Two Longterm Studies of Seasonal Variation in Depressive Symptoms among Community Participants

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Abstract

Background—There is evidence that seasonal variation in depressive symptoms is common in the population. However, research is limited by a reliance on longterm retrospective methods.

Methods—Seasonal patterns were tested in two samples of community participants recruited in separate prospective studies in the Midwestern ($n = 556$ males/females) and Pacific Northwestern ($n = 206$ males) United States. Participants completed self-report measures of depressive symptoms 10–19 times from ages 14–36 years ($n = 8,316$ person observations). These data were compared with local meteorological conditions (e.g., solar radiation) recorded across the 2 weeks prior to each self-report.

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Results—In within-subjects analyses, participants' depressive symptoms and the probability of clinically significant symptoms varied with the time of year, as hypothesized (highest in the weeks of early Winter; lowest in early Fall). However, effects sizes were modest and were not explained by recent sunlight or other meteorological conditions.

Limitations—Samples were not nationally representative. Participants did not complete retrospective reports of seasonal depression or measures of current vegetative symptoms.

Conclusions—Neither time of the year or recent seasonally linked meteorological conditions were powerful influences on depressive symptoms experienced by community populations in relevant geographic regions. Prior studies may have overestimated the prevalence and significance of seasonal variation in depressive symptoms for the general population.

Keywords

adolescence; community; longitudinal; seasonal depression

Introduction

Major depressive disorder with a seasonal pattern, or seasonal affective disorder (SAD), is defined by a pattern of onsets and remissions of episodes at characteristic times of the year (American Psychiatric Association, 2000). Based on stringent criteria in a nationally representative sample, estimated lifetime prevalence of major depression with a seasonal pattern is 0.4%, and 1.0% for minor and major seasonal depression combined (Blazer et al., 1998). Community studies using self-report instruments have suggested many if not most people experience some degree of seasonal changes in mood and behavior—or "seasonality" (Magnusson, 2000). For example, in one large random sample of adults, 92% reported noticing such seasonal changes, 27% reported problematic seasonal changes, and 4–10% reported a degree of seasonal change and impairment interpreted as characteristic of SAD (Kasper et al., 1989).

The evidence base on SAD and seasonality is seriously limited by a reliance on face-valid self-report methodology and longterm retrospection. That is, participants are asked to retrospect across years of their lives to characterize past emotional and behavioral states and identify temporal patterns in the duration and timing of these states. The validity of even much simpler acts of recollection is widely questioned, and retrospection regarding emotional states is especially prone to error (Henry et al., 1994; Rogler et al., 1992). Indeed, several studies cast doubt on the validity of self-reported seasonality. For example, although seasonality is stable by definition (i.e., a pattern requires 2 years), self-reports of trait seasonality are associated with current weather conditions, neuroticism, and generalized tendencies to attribute distress to external causes (Jang et al., 1997; Murray et al., 1995; Sigmon et al., 2009).

Prospective, repeated measures studies are relatively rare but are well suited to unbiased identification of seasonal patterns in depressive symptoms. For example, nonpatients showed expected symptom patterns when assessed repeatedly during a single year (Harmatz et al., 2000). In contrast, a 3-year prospective community study found decreased behavioral engagement in winter compared to summer but no change in negative affect (Murray et al., 2001). Shorter-term studies using meteorological records found nonpatients' daily moods varied not only with day length or sunlight intensity—consistent with putative chronobiological mechanisms of SAD (Rohan et al., 2009a)—but also independently with temperature, precipitation, atmospheric pressure, and/or wind (Denissen et al., 2008; Keller et al., 2005; Klimstra et al., 2011).

Several design innovations are needed to further advance the evidence base on seasonality in depressive symptoms. First, cohorts must be followed for several years so that the long-term patterns implied by the notion of seasonality can be identified. Second, to establish clinical significance, studies should use outcomes with known clinical metrics (vs. daily mood). Third, precise measures of time of year are preferable over the crude and potentially arbitrary (with respect to biological mechanisms) scale points of Fall, Winter, Spring, and Summer. Fourth, measurement of meteorological conditions will clarify the extent to which seasonality shares common environmental influences with SAD, as is assumed. Finally, studies of adolescents are surprisingly rare (e.g., Nillni et al., 2009).

Thus, we examined the extent to which depressive symptoms covaried with day of year and coincident meteorological conditions in two long-term longitudinal cohort studies. Participants were recruited from geographic regions that show a range of meteorological conditions with clear seasonal patterns. Additionally, at the time of data collection, participants and investigators were blind to the present research focus, thus minimizing several forms of bias.

We hypothesized that participants levels of depressive symptoms and probabilities of clinically significant symptoms would show a seasonal pattern (winter > fall = spring > summer months), and that this effect largely would be explained by negative associations between solar radiation and depressive symptoms. We also explored whether other meteorological variables would further explain any seasonal patterns; generally, we expected depressive symptoms would be associated positively with precipitation, wind, humidity, and atmospheric pressure, and negatively with temperature. Finally, we explored whether symptom seasonality would differ by gender or developmental period (adolescence vs. early adulthood).

Method

Participants

Oregon Youth Study (OYS; Capaldi and Patterson, 1989)—Boys ($n = 206$) were recruited in entire fourth-grade classrooms in 1984–1985 from six schools in neighborhoods with higher-than-average delinquency rates. Families were representative of the medium-sized metropolitan area: 90% of boys were White, and most families were classified as low socioeconomic status (SES). Participants' depressive symptoms were assessed annually up to 19 times from ages 14–36 years ($n=3,476$ person observations). Annual participation rates have been 90% or more of living participants (three are deceased).

Family Transitions Project (FTP; Conger and Elder, 1994)—Boys and girls ($n = 559$) were recruited during Grades 7 or 9 from eight rural counties in Iowa in 1989 and 1991, and came from two- (81%) or single-parent (19%) families. Families were primarily lower-middle SES or middle SES, and 99% were White (minority families comprised only 1% of this rural area's population). A total of 263 boys and 293 girls ($n = 556$) completed measures of depressive symptoms at up to 10 waves (77–94% of the sample per wave) between ages 14–33 years ($n = 4,840$ person observations). Retention of living participants (10 are deceased) at the final wave of the study was 89%.

Both the OYS and FTP were approved by the Institutional Review Boards at the original sites, and written informed consent was obtained from all participants and legal guardians when relevant.

Measures

Dates of self-reports and participant locations—Interviews (dated) and mailing records were used to establish participant residence when they self-reported depressive symptoms. Zip codes and international city locations were used to extract latitude and meteorological data.

Depressive symptoms—OYS participants completed the 20-item Center for Epidemiologic Studies Depression (CES-D) scale (Radloff, 1977). Participants used a 4-point scale (*rarely or none of the time [0–1 day]* coded '0' to *most or all of the time [5–7 days]* coded '3') to indicate how often they felt symptoms (e.g., sadness, poor appetite, poor sleep) in the past week. Binary 'caseness' variables (no '0' and yes '1') also were created at each assessment using clinical cut-off scores of 22 for adolescent males (age < 20 years) and 16 for adults (Lewinsohn et al., 1998; Radloff, 1977).

FTP participants completed the 13-item depression scale of the Symptom Checklist 90 (SCL-90; these items are identical to those on the SCL-90-R [Derogatis, 1994]). Participants used a 5-point scale (*not at all* '0' to *extremely* '4') to indicate how much each problem (e.g., crying, low energy) bothered them in the past week; the scale does not include items on sleep or appetite. Binary caseness variables at each assessment were created based on t scores (90th percentile; derived from SCL-90-R nonpatient norms [Derogatis, 1994]) of 63 or greater; for male and female adolescents (age < 20 years) and male and female adults, cut-offs were 1.39, 2.03, 0.70, and 1.11, respectively. In both samples, depressive symptoms were log transformed to reduce skewness.

Age and age group—Exact age at each assessment was calculated from birth and assessment dates. For analyses of caseness, an age-group variable recorded whether each observation used the adolescent '0' or adult '1' cutoff, in order to control for the otherwise artificial shift in rates of caseness that would be modeled at age 20 years when cutoff scores change.

Day of year—We expected the association between day of year and depressive symptoms to follow a cubic function, given the periodicity of the seasonal time scale.[1] Thus, dates of self-reported depressive symptoms were recoded to a scale of 0 to 364 centered at the Winter Solstice (i.e., December 22 of 1 year '0' to December 21 of the next '364'). Day of year then was divided by 100 so that this value, its square, and its cube—which were used to model linear, quadratic, and cubic effects, respectively—were on similar scales.

Meteorological conditions—Residence location and dates of self-reported depressive symptoms were used to identify and extract local recordings of coincident sunlight intensity (surface downward solar radiation; watts/meter^2), precipitation (kilograms/meter^2), temperature (Kelvin), atmospheric pressure (pascals), wind speed (meters/second), and specific humidity (kg of water vapor/kg of total air) from the North American Land Data Assimilation System (NLDAS) project (Mitchell et al., 2004). These data are derived from surface weather stations, ground-based radar, and other measurements. The product is hourly and gridded at 0.125° resolution (~13 km^2) for the entire continental U.S. Downward solar radiation provides an integrated measure of sunlight that accounts for time of day, time of year, and cloud cover. Day of the year, solar radiation, and day length are functionally

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[1]Relations between day of the year and solar phenomena (day length and radiation) are sinusoidal. However, our data indicate that polynomial functions for day of the year explains essentially all variation, R^2 = .993, in the sin function for day of the year and are more interpretable and amenable to statistical modeling. The polynomial function also does not assume that the relations between day of the year and depressive symptoms are identical to the invariant sinusoidal association between day of the year and solar phenomena; for example, Fall/Winter increases in depressive symptoms do not have to be symmetrical to Spring/Summer decreases.

related, and too highly correlated to be considered simultaneously in study models.[2] Therefore, solar radiation was used to index both day length and local deviations (e.g., due to cloud cover) from invariant annual astronomical cycles. Daytime recordings of each meteorological condition were averaged across the 14-day period ending with the participants' self-reported depressive symptoms (consistent with Molin et al., 1996); slope scores also were recorded to reflect 14-day linear changes in conditions. Meteorological conditions for participants residing outside the continental U.S. (< 1% of observations) were derived from National Centers for Environmental Prediction (NCEP) North American Regional Reanalysis (NARR) data (Mesinger et al., 2006) that are available once every 3 hours for every 0.375° (40 km^2) grid of land, or the NCEP/National Center for Atmospheric Research (NCAR) reanalysis (Kalnay et al., 1996) that provides global coverage 4 times daily at 2.5° resolution. For each time and location, the highest resolution data among these sources was used (> 99% was NLDAS).

Data Analyses

Models were run separately by sample. Given the nested data, the models were run using the xt commands for mixed models in Stata 11. For the transformed continuous measure of depressive symptoms, two-level random coefficient mixed models were used. For the caseness outcome, two-level random intercept logistic mixed models were used. Both models are comparable to their ordinary least squares regression counterpart with the addition of random effects. Age, day-of-year functions (day, day^2 and day^3), and meteorological predictors were all time-varying (Level 1) covariates; gender was a time-invariant (Level 2) covariate in FTP models. The continuous outcome was modeled by first checking for random effects of age and the day-of-year functions in addition to a random intercept; correlations among random effects were assumed to be zero.

Next, we tested 14-day intercepts and slopes for each meteorological variable individually as a predictor of depressive symptoms, controlling for the effects of day of year, age, and gender. We also tested random effects for these variables to allow for possible participant differences in the sensitivity of depressive symptoms to meteorological conditions. Finally, interactions that day-of-year or meteorological variables had with gender or age were tested.

Results

Descriptive Statistics

A range of depressive symptoms was observed in both samples. Consistent with other studies using repeated measures (Rohde et al., 2013), cumulative rates of caseness (clinically significant symptoms) were high (61.2% in OYS males; 38.6% and 40.1% in FTP males and females, respectively; see supplemental table for means and caseness at each assessment). Geographic mobility was limited; 85.6% of OYS observations occurred when participants were living in Oregon or Washington, and 82.4% of FTP observations occurred in Iowa.

The design relied on variation in the times of observation for individuals. Thus, we examined the extent to which annual dates of observation were clustered at specific times of the year for each participant. OYS and FTP participants were observed in a mean (*SD*) of 7.04 (1.39) and 6.01 (1.21) unique months, respectively. In OYS and FTP, 99% and 93% of the samples, respectively, were observed in 4 or more different months; 65% and 31%, respectively, were observed in 7 or more different months.

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[2]E.g. in the Oregon sample solar radiation and day length were correlated, $r = .923$, $p < .001$, and together day of the year and solar radiation explained 97.5% of the variance in day length.

Seasonal Influences on Depressive Symptoms

Table 1 Model 1 shows the results of the primary model in the FTP sample. Model fit was improved when linear effects of age and day of year were random (log likelihood ratio tests $\chi^2[1] = 216.22$ and 7.62, respectively, $p < .01$). Significant effects of gender and age indicated that symptom levels were higher among females and tended to decrease and then level off with age. Significant within-subjects effects of day of year were consistent with the hypothesized cyclical effects of season on depressive symptoms; Panel 1 of Figure 1 shows the expected pattern of highs and lows in the Winter and Summer/Fall months, respectively, although appropriate scaling (Panel 2) highlights the visually imperceptible nature of the association.

In the model for OYS, fit was improved when age and age^2 effects were random ($\chi^2[1] = $ 93.41 and 19.08, respectively, $p < .001$); age effects were similar to those in the FTP model. Although only the linear effects of day of year on depressive symptoms were significant (Table 1 Model 1), coefficient estimates and the curve in Panels 3–4 of Figure 1 followed the same pattern observed in FTP. Models were not improved when latitude was controlled or tested in interaction with day of year.

We then considered whether the within-subjects effects of day of year would differ for males versus females (cross-level interaction) and, in a separate model, whether effects of day of year would differ in adolescence versus adulthood (within-subjects interaction). Neither of these models yielded significant interaction effects; thus, differences in the effects by gender and developmental stage were not supported.

Next, in separate models, we tested whether 14-day level and slope of each meteorological variable would explain effects of day of year, controlling for age and (in FTP only) gender. None of the meteorological predictors were significant (see supplemental table). The effects of day of year remained significant in all of these models except one; when average humidity was modeled in FTP, cubic effects of day of year were not significant. Next, none of the random effects of meteorological variables were significant for the FTP sample. In OYS, random effects were significant only for slopes of temperature, precipitation, and humidity. As a final step in these models, we tested whether individual meteorological predictors were significant when day-of-the-year variables were omitted; none of these models were significant.

Seasonal Influences on Clinically Significant Depressive Symptoms

The models depicted in Table 1 Model 1 were rerun in Model 2 to test seasonal patterns in the probability of caseness. Age, age group, and gender (FTP only) were controlled. The effects of day of year were significant for the FTP sample only; however, the overall models were not significant for either sample. None of the meteorological variables significantly predicted this outcome.

Seasonal Influences among Individuals Vulnerable to Depression

Given the surprisingly modest patterns identified when testing a priori hypotheses, we conducted exploratory analyses to determine whether seasonal influences might be more powerful among individuals vulnerable to depression. To this end, the primary models described above were rerun only for participants who reported clinically significant symptoms at least once during the study. As shown in Table 1, findings from Model 3 regarding day of year were similar to those in Model 1. In this subsample, the addition of meteorological variables did not improve prediction of depressive symptoms.

Discussion

This study circumvented the ways that longterm retrospection may have limited prior research on seasonal depression. As hypothesized, participants reported higher depressive symptoms in Winter than in the late Summer to early Fall. However, this trend was of modest magnitude and of limited clinical significance, given that time of year was not a powerful predictor of the probability of clinically significant symptoms. In contrast, prior studies of predictors of (non-seasonal) depression in these two samples have yielded much stronger effects (e.g., Ge et al., 2006; Stoolmiller et al., 2005).

In the context of adequate statistical power and a replication design, the null effects reported here are noteworthy. First and foremost, we found little evidence that depressive symptoms vary appreciably with either broad or recent seasonal solar radiation trends, as indexed by day-of-year and meteorological records, respectively. This is surprising given that sunlight intensity and day length play central roles in biological models of SAD (Rohan et al., 2009a) and seasonality. Second, depressive symptoms did not vary with recent levels or change in precipitation, temperature, humidity, wind speed, or atmospheric pressure; although OYS participants differed from one another in terms of the sensitivity of depressive symptoms to changes in temperature, precipitation, and humidity. Psychological explanations are typically invoked to explain the effects such meteorological conditions have on daily mood. Our findings do not support that previously observed effects on daily mood extend to the more prolonged severe disruptions tapped by our measures of depressive symptoms. Third, given that solar radiation and other meteorological conditions did not explain effects of day of year on depressive symptoms, other seasonally patterned influences (e.g., psychosocial stress, illness) may explain these effects. Fourth, gender and developmental period did not moderate observed effects.

Taken together, our findings challenge the notion that most or even many individuals show more than modest fluctuations in depressive symptoms as a function of the time of year or in response to meteorological conditions. Findings build on those of a cross-sectional, cross-national study of over 6,500 individuals that found no consistent effect of month or season of administration on Beck Depression Inventory scores (Michalak et al., 2004). Given the dearth of previous long-term prospective studies of seasonality and the methodological weaknesses of prior epidemiological studies, a clear implication of the present findings is that the prevalence of seasonality in depressive symptoms has been overestimated. It remains possible that unobserved subpopulations of individuals with opposing patterns of seasonal sensitivity were masked by our examination of average effects, as suggested by other studies (Klimstra et al., 2011). The significant variation in some of the effects of day of the year on Iowa participants' depressive symptoms (i.e., random effects) hinted at this possibility but was not replicated.

The present study had some limitations. The conclusions would have been enhanced had retrospective self-report measures of seasonality in depressive symptoms and episodes also been collected. Also, our outcome measures did not capture the sleep and appetite symptoms that may be typical of SAD. This is important given that Young and colleagues' (2008) dual-vulnerability model suggests seasonal changes in these vegetative symptoms interact with cognitive vulnerabilities to explain the development of SAD. On the other hand, if seasonality is common in the population, we would have expected to see prominent seasonal variation in the cognitive-affective symptoms that were measured presently. Next, the samples were not nationally representative and were recruited based on some contextual risks for maladjustment. However, there is no reason to suspect participants were atypical with respect to human seasonal responsiveness or biological mechanisms in particular. Finally, participants initially resided in the Midwestern and Northwestern United States and

showed limited geographic mobility. Thus, findings may not generalize to people living at extreme latitudes where SAD is more prevalent. The present study regions, however, are representative of significant portions of the populated world in terms of the seasonal sunlight changes and range of pleasant to decidedly "gloomy" weather conditions that occur. Therefore, findings are highly relevant to the epidemiology of symptom seasonality.

Clinical Implications

The present findings do not contradict research showing that SAD exists or that many individuals experience decreases in positive mood states and behavioral engagement in response to weather and other seasonal changes. Yet, these latter experiences may be distinct from depression. Individuals' tendencies to recollect seasonally patterned depressive symptoms may be influenced by a history of seasonal variation in positive mood states in combination with factors such as current depressive symptoms and public awareness of SAD.

Valid methods of identifying seasonal depression will advance etiological studies, but may not affect treatment recommendations at this time, as there is evidence for the efficacy of cognitive behavioral therapy, light therapy, and antidepressant medication both for seasonal and nonseasonal depression (Even et al., 2008; Lam et al., 2006; Rohan et al., 2009b). Cognitive behavioral therapy stands out for its promise in preventing recurrence of seasonal depression and may be especially appropriate when patients' negative expectations about the effects of the seasons on their mood and behavior are significant.

Supplementary Material

Refer to Web version on PubMed Central for supplementary material.

Acknowledgments

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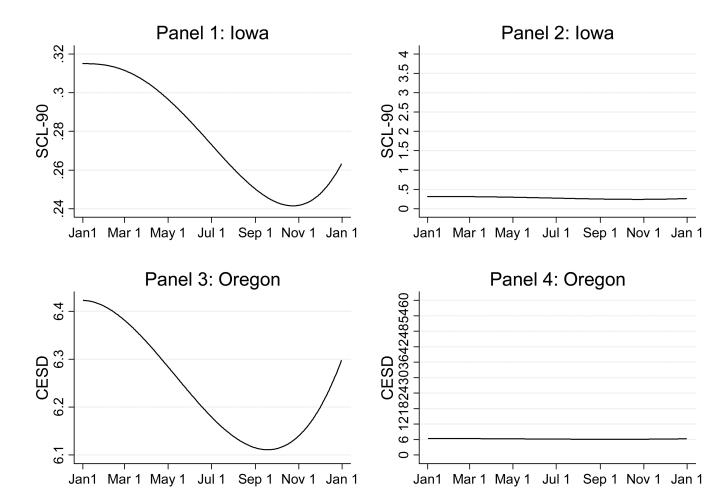
Panel 1: Iowa

Panel 2: Iowa

Panel 3: Oregon

Panel 4: Oregon

Figure 1.
Model estimated depressive symptoms by day of year.

Table 1

Primary Prediction Models for Depressive Symptoms Outcomes in the Oregon and Iowa Subsamples

	Oregon			Iowa		
	B	*SE*	β	*B*	*SE*	β
Model 1: Full Sample Continuous Outcome						
Age	−0.01	0.00 ***	−0.14	−0.02	0.00 ***	−0.01
Age²	0.00	0.00 **	0.04	0.00	0.00 ***	−0.05
Gender	--	--	--	0.41	0.05 ***	−0.64
Day of year	0.07	0.02 **	0.14	0.43	0.11 ***	−0.67
Day²	−0.04	0.02	−0.28	−0.26	0.07 ***	0.97
Day³	0.01	0.01	0.15	0.04	0.01 **	−0.45
Intercept	2.16	0.04 ***		−1.44	0.06 ***	
Likelihood-ratio test	150.32	***		367.43	***	
ICC	.434			.429		
Observations	3476			4840		
N	206			556		

	OR	*SE*	β†	OR	*SE*	β†
Model 2: Full Sample Binary Outcome						
Age group	0.52	0.22 **	0.75	0.52	0.27 *	0.73
Age	1.01	0.01	1.09	1.06	0.03 *	0.96
Age²	1.00	0.00	0.87	0.99	0.00	0.84
Gender	--	--	--	0.95	0.22	0.92
Day of year	1.02	0.08	1.03	6.17	0.66 **	16.87
Day²	0.98	0.12	0.92	0.32	0.42 **	0.01
Day³	1.00	0.03	1.05	1.21	0.08 *	7.26
Likelihood-ratio test	3.10			11.75		
Observations	3476			4840		

		Oregon			Iowa		
		B	SE	β	B	SE	β
	N	206			556		
		B	SE	β	B	SE	B
	Age	−0.01	0.00 **	−0.09	−0.01	0.01 *	−0.01
	Age^2	0.00	0.00	0.02	0.00	0.00 *	−0.05
	Gender	--	--	--	0.32	0.06 ***	−0.49
	Day of year	0.06	0.03 *	0.13	0.58	0.20 **	−0.91
Model 3: Clinical Subsample Continuous Outcome	Day^2	−0.02	0.03	−0.14	−0.35	0.13 **	1.29
	Day^3	0.00	0.01	0.02	0.06	0.02 *	−0.58
	Intercept	2.42	0.04 ***		−0.91	0.10 ***	
	Likelihood-ratio test	72.72	***		159.75	***	
	Observations	2145			1903		
	N	126			212		

Note. B = unstandardized beta coefficient; SE = standard error; $β$ = standardized beta coefficient; OR = odds ratio; ICC = Intraclass correlation coefficient.

[†] coefficients only standardized for the covariate.

[***] $p < .001$.

[**] $p < .01$.

[*] $p < .05$.