Journal of Experimental Psychology: General

Of Pandemics, Politics, and Personality: The Role of Conscientiousness and Political Ideology in the Sharing of Fake News

M. Asher Lawson and Hemant Kakkar

Online First Publication, October 25, 2021. http://dx.doi.org/10.1037/xge0001120

AMERICAN PSYCHOLOGICAL ASSOCIATION

Journal of Experimental Psychology: General

ISSN: 0096-3445

https://doi.org/10.1037/xge0001120

Of Pandemics, Politics, and Personality: The Role of Conscientiousness and Political Ideology in the Sharing of Fake News

M. Asher Lawson and Hemant Kakkar

Management and Organizations Academic Area, The Fuqua School of Business, Duke University

Sharing misinformation can be catastrophic, especially during times of national importance. Typically studied in political contexts, the sharing of fake news has been positively linked with conservative political ideology. However, such sweeping generalizations run the risk of increasing already rampant political polarization. We offer a more nuanced account by proposing that the sharing of fake news is largely driven by low conscientiousness conservatives. At high levels of conscientiousness there is no difference between liberals and conservatives. We find support for our hypotheses in the contexts of COVID-19, political, and neutral news across eight studies (six preregistered; two conceptual replications) with 4,642 participants and 91,144 unique participant-news observations. A general desire for chaos explains the interactive effect of political ideology and conscientiousness on the sharing of fake news. Furthermore, our findings indicate the inadequacy of fact-checker interventions to deter the spread of fake news. This underscores the challenges associated with tackling fake news, especially during a crisis like COVID-19 where misinformation impairs the ability of governments to curtail the pandemic.

Keywords: conscientiousness, conservatives, COVID-19, fake news, liberals

Supplemental materials: https://doi.org/10.1037/xge0001120.supp

Since the 2016 U.S. Presidential Election, researchers have found evidence of the far-reaching malicious impact of misinformation on public opinion (Allcott & Gentzkow, 2017; Grinberg et al., 2019; Guess et al., 2020; Lazer et al., 2018; Vosoughi et al., 2018). However, the current "fake news pandemic" (Rajan, 2020) has implications beyond voting preferences. False and misleading advice on COVID-19 has both spread the virus and claimed lives: Hundreds of people died in Iran after ingesting methanol in attempts to combat COVID-19, after reading erroneous medical advice on social media (The Associated Press, 2020).

Given the myriad negative consequences associated with the spread of misinformation, a growing body of research has examined factors that encourage sharing fake news (Guess et al., 2019). Factors such as less analytical thinking (Pennycook & Rand, 2020, 2019a), prior exposure to a false story (Pennycook et al., 2018), and the absence of tags disputing a story (Pennycook, Bear, et al., 2020) have all been shown to predict the spread of fabricated stories. However, across multiple studies a positive association has emerged between conservative political ideology and dissemination of falsities (Bago et al., 2020; Bail et al., 2018; Grinberg et al., 2019; Guess et al., 2020; Lazer et al., 2018; Pennycook et al., 2018; Pennycook & Rand, 2019a, 2019b, 2020; Scheufele & Krause, 2019; Vosoughi et al., 2018). Similarly, others have found that political conservatives flock to factually dubious news websites more than liberals (Guess et al., 2020). Research on "prebunking"—interventions aimed at preemptively combating the spread of misinformation—has found that although both liberals and conservatives learned to spot and resist misinformation at equal rates, conservatives rated fake headlines as more reliable than liberals prior to intervention (Basol et al., 2020; Roozenbeek & van der Linden, 2019).

Further research casts light on why conservative political ideology might be associated with sharing fake news. Political conservatives are more likely to endorse specific conspiracy theories and espouse conspiratorial worldviews in general (van der Linden et al., 2021). A tendency toward conspiratorial thinking could lead to misinformation seeming more feasible and relevant. Additionally, conservatives are more likely to associate the mainstream media with fake news, and dismiss contra-partisan news outlets (e.g., CNN) as fake at a higher rate than liberals (van der Linden et al., 2020). This greater distrust in ideologically opposed media could further increase the endorsement of ideologically concordant falsehoods among conservatives. In a similar vein, Jost et al. (2018) proposed that conservatives might be more motivated to endorse false content to satisfy their higher need for shared reality. Overall, this corpus of research provides convincing empirical evidence that, on average, conservatives share

M. Asher Lawson https://orcid.org/0000-0001-5782-3583

Hemant Kakkar https://orcid.org/0000-0002-1483-3738

All data and materials are available at https://osf.io/ahdsf/. This work was presented at the Academy of Management (2020) and the Society for Personality and Social Psychology's (2021) annual meetings, and was available prior to publication on PsyArxiv (https://psyarxiv.com/ves5m/).

Correspondence concerning this article should be addressed to Hemant Kakkar, Management and Organizations Academic Area, The Fuqua School of Business, Duke University, 100 Fuqua Drive, Durham, NC 27708, United States. Email: hemant.kakkar@duke.edu

misinformation at a higher rate than liberals. This has led some to conclude that individuals associated with conservative political ideology are more prone to disseminating misinformation (Bago et al., 2020; Bail et al., 2018; Grinberg et al., 2019; Guess et al., 2020; Lazer et al., 2018; Pennycook et al., 2018; Pennycook & Rand, 2019a, 2019b, 2020; Scheufele & Krause, 2019; Vosoughi et al., 2018).

However, such a conclusion runs the risk of denigrating all conservatives. Despite the main effect of conservative political ideology, it is possible that the effects are more nuanced and that the above generalization is not applicable to everyone who identifies as a conservative. Although conservatives' emphasis on group loyalty and asserting the dominance of their group over others (Jost et al., 2003; Kugler et al., 2014) might increase their tendency to share fake news, conservatives' broad moral value system also respects authority, tradition, and the sanctity of institutions (Graham et al., 2009; Jost et al., 2018). Hence, conservatives who regard themselves as guardians of such values will be less likely to engage in the spread of rumors that could compromise the system. In short, conservatives are a hetereogenous group. Additionally, both liberals and conservatives display equal levels of partisan bias with regards to ideologically concordant information (Ditto et al., 2019). Thus, it is unlikely that the members of one political group are on average more likely to engage in motivated cognition than the other. Therefore, making general statements about all conservatives behaving in a counterproductive manner in our opinion runs an unnecessary risk of furthering political polarization and intergroup conflict. A more balanced approach in studying the role of political attitudes in the sharing of fake news is required not only to combat political polarization, but also to allow for greater political diversity in scholarly research (Duarte et al., 2015). To that end, we draw on principles from the personality literature, as personality is a key determinant of individual behavior beyond one's political or general attitudes, to offer a more nuanced understanding of the relationship between conservative political ideology and the sharing of fake news.

The Moderating Role of Conscientiousness

One of the most studied and accepted personality taxonomies—The Big Five (McCrae & Costa, 2008; Soto & John, 2017)—organizes personal traits into five factors: Open-mindedness, Conscientiousness, Extraversion, Agreeableness, and Negative emotionality. The Big Five personality traits have been shown to influence a range of attitudes and behaviors including prejudice (Sibley et al., 2012), cooperation (Koole et al., 2001), intergroup conflict (Koehn et al., 2019), and morality (Tybur et al., 2009). Despite the generality of all five personality factors in predicting individual behavior, we limit our focus to conscientiousness. Conscientiousness is defined as "individual differences in the propensity to follow socially prescribed norms for impulse control, to be goal directed, to plan, and to be able to delay gratification and to follow norms and rules" (Roberts et al., 2009, p.369). Conscientiousness consists of a broad domain of positive traits such as orderliness, impulse control, conventionality, reliability, industriousness, and virtue (Roberts et al., 2009). Unsurprisingly, highly conscientious individuals follow the rules of society, maintain social decorum, and think before acting (Jackson et al., 2010). Higher conscientiousness is associated with less drug abuse (Bogg & Roberts, 2004), less counterproductive behavior at work (Roberts et al., 2007), better management of relationships that are prone to conflict (Finkel & Campbell, 2001), and greater exhibition of integrity and moral character (Cohen et al., 2014).

Individuals who are high on conscientiousness tend to be orderly, and their ability to delay gratification confers a whole host of life benefits. Specifically, conscientiousness leads to better control over one's behavior. Such people report higher levels of perceived control over intended actions (Courneya et al., 1999), a stronger locus of control (Saklofske et al., 2007), and better self-management (Gerhardt et al., 2007). Conscientiousness is strongly related with guilt proneness but not guilt experience—highly conscientiousness people feel guilt more acutely when they misbehave, but are less likely to act in a way that triggers the experience of guilt (Fayard et al., 2012). Importantly, a recent meta-analysis found guilt proneness to be one of the most reliable indicators of prosocial behavior (Thielmann et al., 2020). In sum, both the dimensions of orderliness and impulse control suggest that individuals who are associated with greater conscientiousness should be less inclined to behave irresponsibly by sharing fake news, which sows disorder and reflects a lack of personal control (Petrocelli et al., 2020).

Although political conservatives on an average might be more inclined to share fake news relative to political liberals, conservatives who are higher in conscientiousness and thus more likely to control their impulses should be less likely to share fake news. We argue that the diligent and prudent attributes of highly conscientious individuals will curtail the tendency to propagate falsehoods. Such individuals will rely less on their subjective assessment of news accuracy and expend greater effort and care in discerning news veracity, which may reduce the likelihood of sharing false stories. Highly conscientious people are also more likely to engage in like-minded discussions with contra-partisans (Boulianne & Koc-Michalska, 2021), which might discourage them from participating in negative partisanship behaviors. Conversely, low conscientiousness conservatives who are reflexively driven by their partisan loyalties and assertion of their group's supremacy over others and less restrained by other conservative values will be more likely to share fake news. Accordingly, we predict a negative interaction between political ideology and conscientiousness on the likelihood of sharing fake news, such that higher conscientiousness weakens the positive association between conservative political ideology and the sharing of fake news. We expect no differences in sharing behaviors among liberals and conservatives at higher levels of trait conscientiousness.

Low Conscientiousness Conservatives and the Desire for Chaos

We further contend that the behavior of low conscientiousness conservatives is motivated not only by vehemently promoting the interests of their group, but also by denigrating other rival groups. Such a staunch inclination to elevate one's group at the expense of other political outgroups is an act of negative partisanship—a reality that has become increasingly common due to the exponential rise of intense political polarization since the start of the 21st century (Abramowitz & Webster, 2016; Van den Bos et al., 2007; Westwood et al., 2018). Because conservatives generally score higher on social dominance orientation—a set of beliefs that acknowledges and supports hierarchical differences in society

(Kugler et al., 2014)—they may be more likely to criticize other groups to defend their own (Jost et al., 2003). Relative to liberals, conservatives are also more vigilant in perceiving social threats to their group (van Leeuwen & Park, 2009), which can further increase their tendency to actively denounce other groups and outgroup members. This desire to promote the status of one's group at the expense of other groups and outgroup members can lead to a generally hostile mindset, labeled a "need for chaos" (Arceneaux et al., 2021). The need for chaos is described as a drive to disrupt and destroy the existing order or established institutions in an attempt to secure the superiority of one's own group over others. Such a mindset is especially salient when dominance-oriented individuals feel they are being marginalized and rejected by the broader cultural environment (Arceneaux et al., 2021; Krizan & Johar, 2015; Twenge & Campbell, 2003).

Given the lack of orderliness, diligence, and self-control associated with low conscientiousness individuals, coupled with the high social dominance orientation and group loyalty among conservatives, we contend that low conscientiousness conservatives will be more likely to entertain beliefs and engage in behaviors that seek to cause chaos as a means to defend their group. Indeed, existing research has shown that people are more willing to believe and share outlandish conspiracy theories when it helps them to achieve a positive image of their group, its dominance, and its existence (Douglas et al., 2019; Roberts et al., 2009). Likewise, the desire to cause chaos also leads to less support for outgroups such as immigrants, and a greater desire to increase one's social status and alter the current power structure, especially when political polarization is rampant (Arceneaux et al., 2021; Van Bavel & Pereira, 2018). Consequently, we predict that the interaction effect of conservative political ideology and conscientiousness on the sharing of fake news will be mediated by this desire for chaos.

Beyond Accuracy Beliefs and News Concordance

Furthermore, recent research has highlighted that the dissemination of fake news is largely driven by people's inattention to accuracy. Once accuracy beliefs are primed either implicitly or explicitly, individuals are relatively more judicious when it comes to the sharing of fake news (Pennycook, McPhetres, et al., 2020; Pennycook et al., 2021). However, our proposed effect, where low conscientiousness conservatives share fake news as a result of an elevated desire for chaos, is indicative of a motivated process. Specifically, when low conscientiousness conservatives perceive fake news as a means of furthering their social goals (Douglas et al., 2017) and sowing seeds of destruction (Arceneaux et al., 2021), the accuracy of news stories should play a smaller role in determining their intentions to share such stories. In other words, people who pursue general destruction to defend their ingroup should indicate higher subjective assessments of the accuracy of fake news, as long as it serves the agenda of their group, which in turn will predict the sharing of such news. Thus, when motivated to believe false information as accurate, priming individuals with accuracy beliefs might not be enough to deter the spread of misinformation. Rather, such motivated individuals will perceive false news as subjectively more accurate and hence share falsehoods at a higher rate regardless of accuracy primes. Thus, in all our studies and consistent with recent work that has demonstrated reduced sharing of misinformation after participants are reminded of news accuracy (Pennycook, McPhetres, et al., 2020), we test our interaction hypothesis after priming individuals with accuracy beliefs. Additionally, to further demonstrate that sharing of fake news is a motivated process, we also examine whether subjective accuracy mediates the interactive effect of political ideology and conscientiousness on the sharing of fake news. In so doing, we offer a conservative test of our hypothesis and find evidence extending beyond a well-known cognitive mechanism based on accuracy, in favor of a motivated process as the underlying driver of the observed behavior.

Finally, we also examine whether the sharing of fake news is explained by *news concordance*—that is, are stories that align with people's political ideologies shared more—or *news discordance*—that is, do people share news less when it is misaligned with their political ideology. Previous research has suggested that news concordance increases the proliferation of fake news as opposed to discordance reducing it (Bago et al., 2020; Martel et al., 2019). However, much of that research was conducted in the domain of political news with analysis lacking comparison to a neutral condition. Without a neutral condition, one cannot reliably distinguish whether concordance increases or discordance decreases the proliferation of fake news. Using COVID-19 as backdrop, we examine the effects of news concordance and discordance on the sharing of fake news.

Overview of Studies

Across eight studies (six preregistered and two conceptual replications), consisting of 4,642 participants and 91,144 unique participant-news observations, we examine the joint effect of political ideology and personality on fake news sharing behavior. To offer greater robustness and confidence in our findings, we also account for existing explanations focused on deliberative thinking, by including participants' scores on the Cognitive Reflection test (Frederick, 2005) in our regression analyses. Our studies primarily sampled online participants from Amazon's Mechanical Turk platform (MTurk), which has been shown to be a reliable resource for research into political ideology (Coppock, 2019; Krupnikov & Levine, 2014; Mullinix et al., 2015). Furthermore, researchers have suggested that online samples such as MTurk are representative of the people who are most likely to come across fake news stories—and therefore constitutes the appropriate population in which to study fake news behavior (Pennycook & Rand, 2019b). However, given the documented decrease in data quality on MTurk since 2018 (Chmielewski & Kucker, 2020; Kennedy et al., 2020), we implemented a battery of additional steps to ensure the quality of our MTurk samples, and included an additional replication of our interactive effect of political ideology and conscientiousness on the sharing of fake news using a different platform (Prolific) in our online supplemental materials (see Study S1).

Study 1 tested whether conscientiousness and political ideology interactively predict the sharing of fake news related to COVID-19. Study 2 replicated this effect using political stimuli from the 2016 Presidential election. Studies 3–4 demonstrated the robustness of our effects in the presence of explicit fact-checker warnings, and whether participants updated their intended behaviors in response to such warnings. Studies 5–6 revealed desire for chaos as the mechanism driving the interaction. We preregistered six of the eight studies (Study 2 and Study S2 were conceptual replications), determining sample sizes, exclusion criteria, and the analytical approach in advance. The study materials, preregistration,

data, and analysis code are available at https://osf.io/ahdsf/. All research was approved by the Institutional Review Board of Duke University.

Sample Selection and Exclusion Criteria

Across all our studies, we restricted our samples to only U.S. participants on both Qualtrics and CloudResearch (Litman et al., 2017), and those who were not using mobile devices to respond to our surveys. We further excluded non-U.S. IP addresses using a manual lookup of the sample IPs. This latter manual exclusion was not preregistered. In the online supplemental materials, we provide results using no exclusion criteria, as well as extended exclusion criteria that removed participants flagged as suspicious by the application developed in Kennedy et al. (2020) to mitigate concerns about data quality among online samples. None of our conclusions were affected (Table S21 in the online supplemental materials). Further, using the CloudResearch platform we limited potential participants to those with a HIT acceptance rate of 94% or more. Automatic bots that produce low quality data will likely have lower acceptance rates. We also limited access to users who had completed between 50 and 50,000 HITs on MTurk and posted all of our HITs on a weekday between 10am and 11am EST to ensure participants were more representative of the broader U.S. population. Finally, we included a comprehension check, where participants were asked to identify which item was soft out of a list of several items (e.g., "toaster," "pillow, "road"). Failure to identify "pillow" as the correct answer would screen out potential bots and low quality participants. Overall, our sample selection procedure identified U.S. participants, and we took several steps to screen out bots and low-quality data.

Study 1

Study 1 tested our key hypothesis: whether conscientiousness and political ideology interactively predict the tendency to share fake news. We also examined if news concordance increases the sharing of fake news, news discordance decreases it, or both, in comparison to a control condition containing ideologically-neutral news.

Method

The preregistration including information on sample size, exclusion criteria, hypotheses and analysis can be found at https://osf.io/mvu2r/.

Participants

We recruited 499 participants from MTurk. We excluded 11 participants for having a non-U.S. IP address resulting in a final sample of 488 participants (M_{age} = 39.6y, 55.1% females, 1% nonbinary). Participants were paid $1.01.

Procedure

Participants rated the accuracy of 12 real and 12 fake COVID-19 news stories, and how likely they were to share them on social media. Within each group of 12 real and fake stories, four were conservative-leaning, four liberal-leaning, and four neutral. In all, there were eight each of conservative-leaning, liberal-leaning and neutral news stories. These stories were found by searching fake news fact-checker websites (e.g., Snopes) and mainstream media outlets. An example of a fake conservative-leaning headline is "Even during coronavirus crisis, liberal media cannot resist spreading lies," whereas an example of a fake liberal-leaning headline is "Trump is planning to bail out his corporate pals for pandemic losses—and leave taxpayers with the bill: columnist." The full list of news stories can be found at https://osf.io/ahdsf/.

We randomly presented the 24 news stories and also counterbalanced such that half of the participants responded to the personality questionnaire before seeing the news stories, and the other half after seeing the news stories. After rating the 24 stories and responding to the personality measure, participants answered items corresponding to the cognitive reflection test (CRT), their attitudes toward COVID-19, political ideology, and other demographic information.

Pretest

To validate our classification of stories by ideological content, we conducted a pretest (N = 99) where participants rated the ideological alignment of all of the 24 stories. Following prior research, participants were asked to assume the stories were entirely accurate (Pennycook & Rand, 2019b) and then judge "how favorably aligned would it be with Democrats' versus Republicans' ideology" (on a 5-point scale from *more favorable to Democrats* to *more favorable to Republicans*). This methodology aimed to separate participants' views of the veracity of stories from their views regarding the ideological alignment of stories (i.e., a Republican might view a Democratic-leaning story as very untrustworthy, but still be able to identify that it is aligned with Democratic ideology). Participants' ratings supported our classification of stories (M_{Dem} = 2.27, $M_{Neutral}$ = 2.98, M_{Rep} = 3.48): Liberal-leaning stories were more aligned with Democratic ideology, and conservative-leaning headlines were more aligned with Republican ideology. Neutral stories were rated very close to the center of the scale[1] and were significantly different from both Democratic aligned (t = 6.23, $p < .001$) and Republican favoring (t = −6.07, $p < .001$) news stories. Although the perceived ideological alignment of the Democrat-consistent stories (i.e., their distance from the neutral stories) was slightly greater than the perceived alignment of Republican-consistent stories, this difference was not statistically significant (t = 1.71, p = .090). Overall, analysis of the pretest supported our ideological classification of the news stories.

Measures

We measured participants' personality using the BFI-2 personality inventory (Soto & John, 2017). This 60-item questionnaire classifies individuals on Big-5 personality factors using 12 items for each factor based on their responses on a five-point scale ranging from *Disagree strongly* to *Agree strongly*. We aggregated participants' responses into the five factors of open-mindedness (α = .87), conscientiousness (α = .89), extraversion (α = .85), agreeableness (α = .87), and negative emotionality (α = .92). Political orientation was measured using two instruments. First, participants

[1] A one sample t test comparing the neutral ratings to the null hypothesis M = 3 was not significant (t = −.493, p = .623), supporting the neutral classification of these stories.

reported their political ideology on a seven-point scale ranging from 1 = *Very liberal* to 7 = *Very conservative* (Graham et al., 2009). Thus, higher values on this measure represented individuals who identified more with conservative ideology. Second, participants in a forced choice indicated: "If you had to choose between Democrats and Republicans, who would you prefer?" This measure is consistent with existing procedures (Pennycook, Bear, et al., 2020; Pennycook et al., 2018; Pennycook & Rand, 2019b). Our primary analysis relied on participants' ideology as measured by identification on a left-right continuum (Jost, 2006; Jost et al., 2008); however, obtaining an additional measure of partisanship as a proxy for left-right ideology allowed us to distinguish between politically concordant and discordant news stories. Although scholars have rightly noted that partisan commitments differ from ideology and are associated with party affiliation and political identities (e.g., Krishna & Sokolova, 2017; Oyserman & Schwarz, 2017), empirically the two are highly correlated in the contemporary, polarized United States (Jost, 2017). We report our results with the continuous measure of ideology (in line with our preregistration), but the results remained consistent and significant with the categorical measure as well.

Before indicating their tendency to share a news story, participants first rated their subjective accuracy of news stories on a four-point continuous scale (1 = *Not at all accurate* to 4 = *Very accurate*) by responding to the following item, "To the best of your knowledge, how accurate is the claim in the above headline?" Directing participants' focus to news accuracy in this manner reduces the likelihood of respondents sharing fake news due to inattention (Pennycook, McPhetres, et al., 2020). Following this, participants indicated their desire to share the news using a single item: "Would you consider sharing this story online (e.g., through Facebook or Twitter)?" The responses for this item were *No*, *Maybe*, and *Yes*. Consistent with prior research (Pennycook, Bear, et al., 2020; Pennycook et al., 2018; Pennycook & Rand, 2019b), the variable 'share' took the value of 0 if a respondent indicated *No*, and 1 if they indicated *Maybe* or *Yes*. We followed this convention for two primary reasons. First, to facilitate meaningful comparison of our results to that of extant research. Second, both *Maybe* and *Yes* indicate an openness to sharing a falsehood.

We also included several control variables. We collected responses to the CRT (Frederick, 2005), a three-item scale measuring the extent to which participants' cognitive styles were deliberative. An example item of the CRT is the bat-and-ball problem, "A bat and a ball cost $1.10 in total. The bat costs $1.00 more than the ball. How much does the ball cost? (in cents)." We also measured the CRT-2, a non-numeric test of respondents' cognitive styles. The CRT-2 (Thomson & Oppenheimer, 2016) is a four-item scale including items such as, "If you're running a race and you pass the person in second place, what place are you in?" This scale is similar to CRT but removes the heavy dependency on numeracy. We collapsed these two scales to make a seven-item measure of general cognitive reflection, with scores ranging from 0 to 7. We measured participants' attitudes toward the COVID-19 virus to account for any variance in sharing of news based on such attitudes. Our items fell into two categories: how personally worried they were and their beliefs about the extent to which the virus was dangerous. We collapsed all six items to create a single measure of attitudes toward COVID-19, where a higher score indicated a greater perception of COVID-19 severity (α = .92). We added two additional questions—the extent to which the threat of the virus was exaggerated, and support for federally mandated quarantining—to account for any partisan differences. Finally, we controlled for participants' age and gender (1 if male, 0 otherwise).

Results

Because each participant saw 24 different news stories of varying ideological content, we wanted to take into account any variance based on story level factors (i.e., the ideological content of the story, and its alignment with the participant's ideology). Thus, we preregistered our intention to perform analysis at the news story level, resulting in 11,712 observations. Notably, all of our results replicated when using linear regression to predict the total number of stories shared out of 24 (see Table S22 in the online supplemental materials for further details). Our key dependent variable—the sharing of news—was categorical in nature, hence we ran logistic regression analysis.[2] To provide a more conservative test of our hypothesis, we used cluster-robust standard errors that only assume independence between clusters (Liang & Zeger, 1986). Our main variables (political ideology and conscientiousness) were not mean-centered: This choice does not affect statistical inference regarding the significance of the interaction term, which was our effect of interest. We kept this procedure constant across studies.

Participants who identified as conservative were more likely to share news stories (b = .154, $p < .001$, model 1, Table 1), whereas participants who were higher on conscientiousness were less likely to do so (b = −.648, $p < .001$, model 1). These effects were robust to the inclusion of several control variables (model 4), including cognitive reflection.

Interaction Effect

The interaction between political ideology and conscientiousness was statistically significant and negative, with (b = −.191, $p < .001$, model 5) and without (b = −.231, $p < .001$, model 2) the inclusion of control variables. The interaction was also robust after including interactions between political ideology and all of the other big five personality variables (see Table S1 in the online supplemental materials). Figure 1 plots the interaction pattern from model 2. We used simple slope analysis across all our studies to interpret the interaction. The resulting analysis revealed that the effect of political ideology on the likelihood of sharing a story was significant at one standard deviation below the mean level of conscientiousness (b = .07, $p < .001$), but not at one standard deviation above the mean (b = −.01, p = .44).

We also found that political ideology and conscientiousness interactively predicted participants' accuracy beliefs both without (b = −.070, $p < .001$, see Table S3 in the online supplemental materials) and with control variables (b = −.062, $p < .001$). We next examined whether subjective accuracy mediated the interactive effect of political ideology and conscientiousness on the sharing of fake news

[2] We preregistered our intentions to estimate our regression models using Generalized Estimation Equations. However, when following scholarly guidelines to minimize the Quasi Information Criterion (QIC), as per Hardin and Hilbe (2003), this procedure implied an "independence" correlation structure within clusters, which is equivalent to logistic regression.

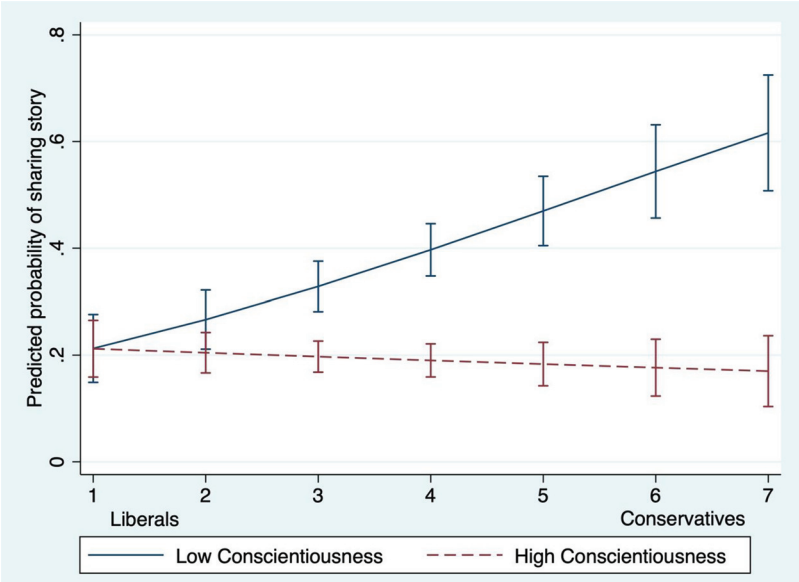
Table 1
Logistic Regression Models Predicting Binary Outcome 'Share' With Clustered Standard Errors, Unstandardized Betas, and 95% Confidence Intervals (Study 1)

Model	1	2	3	4	5	6	7	8
Political ideology (PI)	.154*** [.08, .23]	1.026*** [.54, 1.51]		.118** [.04, .20]	.844*** [.44, 1.25]			
Conscientiousness (C)	−.648*** [−.84, −.45]	.229 [−.22, .67]		−.687*** [−.91, −.46]	.023 [−.40, .45]	−.677*** [−.91, −.45]	−.622*** [−.86, −.38]	−.491*** [−.74, −.24]
PI × C		−.231*** [−.35, −.11]			−.191*** [−.29, −.09]			
Fake (F)			−.429*** [−.49, −.36]	−.446*** [−.51, −.38]	−.453*** [−.52, −.38]	−.444*** [−.51, −.38]	−.444*** [−.51, −.38]	.424 [−.21, 1.06]
Extraversion			.451*** [.26, .65]	.471*** [.28, .66]	.478*** [.29, .67]	.512*** [.32, .71]	.512*** [.32, .71]	.515*** [.32, .71]
Agreeableness			−.298* [−.53, −.06]	−.056 [−.30, .19]	−.031 [−.28, .22]	−.065 [−.31, .18]	−.065 [−.31, .18]	−.065 [−.31, .18]
Negative emotionality			.144 [−.03, .32]	−.034 [−.22, .15]	−.013 [−.20, .17]	−.016 [−.20, .17]	−.016 [−.20, .17]	−.015 [−.20, .17]
Openness			−.508*** [−.70, −.31]	−.320** [−.52, −.12]	−.315** [−.52, −.11]	−.396*** [−.59, −.20]	−.396*** [−.59, −.20]	−.398*** [−.59, −.20]
Attitudes towards COVID-19			.266*** [.11, .42]	.329*** [.17, .49]	.315*** [.14, .49]	.277*** [.12, .43]	.277*** [.12, .43]	.279*** [.12, .44]
General cognitive reflection			−.195*** [−.26, −.13]	−.180*** [−.24, −.12]	−.176*** [−.24, −.11]	−.193*** [−.26, −.13]	−.193*** [−.26, −.13]	−.194*** [−.26, −.13]
Age			.001 [−.01, .01]	−.000 [−.01, .01]	.000 [−.01, .01]	.002 [−.01, .01]	.002 [−.01, .01]	.002 [−.01, .01]
Male			.165 [−.12, .45]	.103 [−.18, .38]	.103 [−.18, .39]	.125 [−.16, .41]	.125 [−.16, .41]	.126 [−.16, .41]
Education			.117 [−.01, .24]	.116 [−.01, .24]	.092 [−.04, .22]	.109 [−.02, .24]	.108 [−.02, .24]	.109 [−.02, .24]
News - Liberal			−.143* [−.25, −.03]	−.148* [−.26, −.03]	−.151* [−.27, −.03]			
News - Conservative			−.241*** [−.34, −.14]	−.250*** [−.35, −.15]	−.254*** [−.36, −.15]			
Counterbalance			.061 [−.21, .33]	.012 [−.25, .28]	−.030 [−.30, .24]	.019 [−.25, .29]	.019 [−.25, .29]	.019 [−.25, .29]
Concordant						.017 [−.09, .12]	−.011 [−.53, .51]	−.117 [−.77, .53]
Discordant						−.428*** [−.54, −.32]	.291 [−.29, .87]	.495 [−.15, 1.14]
Concordant × C							.007 [−.13, .15]	−.054 [−.22, .11]
Discordant × C							−.192* [−.35, −.04]	−.297*** [−.47, −.12]
Concordant × F								−.027 [−.86, .81]
Discordant × F								−.618 [−1.41, .17]
C × F								−.336*** [−.51, −.17]
Concordant × C × F								.205 [−.02, .43]
Discordant × C × F								.284** [.07, .50]
Intercept	.999* [.22, 1.78]	−2.329* [−4.14, −.52]	−.175 [−1.63, 1.28]	.618 [−1.07, 2.31]	−2.142* [−4.28, −.00]	1.339 [−.31, 2.99]	1.126 [−.52, 2.77]	.791 [−.87, 2.45]
N	11,712	11,712	11,712	11,712	11,712	11,712	11,712	11,712
AIC	13,376.6	13,142.2	12,716.1	12,350.7	12,201.2	12,371.7	12,366.7	12,313.7
BIC	13,398.7	13,171.6	12,819.2	12,468.6	12,326.5	12,482.3	12,492.0	12,475.8
Log likelihood	−6,685.3	−6,567.1	−6,344.0	−6,159.4	−6,083.6	−6,170.9	−6,166.4	−6,134.8

Note. AIC = Akaike information criterion; BIC = Bayesian information criterion.
* $p < .05$. ** $p < .01$. *** $p < .001$.

(Figure 2; Table S2 in the online supplemental materials). A moderated mediation analysis using a WLSMV estimator (a robust version of diagonally weighted least squares) and Huber-White robust standard errors revealed conservative political ideology had a significant positive indirect effect via subjective accuracy at low levels of conscientiousness ($b = .018$, $p < .001$, 95% CI [.015, .021]), but a negative indirect effect at high levels of conscientiousness ($b = -.013$, $p < .001$, 95% CI [−.016, −.009]). In short, the moderated mediation

Figure 1
The Predicted Probability of Sharing a News Story for High and Low Conscientiousness Respondents Across Political Ideology in Study 1



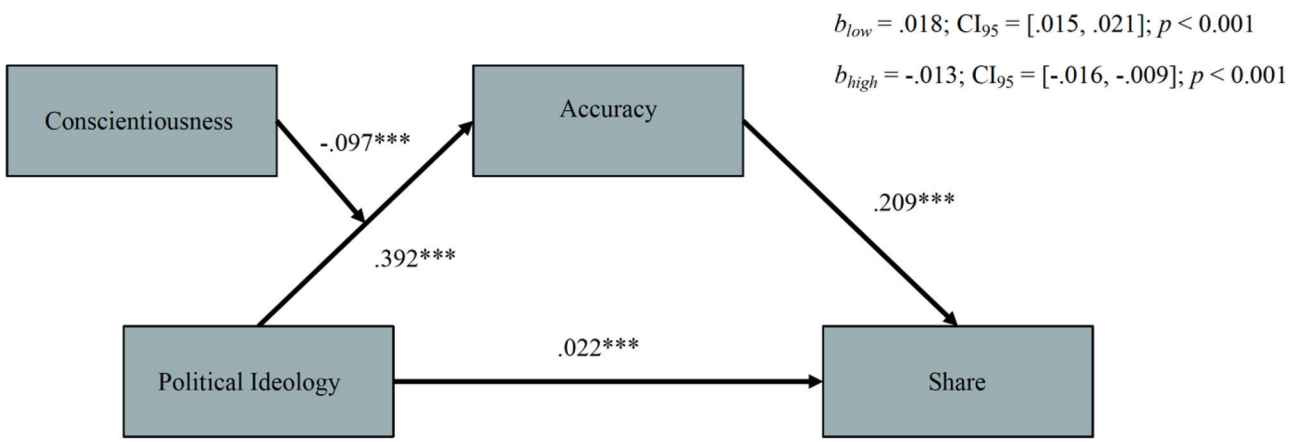
Note. Error bars represent 95% confidence intervals. These are calculated using the error intervals of the coefficients, rather than prediction intervals that are based on the error variance of fitting. This is true for all of the figures. See the online article for the color version of this figure.

model evidenced the role of a motivated process leading to higher perceived accuracy, which in turn increased the likelihood of sharing falsehoods. If inattention was driving these results, we would not have expected political ideology and conscientiousness to interactively predict accuracy perceptions.

News Concordance Versus Discordance

We next conducted preregistered tests on the relationship between the ideological concordance of the news and its likelihood of being shared. We classified news as concordant or discordant news based on the news content (Democrat-leaning, Republican-leaning or neutral) and whether the participant indicated that they preferred the Democratic or Republican party. The dummy variable *concordant* took a value of 1 if respondents indicated that they preferred the Republican party and the news was classified as Republican-leaning (e.g., Breitbart), or if they indicated a preference for Democrats, and the news was Democrat-leaning (e.g., The Raw Story). Similarly, the dummy variable *discordant* was 1 if the news

Figure 2
Moderated Mediation Analysis



Note. We used a WLSMV estimator and Huber-White robust standard errors. *** $p < .001$. See the online article for the color version of this figure.

leaned toward the opposing party. This procedure was similar to prior work studying the effects of the ideological concordance of fake news (Pennycook et al., 2018; Pennycook & Rand, 2019b). However, unlike past work our concordance and discordance categories were defined relative to the neutral category, where the news was not laden with any political content. This enabled us to tease apart the effects of news alignment with political ideology beyond the difference between concordance and discordance.

We tested whether the alignment of a news story with a participant's political beliefs predicted their likelihood of sharing the story. There was no main effect of politically 'concordant' news on the likelihood of sharing ($b = .017$, $p = .76$, model 6), but there was a negative main effect of the news being 'discordant' ($b = -.428$, $p < .001$, model 6). This suggests that differences in participants' sharing of political news was driven by a reduced propensity to share discordant news rather than an enhanced propensity to share concordant news. Previous designs lacking a neutral news condition have precluded this insight. Model 7 examined the interaction between conscientiousness and the news-alignment variables. This was intended to test whether conscientiousness had different effects on the likelihood of sharing a news story contingent on ideological alignment. We found a significant main effect of conscientiousness ($b = -.622$, $p < .001$), no significant main effects of *concordant* or *discordant*, but a significant negative interaction between discordant and conscientiousness ($b = -.192$, $p = .016$; see Figure S3 in the online supplemental materials). Combined with the results of model 6, this suggests that higher conscientiousness participants were driving the negative effect of news discordance on sharing behavior. In other words, highly conscientious people were averse to sharing discordant political news and not biased toward sharing concordant news.

We also estimated all of our regression models using subjective assessments of accuracy as the dependent variable (see Table S3 in the online supplemental materials). Notably, for each of our studies we report this analysis for subjective accuracy in the online supplemental materials. Additionally, for all studies we tested whether conscientiousness and other control variables differed based on political affiliation. We found no differences in conscientiousness based on party affiliation, except for Study S1 and Study 6 where Republicans were slightly higher on conscientiousness (see Table S20 in the online supplemental materials).

Discussion

Overall, Study 1 revealed that the effect of political orientation on the sharing of fake news is more nuanced, such that higher conscientiousness weakened the positive association between conservative ideology and the sharing of fake news: the observed association was largely driven by low conscientiousness conservatives. Furthermore, unlike past work (Pennycook et al., 2018; Pennycook & Rand, 2019b), we found an aversion to discordant news rather than a preference for concordant news as the motivation behind the dissemination of news.

Several scholars have suggested analyzing data from multiple online panels to test the generalizability and robustness of findings (Chandler et al., 2019; Paolacci & Chandler, 2014; Stewart et al., 2017). We therefore performed an additional replication of these results using the online panel Prolific to test our interaction effect in a different population (see Study SI in the online supplemental materials). We replicated the negative interaction between political ideology and conscientiousness on the sharing of fake news in the Prolific sample, both with and without control variables (Table S4 in the online supplemental materials). Additionally, we also ran a study testing our core interaction hypothesis without asking participants to report the subjective accuracy of news stories (Study S2 in the online supplemental materials). Using 20 new false stories, we again replicated the interactive effect of political ideology and conscientiousness on the likelihood of sharing fake news (Table S5 in the online supplemental materials).

Study 2

Study 1 found support for our interaction hypothesis with respect to COVID-19 news stories. However, the use of COVID-19 stories raises concerns regarding the generalizability of our findings with respect to the political news that has been the main focus of investigation when it comes to the sharing of fake news. Thus, in Study 2, we performed a conceptual replication of our core result using political news stimuli from the 2016 Presidential election taken from previous research (Pennycook, Bear, et al., 2020).

Method

Participants

A total of 502 US MTurk participants completed the study in exchange for $1.01. 18 participants were excluded for having a non-U.S. IP address. The final sample consisted of 484 participants (M_{age} = 39.3y, 46.3% females, .6% nonbinary).

Procedure

The procedure for this study was similar to Study 1 except participants rated the accuracy and their tendency to share 12 real and 12 fake political news stories taken from Pennycook and colleagues (Pennycook, Bear, et al., 2020). Within each group of 12 real and fake stories, six were conservative-leaning and six were liberal-leaning. In all there were 12 each of conservative-leaning and liberal-leaning. We randomly presented the 24 news stories and counterbalanced the positions of the stories such that half of the participants responded to personality questionnaire before seeing the news stories, and the other half after seeing the news stories. Following this, participants reported their political orientation, CRT, and other demographic information.

Measures

Our measures were similar to those of Study 1, with exceptions listed below. Personality was measured using the same BFI-2 personality inventory for the five factors; open-mindedness ($\alpha = .89$), conscientiousness ($\alpha = .91$), extraversion ($\alpha = .87$), agreeableness ($\alpha = .85$), and negative emotionality ($\alpha = .91$). We used the same continuous measure of political orientation. For the categorical measure, participants chose whether they voted/would have voted for Donald Trump or Hilary Clinton in the 2016 election (Pennycook & Rand, 2020). The measurement and coding of our dependent variables was identical to Study 1. We included a similar set of control variables as Study 1, with the exclusion of attitudes toward COVID-19.

Results

We followed an identical modeling procedure to Study 1. The results are reported in Table 2. Conscientiousness had a negative main effect ($b = -.864$, $p < .001$, model 1) and political ideology had a positive main effect ($b = .256$, $p < .001$, model 1) on the likelihood of sharing a news story both without (model 1) and with control variables (model 4). More importantly, political ideology and conscientiousness negatively interacted to predict the likelihood of sharing a news story. The interaction effect ($b = -.292$, $p < .001$, model 2) was robust to the inclusion of control variables ($b = -.158$, $p = .005$, model 5). The simple effect of conservative political ideology on the likelihood of sharing a story was significant at low values of conscientiousness ($b = .081$, $p < .001$), but not when conscientiousness was high ($b = -.006$, $p = .33$, Figure 3).

We also found support for the moderated mediation model (Table S7 in the online supplemental materials). There was a positive indirect effect of a more conservative political ideology via subjective accuracy at low levels of conscientiousness ($b = .030$, $p < .001$, 95% CI [.027, .033]), but a negative indirect effect at high levels of conscientiousness ($b = -.010$, $p < .001$, 95% CI [−.013, −.007]).

As in Study 1, we investigated the effect of the political alignment of a news story on its likelihood of being shared. We found a significant positive effect of news being politically concordant on the likelihood of sharing a story ($b = .439$, $p < .001$, model 6), and a significant interaction between conscientiousness and news being

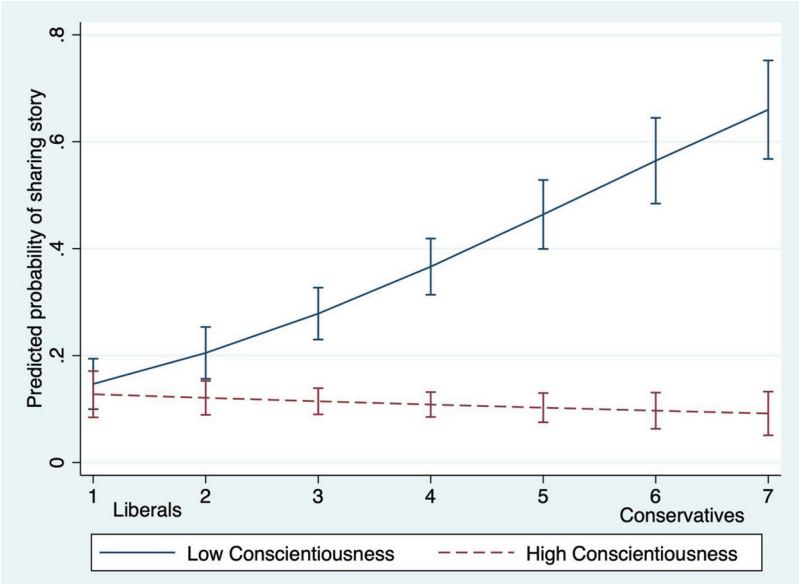
Table 2

Logistic Regression Models Predicting Binary Outcome 'Share' With Clustered Standard Errors, Unstandardized Betas, and 95% Confidence Intervals (Study 2)

Model	1	2	3	4	5	6	7	8
Political ideology (PI)	.256*** [.17, .34]	1.319*** [.93, 1.71]		.133** [.05, .22]	.725*** [.30, 1.15]			
Conscientiousness (C)	−.864*** [−1.07, −.66]	.190 [−.21, .59]		−.777*** [−1.05, −.51]	−.207 [−.73, .31]	−.781*** [−1.05, −.51]	−.924*** [−1.21, −.64]	−.760*** [−1.06, −.46]
PI × C		−.292*** [−.39, −.19]			−.158** [−.27, −.05]			
Fake			−.598*** [−.70, −.50]	−.628*** [−.73, −.53]	−.638*** [−.74, −.53]	−.624*** [−.72, −.52]	−.624*** [−.72, −.52]	.810* [.17, 1.45]
Extraversion			.451*** [.21, .69]	.459*** [.22, .70]	.424*** [.18, .67]	.534*** [.29, .78]	.533*** [.29, .77]	.530*** [.29, .77]
Agreeableness			−.217 [−.47, .03]	−.003 [−.27, .26]	.019 [−.25, .29]	−.054 [−.31, .20]	−.054 [−.31, .20]	−.055 [−.31, .20]
Negative emotionality			.337** [.14, .54]	.046 [−.19, .29]	.079 [−.16, .32]	.010 [−.24, .26]	.011 [−.24, .26]	.013 [−.23, .26]
Openness			−.516*** [−.76, −.28]	−.296* [−.53, −.06]	−.255* [−.50, −.01]	−.393*** [−.62, −.16]	−.392*** [−.62, −.16]	−.392*** [−.62, −.16]
General cognitive reflection			−.234*** [−.30, −.16]	−.220*** [−.29, −.15]	−.202*** [−.27, −.13]	−.238*** [−.31, −.17]	−.238*** [−.31, −.17]	−.239*** [−.31, −.17]
Age			−.014* [−.03, −.00]	−.012 [−.03, .00]	−.010 [−.02, .00]	−.010 [−.02, .00]	−.010 [−.02, .00]	−.010 [−.02, .00]
Male			.351* [.02, .68]	.236 [−.08, .56]	.219 [−.11, .54]	.248 [−.08, .57]	.248 [−.08, .57]	.248 [−.08, .57]
Education			.336*** [.18, .50]	.356*** [.20, .51]	.324*** [.16, .48]	.360*** [.20, .52]	.361*** [.20, .52]	.362*** [.20, .52]
News - Conservative			−.048 [−.15, .05]	−.050 [−.15, .05]	−.051 [−.16, .05]			
Counterbalance			.061 [−.26, .38]	−.014 [−.32, .30]	−.025 [−.34, .29]	.026 [−.29, .34]	.025 [−.29, .34]	.023 [−.29, .34]
Concordant						.439*** [.35, .53]	−.526 [−1.08, .03]	−.526 [−1.16, .10]
Concordant × C							.262*** [.11, .41]	.274** [.11, .44]
Concordant × F								.042 [−.67, .76]
C × F								−.375*** [−.56, −.19]
Concordant × C × F								−.040 [−.25, .17]
Intercept	1.057** [.32, 1.80]	−2.758*** [−4.35, −1.16]	−.444 [−2.14, 1.25]	1.074 [−.91, 3.06]	−1.278 [−4.13, 1.58]	1.676 [−.28, 3.63]	2.200* [.21, 4.19]	1.561 [−.46, 3.58]
N	11,616	11,616	11,616	11,616	11,616	11,616	11,616	11,616
AIC	11,169.6	10,808.7	10,393.1	10,007.2	9,924.7	10,018.7	10,003.9	9,969.6
BIC	11,191.7	10,838.2	10,481.4	10,110.3	10,035.1	10,114.4	10,106.9	10,094.7
Log likelihood	−5,581.8	−5,400.4	−5,184.5	−4,989.6	−4,947.3	−4,996.3	−4,987.9	−4,967.8

Note. AIC = Akaike information criterion; BIC = Bayesian information criterion.
* $p < .05$. ** $p < .01$. *** $p < .001$.

Figure 3
The Predicted Probability of Sharing a News Story for High and Low Conscientiousness Respondents Across Political Ideology in Study 2



Note. Error bars represent 95% confidence intervals. See the online article for the color version of this figure.

politically concordant ($b = .262$, $p < .001$, model 7). These results replicate previous findings that found concordant news to be shared more than discordant news (Bago et al., 2020; Martel et al., 2019). However, this positive relationship between news concordance and its sharing should be interpreted with caution *without a neutral control condition*. The control condition in Study 1 allowed us to test the composition of this effect: that is, whether concordance increased or discordance decreased the sharing of fake news. Without a control condition one would invariably infer that concordance increases sharing of fake news.

Discussion

Study 2 replicated our findings from Study 1 using political stimuli that have been used in prior research. Taken together, Studies 1–2, and S1-2 demonstrate the robustness of our effects across four large samples from two online platforms (MTurk and Prolific) and four sets of news stimuli (two related to COVID-19, one containing both COVID-19 and neutral news, and one set of political stimuli).

Study 3

Having found support for our hypotheses across both COVID-19 and political news, in Study 3 we tested whether our effects on the sharing of misinformation persist *even* in the presence of explicit fact-checker warnings. If low conscientiousness conservatives read and trusted fact-checker warnings, we would expect the effects to disappear. If, however, low conscientiousness conservatives share fake news as an act of negative partisanship against their political outgroup—a motivated process—we would not expect fact checker warnings to have a large impact on their behavior. Accordingly, participants were explicitly warned with a message analogous to those used by Facebook's fact-checking system whether a story was true or false.

Method

We preregistered our study design, sample size, exclusion criteria, hypotheses and analysis (https://osf.io/bv2yg/).

Participants

Four hundred ninety-three U.S. MTurk participants accepted taking part in our study for a payment of $1.26. 14 participants were removed after data collection for having a non-U.S. IP address, resulting in a final sample of 479 participants (M_{age} = 40.4 y, 51.6% women).

Procedure

The study design was similar to Study 1, with the exception that stories were (accurately) tagged with fact-checker warnings. The 12 fake COVID-19 news stories had the warning "Disputed by 3rd Party Fact-Checkers. Learn why this is disputed." The 12 real COVID-19 news stories had the message "Supported by 3rd Party Fact-Checkers. Learn why this is supported." Similar to Study 1, participants indicated their accuracy perceptions of the news, their willingness to share it, self-reported personality, answered CRT items, reported their attitudes toward the COVID-19 virus, and submitted demographic information.

Measures

Our measures were identical to Study 1. Personality was measured using the BFI-2 personality inventory (Soto & John, 2017) and the constructs displayed reasonable reliability—open-mindedness

(α = .88), conscientiousness (α = .89), extraversion (α = .84), agreeableness (α = .85), and negative emotionality (α = .91). Similarly to Study 1, a higher score on attitudes toward COVID-19 indicated a greater perception of COVID-19 severity (α = .91). We also controlled for CRT and demographic information.

Results

Consistent with our theory, we preregistered several hypotheses, including a negative two-way interaction between conscientiousness and political ideology. We additionally predicted a negative effect of a 'false warning' on the likelihood of sharing a story and a negative two-way interaction between conscientiousness and a false warning, such that a false warning will have a greater negative impact on likelihood of sharing when conscientiousness is higher.

As in our other studies, we performed analysis at the news level using logistic regression with standard errors clustered within each individual. The results are presented in Table 3. We found support for our first hypothesis both with (model 4) and without the control variables (model 1). Political ideology positively predicted the sharing of fake news (b = .128, $p < .001$, model 1) whereas the coefficient for conscientiousness was negative (b = −.560, $p <$ .001, model 1).

The interaction effect was significant with (b = −.114, p = .008, model 5) and without (b = −.170, $p < .001$, model 2) the inclusion of control variables. The effect of political ideology on the likelihood of sharing a story was significant at low levels of conscientiousness (b = .056, $p < .001$), but not at high levels (b = −.002, p = .76, Figure 4). We further found a significant negative interaction between conscientiousness and a more conservative political ideology on accuracy (b = −.064, $p < .001$, Table S11 in the online supplemental materials). This is a noteworthy finding as the participants were provided explicit judgments of accuracy, yet the interaction of conscientiousness and political ideology *still* played a key role in their subjective evaluation of news stories. In support of our theory, we again found a positive indirect effect of a more conservative political ideology via subjective accuracy at low levels of conscientiousness (b = .016, $p < .001$, 95% CI [.013, .019]), and a negative indirect effect at high levels of conscientiousness (b = −.013, $p < .001$, 95% CI [−.016, −.009]).

Stories that were associated with a false warning were shared at a lower rate than stories that were indicated as true (b = −.868, $p < .001$, model 3). Moreover, in support of our second interaction hypothesis, we found a significant negative interaction between conscientiousness and a story being explicitly labeled as false (b = −.529, $p < .001$). In other words, the fact-checker warnings were more likely to be effective in curtailing the spread of misinformation for people higher in conscientiousness.

Discussion

In Study 3, we provided participants with explicit warnings regarding the veracity of each story they were asked to evaluate. We found that the interactive effect of political ideology and conscientiousness on the sharing of fake news persisted even in the face of these warnings, consistent with our contention that low conscientiousness conservatives are motivated by furthering the interests of their group, regardless of the news veracity. The significant interaction despite the fact check warnings not only highlights the robustness of this effect, but also points out the inability of such warnings to curtail the spread of misinformation, especially among low conscientiousness conservatives.

Study 4

Study 4 aimed to offer an even stricter test of our hypothesis that conscientiousness weakens the link between political ideology and sharing of misinformation. After participants decided to share a fake story, we explicitly warned them that the news was probably false and asked if they would still like to share the story. If the effect was driven by lazy or inattentive responding, we would expect it to be mitigated when participants had another chance to indicate their intentions to share the falsehoods. But if low conscientiousness conservatives willingly engage in sharing misinformation, they will be unlikely to dynamically update their preferences in response to explicit fact-checker warnings.

Method

We preregistered our study design, sample size, exclusion criteria, hypotheses, and analysis (https://osf.io/ku6fq/).

Participants

Of the 997 US MTurk participants that completed our study, 30 were excluded for not having a U.S. IP address, resulting in a final sample consisting of 967 participants (M_{age} = 39.2y, 51.8% female). Participants were paid $1.51.

Procedure

The study design was similar to Study 1, with the difference that subjects were offered a chance to revise their decision to share a fake news story by informing them that the news was likely false. Thus, the first part of the study was identical to Study 1 and the second part included an intervention only for the fake news stories that participants were willing to share. We hoped that offering targeted interventions for fabricated news might be a more effective strategy to dissuade people from sharing falsehoods. As per our other studies, participants indicated the stories' subjective accuracy and their likelihood of sharing the news for each of the 24 news stories on COVID-19. These stimuli were identical to Studies 1 and 3 and comprised of neutral, liberal, and conservative-leaning stories. After answering questions about the stories, participants responded to the BFI-2 personality inventory and CRT measures. At this point, we introduced our intervention only for fake news stories that participants indicated they were willing to share.

Participants saw the following message, "Prior to making your ultimate decision regarding your views of the accuracy and your likelihood of sharing these stories, you have the chance to see the results of an independent fact checker. If none of the stories you indicated you may share were disputed by third party fact checkers, the survey will automatically move on to asking questions about yourself." Participants then learned whether the news stories that they were willing to share were disputed by 3rd party fact-checkers and asked if they would still like to share them. Finally, participants reported their political ideology, their attitudes toward COVID-19, and demographic information. Given this study

Table 3
Logistic Regression Models Predicting Binary Outcome 'Share' With Clustered Standard Errors, Unstandardized Betas, and 95% Confidence Intervals (Study 3)

Model	1	2	3	4	5	6	7	8
Political ideology (PI)	.128*** [.07, .19]	.777*** [.43, 1.13]		.083* [.02, .15]	.525** [.18, .87]			
Conscientiousness (C)	−.560*** [−.73, −.39]	.093 [−.26, .45]		−.369*** [−.56, −.18]	.040 [−.31, .39]	−.366*** [−.55, −.18]	−.333*** [−.52, −.15]	−.153 [−.36, .05]
PI × C		−.170*** [−.26, −.08]			−.114** [−.20, −.03]			
False warning (FW)			−.868*** [−.98, −.76]	−.881*** [−.99, −.77]	−.887*** [−1.00, −.78]	−.883*** [−.99, −.77]	−.884*** [−.99, −.78]	.624 [−.24, 1.48]
Extraversion			.336*** [.15, .53]	.322*** [.13, .51]	.325*** [.14, .51]	.349*** [.16, .54]	.349*** [.16, .54]	.352*** [.16, .54]
Agreeableness			−.425*** [−.64, −.21]	−.338** [−.55, −.12]	−.338** [−.55, −.13]	−.367*** [−.58, −.15]	−.367*** [−.58, −.15]	−.368*** [−.59, −.15]
Negative emotionality			.108 [−.06, .27]	.016 [−.17, .20]	−.002 [−.18, .18]	−.007 [−.19, .17]	−.007 [−.19, .17]	−.007 [−.19, .17]
Openness			−.277** [−.46, −.09]	−.162 [−.35, .03]	−.131 [−.33, .06]	−.218* [−.41, −.03]	−.217* [−.41, −.03]	−.215* [−.41, −.02]
Attitudes towards COVID-19			.227** [.08, .37]	.278*** [.13, .43]	.266*** [.11, .42]	.250** [.10, .40]	.250** [.10, .40]	.252** [.10, .40]
General cognitive reflection			−.140*** [−.20, −.08]	−.121*** [−.18, −.06]	−.110*** [−.17, −.05]	−.134*** [−.19, −.08]	−.134*** [−.19, −.08]	−.134*** [−.19, −.08]
Age			−.004 [−.01, .01]	−.003 [−.01, .01]	−.003 [−.01, .01]	−.001 [−.01, .01]	−.001 [−.01, .01]	−.002 [−.01, .01]
Male			.517*** [.26, .78]	.464*** [.21, .72]	.433** [.17, .69]	.458*** [.20, .72]	.458*** [.20, .72]	.461*** [.20, .72]
Education			.243*** [.13, .36]	.217*** [.10, .33]	.212*** [.10, .33]	.224*** [.11, .34]	.224*** [.11, .34]	.223*** [.11, .34]
News - Liberal			−.129* [−.24, −.02]	−.131* [−.25, −.02]	−.132* [−.25, −.02]			
News - Conservative			−.146** [−.24, −.05]	−.148** [−.25, −.05]	−.149** [−.25, −.05]			
Counterbalance			−.081 [−.33, .17]	−.060 [−.31, .19]	−.051 [−.30, .20]	−.058 [−.31, .19]	−.059 [−.31, .19]	−.061 [−.31, .19]
Concordant						.098 [−.01, .20]	−.268 [−.81, .28]	−.281 [−.94, .37]
Discordant						−.393*** [−.49, −.29]	.436 [−.09, .96]	.228 [−.43, .88]
Concordant × C							.095 [−.05, .24]	.009 [−.16, .18]
Discordant × C							−.220** [−.36, −.08]	−.232** [−.40, −.06]
Concordant × FW								−.496 [−1.39, .40]
Discordant × FW								.068 [−.76, .90]
C × FW								−.529*** [−.76, −.30]
Concordant × C × FW								.348** [.11, .59]
Discordant × C × FW								.146 [−.08, .38]
Intercept	.965** [.27, 1.66]	−1.531* [−2.99, −.07]	−.129 [−1.61, 1.35]	.316 [−1.38, 2.01]	−1.315 [−3.42, .79]	.995 [−.59, 2.58]	.868 [−.73, 2.46]	.352 [−1.27, 1.97]
N	11,496	11,496	11,496	11,496	11,496	11,496	11,496	11,496
AIC	14,037.3	13,884.6	12,971.0	12,830.9	12,773.5	12,790.3	12,774.2	12,677.6
BIC	14,059.3	13,914.0	13,073.9	12,948.5	12,898.5	12,900.5	12,899.1	12,839.3
Log likelihood	−7,015.6	−6,938.3	−6,471.5	−6,399.5	−6,369.8	−6,380.1	−6,370.1	−6,316.8

Note. AIC = Akaike information criterion; BIC = Bayesian information criterion.
* $p < .05$. ** $p < .01$. *** $p < .001$.

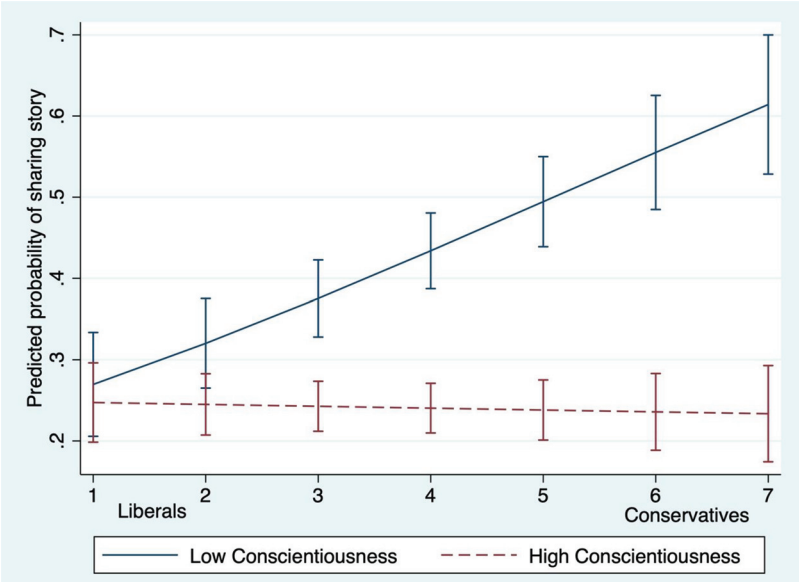
design, we expected to replicate our interaction effect at story level before the intervention. Following the intervention, we examined the average change in participants' willingness to continue sharing false stories despite the warning. This latter analysis was performed at the participant level.

Measures

Our independent measures were identical to Studies 1 and 3. Personality was measured using the BFI-2 personality inventory's five factors (Soto & John, 2017)—open-mindedness (α = .87),

Figure 4

The Predicted Probability of Sharing a News Story for High and Low Conscientiousness Respondents Across Political Ideology in Study 3



Note. Error bars represent 95% confidence intervals. See the online article for the color version of this figure.

conscientiousness (α = .88), extraversion (α = .85), agreeableness (α = .86), and negative emotionality (α = .90). As per our preregistration, we created a variable *change*, which indicated the extent to which participants changed their mind to share fake stories after seeing an independent third-party fact-checker warning. For example, consider a situation where a participant had selected *Maybe* or *Yes* to the sharing question for three fake stories initially, and after seeing the warning participant chose to share 1 of those stories again. The variable *change* in this case would take the value (3 − 1)/3 = .67. Thus, this variable ranged from 0 to 1, with 0 representing no change in preference and 1 representing total change in preference to not share fake news. Hence, a higher value on this variable represented a *greater reluctance* to share fake news. Overall, this variable captured aggregate reduction in the sharing of fake news at a participant level.

Similar to Study 1, we created a composite measure reflecting COVID-19 attitudes (α = .93), and also controlled for CRT and other demographic information.

Results

Consistent with our other studies, we replicated all of our findings before the fact-checker intervention. There was a negative interaction between conscientiousness and political ideology (see Figure 5). This was robust to the inclusion of controls ($b = -.162$, $p < .001$, Table S12 in the online supplemental materials) and the inclusion of all other personality interactions ($b = -.117$, $p = .007$, Table S13 in the online supplemental materials).

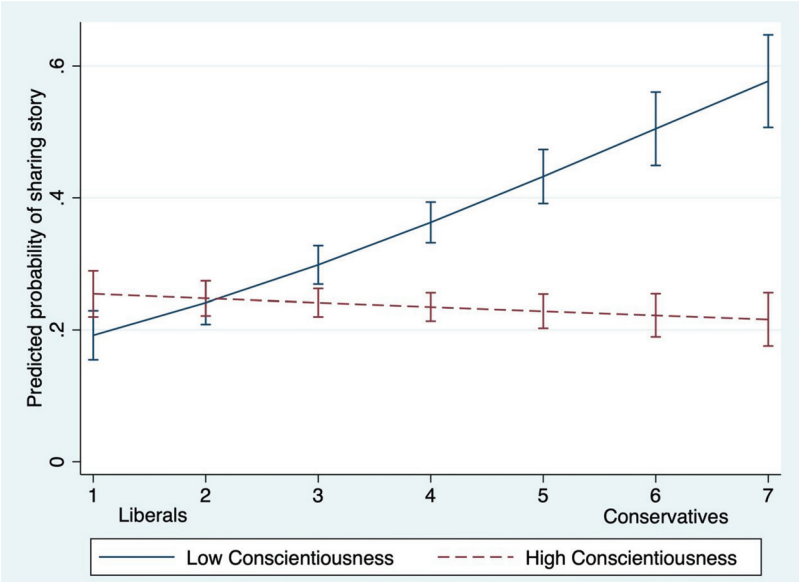
Next, we present our analysis examining whether fact checker warnings revised individuals' tendencies to share fake news. Of the 651 participants who initially indicated their willingness to share one or more fake story, their average willingness to share false information reduced from 4.58 to 2.78 stories ($p < .001$) after seeing the fact-checker warning. This is encouraging as it shows that fact checkers were somewhat effective in reducing the sharing of misinformation. We next examined whether conscientiousness and political ideology predicted this change.

Conscientiousness had a positive main effect on participants' tendency to update their choices ($b = .06$, $p = .003$, model 1, Table 4), whereas political ideology was negatively associated with changing ($b = -.059$, $p < .001$, model 1). We also observed a significant positive interaction between political ideology and conscientiousness ($b = .034$, $p = .001$, model 2). This interaction was robust to the inclusion of control variables ($b = .024$, $p = .024$, model 5). The simple effect of political ideology on 'change' was significant at one standard deviation below mean conscientiousness ($b = -.084$, $p < .001$), and also at one standard deviation above ($b = -.033$, $p = .003$). However, the two slopes were significantly different ($b = -.051$, $p = .001$), suggesting that low conscientiousness conservatives changed their news sharing behavior significantly less than high conscientiousness conservatives (see Figure 6). Overall, our hypothesis was supported: Political ideology and conscientiousness interactively predicted participants' responsiveness to fact-checker warnings.

Discussion

Study 4 revealed that conscientiousness and political ideology jointly predicted participants' tendency to update their fake news sharing behavior in response to fact checker warnings. While political conservatives were less likely to change their behavior in general, this was even more pronounced for low conscientiousness conservatives. These results are consistent with our assertion that

Figure 5
The Predicted Probability of Sharing a News Story for High and Low Conscientiousness Respondents Across Political Ideology in Study 4



Note. Error bars represent 95% confidence intervals. See the online article for the color version of this figure.

conservatives share fake news to further the interests of their group, and this tendency is enhanced for conservatives who are low in conscientiousness. Importantly, we also replicated the results from the other studies preintervention.

Study 5

Having established the robustness of this effect to fact checker interventions, we next examined whether low conscientiousness conservatives exhibited a greater need for chaos. Accordingly, we tested the interactive effect of political ideology and conscientiousness on participants' desire for chaos. At the same time, this study also ruled out other mechanisms that could be pertinent, such as support for social and economically conservative issues, support for Donald Trump, trust in the media, and whether the time participants spent on social media accounted for any meaningful variance.

Method

The preregistration containing information on study design, sample size, exclusion criteria, hypotheses and analyses can be found at https://osf.io/6rsx5/.

Participants

On MTurk, we recruited 300 U.S. participants in exchange for $1.01. We excluded 15 participants for not having a U.S. IP address, resulting in a final sample of 285 participants (M_{age} = 40.1 y, 51.2% female).

Procedure

Participants first responded to the personality questionnaire and indicated their political ideology, before responding to the other scales. We randomly administered five different measures—desire for chaos (Petersen et al., 2020), support for social and economically conservative issues (Everett, 2013), support for Donald Trump, trust in the mainstream media, and time spent on social media. Following this, participants reported their demographics.

Measures

Consistent with the other studies, political ideology was measured using a one item continuous measure and personality via the BFI-2 inventory. The reliability coefficients for the Big Five factors were acceptable; open-mindedness (α = .88), conscientiousness (α = .89), extraversion (α = .83), agreeableness (α = .84), and negative emotionality (α = .90).

We measured desire for chaos using an eight-item scale (Petersen et al., 2020) on a five-point scale ranging from *Strongly Disagree* to *Strongly Agree*. This scale allowed us to measure people's desire for anarchy (α = .95). Another possible mechanism we wanted to rule out was whether low conscientiousness conservatives reported greater support for traditional social and economic conservative issues. To that end, we measured participants' alignment with both social and economic conservatism via the 12-item SECS scale (Everett, 2013). This scale captures participants opinion on a number of social (e.g., abortion; α = .85) and economic issues (e.g., welfare benefits; α = .68). We also measured support for Donald Trump as another alternate mechanism using a four-item scale. An example item was "President Trump has been very effective in his presidency" (α = .97).

It is important to note that participants' support for social and economically conservative issues may have some overlap with political ideology, but they are not identical constructs. Support for social and economic conservatism reflects beliefs associated with specific policies, rather than one's identification with a position on

Table 4
The Change of Intent to Share Fake News Stories Using Linear Regression With 95% Confidence Intervals (Study 4)

Model	1	2	3	4	5
Political ideology (PI)	−.059***	−.188***		−.041***	−.132**
	[−.07, −.04]	[−.27, −.11]		[−.06, −.02]	[−.21, −.05]
Conscientiousness (C)	.060**	−.068		.031	−.056
	[.02, .10]	[−.15, .02]		[−.02, .09]	[−.15, .04]
PI × C		.034**			.024*
		[.01, .05]			[.00, .04]
Extraversion			−.005	−.005	−.000
			[−.05, .04]	[−.05, .04]	[−.04, .04]
Agreeableness			.059*	.043	.044
			[.01, .11]	[−.01, .09]	[−.01, .09]
Negative emotionality			.045*	.045	.045
			[.00, .09]	[−.00, .09]	[−.00, .09]
Openness			.102***	.066**	.059*
			[.06, .15]	[.02, .11]	[.01, .11]
Attitudes towards COVID-19			.005	−.007	−.005
			[−.03, .04]	[−.04, .03]	[−.04, .03]
General cognitive reflection			.036***	.029***	.027***
			[.02, .05]	[.01, .04]	[.01, .04]
Age			−.001	−.000	−.000
			[−.00, .00]	[−.00, .00]	[−.00, .00]
Male			−.085**	−.081*	−.080*
			[−.15, −.02]	[−.14, −.02]	[−.14, −.02]
Intercept	.550***	1.038***	−.219	.050	.397
	[.38, .72]	[.70, 1.37]	[−.56, .12]	[−.34, .44]	[−.10, .89]
N	651	651	651	651	651
AIC	614.3	605.5	606.6	586.1	582.9
BIC	627.8	623.4	646.9	635.4	636.6
Log likelihood	−304.2	−298.8	−294.3	−282.1	−279.5

Note. AIC = Akaike information criterion; BIC = Bayesian information criterion.
* $p < .05$. ** $p < .01$. *** $p < .001$.

the left-right continuum. While in our sample support for these policies and institutions (e.g., Abortion, Welfare benefits) correlated highly with left-right identification, this does not have to be the case. These correlations likely reflect the polarized climate in the United States, where people uncritically align with the views of their party/political group. Regardless, we primarily wanted to test whether political ideology and conscientiousness interactively predicted need for chaos, rather than specific policy issues that might be important to the majority of conservatives including those high in conscientiousness.

Finally, we measured participants' trust in the mainstream media. Prior research has established that there is declining trust in mainstream media (Allcott & Gentzkow, 2017; van der Linden et al., 2020), and this has been touted as a possible reason for the advent of the present fake news pandemic (Allen et al., 2020). Four items assessed participants' trust in the mainstream media. Sample item: "What is reported in the news is often not true" ($\alpha = .83$). We also controlled for time spent on social and digital media in our analysis. Participants indicated the number of hours they spent on social media in a day and their reliance on news sources—adding up to 100%—from the following choices; social media, digital media, print press, friends and family, and other.

Results

We interacted political ideology and conscientiousness to predict need for chaos, social conservatism, economic conservatism, support for Donald Trump, and trust in the mainstream media while controlling for participants' media diet composition, other personality factors and demographic information. The results are reported in Table 5. We found a statistically significant negative interaction of political ideology and conscientiousness on desire for chaos ($b = -.136$, $p < .001$, model 1) indicating that low conscientiousness conservatives reported an elevated desire for chaos (see Figure 7). The interaction of political ideology and conscientiousness was also statistically significant for trust in the mainstream media ($b = -.088$, $p = .017$), suggesting contrary to our expectations that low conscientiousness conservatives reported greater trust in the mainstream media. Notably, the political ideology-conscientiousness interaction did not predict social conservatism ($b = .072$, $p = .30$), economic conservatism ($b = .003$, $p = .96$), or support for Trump ($b = .046$, $p = .23$).

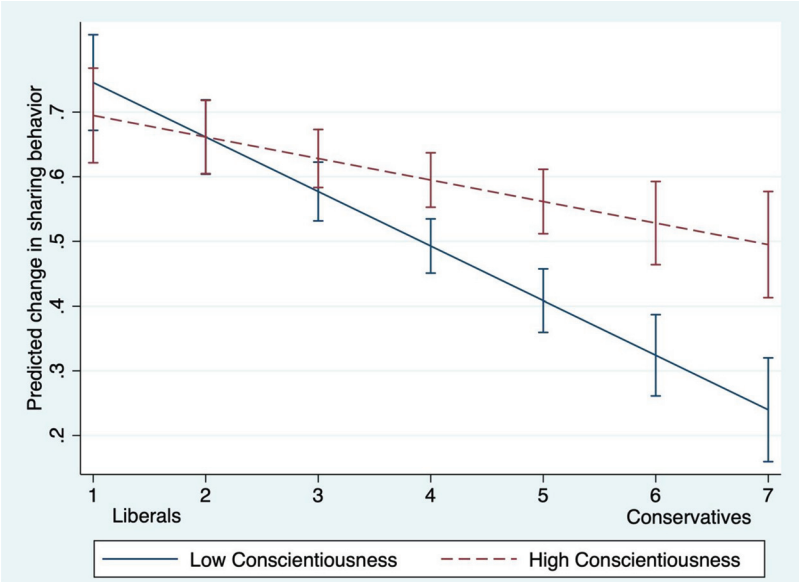
Discussion

Study 5 revealed that political ideology and conscientiousness interactively predicted desire for chaos. Specifically, low conscientiousness conservatives reported a greater desire for chaos compared with high conscientiousness conservatives and liberals. At the same time, this study also ruled out other alternate explanations.

Study 6

Study 6 tested the full model to ascertain whether desire for chaos mediated the interactive effect of political ideology and conscientiousness on the sharing of fake news.

Figure 6
The Predicted Change in Sharing Behavior for High and Low Conscientiousness Respondents Across Political Ideology in Study 4



Note. Error bars represent 95% confidence intervals. See the online article for the color version of this figure.

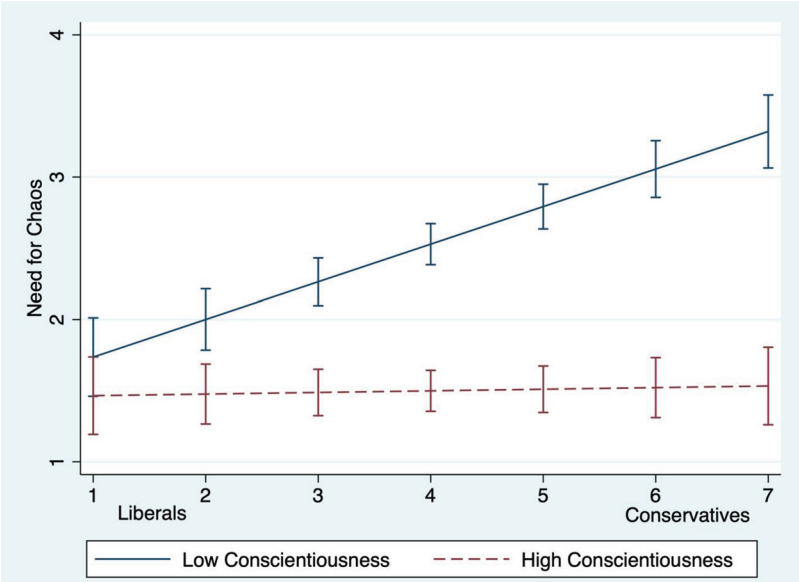
Method

The preregistered information on study design, sample size, exclusion criteria, hypotheses, and analyses is available at https://osf.io/98wx2/.

Participants

Four hundred ninety-eight U.S. MTurk workers signed up for a payment of $1.01. We excluded the responses of 7 participants for a non-U.S. IP address. Our final sample consisted of 491

Figure 7
The Predicted Desire for Chaos for High and Low Conscientiousness Respondents Across Political Ideology in Study 5



Note. Error bars represent 95% confidence intervals. See the online article for the color version of this figure.

Table 5
The Relationship Between the Interaction Between Political Ideology and Conscientiousness and Various Dependent Variables Using Linear Regression With 95% Confidence Intervals (Study 5)

Model	1	2	3	4	5
Dependent variable	Desire for chaos	Social conservatism	Economic conservatism	Support for Trump	Trust in the media
Political ideology (PI)	.581***	.544*	.597**	.409**	.126
	[.34, .82]	[.03, 1.06]	[.19, 1.00]	[.12, .70]	[−.15, .40]
Conscientiousness (C)	.318*	−.476	.223	−.196	.269
	[.02, .61]	[−1.10, .15]	[−.27, .71]	[−.54, .15]	[−.06, .60]
PI × C	−.136***	.072	.003	.046	−.088*
	[−.20, −.07]	[−.06, .21]	[−.10, .11]	[−.03, .12]	[−.16, −.02]
Extraversion	.085	.352*	.149	.129	−.022
	[−.06, .23]	[.04, .66]	[−.10, .40]	[−.05, .30]	[−.19, .15]
Agreeableness	−.556***	.680***	.039	−.190	.263**
	[−.73, −.38]	[.31, 1.05]	[−.25, .33]	[−.40, .01]	[.07, .46]
Negative emotionality	.003	−.082	−.046	−.009	.009
	[−.14, .15]	[−.39, .22]	[−.29, .19]	[−.18, .16]	[−.15, .17]
Openness	−.208**	−.221	−.026	−.112	−.181*
	[−.34, −.07]	[−.51, .06]	[−.25, .20]	[−.27, .05]	[−.33, −.03]
Age	−.004	.012	.010	.002	−.001
	[−.01, .00]	[−.00, .03]	[−.00, .02]	[−.01, .01]	[−.01, .01]
Male	.280**	−.487*	−.215	−.310**	.088
	[.08, .48]	[−.90, −.07]	[−.54, .11]	[−.54, −.08]	[−.13, .31]
Social media time	.088***	−.041	−.101**	.035	.028
	[.05, .13]	[−.13, .05]	[−.17, −.03]	[−.01, .08]	[−.02, .08]
Social media proportion of news	−.001	.000	.003	.001	−.002
	[−.01, .00]	[−.01, .01]	[−.01, .01]	[−.00, .01]	[−.01, .00]
Intercept	2.921***	2.907	3.283**	1.698*	2.348**
	[1.50, 4.34]	[−.09, 5.91]	[.91, 5.65]	[.03, 3.37]	[.75, 3.95]
N	285	285	285	285	285
AIC	656.8	1,084.4	949.9	750.9	726.4
BIC	700.6	1,128.2	993.8	794.8	770.2
Log likelihood	−316.4	−530.2	−463.0	−363.5	−351.2

Note. AIC = Akaike information criterion; BIC = Bayesian information criterion.
* $p < .05$. ** $p < .01$. *** $p < .001$.

participants (M_{age} = 40.4 y, 54.4% female). Participants were paid $1.01.

Procedure

The procedure was identical to that of Study 1, with the exception that we measured three additional variables; desire for chaos, trust in the mainstream media, and participants' media use. Participants faced either the 24 news stories or BFI-2 personality inventory in a counterbalanced order, before responding to the desire for chaos, trust in the mainstream media, and social media use measures in a randomized order. Participants then reported their political ideology, attitudes toward COVID-19, and demographic information.

Measures

We measured our independent variables in line with Studies 1–5. All reliability coefficients for the Big Five personality dimensions fell in the acceptable range; open-mindedness (α = .88), conscientiousness (α = .90), extraversion (α = .85), agreeableness (α = .85), and negative emotionality (α = .90). We measured our mediating variable desire for chaos using the same 8-item scale as in Study 5 (α = .91) (Petersen et al., 2020). We included participants' concern with the virus COVID-19 (α = .93) and trust in the mainstream media as control variables (α = .86).

Results

We replicated all of our findings for the sharing of fake news (see Table 6, Table S16 in the online supplemental materials). The interaction of political ideology and conscientiousness on the sharing of fake news was statistically significant without (b = −.116, p = .023, model 2) and with the inclusion of control variables (b = −.107,
p = .04, model 5). The interaction pattern revealed that the simple slope was significant at both low (b = .07, $p < .001$) and high values of conscientiousness (b = .02, p = .002, Figure 8). A possible reason for the significant simple slope observed at high values of conscientiousness was Republicans' higher average levels of conscientiousness in Study 6 (Table S20 in the online supplemental materials). However, the difference in the two slopes was negative and statistically significant (b = −.05, p = .001), implying that low conscientiousness conservatives were even more likely to share falsities.

Additionally, we replicated our findings from Study 5 with a statistically significant interaction of political ideology and conscientiousness in predicting desire for chaos both without (b = −.123, $p < .001$, Table S17 in the online supplemental materials) and with control variables for the other four personality dimensions, trust in the mainstream media, attitudes toward COVID-19,

Table 6
Logistic Regression Models Predicting Binary Outcome 'Share' With Clustered Standard Errors, Unstandardized Betas, and 95% Confidence Intervals (Study 6)

Model	1	2	3	4	5	6	7
Political ideology (PI)	.249***	.683***		.253***	.658**	.263***	.400*
	[.18, .32]	[.29, 1.08]		[.17, .33]	[.25, 1.06]	[.18, .35]	[.00, .80]
Conscientiousness (C)	−.574***	−.131		−.504***	−.103	−.236*	−.109
	[−.76, −.39]	[−.55, .29]		[−.73, −.27]	[−.55, .34]	[−.47, −.00]	[−.55, .33]
PI × C		−.116*			−.107*		−.036
		[−.22, −.02]			[−.21, −.00]		[−.14, .07]
Fake			−.235***	−.246***	−.248***	−.263***	−.263***
			[−.29, −.18]	[−.31, −.19]	[−.31, −.19]	[−.33, −.20]	[−.33, −.20]
Extraversion			.310**	.317**	.318**	.245*	.246*
			[.12, .50]	[.11, .52]	[.11, .52]	[.04, .45]	[.04, .45]
Agreeableness			−.200	−.026	.006	.218	.225
			[−.45, .05]	[−.28, .22]	[−.25, .26]	[−.04, .48]	[−.04, .49]
Negative emotionality			.255**	.135	.155	.216*	.220*
			[.06, .45]	[−.06, .33]	[−.04, .35]	[.02, .42]	[.02, .42]
Openness			−.399***	−.178	−.161	−.011	−.007
			[−.60, −.20]	[−.38, .02]	[−.37, .05]	[−.22, .20]	[−.22, .21]
Attitudes towards COVID-19			.055	.181*	.176*	.071	.074
			[−.10, .21]	[.03, .34]	[.02, .33]	[−.09, .23]	[−.09, .24]
Age			−.002	−.004	−.004	.001	.001
			[−.01, .01]	[−.02, .01]	[−.02, .01]	[−.01, .01]	[−.01, .01]
Male			.453**	.340*	.346*	.282*	.285*
			[.16, .75]	[.05, .63]	[.06, .63]	[.01, .56]	[.01, .56]
Education			.018	.029	.017	−.076	−.078
			[−.11, .14]	[−.10, .16]	[−.11, .14]	[−.20, .05]	[−.20, .05]
News - Liberal			−.046	−.048	−.048	−.051	−.051
			[−.16, .06]	[−.16, .07]	[−.16, .07]	[−.17, .07]	[−.17, .07]
News - Conservative			−.161***	−.169***	−.170***	−.180***	−.180***
			[−.25, −.07]	[−.26, −.07]	[−.27, −.07]	[−.28, −.08]	[−.28, −.08]
Counterbalance			−.080	−.155	−.165	−.070	−.076
			[−.35, .19]	[−.42, .11]	[−.43, .10]	[−.33, .19]	[−.34, .18]
Desire for chaos						.725***	.716***
						[.55, .90]	[.54, .89]
Trust in media						.235**	.226**
						[.08, .39]	[.07, .39]
Intercept	.255	−1.386	−.506	−1.104	−2.780*	−5.108***	−5.594***
	[−.45, .96]	[−3.02, .24]	[−2.06, 1.05]	[−2.78, .57]	[−5.15, −.40]	[−6.89, −3.33]	[−7.94, −3.24]
N	11,784	11,784	11,784	11,784	11,784	11,784	11,784
AIC	13,135.9	13,079.6	13,341.5	12,866.8	12,822.1	12,260.2	12,257.4
BIC	13,158.0	13,109.1	13,437.4	12,977.4	12,940.1	12,385.6	12,390.2
Log likelihood	−6,564.9	−6,535.8	−6,657.7	−6,418.4	−6,395.1	−6,113.1	−6,110.7

Note. AIC = Akaike information criterion; BIC = Bayesian information criterion.
* $p < .05$. ** $p < .01$. *** $p < .001$.

age, gender, time spent on social media, and the proportion of news received from social media ($b = -.083$, $p < .001$, Table S17 in the online supplemental materials).

Finally, we tested a preregistered moderated mediation model (Figure 9, Table S18 in the online supplemental materials), where political ideology was the independent variable, conscientiousness was the moderator, desire for chaos was the mediator, and the likelihood of sharing a story was the dependent variable. We found that there was a positive indirect effect of political ideology on the likelihood of sharing a news story via the mediator of desire for chaos at low levels of conscientiousness ($b = .034$, $p < .001$, 95% CI [.031, .037]), but that the indirect effect was not significant for high levels of conscientiousness ($b = -.001$, $p = .279$, 95% CI [−.002, .0001]). Note also that in model 8 (see Table 6) with the inclusion of desire for chaos, the interaction between political ideology and conscientiousness was no longer statistically significant ($b = -.036$, $p = .49$). This gives further weight to our contention that the interactive effect of political ideology and conscientiousness on the likelihood of sharing stories was mediated by a desire for chaos.

Discussion

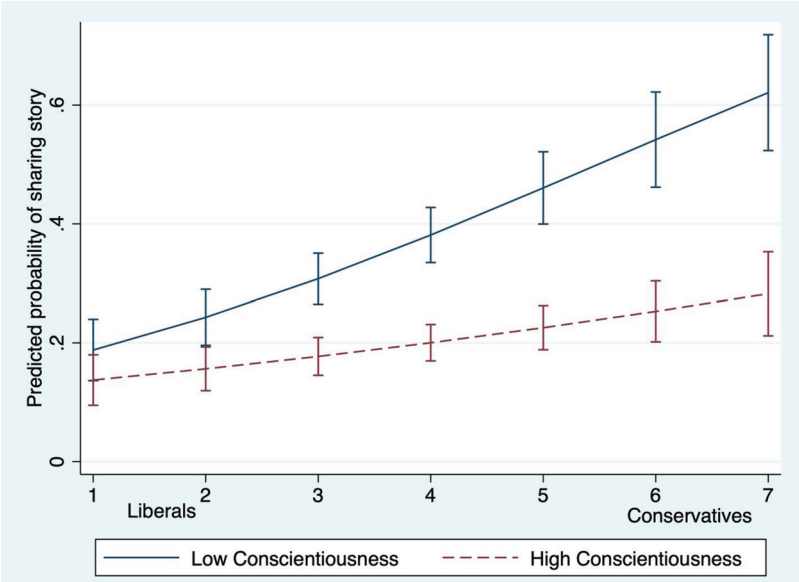
Study 6 further replicated the interactive effect of political ideology and conscientiousness on people's desire for chaos. Furthermore, low conscientiousness conservatives' need for chaos emerged as the underlying mechanism that mediated the interactive effect of political ideology and conscientiousness on the sharing of fake news. All of these analyses were robust to controlling for trust in the media.

General Discussion

Our goal was to demonstrate that trait conscientiousness weakens the positive effect of conservative political ideology on the spread of misinformation. At the same time, we focused on news

Figure 8

The Predicted Probability of Sharing a News Story for High and Low Conscientiousness Respondents Across Political Ideology in Study 6



Note. Error bars represent 95% confidence intervals. See the online article for the color version of this figure.

stories related to COVID-19, as it provided a context with the potential to harm anyone, irrespective of their political beliefs. This allowed us to shed light on why even in a pandemic, misinformation may spread like wildfire, impairing our ability to combat the virus.

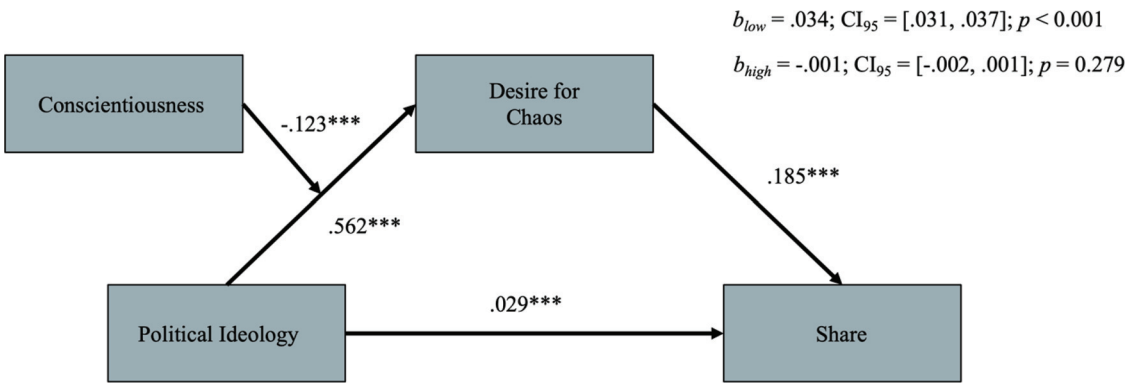
Across eight studies with a sample of 4,642 participants and 91,144 news level observations, we found that political ideology and conscientiousness interactively predicted the likelihood of sharing fake news, such that differences in sharing behavior across party lines were driven by conservatives who were low in conscientiousness. We found no differences among liberals and conservatives at high levels of conscientiousness. This effect was consistent both for news articles regarding COVID-19 (Studies 1, S1, S2, 3, 4, and 6), the 2016 Presidential Election (Study 2), and neutral topics (Study S2), and across samples from two different online populations, MTurk (Studies 1–4 and 6), and Prolific (Study S1). The interaction was robust to alternative explanations based on inattention, accuracy primes, and also to the inclusion of several control variables, including cognitive reflection and the interaction of political ideology with other personality factors. This underscores the reliability of our effect in predicting the dissemination of misinformation.

Theoretical Contributions

Our findings advance the literature by moving beyond studying the main effect of conservative political ideology on the sharing of fake news, and thus providing a more nuanced and balanced opinion. In this way, our theory and findings avoid denigrating all conservatives and offer empirical evidence that aligns with the general observation of greater heterogeneity among conservatives. Notably, the size of our interaction effect across our studies is also of practical relevance. In Table S22 in the online supplemental materials, we used linear regression to predict the total number of stories participants' shared, and report regression predictions for "Slightly Liberal" and "Slightly Conservative" participants at low and high levels of conscientiousness. At high levels of conscientiousness, the predicted proportion of stories shared varied little between these political groups (.004–.047 across Studies 1–4 and 6), but at low levels of conscientiousness the gap widened vastly (.113–.166). This gap was even more pronounced for those who were at the more extreme ends of political ideology. In summary, disparities in sharing behavior across the political aisle were nearly exclusively driven by low conscientiousness respondents. To curtail the spread of misinformation, policymakers should focus on low conscientiousness conservatives.

Furthermore, we find that our effects were driven by an indiscriminate desire to cause chaos, as opposed to alternative explanations based on the promotion of social or economic conservative policies, supporting Donald Trump, distrust in the mainstream media, or time spent on social media (Studies 5 and 6). This explanation underscores the negative effects of political polarization, which leads a subset of individuals (low conscientiousness conservatives) to promote the interests of their political group by actively subverting other groups through the sharing of falsehoods—a form of radical political activism (Moskalenko & McCauley, 2009). Our data suggest that higher trait level conscientiousness may be a key factor in mitigating the likelihood of political conservatives engaging in such negative partisanship behaviors. Additionally, we did not find that respondents have a preference for sharing ideologically concordant news, but rather they have an *aversion* to sharing discordant news (Studies 1 and 2). Comparing

Figure 9
Serial Moderated Mediation Analysis (Study 6)



Note. We used a WLSMV estimator and Huber-White robust standard errors. *** $p < .001$. See the online article for the color version of this figure.

participants' behavior concerning our politically neutral COVID-19 news stories to overtly political news allowed us to tease apart this effect.

More broadly, our findings demonstrate the importance of studying personality in relation to understanding fake news as a phenomenon. Our focus on a single personality trait responds to calls in the personality literature to identify the impact of specific traits rather than investigating all Big 5 personality attributes simultaneously (Moskalenko & McCauley, 2009; Paunonen & Ashton, 2001), but future research can benefit from studying the relationship between other facets of personality and the sharing of fake news. We observed a positive main effect of extraversion on the likelihood of sharing stories, and some evidence of a negative effect of open-mindedness.

A more worrying aspect of our findings is the inadequacy of fact checker interventions. Although in Study 4 we found some reduction in the sharing of fake news after participants learned the news that they were willing to share was false, it was still not sufficient to stop low conscientiousness conservatives from widely spreading falsehoods. Another relevant comparison comes from comparing the results of Study 1 and Study 3, which used identical news stories, but Study 3 explicitly attached either fake or real tags to stories (Table S24 in the online supplemental materials). We found that attaching a real tag to a story was more effective in increasing participants' likelihood of sharing a story than fake warnings were in deterring sharing behavior. This again suggests that individuals who are motivated to defend their ingroup at the expense of the outgroup will perceive higher subjective accuracy of news, regardless of a story's true accuracy. Notably, the interaction effect between political ideology and conscientiousness was robust to removing ambiguity regarding the veracity of news in both Studies 3 and 4.

Overall, these findings troublingly suggests that low conscientiousness conservatives were deliberately choosing to propagate false stories as a means to further the superiority of their group and were particularly unwilling to acknowledge the falseness of COVID-19 stories, despite insistent warnings. This poses a serious challenge for policy interventions aimed at combating misinformation and effectively promoting public health information. Whereas previous intervention designs have focused on emphasizing the veracity of news (Pennycook, McPhetres, et al., 2020; Pennycook & Rand, 2019a) or preserving access to reputable news (Allen et al., 2020), more effective interventions may need to target aspects of individuals' personality such as conscientiousness to reduce the spread of fake news (Magidson et al., 2014; Roberts et al., 2017).

Limitations and Future Directions

Despite providing extensive experimental evidence that higher trait conscientiousness weakens the positive association between conservative political ideology and the sharing of fake news, our work also suffers from several limitations that could be addressed in future research. Research has shown that MTurk samples can be liable to sampling and data quality issues, and thus may be prone to biased interpretations (Chandler et al., 2019; Stewart et al., 2017). We partially mitigated this concern by replicating our results in a separate sample on Prolific, but to further establish the generalizability of our results it may be necessary to collect a nationally representative sample of participants. However, given that only a subset of the broader population are consistently exposed to fake news, specifically those who spent a lot of time on the Internet, social media, and public forums, an MTurk sample may constitute the relevant population of study, as individuals present on this platform spend a lot of time working online (Pennycook & Rand, 2019b). Similarly, whereas we presented participants with a mixture of concordant, discordant, and neutral news items, this may not be reflective of the context in which people face news in social media environments (Pennycook & Rand, 2019b). Given the increasing prevalence of online echo chambers, it may be more likely that individuals are overexposed to a subset of politically concordant stories (Bail et al., 2018; Bakshy et al., 2015). This could have implications for how people respond to the diverse stories they encounter online.

Another potential limitation is that we primarily relied on a single item measure of political ideology, as per Graham et al. (2009). It is possible that other measures of ideology might display different relationships with fake news sharing behavior. For example, in a different context, researchers have shown that anger leads

to support for economic but not sociocultural conservatism (Kettle & Salerno, 2017). Yet although political views are multifaceted, a single liberal-conservative continuum has been shown to have adequate predictive validity across a wide range of issues (Jost, 2006; Jost et al., 2008). Further, the present research is primarily concerned with an interaction effect—namely, that higher conscientiousness weakens the positive association between more conservative political ideology and the sharing of fake news. Future research could expand the range of ideological measures to tease apart what specific values and facets of ideology lead to the sharing of fake news and how those facets interact with conscientiousness.

Finally, we find support for desire for chaos as the mechanism using correlational mediation analysis. This method has limitations that have been noted in the literature extensively (Bullock et al., 2010; Green et al., 2010; Imai et al., 2011; Rohrer, 2018). Establishing causality is particularly difficult when studying the interaction of political ideology and personality—both relatively stable factors that are not easily manipulable. Thus, our analysis of the mechanism underpinning the interactive effect of political ideology and conscientiousness should be interpreted with caution.

Conclusion

Dr. Anthony Fauci, the director of the National Institute of Allergy and Infectious Diseases, publicly disagreed with President Trump's stance on the Coronavirus, leading to malicious backlash. Fake stories discrediting his expert opinion were shared millions of times on social media (Alba & Frenkel, 2020). The reconciliation of public opinion with scientific experts is paramount (Scheufele & Krause, 2019) and becomes even more critical during a crisis. For if expert opinions are dismissed as fake news, and such recommendations are ignored, the consequences can be palpable. Hence, it is not surprising that the United States catapulted into the number one location of COVID-19 infections in the world, with more than half a million dead and the count still rising. Our work provides further understanding of what leads to the spread of such malicious falsehoods and offers a sobering perspective on the prospects of interventions aimed at counteracting the dissemination of fake news.

Context

Conservative political ideology has been identified as a key predictor of sharing fake news. However, despite this empirical evidence, blaming all individuals on one side of the political divide runs the risk of further polarization and intergroup conflict that is already too high. We offer a more balanced perspective by demonstrating that low conscientiousness conservatives drive differences in the spread of misinformation across party lines. This effect was explained by their greater desire for chaos. There were no differences among liberals and conservatives who were high on conscientiousness. Our work therefore avoids denouncing all individuals who identify with conservative values. Finally, by testing our theory in the context of COVID-19, our results highlight the inadequacy of fact-checker interventions to curtail the dissemination of fake news and the partisan tribalism that drives these effects. This research advances our understanding of partisan differences and the role of personality in the sharing of fake news, offering insights that can guide the design of effective interventions.

References

Abramowitz, A. I., & Webster, S. (2016). The rise of negative partisanship and the nationalization of U.S. elections in the 21st century. *Electoral Studies*, *41*, 12–22. https://doi.org/10.1016/j.electstud.2015.11.001

Alba, D., & Frenkel, S. (2020, March 28). Medical expert who corrects Trump is now a target of the far right. *The New York Times*. https://www.nytimes.com/2020/03/28/technology/coronavirus-fauci-trump-conspiracy-target.html

Allcott, H., & Gentzkow, M. (2017). Social media and fake news in the 2016 election. *The Journal of Economic Perspectives*, *31*(2), 211–236. https://doi.org/10.1257/jep.31.2.211

Allen, J., Howland, B., Mobius, M., Rothschild, D., & Watts, D. J. (2020). Evaluating the fake news problem at the scale of the information ecosystem. *Science Advances*, *6*(14), eaay3539. https://doi.org/10.1126/sciadv.aay3539

Arceneaux, K., Gravelle, T. B., Osmundsen, M., Petersen, M. B., Reifler, J., & Scotto, T. J. (2021). Some people just want to watch the world burn: The prevalence, psychology and politics of the 'Need for Chaos'. *Philosophical Transactions of the Royal Society of London. Series B, Biological Sciences*, *376*(1822), 20200147. https://doi.org/10.1098/rstb.2020.0147

Bago, B., Rand, D. G., & Pennycook, G. (2020). Fake news, fast and slow: Deliberation reduces belief in false (but not true) news headlines. *Journal of Experimental Psychology: General*, *149*(8), 1608–1613. https://doi.org/10.1037/xge0000729

Bail, C. A., Argyle, L. P., Brown, T. W., Bumpus, J. P., Chen, H., Hunzaker, M. B. F., Lee, J., Mann, M., Merhout, F., & Volfovsky, A. (2018). Exposure to opposing views on social media can increase political polarization. *Proceedings of the National Academy of Sciences of the United States of America*, *115*(37), 9216–9221. https://doi.org/10.1073/pnas.1804840115

Bakshy, E., Messing, S., & Adamic, L. A. (2015). Political science. Exposure to ideologically diverse news and opinion on Facebook. *Science*, *348*(6239), 1130–1132. https://doi.org/10.1126/science.aaa1160

Basol, M., Roozenbeek, J., & van der Linden, S. (2020). Good news about bad news: Gamified inoculation boosts confidence and cognitive immunity against fake news. *Journal of Cognition*, *3*(1), 2. https://doi.org/10.5334/joc.91

Bogg, T., & Roberts, B. W. (2004). Conscientiousness and health-related behaviors: A meta-analysis of the leading behavioral contributors to mortality. *Psychological Bulletin*, *130*(6), 887–919. https://doi.org/10.1037/0033-2909.130.6.887

Boulianne, S., & Koc-Michalska, K. (2021). The role of personality in political talk and like-minded discussion. *The International Journal of Press/Politics*. Advance online publication. https://doi.org/10.1177/1940161221994096

Bullock, J. G., Green, D. P., & Ha, S. E. (2010). Yes, but what's the mechanism? (don't expect an easy answer). *Journal of Personality and Social Psychology*, *98*(4), 550–558. https://doi.org/10.1037/a0018933

Chandler, J., Rosenzweig, C., Moss, A. J., Robinson, J., & Litman, L. (2019). Online panels in social science research: Expanding sampling methods beyond Mechanical Turk. *Behavior Research Methods*, *51*(5), 2022–2038. https://doi.org/10.3758/s13428-019-01273-7

Chmielewski, M., & Kucker, S. C. (2020). An MTurk crisis? Shifts in data quality and the impact on study results. *Social Psychological & Personality Science*, *11*(4), 464–473. https://doi.org/10.1177/1948550619875149

Cohen, T. R., Panter, A. T., Turan, N., Morse, L., & Kim, Y. (2014). Moral character in the workplace. *Journal of Personality and Social Psychology*, *107*(5), 943–963. https://doi.org/10.1037/a0037245

Coppock, A. (2019). Generalizing from survey experiments conducted on Mechanical Turk: A replication approach. *Political Science Research and Methods*, *7*(3), 613–628. https://doi.org/10.1017/psrm.2018.10

Courneya, K. S., Bobick, T. M., & Schinke, R. J. (1999). Does the theory of planned behavior mediate the relation between personality and exercise behavior? *Basic and Applied Social Psychology*, *21*(4), 317–324. https://doi.org/10.1207/S15324834BASP2104_5

Ditto, P. H., Liu, B. S., Clark, C. J., Wojcik, S. P., Chen, E. E., Grady, R. H., Celniker, J. B., & Zinger, J. F. (2019). At least bias is bipartisan: A meta-analytic comparison of partisan bias in liberals and conservatives. *Perspectives on Psychological Science*, *14*(2), 273–291. https://doi.org/10.1177/1745691617746796

Douglas, K. M., Sutton, R. M., & Cichocka, A. (2017). The psychology of conspiracy theories. *Current Directions in Psychological Science*, *26*(6), 538–542. https://doi.org/10.1177/0963721417718261

Douglas, K. M., Uscinski, J. E., Sutton, R. M., Cichocka, A., Nefes, T., Ang, C. S., & Deravi, F. (2019). Understanding conspiracy theories. *Political Psychology*, *40*(S1), 3–35. https://doi.org/10.1111/pops.12568

Duarte, J. L., Crawford, J. T., Stern, C., Haidt, J., Jussim, L., & Tetlock, P. E. (2015). Political diversity will improve social psychological science. *Behavioral and Brain Sciences*, *38*, 1–13. https://doi.org/10.1017/S0140525X14000430

Everett, J. A. C. (2013). The 12 item Social and Economic Conservatism Scale (SECS). *PLoS ONE*, *8*(12), e82131. https://doi.org/10.1371/journal.pone.0082131

Fayard, J. V., Roberts, B. W., Robins, R. W., & Watson, D. (2012). Uncovering the affective core of conscientiousness: The role of self-conscious emotions. *Journal of Personality*, *80*(1), 1–32. https://doi.org/10.1111/j.1467-6494.2011.00720.x

Finkel, E. J., & Campbell, W. K. (2001). Self-control and accommodation in close relationships: An interdependence analysis. *Journal of Personality and Social Psychology*, *81*(2), 263–277. https://doi.org/10.1037/0022-3514.81.2.263

Frederick, S. (2005). Cognitive reflection and decision making. *The Journal of Economic Perspectives*, *19*(4), 25–42. https://doi.org/10.1257/089533005775196732

Gerhardt, M. W., Rode, J. C., & Peterson, S. J. (2007). Exploring mechanisms in the personality–performance relationship: Mediating roles of self-management and situational constraints. *Personality and Individual Differences*, *43*(6), 1344–1355. https://doi.org/10.1016/j.paid.2007.04.001

Graham, J., Haidt, J., & Nosek, B. A. (2009). Liberals and conservatives rely on different sets of moral foundations. *Journal of Personality and Social Psychology*, *96*(5), 1029–1046. https://doi.org/10.1037/a0015141

Green, D. P., Ha, S. E., & Bullock, J. G. (2010). Enough already about "black box" experiments: Studying mediation is more difficult than most scholars suppose. *The Annals of the American Academy of Political and Social Science*, *628*(1), 200–208. https://doi.org/10.1177/0002716209351526

Grinberg, N., Joseph, K., Friedland, L., Swire-Thompson, B., & Lazer, D. (2019). Fake news on Twitter during the 2016 U.S. presidential election. *Science*, *363*(6425), 374–378. https://doi.org/10.1126/science.aau2706

Guess, A. M., Nyhan, B., & Reifler, J. (2020). Exposure to untrustworthy websites in the 2016 U.S. election. *Nature Human Behaviour*, *4*(5), 472–480. https://doi.org/10.1038/s41562-020-0833-x

Guess, A., Nagler, J., & Tucker, J. (2019). Less than you think: Prevalence and predictors of fake news dissemination on Facebook. *Science Advances, 5*, *5*(1), eaau4586. https://doi.org/10.1126/sciadv.aau4586

Hardin, J. W., & Hilbe, J. M. (2003). *Generalized estimating equations.* Wiley.

Imai, K., Keele, L., Tingley, D., & Yamamoto, T. (2011). Unpacking the black box of causality: Learning about causal mechanisms from experimental and observational studies. *The American Political Science Review*, *105*(4), 765–789. https://doi.org/10.2307/23275352

Jackson, J. J., Wood, D., Bogg, T., Walton, K. E., Harms, P. D., & Roberts, B. W. (2010). What do conscientious people do? Development and validation of the Behavioral Indicators of Conscientiousness (BIC). *Journal of Research in Personality*, *44*(4), 501–511. https://doi.org/10.1016/j.jrp.2010.06.005

Jost, J. T. (2006). The end of the end of ideology. *American Psychologist*, *61*(7), 651–670. https://doi.org/10.1037/0003-066X.61.7.651

Jost, J. T. (2017). Asymmetries abound: Ideological differences in emotion, partisanship, motivated reasoning, social network structure, and political trust. *Journal of Consumer Psychology*, *27*(4), 546–553. https://doi.org/10.1016/j.jcps.2017.08.004

Jost, J. T., Glaser, J., Kruglanski, A. W., & Sulloway, F. J. (2003). Political conservatism as motivated social cognition. *Psychological Bulletin*, *129*(3), 339–375. https://doi.org/10.1037/0033-2909.129.3.339

Jost, J. T., Nosek, B. A., & Gosling, S. D. (2008). Ideology: Its resurgence in social, personality, and political psychology. *Perspectives on Psychological Science*, *3*(2), 126–136. https://doi.org/10.1111/j.1745-6916.2008.00070.x

Jost, J. T., van der Linden, S., Panagopoulos, C., & Hardin, C. D. (2018). Ideological asymmetries in conformity, desire for shared reality, and the spread of misinformation. *Current Opinion in Psychology*, *23*, 77–83. https://doi.org/10.1016/j.copsyc.2018.01.003

Kennedy, R., Clifford, S., Burleigh, T., Waggoner, P. D., Jewell, R., & Winter, N. J. G. (2020). The shape of and solutions to the MTurk quality crisis. *Political Science Research and Methods*, *8*(4), 614–629. https://doi.org/10.1017/psrm.2020.6

Kettle, K. L., & Salerno, A. (2017). Anger promotes economic conservatism. *Personality and Social Psychology Bulletin*, *43*(10), 1440–1454. https://doi.org/10.1177/0146167217718169

Koehn, M. A., Jonason, P. K., & Davis, M. D. (2019). A person-centered view of prejudice: The Big Five, Dark Triad, and prejudice. *Personality and Individual Differences*, *139*, 313–316. https://doi.org/10.1016/j.paid.2018.11.038

Koole, S. L., Jager, W., van den Berg, A. E., Vlek, C. A. J., & Hofstee, W. K. B. (2001). On the social nature of personality: effects of extraversion, agreeableness, and feedback about collective resource use on cooperation in a resource dilemma. *Personality and Social Psychology Bulletin*, *27*(3), 289–301. https://doi.org/10.1177/0146167201273003

Krishna, A., & Sokolova, T. (2017). A focus on partisanship: How it impacts voting behaviors and political attitudes. *Journal of Consumer Psychology*, *27*(4), 537–545. https://doi.org/10.1016/j.jcps.2017.07.005

Krizan, Z., & Johar, O. (2015). Narcissistic rage revisited. *Journal of Personality and Social Psychology*, *108*(5), 784–801. https://doi.org/10.1037/pspp0000013

Krupnikov, Y., & Levine, A. (2014). Cross-sample comparisons and external validity—ProQuest. *Journal of Experimental Political Science*, *1*(1), 59–80. https://doi.org/10.1017/xps.2014.7

Kugler, M., Jost, J. T., & Noorbaloochi, S. (2014). Another look at moral foundations theory: Do authoritarianism and social dominance orientation explain liberal-conservative differences in "moral" intuitions? *Social Justice Research*, *27*(4), 413–431. https://doi.org/10.1007/s11211-014-0223-5

Lazer, D. M. J., Baum, M. A., Benkler, Y., Berinsky, A. J., Greenhill, K. M., Menczer, F., Metzger, M. J., Nyhan, B., Pennycook, G., Rothschild, D., Schudson, M., Sloman, S. A., Sunstein, C. R., Thorson, E. A., Watts, D. J., & Zittrain, J. L. (2018). The science of fake news. *Science*, *359*(6380), 1094–1096. https://doi.org/10.1126/science.aao2998

Liang, K.-Y., & Zeger, S. L. (1986). Longitudinal data analysis using generalized linear models. *Biometrika*, *73*(1), 13–22. https://doi.org/10.1093/biomet/73.1.13

Litman, L., Robinson, J., & Abberbock, T. (2017). TurkPrime.com: A versatile crowdsourcing data acquisition platform for the behavioral sciences. *Behavior Research Methods*, *49*(2), 433–442. https://doi.org/10.3758/s13428-016-0727-z

Magidson, J. F., Roberts, B. W., Collado-Rodriguez, A., & Lejuez, C. W. (2014). Theory-driven intervention for changing personality: Expectancy

value theory, behavioral activation, and conscientiousness. *Developmental Psychology*, *50*(5), 1442–1450. https://doi.org/10.1037/a0030583

Martel, C., Pennycook, G., & Rand, D. (2019). *Reliance on emotion promotes belief in fake news.* https://doi.org/10.31234/osf.io/a2ydw

McCrae, R. R., & Costa, P. T., Jr. (2008). The five-factor theory of personality. In: O. P. John, R. W. Robins, & L. A. Pervin (Eds.), *Handbook of personality: Theory and research* (pp. 159–181). The Guilford Press.

Moskalenko, S., & McCauley, C. (2009). Measuring Political Mobilization: The Distinction Between Activism and Radicalism. *Terrorism and Political Violence*, *21*(2), 239–260. https://doi.org/10.1080/09546550902765508

Mullinix, K. J., Leeper, T. J., Druckman, J. N., & Freese, J. (2015). The generalizability of survey experiments. *Journal of Experimental Political Science*, *2*(2), 109–138. https://doi.org/10.1017/XPS.2015.19

Oyserman, D., & Schwarz, N. (2017). Conservatism as a situated identity: Implications for consumer behavior. *Journal of Consumer Psychology*, *27*(4), 532–536. https://doi.org/10.1016/j.jcps.2017.08.003

Paolacci, G., & Chandler, J. (2014). Inside the Turk: Understanding Mechanical Turk as a participant pool. *Current Directions in Psychological Science*, *23*(3), 184–188. https://doi.org/10.1177/0963721414531598

Paunonen, S. V., & Ashton, M. C. (2001). Big five factors and facets and the prediction of behavior. *Journal of Personality and Social Psychology*, *81*(3), 524–539. https://doi.org/10.1037/0022-3514.81.3.524

Pennycook, G., & Rand, D. G. (2019a). Fighting misinformation on social media using crowdsourced judgments of news source quality. *Proceedings of the National Academy of Sciences of the United States of America*, *116*(7), 2521–2526. https://doi.org/10.1073/pnas.1806781116

Pennycook, G., & Rand, D. G. (2019b). Lazy, not biased: Susceptibility to partisan fake news is better explained by lack of reasoning than by motivated reasoning. *Cognition*, *188*, 39–50. https://doi.org/10.1016/j.cognition.2018.06.011

Pennycook, G., & Rand, D. G. (2020). Who falls for fake news? The roles of bullshit receptivity, overclaiming, familiarity, and analytic thinking. *Journal of Personality*, *88*(2), 185–200. https://doi.org/10.1111/jopy.12476

Pennycook, G., Bear, A., Collins, E. T., & Rand, D. G. (2020). The implied truth effect: Attaching warnings to a subset of fake news headlines increases perceived accuracy of headlines without warnings. *Management Science*, *66*(11), 4944–4957. https://doi.org/10.1287/mnsc.2019.3478

Pennycook, G., Cannon, T. D., & Rand, D. G. (2018). Prior exposure increases perceived accuracy of fake news. *Journal of Experimental Psychology: General*, *147*(12), 1865–1880. https://doi.org/10.1037/xge0000465

Pennycook, G., Epstein, Z., Mosleh, M., Arechar, A. A., Eckles, D., & Rand, D. (2021). Shifting attention to accuracy can reduce misinformation online. *Nature*, *592*(7855), 590–595. https://doi.org/10.31234/osf.io/3n9u8

Pennycook, G., McPhetres, J., Zhang, Y., Lu, J. G., & Rand, D. G. (2020). Fighting COVID-19 misinformation on social media: Experimental evidence for a scalable accuracy-nudge intervention. *Psychological Science*, *31*(7), 770–780. https://doi.org/10.1177/0956797620939054

Petersen, M. B., Osmundsen, M., & Arceneaux, K. (2020). *A "need for chaos" and the sharing of hostile political rumors in advanced democracies*. [Preprint]. PsyArXiv. https://doi.org/10.31234/osf.io/6m4ts

Petrocelli, J. V., Watson, H. F., & Hirt, E. R. (2020). Self-regulatory aspects of bullshitting and bullshit detection. *Social Psychology*, *51*(4), 239–253. https://doi.org/10.1027/1864-9335/a000412

Rajan, A. (2020, March 14). Coronavirus and a fake news pandemic—BBC News. *BBC*. https://www.bbc.com/news/entertainment-arts-51858555

Roberts, B. W., Harms, P. D., Caspi, A., & Moffitt, T. E. (2007). Predicting the counterproductive employee in a child-to-adult prospective study. *Journal of Applied Psychology*, *92*(5), 1427–1436. https://doi.org/10.1037/0021-9010.92.5.1427

Roberts, B. W., Hill, P. L., & Davis, J. P. (2017). How to change conscientiousness: The sociogenomic trait intervention model. *Personality Disorders*, *8*(3), 199–205. https://doi.org/10.1037/per0000242

Roberts, B. W., Jackson, J. J., Fayard, J. V., Edmonds, G., & Meints, J. (2009). Conscientiousness. *Handbook of individual differences in social behavior* (pp. 369–381). Guilford Press.

Rohrer, J. M. (2018). Thinking clearly about correlations and causation: graphical causal models for observational data. *Advances in Methods and Practices in Psychological Science*, *1*(1), 27–42. https://doi.org/10.1177/2515245917745629

Roozenbeek, J., & van der Linden, S. (2019). Fake news game confers psychological resistance against online misinformation. *Palgrave Communications*, *5*(1), 1–10. https://doi.org/10.1057/s41599-019-0279-9

Saklofske, D. H., Austin, E. J., Galloway, J., & Davidson, K. (2007). Individual difference correlates of health-related behaviours: Preliminary evidence for links between emotional intelligence and coping. *Personality and Individual Differences*, *42*(3), 491–502. https://doi.org/10.1016/j.paid.2006.08.006

Scheufele, D. A., & Krause, N. M. (2019). Science audiences, misinformation, and fake news. *Proceedings of the National Academy of Sciences of the United States of America*, *116*(16), 7662–7669. https://doi.org/10.1073/pnas.1805871115

Sibley, C. G., Osborne, D., & Duckitt, J. (2012). Personality and political orientation: Meta-analysis and test of a Threat-Constraint Model. *Journal of Research in Personality*, *46*(6), 664–677. https://doi.org/10.1016/j.jrp.2012.08.002

Soto, C. J., & John, O. P. (2017). The next Big Five Inventory (BFI-2): Developing and assessing a hierarchical model with 15 facets to enhance bandwidth, fidelity, and predictive power. *Journal of Personality and Social Psychology*, *113*(1), 117–143. https://doi.org/10.1037/pspp0000096

Stewart, N., Chandler, J., & Paolacci, G. (2017). Crowdsourcing samples in cognitive science. *Trends in Cognitive Sciences*, *21*(10), 736–748. https://doi.org/10.1016/j.tics.2017.06.007

The Associated Press. (2020, March 27). In Iran, false belief a poison fights virus kills hundreds. *The New York Times*. https://www.nytimes.com/aponline/2020/03/27/world/middleeast/ap-ml-virus-outbreak-iran-a-deadly-drink.html

Thielmann, I., Spadaro, G., & Balliet, D. (2020). Personality and prosocial behavior: A theoretical framework and meta-analysis. *Psychological Bulletin*, *146*(1), 30–90. https://doi.org/10.1037/bul0000217

Thomson, K. S., & Oppenheimer, D. M. (2016). Investigating an alternate form of the cognitive reflection test. *Judgment and Decision Making*, *11*(1), 15.

Twenge, J. M., & Campbell, W. K. (2003). Isn't it fun to get the respect that we're going to deserve?" Narcissism, social rejection, and aggression. *Personality and Social Psychology Bulletin*, *29*(2), 261–272. https://doi.org/10.1177/0146167202239051

Tybur, J. M., Lieberman, D., & Griskevicius, V. (2009). Microbes, mating, and morality: Individual differences in three functional domains of disgust. *Journal of Personality and Social Psychology*, *97*(1), 103–122. https://doi.org/10.1037/a0015474

Van Bavel, J. J., & Pereira, A. (2018). The partisan brain: An identity-based model of political belief. *Trends in Cognitive Sciences*, *22*(3), 213–224. https://doi.org/10.1016/j.tics.2018.01.004

Van den Bos, K., Euwema, M. C., Poortvliet, P. M., & Maas, M. (2007). Uncertainty management and social issues: Uncertainty as an important determinant of reactions to socially deviating people. *Journal of Applied Social Psychology*, *37*(8), 1726–1756. https://doi.org/10.1111/j.1559-1816.2007.00235.x

van der Linden, S., Panagopoulos, C., & Roozenbeek, J. (2020). You are fake news: Political bias in perceptions of fake news. *Media Culture & Society*, *42*(3), 460–470. https://doi.org/10.1177/0163443720906992

Linden, S., Panagopoulos, C., Azevedo, F., & Jost, J. T. (2021). The paranoid style in American politics revisited: An ideological asymmetry in conspiratorial thinking. *Political Psychology*, *42*(1), 23–51. https://doi.org/10.1111/pops.12681

van Leeuwen, F., & Park, J. H. (2009). Perceptions of social dangers, moral foundations, and political orientation. *Personality and Individual Differences*, *47*(3), 169–173. https://doi.org/10.1016/j.paid.2009.02.017

Vosoughi, S., Roy, D., & Aral, S. (2018). The spread of true and false news online. *Science*, *359*(6380), 1146–1151. https://doi.org/10.1126/science.aap9559

Westwood, S. J., Iyengar, S., Walgrave, S., Leonisio, R., Miller, L., & Strijbis, O. (2018). The tie that divides: Cross-national evidence of the primacy of partyism. *European Journal of Political Research*, *57*(2), 333–354. https://doi.org/10.1111/1475-6765.12228

Received September 30, 2020
Revision received May 8, 2021
Accepted June 18, 2021 ■