An Investigation Into the Factor Structure of the Attitudes to Suicide Prevention Scale

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Abstract. *Aim:* The aim of this study was to investigate the factor structure of the Attitudes to Suicide Prevention Scale (ASPS). *Method:* The ASPS was distributed to all staff in a UK National Health Service Trust (N = 957). We conducted an exploratory factor analysis followed by a confirmatory factor analysis by splitting the data 60/40 into training and testing subsets. A multiple regression analysis was carried out to investigate whether the overall scale score varied as a function of professional role, age, and gender and whether respondents had completed suicide prevention training or not. *Results:* Two items displaying poor item-scale correlation were excluded from the factor analysis and a further item was excluded as it was based on different anchor points. For the remaining 11 items, no adequate factor structure emerged. The scale total demonstrated statistically significant differences in attitudes between staff groups (defined by attendance at suicide awareness or prevention training, by gender, and by level of patient contact), but not between groups defined by age range. Generally, however, there were positive attitudes across all Trust staff. *Limitations:* This study had a low response rate (24%) and was cross-sectional which limits the conclusions that could be drawn. Furthermore, other areas such as convergent validity and test–retest reliability were not examined. *Conclusion:* Our findings found no satisfactory factor structure for the ASPS. Further scale development would be beneficial.

Keywords: staff attitudes, suicide prevention, clinician, ASP

Approximately 800,000 people die by suicide each year (World Health Organization, 2017). A third of those who are lost to suicide are individuals who had been in contact with mental health services in the 12 months prior to death (Luoma, Martin, & Pearson, 2002). Every clinical encounter is an opportunity to potentially prevent a suicide and, clearly, mental health services have a central role and responsibility in suicide prevention.

The opportunity to identify those at risk of suicide does of course extend beyond contact with mental health services. For example, an estimated 77% of those who die by suicide had attended their general practitioner (GP) service in the 12 months prior to death (National Confidential Inquiry Into Suicide and Homicide by People with Mental Illness, 2016). Identification of risk can itself be problematic. A recent study highlighted the high rates of misclassification between suicidal behaviors and nonsuicidal self-directed violence and the potential impact of this on risk assessment, management, and interventions (Cwik & Teismann, 2017). The authors found that rates of misclassification were largely independent of length of professional experience among psychologists, but they identified particular biases when classifying the behavior of female patients and those with a diagnosis of borderline personality disorder (e.g., suicidal behavior of female patients was significantly more often interpreted as nonsuicidal self-directed violence [30.5%] compared with male patients [52.6%]).

The response of all health professionals to people at risk of suicide is of vital importance and it is likely to be influenced by their attitudes toward suicide and more specifically toward suicide prevention. Furthermore, beliefs and attitudes can negatively impact upon the effectiveness of suicide risk assessment and management (Herron, Ticehurst, Appleby, Perry, & Cordingley, 2001; Valente, 2011); for example, health professionals' beliefs about the preventability of suicide is likely to influence how risk is assessed and managed (Ramberg, Di Lucca, & Hadlaczky, 2016). Attitudes toward responsibility are also likely to affect engagement with assessing risk, willingness to access training in risk management (Herron et al., 2001), or influence risk assessment and management skills (Brunero, Smith, Bates, & Fairbrother, 2008).

Numerous studies have investigated the attitudes of health-care staff toward suicide prevention (Brunero et al.,

https://doi.org/10.1027/0227-5910/a000608

2008; Herron et al., 2001; Nebhinani, Gaikwad, & Tamphasana, 2013), often using the Attitudes to Suicide Prevention Scale (ASPS) developed by Herron and colleagues (2001). In the original design of the scale, factor analysis was performed on 28 items that were reduced to 14 when items with poor factor loadings were dropped. Following validation of the scale by Herron et al. (2001), only one subsequent study has sought to replicate the original findings of the internal validity of the scale (Brunero et al., 2008), but to our knowledge, no studies have attempted to investigate the factor structure of the ASPS. Thus, despite its frequent use, the psychometric soundness of the ASPS has received little attention.

Herron et al. (2001) found that attitudes toward suicide prevention differed significantly between the four groups of health professionals they investigated: GPs, accident and emergency nurses, psychiatrists in training, and community psychiatric nurses. They concluded that more positive attitudes were associated with being a mental health professional, working in the community, and having had previous training in suicide risk assessment. Herron et al. (2001) suggested that some negative attitudes could result in the underestimation of risk and recommended that negative attitudes should be assessed and targeted in training designed to improve the management of suicide risk. More recently, Nebhinani et al. (2013) used the ASPS to study the attitudes of 308 nursing students. While nearly half of their sample had positive attitudes toward working with suicidal patients, half also considered suicide prevention efforts to be ineffectual. Nebhinani and colleagues (2013) concluded that this highlighted the need for further training in suicide prevention, recommending regular educational and training programs on suicide assessment, risk reduction, and prevention of suicide, supervision, and ongoing support for new staff and student nurses. Previous studies have investigated the impact of training on attitudes to suicide prevention although sample sizes have been small (Appleby et al., 2000; Brunero et al., 2008; Ramberg et al., 2016) and therefore larger-scale research in this area would be beneficial.

The Current Study

This study emerged out of discussions within a UK National Health Service (NHS) Trust about the need to prioritize suicide prevention, as has been identified within the NHS more widely (The Mental Health Taskforce, 2016). As part of this effort, a survey of the attitudes of Trust staff to suicide prevention was conducted. While previous studies using the ASPS focused on health professionals, in this study we investigated attitudes across the entirety of Trust staff, consistent with local and national policy initiatives that highlight suicide prevention as everybody's business (Mathieson & Twiselton, 2014; Public Health England, 2016).

In sum, this study had three aims: (a) to investigate the internal consistency of the ASPS and its factor structure; (b) to investigate whether differences in attitudes to suicide prevention existed between staff members with different vocational roles (as defined by their contact with patients) or as a function of age or gender; and (c) to explore whether there was an association between attendance at training in suicide awareness or prevention, and attitudes to suicide prevention.

Method

Participant Recruitment and Procedure

The NHS Trust studied provides community and mental health services to a population of half a million people and employs around 4,000 staff. The clinical services are divided into four Care Groups – Mental Health (community mental health teams, crisis teams, and primary care psychological therapy), Community Services (e.g., district nursing, occupational therapy, physiotherapy, cardiac rehabilitation), Children and Families (e.g., health visiting, school nursing, child and adolescent mental health), and Specialist Services (e.g., learning disability, specialist dentistry, neurology, diabetes) – with a fifth group covering Corporate Services.

An anonymous online questionnaire was distributed in December 2016 to all Trust employees (≈ 4,000) via the Trust newsletter, which was delivered electronically to all employees. Links to the questionnaire were also distributed via e-mails through the communication channels of each Care Group within the Trust. Since this was designed as a service audit, NHS ethical approval was not required. As part of the introduction to the questionnaire, participants were advised of the subject matter to be addressed, that they would not be identifiable, and they were asked to contact the Suicide Prevention Project Lead for the Trust if they had any questions or concerns.

Measures

Demographics

All participants were asked to respond to questions on age, gender, suicide prevention or awareness training attended, Care Group, geographical work base, and level of patient contact offered by vocational role. Level of patient contact was defined by three categories: clinical staff with patient contact (e.g., those employed in clinical roles); nonclinical staff with some patient contact (e.g., estates, facilities, and administration); and staff with no patient contact (e.g., support services, governance, IT, nonexecutive directors).

Attitudes to Suicide Prevention

The ASPS (Herron et al., 2001), is a 14-item questionnaire (see Table 2) that asks people to rate their attitudes on a 5-point Likert scale from 1 to 5 anchored at *strongly disagree, disagree, uncertain, agree*, and *strongly agree*. Two items (Items 4 and 14) are reverse-scored and one item (Item 14) is anchored at *none, few, many, most, all*. A lower score on the ASPS indicates more positive attitudes toward suicide prevention.

Data Analysis

Originally, we conducted an exploratory factor analysis (EFA) in SPSS version 22 using minimum residual extraction with an oblimin rotation, and applied the Kaiser–Guttman criteria (eigenvalues > 1) for retaining items. Following initial reviewer comments and further discussion within the research team, several issues arose. First, the original validation paper for the ASPS does not report the factor structure or the item loadings resulting from their principal components analysis (PCA) of the scale. The first author (DS) contacted the corresponding author of the original paper by Herron et al. (2001) to make enquiries regarding the results of the original PCA of the ASPS; however, details beyond those included in the paper were unfortunately no longer available (L. Appleby, personal communication, Jan 21, 2019). The use of a total score for the ASPS (Brunero et al., 2008; Herron et al., 2001) appears to assume a single-factor solution, as had we; however, there is no published record of such a structure having been validated. Furthermore, a single-factor structure may be somewhat surprising, given that the initial pool of items generated by Herron et al. (2001) prior to PCA could be grouped into six themes: the accuracy of suicide risk assessment in clinical practice; the interpretation of expressions of suicidal intent; the responsibility of a clinician in preventing suicide; the practicality of preventing suicide in clinical practice; the preventability of suicide in general; and the impact of nonclinical factors on suicide rates. Additionally, the original validation of the ASPS by Herron et al. (2001) was carried out using PCA, which, while frequently used interchangeably with EFA, has different objectives and results in different outcomes from EFA (Costello & Osborne, 2005). In the absence of a validated factor structure to confirm, we decided to first use EFA to investigate the factor structure of the ASPS, then validate the factor structure that emerged from our EFA using confirmatory factor analysis (CFA) in a subset of the sample. Data were randomly divided into training and testing subset samples, comprising 60% and 40% of the dataset, respectively. EFA was conducted using the Psych package (Revelle, 2018) in R, with a minimum residual extraction method and oblimin rotation, to allow for correlation between factors. As data are ordinal and not continuous, we used polychoric correlations instead of Pearson's correlations to reduce the likelihood of overfitting, as recommended by Van der Eijk & Rose (2015) and Watkins (2018). First, parallel analysis (PA) was conducted on the training sample in order to obtain a recommendation of the number of factors to retain. PA indicated that two factors should be retained and, consequently, we conducted an EFA specifying two factors. Visual inspection of data using histograms of responses to individual items showed the data were not normally distributed, therefore the EFA was conducted on the covariance matrix instead of the correlation matrix, as this is less affected by issues of dispersion and violations of multivariate normality (Tinsley & Tinsley, 1987; Yong & Pearce, 2013). Items with loadings below .3 were suppressed. Inspection of inter-item correlations demonstrated that Items 7 ("It is easy for people not involved in clinical practice to make judgments about suicide prevention") and 9 ("People have the right to take their own lives") did not correlate well with any of the other items in the scale, and therefore they were removed. Item 14 ("What proportion of suicides do you consider preventable?") was also removed prior to factor analysis, as this item is not on the same scale as the other items. The ratio of participants to items was approximately 50:1 for the EFA and 37:1 for the CFA (ratios of greater than 10:1 are considered acceptable, with greater than 30:1 desirable; Yong & Pearce, 2013). RMarkdown of the analysis code is available from the corresponding author. Internal consistency for the ASPS was calculated (Cronbach's α and McDonald's ω). A multiple linear regression was performed with the total score for the scale as the dependent variable and attendance at training, gender, work role, and age range as the independent variables. The regression was conducted using SPSS 22 for Windows. The α value for all tests was .05.

Results

Participants

In total, 1,012 staff members returned the questionnaire (Table 1), a response rate of approximately 25%. Of the 1,012 respondents, 797 identified as female, 154 as male, five preferred not to state their gender, and one identified

Table 1. Number of respondents by care group and vocational role

	Vocational role			
Care group	Number of clinical staff (%)	Number of nonclinical staff but with some patient contact (%)	Number of staff with no patient contact (%)	Total number of staff (%)
Specialist services	82 (9)	15 (2)	10 (1)	107 (11)
Mental health	292 (31)	38 (4)	15 (2)	345 (36)
Community health	209 (29)	21 (2)	8 (1)	238 (25)
Corporate services	5 (1)	19 (2)	130	154 (16)
Children and families	88 (9)	15 (2)	10 (1)	113 (12)
Total (%)	676 (71)	108 (11)	173 (18)	957 (100)

Table 2. Mean scores per item

Item	*M*	*SD*
Q1. I resent being asked to do more about suicide	1.69	0.798
Q2. Suicide prevention is not my responsibility	1.66	0.823
Q3. Making more funds available to the appropriate health services would make no difference to the suicide rate	2.11	1.010
Q4. Working with suicidal patients is rewarding (R)	2.63	0.750
Q5. If people are serious about ending their life by suicide, they don't tell anyone	2.65	1.011
Q6. I feel defensive when people offer advice about suicide prevention	1.88	0.758
Q7. It is easy for people not involved in clinical practice to make judgments about suicide prevention	3.26	0.948
Q8. If a person survives a suicide attempt, then this was a ploy for attention	1.76	0.796
Q9. People have the right to take their own lives	3.23	0.862
Q10. Since unemployment and poverty are the main causes of suicide, there is little that an individual can do to prevent it	1.80	0.663
Q11. I don't feel comfortable assessing someone for suicide risk	2.95	1.289
Q12. Suicide prevention measures are a drain on resources, which would be more useful elsewhere	1.66	0.697
Q13. There is no way of knowing who is going to end their life by suicide	2.82	1.012
Q14. What proportion of suicides do you consider preventable? (R)	2.86	0.779
Total	32.96	0.198

as transgender. Of the respondents, 55 failed to complete the ASPS and were excluded from the analysis. This left 957 respondents who completed the ASPS (Herron et al., 2001) yielding a final response rate of approximately 24%. Table 1 provides a breakdown of respondents by vocational role and by care group.

The means and standard deviations for individual items of the ASPS are summarized in Table 2.

Factor Analysis

The Kaiser–Meyer–Olkin measure of sampling adequacy was 0.879, and Bartlett's test of sphericity result was significant (2,312, $df = 55$, $p < .001$), both indicating that the 11 items were suitable for factor analysis.

Internal Consistency

Cronbach's α for the 14-item ASPS for this study was .76. This compares with .77 reported in the validation study by Herron et al. (2001) and .76 reported by Brunero et al. (2008). With items Q7, Q9, and Q14 removed, Cronbach's α for the remaining 11 items was .79. McDonald's ω was calculated as .79 for the original 14 items and .81 with Q7, Q9, and Q14 removed.

EFA Results

Parallel Analysis

Examination of the loadings matrix for a two-factor solution, as suggested by parallel analysis, indicated that Items 4 ("Working with suicidal patients is rewarding") and 8 ("If a person survives a suicide attempt, then this was a ploy for attention") did not load. The BIC and RMSEA model fit indices suggested that the two-factor model was an acceptable fit, RMSEA= .064 (90% CI [051, .077]), as values below .07 are classed as acceptable (Steiger, 2007), BIC= -104.6. The Tucker–Lewis index was .94. The chi-square test result was highly significant and therefore did not indicate a good fit, $\chi^2(34) = 109.68$, $p < .001$; however, when sample size is large, chi-square tests can reject even correctly fitted factor models (Bentler & Bonett, 1980; Jöreskog & Sörbom, 1996).

One-Factor Model

As previous work has assumed a single-factor structure for the ASPS, we also fitted a one-factor model and compared this with the two-factor model suggested by parallel analysis, using an ANOVA. There was a statistically significant difference between the one- and two-factor models, $p < .001$, and examination of the BIC model fit statistics indicated that the one-factor model was a better fit (two-factor BIC = –104.60 vs. one-factor BIC = –114.16). The RMSEA for the one-factor model was not acceptable, RMSEA = .071 (90% CI [0.059, 0.082]) and the Tucker–Lewis index was .92. The chi-square test result was significant, indicating poor fit, $\chi^2(44) = 163.15$, $p = <.001$. Given the six themes involved in initial item generation by Herron et al. (2001), a one-factor model would be conceptually surprising, as qualitatively different items are then grouped together on a single factor.

CFA Results

What we can conclude from these analyses is that there is no factor structure that satisfies the requirements of both statistical and conceptual fit, for the current set of items. Neither model is a good statistical fit on any of the fit indices.

We have two "competing" models: the conceptual fit model (two factors) and the statistical fit model (one factor). We used the testing sample to estimate both of these models in a new, independent sample, using CFA to see if support for either of the competing factor solutions could be found. CFA was conducted using the Lavaan package (Rosseel, 2012) for R. Diagonal weighted least squares (DWLS) was used to estimate the factor structure, as this is less biased for ordinal data (Li, 2016).

One-Factor Model

The RMSEA for the one-factor model was not acceptable, RMSEA = .075 (90% CI [0.061, 0.088]) and the Tucker–Lewis index was .96. The chi-square test result was significant, indicating poor fit, $\chi^2(44) = 144.65$, $p = <.001$.

Two-Factor Model

The RMSEA for the t wo-factor model was not acceptable, RMSEA = .080 (90% CI [0.063, 0.098]) and the Tucker–Lewis index was .96. The chi-square test result was significant, indicating poor fit, $\chi^2(26) = 94.08$, $p = <.001$.

It should be noted, however, that the cut-off that we used of .07 for the acceptability of model fit is purposefully stringent (Steiger, 2007). MacCallum, Browne, and Sugawara (1996) have suggested a graded approach whereby a value of < .05 indicates close fit, .05–.08 indicates fair fit, .08–.10 indicates mediocre fit, and values above 0.10 indicate poor fit. By these criteria, our CFA RMSEA values for both models could be considered to indicate a fair fit.

Multiple Linear Regression

A multiple linear regression was run to predict total scale score (of the 11-item ASPS) from gender, attendance at training, age range, and role. Given the limited support for

Table 3. Summary of multiple regression analysis

Variable	B	SE_β	β	p
Constant	24.804	.972	–	< .0005
Gender (female, male)	–1.087	.438	–.072	.013
Training attendance (no, yes)	–4.883	.353	–.419	< .0005
Age range				
18–24 vs. 25–34	.455	.895	.030	.611 (*ns*)
18–24 vs. 35–44	1.330	.870	.102	.127 (*ns*)
18–24 vs. 45–54	1.357	.852	.116	.112 (*ns*)
18–24 vs. 55–64	2.219	.885	.156	.012
18–24 vs. 65–75	3.207	1.535	.069	.037
Work role: clinical vs. no contact	1.283	.518	.073	.013
Work role: clinical vs. nonclinical, some contact	1.515	.442	.105	.001

Note. B = unstandardized regression coefficient. SE_β = standard error of the coefficient. β = Standardized coefficient. *ns* = not significant at $p > .05$.

the unidimensionality of the scale, these results should be treated with caution.

There was linearity as assessed by partial regression plots and a plot of studentized residuals against the predicted values. There was homoscedasticity, as assessed by visual inspection of a plot of studentized residuals versus unstandardized predicted values. There was no evidence of multicollinearity, as assessed by tolerance values greater than 0.1. There were two outliers with studentized deleted residuals greater than ±3 standard deviations; however, there were no leverage values greater than 0.2, and no values for Cook's distance above 1. Because the results did not differ substantially with these outliers removed, they were included in the analysis. The assumption of normality of the residuals was met, as assessed by a Q-Q Plot.

The multiple regression model statistically significantly predicted scale total, $F(9, 941) = 32.537$, $p < .0001$, adj. $R^2 = .230$. However, this would indicate a small effect size (Cohen, 1988). Prior attendance at suicide awareness or suicide risk training, gender, and work role based on level of patient contact all added statistically significantly to the prediction ($p < .05$). However, age range only became significant from the age range of 55–64 years and older. Regression coefficients and standard errors can be found in Table 3.

Discussion

The use of the scale total, both in this study and in previous studies, should be treated with caution given that we were unable to verify a factor structure for the ASPS. This study did support the findings of previous studies (Brunero et al., 2008; Herron et al., 2001), that the ASPS demonstrates good internal consistency. However, data from this study indicate that the internal reliability of the scale would be improved by removing two of the 14 questions, namely, Q7 ("It is easy for people not involved in clinical practice to make judgments about suicide prevention") and Q9 ("People have the right to take their own lives"). This fits with informal feedback from participants in the survey that the meaning of Q7 is not clear and that a negative response to this question would not necessarily imply a negative attitude to suicide prevention. Question 9 may also be confounded given the debate surrounding voluntary euthanasia versus suicide prevention.

It is noteworthy that a method of factor extraction that has been frequently used for the validation of Likert scale measures is to apply the Kaiser criteria (eigenvalues greater than 1) supplemented by visual inspection of the scree plot of eigenvalues. Applying this approach to the dataset from the current study would indicate a one-factor solution. This would be misleading and would not be supported by either a theoretical construct or by the more appropriate factor analysis procedure detailed herein. This should serve as a note of caution when selecting previously validated scales for research purposes and also supports the growing call for replication studies into scale validation.

Attitudes were found to be more positive among those who had attended suicide awareness or prevention training compared with those who had not attended training. It is important to note that this was a cross-sectional study and furthermore it is possible that the staff with a more positive attitude to suicide prevention would be more likely to seek out and attend training. Therefore, these findings do not provide evidence that training promotes a positive change in attitude; however, other studies (Appleby et al., 2000; Brunero et al., 2008; Ramberg et al., 2016) have specifically investigated this link and provide some limited evidence that this may be the case.

The findings from the present study suggest that attitudes to suicide prevention were more positive (i.e., scores on the ASP scale are lower) among staff groups with greater patient contact. It should be stressed, however, that overall attitudes were positive, in that none of the three staff groupings reported mean total scores higher than the midpoint for the scale (which would indicate more negative attitudes).

Males in this sample were found to have significantly more positive attitudes than females. This was contrary to previous findings from Brunero et al. (2008), who reported no difference on total score of the ASPS based on gender.

Herron et al. (2001) and Brunero et al. (2008) found no significant association between ASPS total score and age. Nebhinani et al. (2013) also found no significant difference in attitudes between different age ranges although they did note that in the population they studied, the overall age range was quite narrow. The current study found that the mean total scale scores for the five age ranges increased through the age bandings, suggesting a more negative attitude with increasing age. However, age range only became significantly predictive of total ASPS score with a negative correlation from the 45–54 age range onward. Age bandings were used in this study as another means of ensuring confidentiality; however, if actual age in years had been collected the results may have been more illustrative.

As Herron et al. (2001) made clear when developing the ASPS, attitudes deemed more negative (and therefore with higher scores on the scale) are not implied to be incorrect. However, they hypothesized that responses deemed more negative to suicide prevention may be indicators of behaviors that are less effective in managing those at risk of suicide. They gave examples from their findings of (a) a group that was most likely to believe that people who are serious about dying by suicide will not tell anyone and (b) a group who reported most agreement that nonfatal self-harm is a "ploy for attention," and made the suggestion that such attitudes could result in the underestimation of risk in people with suicidal ideas or recent self-harm. However, future research might investigate the extent to which all of the items (e.g., "I don't feel comfortable assessing someone for suicide risk") are actually measuring attitudes toward suicide prevention.

Limitations

Although our sample size was large, it is important to note that the response rate was low (24%), and therefore it is possible that people with more negative attitudes did not complete the survey. Furthermore, the response rate itself is only an estimate because, owing to the method of recruitment, it is not known exactly how many people from the total staff employed by the Trust received the invitation to complete the survey. Unfortunately, we do not have data on the nonresponders and thus were not able to explore how representative our sample was of the total workforce. Our sample differed from that of the original scale development study. Our sample includes all NHS Trust staff rather than just health professionals and therefore it is possible that this has introduced measurement variance, that is, the scale may not be reflecting the same construct across the different samples (Hussey & Hughes, 2018). The majority of responders (79%) were female and this may need to be taken into account before generalizing the results. The cross-sectional nature of this study limits the conclusions that can be drawn, thus no inferences can be made about how the attitudes reported in the survey affect the interactions between staff and those at risk of suicide. The scope of this study limited the investigation of the validity of the scale. For instance, convergent and test–retest validity were not examined.

Conclusion

This study did not yield a satisfactory factor structure for the ASPS, and as the unidimensionality of the scale has not been confirmed, use of the scale's total score should be treated with caution. Further attention to scale development would be beneficial, to ensure statistical and conceptual fit in the factor structure. Researchers and evaluators might wish to consider using alternative existing scales to assess attitudes toward suicide prevention; including scales which focus on attitudes and knowledge more broadly (e.g., Batterham, Calear, & Christensen, 2013; Kishi, Kurosawa, Morimura, Hatta, & Thurber, 2011; Kodaka, Postuvan, Inagaki, & Yamada, 2011; Scocco, Castriotta, Toffol, & Preti, 2012). It could be hypothesized that there are benefits to an organization in the act itself of carrying out a survey of this type. Enquiring about attitudes to suicide prevention could help individuals reflect on their own beliefs in a beneficial way and help strengthen the message that it is important that all staff is aware of suicide risk and that suicide prevention is indeed everyone's business.

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History
Received September 19, 2018
Revision received March 26, 2019
Accepted March 30, 2019
Published online July 16, 2019

Authorship
All authors contributed to the writing of this submission. GL provided statistical consultation. We would like to acknowledge the work of the host trust for questionnaire distribution and for data collection.

Funding
DMS is completing a PhD funded by Cumbria Partnership NHS Foundation Trust.

OJK is supported by a fellowship from an FWO Odysseus grant (Myin-Germeys, FWO GOF8416N).

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