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Fear of Adverse Mental Health Treatment Experiences: Initial Psychometric Properties of a Brief Self-Report Measure

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AQ: 1

Many are apprehensive about mental health care, which potentially affects engagement in recovery processes as well as health outcomes. This article introduces a tool to assess fear of adverse mental health treatment experiences from the client's perspective. In a sample of 656 adults receiving mental health services at community agencies, this study is an initial exploration into the validity of a scale assessing fears associated with commonly experienced coercive or disorganized interventions. Factor analyses supported the construct validity of the 10-item Fear of Adverse Treatment Experiences Scale. It significantly discriminates based upon service characteristics, gender, history of victimization, and past experiences with coercive or disorganized interventions, with higher levels of fear reported by users of traditional mental health services, former inpatients who had their voluntary admission status changed, males, people with history of childhood abuse, and people with certain forms of criminal justice involvement.



Keywords: [treatment apprehension](#), [mental health](#), [exploratory factor analysis](#), measurement, multivariate methods, [nonparametric methods](#), [psychometrics](#), rating scales, self-reports, statistical computation

AQ: 2

The realized risks and benefits of psychiatric treatment strike a tough balance. Despite growing numbers of people diagnosed with mental disorders, increased public awareness, and increased regard for person-centered care, many people in recovery remain apprehensive about forging an alliance with mental health professionals. Fear of adverse mental health treatment experiences may be a lasting effect of previous experiences in mental health care, or it could be ingrained in preexisting beliefs new service users bring to the treatment setting. Either way, anxieties about mental health care play a role in how a service user engages in the recovery process and potentially affects health outcomes in any service environment. The purpose of this article is to explore psychometric properties of a tool to assess fear of adverse treatment experiences as a means of informing service engagement efforts directly through potential client perspectives.

A wide range of service recipients may harbor fear of excessive or inappropriate mental health care; such fear is influenced by many factors. Historically, no helping profession has been so linked to coercion as psychiatry, and the association persists. A spectrum of treatment pressures common to modern mental health practices is recognized, ranging from informal coercion to legally sanctioned compulsory treatment (Jaeger & Rossler, 2010). Within

institutionalized environments, coercive measures are common and have had adverse consequences (Frueh et al., 2005; Robins, Sauvageot, Cusack, Suffoletta-Maierle, & Frueh, 2005). In a longitudinal study of state hospital inpatients ($N = 109$), 22.9% of participants described some form of coercive care when recalling their first major contact with psychiatric services (Pescosolido, Gardner, & Lubell, 1998), and past experiences in involuntary circumstances may affect willingness to engage in care (Swartz, Swanson, & Hannon, 2003). Though the topic is not well-established in the literature, unresolved trauma related to previous victimization, abuse, or incarceration may heighten sensitivities toward future interventions (Reddy & Spaulding, 2010). The allied mental health professions are gradually recognizing more holistic, "person-centered" methods of service delivery, yet coercive treatment modalities are increasingly represented in mental health care (Monahan et al., 2001).

Adding to the issue of recognized coercion in treatment, fear may be engendered by the characteristic disorganization of mental health systems (Aron et al., 2009). The mental health professions, already slow to implement interventions grounded in emerging evidence, struggle to accommodate growing needs into their current service infrastructure. Comprehensive, compassionate services have not been available, accessible, and affordable, and many—some sources say over half—of those who carry psychiatric diagnoses experience barriers to adequate, well-coordinated care (Sundararaman, 2009). The piecemeal funding of mental health services results in a fragmentation of care, likely to persist in the face of continual policy and administrative changes (Richman, Grossman, & Sloan, 2010); such disorganization overburdens service workers and complicates outcomes. Potential service recipients are wary of seeking help at agencies too overwhelmed to provide proper treatment.

While good care is available to some, inappropriate treatment is a reality for many service users, and disparities continue. Disem-

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powered ethnic or cultural groups are apprehensive of mental health treatment, or are at risk of receiving unsuitable care (Barnes, 2004; Cusack et al., 2007; Kuno & Rothbard, 2002; Sussman, Robins, & Earls, 1987). Gender disparities in differential use of seclusion and restraint persist, with males tending to be secluded or restrained more upon threatening violence in comparison with females, who are more likely to receive such treatment only after committing actual violence (Clark et al., 2005; Kaltiala-Heino, Tuohimäki, Korkeila, & Lehtinen, 2003). Mothers diagnosed with mental disorders have been found to be at greater risk of coercive care due to efforts to protect children (Miller, 1992). Age may also play a role in attitude toward treatment, as younger and older adults have been reported to view it more negatively than middle-aged adults (Leaf, Bruce, Tischler, & Holzer, 1987). Grubaugh, Frueh, Zinzow, Cusack, and Wells (2007), however, examined perceptions of care and safety in an inpatient setting, and found no significant differences associated with race, gender, or age. The relationships between these demographics and fear of treatment are unclear, especially since much of this research on fear of treatment examines a broad range of possible consequences of care (e.g., stigmatization), not simply the risk of receiving services inappropriate for the presenting need.

Unaddressed fear of treatment erodes therapeutic alliances and affects health outcomes for those who would otherwise benefit from added support. Previous research on treatment fears has shown that it delays help-seeking behaviors (Deane & Chamberlain, 1994; Deane & Todd, 1996; Kushner & Sher, 1989, 1991). Underutilization of suitable mental health treatment equates to less advocacy and restricted access to other vital services, including medical care and housing. Furthermore, reduced contact with coordinated services leads to heavier reliance on urgent care, the nature of which often reaffirms negative experiences. Positive outcomes are maximized by facilitating regular contact with providers and services matched to the service user's goals.

In the spirit of honoring recovery as starting "where you're at," early rapport-building and assessment of attitudinal barriers toward treatment are key strategies to enhance therapeutic alliance. Treatment apprehension can be mitigated in several ways—from seeking support from peers, creating advance directives, or having access to a variety of complementary, wellness-promoting resources. All of these foster a sense of empowerment, and can serve as protective measures against feared treatment, but common resource constraints challenge the implementation of these strategies. While quality measures of treatment-related fear have been developed (Kushner & Sher, 1989; Park, Attenweiler, & Rieck, 2012; Pipes, Schwarz, & Crouch, 1985), their length, scope, and target populations pose logistical challenges to practitioners working with already-established service users. For example, Kushner and Sher (1989) introduced the Thoughts About Psychotherapy Survey (TAPS), a modification of an earlier effort by Pipes, Schwarz, and Crouch (1985). Initially examined in nonclinical samples, the TAPS relates to important factors in treatment aversion, such as therapist responsiveness, image concerns, and coercion concerns. Later analyses with a clinical sample revealed a two-factor structure conflating coercive pressures with "fear of change" (Zartaloudi & Madianos, 2010). At 19 items, the TAPS targets nuanced interpersonal and intrapsychic phenomena which may occur in the context of therapeutic relationships, but it does not speak to organizational or systemic issues which may ad-

versely affect service user engagement, and may have limited utility in agencies where one-to-one emotional support is not the only service provided. Park, Attenweiler, and Rieck (2012) developed an inventory of mental health treatment fears; however, it was designed for college students. Given the ongoing reality of problematic mental health treatment environments and limited research focused on the construct of mental health treatment-related fear, this article introduces a more targeted instrument to address the concerns of current mental health service users in multiple settings: the Fear of Adverse Treatment Experiences Scale (FATES).

Method

Sample and Procedures

The following analysis utilizes archived data, collected during a randomized controlled trial on the effectiveness of mental health services from 1996 to 2001. Adults ($n = 787$) seeking mental health services for the first time in the San Francisco Bay area of California were asked to participate. Participants were recruited from two types of service locations where they themselves had recently initiated services: traditionally run community mental health agencies—which focus on clinically oriented interventions such as case management, in- and outpatient treatments, and medication services—and self-help agencies functioning as drop-in centers, staffed and governed by former clients to provide resources, social support, and shelter in an understanding and welcoming environment. Six hundred and 73 adults (86%) consented to participate in the trial. No significant differences were found in the characteristics of study participants and refusal groups (gender, ethnicity, and housing, or in site where enrolled). The sample was well-balanced with regard to age, gender, and housing status. The mean age was 39 years and 54% of the sample identified as male. A third of the sample was homeless, while another third reported stable housing, and the remaining reported inadequate living arrangements. Ethnic group representation was not as well-balanced within the sample, with 54% identifying as Caucasian, 29% African American, 10% Hispanic, and 3% Asian (Segal, Silverman, & Temkin, 2010, 2011).

After giving informed consent, participants ($n = 673$) were randomly assigned: Of 447 new clients at community mental health agencies, 259 were referred to a combination of self-help and community mental health agency services, while 188 continued outpatient treatment at community mental health agencies only. The remaining 226 participants were new self-help service users and continued participation at these agencies. Participants provided information on demographics, history of mental health service use, and recovery-oriented measures of psychosocial functioning. Though randomization was integral to the original study of new service users, the current study on fear of adverse treatment experiences does not depend on such a design and uses archival data from the initial project, without regard for assignment.

Measures

The FATES was developed in preparation for the randomized controlled trial. Scale items were generated using a deductive approach (Hinkin, 1998), drawing from theoretical and empirical

work (Campbell & Schraiber, 1989) as well as the expertise of researchers and leaders of mental health consumer groups in northern California. To further refine the response items, we pilot-tested the scale at local self-help agencies. The resulting product was a 10-item scale of Likert response items, ranging from 1 (*not at all worried about a situation occurring at given agency*) to 5 (*extremely worried about a situation occurring at given agency*). Scale items (see the Appendix) address, for example, worry over staff initiating hospitalization for psychiatric reasons, or staff lacking the training to provide desired services. During the trial, the FATES was administered at the baseline interview only.

Following conclusion of the trial, question items were divided by researchers into two subscales for secondary analysis; these were presented as the Fear of Coerced Care Scale and Fear of Inadequate Care Scale (Segal, Hodges, & Hardiman, 2002). Though the internal consistency of each scale was high (.86 and .90, respectively), the purpose of the work was not to investigate their validity for interpretation in this sample, and no factor analyses were reported. Rather, these subscales were used as part of a larger investigation of factors affecting clients' initial decisions to use the services at the self-help agencies and community mental health agencies—these were the sites at which they were recruited for participation, not where they were assigned to receive services during the trial. In contradiction to the authors' hypotheses on attitudinal barriers, it was found that participants who initially sought treatment at community mental health agencies scored higher on the Fear of Coerced Care Scale than those who sought treatment at self-help agencies. Another surprising result was that participants who initially sought services at self-help agencies tended to score higher on the Fear of Inadequate Care Scale. The authors concluded that clients' realized need for services could serve not only as a significant motivator toward accessing them, but also bring into consciousness the myriad of challenges associated with seeking help in traditional mental health settings (Segal, Hardiman, & Hodges, 2002).

Though the FATES was previously presented in the form of these two scales, this article seeks to establish baseline evidence supporting the use of total scores from the originally developed scale as a stand-alone measure of fear of adverse treatment experiences. Use of the de-identified dataset was granted exemption by the University of California, Berkeley, Committee for Protection of Human Subjects (Protocol 2014–01-5916). The present analysis synthesizes a traditional and nonparametric approach. Following an analysis of missing values, correlation matrices were estimated. Dimension reduction was achieved utilizing principal components analysis and principal axis factoring, for comparison. Internal consistency of the FATES was determined and finally, nonparametric tests of proportion equality were used to compare responses between groups within the sample.

Results

Response

Among respondents, summated scores on the 10-item FATES ranged from 10 to 50 (maximum). Response patterns were positively skewed with a median total score of 17 ($M = 18.35$, $SD = 8.29$). The majority of respondents indicated relatively low fear of adverse mental health treatment experiences, with 27.4% of par-

ticipants included in final analyses ($n = 180$) reporting no fear at all. Fourteen individuals, however, endorsed a 4 or greater on all response items, indicating significant fear. Participants were least concerned that police would be notified as a result of seeking help at an agency ($M = 1.57$, $SD = .975$). Participants reported the highest levels of fear over staff being too busy to help ($M = 2.09$, $SD = 1.18$).

Missing Values

Given the potentially sensitive nature of some items, we paid special attention to missing data. The instrument was administered via face-to-face interview, so missing items in our dataset indicate participant refusal. A total of 13.4% of the sample ($n = 90$) skipped one or more items on the scale. Little's Missing Completely at Random test (Little, 1988) was conducted in SPSS 21 to identify any systematic data loss, and the significant result ($\chi^2 = 277.87$, $df = 159$, $p < .001$) indicated that data loss was dependent on other respondent characteristics. The most frequent pattern of missing data ($n = 52$) involved items 9 (staff restricting freedom) and 10 (staff calling police). As these were the final items on the FATES, one plausible explanation for this pattern could have been respondent fatigue. Researchers involved in the original trial, however, reported that fatigue was very rare. Given the especially sensitive nature of these items and that the instrument was administered midway through the baseline interview schedule, we investigated the skip pattern further using the chi-square test of independence. There was no loss associated with gender, age, history of abuse, or recent involuntary hospitalization. Significant differences were found based on ethnic group; while comprising 26.9% of the sample ($n = 181$), 54.7% of individuals with this skip pattern were African American. The skip pattern reveals a potentially concerning phenomenon in a group actively seeking services; it is important to consider what environmental or relational factors may relate to participants concealing information on fears of restricting freedom or police involvement. Overall, 17 participants (2.5%) missed more than 20% of response items and were therefore dropped from the analysis (Downey & King, 1998; Hawthorne & Elliott, 2005). Neither gender, ethnicity, nor recent incarceration or involuntary hospitalization were strongly associated with this elevated level of missing responses. It was, however, contingent upon age category, $\chi^2(2, N = 655) = 18.22$, $p < .001$; younger adults (ages 18–35) were less likely to miss three or more items, while older adults (ages 55–80) were more likely. To conserve risk of bias while retaining the vast majority of original data, person-level mean substitution was used in the remaining missing data; that is, where participants missed two or fewer items, the mean of the completed items was calculated individually for each respondent and imputed.

Estimation of the Correlation Matrix

Polychoric correlations were calculated in FACTOR (Lorenzo-Seva & Ferrando, 2006) to estimate associations between item pairs. The use of this matrix is supported with ordinal scales and when univariate descriptive statistics are highly skewed or kurtotic (Flora & Curran, 2004; Muthén & Kaplan, 1985, 1992); they are robust against moderately nonnormal latent response distributions. See Table 1 for skewness and kurtosis values of the 10-item



Table 1
Descriptive Statistics of the 10-Item FATES

Item	N	M	SD	Skewness	Kurtosis
1	656	1.740	1.165	1.533	1.251
2	656	2.072	1.184	.791	-.418
3	656	1.919	1.143	1.086	.211
4	656	1.856	1.201	1.271	.464
5	656	1.997	1.144	.921	-.128
6	656	1.854	1.091	1.212	.735
7	656	2.093	1.176	.722	-.481
8	656	1.869	1.074	1.144	.598
9	656	1.793	1.084	1.273	.682
10	656	1.567	.975	1.892	3.002

Note. According to Muthén and Kaplan (1985, 1992), Pearson correlations are indicated when both indices of skewness and kurtosis for ordinal items are lower than 1; given the variability of these indices on the FATES items, we incorporated polychoric correlations in the analysis. FATES = Fear of Adverse Treatment Experiences Scale.

FATES. The distribution of polychoric correlations among the 45 pairs of response items ranged from .45 to .80 and is shown in Table 2.

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Though the Pearson’s covariance matrix is best calculated on continuous data, it is the foundation for more traditional factor analytic approaches, which will be used later in the report.

As expected, in comparison with the polychoric matrix, the Pearson’s matrix showed consistently lower correlations in the pairs of scale items, ranging from .31 to .73 (see Table 2). Adequacy of the matrices was assessed using the Kaiser-Meyer-Okin (KMO) measure of sampling adequacy and Bartlett’s Test of Sphericity. The KMO index at .918 is considered very good for the purposes of data reduction. Bartlett’s test was highly significant ($\chi^2 = 3846.59, df = 45, p < .001$), indicating that the population matrix was not an identity matrix. Both approaches indicate suitability for data reduction and the differences between the matrices are largely additive—the size and direction of correlations being relatively similar.

Dimension Reduction

Several reduction procedures are available depending on the goals of analysis. We utilized two procedures to reduce the dimen-

sionality of the FATES and compared the results. Principal components analysis (PCA), an approach adapted to data which violates assumptions required for factor analysis, was conducted based on the polychoric correlation matrix. Then, principal axis factoring was used as a comparison and based on the Pearson’s covariance matrix.

Principal component analysis. To reduce the correlated observed variables to a smaller set of independent composite variables, we applied PCA to the polychoric matrix in FACTOR (Lorenzo-Seva & Ferrando, 2006). Eigenvalues were examined for a sharp drop in component variance, and Kaiser’s Rule was used to determine the number of components to be retained; that is, all components with eigenvalues greater than one. The first eigenvalue was 6.73, explaining 67.3% of the variance, and the second was .961, explaining 9.6% of the variance. The ratio between the first and second eigenvalues and the fact that only the first eigenvalue was greater than one signifies unidimensionality of the scale. We followed up with the Hull method, a convex hull-based heuristic used to balance a model’s goodness-of-fit and degrees of freedom while selecting the number of common factors (Lorenzo-Seva, Timmerman, & Kiers, 2011). Comparable to the scree test used in traditional factor analytic approaches, the Hull method has shown in simulation studies to be less likely to over- or underestimate the number of common factors. In its execution, we entered the range of number of dimensions under consideration (0–2), the comparative fit index as a goodness-of-fit measure, and the degrees of freedom to confirm that the scale can be conceptualized in terms of a single dominant component.

Principal axis factoring. Responses on the FATES were then analyzed in SPSS version 20, using the principal-axis method to extract factors. Communalities were examined to assess the extent each item correlated with the rest of the items on the scale. These were moderate, with the lowest after extraction valued at .475, indicating a fair amount of homogeneity between variables. Eigenvalues and scree plots were examined for a sharp drop in variance, and Kaiser’s Rule was again used to determine the number of factors to be retained. The first initial eigenvalue was 5.673, explaining 56.72% of the variance, and the second was 1.148, explaining 11.49% of the variance. The scree test suggested that the first two factors were meaningful. Given our objective to determine whether the scale contains a general factor, we retained

Table 2
Correlation Matrices of the 10-Item FATES (N = 656)

	Item	Pearson’s matrix										
		1	2	3	4	5	6	7	8	9	10	
Polychoric matrix	1	1										
	2	.65	1									
	3	.68	.72	1								
	4	.63	.57	.72	1							
	5	.52	.64	.62	.57	1						
	6	.46	.64	.61	.55	.77	1					
	7	.48	.68	.60	.57	.74	.77	1				
	8	.45 ^a	.64	.62	.52	.74	.73	.80 ^b	1			
	9	.69	.66	.74	.68	.60	.63	.62	.64	1		
	10	.62	.54	.67	.64	.61	.58	.62	.63	.78	1	

Note. All correlations are significant at the .001 level. FATES = Fear of Adverse Treatment Experiences Scale.
^a Weakest correlation. ^b Strongest correlation.

these factors using Quartimax (orthogonal) rotation to maximize the sum of squared loadings and clarify their interpretation. After Quartimax rotation, the first eigenvalue was 5.246, explaining 52.46% of the variance, and the second was .812, explaining 8.11% of the variance. We also attempted the Varimax method as a challenge to the unidimensional hypothesis. This challenge did not visibly yield two separate strong factors.

Next, we assessed factor loadings of scale items. The loading range for Factor 1 was .644 to .794, and for Factor 2 was $-.337$ to .418. Correlations of .32 or less (indicating less than 10% overlapping variance) were considered too weak to load on the given factor (Tabachnick & Fidell, 2001). Though two factors were originally extracted, multiple cross-loadings weakened their convergent and discriminant validity. Four items loaded moderately on Factor 2; these were the four items most distinctly associated with the original Fear of Inadequate Care scale. Factor loadings are displayed in Table 3.

Then, we forced all items to load onto a single factor for comparison. Communalities were examined to assess the extent each item correlated with the rest of the items on the scale. For the most part, these were moderate, with the lowest after extraction valued at .352, generally indicating a fair amount of homogeneity between variables with the exception of Item 1.

With the lack of clear hypothesis testing in data reduction procedures, the two methods produced similar results. No approach visibly yielded two separate strong factors. Given the evaluation and interpretation of both PCA and factor analytic methods, we report the solution generated through PCA with a single component, as it is more aptly suited to the nature of the data and yields a parsimonious result.

Internal Consistency

To assess internal consistency of the single-factor FATES, ordinal reliability theta (Zumbo, Gadermann, & Zeisser, 2007) was calculated as

$$\theta = \frac{p}{p-1} \left(1 - \frac{1}{\lambda_1} \right),$$

where p represents the number of items entered into the polychoric matrix and λ_1 is the first and largest eigenvalue, or the amount of

Table 3
Factor Loadings, Initial Communalities (h^2), and Percentages of Variance for Principal Axis Factor Extraction and Quartimax Rotation on FATES Items

Item	Factor 1	Factor 2	h^2
1	.644	-.337	.444
2	.722	.044	.550
3	.794	-.190	.599
4	.670	-.189	.446
5	.735	.324	.593
6	.716	.361	.584
7	.752	.416	.656
8	.724	.418	.625
9	.789	-.180	.600
10	.682	-.099	.495
% variance	52.46	8.11	

Note. FATES = Fear of Adverse Treatment Experiences Scale.

variance explained by the first component. This was followed up by computation of Cronbach’s alpha to represent a more negatively biased estimate. When calculated on the polychoric matrix, ordinal theta was nearly identical to Cronbach’s alpha; these were .9460 and .9459, respectively (alpha reduced to .9145 on the Pearson’s matrix). Removing any item from either of the correlation matrices reduced the values of theta and alpha, strengthening the conclusion that this instrument is well represented as a single component. In summary, it appears that the most parsimonious model consists of a single dominant component, and with such strong convergent validity it is reasonable to combine the response items into a single scale that sufficiently covers the construct of fear of adverse treatment experiences in mental health care settings.

Group Comparisons

We sought to further validate the FATES by examining score patterns among groups with characteristics theoretically linked to fear of treatment, utilizing nonparametric tests of proportion equality to account for nonnormality and the failure of transformations of total fear scores (de Winter & Dodou, 2010). We anticipated that identification as part of an underpowered demographic group and history of involuntary treatment, incarceration, or abuse would translate to increased total scores on the FATES. Increased resiliency factors (i.e., empowerment) were expected to negatively impact total scores. We had no a priori hypotheses regarding responses based on service condition—that is, whether participants were enrolled from self-help agencies or from community mental health agencies.

The Mann–Whitney–Wilcoxon’s tests (see Table 4) showed that service condition made a difference in median total fear scores, with the community mental health agency participants scoring higher than those from self-help agencies ($Z = -3.766, p < .001$). Males and females significantly differed in their median fear scores, but unexpectedly, males scored higher than females ($Z = -3.338, p = .001$). As hypothesized, those who reported experiencing sexual abuse in childhood had significantly higher median scores than those who did not report such abuse ($p = .011$), but this effect was not consistent for those who reported only experiencing physical abuse ($p = .130$). We found no significant increase in scores for those respondents on parole ($p = .265$) or probation ($p = .406$); however, there was a highly significant increase in scores for those who reported that the police had been called on them in the previous month ($p < .001$). Neither the involuntary nature of previous psychiatric admission nor court mandate had an impact on scores; however, in partial support of our hypotheses, for those respondents who reported that their last admission was changed from voluntary to involuntary status, FATES scores were significantly elevated ($p = .016$). Finally, in contradiction to our hypothesis on the impact of resiliency factors, we did not see a significant difference between respondents who believed they were empowered and those who did not.

The Kruskal-Wallis equality of proportions rank test (see Table 5) showed a significant difference, as expected, in total FATES scores between the five categories of ethnic groups surveyed, $\chi^2(5, N = 656) = 18.267, p = .001$, with Caucasians endorsing the lowest levels of fear (mean rank 310.83) and Native Americans endorsing the highest (mean rank 467.83). Post hoc analysis with

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Table 4
Mann–Whitney Tests for Comparisons of Median FATES Scores

Characteristic	<i>n</i>	Mean rank	<i>Z</i>	<i>p</i>
Service condition (<i>N</i> = 406)			−3.766	.000*
Self-help agencies	221	183.26		
CMHAs	185	227.08		
Gender (<i>N</i> = 656)			−3.338	.001*
Male	349	351.42		
Female	307	302.45		
Abuse history				
Sexual (<i>N</i> = 652)			−2.553	.011*
Y	247	350.35		
N	405	311.95		
Physical (<i>N</i> = 654)			−1.516	.130
Y	312	316.92		
N	342	339.10		
Either (<i>N</i> = 656)			−1.694	.090
Y	361	339.71		
N	295	314.78		
Both (<i>N</i> = 651)			−2.450	.014*
Y	198	353.02		
N	453	314.19		
Police involvement				
Currently on probation (<i>N</i> = 656)			−.831	.406
Y	112	341.90		
N	544	325.74		
Currently on parole (<i>N</i> = 656)			−1.115	.265
Y	15	275.13		
N	641	329.75		
Someone called police on you (<i>N</i> = 650)			−3.812	.000*
Y	136	379.50		
N	514	311.21		
Mental health system involvement				
Nature of last admission			−1.173	.241
Voluntary	169	175.67		
Involuntary	195	188.42		
Changed to involuntary after voluntary admission			−2.402	.016*
Y	31	103.15		
N	137	80.28		
Stayed in hospital by court order			−1.197	.231
Y	51	97.64		
N	161	109.31		
Resiliency factors				
Are empowered (<i>n</i> = 314)			−1.491	.136
Y	195	151.56		
N	119	167.24		

Note. The Mann–Whitney test examines the differences in ranked positions of median FATES scores in two groups subject to comparison. Groups with comparatively lower mean ranks have higher concentrations of lower scores, while comparatively higher mean ranks signify a higher concentration of high scores. Service condition excludes those participants assigned to the joint service condition in the original randomized control trial. Due to occasional missing data, *N*s range from 650 to 656 in comparisons of groups classified by gender, abuse history, and police involvement. Response rate was lower in categories of mental health system involvement and empowerment, as not all participants had history of hospitalization, and participants were not asked to report on their own empowerment if they did not demonstrate familiarity with the term. FATES = Fear of Adverse Treatment Experiences Scale; CMHAs = community mental health agencies; Y = yes; N = no.

* Significant at $p < .05$.

pairwise Mann–Whitney tests was conducted, applying a Bonferroni adjustment at a .005 significance level (a result of dividing the critical p value of .05 by 10 intergroup comparisons), with the only significant finding between Caucasians and Native Americans ($Z = -3.804$, $p < .001$). However, due to small sample size this difference is uninterpretable. Our hypotheses on the associations between history of oppressive circumstances (e.g., involuntary treatment, assault) and FATES scores were generally supported by the Kruskal–Wallis tests. A significant difference in total scores

was found between those who had been imprisoned within the last year, within the last month, and not at all, $\chi^2(3, N = 655) = 9.14$, $p = .010$). Follow-up tests showed a difference approaching significance (at $\alpha = .017$, the critical p value of .05 divided by three comparisons) between those who had spent time in jail in the last year and those who had not ($Z = -2.360$, $p = .018$). There was a significant difference in total scores between those who had been beaten, mugged, or raped within the last year, within the last month, and not at all, $\chi^2(3, N = 655) = 9.14$, $p = .010$). Follow-up

Table 5
Kruskal-Wallis Tests for Comparisons of Median FATES Scores

Characteristic	<i>n</i>	Mean rank	χ^2	<i>df</i>	<i>p</i>
Age (<i>N</i> = 655)					
18–35	209	334.44			
35–55	408	331.06			
55–80	38	259.75	5.408	2	.067
Ethnicity (<i>N</i> = 656)					
Caucasian ^a	382	310.83			
African American	177	353.45			
Hispanic	57	319.49	18.267	4	.001 ^a
Asian	19	324.32			
Native American ^a	21	467.83			
Police involvement					
Spent time in jail/prison (<i>N</i> = 655)					
No	488	315.29			
Yes, within 12 months	120	360.00	9.139	2	.010 ^b
Yes, within 1 month	47	378.23			
Mental health system involvement					
Involuntary hospitalization (<i>N</i> = 656)					
No	529	319.28			
Yes, within 12 months	84	363.49	6.686	2	.035 ^b
Yes, within 1 month	43	373.53			
Abuse history					
Beaten/mugged/stabbed/raped (<i>N</i> = 656)					
No ^c	555	310.81			
Yes, within 12 months ^c	80	429.66	32.259	2	.000 ^c
Yes, within 1 month ^c	21	410.62			

Note. The Kruskal-Wallis test examines the differences in ranked positions of median FATES scores in three or more groups subject to comparison. Groups with comparatively lower mean ranks have higher concentrations of lower scores while comparatively higher mean ranks signify a higher concentration of high scores. Due to occasional missing data, *Ns* range from 655 to 656. FATES = Fear of Adverse Treatment Experiences Scale. ^a The *p* values were significant after Bonferroni correction ($p_c = .005$); pairwise comparison of Caucasian and Native American participants. ^b Familywise comparison of scores found initial significant difference, but pairwise comparisons showed no significant differences in levels of police involvement as well as levels of involuntary hospitalization. ^c The *p* values were significant after Bonferroni correction ($p_c = .017$); pairwise comparisons of those who had been beaten, mugged, or raped within the last year versus not at all, and within the last month versus not at all.

tests showed significant differences (also at Bonferroni-adjusted $\alpha = .017$) between those who had been assaulted in the last year and those who had not ($Z = -5.30, p < .001$), as well as between those who had been assaulted in the last month and those who had not ($Z = -2.409, p = .016$). There was an initial significant difference in total scores between those who had been hospitalized against their will within the last year, within the last month, and not at all, $\chi^2(3, N = 656) = 9.14, p = .010$, but post hoc comparisons at Bonferroni-adjusted $\alpha = .017$ yielded no significant differences between pairs. Finally, the analysis was conducted with no expectations regarding the specific impact of age on FATES score, and when dividing the sample into three age intervals, we found no significant differences between groups.

Discussion

The primary objective of this study was to explore evidence supporting the validity of a unidimensional measure of fear of adverse treatment experiences in adults seeking mental health services. Factor analyses support the construct validity of the 10-item FATES in this sample of adults seeking mental health services. In attempting to consider whether the level of fear a service user may have of adverse mental health treatment experiences as reflected in the FATES score can be linked to social or

demographic characteristics, the results suggest that the FATES significantly discriminates based upon service agency characteristics, individual demographics, and past experiences with coercive or disorganized interventions.

While comparing the median scores between groups, we found many expected results, including higher scores among those respondents with history of childhood abuse, involuntary hospitalization, incarceration, and recent victimization. People with histories in oppressive circumstances may be sensitive to treatment that may be perceived as unhelpful, or a further restriction of personal liberties. In the case of childhood victimization, however, this link does not appear to be strongly established, especially as the literature on abuse history and help-seeking seems to emphasize treatment adherence—a different construct—and is generally focused on the psychological characteristics (vs. treatment characteristics)

Table 6
Fear of Adverse Treatment Experiences Scale (FATES)

1	2	3	4	5
Not at all worried	Not too worried	Somewhat worried	Very worried	Extremely worried



of people in recovery from substance abuse (Schober & Annis, 1996). The link between history of victimization and fear of adverse mental health treatment experiences among mental health service users may be mediated by general mistrust or other troubling perceptions and beliefs, and is worthy of further study. In partial alignment with expectations, categorization by ethnic group led to an initially significant general difference in FATES scores, but follow-up analyses showed that this significance only carried to the lowest-scoring group (Caucasians), and the highest-scoring group (Native Americans). With such a small subsample of the latter group, these results are not interpretable. Finally, consistent with perceptions of the voluntary nature of services and less authoritarian atmosphere at self-help agencies, total scores discriminated between participants who attended self-help agencies and those who sought more traditional services, with the former group endorsing lower levels of fear.

We also found results that contradicted initial hypotheses. For one, men tended to score higher than women on the FATES. Compatible with the observations of Kaltiala-Heino et al. (2003), this could be a result of inflated perceptions of risk associated with men in treatment, leading to a pattern of excessive or otherwise inappropriate clinical management of behavior or symptoms. Personal empowerment, as reported by the respondents, did not make a difference in total scores on the FATES, but it is important to note the general lack of consensus on this complex construct (Segal, Silverman, & Temkin, 1995). To better establish evidence for discriminant validity and examine empowerment as a protective factor against fear of adverse treatment experiences, future work would benefit from the inclusion of a well-regarded personal empowerment measure. As no differences were found between respondents who had most recently been involuntarily admitted and those who were voluntary, a most interesting finding concerned the significant elevation in levels of fear for those whose hospitalization status was changed from voluntary to involuntary midstay. This is perhaps the clearest illustration of the effects of institutional betrayal as perceived by those who put faith in voluntary services that unfortunately could not meet their presented mental health needs.

Strengths of this scale include its origin in a service user-directed, recovery-focused study, as well as its brevity and ease of use. Some limitations should be noted, and results should be interpreted cautiously until more evidence for the validity of the FATES is generated. First, this study was unlikely to have included those potential participants with the most concerning fears of adverse mental health treatment experiences who presumably have self-selected out of treatment options altogether. On balance, the trial's inclusion of alternative support services, the self-help agencies, is likely to yield valuable information from participants whose apprehension of traditional services may have precluded their participation in the original randomized controlled trial. Second, while we have considered unique demographic characteristics such as ethnic background, external validity is compromised as the respondents were predominantly Caucasian. The FATES, having been administered at baseline interview only, has limited utility in causal inferences and affects the reliability of scores. Future studies should utilize this scale in tandem with previously established measures of treatment-related fear in more ethnically diverse samples. It is important to consider that treatment aversion may also be influenced by factors not captured by this scale, such as stigma and

familiarity with mental health systems of care. Some opportunities for further validation of the FATES may be found in studies of treatment adherence and termination, therapeutic rapport, or the characteristics of people who have chosen not to engage in conventional mental health care. Given previous use of these scale items to predict patterns of help-seeking (Segal, Hardiman, & Hodges, 2002), future exploration of self-perceived need for services in tandem with fear of services may be useful in clinical or research applications.

In an uncertain landscape of services available to increasingly eligible service recipients, a user-friendly measure of fear such as the FATES can support evaluative or clinical services in any number of mental health settings. These may include primary and urgent care centers, which commonly serve as a first or even regular point of contact for mental health consumers, particularly those who are not familiar or are uncomfortable with mental health resources. The tool requires no specialized training and can be administered by psychiatric rehabilitation practitioners from a variety of educational backgrounds, having utility for instance within the context of a clinician's intake interview or with peer providers. However, until more supportive evidence has been generated for use of the scale, some practical considerations should also be noted. First, with respect to the specific way we have handled participant nonresponse on certain items, practitioners interested in utilizing the instrument should keep in mind that the results reported in this study would only generalize to instances where the same exclusionary and imputation procedures (i.e., excluding those with less than 80% response rate and imputing the mean score for otherwise missing items) are followed. Second, nonparametric tests, as used in this study, lead to slightly different interpretation of results—in group comparisons, we note that the lower median FATES scores are clustered in those comparison groups with lower ranks, while the higher scores are clustered in those groups showing higher ranks. This is a less-intuitive result than what is yielded in parametric counterparts, and because the data are ranked, some information about the magnitude of score differences is lost—it cannot be assumed. Furthermore, nonparametric tests are more susceptible to Type II error, where there may be a difference between comparison groups but the test does not have adequate power to detect it. Finally, as with any assessment, use of the FATES should be grounded in good clinical reasoning, with respect for client capacity and comfort, as well as the treatment environment and therapeutic relationship. These considerations point to a need for practitioner sensitivity in expectations for client performance, as well as handling and interpretation of scores.

Nevertheless, the FATES could prove especially useful for those professionals who are often charged with facilitating an entry process to an often confusing network of survival resources, medical assistance, vocational support services, and mental health care (Robiner, 2006). With more research on client preferences, adverse experiences, and service outcomes, the tool may bolster practitioners' insight on the best-matched services for the person seeking support, and help them anticipate challenges which may arise out of limited or aversive treatment options. Ultimately, this is a measure created with high regard for service user empowerment and has strong potential to supplement the consumer voice in the changing face of service provision.

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Appendix



AQ: 5

Fear of Adverse Treatment Experiences Scale Items

1. Before you came to this agency, how worried were you that the counselors/staff would have you hospitalized for psychiatric reasons?
2. Before you came to this agency, how worried were you that counselors/staff wouldn't really listen to you?
3. Before you came to this agency, how worried were you that the counselors/staff would make you do things you didn't like to get the things you wanted?
4. Before you came to this agency, how worried were you that the counselors/staff would make you take medications you didn't want?
5. Before you came to this agency, how worried were you that counselors/staff didn't have enough training to give you the services you wanted?
6. Before you came to this agency, how worried were you that counselors/staff would have too many
7. Before you came to this agency, how worried were you that counselors/staff would be too busy to give you the help that you wanted?
8. Before you came to this agency, how worried were you that the agency would be too disorganized to be able to help you?
9. Before you came to this agency, how worried were you that the counselors/staff would restrict your freedom?
10. Before you came to this agency, how worried were you that the counselors/staff would call the police on you?

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AUTHOR QUERIES

AUTHOR PLEASE ANSWER ALL QUERIES

1

AQau—Please confirm the given-names and surnames are identified properly by the colors.

■ = Given-Name, ■ = Surname

The colors are for proofing purposes only. The colors will not appear online or in print.

AQ1—Author: Please be sure to provide the name of the department(s) with which you and your coauthors are affiliated at your respective institutes if you have not already done so. If you are affiliated with a governmental department, business, hospital, clinic, VA center, or other nonuniversity-based institute, please provide the city and U.S. state (or the city, province, and country) in which the institute is based. Departments should be listed in the author footnote only, not the byline. If you or your coauthors have changed affiliations since the article was written, please include a separate note indicating the new department/affiliation: [author's name] is now at [affiliation].

AQ2—Author: Please limit keywords to 3-5 words or phrases.

AQ3—Author: Please cite Table 6 after the Table 5 citation.

AQ4—Author: Please provide page range for Richman et al., 2010, chapter.

AQ5—Author: Is the Appendix title ok as added here? If not, please provide an appropriate title for the appendix.
