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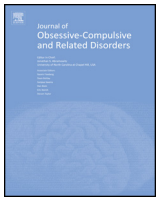

Short communication

Obsessive compulsive disorder prevalence may not increase with latitude: A re-analysis and extension of Coles et al.

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ABSTRACT

A recently published correlational analysis has suggested a linear increase in the prevalence of Obsessive-Compulsive Disorder (OCD) as a function of increasing latitude. This was attributed to a misalignment of the circadian system leading to increased obsessive tendencies and perseverative thought. The present article sought to evaluate these claims in the data reported by those authors using a fully Bayesian meta-analytic approach testing both linear and non-linear effects. Findings were then validated against a sample of independently coded studies. Whereas re-analysis of the original data provided strong evidence in favour of a non-linear relation between latitude and OCD prevalence, analysis of the independently coded studies provided evidence against any statistical relation. In summary, the link between OCD prevalence and latitude is likely to be weaker than previously thought. Further research is required before firm conclusions can be drawn.

Environmental factors such as latitude and light exposure are associated with differences in circadian preference, with an evening orientation associated with higher latitude, longer days, lower levels of light exposure during the day and greater light exposure at night (Adan et al., 2012; Randler & Rahafar, 2017). Coles, Wirshba, Nota, Schubert, and Grunthal (2018) recently presented a correlational analysis suggesting a linear increase in the prevalence of Obsessive-Compulsive Disorder (OCD) as a function of increasing latitude. The authors suggest that delayed bedtime and subsequent misalignment of the circadian system is related to increased obsessive tendencies and perseverative thought (Coles, Schubert, & Sharkey, 2012; Hewig, Hagemann, Seifert, Naumann, & Bartussek, 2005; Nota & Coles, 2015; Reinberg, Ashkenazi, & Smolensky, 2007; Schubert & Coles, 2013, 2015; Stokkan & Reiter, 1994). Their findings are compelling and fit into a broader framework linking OCD to sleep difficulties (Coles et al., 2012; Nota, Sharkey, & Coles, 2015; Schubert & Coles, 2013). For instance, individuals with OCD have shorter sleep duration and a higher prevalence of Delayed Sleep Phase Disorder (DSPD; Cox & Olatunji, 2016; Nota et al., 2015). Elevated OCD symptoms are also found in community participants with DSPD (Schubert & Coles, 2015). As natural light exposure varies based on latitude, and as higher latitudes are associated with eveningness (Randler & Rahafar, 2017), it is understandable to draw a connection between latitude and OCD prevalence.

However, it struck us that Coles et al. (2018) conducted their analyses without using standard meta-analytic techniques, which would have accounted for the imprecision of the included estimates. This allows for the potential that their findings are contaminated by small-sample effects. Inspection of their figure also implies a non-linear relation between prevalence and latitude that might qualify their claims. The present manuscript reconsiders their findings using the data reported in their original study and validates them against data coded from a larger, independent meta-analysis of the global prevalence of OCD, conducted by Fawcett, Power, & Fawcett (2020). Although the primary focus of the latter meta-analysis was in estimating the global prevalence of OCD (and evaluating gender differences therein), it was nonetheless possible to code the appropriate latitude for each study after the fact, permitting comparison to Coles et al. (2018). The decision to validate their findings against an independent sample of studies was driven by the belief that a reliable association should be resilient to variation in the particular search or coding parameters employed by a given team. Further, the comparison of two independently coded samples of studies provides an important opportunity to identify coding errors or omissions in either analysis.

1. Method

We first verified the results of Coles et al. (2018) using the data reported in Table 2 of that study. There was one exception: The effect

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Table 1
Bayes factors comparing each model as a function of data set (Coles et al. Fawcett et al.) and dependent measure (Latitude, Max. Daylight Hours). Bayes factors are provided in favour of the supported model, with the supported model named in parentheses below.

Dependent Measure	Null/Linear	Null/Non-Linear	Linear/Non-Linear
Coles et al. (2018)			
Latitude	4.6 *(Linear)*	182.0 *(Non-Linear)*	41.1 *(Non-Linear)*
Max. Daylight Hours	12.4 *(Linear)*	1051.5 *(Non-Linear)*	81.5 *(Non-Linear)*
Fawcett et al. (2020)			
Latitude	5.4 *(Null)*	3.0 *(Null)*	1.7 *(Non-Linear)*
Max. Daylight Hours	5.6 *(Null)*	1.7 *(Null)*	3.2 *(Non-Linear)*

reported for Sahin (1997) was excluded because we were unable to procure the source to verify its sample size.[1] This resulted in 23 estimates.

We followed these models with comparable analyses based on an independent meta-analysis that aimed to provide a worldwide estimate of the prevalence of OCD in men and women (Fawcett et al., 2020). In the interest of expediency, we direct interested readers to that source – but provide a brief synopsis of our search and coding protocol here. A search of the online resources *PsycINFO* and *PubMed* was conducted using a Boolean search phrase with the keywords *obsessive compulsive disorder*, *epidemiology*, and *prevalence*. Studies were included if they assessed a representative community sample with a mean age of 18+, used a diagnostic interview according to DSM or ICD criteria, and reported current, period, or lifetimes estimates of OCD. Studies were excluded if diagnoses were based on retrospective chart reviews, clinical records or self-report surveys, if they reported non-community samples or targeted special populations, or if respondents were from restricted age ranges or were all over the age of 65. Although gender and measurement window were recorded, they were orthogonal to the current research question, so we used only the lifetime prevalence estimates from the mixed samples (those not differentiating between male and female) reported by Fawcett et al. (2020). This resulted in 30 estimates.

In line with Coles et al. (2018), latitude for each study was extracted from the World Atlas webpage (www.worldatlas.com), calculated as absolute degrees from the equator. The specific city or region was used to generate latitude wherever possible, with the capital city of the relevant country used otherwise.

2. Quality ratings

To assess study quality within each analysis, the fourth author generated a 10-point checklist by combining two standardized assessment tools (Giannakopoulos, Rammelsberg, Eberhard, & Schmitter, 2012; Knight et al., 2012). The resulting items are summarized below:

1. Was the target population clearly defined and were demographic characteristics provided? (e.g., age, sex, ethnicity, income; not reported/only one of the above listed = 0, two or more of the above listed = 1)
2. Were the eligibility criteria clearly specified? (neither specified = 0, inclusion or exclusion criteria specified = 1)
3. Was either of the following ascertainment methods used? (must be one or the other)
   I. Probability sampling OR
   II. Entire population surveyed
   (unclear/no, or convenience sample used = 0, yes = 1).
4. Was the response rate adequate? (e.g., below 70%/not reported = 0, 70% or higher = 1)
5. Was information included about people who completed the study versus those who refused? For instance, did they differ on any demographic variables? (no/not reported = 0, yes = 1)
6. Was the sample representative of the target population? (no/unclear = 0, yes = 1)
7. Were data collection methods standardized? (no/unclear = 0, yes = 1)
8. Were validated criteria used to assess for the presence/absence of disorder? (e.g., validated scale or diagnostic tool; no/unclear = 0, yes = 1)
9. Who administered the diagnostic interview? (trained lay person/not reported = 0, trained clinician/researcher/allied mental health worker or trainee = 1)
10. Were the estimates of prevalence given with confidence intervals or standard errors (not reported = 0, reported = 1)

As depicted in Table 2, the overall mean quality rating was 7.90 (SD = 1.18).[2] The mean quality rating for the 19 studies included by Coles et al. was 7.58 (SD = 1.26), which was numerically lower than the mean quality rating for the 24 studies included by Fawcett et al., which was 8.13 (SD = 1.04). No statistical comparison was made of these figures because the two samples were not independent. Nonetheless, the mean quality rating for the 5 studies included only by Coles et al. was 6.8 (SD = 1.3), which was lower than the mean quality rating for the 6 studies included only by Fawcett et al. which was 8.5 (SD = 0.7; difference = 1.7, $CI_{95\%}$ [0.5, 2.8]). This difference was driven by two studies included only by Coles et al. that had the lowest quality ratings of the evaluated studies (i.e., 5 and 6). These studies were also influential within the linear model, owing to the fact that they reflected one of the lowest prevalence estimates and lowest latitudes (0.3% at latitude 7.3° for Beyero et al., 2004) and one of the highest prevalence estimates and highest latitudes (3.5% at latitude 47.2° for Angst et al., 2004) within that sample. Excluding these studies from our re-analysis of Coles et al. causes the linear relation described below to become only marginally supported.

3. Analytic plan

Our analyses used fully-Bayesian multilevel binomial regression models implemented using *brms 2.9.0* (Bürkner, 2017, 2018) within *R 3.5.2* (R Core Team, 2013). Effect size calculations were unnecessary as the models themselves were fit using the sample size and the number of individuals with OCD in that sample. Prevalence was estimated within the model as logit-transformed proportions (Cooper, Hedges, & Valentine, 2009), but has been back-transformed and reported as a percentage for ease of interpretation. Each model incorporated random intercepts accounting for variability across samples. Bayes factors were calculated for each analysis evaluating evidence for inclusion of the relevant variable, although we adopt a holistic view and have reported predictions from the model incorporating the relevant moderator alongside those values. Models were fit and evaluated for convergence using standard metrics (e.g., R-hat < 1.01; Gelman & Hill, 2007) and practices. For the sake of exposition, we direct interested readers to other sources for information on our modelling approach (e.g., Fawcett et al., 2020; Fawcett, Fairbrother, Fawcett, & White, 2018; Fawcett, Fairbrother, Cox, White, & Fawcett, 2019; see also, Fawcett, Lawrence,

[1] To determine whether this exclusion impacted our conclusions, we conducted the relevant analysis two additional times – once including Sahin (1997) and assuming a sample size equal to the smallest of the remaining studies (N = 483; Wittchen, Essau, Von Zerssen, Krieg, & Zaudig, 1992) and again assuming a sample size equal to the largest of the remaining studies (N = 25,180; Mohammadi et al., 2004). Both models produced results similar to those reported in-text.

[2] Sahin (1997) was excluded from the following calculations because we were unable evaluate it.

Table 2
Quality ratings for each study provided on a 10-point scale (see text for details). A checkmark or "X" indicates the study was included or excluded, respectively, from the analysis corresponding to the relevant column. Comments are provided explaining coding decision, as appropriate.

First Author	Year	Coles	Fawcett	Both	Comments	Quality Score
Abou-Saleh	2001	✓	✓	✓		9
Alhasnawi	2009	✓	✓	✓		7
Andrade	2002	✓	✓	✓		8
Angst	2004	✓	✗	✗	Fawcett et al. excluded this study because it only required 1 of 9 criterial OC-symptoms for Criteria A of the SPIKE Interview, which does not conform to DSM-IV OCD criteria as stated.	5
Beyero	2004	✓	✗	✗	Fawcett et al. excluded this study both because the population under consideration was a semi-nomadic community not comparable in our view to the other included samples, and also because the sample included adolescents.	6
Bijl	1998	✗	✓	✗		8
Bland	1988	✓	✓	✓		8
Canino	1987	✗	✓	✗		9
Caraveo-Anduaga	2004	✓	✓	✓		6
Chen	1993	✗	✓	✗		8
Cho	2007	✗	✓	✗		10
Cho	2010	✗	✓	✗		8
Chong	2012	✗	✓	✗		9
El-Wasify	2011	✓	✗	✗	Fawcett et al. excluded this study because it included adolescents.	8
Grabe	2000	✗	✓	✗		8
Gureje	2006	✓	✓	✓		8
Hwu	1989	✓	✓	✓	Fawcett et al. coded this study as a series of separate, regional estimates to remain consistent with Karno et al. See text for details.	8
Karno	1988	✓	✓	✓		7
Keqing	2008	✓	✓	✓		9
Kringlen	2001	✓	✓	✓		9
Lee	1990	✗	✓	✗		8
Mohammadi	2004	✓	✓	✓		8
Ruscio	2010	✓	✓	✓		6
Sahin	1997	✓	✗	✗	Fawcett et al. excluded this study because we were unable to access the original article and did not have the appropriate sample size.	–
Stein	1997	✓	✗	✗	This study reports only current prevalence and is therefore ineligible for inclusion in either analysis. We speculate that Coles et al. miscoded the current estimate as a lifetime estimate.	7
Szádóczky	1998	✓	✓	✓		7
Vicente	2006	✓	✗	✗	Fawcett et al. excluded this study because it included adolescents.	8
Wells	1989	✗	✓	✗		8
Williams	2010	✗	✓	✗		9
Wittchen	1992	✓	✓	✓		10

& Taylor, 2016; Fawcett & Ozubko, 2016).

One benefit of using a Bayesian framework – beyond drawing conclusions pertaining to null findings – is the ability to supplement the model with mildly informative expert knowledge. To this end, our prior expectations relating to the intercept of each model assumed that the average prevalence in a typical sample should range somewhere between 0.6% and 27.0%. We further assumed that the logit-transformed slopes and standard deviations pertaining to random effects would most likely range between −2 and 2 on the logit scale; this broadly reflects the belief that after accounting for variability across samples the "true" prevalence within any given sample might vary anywhere from < 0.1% to 73.0%. Values outside this range remained possible, albeit unlikely.

For each analysis, we fit three models. The first was an intercept model including no slope for latitude. This model provided a base estimate of the prevalence of OCD and served as a baseline against which to compare our remaining models. We next fit a linear model representing a meta-analytic analog to the analysis reported by Coles et al. (2018). This should reveal evidence for any linear effect of latitude on the prevalence estimates after accounting for the imprecision of those estimates. Our final model used thin-plate regression splines (Wood, 2003) to fit a non-linear relation between prevalence and latitude. This model includes a linear term in addition to the non-linear term. Importantly, the "smoothness" of the resulting non-linear component is estimated like any other parameter and would reduce to a linear relation should there be no evidence of non-linearity.

In addition to interpreting the parameters produced by these models, we also quantified the evidence for each using Bayes Factors (BF; Kass & Raftery, 1995). BFs are calculated as a ratio of the marginal likelihoods of two competing models – most often a model containing a particular predictor and a model excluding that predictor. The BF itself reflects the relative support for one model over the other, such that a BF of 5 in favour of a particular model would mean that the model was 5 times more likely than the competing model. For reference, BFs greater than 3 and 10 reflect weak and strong evidence (respectively) in favour of the relevant model, whereas values between 1 and 3 reflect very weak or equivocal evidence (Jeffreys, 1961; Kass & Raftery, 1995).[3] Importantly, whereas many common statistical approaches are capable only of providing evidence in *favour* of a particular statistical hypothesis, BFs (and other Bayesian techniques) are also capable of providing evidence *against* a particular statistical hypothesis.

Latitude was standardized prior to analysis, although this transformation has been reversed in our figures to ease interpretation. We opted to focus our analyses on latitude – but models applied to maximum daylight hours produced similar results in each case and are summarized in Table 1.

4. Results

Re-analysis of Coles et al. As depicted in the upper panel of Fig. 1 and summarized in Table 1, there is compelling evidence of a non-linear relation between OCD prevalence and latitude within this sample that is supported over both the intercept-only model (BF = 182.01) and the linear model (BF = 41.07) – which itself is weakly supported over the intercept-only model (BF = 4.60). From Fig. 1, it appears that OCD

[3] For simplicity, we report all Bayes factors in terms of evidence *in support of* a particular model – be it the full model or the reduced model. This ensures that the reported values are always greater than 1.

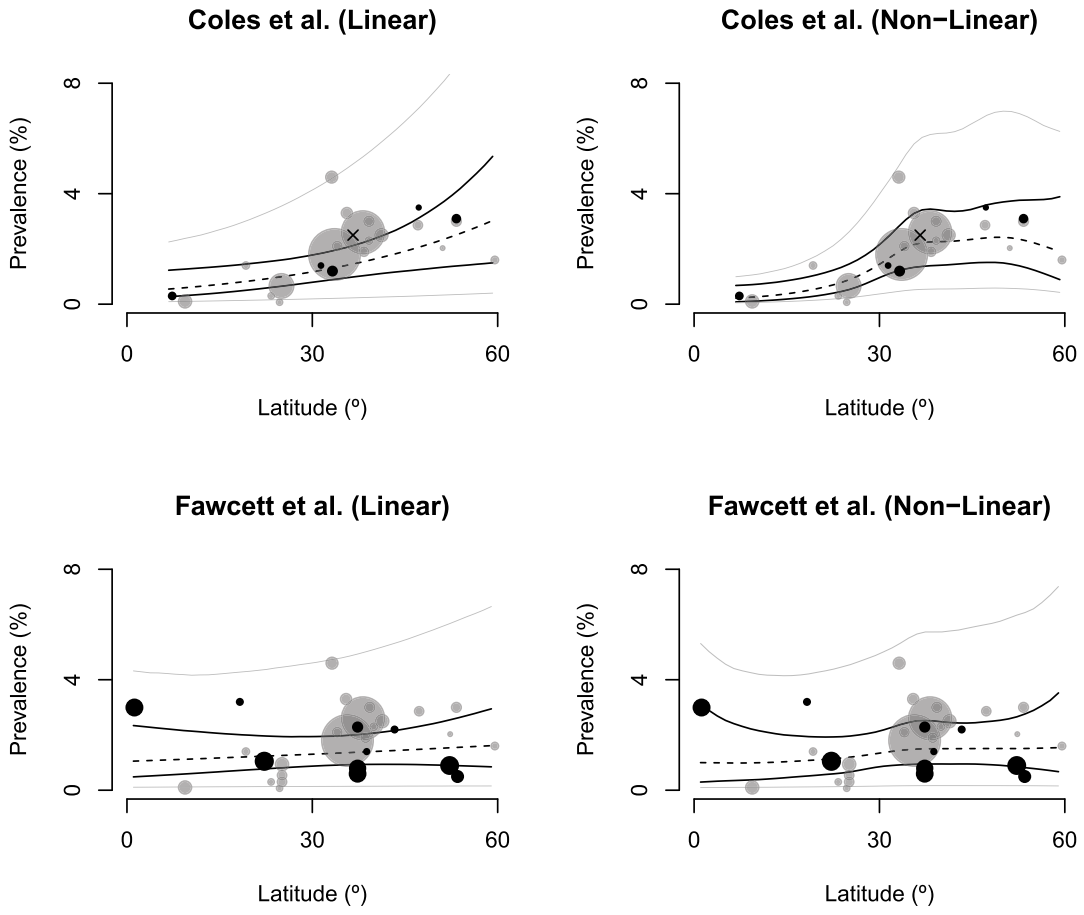

Fig. 1. Prevalence estimates (%) of Obsessive-Compulsive Disorder (OCD) as a function of Latitude (°) plotted separately for the linear and non-linear models based on data from Coles et al. (2018) or Fawcett et al. (2020). The dotted and solid lines reflect the predictions and 95% confidence intervals from the relevant model, respectively. The thin, outer lines reflect 95% prediction intervals, representing the "true" prevalence expected from a new, hypothetical sample similar to those included in the analysis; prediction intervals reflect the degree of between-study heterogeneity in the sample (Higgins, Thompson, & Spiegelhalter, 2009). Marker size is scaled to reflect the relative sample size of the relevant study. Grey circles reflect studies common to all analyses; black circles reflect studies specific to that particular analysis. Sahin (1997) is marked with an 'X'.
(Figure as File Appended During Submission and Should Appear Below Manuscript).

prevalence increases in a semi-linear fashion until a latitude of between 35° and 40° after which the prevalence remains constant. That said, model estimates within the upper range of the current latitudes are quite imprecise and therefore caution should be taken in interpreting this figure as more data are required to draw firm conclusions. Heterogeneity – as depicted in Fig. 1 by prediction intervals – was also estimated to be quite high, suggesting that the "true" underlying prevalence varied substantially between samples even after accounting for latitude.

To further evaluate the nature of any linear relation, we also investigated posterior samples for the slope of the linear model. The slope (measured in log-odds) was 0.43, $CI_{95\%}$ [0.11, 0.71], with 99.4% of the posterior samples greater than 0 (corresponding to a Bayesian *p*-value of .006).

Re-analysis of Fawcett et al. Having replicated and extended the findings of Coles et al. (2018), we next validated them against Fawcett et al. (2020). This sample included a different array of studies, discussed at length in the following section. Again, inclusion decisions had been made independently – without considering latitude as a predictor.

As depicted in the lower panel of Fig. 1 and summarized in Table 1, there is weak evidence *against* a non-linear relation between OCD prevalence and latitude within this sample. The intercept-only model was supported over both the linear model ($BF = 5.37$) and the non-linear model ($BF = 3.03$), with very weak evidence favouring the non-linear model over the linear model ($BF = 1.73$). These results suggest no evidence of a relation between latitude and OCD prevalence, although it is perhaps worth noting (with suitable caution) that the non-linear model is still a better fit than the linear model, even though it has little explanatory power. As in the preceding model, heterogeneity was quite high, suggesting variability in the "true" prevalence of OCD across samples (for further analyses of heterogeneity in these studies, see Fawcett et al., 2020).

To further evaluate the nature of any linear relation, we also investigated posterior samples for the slope from the linear model. The slope (measured in log-odds) was 0.10, $CI_{95\%}$ [-0.19, 0.37], with only 76.1% of the posterior samples greater than 0 (corresponding to a Bayesian *p*-value of .239).

Post-hoc Power Analyses. So long as BFs are interpreted as a continuous metric of evidence supporting one model over another, the concept of statistical power does not apply in the same way it does to Frequentist approaches (e.g., Kruschke & Liddell, 2018; Rouder, 2014). However, we nonetheless recognize the value in understanding the probability of identifying compelling evidence favoring a linear relation should one be present in the data; this is particularly true given that our

re-analysis of Fawcett et al. (2020) found evidence *against* such a relation.[4] For that reason, we undertook a series of simulations, each generating a sample of studies with latitudes and sample sizes based on either Coles et al. (2018) or Fawcett et al. (2020), but with prevalence estimates generated from a population with an average prevalence of 1.4% and a linear relation between prevalence and latitude that was small, medium or large based on the norms provided in Table 1 of Chen, Cohen and Chen (2010; we assumed a P_0 of 0.02, corresponding roughly to the predicted prevalence in the current model for a study at the average latitude). Between-study heterogeneity was set using the estimate from our re-analysis of Coles et al. We simulated data in this manner 500 times for each of Coles et al. (2018) and Fawcett et al. (2020), fitting a separate meta-analysis to each simulated sample. We then calculated the proportion of these simulated meta-analyses (a) providing at least moderate evidence (BF $>$ 3) favoring a linear relation; and, (b) for which at least 95% of the posterior samples for the slope in our linear model were greater than 0 (corresponding to a Bayesian *p*-value of $<$ .05).

According to our simulations, the statistical power for detecting a large or medium linear relation between prevalence and latitude was > 99% using either BFs or Bayesian *p*-values for either set of studies. For reference, the correlation reported by Coles et al. (r = .64) would be considered large. However, if a small effect size were assumed – power was reduced to 74% and 93% for Coles et al. and to 78% and 95% for Fawcett et al. based on BFs or Bayesian *p*-values, respectively. For reference, the odds ratio reported in our re-analysis of Coles et al. (OR = 1.54) and Fawcett et al. (OR = 1.10) would both be considered small. In summary, the present analyses are likely to be adequately powered to detect relations of small-to-moderate magnitude but may be underpowered for particularly small effects.

Importantly, the probability of any simulation finding even weak evidence *against* a linear relation (defined as a BF $>$ 3 in favour of the *Null* when a linear relation was actually present) was rare ($<$ 4%), and a BF of 5 or more in favour of the *Null* (as observed for our re-analysis of Fawcett et al.) occurred in less than 1% of all simulated samples, even assuming a small effect. For that reason, we do not believe the present findings are easily attributed to a lack of statistical power, unless a very small effect were assumed.[5]

5. Discussion

The present analyses extend and challenge the findings reported by Coles et al. (2018). Using their reported data, we observed a non-linear relation between OCD prevalence and latitude. Taken at face value, this implies either a minimal latitude or level of circadian disruption necessary to induce the measured pathology beyond which prevalence does not increase. It could also imply regional customs that change between latitudes of 30° and 40°. Regardless of the explanation, any major theoretical framework seeking to interpret this relation must consider its inherent non-linearity – or better characterize its function. However, such efforts are frustrated by the fact that an independent sample of studies failed to support any (linear or non-linear) relation. Given differences observed between these analyses emerge due to differences in the included studies, it is important to consider those studies that differ between our models.

As summarized in Table 2, Sahin (1997) was excluded from Fawcett et al. (2020) on the account that we were unable to access their article. We further excluded Beyero et al. (2004), El-Wasify et al. (2011), and Vincente et al. (2006) on the basis that they included adolescents in their sample; Beyero et al. (2004) also studied a special population (semi-nomadic community) not comparable to our other samples. Coles et al. (2018) used neither age specifiers nor special populations in their exclusionary criteria and thus would not have excluded those studies. Angst et al. (2004) was excluded because their measure (the SPIKE Interview) required only one of nine OC-symptoms (e.g., over cautiousness with regard to money), which did not conform to DSM-IV criteria for OCD. Lastly, Stein, Forde, Anderson, and Walker (1997) was excluded due to their only reporting 1-month prevalence (p. 1121), which we believe to be mislabelled as lifetime prevalence by Coles et al. (2018).

Studies unique to Fawcett et al. (2020) include: Bijl, Ravelli, and van Zessen (1998), Canino et al. (1987), Chen et al. (1993), Cho et al. (2007), Cho et al. (2010), Chong et al. (2012), Grabe et al. (2000), Lee et al. (1990), Wells, Bushnell, Hornblow, Joyce, and Oakley-Browne (1989), and Williams et al. (2010). Based on the inclusion/exclusion criteria in Coles et al. (2018), we could see no obvious reason for their exclusion. However, it is possible that some or all of the aforementioned studies did not appear in their search. There was also a slight coding variation for one article. Coles et al. (2018) used an aggregate estimate for Hwu, Yeh, and Chang (1989), likely because the three reported estimates provided only a single region (metropolitan Tapei, small towns and villages); Fawcett et al. (2020) kept these estimates separate, addressing their dependency via random effects.

Notably, many of the additional studies included by Fawcett et al. (2020) reflect moderate sample sizes with relatively low prevalence estimates distributed across the range of observed latitudes. As a result, they counteract the otherwise observed effects. The notable exception is Chong et al. (2012), which is characterized by a moderate prevalence estimate but a low latitude, presenting a point of high statistical leverage. Results do change slightly if this data point were removed: Although weak evidence remains against a linear relation (BF = 3.07), there is now equivocal evidence of a non-linear relation relative to the intercept-only model (BF = 1.40) with a weak preference for a non-linear over a linear relation (BF = 4.50). Importantly, there is no compelling reason to exclude this study.

The potential relation between latitude and OCD symptomology is an important consideration with converging support from studies demonstrating a link between latitude and circadian rhythmic expressions (Adan et al., 2012; Randler & Rahafar, 2017) and an association between circadian rhythm irregularities and OCD (Coles et al., 2012; Cox, Tuck, & Olatunji, 2018; Drummond et al., 2012; Monteleone, Natale, Fuschino, & Maj, 1997; Nota et al., 2015; Schubert & Coles, 2013, 2015; Turner et al., 2007). Furthermore, there is evidence from case studies that phototherapy, which can phase-shift human circadian rhythms (Rosenthal et al., 1990), may improve OCD symptoms. Case studies of

[4] We thank an anonymous reviewer for recommending the inclusion of the power analyses.

[5] Another concern is that precise locations were unavailable for some samples, resulting in the use of capital cities for the purpose of calculating latitude. While necessary, this sort of substitution could obscure a statistical relation. To evaluate support for this concern in the present data set, the power simulations were repeated with the exception that following data generation – but preceding analysis – the latitude for each sample having used the capital city was replaced by a random latitude sampled uniformly from between the minimum and maximum possible latitudes for that country. This process would emulate a similar (if not greater) degree of measurement error as observed in the analyzed data. Simulations were undertaken only for the re-analysis of Fawcett, Power, & Fawcett, 2020. Results were largely unchanged. Statistical power for detecting a large or medium linear relation remained > 99% using either BFs or Bayesian *p*-values. However, statistical power to detect a small effect was reduced to 74% and 93% based on BFs or Bayesian *p*-values, respectively. The probability of finding evidence *against* a linear relation remained rare, with BFs of 3 or more and 5 or more occurring in 3% and 1% of the simulated samples when assuming a small effect, respectively. The diminutive nature of this reduction in statistical power owes largely to the fact that the median range of possible latitudes (maximum – minimum) for a given sample subtended only 4.2°. In short, it is unlikely that measurement error arising from the use of capital cities obscured a linear relation in the data reported by Fawcett, Power, & Fawcett, 2020. Further, given that this issue was common to both data sets, it is unclear why it would mask such a relation in one, but not the other.

individuals with seasonal forms of OCD (e.g., occurring exclusively in autumn or winter) or with comorbid Seasonal Affective Disorder (SAD), show a reduction or complete remission when treated with phototherapy (Brinkhuijsen, Koenegracht, & Meesters, 2003; Hoflich, Kasper, & Moller, 1992; Sinha, Bakhla, Patnaik, & Chaudhury, 2014; although see; Yoney & Pigott, 1991). Seasonal variation in OCD symptoms has been demonstrated across both youth and adult populations, with significantly lower symptom counts between August and October (Kovalenko et al., 2000) and increased OCD prevalence in autumn compared to summer (De Graaf, Van Dorsselaer, Ten Have, Schoemaker, & Vollebergh, 2005). Reduced OCD symptomatology in the summer months may be reflective of annual variation in serotonin levels, with bright light increasing serotonin (Kovalenko et al., 2000).

In a review of research published within the last 5 years on the relation between circadian rhythms and OCD, Cox and Olatunji (2019) found the link to be inconsistent. For instance, Kani et al. (2018) observed only a marginal difference in the frequency of the eveningness chronotype in OCD patients under treatment versus healthy controls and no significant relation between chronotype and OCD symptom severity. In adolescents, baseline chronotype was not predictive of OCD symptoms at follow-up (Alvaro, Roberts, Harris, & Bruni, 2017). Cox and Olatunji (2019) highlight several important recommendations for future research in this area to help clarify the relation between circadian rhythms and OCD, including increased use of prospective designs, clinical OCD populations, and multiple indicators of circadian rhythms versus single-method approaches, as well as decreased reliance on undergraduate samples and self-report measures of circadian rhythms. Our findings add to this existing literature suggesting that current evidence of a relationship between OCD prevalence and latitude is weak, and that further research is required before firm conclusions can be drawn from the extant epidemiological data.

CRediT authorship contribution statement

Jonathan M. Fawcett: Conceptualization, Data curation, Formal analysis, Methodology, Visualization, Writing - original draft, Writing - review & editing. **Rachelle M. Wakeham-Lewis:** Data curation, Writing - original draft, Writing - review & editing. **Sheila Garland:** Writing - original draft, Writing - review & editing. **Emily J. Fawcett:** Conceptualization, Data curation, Project administration, Writing - original draft, Writing - review & editing.

Declaration of competing interest

None.

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