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Is there a correlation between transgender identity and screen time?

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Abstract
The number of individuals identifying as transgender, or gender non-binary, has risen in recent decades. One hypothesis holds that this increase is attributable to the influence of social media. To test whether transgender identification correlated with screen time and/or social media use, while controlling for relevant covariates, we used data from the Florida version of the Center for Disease Control (CDC) Youth Risk Behavior Survey (YRBS) administered in 2019 and 2021 ($N=9055$), as well as national data from the 2023 YRBS ($N=9581$). Results showed that screen time had a very small but statistically significant association with Transgender identification among adolescents in the 2019 and 2021 data. In the 2023 YRBS data, no association was found between social media use and transgender identity, except among natal females, where greater social media usage was associated with a lower probability of transgender identity. The effect sizes of these relationships are well below accepted benchmarks of practical or theoretical significance. We conclude that that the relationship between social media use and transgender identity is unclear, and offer diverging interpretations of these small or absent effects. After noting the substantial limitations of the present data, we discuss methodological considerations and refinements that would offer a stronger test of the purported relationship between social media use and gender diverse identities.

Keywords Gender variance · Epidemiology · Social contagion

The past decade has seen a marked increase in public discourse, and academic research, surrounding gender identity and persons who identify as transgender (e.g., Twenge et al., 2024). For most individuals, gender identity (woman/man; girl/boy) aligns with biological sex (female/male), but this is not true for all persons. Transgender and non-binary identities occur when a persons' experienced gender, sense of self, and often public self-expression, do not align with their natal sex, or the sex/gender documented at birth (Coleman et al., 2022). Estimates vary, but high-quality representative samples indicate that persons identifying as either Transgender or gender non-binary comprise somewhere between 0.3% of the Canadian (Statistics Canada, 2022), and 0.5% of the US (Herman et al., 2022), populations respectively. General population samples show transgender identification is more common among younger generations (~1.5% of respondents) (Diamond, 2020; Herman et al., 2022; Statistics Canada, 2022; Twenge, et al., 2024), and similar increases in the number of individuals accessing gender related health services have been documented in the past two decades (e.g., Kaltiala et al., 2019; Thompson et al., 2022; Zucker, 2017).

Concurrent with increases in the number of individuals identifying as Transgender has been a shift in the natal sex ratio. Prior to the 2000s, natal males outnumbered natal females roughly 2:1 among those seeking therapeutic treatment with gender related concerns (Zucker, 2017). This ratio has now been reversed, with more natal females identifying as transgender (and gender non-binary) than natal males (Cass, 2024; Thompson et al., 2022; Twenge et al., 2024; Zucker, 2017). Increases in transgender identification, and changes in male/female ratio, have even been documented in cultural locales that are repressive toward gender and sexual orientation diversity, such as Iran (Talaei et al., 2022), indicating these trends extend beyond Euro-America.

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Understanding demographic shifts in persons identifying as transgender requires recognition that gender dysphoria is not a singular phenomenon, and involves differential age of onset, sexual orientations, and patterns of childhood behavior (Bailey & Blanchard, 2017; Lawrence, 2010). For most of the 20th century, Transgender identities were relatively rare, with meta-analytic prevalence estimates of 6.8/100,000 males and 2.6/100,000 females accessing specialty gender services (Arcelus et al., 2015). Contemporary prevalence estimates are markedly higher for those seeking medical and mental health services (Cass, 2024; Zucker, 2017), and higher yet for those who self-identify as transgender (Twenge et al., 2024). One type of gender dysphoria involves early childhood onset (i.e., < 12 years old, but often much younger), is strongly associated with early expression of sex-atypical behavior and preferences, and (eventual) same-sex attraction relative to natal sex (Lawrence, 2010; Ristori & Steensma, 2016; Zucker, 2005). Late onset gender dysphoria, emerging after puberty, has been observed among natal males for several decades. This type of gender dysphoria sometimes emerges in adolescence, and sometimes well into adulthood, and is not associated with exclusive same-sex attraction or marked sex-atypicality in childhood (Blanchard, 1989; Lawrence, 2010). More recently, epidemiological shifts have occurred in the number of adolescent and young adult natal females identifying as transgender (e.g., Twenge et al., 2024), and seeking gender-related treatment (Cass, 2024). Many of these adolescent females differ from those with early onset presentation, because they do not display strong sex-atypical behaviors and preferences in childhood and the disclosure of their transgender identities comes as a surprise to their families (Littman, 2018). Numerous explanations have been proposed to explain the increases in transgender identity, as well as the shift in natal sex-ratio. Transgender identities have become less pathologized by clinicians, and less stigmatized in the general public, than in previous generations, which has in turn led to greater access to care such as puberty blockers and cross-sex hormones intended to lessen individuals' distress and affirm their gender identity (e.g., Coleman et al., 2022; Kaltiala et al., 2019; Thompson et al., 2022). In short, this explanation asserts that sociocultural changes *reveal* transgender persons who are now more comfortable acknowledging and expressing their identities than they were in the past. Other explanations also emphasize recent sociocultural changes, but instead argue that peer influence and social salience of transgender persons in some way *cause* (or perhaps induce) certain individuals to adopt these identities. Transgender issues are highlighted in numerous forms of media (positively or negatively), and there has been a proliferation of websites devoted to gender dysphoria and transgender care (Zucker, 2019). Some scholars have suggested that peer and sociocultural influence are particularly relevant factors for explaining the rising number of natal females whose transgender identification emerges in adolescence (Littman, 2018; Marchiano, 2017) in the absence of a childhood history of gender nonconforming behavior (see also Lawrence, 2010; Zucker & Lawrence, 2009). Some explanations go even further, suggesting that many adolescent females with existing psychosocial vulnerabilities, or perhaps simply discomfort about their developing bodies, are influenced by social media messages that encourage these individuals to interpret this discomfort as evidence that they are transgender (Shrier, 2020; Littman, 2018).

Suggesting that developing a transgender identity can be socially influenced, or is perhaps socially contagious, echoes research examining rising rates of depression and anxiety among North American adolescents. Some scholars have attributed this rise in mental health issues in part to the influence of social media (e.g., Twenge, 2020; Twenge et al., 2024b), whereas other scholars have concluded that the evidence for this connection is thin (e.g., Heffer et al., 2019; Jensen et al., 2019; Yang & Feng, 2024).

The idea that increases in transgender identification among younger generations is driven in part by peer or social influence is controversial, being rejected by numerous researchers (e.g., Bauer et al., 2022; Indremo et al., 2022; Restar, 2020; Turban et al., 2022). As with any controversy, other experts are less dismissive of the possibility that peer and sociocultural factors play a role (Cass, 2024; Littman, 2018; Littman et al., 2024). The precise mechanisms for the purported social contagion of transgender identity are presently unclear, varying from highly specific vectors such as direct peer interaction promoting transgender identification in existing social groups (Littman, 2018), to the much broader suggestion that the prominence of transgender issues in media spaces (Zucker, 2019) means that greater social media use precedes questioning one's own gender identity and perhaps even transgender identification (Marchiano, 2017; Shrier, 2020). A very simplistic prediction flowing from the latter assertion is that transgender identity will be correlated with screen time and social media use. Here, we examine the purported relationship between transgender identification and social media use in the Youth Risk Behavior Survey (YRBS), nationally representative data collected by the Center for Disease Control (CDC). These cross-sectional data are of course incapable of revealing whether social media use *causes* transgender identity but could in principle reveal whether the two variables are correlated, a connection found in some other large-scale datasets (e.g., Nagata et al., 2024).

No formal predictions were made (in fact, the co-authors disagreed about the likely direction of results, making this a cheerful adversarial collaboration of sorts), but several relevant covariates were identified for our preregistered analysis. For Study 1, which used YRBS data collected in Florida during 2019 and 2021 waves, we examined whether time on spent engaging with digital media was associated with transgender identification while also controlling for the influence of age (e.g., Herman et al., 2022), sex (Thompson et al., 2022; Zucker, 2017), race (Twenge et al., 2024), as well as indicators of (internalizing) depression and self-harm (de Graaf et al., 2022) and (externalizing) delinquency type behaviors (e.g., van der Miesen et al., 2020). National 2023 YRBS data included new questions about adverse childhood experiences (ACEs), so Study 2 mirrored the above analysis but added ACEs as a covariate given that such experiences are frequently elevated among those who identify as transgender (e.g., Kreski & Hayes, 2025; Tran et al., 2023). We generally employed a smallest effect size of interest (SESOI) equivalent to $r = \pm 0.10$ (or odds ratio of ≥ 1.44 or ≤ 0.69). This was employed because in large sample studies, noise or "crud" results can often become statistically significant, even if these reflect tiny, chance correlations (Ferguson & Heene, 2021). This SESOI was not mentioned in the preregistration for Study 1 due to an omission, but was included in our Study 2 preregistration, and is a routine effect size benchmark for our lab. We nonetheless report and interpret statistical significance and whether the effect size exceeded our SESOI in the interests of completeness.

This SESOI is included due to the observation that, in large sample studies, nonsense correlations often become "statistically significant" due to methodological noise or chance, rather than due to real effects (Ferguson & Heene, 2021). It is recognized that reliance on the p-value is insufficient for hypothesis support and likely to result in false positive conclusions (Anvari & Lakens, 2021). In Ferguson and Heene's (2021), analysis it was demonstrated that 2 out of 3 nonsense correlations (i.e., correlations between variables specifically chosen to have no theoretical relationships) were detected as "statistically significant" at $p < .05$ or lower. False positive correlations remained high up to the level of $r = .20$. They suggested adopting $r = .10$ as a minimum SESOI, with $r = .20$ reserved for clinical/practical significance. We have adopted this threshold ourselves.

Study 1

Methods

Participants

The Youth Risk Behavior Survey (YRBS) is administered in American schools by the Center for Disease Control (CDC) on a biennial basis, and Study 1 utilized archival data from 2019 to 2021 collection years in Florida. A total of 10,375 students were available in the data (5703 from 2019; 4672 from 2021). Data were considered for participants who provided complete responses to questions evaluating sex, age, transgender identification, and screentime usage, as well as providing sufficient responses to calculate indices of depression/self-harm and delinquency. This inclusion criteria resulted in 9055 total participants (4605 females, 4450 males), with a mean age of approximately 16 years. Most participants identified as white (64.3%), with the remainder belonging to other racial/ethnic categories, and 140 participants (1.5%) identified themselves as being transgender.

Materials

The YRBS is designed and administered by the CDC to monitor health and wellbeing indicators among American youth. Our inclusion criteria and analysis plan were preregistered (https://aspredicted.org/GPW_GK2) before exam ining the data, and involved using binary logistic regression to predict transgender identification using the variables of age, sex, race (white vs. non-white), depressive self-harm, delinquency, and screen time.

Biodemographic variables

Sex was evaluated with the question "what is your sex?" and participants selected from binary options (Female/ Male). Age was measured with seven response options ranging from "12 years old or younger" to "18 years old or older." Race was evaluated with 7 response options (American Indian or Alaska Native; Asian; Black or African American; Hispanic or Latino; Middle Eastern or North African; Native Hawaiian or Pacific Islander; White). This variable was dichotomized to capture participants who were White only (64.3% of respondents), or selected other/additional response options (35.7%).

Transgender identity

Our main outcome variable was transgender identification, measured by the question: "Some people describe themselves as transgender when their sex at birth does not match the way they think or feel about their gender. Are you transgender?" with the response options of: No, I am not transgender; Yes, I am transgender; I am not sure if I am transgender; and I do not know what this question is asking. Only participants who answered in the affirmative (i.e., "Yes, I am transgender") were coded as being transgender, those who responded with uncertainty about their identity or the meaning of the question were excluded, and the remainder were treated as cisgender.

Screen time

Screen time, our proxy for social media engagement, was measured with a single item: "On an average school day, how many hours do you spend in front of a TV, computer, smart phone, or other electronic device watching shows or videos, playing games, accessing the Internet, or using social media (also called "screen time")? (Do not count time spent doing schoolwork)". In 2019, the item for screen use was slightly different focused more on video games and using computers for purposes other than school: "On an average school day, how many hours do you play video or computer games or use a computer for something that is not school work? (Count time spent playing games, watching videos, texting, or using social media on your smartphone, computer, Xbox, PlayStation, iPad, or other tablet)". There were six response options, ranging from "Less than 1 hour per day" to "5 or more hours per day."

Depressive self-harm

A composite variable was created to capture depressive self-harm. Here we combined responses about feeling sad or hopeless for more than two weeks in a row in the past 12 months (No/Yes), seriously considering attempting suicide (No/Yes), making a plan to attempt suicide (No/Yes), and number of suicide attempts in the past 12 months (0, 1, 2–3, 4–5, 6 or more times). A simple tally of "yes" responses including for suicide attempts was calculated. Out of concern this might introduce bias, an alternative system of *z*-scores was calculated for each question, then summed to create a composite measure such that high scores indicated greater depressive self-harm ($\alpha=0.74$ for these four items). Regression results were virtually identical with either approach, and we report present results using the composite *z*-scores.

Delinquency

Delinquency was calculated from five items (driving while intoxicated, weapon carrying generally and on school property, fighting generally and on school property). These item responses were converted to z-scores and summed ($\alpha=0.69$).

Procedures

Data collection procedures for the YRBS are described by the CDC (CDC, 2025; https://www.cdc.gov/yrbs/methods/index.html), and data for the Florida YRBS are publicly available by request from the Florida department of health at https://www.floridahealth.gov/statistics-and-data/survey-data/index.html, meaning the data are openly available for replication by other researchers. All data analysis were performed with opensource jamovi software (The jamovi project, 2025).

Results

Among all participants, 140 (1.5%) identified as transgender. This number was comparable among female participants (81/4605 = 1.8%) and male participants (59/4450 = 1.3%), and these proportions did not significantly differ, χ^2 (1) = 2,79, $p=.095$. Table 1 reports our main preregistered analyses (Model 1), a binary logistic regression using simultaneous entry of all study variables to predict transgender identification. The overall model was significant, χ^2 (6) = 173.0, $p<.001$, Nagelkerke $R^2=0.129$. In terms of simple statistical significance, screen time was positively associated with the probability of transgender identification (OR = 1.13, $p=.008$) while controlling for age, sex, race, depressive self-harm, and delinquency. However, this value did not exceed or come close to our SESOI of OR = 1.44. Three of the latter variables were significantly associated with transgender identity such that depressive self-harm and delinquency increased the odds, whereas being non-White lowered the odds. However, as with screen time, delinquency did not exceed our SESOI.

In addition to our preregistered analysis, two additional models were run as robustness checks, splitting the data by sex indicated by the participant. Binary logistic regression was again used to predict transgender identification using screen time, age, race, depressive self-harm, and delinquency (Table 2, Models 2 & 3). Among females, the overall model was significant, χ^2 (5) = 92.4, $p<.001$, Nagelkerke $R^2=0.122$, with screen time, race, and depressive self-harm all being significant predictors, although screen time did not exceed our SESOI once again. An identical model was run among males, which was also significant, χ^2 (5) = 81.3, $p<.001$, Nagelkerke $R^2=0.138$, but showed only depressive self-harm and delinquency were significant predictors of transgender identification, with only depressive self-harm exceeding our SESOI.

Study 2

Participants

Data for the 2023 National YRBS were used for Study 2, with 20,103 students available for analyses. We focussed on participants who provided complete responses about their sex, age, transgender identification, and social media usage,

Table 1 Descriptive statistics and group comparisons for combined 2019 and 2021 Florida YRBS data

	Transgender	Cisgender			
	M (*SD*)	*M* (*SD*)	Levene's Test	*t*-test (two-tailed)[a]	*d* (95%CI)
Sample (*n* = 9055)	***n* = 140**	***n* = 8915**			
Age	15.72 (1.49)	15.88 (1.23)	$F = 12.69$, $p < .001$	$t_{(142)} = -1.27$, $p = .207$	−0.12 (−0.30, 0.04)
Screen Time	4.97 (1.85)	4.33 (1.96)	$F = 4.12$, $p = .042$	$t_{(144)} = 4.04$, $p < .001$	0.33 (0.16, 0.49)
Depressive Self Harm	2.35 (2.07)	0.725 (1.21)	$F = 155.79$, $p < .001$	$t_{(140)} = 9.24$, $p < .001$	0.96 (0.76, 1.16)
Delinquency	−0.537 (4.89)	−1.43 (2.05)	$F = 168.54$, $p < .001$	$t_{(140)} = 2.17$, $p = .032$	0.24 (0.07, 0.41)
Females (*n* = 4605)	***n* = 81**	***n* = 4524**			
Age	15.78 (1.40)	15.84 (1.22)	$F = 3.38$, $p = .066$	$t_{(4603)} = -0.45$, $p = .650$	−0.05 (−0.27, 0.17)
Screen Time	4.94 (1.81)	4.27 (2.04)	$F = 7.99$, $p = .005$	$t_{(83.7)} = 3.31$, $p < .001$	0.35 (0.11, 0.55)
Depressive Self Harm	2.49 (1.97)	0.95 (1.31)	$F = 49.18$, $p < .001$	$t_{(81.3)} = 7.04$, $p < .001$	0.93 (0.66, 1.19)
Delinquency	−0.75 (5.38)	−1.35 (1.93)	$F = 158.99$, $p < .001$	$t_{(80.4)} = 1.00$, p 0.319	0.15 (−0.07, 0.37)
Males (*n* = 4450)	***n* = 59**	***n* = 4391**			
Age	15.65 (1.62)	15.93 (1.25)	$F = 10.74$, $p < .001$	$t_{(58.9)} = -1.33$, $p = .188$	−0.20 (−0.45, 0.07)
Screen Time	5.017 (1.93)	4.40 (1.87)	$F = 0.01$, $p = .929$	$t_{(4448)} = 2.51$, $p = .012$	0.33 (0.07, 0.59)
Depressive Self Harm	2.15 (2.21)	0.50 (1.05)	$F = 127.97$, $p < .001$	$t_{(58.3)} = 5.75$, $p < 001$	0.96 (0.64, 1.26)
Delinquency	−0.24 (4.16)	−1.52 (2.16)	$F = 41.73$, $p < .001$	$t_{(58.4)} = 2.36$, $p = 022$	0.39 (0.12, 0.65)

[a]Welch's *t* reported when Levene's test showed unequal variances

Table 2 Binary logistic regression predicting transgender identification

Model	Predictor	(B)	SE	*p*	OR (95% CI)
1 (All participants)	Screen Time	0.124	0.047	0.008	1.13 (1.03, 1.24)
N = 9055	Age	−0.043	0.069	0.527	0.96 (0.84, 1.10)
	Sex[a]	0.071	0.391	0.696	1.07 (0.75, 1.53)
	Race[b]	−0.378	0.177	0.032	0.66 (0.48, 0.97)
	Depressive self-harm	0.472	0.038	< 0.001	1.60 (1.49, 1.73)
	Delinquency	0.080	0.027	0.003	1.08 (1.03, 1.14)
2 (Females)	Screen Time	0.128	0.060	0.033	1.14 (1.01, 1.28)
n = 4605	Age	0.019	0.094	0.834	1.02 (0.85, 1.28)
(81 transgender)	Race[b]	−0.523	0.231	0.023	0.59 (0.38, 0.93)
	Depressive self-harm	0.439	0.049	< 0.001	1.55 (1.41, 1.71)
	Delinquency	0.082	0.042	0.052	1.09 (0.99, 1.18)
3 (Males)	Screen Time	0.127	0.076	0.093	1.14 (0.98, 1.32)
n = 4450	Age	−0.114	0.102	0.260	0.89 (0.73, 1.09)
(59 transgender)	Race[b]	−0.198	0.280	0.479	0.82 (0.47, 1.42)
	Depressive self-harm	0.518	0.058	< 0.001	1.68 (1.50, 1.88)
	Delinquency	0.075	0.036	0.037	1.08 (1.00, 1.16)

[a] Sex coded as 0 = Female, 1 = Male

[b] Race coded as 0 = White, 1 = non-White

as well as sufficient responses to questions about depression/self-harm, delinquency, and ACEs (i.e., ≤ three missing responses per composite). This inclusion criteria resulted in 9581 total participants (4677 females, 4904 males), with a mean age of approximately 16 years. Most participants identified as white (56.1%), with the remainder belonging to other racial/ethnic categories, and 291 participants (3.0%) identified themselves as being transgender.

Materials

The YRBS materials and data are available from the CDC (CDC, 2025; https://www.cdc.gov/yrbs/methods/index.html). Our inclusion criteria and analysis plan were pre-registered (https://aspredicted.org/hnmz-mkjh.pdf) before examining the data, and involved using binary logistic regression to predict transgender identification using the variables of age, sex, race (white vs. non-white), depressive self-harm, delinquency, ACEs, and social media use.

Biodemographic variables

Biodemographic variables (age, sex, race, transgender identity) were evaluated in an identical fashion to Study 1. Race was dichotomized to capture participants who were White only (56.1% of respondents) or selected other/additional response options (43.9%).

Delinquency

As in Study 1, Delinquency was calculated from five items (driving while intoxicated, weapon carrying generally and on school property, fighting generally and on school property). These item responses were converted to z-scores and averaged ($\alpha = 0.62$).

Social media use

Social media use was measured with a single item: "How often do you use social media?" There were eight response options, ranging from "I do not use social media" to "More than once an hour."

Depressive self-harm

A composite variable was created to capture depressive self-harm using six questions, a slightly expanded metric compared to Study 1. Here we combined responses about feeling sad or hopeless for more than two weeks in a row in the past 12 months (No/Yes), seriously considering attempting suicide (No/Yes), making a plan to attempt suicide (No/Yes), number of suicide attempts in the past 12 months (0, 1, 2–3, 4–5, 6 or more times), whether a suicide attempt resulted in injury (no attempt, attempt but no injury, yes), and responses to the question "During the past 30 days, how often was your mental health not good? (Poor mental health includes stress, anxiety, and depression)" (Never, Rarely, Sometimes, Most of the time, Always). Z-scores were calculated for each question, then averaged to create a composite such that higher scores indicated greater depressive self-harm ($\alpha = 0.79$ for these six items).

ACEs

ACE scores were calculated consistent with CDC recommendations (Swedo et al., 2024) to create a cumulative count of eight adverse childhood experiences, dichotomized as either absent or present. This includes questions about emotional abuse, physical abuse, sexual abuse, physical neglect, witnessing intimate partner violence, parent/guardian substance abuse, parent/guardian with poor mental health, and separation from a parent/guardian due to incarceration. ACE scores range from 0 to 8, and responses to these eight items showed appreciable reliability ($\alpha = 0.71$).

Results

Among all participants, 291 (3.0%) identified as transgender. This number was significantly higher among female participants (189/4677 = 4.0%) compared to male participants (102/4904 = 2.1%), χ^2 (1) = 31.3, $p < .001$. Table 3 reports differences between transgender and cisgender participants on all study variables, with follow-up analysis stratified by natal sex. In the overall sample, transgender participants did not use social media more than their cisgender counterparts ($d = -0.01$), but reported significantly more depressive self-harm ($d = 0.91$), delinquency ($d = 0.28$), and ACEs ($d = 0.70$). Among natal females, those who identified as transgender used social media less frequently than those who were cisgender ($d = -0.20$), but reported more depressive self-harm ($d = 0.59$) and ACEs ($d = 0.46$). Similarly, among natal males, those who identified as transgender reported significantly more depressive self-harm ($d = 1.36$), delinquency ($d = 0.52$), and ACEs ($d = 1.00$).

Our main preregistered analysis (Table 4, Model 1) was a binary logistic regression using simultaneous entry of all study variables to predict transgender identification. The overall model was significant, χ^2 (7) = 220, $p < .001$, Nagelkerke $R^2 = 0.096$. Social media use was negatively associated with the probability of transgender identification (OR = 0.944, $p = .069$) while controlling for age, sex, race, depressive self-harm, delinquency, and ACEs. This value did not exceed or come close to our SESOI (OR $\geq$ 1.44 or $\leq$ 0.69), nor did it meet traditional significance benchmarks. The latter three variables were significantly associated with transgender identity such that depressive suicidality and delinquency increased the odds, whereas being non-White lowered the odds.

Two additional preregistered models were run as robustness checks, splitting the data by natal sex. Binary logistic regression again predicted transgender identification using screen time, age, race, depressive self-harm, delinquency, and ACEs (Table 4, Models 2 & 3). Among females, the overall model was significant, χ^2 (6) = 79.4, $p < .001$, Nagelkerke $R^2 = 0.059$, with social media use, race, depressive self-harm, and ACEs all being significant predictors. Social media use did not exceed our SESOI, and was related to a decreased odds of transgender identity. An identical model was run among males, which was also significant, χ^2 (6) = 143, $p < .001$, Nagelkerke $R^2 = 0.157$, showing only race, depressive self-harm, and ACEs as significant predictors of transgender identification.

Table 3 Descriptive statistics and group comparisons for 2023 YRBS data

	Transgender M (SD)	Cisgender M (SD)	Levene's Test	*t*-test (two-tailed)[a]	*d* (95%CI)
Sample (n=9581)	***n*=291**	***n*=9290**			
Age	15.96 (1.23)	16.06 (1.21)	F=1.91, p=.167	$t_{(9579)}=-1.33$, p=.183	−0.08 (−0.20, 0.04)
Social Media Usage	6.12 (1.88)	6.14 (1.89)	F=0.28, p=.598	$t_{(9579)}=-0.14$, p=.888	−0.01 (−0.13, 0.11)
Depressive Self Harm	0.62 (0.83)	−0.01 (0.68)	F=34.34, p<.001	$t_{(302)}=15.22$, p<.001	0.91 (0.79, 1.03)
Delinquency	0.15 (1.04)	−0.01 (0.54)	F=67.81, p<.001	$t_{(295)}=2.55$, p=.011	0.28 (0.16, 0.40)
ACEs	4.24 (1.87)	2.94 (1.86)	F=1.13, p=.289	$t_{(9579)}=11.75$, p<.001	0.70 (0.58, 0.82)
Females (n=4677)	***n*=189**	***n*=4488**			
Age	15.96 (1.20)	16.00 (1.21)	F=2.73, p=.098	$t_{(4675)}=-0.375$, p=.708	0.03 (−0.17, 0.12)
Social Media Usage	6.04 (1.98)	6.38 (1.70)	F=0.51, p=.474	$t_{(4675)}=-2.63$, p=.009	−0.20 (−0.34, −0.05)
Depressive Self Harm	0.62 (0.76)	0.18 (0.74)	F=0.94, p=.332	$t_{(4675)}=7.99$, p<.001	0.59 (0.45, 0.74)
Delinquency	0.02 (0.88)	−0.07 (0.41)	F=36.62, p<.001	$t_{(191)}=1.39$, p=.167	0.20 (0.06, 0.35)
ACEs	4.21 (1.76)	3.32 (1.96)	F=4.52, p=.034	$t_{(208)}=6.80$, p<.001	0.46 (0.31, 0.61)
Males (n=4904)	***n*=102**	***n*=4802**			
Age	15.95 (1.29)	16.11 (1.20)	F=0.01, p=.909	$t_{(4902)}=-1.33$, p=.185	−0.13 (−0.33, 0.06)
Social Media Usage	6.28 (1.68)	5.92 (2.02)	F=2.12, p=.145	$t_{(4902)}=1.77$, p=.077	0.18 (−0.02, 0.37)
Depressive Self Harm	0.61 (0.95)	−0.18 (0.57)	F=89.2, p<.001	$t_{(103)}=8.33$, p<.001	1.36 (1.16, 1.56)
Delinquency	0.38 (1.27)	0.04 (0.63)	F=50.3, p<.001	$t_{(102)}=2.66$, p=.009	0.52 (0.32, 0.71)
ACEs	4.28 (2.05)	2.58 (1.69)	F=15.33, p<.001	$t_{(104)}=8.32$, p<.001	1.00 (0.81, 1.20)

[a]Welch's *t* reported when Levene's test showed unequal variances

Discussion

Analyzing data from the YRBS survey administered in Florida schools during 2019 and 2021 showed that screen time had a very small positive relationship with Transgender identification, even controlling for relevant covariates such as age, race, depressive self-harm, and delinquency, though the effect size did not exceed our smallest effect size of interest threshold. The 2023 national YRBS data showed that among all participants, social media use was not related to transgender identification when controlling for age, race, depressive self-harm, delinquency, and ACEs. In these more recent data, the only significant association between social media use and transgender identification was found among natal females, where greater use was related to *lower* odds of identifying as transgender. Additionally, the effect sizes of relationships between screen time and/or social media use and transgender identity did not exceed our SESOI (OR≥1.44 or ≤0.69), which means there is a very high chance these observations are due to noise, not a real effect. Although our study was not aimed at explicating the relationships between mental health, delinquency, or ACEs (Study 2), and transgender identification, our results are broadly consistent with other studies showing that adolescents who identify as transgender are more likely to be white (Twenge et al., 2024), to report elevated depressive self-harm (de Graaf et al., 2022) and (externalizing) delinquency type behaviors (e.g., van der Miesen et al., 2020), as well as analysis of the same 2023 CDC YRBS data showing elevated ACEs among transgender adolescents (Kreski & Hayes, 2025). These cross-sectional data preclude understanding why these associations exist, and we note that only race and depressive self-harm were reliably larger than our SESOI.

The focus of the present work was to understand whether screen time or social media use were associated with the probability of transgender identity, as has been proposed by some scholars seeking to explain contemporary increases in the number of adolescents identifying as transgender (Littman, 2018; Marchiano, 2017). If true, the present data would show a clear association between social media use and transgender identity. Our results provide either very weak (Study 1) or no (Study 2) evidence for the social contagion hypothesis for transgender identity. Alternatively, it is reasonable to interpret the results as showing no practical association between screen time and transgender identification, and hence evidence *against* the core prediction of the social contagion hypothesis. At best, one might view the present data as equivocal. The social contagion hypothesis may yet be viable, but requires better study designs that are not hampered by the salient limitations and caveats that temper any conclusions drawn from the data we present. We discuss diverging interpretations of the data before making recommendations about future work in this area.

One perspective on the present findings, favored by the second author, is that they represent evidence *against* the social contagion hypothesis. This interpretation relies on the small association between screen time and Transgender identification in Study 1 (OR = 1.13), equivalent to a

Table 4 Binary logistic regression predicting Transgender identification in 2023YRBS

Model	Predictor	(B)	SE	*p*	OR (95% CI)
1 (All participants)	Social media use	−0.058	0.032	0.069	0.94 (0.89, 1.01)
N=9581	Age	−0.063	0.050	0.210	0.94 (0.85, 1.04)
	Sex[a]	0.282	0.129	0.036	1.33 (1.02, 1.73)
	Race[b]	−0.503	0.129	<0.001	0.60 (0.47, 0.78)
	Depressive Self-harm	0.631	0.078	<0.001	1.88 (1.61, 2.19)
	Delinquency	0.048	0.083	0.567	1.05 (0.89, 1.23)
	ACEs	0.182	0.035	<0.001	1.20 (1.12, 1.29)
2 (Females)	Social media use	−0.129	0.039	<0.001	0.88 (0.81, 0.95)
n=4677	Age	−0.018	0.062	0.773	0.98 (0.87, 1.12)
(189 transgender)	Race[b]	−0.454	0.160	0.005	0.64 (0.46, 0.87)
	Depressive Self-harm	0.516	0.099	<0.001	1.68 (1.38, 2.03)
	Delinquency	−0.009	0.138	0.946	0.99 (0.76, 1.29)
	ACEs	0.125	0.042	0.003	1.13 (1.04, 1.23)
3 (Males)	Social media use	0.054	0.057	0.338	1.06 (0.95, 1.18)
n=4904	Age	−0.156	0.087	0.074	0.86 (0.72, 1.02)
(102 transgender)	Race[b]	−0.558	0.219	0.011	0.57 (0.37, 0.88)
	Depressive Self-harm	0.891	0.127	<0.001	2.44 (1.90, 3.13)
	Delinquency	−0.017	0.109	0.876	0.98 (0.79, 1.22)
	ACEs	0.301	0.060	<0.001	1.35 (1.20, 1.52)

[a] Sex coded as 0 = Female, 1 = Male

[b] Race coded as 0 = White, 1 = non-White

correlation of r=.034 (Effect Size Converter, 2024), and null (or opposite) effects in Study 2. Such small effects are below benchmarks for practical significance in psychological research (i.e., $r \geq .10$), and estimates of this size are frequently attributable to methodological noise rather than genuine associations (Ferguson & Heene, 2021). The same cannot be said for the covariates of race (OR = 0.60–0.66) and depressive self-harm (OR = 1.60–1.88), which represent correlations of r=−.11 – −0.14 and 0.13–0.17 respectively (Effect Size Converter, 2024). These factors are more prominent correlates of transgender identity, aligning with other research showing that such identities are more common among White adolescents in the United States (Twenge et al., 2024), and that gender diversity is correlated with elevated depression and suicidal ideation or attempts (e.g., de Graaf et al., 2022). Proposing that social media engagement induces adolescents to identify as transgender is an extraordinary claim. If social media is a salient factor in explaining the dramatic rise in the number of transgender and gender non-binary adolescents, this would show up in high quality datasets. Hence, the present association is unconvincing because it is indistinguishable from statistical noise (Study 1), or largely absent (Study 2). Related claims are made about the widespread impact of social media use on mental health, and here the data can be similarly underwhelming (Ferguson, 2024; Heffer et al., 2019; Jensen et al., 2019; Yang & Feng, 2024).

An alternative interpretation, favored by the first author, is that the present study may show (weak) evidence consistent with the social contagion hypothesis (Study 1), with Study 2 unable to show any meaningful associations between transgender identity and social media use due to ceiling effects in the latter. More importantly, methodological and measurement issues mean that the CDC data are not capable of adequately testing whether identifying as transgender is plausibly socially influenced. In short, the social contagion hypothesis is viable, but requires better tests. First, no cross-sectional dataset can determine causality. Even if the present data showed a relationship between social media use and transgender identity, this association alone wouldn't show whether exposure to media focussed on transgender issues precedes later identity, or if transgender adolescents simply sought out online support unavailable in their families or communities (e.g., Coyne et al., 2023; Nagata et al., 2024). Longitudinal designs would better establish temporal sequence, but these would also require measures that go beyond general screen time and social media use to probe the *type* of content consumed by adolescents. Additionally, the social contagion hypothesis has been proposed to apply predominantly to natal females, especially those with pre-existing mental health concerns, displaying no previous discomfort with their sex/gender, and suddenly declaring transgender or non-binary identities alongside others in their social milieu (Littman, 2018; Marchiano, 2017). Although we stratified analyses by reported sex, the YRBS measures this variable with a single question ("what is your sex?") and two response options (female; male). It is impossible to know whether transgender respondents interpreted this question in relation to their sex documented at birth, or to their current self-identification. Ideally, separate measures of sex (female/male/intersex) and gender (girl/boy/non-binary, etc.) should be available to ensure accurate classification of transgender identity vis-à-vis natal sex. The present measure of Transgender identity forced a binary categorization scheme on a group of gender diverse individuals who very likely show high variance in their gender identities and expression. Next, the measure of screen time used in Study 1 was crude at best, combining numerous mediums (TV, computer, smart phone) and activities (tv shows, video games, social media, etc.) into one metric. Study 2 improved upon this because the 2023 YRBS asks directly about social media use, but overall usage was high, averaging "Several times a day" in all groups. In short, using social media throughout the day is normative among adolescents, as is frequent switching between social media platforms (e.g., Wangqu et al., 2024), and the presently used self-report metric does not indicate the most widely encountered themes and topics. An adequate test of the social contagion hypothesis would require more precise measurement of both the types of media used, and the nature of this content. The hypothesis that social influence can result in transgender identification (Littman, 2018; Marchiano, 2017) does not predict that adolescent females would question or modify their gender identities after watching TV series, playing video games, or generally engaging with social media, but instead suggests that such changes occur after being immersed in forums (or peer groups) focussing on gender identity/expression and transition. Even precise cross-sectional measurements would suffer the same reverse-causality issues discussed above, so longitudinal data of this nature would be informative (e.g., Rawee et al., 2024). Even better study designs would track connections among participants longitudinally, given that social contagion is hypothesized to spread in local milieus. Here longitudinal data are vital to differentiate between transgender identities becoming more common in *existing* social clusters (e.g., friendship groups, classrooms, or even schools), as opposed to social groupings of likeminded gender variant adolescents coalescing well after a transgender identity is established.

We urge readers to be cautious in interpreting our findings—the weak or null effects leave the plausibility of the social contagion hypothesis unclear. Instead, we want to draw attention to study designs that would offer stronger tests of this hypothesis, with the hope that better data brings understanding to a scientific issue that impacts a growing number of gender diverse adolescents and their families. These methodological considerations are necessary given that the hypothesis cannot be tested experimentally, as can be done for the more general relationship between social media use and mental health, albeit with caveats (Ferguson, 2024). First, natal sex should be measured independent from current gender identity to precisely know participants' (trans)gender status, and measures of gender identity should better capture the diversity of this trait. Identifying age at which transgender identification first emerged would also be desirable. Next, recalled measures of gender-typed behavior and play in childhood (e.g., Bailey & Zucker, 1995) should be employed, giving insight into whether sex-atypical behavior, interests, and/or identity have been longstanding or more recent. Given that same-sex attraction is a widely documented correlate of sex-atypical behavior in childhood (Bailey & Zucker, 1995; Bailey et al., 2016), and same-sex attraction is related to childhood onset gender dysphoria (Lawrence, 2010; Ristori & Steensma, 2016; Zucker, 2005), participants' sexual orientation should also be measured. Here we note that the YRBS does measure sexual orientation *identity* (heterosexual, gay or lesbian, bisexual), but identity-based measures create ambiguity regarding who transgender respondents find sexually attractive (e.g., a transgender man may identify as "heterosexual" because he is attracted to women, but this would constitute same-sex attraction relative to natal sex). It would be preferable to measure sexual attraction to males and females separately (e.g., Legate & Rogge, 2019). Last, as noted above, precisely measuring digital media platforms, types of media consumed, and prominent transgender individuals in friendship circles, classrooms, or schools, is vital. Using the variables outlined above, along with relevant covariates such as depressive self-harm, researchers would be in the position to adequately test the social contagion hypothesis as outlined by Littman (2018). This hypothesis likely does not apply to natal males or females who are same-sex attracted, display sex-atypical behavior in childhood, and express early wishes to behave/identify as the opposite sex (Bailey & Zucker, 1995; Zucker et al., 2011; Zucker, 2017). Similarly, adolescent males who report either opposite sex attractions (or occasionally bisexuality) and begin cross-gender behavior and identification in adolescence or early adulthood are less likely to be influenced by social factors (Lawrence, 2010; Zucker & Lawrence, 2009). The social contagion hypothesis is largely constrained to natal females who have not previously expressed gender non-conformity (Littman, 2018), so a proper test of the hypothesis requires isolating this group and comparing them to their cisgender counterparts to understand the purported correlation between social media engagement and peer influences, and transgender identification (see also Bailey & Blanchard, 2017). This comparison is impossible in the absence of adequate measurement. At the same time, it is warranted to point out that the social contagion hypothesis' constraint to natal females (e.g., Littman, 2018) was necessarily post-hoc, and that it is a theory constrained to a particular group only *after* the ecological observation of behaviors in that group, not an a priori theory. Such post-hoc fitting of data to a theory raises the high potential of the ecological fallacy, making such a theory weaker than one developed a priori. Nonetheless, we believe it is an interesting theory worthy of further exploration, particularly using carefully controlled and preregistered analyses.

It is further worth noting that even if theories of social contagion via social media prove falsified by good data, this does not mean some version of social contagion may not occur through some other means (such as via parents, teachers, medical professionals, school counsellors, real-life peer groups, news media, etc.) However, we implore scholars to develop rigorous, falsifiable, specific theories that can be tested with preregistered studies, which we observe has not occurred to date. Beyond these empirical considerations are relevant ethical ones. First, our present data analysis was merely an attempt to understand whether social media use correlates with transgender identity in large national datasets. As such it would be inappropriate to use these results to justify any specific government, school, or transgender healthcare policy. Second, empirical tests of the social contagion hypothesis, and data refuting or supporting its plausibility, have no bearing on whether transgender persons should be socially accepted or offered compassionate evidence-based care. The former is a scientific endeavour, and we have noted the muted evidence regarding the hypothesis, whereas the latter considerations pertain to social values and human rights which are not determined by any scientific result.

Limitations

We recognize several other limitations of our data, though some of these are discussed above in detail. First, our data are cross-sectional and no causal inferences can be made. Not being longitudinal in nature, these data do not allow us to test long-term associations between social media and transgender identity, although it is worth noting that in most realms of media effects, longitudinal designs tend to produce weaker, not stronger effect sizes than cross-sectional designs (Drummond et al., 2020). Nonetheless, longitudinal designs would certainly be welcome in this area. Second, it is possible that the operational variables here may miss some nuances. For instance, more nuanced and sophisticated measures of gender identity would be valuable. So too would variables of social media use that move away from time spent and, instead, focus on how youth use social media (e.g., passive vs. active, gender-related content vs. general content, etc.).

Conclusion

The number of transgender adolescents has grown in recent decades (Thompson et al., 2022; Twenge et al., 2024), and the reasons for these trends remain unclear (Zucker, 2019). Here we report no practical associations between screen time (Study 1) or social media use (Study 2) and Transgender identity over and above covariates such as age, race, and depressive self-harm. The effect sizes did not meet our SESOI threshold, and in Study 2 were opposite the predictions of the social contagion hypothesis, and both effects are best interpreted as methodological noise. Given our data, claims that social contagion leads to widespread transgender identity (e.g., Shrier, 2020) should be tempered until stronger evidence is available. So too should claims that the hypothesis is entirely debunked (Serano, 2023). Our findings come with many limitations and should not be overinterpreted in the absence of stronger tests of the purported association between social media and transgender identity.

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Data availability Archival Data from the Youth Risk Behavior Survey (YRBS) are available from the Center for Disease Control (CDC, 2025; https://www.cdc.gov/yrbs/methods/index.html). Data for the Florida YRBS are publicly available by request from the Florida department of health at https://www.floridahealth.gov/statistics-and-data/survey-data/index.html. Our pre-registered data analysis plan can be found at https://aspredicted.org/GPW_GK2.

Declarations

Ethical approval and informed consent Ethical approval and procedures, including informed consent, are described by the CDC (Center for Disease Control (CDC, 2025; https://www.cdc.gov/yrbs/methods/index.html). Secondary data analysis of archival data is exempt from local IRB approval.

Conflict of interest The authors declare that they have no conflict of interest.

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