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The Longitudinal Association Between Social-Media Use and Depressive Symptoms Among Adolescents and Young Adults: An Empirical Reply to Twenge et al. (2018)

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Abstract
Research by Twenge, Joiner, Rogers, and Martin has indicated that there may be an association between social-media use and depressive symptoms among adolescents. However, because of the cross-sectional nature of this work, the relationship among these variables over time remains unclear. Thus, in this longitudinal study we examined the associations between social-media use and depressive symptoms over time using two samples: 594 adolescents (M_{age} = 12.21) who were surveyed annually for 2 years, and 1,132 undergraduate students (M_{age} = 19.06) who were surveyed annually for 6 years. Results indicate that among both samples, social-media use did not predict depressive symptoms over time for males or females. However, greater depressive symptoms predicted more frequent social-media use only among adolescent girls. Thus, while it is often assumed that social-media use may lead to depressive symptoms, our results indicate that this assumption may be unwarranted.

Keywords
depressive symptoms, social media, longitudinal, bidirectional associations, adolescents, young adults

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Social media has allowed youths to stay connected and socialize with peers in unprecedented ways. Indeed, this dramatic shift in daily communication has garnered considerable interest from researchers, who are eager to identify how social media has affected today's youth. Recent research, particularly, has focused on how social-media use is related to negative outcomes, such as depressive symptoms. Interest in this question has permeated media reports and led to the emergence of terms such as *Facebook depression*, which has been used to warn parents of the dangers associated with social-media use (O'Keeffe & Clarke-Pearson, 2011). Drawing implications from this early research is premature, however. Research on the link between depressive symptoms and social-media use in particular has been primarily cross-sectional, or involving studies with multiple time points that did not allow for an assessment of change over time (e.g., Booker, Kelly, & Sacker, 2018; Nesi & Prinstein, 2015). Thus, these studies have not been able to assess temporal order (i.e., they have not simultaneously assessed whether depressive symptoms predict greater social-media use at a later point in time and/or whether social-media use predicts more depressive symptoms at a later point in time).

Notably, a recent study by Twenge, Joiner, Rogers, and Martin (2018) investigated social-media use and mental health among adolescents using two large-scale, nationally representative samples. Their results demonstrated that between 2010 and 2015, average levels of social-media use, depressive symptoms, and rates of suicide increased, especially among females. Moreover,

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among females, social-media use was positively correlated with depressive symptoms, whereas more in-person interactions were uniquely correlated with lower levels of depressive symptoms over time for both sexes. The authors concluded that social-media use and other-screen activities may be driving increases in poor mental health among adolescents. Of concern, this conclusion has been drawn on as a key explanation for the reported increases in adolescent mental health problems in the United Kingdom, United States, and elsewhere (e.g., see Gunnell, Kidger, & Elvidge, 2018).

These results should be interpreted with caution, however. Twenge and colleagues (2018) were not able to ascertain from their data the direction of effects between social-media use and depressive symptoms because they did not assess the same participants over time. In other words, the concurrent associations that Twenge et al. (2018) found could have been from social-media use leading to higher levels of depressive symptoms over time, higher levels of depressive symptoms leading to higher social-media use over time, a bidirectional effect, or an omitted third variable that influences both social-media use and depressive symptoms. This is a critical limitation to the study.

Twenge and colleagues (2018) claim, however, that some research supports their hypothesis of a causal link from social media to adjustment difficulties. Specifically, they cite a longitudinal study (Shakya & Christakis, 2017) and an experience-sampling-method study (Kross et al., 2013), both of which found that higher frequency of Facebook use at one time point predicted subsequent increases in negative emotions. They also reported results from a short-term experimental study that found that giving up Facebook for a week increased positive emotions (Tromholt, 2016; see also Sagioglou & Greitemeyer, 2014). But, with the exception of Kross et al. (2013)—whose 14-day study supported Twenge et al.'s (2008) view that social-media use leads to reductions in emotional well-being, but not vice versa—these studies tested only one direction of effects (i.e., from social-media use to mental health). Of note, Twenge and colleagues (2018) did not cite results of another experience-sampling-method study similar to Kross et al. (2013), Jelenchick, Eickhoff, and Moreno (2013), that found no across-time associations between Facebook use and depressive symptoms. Other longitudinal studies include Hökby et al. (2016), who found a link between Internet use and negative consequences, but did not find an association between frequency of socializing on the Internet and negative consequences. In addition, Nesi, Miller, and Prinstein (2017) found that depressive symptoms and frequency of social-media use were unrelated, but tested only one direction (i.e., from depressive symptoms to social-media use). Overall, existing longitudinal evidence regarding the relationship between social-media use and depressive symptoms remains inconsistent, and the effect sizes within these studies were generally small.

As Daly (2018) pointed out, another limitation of Twenge et al. (2018) is that their measure of social-media use "failed to assess the frequency of social networking website use beyond 'almost every day,' a very broad category that applied to more than 85% of the female sample in 2015" (p. 1). Thus, Daly suggested that "the potential contribution of the full range of time spent on social media and screen activities (e.g., as measured in hours/minutes per day) to recent trends in depression among adolescent girls remains unknown" (p. 1).

Another concern is the exclusion of covariates in the primary analyses. Twenge et al. (2018) found (for girls only) an association between social media and depressive symptoms even after important covariates were added—including in-person social interaction and demographics (grade, race, and region). In separate analyses, however, they also examined the link between depressive symptoms and other-screen activities (television use), and depressive symptoms and nonscreen activities (print-media use, sports/exercise, religious-service attendance, and paid work), including the same covariates listed above. However, other-screen and nonscreen activities are likely to be correlated with both social-media use and depressive symptoms. A more conservative approach when attempting to uncover unique variance shared by social media and depression may be to consider other-screen and nonscreen activities as covariates (in addition to in-person social interaction).

Given the limitations of Twenge et al.'s (2018) study, the small number of studies that have tested the directions of effects between social-media use and depressive symptoms, and the mixed results of the studies that have assessed bidirectionality, it is clear that this body of literature would benefit from further analysis of current longitudinal data. The present study addressed these limitations through a longitudinal cross-lagged analysis (separately for males and females) of the relation between frequency of daily social-media use and a well-validated measure of depressive symptoms in a recent sample of young adults (six waves between 2010 and 2016) and adolescents (2 waves between 2017 and 2018). We tested both directions of effects (from social-media use to later depressive symptoms and from depressive symptoms to later social-media use). Furthermore, we included a wide range of covariates, including demographics and in-person social interaction, as well as other-screen and nonscreen activities.

Method

Participants

The adolescent sample comprised 597 students (50.8% female) from elementary schools in Ontario, Canada, who were surveyed annually for 2 years beginning in 2017. At Time 1, all participants were in the 6th, 7th or 8th grade (M_{age} = 12.21, *SD* = .931), with 95.7% born in Canada. Participant race included 85.7% White, 5.9% Latin American/Hispanic, 4.7% Black, 3% Filipino, and 1.5% Asian. Mean levels of parental education fell between "some college, university, or apprenticeship program" and "completed a college/apprenticeship and/or technical diploma."

The young-adult sample consisted of 1,132 undergraduates (70.5% female) who were enrolled at a university in Ontario, Canada. Participants were surveyed annually for 6 years beginning in 2010. At Time 1, all participants were in their first year of university (M_{age} = 19.06, *SD* = 1.05), with 87.5% born in Canada. Students born in Canada were asked, "Other than Canadian, is there another culture or ethnic background that your family belongs to?" Most common responses included British (17%), Italian (15%), French (8%), and German (8%). International students (11.8%) were predominantly from Asia (4%), the European Union (2%), Africa (1%), and the Caribbean (1%; students were not asked about race). Parental education levels were similar to the adolescent sample.

Procedure

Adolescents were invited to participate in the study through visits to schools. Surveys were completed in classrooms during school hours. Participants received a gift (e.g., backpack) as compensation. The young adults were invited to participate in the study through posters, announcements, Web site postings, and student-residence visits. Participants were given course credit or money as compensation. In both samples, all students who participated in the first assessment were invited to participate again at each follow-up assessment. The University Ethics Board approved both studies. Young adults provided informed consent prior to participation while adolescents provided informed assent and their parents provided informed consent.

Primary measures

Depressive symptoms. Depressive symptoms were assessed at each time point using the Center for Epidemiologic Studies Depression Scale (CES-D; Radloff, 1977) for the young adults and the CES-D Scale for Children for the adolescents (Weissman, Orvaschel, & Padian, 1980). Responses on these 20-item scales were measured on a 5-point scale ranging from 1 (*none of the time*) to 5 (*most of the time*). Responses were averaged across the 20 items for each sample to create composite measures (α = .793 and .830 for Times 1 and 2 for adolescents, and ranged from .784 to .811 for young adults).

Social-media use. In both samples, frequency of social-media use was assessed at each time point with two items—one for weekday use and one for weekend use (i.e., "On an average weekday, and day on the weekend, how many hours do you spend going on social media?"). Response options were as follows: 1 (*not at all*), 2 (*less than 1 hour*), 3 (*1–2 hours*), 4 (*3–4 hours*), 5 (*5 or more hours*). The average scores for weekend and weekday items were combined to form composite social-media scales (*r* = .880 at both time points for adolescents, and ranged from .700 to .783 for young adults).

Covariates

To be as consistent as possible with Twenge et al. (2018), we included the following covariates: demographics (age, born inside or outside Canada, parental education), other-screen activities (video-game and television use), and nonscreen activities (homework, exercise, sports, in-person social interaction, club involvement).

Demographics. In both samples, age, sex, born inside or outside of Canada, and parental education (one item per parent; 1 = *did not finish high school* to 6 = *professional degree*, averaged for both parents; *r* = .452 and *r* = .440 for the adolescent sample and the young adult sample, respectively) were assessed at Time 1.

Other-screen activities. In the adolescent sample, frequency of video-game play and television use were each assessed at Time 1 with two items (i.e., "On an average weekday, and day on the weekend, how many hours do you spend playing or doing this activity?"). Response options ranged from 1 (*not at all*) to 5 (*5 or more hours*). For each measure, the two items were combined to create a composite measure, *r*s = .848 and .703 for video games and television, respectively. Frequency of television use at Time 1 was assessed the same way for the young adults (*r* = .590 for the composite measure). Video-game use for the young adults was assessed at Time 1 by asking "How often do you play the following kinds of video/computer games?" and presenting participants with a list of 12 different types of games (e.g., sports, puzzles, and so forth). For each type, participants responded on a scale ranging from 1 (*not at all*) to 5 (*5 or more hours*). Responses were combined to create a composite measure (α = .824).

Nonscreen activities. The amount of time participants spent on homework and in-person social interaction each were measured at Time 1 only among adolescents with 2 items: "On an average weekday, and day on the weekend, how many hours do you spend doing that activity?"; 1 (*not at all*) to 5 (*5 or more hours*). The two items were combined to create a composite measure (r = .633 and .571 for homework and in-person social interaction, respectively). At Time 1 in both samples, physical activity was assessed using three items, which assessed how many times they "participated in a physical activity on [their] own or with a team" that was of (a) high intensity, (b) moderate intensity, or (c) low intensity. Responses were on a 5-point scale from 1 (*every day*) to 5 (*not at all*). The three items were averaged to create a composite variable. For adolescents, involvement in sports and clubs was measured at Time 1 with the question, "Since the beginning of summer last year, how often do you do these activities?" Participants indicated how much they engaged in 19 sports (e.g., swimming, basketball, and so forth) and 12 organized nonsport activities (e.g., YMCA/YWCA) on a scale from 1 (*never*) to 4 (*several times a week*). The responses for both sports and clubs were averaged to create a composite measure. For young adults, club involvement was assessed at Time 1 with the question, "Since September 2010 [the start of university], how often have you participated in school or community clubs that are not sports clubs" (1 = *never*; 6 = *several times a week*). Participation in sports was measured at Time 2 with the question, "How often have you participated in sports?" (1 = *never*, 6 = *several times a week*).

Results

Analyses were carried out using an autoregressive cross-lagged path analysis in MPlus 7 (Muthén & Muthén, 2012). Two models were run, one with the adolescents and one with the young adults. Depressive symptoms and social-media use were measured across 2 years for the adolescents and 6 years for the young adults. Across all time points, we included cross-lag paths, autoregressive paths (i.e., within each variable), and concurrent associations among all variables within each wave. Covariates included age, parental education, whether participants were born in Canada (yes/no), other-screen activities (video-game and television use), and nonscreen activities (homework, in-person social interaction, exercise, sports involvement, and club involvement; see Table 1). Thus, correlations were specified between each of the covariates and each variable at Time 1 and paths were estimated between the covariates and each variable at Time 2. Any significant cross-lagged paths (i.e., between social-media use at Time 1 and depressive symptoms at Time 2, or vice versa), therefore, accounted for covariates, previous scores on the outcome variables, correlations among variables within a wave, and any other predictors in the model (i.e., estimating unique relationships among study variables). Model fit was assessed only for the young adults, as the models for adolescents were saturated and thus model fit was irrelevant.

Given that Twenge et al. (2018) found sex differences (see also Steers, Wickham, & Acitelli, 2014), we conducted models separately for males and females. Furthermore, we ran each model with and without the covariates to determine whether they affected the pattern of results. Each model was first run without any

Table 1. Means and Standard Deviations of Study Variables for Both Samples

	Young adults		Adolescents	
Variable	*M*	*SD*	*M*	*SD*
Depressive symptoms 1	2.11	0.65	1.73	0.62
Depressive symptoms 2	2.09	0.68	1.84	0.67
Depressive symptoms 3	2.10	0.70		
Depressive symptoms 4	2.04	0.68		
Depressive symptoms 5	1.99	0.70		
Depressive symptoms 6	1.94	0.68		
Social media 1	3.45	0.98	2.62	1.27
Social media 2	3.28	0.99	3.04	1.36
Social media 3	3.19	0.95		
Social media 4	3.18	0.95		
Social media 5	3.10	0.92		
Social media 6	3.01	0.89		
Age 1	19.06	0.93	12.22	0.93
Sex 1 (% female)	70.5%		50.8%	
Born in Canada 1 (%)	87.5%		95.7%	
Parent education 1	3.65	1.29	3.76	1.15
Video games 1	1.28	0.43	2.79	1.37
Television 1	3.31	0.98	2.69	1.02
Exercise 1	2.58	0.91	3.27	1.02
Sports 1	1.97	1.20	1.52	0.47
Clubs 1	1.86	1.39	1.25	0.35
Homework 1			2.23	0.85
In-person interaction 1			3.12	1.18

Note: For the young-adult sample, N = 1,132 and numbers 1 to 6 refer to the years 2010–2016, respectively. For the adolescent sample, N = 597 and numbers 1 and 2 refer to the years 2017 and 2018, respectively. To improve interpretation, the scale ranges for depressive symptoms and social-media use are as follows: Depressive symptoms: a 5-point scale ranging from 1 (*none of the time*) to 5 (*most of the time*). Social-media use: 1 (*not at all*), 2 (*less than 1 hour*), 3 (*1–2 hours*), 4 (*3–4 hours*), 5 (*5 or more hours*).

covariates. Demographic covariates were then added to the models (e.g., age), followed by all other covariates (e.g., frequency of video-game play). Because the young-adult sample did not include two of the covariates (homework and in-person social interactions), the adolescent models were also run with and without those two covariates (i.e., included to be consistent with Twenge et al., 2018, but also excluded to be consistent with our young-adult sample).

Adolescents

Among males, there was a unidirectional relationship between social-media use and depressive symptoms, such that greater social-media use predicted more depressive symptoms over time, β = .152, *B* = .066, *SE* = .030, *p* = .030, 95% confidence interval (CI) = [.006, .125]. When demographic variables were added to this model (e.g., age), however, the relationship between social-media use and depressive symptoms over time was no longer significant, β = .130, *B* = .056, *SE* = .031, *p* = .071, 95% CI = [–.010, .248]. When all covariates were added to the model, the results remained consistent. Specifically, social-media use at Time 1 did not predict depressive symptoms at Time 2 (β = .145, 95% CI = [.000, .251], *p* = .051) and depressive symptoms at Time 1 did not predict social-media use at Time 2, β = .094, 95% CI [–.033, .220], *p* = .145 (see Table 2 for full results). Among the covariates, only exercise (Time 1) predicted more social-media use at Time 2. When homework and in-person social interactions were added to this model, the results remained consistent.

Among females, higher depressive symptoms at Time 1 predicted more-frequent social-media use at Time 2, β = .108, *B* = .236, *SE* = .115, *p* = .041, 95% CI = [.009, .462], but not vice versa. Next, we ran a model that included only demographic covariates (e.g., age) and then a model that included all covariates. The results remained consistent across both models in that higher depressive symptoms at Time 1 predicted more-frequent social-media use at Time 2, β = .131, 95% CI = [.026, .236], *p* = .015, but not vice versa, β = –.043, 95% CI = [–.159, .073], *p* = .468 (see Table 2). Furthermore, lower parental education and older age predicted greater levels of depressive symptoms at Time 2 (see Table 2). Less video-game use also predicted greater Time 2 social-media use. When homework and in-person social interactions were added to this model, the results remained consistent; however, in-person interactions with friends was also a predictor of more

Table 2. Autoregressive Cross-Lagged Model Results for Adolescents

	Time 1→Time 2					
	Females			Males		
Structural path	*B*	β	95% CI	*B*	β	95% CI
Soc.Media1→Dep.2	–0.022	–0.043	[–0.159, 0.073]	0.062	0.145	[0.000, 0.288]
Dep.1→Soc.Media.2	0.287	0.131*	[0.026, 0.236]	0.240	0.094	[–0.033, 0.220]
Soc.Media1→Soc.Media.2	0.658	0.612***	[0.513, 0.711]	0.499	0.459***	[0.340, 0.577]
Dep.1→Dep.2	0.573	0.546***	[0.454, 0.638]	0.369	0.365***	[0.238, 0.492]
Video game.1→Dep.2	0.058	0.109	[–0.002, 0.219]	0.032	0.070	[–0.076, 0.216]
TV.1→Dep.2	0.053	0.083	[–0.017, 0.184]	0.009	0.018	[–0.122, 0.158]
Exercise.1→Dep.2	–0.034	–0.053	[–0.159, 0.053]	0.007	0.014	[0.135, 0.163]
Age.1→Dep.2	0.195	0.283***	[0.174, 0.391]	0.043	0.078	[–0.066, 0.221]
Canadian.born.1→Dep.2	0.279	0.084	[–0.044, 0.211]	0.149	0.062	[–0.077, 0.202]
Clubs.1→Dep.2	0.020	0.011	[–0.102, 0.124]	–0.053	–0.035	[–0.171, 0.101]
Sports.1→Dep.2	–0.114	–0.079	[–0.188, 0.029]	0.095	0.094	[–0.047, 0.234]
Parent.educ.1→ Dep.2	–0.085	–0.143*	[–0.284, –0.002]	0.007	0.016	[–0.134, 0.166]
Video.game.1→Soc.Media.2	–0.127	–0.114*	[–0.219, –0.009]	–0.096	–0.082	[–0.215, 0.051]
TV.1→Soc.Media.2	–0.008	–0.006	[–0.102, 0.091]	–0.140	–0.112	[–0.232, 0.009]
Exercise.1→Soc.Media.2	0.112	0.083	[–0.024, 0.190]	0.211	0.171*	[0.041, 0.301]
Age.1→Soc.Media.2	–0.005	–0.004	[–0.112, 0.105]	0.080	0.057	[–0.069, 0.183]
Canadian.born.1→Soc.Media.2	0.612	0.088	[–0.025, 0.201]	0.248	0.041	[–0.082, 0.165]
Clubs.1→Soc.Media.2	0.198	0.051	[–0.050, 0.152]	–0.417	–0.111	[–0.228, 0.007]
Sports.1→Soc.Media.2	–0.119	–0.040	[–0.148, 0.069]	–0.031	–0.012	[–0.133, 0.108]
Parent.educ.1→Soc.Media.2	–0.080	–0.064	[–0.198, 0.069]	–0.061	–0.059	[–0.187, 0.069]

Note: B = unstandardized beta weights; β = standardized beta weights; CI = standardized confidence intervals; Dep. = depressive symptoms; Soc.Media = social media; Parent.educ. = parent education. Numbers 1 and 2 represent Times 1 and 2, respectively.
*p < .05. **p < .01. ***p < .001.

Table 3. Autoregressive Cross-Lagged Model Results for Young Adults

	Time 1→Time 2					
	Females			Males		
Structural path	*B*	β	95% CI	*B*	β	95% CI
Soc.Media.1→Dep.2	0.015	0.011	[–0.018, 0.039]	0.024	0.014	[–0.029, 0.056]
Dep.1→Soc.Media.2	–0.006	–0.008	[–0.039, 0.023]	–0.016	–0.022	[–0.071, 0.027]
Soc.Media.1→Soc.Media.2	0.518	0.511***	[0.452, 0.570]	0.625	0.505***	[0.405, 0.606]
Dep.1→Dep.2	0.485	0.477***	[0.414, 0.540]	0.565	0.555***	[0.458, 0.651]
Video.game.1→Dep.2	0.063	0.022	[–0.047, 0.091]	0.029	0.015	[–0.098, 0.128]
TV.1→Dep.2	0.072	0.073*	[0.003, 0.143]	–0.080	–0.079	[–0.190, 0.032]
Exercise.1→Dep.2	0.052	0.050	[–0.024, 0.124]	–0.028	–0.024	[–0.143, 0.094]
Age.1→Dep.2	–0.018	–0.017	[–0.091, 0.056]	–0.096	–0.090	[–0.191, 0.011]
Canadian.born.1→Dep.2	0.092	0.030	[–0.039, 0.100]	–0.194	–0.064	[–0.174, 0.045]
Clubs.1→Dep.2	–0.059	–0.091**	[–0.158, –0.023]	–0.009	–0.010	[–0.108, 0.088]
Sports.1→Dep.2	–0.031	–0.056	[–0.129, 0.017]	–0.089	–0.168**	[–0.284, –0.052]
Parent.educ.1→ Dep.2	–0.009	–0.012	[–0.083, 0.058]	–0.140	–0.186**	[–0.298, –0.075]
Video.game.1→Soc.Media.2	0.148	0.074*	[0.005, 0.142]	0.006	0.004	[–0.107, 0.115]
TV.1→Soc.Media.2	0.052	0.074*	[0.007, 0.142]	–0.002	–0.002	[–0.110, 0.105]
Exercise.1→Soc.Media.2	0.007	0.010	[–0.065, 0.084]	0.101	0.120*	[0.001, 0.238]
Age.1→Soc.Media.2	–0.023	–0.032	[–0.106, 0.042]	0.007	0.010	[–0.089, 0.109]
Canadian.born.1→Soc.Media.2	0.035	0.017	[–0.052, 0.086]	–0.053	–0.025	[–0.132, 0.082]
Clubs.1→Soc.Media.2	–0.015	–0.033	[–0.101, 0.035]	0.061	0.094	[–0.002, 0.190]
Sports.1→Soc.Media.2	–0.009	–0.023	[–0.096, 0.050]	–0.053	–0.139*	[–0.254, –0.025]
Parent.educ.1→Soc.Media.2	–0.003	–0.005	[–0.076, 0.065]	–0.001	–0.002	[–0.114, 0.111]

Note: *B* = unstandardized beta weights; β = standardized beta weights; CI = standardized confidence intervals; Soc.Media = social media; Dep. = depressive symptoms; Parent.educ. = parental education. Numbers 1 and 2 represent Times 1 and 2, respectively. Given that the pattern of results was consistent across time, only the results from Times 1 and 2 are shown.
*$p < 0.05$. **$p < 0.01$. ***$p < 0.001$.

social-media use over time, β = .112, *B* = .129, *SE* = .062, $p = .038$, 95% CI = [.007, .251].

Young adults

For males and females, we first assessed whether the pattern of results was invariant across time. Across-time invariance was tested by comparing a model in which all cross-lagged paths were constrained to be equal across time to the unconstrained model in which paths were free to vary. For males, the Chi-Square Difference Test of Relative Fit indicated that the unconstrained model was not a significantly better fit than the constrained model, $\chi^2_{diff}(8) = 11.499$, $p = .174$, suggesting that the pattern of results was consistent across the 6 years. Thus, all subsequent analyses were based on the model in which cross-lagged paths were constrained to be equal over time, as this was the more parsimonious model. The constrained model fit was good, $\chi^2(28) = 33.441$, $p = .220$, comparative fit index (CFI) = .993 and root-mean-square error of approximation (RMSEA) = .024, 90% CI = [.000, .051], $p = .943$ (Hu & Bentler, 1999). There were no significant bidirectional or unidirectional relationships between social-media use and depressive symptoms over time (*p*s > .05). Next, we ran a model that included only demographic covariates (e.g., age) and then a model that included all covariates. The results remained consistent across both models. Specifically, social-media use at Time 1 did not predict depressive symptoms at Time 2, β = .014, 95% CI = [–.029, .056], $p = .528$, and depressive symptoms at Time 1 did not predict social-media use at Time 2, β = –.022, 95% CI = [–.071, .027], $p = .370$ (see Table 3 for full results). Furthermore, lower parental education predicted greater levels of depressive symptoms at Time 2 (see Table 3). Greater involvement in sports predicted fewer depressive symptoms, and finally less involvement in sports and more exercise predicted greater social-media use.

Among females, the Chi-Square Difference Test of Relative Fit indicated that the unconstrained model was not a significantly better fit than the constrained model, $\chi^2_{diff}(8) = 9.416$, $p = .308$, suggesting that the pattern of results was consistent across the 6 years. Thus, all subsequent analyses were based on the models that were constrained over time, as this was the more parsimonious

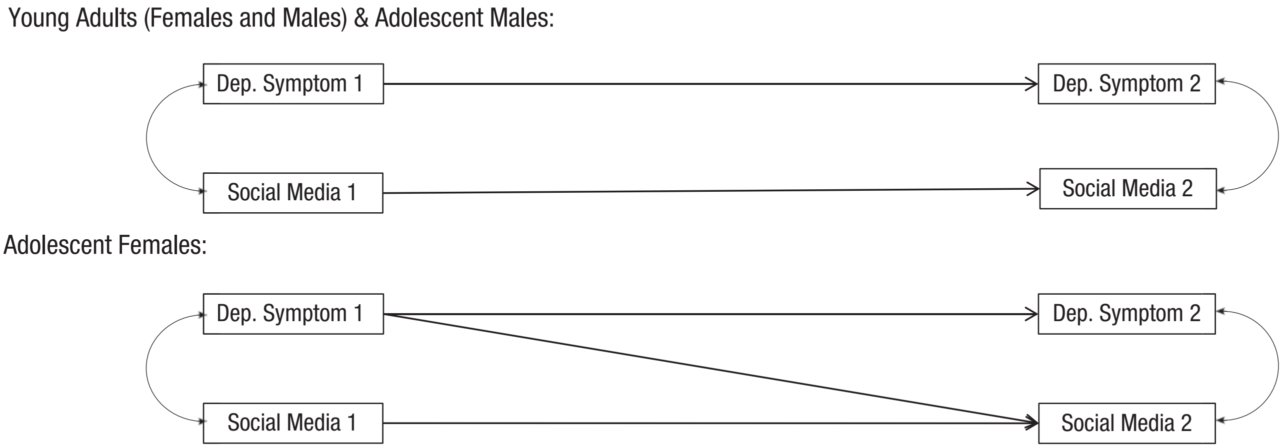

Fig. 1. Significant results from the full model. Dep. Symptom = depressive symptoms. Numbers 1 and 2 represent Times 1 and 2, respectively. Results from female and male young adults and male adolescents are depicted in the same model because they had the same findings (no significant cross-lag associations). Given that the results for the young-adult sample were invariant across time, only the paths from Time 1 to Time 2 are shown. Results from the covariates can be obtained from Tables 2 and 3.

model. The constrained model fit was good, $\chi^2(28)$ = 23.392, p = .713, CFI = 1.00 and RMSEA = .000, 90% CI = [.000, .021], p = 1.00. There were no significant bidirectional or unidirectional relationships found between social-media use and depressive symptoms over time for all three models (i.e., without demographics or covariates, with demographics, and with all covariates; ps > .05; see Table 3). Therefore, social-media use at Time 1 did not predict depressive symptoms at Time 2, β = .011, 95% CI = [–.018, .039], p = .467, and depressive symptoms at Time 1 did not predict social-media use at Time 2, β = –.008, 95% CI = [–.039, .023], p = .604 (see Table 3). For the model with all covariates, watching television was associated with more depressive symptoms and more social-media use over time. Furthermore, greater club involvement was associated with fewer depressive symptoms, while more video-game use predicted more social-media use over time.

Discussion

The current longitudinal study extends earlier work by using a cross-lagged design to investigate the associations between depressive symptoms and social-media use across time. Although there is speculation that the direction of the relationship is from social-media use to increased depressive symptoms, longitudinal studies, testing both directions simultaneously, are required in order to infer temporal order. Furthermore, our measure of social-media engagement captured average daily frequency rather than average yearly frequency (see Twenge et al., 2018). Finally, we investigated these associations among both adolescents and young adults, and controlled for many of the covariates that were highlighted by Twenge and colleagues (e.g., demographics and in-person social interaction), as well as other potentially important third variables (e.g., homework, exercise, club involvement, television, and so forth).

Critically, in the current study, social-media use did not predict future depressive symptoms in adolescent females. This key finding challenges Twenge et al's (2018) claim that the positive association found in their sample among adolescent girls between social-media use and depression was likely from social-media use leading to greater depression. Our results also revealed that there were no significant associations between Time 1 depressive symptoms and Time 2 social-media use in either the young adult or adolescent male sample. Among female adolescents, however, greater depressive symptoms at Time 1 predicted more engagement in social media at Time 2 (see Figure 1). Again, these findings are not consistent with the hypothesis that social-media use "causes" depressive symptoms. In fact, our results suggest that there are no significant relationships between social-media use and depressive symptoms over time among young adults and adolescent males, and that adolescent girls who experience depressive symptoms tend to use more social media across time (rather than vice versa).

Given the larger betas and bivariate correlations found among adolescents compared to young adults (see Tables 2 and 3, and Tables 1 and 2 in the Supplemental Material available online), our results highlight that the relationship between depressive symptoms and social-media use appears to be stronger among adolescents. In fact, adolescents may use social media in different ways than young adults. For example, Frison and Eggermont (2015) found that late adolescents (19-year-olds) used

Facebook in a more private way (e.g., sending private messages to friends) than young adolescents, who are more likely to post pictures to their timeline which all of their friends can view. It also could be that young adults are more likely than adolescents to use social media to keep up with friends that they no longer see face to face, while adolescents who are feeling down may turn to social-media use because they expect that it will make them feel better (see Sagioglou & Greitemeyer, 2014). More research clarifying these potential developmental differences is important.

Of note, we found a marginally significant relationship between greater social-media use at Time 1 and depressive symptoms at Time 2 among adolescent males when all covariates were included. The effect size for this path (β = .15) was similar to the effect size of the significant paths found in our study. While this result was nonsignificant in our final model with covariates (as well as in our young-adult sample), more research is needed before conclusions about this relationship in males can be drawn. Furthermore, although Twenge et al. (2018) suggested that other-screen activities (e.g., video games) may partially explain increases in depressive symptoms over time, we did not find any significant associations between video gaming and depressive symptoms over time, nor did we find that in-person interactions predicted lower depressive symptoms over time (as suggested by Twenge et al., 2018).

The current study is not without limitations. First, while longitudinal data provides stronger grounds for causal inference than cross-sectional data (i.e., temporal ordering of variables can be established with the former but not the latter), third variable effects still are possible with longitudinal designs (i.e., without random assignment to conditions, it is not possible to rule out the possibility that some unmeasured, extraneous variable is responsible for an association).

Furthermore, we measured only frequency—not quality—of social-media use. Measures that capture quality of social-media use (e.g., time of day when social media is used, using social media for social comparisons, social-media addiction, active versus passive engagement, and types of social-media use) may yield different results in terms of their associations with depressive symptoms (e.g., Frison & Eggermont, 2015; Nesi & Prinstein, 2015; Teppers, Luyckx, Klimstra, & Goossens, 2014). Our study also focused specifically on depressive symptoms, but other potential predictors or outcomes also are important (e.g., victimization, sleep problems). Another limitation of the current study is that the participants came from relatively homogeneous samples, in contrast to Twenge et al. (2018) who analyzed two large, nationally representative samples. Thus, our results may not generalize to populations from other regions with greater ethnic diversity.

Another reason to exercise caution with regard to the panic around social media “causing” depression, and the application of this body of literature to policies directed at adolescents and their parents, is related to small effect sizes. While small standardized coefficients are to be expected among longitudinal models, given that they control for previous scores (within variables), covariates, and correlations among variables within each wave (Adachi & Willoughby, 2015), in this body of literature, the magnitude of even cross-sectional associations are small. For instance, in a meta-analysis of 61 primarily nonlongitudinal studies (Huang, 2017), the mean correlation between frequency of social-media use and psychological well-being was $r = -.07$, a trivial effect (see also Booker et al., 2018).

Overall, the findings from this study suggest that the relationship between depressive symptoms and social-media use is complex. Among young adults, we found no evidence of a relationship between social-media use and depressive symptoms over time. Furthermore, adolescent girls with higher depressive symptoms reported greater social-media use over time. Thus, while it may be common in popular media to suggest that social-media use might cause depression (e.g., Ghaemi, 2018; Hasselton, 2017; Twenge, 2017), our results suggest that this claim may be premature. Indeed, recent policy debates regarding ways to tackle the mental-health effects of social-media use may be based on the false premise that social-media use leads to increases in depression (Royal Society for Public Health, 2018). Replications of these findings, of course, will be important to establish more definitively the direction of the effect between social-media use and depressive symptoms over time.

Action Editor

Kelly L. Klump served as action editor for this article.

Author Contributions

All authors developed the study concept, contributed to the study design, and were involved in collecting data for the study. T. Willoughby and T. Heffer performed the data analysis. All authors interpreted the data, cowrote and revised the manuscript, and approved the final version for submission.

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Supplemental Material

Additional supporting information can be found at http://journals.sagepub.com/doi/suppl/10.1177/2167702618812727

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