U^{\text{WORK}} < 0.$$

Therefore,

$$\begin{aligned} P(\text{FILL} = 1) &= P(\Delta U^{\text{FILL}} > 0) = P(\boldsymbol{\alpha}'\boldsymbol{x}_{\text{Fill}} + \delta > 0) = P(\delta < \boldsymbol{\alpha}'\boldsymbol{x}_{\text{Fill}}) \\ &= F(\boldsymbol{\alpha}'\boldsymbol{x}_{\text{Fill}}). \end{aligned}$$

Similarly,

$$\begin{aligned} P(\text{WORK} = 1) &= P(\Delta U^{\text{WORK}} > 0) = P(\boldsymbol{\beta}'\boldsymbol{x}_{\text{Work}} + \varepsilon > 0) = P(\varepsilon < \boldsymbol{\beta}'\boldsymbol{x}_{\text{Work}}) \\ &= F(\boldsymbol{\beta}'\boldsymbol{x}_{\text{Work}}). \end{aligned}$$

Given the assumed standard normal distributions of δ and ε, probit regression was used to produce estimates of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$.

It should also be noted that the decision to fill a prescription may be related to the decision to return to work (e.g. an employee more likely to fill a prescription might also be more likely to return to work). In this case, a more complex modeling strategy is needed to adjust for the correlation between the two decisions. This possibility was explored using a seemingly unrelated bivariate probit. The results did not provide a compelling reason to believe the decisions were highly correlated. In fact, we were unable to reject the null hypothesis that the correlation between δ and ε was 0 ($p < 0.44$). Additional results from this auxiliary analysis are available from the authors.

RESULTS

The descriptive statistics (Table 1) indicated that with regard to age and management status, there was a significant difference in the characteristics of men and women who had depression-related short-term disability claims. Compared to women, there were significantly larger proportions of men who were older (56+ years) ($\chi^2 = 19.11$, df $= 1$, $p < 0.0001$) and in management positions ($\chi^2 = 11.81$, df $= 1$, $p < 0.001$).

At the same time, there were no significant differences either in the number of depressive symptoms reported or comorbidities between the two groups. However, on average men remained on short-term disability significantly longer than women (t-statistic $= -2.18$, df $= 228.57$, $p < 0.030$). They were also less likely to return to work at the end of their short-term disability episode ($\chi^2 = 5.56$, df $= 1$,

$p < 0.019$). At the same time, men were also significantly more likely to use antidepressants ($\chi^2 = 3.96$, df $= 1$, $p < 0.047$). But, once antidepressants were used, there were no differences in the patterns of use between men and women (Table 2).

Factors Associated with Antidepressant Use

The results from the probit models indicated there were both similarities and differences between men and women with regard to the factors influencing antidepressant use (Table 3). Symptoms played a significant and positive role for both men ($\mathrm{d}F/\mathrm{d}x = 0.043$, 95% CI 0.0094–0.076) and women ($\mathrm{d}F/\mathrm{d}x = 0.025$, 95% CI 0.013–0.037). Job duration was also significantly and positively associated with antidepressant use for men ($\mathrm{d}F/\mathrm{d}x = 0.11$, 95% CI 0.000345–0.021) and women ($\mathrm{d}F/\mathrm{d}x = 0.0074$, 95% CI 0.0027–0.012). In both cases, the probability of antidepressant use grew with the length of tenure with the company.

In contrast, experience with a prior short-term disability episode ($\mathrm{d}F/\mathrm{d}x = 0.16$, 95% CI 0.071–0.24) and depression only ($\mathrm{d}F/\mathrm{d}x = 0.12$, 95% CI 0.063–0.18) were also significantly associated with women's use of antidepressants. These findings suggest that the probability of antidepressant use increased either when a woman had a prior short-term disability episode or had depression only.

Factors Associated with Return to Work

The results of the probit models for return to work (Table 4) suggested that the probability of return to work declined as the number of depressive symptoms a claimant reported increased for both men ($\mathrm{d}F/\mathrm{d}x = -0.049$, 95% CI -0.066, -0.0042) and women ($\mathrm{d}F/\mathrm{d}x = -0.036$, 95% CI -0.046, -0.025). A similar pattern was observed in relationship to complexity. If either men or women augmented antidepressant use, their probability of returning to work significantly decreased (for men: $\mathrm{d}F/\mathrm{d}x = -0.64$, 95% CI -0.92, -0.36 and for women: $\mathrm{d}F/\mathrm{d}x = -0.26$, 95% CI -0.44, -0.078).

For men, holding a management position ($\mathrm{d}F/\mathrm{d}x = 0.19$, 95% CI 0.0064, 0.37) was significantly associated with a higher probability of returning to work. In contrast, for women the use of guideline recommended first line antidepressant was related to a significantly greater probability of returning to work ($\mathrm{d}F/\mathrm{d}x = 0.086$, 95% CI -0.016, 0.19) while older age was associated with a lower probability ($\mathrm{d}F/\mathrm{d}x = -0.0036$ 95% CI -0.0074, -0.000096).

Table 1. Study Population Descriptive Characteristics: by Sex.

	Total Sample		By Sex		Statistical Tests of Differences Between Men and Women
	%	*n*	Men	Women	
Total	100	1461	12.6%	87.4%	
Demographic					
Age					
<26 yrs	4.2	61	3.9%	4.3%	$\chi^2 = 0.067$, df $= 1$, $p < 0.80$
26–35 yrs	28.2	406	32.6	27.5	$\chi^2 = 2.018$, df $= 1$, $p < 0.16$
36–55 yrs	63.9	922	54.1	65.3	$\chi^2 = 8.61$, df $= 1$, $p < 0.003$
56+ yrs	3.7	53	9.4	2.9	$\chi^2 = 19.11$, df $= 1$, $p < 0.0001$
Work experience					
Management position	9.5	135	16.5%	8.5%	$\chi^2 = 11.81$, df $= 1$, $p < 0.001$
Average years of employment in company	12.8 yrs (8.7)		11.6 yrs (9.6)	12.9 yrs (8.5)	*t*-test $= 1.71$, df $= 197.46$, $p < 0.088$[a]
Symptom and complexity					
Prior episode within past 12-months	12.3	180	16.3%	11.7%	$\chi^2 = 3.093$, df $= 1$, $p < 0.079$
Depression only	47.3	691	47.3%	47.3%	$\chi^2 = 0.0001$, df $= 1$, $p < 1.00$
With anxiety	24.3	355	26.1	24.04	$\chi^2 = 0.37$, df $= 1$, $p < 0.55$
With stress	9.3	136	8.2	9.5	$\chi^2 = 0.33$, df $= 1$, $p < 0.56$
With adjustment disorder	6.4	94	4.9	6.7	$\chi^2 = 0.83$, df $= 1$, $p < 0.36$
With anxiety & Stress	4.4	64	3.8	4.5	$\chi^2 = 0.17$, df $= 1$, $p < 0.68$
With other mental disorder	4.5	65	7.6	4.0	$\chi^2 = 4.94$, df $= 1$, $p < 0.026$
With physical disorder	3.8	56	2.2	4.1	Fisher's exact $p < 0.302$[b]
Average # of reported symptoms	3.8 symptoms (2.7)		4.0 symptoms (2.7)	3.8 symptoms (2.7)	*t*-test $= -1.016$, df $= 1459$, $p < 0.31$
Characteristics of disability episode					
Average short-term disability days during episode	94.6 days (65.7)		105.2 days (71.3)	93.1 days (64.7)	*t*-test $= -2.18$, df $= 228.57$, $p < 0.030$[a]
Returned to work at end of episode	76.5	1117	69.6%	77.5%	$\chi^2 = 5.56$, df $= 1$, $p < 0.019$

Note: Standard deviations in parentheses.
[a] Based on *t*-test with unequal variances.
[b] Based on Fisher's exact test.

Table 2. Antidepressant Use Patterns Among Antidepressant Users: by Sex.

Variables	Total Sample		By Sex		Statistical Tests of Differences Between Men and Women
	%	n	Men	Women	
Used antidepressants	57.9	846	64.7%	56.9%	$\chi^2 = 3.96$, df $= 1$, $p < 0.047$
Complexity of antidepressant use					
One fill only	13.4	113	14.3%	13.2%	$\chi^2 = 0.10$, df $= 1$, $p < 0.75$
Exclusively used only 1 antidepressant	44.7	378	42.0	45.1	$\chi^2 = 0.40$, df $= 1$, $p < 0.53$
Changed antidepressants	30.3	256	31.9	30.0	$\chi^2 = 0.18$, df $= 1$, $p < 0.67$
Augmented use of 2 antidepressants	11.7	99	11.8	11.7	$\chi^2 = 0.0005$, df $= 1$, $p < 0.98$
Guideline recommended use of antidepressants					
Had antidepressant claim with 30 days of start of short-term disability	70.2	594	65.5%	71.0%	$\chi^2 = 1.44$, df $= 1$, $p < 0.23$
Used a guideline recommended first choice antidepressant	90.9	769	93.3%	90.5%	$\chi^2 = 0.98$, df $= 1$, $p < 0.33$

Table 3. Probit Regression Results for Antidepressant Use Stratified by Sex.

Variables	Overall		Males		Females	
	Marginal Probability Effects (d*F*/d*x*)	95% CI	Marginal Probability Effects (d*F*/d*x*)	95% CI	Marginal Probability Effects (d*F*/d*x*)	95% CI
Demographic variables						
Female	−0.064	−0.15, 0.019				
Age	−0.00041	−0.0045, 0.0037	−0.0042	−0.014, 0.0064	-4.7×10^{-6}	−0.0045, 0.0045
Work experience						
Management	0.040	−0.057, 0.14	0.11	−0.10, 0.32	0.021	−0.088, 0.13
Length of employment	0.0079***	0.0035, 0.012	0.011**	0.00058, 0.021	0.0074**	0.0027, 0.012
Company 1	0.17**	0.036, 0.30	0.20	−0.059, 0.46	0.16*	0.0080, 0.31
Company 2	0.053*	−0.0068, 0.11	0.071	−0.092, 0.23	0.053	−0.012, 0.12
Symptom and complexity variables						
Number of symptoms	0.027***	0.015, 0.038	0.043**	0.0094, 0.076	0.025***	0.013, 0.038
Prior episode in past 12-months	0.14***	0.064, 0.22	0.048	−0.15, 0.25	0.16***	0.071, 0.24
Depression only	0.11***	0.049, 0.16	−0.0072	−0.16, 0.14	0.12***	0.063, 0.18
Pseudo-R^2	0.0563		0.0993		0.053	
N	1188		155		1033	

*$p < 0.10$.
**$p < 0.05$.
***$p < 0.001$.

Table 4. Probit Regression Results for Return to Work Stratified by Sex.

Variables	Overall		Males		Females	
	Marginal Probability Effects (dF/dx)	95% CI	Marginal Probability Effects (dF/dx)	95% CI	Marginal Probability Effects (dF/dx)	95% CI
Demographic variables						
Female	0.050	−0.027, 0.13				
Age	−0.0034*	−0.0069, −0.00013	−0.0029	−0.013, 0.0075	−0.0036*	−0.0074, −0.000096
Work experience						
Management	0.0026	−0.083, 0.088	0.19*	0.0064, 0.37	−0.041	−0.14, 0.057
Length of employment	0.00083	−0.0028, 0.0045	−0.00084	−0.013, 0.011	0.0012	−0.0026, 0.0051
Company 1	−0.010	−0.14, 0.12	−0.055	−0.39, 0.28	−0.0058	−0.15, 0.13
Company 2	−0.0042	−0.057, 0.049	0.041	−0.12, 0.20	−0.018	−0.074, 0.038
Symptom and complexity variables						
Number of symptoms	−0.036***	−0.046, −0.026	−0.035**	−0.067, −0.0042	−0.036***	−0.046, −0.025
Prior episode in past 12-months	−0.081**	−0.16, −0.0016	0.094	−0.074, 0.26	−0.11**	−0.20, −0.021
Depression only	0.013	−0.037, 0.062	0.11	−0.046, 0.27	0.0049	−0.047, 0.057
One fill only	−0.077	−0.23, 0.077	−0.22	−0.68, 0.23	−0.051	−0.21, 0.11
One antidepressant exclusively	−0.12*	−0.25, 0.012	−0.18	−0.61, 0.26	−0.10	−0.24, 0.034
Switched antidepressants	−0.26***	−0.40, −0.11	−0.26	−0.71, 0.19	−0.25***	−0.41, −0.097
Augmented antidepressant use	−0.32***	−0.49, −0.15	−0.64**	−0.92, −0.36	−0.26*	−0.41, −0.096
Guideline recommended use						
First line antidepressant	0.086*	−0.012, 0.18	0.080	−0.25, 0.41	0.086*	−0.016, 0.19
Used within 30 days	0.0038	−0.063, 0.071	0.045	−0.16, 0.25	−0.0076	−0.080, 0.065
Pseudo-R^2	0.0937 ($n = 1188$)		0.1370 ($n = 155$)		0.0970 ($n = 1033$)	

*$p < 0.10$.
**$p < 0.05$.
***$p < 0.001$.

DISCUSSION

Our findings suggest that men affected by disabling depression are a small but potentially significant group. They are more likely to be managers and experienced workers, valuable assets to their companies representing a wealth of accumulated human capital. Yet, they are less likely to return to work at the end of their short-term disability episodes. If they did return to work, their average short-term disability episode was relatively longer. They comprise a group whose replacement will translate into a large friction cost for the employer.

The implications of our probit results suggest different strategies for prevention and treatment for the two groups may be beneficial. This may be particularly important information as companies attempt to plan occupational health programs in these times of fiscal constraints. Rather than create universal programs, they can be targeted to the group for whom they are most effective.

Differences in Antidepressant Use

For example, compared to men a significantly smaller proportion of women used antidepressants. Given the consistent finding in other studies that women generally use health services for mental disorders more often than men (Bland et al., 1997; Katz et al., 1997; Lin et al., 1996; Rhodes et al., 2002), this finding is somewhat counterintuitive. This is especially true since between the groups there were no significant differences in either the average number of reported symptoms or the presence of comorbid mental or physical problems. These two facts would suggest there were no differences in the severity experienced by the two groups. One interpretation of these findings is that antidepressant use is not necessarily related to severity.

Another interpretation of these patterns is that the greater use of antidepressants among men lies in different responses to symptoms. Men may be more responsive to symptoms thereby giving them greater impetus to seek antidepressant treatment and an incentive to continue treatment. For instance, the results from the probit regressions suggest symptoms are significantly associated with use in both groups. But, for men, the addition of one symptom is associated with an increase in the probability of using antidepressants that is almost double that for women. Indeed, Leaf and Bruce (1987) observed that men distressed by symptoms are more likely to use services for mental disorders than women. Interestingly, they observed this was especially true for men with negative attitudes towards treatment.

Another potential explanation might be linked to differences in men and women's responses to treatment alternatives. For instance, relative to women, men may have a greater preference for antidepressants over psychotherapy. This would result in a greater proportion of men using antidepressants – a pattern similar to the one we observed. A test of this hypothesis would be to compare the therapies (e.g. pharmacologic, psychotherapy, or both) chosen by the two groups.

Use of Antidepressants and Disability Outcomes

Though women were less likely to use antidepressants than men, when they did use antidepressants, their patterns of use were associated with significant differences in outcome. The use of guideline recommended first line antidepressants significantly increased their probability of returning to work. These findings suggest that for women, follow-up during the disability episode may be particularly important with a special emphasis on guideline concordant treatment.

Among men, the relationship between guideline use of antidepressants and return to work is not as clear. The estimated associations are positive, but they are small and statistically insignificant at conventional levels. One explanation for this difference may be associated with the timing of treatment. That is, men and women might seek treatment at different points within the acute phase of the depressive episode. While there is no significant difference in reported symptoms, there may be a difference in the length of time during which they are experienced. If men delayed treatment, they could require a longer time to recover regardless of guideline care. Such a pattern would be consistent with our observations. It is also a testable hypothesis that would involve following men and women from the onset of the acute phase of their depressive episodes through to the continuation phase and measuring the point at which they receive treatment and the time to recovery.

Another explanation for our findings could involve interactions among three factors. First, in a previous analysis we observed that during the short-term disability episode, men were significantly more likely to terminate their employment than women (11.9% vs. 7.4%) (Dewa et al., 2001). Second, a greater proportion of men were over 56 years of age. Third, a larger proportion of men were in management positions. Together, these factors could indicate that approaching retirement and being in management positions may influence the decision to retire rather than return to work. This would be especially true if management positions offered a lucrative retirement package. Indeed, other studies have observed that benefits packages influence the decision to return to work (Salkever et al., 2000).

Prior Short-Term Disability Episode and Disability Outcomes

The probit results also indicate that for both groups, the number of symptoms as well as the complexity of antidepressant use were associated with a decreased probability of returning to work. In addition, for women a previous episode of depression-related disability was also an important factor. This corroborates findings reported by Mintz and colleagues (1992) who found that recurrence of depression negatively affected long term work outcomes. These results imply that for women, antidepressant use may not be the sole answer. Rather, prevention of recurrence may be a more important tactic with an emphasis on monitoring and support once they return to work. Indeed, in a review of the literature Hirschfeld (2001) found that sustained treatment with antidepressants was associated with a significant decrease in recurrence.

Potential Significance of Management Status

For men, occupational status was an important element to their return to work; working in a management position increased their probability of returning to work. This might be attributed to a sense of personal control associated with the position (Rosenfield, 1989). This perception of power could be an important draw to return to work. Alternatively, perhaps the circumstances that allowed for them to assume management positions involved protective factors.

Conversely, for women, though not significant, the probability of returning to work decreased if they are in management positions. Based on Rosenfeld's (1989) theory, we would posit that the management position places an additional demand (along side those incurred at home) on the woman and therefore inhibits return to work. A test of this hypothesis would be to compare women with and without demands (e.g. primary caregiver responsibilities) at home.

Unfortunately, this hypothesis cannot be tested using this dataset. It does not contain information about subjects' other social or economic factors. Yet, these could also be associated with return to work. For example, the additional demands placed on a woman who is a single head of house could increase the difficulty in returning to work. Thus, it would be important for future studies to also examine the impacts of environmental factors on work return.

Limitations and Directions for Further Research

It should be noted that as with most administrative database studies, our results are limited by the accuracy of the diagnosis on the claims forms (Browne et al., 1998). Thus, we examined a population identified with depression rather than those confirmed with depression. In addition, unless otherwise indicated, it was assumed that individuals suffered from what CANMAT guidelines identified as *typical* depression. Yet, there are other types of depressive disorders including dysthymia, adjustment disorder, seasonal affective disorder and bipolar disorder. For some of these other disorders such as bipolar disorder, antidepressants would not be the recommended course of treatment. To the extent that workers suffered from disorders such as bipolar disorder, we would not expect to observe patterns consistent with the guidelines we used.

We also focused on only one aspect of treatment for depression and had limited information on comorbid conditions. In future studies it will be important to understand the roles of other treatment modalities such as psychotherapy and whether they affect men and women differently as well as the impact of co-occurring conditions on return to work. Disability management practices and preventive interventions that consider these possibilities are other areas worth exploring.

Furthermore, our study focused on workers who went on depression-related disability leave. The limitations associated with an observational study design make our results more exploratory than definitive. We cannot comment on the precise mechanism that results in return to work. The administrative data limits the extent to which this can be done.

Finally, the workers in our study came from the financial/insurance sector. The period during which the study took place was one of great unrest for the financial/insurance sector. A number of the large institutions were in merger and acquisition negotiations (Canadian Bankers Association, 1997). This left many employees feeling vulnerable and at risk of losing their jobs. To the extent that this additional stress and insecurity exacerbated the incidence of depression-related disability, our results may not generalizable to other more stable sectors. On the other hand, given the current economic downturn and its rising unemployment rate, the study results may be helpful to the increasing number of occupational health departments who are yet to face similar circumstances of our three study companies.

CONCLUSION

Depression-related short-term disability is a problem that needs to be addressed but for which there are no simple solutions. It is a multi-dimensional issue that we are only beginning to understand. This study lends some clarity to the differences between men and women who go on to short-term disability and identifies factors that are related to the successful outcomes of both groups. The next step is to understand how to use this information to shape effective disability management programs.

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INSURANCE STATUS OF DEPRESSED LOW-INCOME WOMEN

Allison A. Roberts

INTRODUCTION

Women are nearly twice as likely as men to suffer a major depressive episode (Kessler et al., 1994). Risk of onset for single mothers is twice that of married mothers and financial hardship also doubles the risk of becoming depressed (Brown & Moran, 1997). If diagnosed, depression can be effectively treated, typically with pharmacotherapy or psychotherapy or some combination of the two (Goldman et al., 1999; Sirey et al., 1999). But a sizable majority of sufferers remain undiagnosed and untreated (Lennon et al., 2001). Such treatment can be prohibitively expensive to patients who lack health insurance, particularly those with few financial resources. Although most low-income women have a safety net in Medicaid, welfare reform's delinking of Medicaid from welfare cash assistance has left uncovered many who are eligible for the benefits (Garrett & Holahan, 2000).

Access to any type of health care in the U.S. – physical or mental – is largely driven by insurance coverage. Being uninsured presents a major obstacle to obtaining care. In 1996, 22.8% of the uninsured were not satisfied their family could get care when needed and 16.6% of the uninsured stated they had gone without needed care; only 5.3% of the insured population stated such difficulty with access to care (Roberts & Ishaya, 2003). Low-income individuals are more likely to be uninsured, imposing a significant barrier to acquiring care. Focusing specifically on the relationship between insurance status and access to mental

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health care, Katz et al. (1997) tested the hypothesis that expanding insurance coverage, especially to low-income Americans, would improve access to care for people with mental health problems. They did so by comparing mental health care utilization of similar sociodemographic groups in Ontario, Canada and the U.S. They found that Americans with the lowest incomes and high morbidity are much less likely to receive services than their Canadian counterparts (who have universal insurance coverage and no explicit limits for psychiatric care). This supports the notion that the lack of health insurance for the low-income group is reducing the likelihood this most at-risk segment of the population will receive the necessary care to treat their depressive disorder.

This study investigates insurance status for low-income women who are suffering from the symptoms of major depressive disorder. I hypothesize that low-income depressed women are disadvantaged in terms of ability to obtain needed health care – particularly because they are more likely to be uninsured. Concern that many in this group lack health insurance arises for several reasons. First, many low-income women are on welfare or have recently exited the welfare rolls. According to the 1997 National Survey of America's Families, 41% of women who left welfare in 1996 had no health insurance in 1997 (Garrett & Holahan, 2000). This is disconcerting since welfare reform provides for Medicaid coverage for up to one year after leaving welfare. Also, recipients of welfare cash benefits are no longer automatically enrolled in Medicaid; they must now apply for the publicly provided insurance separately and statistics show that many are not taking the extra steps to do so (Families USA, 1999). The extra time and effort needed to enroll in Medicaid may be a particular barrier for depressed women, keeping them from seeking out the assistance for which they qualify.

Second, depression is arguably the most treatable of all diagnosed mental illnesses. Those suffering from depression can continue to be functioning, contributing members of society as long as they have access to care. As indicated above, this access begins with insurance coverage. However, length of job tenure and employment status are negatively affected by major depressive episodes (Marcotte et al., 2000; Simon et al., 2000). Since employment status determines an individual's eligibility for employer-provided health insurance, and jobs that experience more mobility often do not include health insurance benefits, the employment consequences associated with mental illness can negatively affect insurance status as well.

Policymakers have had limited success in helping depressed, low-income women improve their economic circumstances and their overall well-being. Lennon et al. (2001) found that low-income mothers who are depressed may be reluctant to reveal their illness, especially to welfare caseworkers, for the fear of having their children taken away. As the New Chance Demonstration Program discovered, "in addition to potentially interfering with employment opportunities and job retention, mental health problems in general, and depression in particular, may reduce the effectiveness of interventions designed to improve education and employment" (Lennon et al., 2001). The World Health Organization ranks depression as one of the world's most undertreated diseases, and as the leading cause of disability among women (Murray et al., 1994). Because the burden of depression on society is so immense, policymakers must devote considerable attention to ensuring that low-income women – the group with the highest current prevalence of depressive disorder – have access to needed mental health care.

BACKGROUND

Many studies have investigated the prevalence of depression among low-income women, the long-term impact of a mother's depression on her children, the impact of depression on labor supply, treatment provided to the depressed (including comparisons of privately insured to publicly insured patients), and welfare reform's detriment to insuring the low-income population. This research is distinct because no other study has focused specifically on depressed women's insurance status. For the reasons mentioned above, low-income depressed women – women who specifically need access to care through health insurance – may be at significantly higher risk of being uninsured. This paper uses data from the 1999 National Health Interview Survey (NHIS), which provides responses to the Composite International Diagnostic Interview – Short Form (CIDI-SF) questionnaire for a large sample of more than 30,000 adults. This data is critical to this research because it enables identification of probable caseness of major depressive episodes (MDE) within the past 12 months. If a woman is suffering from depression she may have never been diagnosed, particularly if she lacks health insurance. Other data require resorting to either self-rated health status or diagnosed illness, preventing the researcher from obtaining unbiased results.

Prevalence of Depression Among Low-Income and Single Women

Many studies highlight the association between poverty and mental health problems. It has long been established that low-income mothers of young children are at high risk for depression because of their sex, their low socioeconomic status, and their status as parents of young children (Hall et al., 1985). Subsequently, Belle (1990) found that unmarried mothers, unemployment, housing problems, and inadequate income were among the stress factors that most highly correlated with depressive symptoms among adult women. Brown and Moran (1997) estimated that risk of onset of depression among single mothers was double that of the married, and the risk of onset was almost double among women experiencing financial hardship compared to those who were not.

Impact of Mother's Depression on her Children

Studies have also investigated the link between a mother's depression and the functioning of her children over time. For example, Goodman (1992) established that children of depressed mothers are primarily at risk for: (1) the development of psychopathology, depression in particular, as well as other adjustment problems; (2) reduced social and emotional competence; and (3) problems with cognitive-intellectual functioning. Links have also been found regarding suicide ideation and suicide behavior: children of depressed mothers were significantly more likely to report suicide thoughts or behaviors when compared to children of well mothers (Klimes-Dougan et al., 1999). Thus, it is clear that appropriately diagnosing and treating depressed women not only affects their own well being, but also the well being of their children.

Impact of Depression on Labor Supply

Depression has been associated with a decline in productivity in the workplace and increased absenteeism from work (Kessler et al., 1999). Research has been conducted to measure the impact of depression on labor market outcomes. Marcotte et al. (2000) found that for women, depression was associated with an earnings loss of over $6000 annually with substantial negative employment effects as well. Whooley et al. (2002) found that "depressive symptoms were associated with a 60% increased adjusted odds of subsequent unemployment and a 90% increased adjusted odds of decreased family income during 5 years of follow-up." On a more positive note, they also reported several treatment trials have established that alleviating depressive symptoms may lead to improved work performance. Berndt et al. (1998) found that depressive severity of chronically depressed individuals has a negative impact on at-work performance, and although Berndt et al. (2000) found that those with depression have about a 20% shorter job tenure than those without any mental disorder, they did find that employees who are diagnosed and treated with pharmacotherapy reap productivity benefits – finding no evidence of difference in average at-work productivity between treated employees and non-depressed employees. Thus, appropriately treating the depressed can reverse some of the negative labor market outcomes of depression, and therefore improve the socioeconomic status of these workers.

Treating the Depressed

Several studies have investigated the rates of diagnosis and success of treatment protocols for major depression. All agree that depression is commonly underdiagnosed in primary care, and that many of those suffering the symptoms of depression do not seek medical help. Szewczyk and Chennault (1997) report that the impact of depression extends beyond the woman to all those around her: relationships with her children and spouse may become strained and employment may become threatened. They indicate that depressed individuals use health care services at twice the rate of nondepressed patients, and 15% of untreated patients with major depression eventually commit suicide with 90% of all suicides associated with depression. Greenberg et al. (1993) report that the economic burden of depression – in terms of treatment, lost work, and lost life – has been estimated at $44 billion annually.

Insurance Status of the Low-Income Population After Welfare Reform

Since the passage of welfare reform in 1996, attention has been paid to the insurance and employment status of welfare leavers. One of the unintended consequences of welfare reform is that many people lose Medicaid coverage and become uninsured; in 1997, an estimated 675,000 low-income people became uninsured as a result of welfare reform (Families USA, 1999). Garrett and Holahan (2000) investigated the health insurance coverage of former welfare recipients. They found that despite federal guarantees of continued Medicaid coverage to welfare leavers, by the time women have been off welfare for a year or more less than one-quarter are receiving Medicaid benefits and about one-half are without insurance. Medicaid is not functioning as a far-reaching safety net in providing health insurance to the poor. U.S. Census Bureau reports indicate that despite the Medicaid program, 44.2% of the 18–64 year olds who are poor (amounting to 7.5 million people) had no health insurance of any kind in 1999 (Mills, 2000).

Summary

This paper adds to the existing literature by investigating the health insurance status of low-income women who have likely suffered a major depressive episode in the past 12 months. The relative risk ratios that these women will be uninsured vs. hold public or private health insurance will be estimated using multinomial logit analysis to control for the other sociodemographic variables that likely contribute to their insurance status. Because these women need access to care – for their employability and potential to earn income, for their childrens' sakes, and for their own emotional well-being – ensuring that they are covered with a comprehensive health insurance plan is critical.

DATA

This study uses data from the 1999 NHIS, which is made publicly available by the National Center for Health Statistics (NCHS). When weighted, the data represent the U.S. civilian, non-institutionalized population. The NHIS provides comprehensive information on respondents' insurance coverage, employment and poverty status, as well as desired sociodemographic characteristics. The 1999 survey year is particularly useful because a sample of over 30,000 adult respondents were asked an additional set of questions which correspond to the 12 month DSM-IV version of the CIDI-SF. As developed and used by Kessler et al. (1994), the responses to these questions can be aggregated to distinguish between probable and non-probable cases of several psychiatric disorders, including MDE.

The CIDI-SF diagnostic screen for MDE has a stem-branch structure, where two initial (stem) questions are asked to determine if the "master symptoms" of MDE are present (Kessler et al., 2000). These two stem questions are: (1) "During the past 12 months, was there ever a time when you felt sad, blue, or depressed for two weeks or more in a row?"; and (2) "During the past 12 months, was there ever a time lasting two weeks or more when you lost interest in most things like hobbies, work, or activities that usually give you pleasure?" If neither stem question is endorsed, the individual is automatically classified as not having MDE and is not asked additional questions about the syndrome. However, when either stem question is endorsed, additional (branch) questions are asked as a follow-up to determine likelihood of a clinically significant case. If three or more of the branch questions are affirmed, respondents are classified as a probable case (Walters et al., 2002). Using this approach the adult respondents in the 1999 NHIS supplement were classified as suffering from depression or not, with the understanding that such classification is due to probable caseness, not clinical diagnosis.

Because this study focuses on the insurance status of adult women, a subset of the NHIS data is chosen such that only females between the ages of 18 and 64 who provided an insurance status and were asked the supplemental diagnostic

Table 1. Descriptive Statistics by Women's Insurance Status – 1999 (Weighted).

Variable Name	Uninsured		Public Only		Some Private	
	Mean	S.D.	Mean	S.D.	Mean	S.D.
Age	35.83	12.5	37.36	13.8	40.14	12.2
Married	0.483	0.50	0.355	0.48	0.680	0.47
Nonwhite	0.276	0.45	0.419	0.49	0.164	0.37
Hispanic	0.250	0.43	0.192	0.39	0.075	0.26
Non-MSA	0.258	0.44	0.248	0.43	0.191	0.39
Yrs School	12.80	3.5	12.03	3.3	15.13	2.8
Employed	0.570	0.50	0.253	0.43	0.759	0.43
Numkids	1.43	1.6	1.48	1.6	1.00	1.3
GoodHlth[a]	0.890	0.31	0.647	0.48	0.942	0.23
Low_inc	0.456	0.50	0.709	0.45	0.093	0.29
Depressd	0.126	0.33	0.215	0.41	0.089	0.28
Low_inc&Depr	0.067	0.25	0.166	0.37	0.011	0.10
Sample size	2331		1338		9589	
Weighted *N*	12,480,203		5,823,416		63,590,803	

Source: Author's calculations from the 1999 NHIS.
[a] GoodHlth = 1 if self-reported physical health status is excellent, very good, or good; 0 if fair or poor.

questions for MDE are included. This leaves a sample dataset with 13,258 total observations.[1] Table 1 provides weighted descriptive statistics for each of the variables used in the multinomial logit analysis, by insurance status. Over 15% of women have no health insurance at all while about 7% are relying solely on publicly provided coverage. Control variables include age, marital status, race, ethnicity, urban vs. rural location, years of education, employment status, number of children in the family, and a control for self-perceived physical health status. Approximately 84.8% of the women have some type of insurance coverage, while 19.1% are considered low-income (defined here as having family income below $20,000 per year),[2] and 10.34% are classified as probable cases of MDE.

While this NHIS estimate for 12-month prevalence of depression using the CIDI-Short Form is less than the National Comorbidity Survey (NCS) female estimate of 12.9%, the NCS 12-month prevalence rate was obtained after administering structured psychiatric interviews (the University of Michigan Composite International Diagnostic Interview, or UM-CIDI). Also, the NCS results are reported for women aged 15–54 only and show "that 12-month disorders are consistently most prevalent in the youngest cohort (age range, 15 to 24 years) and generally decline monotonically with age" (Kessler et al., 1994). Thus, 12-month prevalence rates for the 18–64 age range *should be lower* than 12.9%, so that use of the CIDI-Short Form with the NHIS respondents appears to provide a reasonable

Table 2. Characteristics of Women of Working Age, by Depression and Income Status – 1999 (Weighted).

Characteristic	Healthy, Non-Poor Women	Depressed Women	Low-Income Women	Low-Income and Depressed Women
sample size	8335	1465	3479	564
Uninsured	9.9%	18.6%	36.2%	33.6%
Public insurance only	2.3%	14.8%	26.0%	38.7%
Some private insurance	87.8%	66.6%	37.8%	27.7%
Receiving cash welfare	0.7%	6.5%	11.4%	17.8%
Married	69.3%	51.9%	38.0%	31.7%
Has children	50.6%	46.6%	48.9%	43.4%
Low-income	–	28.8%	–	–
Depressed	–	–	15.7%	–
Nonwhite	16.6%	19.7%	33.2%	32.8%
Hispanic	9.4%	8.3%	18.9%	12.5%
Good physical health	94.9%	76.8%	80.9%	61.1%
High School grad only	26.9%	26.5%	27.6%	27.4%
College grad or higher	28.5%	16.9%	8.2%	6.7%
Employed	75.2%	61.5%	50.3%	43.6%
Average age	39.66 yrs	39.31 yrs	36.93 yrs	38.46 yrs
Avg. number of children	1.12	1.03	1.08	0.96
Avg. years of schooling	15.03 yrs	14.2 yrs	12.76 yrs	12.89 yrs
If employed:				
Paid by the hour	57.4%	68.0%	81.6%	78.2%
Health ins. Offered at work	72.9%	66.1%	45.8%	40.2%

Source: Author's calculations from the 1999 NHIS.

account of cases of major depressive episodes in the female population aged 18–64.

It is also worth investigating the differences in characteristics for women who fit the criteria for: (1) depressed; (2) low-income; and (3) low-income and depressed, vs. those who do not. Table 2 provides a variety of comparative, unadjusted statistics on these groups of women.

The statistics in Table 2 confirm the health insurance status of depressed and low-income women differs significantly from healthy, non-poor women. Comparing depressed women with healthy, non-poor women, the depressed are nearly twice as likely to be uninsured, seven times more likely to hold only public health insurance, and substantially less likely to have any private health insurance. In addition to health insurance status, depressed women differ greatly in terms of welfare recipiency, marital status, educational attainment, physical health status, and employment status. When employed, depressed women are more likely to be paid by the hour and less likely to be offered health insurance through their employer.

Comparing low-income women to the healthy, non-poor women note the striking disparity in insurance status; nearly four times more women in families with annual incomes below $20,000 are uninsured and there are more than ten times as many holding only public insurance. Less than 38% of women in families with incomes below $20,000 are covered by private insurance, compared to nearly 88% of healthy, non-poor women. These statistics provide a clear link between socioeconomic status and insurance status among working age women.

Some additional comparisons across these two groups are enlightening. Because we are using family income we would expect a smaller proportion of women in low-income families to have a spouse, in fact only about half as many are married. But despite their differences in marital status, the proportion of women who are mothers is nearly the same in the low-income group, further exacerbating their financial challenges. The low-income women are also twice as likely to be nonwhite or Hispanic, are more likely to be in poor health, and substantially less likely to have a college degree or be employed. If employed, they are considerably more likely to be paid by the hour and acutely less likely to have been offered health insurance through their employer.

Finally, from Table 2 the characteristics of women who are doubly disadvantaged, low income and depressed, can be compared to the healthy and non-poor. Note the exaggerated tendency for the disadvantaged women to hold public health insurance. Nearly 39% of these women are relying on public aid – the greatest of any of the subgroups investigated. At the same time, these women are substantially more likely to be uninsured and have the lowest rate of private insurance. In addition to insurance status, these women also greatly differ in terms of welfare recipiency, marital status, race, ethnicity, physical health, educational attainment, and employment status. Note that if employed, the group of *low-income and depressed* is the least likely to have been offered health insurance through their employer of any of the subgroups investigated. These simple, unadjusted statistics highlight the necessity to further investigate the health insurance status of low-income, depressed women.

EMPIRICAL ANALYSIS

The empirical analysis in this study determines the extent to which the included factors predict the probability that an adult female will have a particular health insurance status. First, the econometric framework of the multinomial logit (MNL) model used to predict the categorical insurance status is outlined. Then, the results of the model are presented and summarized.

Econometric Framework[3]

Assume that each woman *selects* one of three mutually exclusive alternatives: (i) uninsured (indexed un); (ii) public insurance only (indexed pu); and (iii) some private insurance (indexed pr). Let X_i denote the vector of values of the individual woman's exogenous characteristics explaining her insurance status and β_j the unknown parameter vector. Because there are three possible insurance statuses in the choice set, the probability function of woman i choosing the j^{th} insurance status can be written as:

$$P_{ij} = \frac{\exp(\beta_j' X_i)}{\exp(\beta_{\text{un}}' X_i) + \exp(\beta_{\text{pu}}' X_i) + \exp(\beta_{\text{pr}}' X_i)}.$$

This is the multinomial logit model. The estimated equations provide a set of probabilities for the three insurance statuses for a woman with characteristics X_i. As is, this model suffers from indeterminacy because the number of parameters to be estimated exceeds the number of exogenous characteristics. To remove this indeterminacy and identify the model, it is common to normalize the parameters by arbitrarily setting one vector of parameter values to zero. For the purposes of this study, the parameter values for the insurance status *private* are normalized ($\beta_{\text{pr}} = 0$), so that having private insurance is the base (comparative) status.

Such normalization is convenient as the probability equations become:

$$P(\text{private}) = \frac{1}{1 + \exp(\beta_{\text{un}}' X_i) + \exp(\beta_{\text{pu}}' X_i)},$$

$$P(\text{uninsured}) = \frac{\exp(\beta_{\text{un}}' X_i)}{1 + \exp(\beta_{\text{un}}' X_i) + \exp(\beta_{\text{pu}}' X_i)}, \text{ and}$$

$$P(\text{public}) = \frac{\exp(\beta_{\text{pu}}' X_i)}{1 + \exp(\beta_{\text{un}}' X_i) + \exp(\beta_{\text{pu}}' X_i)}.$$

Then, the *relative* probability of being uninsured to being privately insured is simply:

$$\frac{P(\text{uninsured})}{P(\text{private})} = \exp(\beta_{\text{un}}' X_i),$$

and the *relative* probability of being publicly insured to being privately insured is simply:

$$\frac{P(\text{public})}{P(\text{private})} = \exp(\beta'_{\text{pu}} X_i).$$

Thus, there are explicit formulae for the desired probabilities; the exponentiated value of a coefficient is the *relative risk ratio* for a one-unit change in the corresponding exogenous variable, where that relative risk is measured as the probability of one insurance status (*uninsured, public*) relative to the base insurance status, *private* (StataCorp, 1999).

Empirical Results

The control variables included in the empirical analysis were chosen to capture the characteristics that may influence a woman's desire and/or ability to obtain health insurance coverage. Included are age, marital status, number of children, race, ethnicity, years of schooling, self-reported physical health status, and employment status. The MNL analysis enables determination of the separate effect of each exogenous factor on insurance status, while holding other factors constant (ceteris paribus). In addition to investigating the impact of these controlling factors, the remaining discussion focuses on the probability that low-income, depressed women will be uninsured.

Table 3 provides the weighted maximum likelihood MNL estimation results. The computed parameter coefficients are obtained by assuming that $\beta_{\text{pr}} = 0$. Each of the signs on the control variables are as expected in both the *uninsured* and *public* equations. Women who are between the ages of 18 and 64 who are older, married, more educated, employed, and in good health are *more* likely to hold *private* insurance coverage. Women who are nonwhite, Hispanic, living in non-urban areas, and who have more children are *less* likely, ceteris paribus, to have *private* health insurance. These results are consistent with published statistics on the insured population in the U.S.

To focus on the variables of interest to this study (Low_inc and Depressd), Table 3 shows that adult females in a low-income family and those who have likely suffered a major depressive episode in the past 12 months are significantly *less* likely, ceteris paribus, to have *private* health insurance; they are significantly *more* likely to have *public* insurance coverage and (most problematic) to be *uninsured*. This is worrisome considering these results hold *even after* controlling for age, race, educational attainment, employment, marital status, etc.

Table 3. Maximum Likelihood Estimates of MNL Model of Women's Insurance Status – 1999, Weighted.

Variable Name	β_{un}		β_{pu}
Age	−0.0178		−0.0194
Married	−0.6830		−1.2553
Nonwhite	0.2570		0.6381
Hispanic	0.8508		0.4214
Non-MSA	0.4136		0.2725
YrsSchl	−0.1499		−0.1675
Employed	−0.5938		−1.7390
NumKids	0.1520		0.2595
GoodHlth	−0.2732		−1.4960
Low_inc	1.5880		2.3461
Depressd	0.2864		0.5230
_constant	1.3282		1.8754
Pseudo R^2		0.2740	
Log likelihood		−6048.93	
No. of observations		12,525	

Note: All coefficients are significantly different from zero at $\alpha = 0.05$ or below.

To better comprehend the *magnitude* of the differences in probabilities, the relative risk ratios (RRR) of the uninsured and public insurance statuses – compared to private insurance – are presented in Table 4. The highest risk factors for women being *uninsured* are, in order: (1) Low_inc; (2) Hispanic; (3) living in a non-urban area; and (4) depression. Low-income working age women are 10 times as likely

Table 4. Relative Risk Ratios from Maximum Likelihood Estimates of MNL Model of Women's Insurance Status – 1999, Weighted.

Variable Name	RRR_{un}	RRR_{pu}
Age	0.982	0.981
Married	0.505	0.285
Nonwhite	1.293	1.893
Hispanic	2.341	1.524
Non-MSA	1.512	1.313
YrsSchl	0.861	0.846
Employed	0.552	0.176
NumKids	1.164	1.296
GoodHlth	0.761	0.224
Low_inc	4.894	10.445
Depressd	1.332	1.687

Note: The base insurance status is private.

to have public insurance (relative to private), and are nearly 4.9 times more likely to be uninsured than hold private insurance, ceteris paribus.

A depressed woman is nearly 1.7 times more likely to be publicly insured and over 1.3 times more likely to be uninsured than have private insurance. Thus, if a woman is classified as having suffered a MDE in the past 12 months, she is *less* likely to have *any* health insurance, even after controlling for poverty status and employment status, when compared to women who are not suffering from depression. Perhaps this is due to a lack of interest in taking care of her health or an inability to pursue (or a sense of powerlessness that prevents her from pursuing) insurance options that might be available, like Medicaid. Thus, the MNL results confirm that after controlling for a host of sociodemographic characteristics, adult women in low income families and depressed women are significantly more likely to be uninsured.

The remaining investigation considers women who are doubly disadvantaged – those living in families with low annual incomes AND suffering from depression. When weighted, the data indicate that 3.1% of the women in the sample fall into these two categories simultaneously. This analysis requires re-estimating the MNL model, including an interactive variable to capture the additional effect (if any) of suffering both burdens at the same time. The parameter estimates are provided in Table 5 and the relative risk ratios are given in Table 6.

Note than no parameter estimates experience a change in sign or magnitude when including the interactive variable. Only the *Depressd* variable changes in significance (from $p = 0.005$ without to $p = 0.047$ with the interactive variable). The interactive variable is positive but statistically insignificant in both the *uninsured* and *public* equations, indicating that being *both* low-income and depressed does not appear to have a significant additional negative impact on insurance status. Comparing the relative risk ratios for each variable in Table 6 to those in Table 4, note that inclusion of the interactive variable does not have an impact on the relative probability of being uninsured or publicly insured for those who are low-income or depressed. Thus, the separate effects of low-income and depression on working age women's insurance status are purely additive.

The insignificance of the interactive variable is somewhat surprising. However, it indicates that there may be one piece of good news stemming from these results. It appears that those who are doubly disadvantaged don't incur an additional insurance status penalty beyond just the simple effects of being low-income and depressed. This is most likely driven by the propensity of these needy women to receive attention from someone, be it a caseworker or a family member, to ensure that available coverage for which they are eligible is applied for and obtained. The descriptive statistics provided in the last two columns of Table 2 reveal that those who are low-income *and* depressed are slightly less likely to be uninsured

Table 5. Maximum Likelihood Estimates of MNL Model when Including Interactive Variable, Weighted.

Variable Name	β_{un}	β_{pu}
Age	−0.0178	−0.0194
Married	−0.6829	−1.2555
Nonwhite	0.2572	0.6375
Hispanic	0.8505	0.4218
Non-MSA	0.4138	0.2735
YrsSchl	−0.1500	−0.1676
Employed	−0.5939	−1.7401
NumKids	0.1520	0.2591
GoodHlth	−0.2736	−1.4974
Low_inc	1.5797	2.3173
Depressd	0.2713	0.4294
Low_inc&Depr	0.0657	0.1739
_constant	1.3311	1.8945
Pseudo R^2	0.2741	
Log likelihood	−6048.67	
No. of observations	12,525	

Note: All coefficients are significantly different from zero at $\alpha = 0.05$ or below *except* for the interactive variable of interest.

Table 6. Relative Risk Ratios when Including Interactive Variable, Weighted.

Variable Name	RRR_{un}	RRR_{pu}
Age	0.9824	0.9807
Married	0.5052	0.2849
Nonwhite	1.2933	1.8917
Hispanic	2.3409	1.5247
Non-MSA	1.5125	1.3146
YrsSchl	0.8607	0.8457
Employed	0.5522	0.1755
NumKids	1.1641	1.2958
GoodHlth	0.7607	0.2237
Low_inc	4.8533	10.1482
Depressd	1.3117	1.5363
Low_inc&Depr[a]	1.6078	1.1900

Note: The base insurance status is private.
[a] Low_inc&Depr is statistically insignificant.

and significantly more likely to hold public insurance than the low-income who are not depressed. This points to the likelihood that those who suffer both burdens simultaneously are receiving some assistance to obtain publicly provided insurance.

To test this notion, I computed another MNL model that included an interactive variable – this time for poor physical health AND low-income. The results confirm the hypothesis that especially needy groups are enrolling in Medicaid; this interactive variable is also insignificant and leaves the other parameters unchanged in terms of magnitude and significance.

CONCLUSION

With health care costs escalating, Americans are facing an ever-tougher challenge in accessing and affording needed care. Millions of uninsured individuals in low-income families are eligible for public health insurance, but are not enrolled. Access to health care is limited for those who lack health insurance coverage. Policymakers must ensure that some type of plan covers the poor – those least likely to be able to pay for health care out-of-pocket. Adult women in low-income families who are also suffering the symptoms of depression are doubly disadvantaged and are less likely to be covered by any health insurance. Future research in this area should investigate the means by which insurance coverage could be extended to the more than one-third of low-income, depressed adult women who are currently uninsured. In addition, the quality of mental health care received under publicly provided insurance needs further investigation, as nearly 40% of adult women who are suffering from depression and living in low-income families are relying solely on Medicaid or other state insurance plans. By focusing policy efforts toward assisting these women in obtaining quality health insurance, these disadvantaged women will have greater access to the mental health care they need, not only for their own emotional and physical well-being, but also to help them raise well-adjusted children and to be productive contributors to the labor market and society.

NOTES

1. Due to missing values, a total of 12,525 observations are usable for the multinomial logit analysis.
2. Although this low-income designation is not ideal, the variable designating whether annual family income is above or below $20,000 is the only income variable in the NHIS data that does not have a substantial number (595) of observations with missing values.

A cross-tabulation of this variable's values with poverty status (when known) shows that 94% of those designated low-income (vs. not) have incomes below (above) 150% of their poverty threshold. Thus, the above/below $20,000 annual family income variable appears to be a reasonably good proxy for poverty status and does not have the magnitude of missing observations that the income grouping and poverty status variables have (these variables have over 2700 observations with no detail or missing values altogether).

3. This model is drawn from the work of Maddala (1983) and Greene (1993).

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