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A Longitudinal Test of the Conservative-Liberal Well-Being Gap

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Abstract

In this article, we test if conservatism predicts psychological well-being longitudinally. We based the study on previous findings showing that conservatives score higher on different measures of well-being, such as life satisfaction and happiness. Most explanations in the literature have assumed that conservatism antecedes well-being without considering the alternative—that well-being may predict conservatism. In Study 1, using multilevel cross-lagged panel models with a two-wave longitudinal sample consisting of data from 19 countries (N = 8,740), we found that conservatism did not predict well-being over time. We found similar results in Study 2 (N = 2,554), using random-intercept cross-lagged panel models with a four-wave longitudinal sample from Chile. We discuss the main implications of these results for the literature examining the association between conservatism and well-being.

Keywords

conservatism, well-being, longitudinal analysis, life satisfaction, happiness

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A striking finding in the psychological literature is that conservatives report greater levels of subjective psychological well-being compared with liberals. Although the effect sizes tend to be small, this pattern of results has been observed for different measures of subjective well-being, such as life satisfaction (Napier & Jost, 2008), happiness (Schlenker et al., 2012), and self-reported health (Subramanian & Perkins, 2010).

Different explanations have been put forward for this phenomenon (Butz et al., 2017). System justification theorists have proposed that conservative individuals score higher in system justification beliefs (i.e., a motivation to perceive social systems as fair and legitimate), providing an ideological rationalization of the status quo that alleviates negative psychological consequences brought about by societal inequalities (Harding & Sibley, 2013; Jost & Hunyady, 2002). Others have attributed differences in subjective well-being between conservatives and liberals to personality traits, demonstrating that greater well-being outcomes (e.g., life satisfaction and happiness) correspond to greater agency beliefs and a more positive outlook among conservatives (Schlenker et al., 2012). Furthermore, Jetten et al. (2013) showed that higher socioeconomic status among conservatives enables them to belong to and participate in more social groups and in turn reap the well-being benefits of multigroup membership. Finally, Stavrova and Luhmann (2016) proposed that the conservative-liberal subjective well-being gap can be explained by person–culture fit, illustrating that conservatives only report greater well-being in sociocultural contexts within which conservative political ideology prevails.

Regardless of the differences between these theoretical accounts, the extant literature on the relationship between conservatism and subjective psychological well-being suffers from an important caveat. Most of the studies share the same assumption about the causal order of the relationship between conservatism and subjective well-being—that conservativism predicts well-being (e.g., Briki & Dagot, 2020; Butz et al., 2017; Napier & Jost, 2008; Schlenker et al., 2012; Stavrova & Luhmann, 2016; Subramanian et al., 2009; Subramanian & Perkins, 2010). However, none of these studies have provided empirical evidence for this order. Few studies examining the relationship with proximal measures of conservatism, mainly system justification beliefs, have

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utilized either longitudinal (Godfrey et al., 2017; Vargas-Salfate et al., 2018) or experimental (Li et al., 2020) research designs. This gap in the literature is relevant because two alternative explanations could be proposed. First, third variables might cause both conservatism and subjective well-being (e.g., socioeconomic status; Jetten et al., 2013). And second, psychological well-being might antecede conservatism. People experiencing greater levels of subjective well-being may be more likely to adopt ideologies that protect their subjective and objective status (Subramanian et al., 2009) and also be less attentive to injustice cues that might lead them to adopt conservative ideologies (Napier et al., 2020).

The Present Research

In this research, we examined whether conservatism predicted subjective well-being longitudinally in two studies. Based on previous findings on system justification theory that have utilized longitudinal or experimental designs, we expect that conservatism will predict subjective well-being over time. However, the use of longitudinal data in these studies allowed us to test the alternative hypothesis that subjective well-being predicts conservatism over time. In addition, we used multiple indicators of subjective well-being, capturing both cognitive evaluations and affective reactions (Diener, 2009)—specifically life satisfaction, anxiety, depression, and self-reported health. In both studies, we relied on secondary data and therefore used all indicators available to test our hypothesis. In Study 1, we used a two-wave survey conducted in 19 countries (N = 8,740), and in Study 2, we used a four-wave representative survey conducted in Chile (N = 2,554). In both studies, we controlled for demographic variables to maintain comparability with previous research (e.g., Napier & Jost, 2008; Schlenker et al., 2012), with a special focus on the subjective social status given that this variable has been proposed as an alternative account for the conservative-liberal subjective well-being gap (Jetten et al., 2013). We present results both with control and without control variables. All study materials can be accessed at https://osf.io/z53cq.

Study 1

Method

Sample. Participants were recruited through an international online panel maintained by Nielsen during September 2015 (T1) and then 6 months later (T2) in the context of the Digital Influence Project.[1] The sample was stratified by age, gender, and region based on census data (for more information on this project, see Gil de Zúñiga & Liu, 2017). We selected all cases that participated in both waves, which resulted in a final sample size of 8,740 participants: Argentina (N = 360), Brazil (N = 353), China (N = 387), Estonia (N = 733), Germany (N = 643), Indonesia (N = 305), Italy (N = 579), Japan (N = 574), South Korea (N = 572), New Zealand (N = 605), Philippines (N = 153), Poland (N = 628), Russia (N = 551), Spain (N = 301), Taiwan (N = 426), Turkey (N = 331), United Kingdom (N = 649), Ukraine (N = 101), and the United States (N = 489). We included all countries in the Digital Influence Project, except India and South Africa. Data for these two countries have not been made available to researchers because of the use of nonrepresentative (city-based) samples. The mean age was 45.57 (SD = 14.61), and 51.10% of participants were female.

Measures

Conservatism. We used a 3-item measure to assess conservatism (Pratto et al., 1997), for example: *On economic/political/social issues, where would you place yourself on a scale of 0 to 10?*. Answers to these items ranged from 0 (*strong liberal*) to 10 (*strong conservative*). This scale was highly reliable at both T1 (α = .93) and T2 (α = .93).

Life satisfaction. We used 5 items from the Personal Well-being Index (PWI; Lau et al., 2005). These items measure overall life satisfaction and satisfaction with health, standard of living, safety, and relationships. Answers to these items ranged from 1 (*completely dissatisfied*) to 7 (*completely satisfied*). This scale was highly reliable at both T1 (α = .87) and T2 (α = .87).

Anxiety. We used the Generalized Anxiety Disorder scale (GAD; Spitzer et al., 2006), which includes 7 items assessing the two main criteria of generalized anxiety disorder: excessive anxiety and worry and difficulties controlling worries during the last 2 weeks. Responses ranged from 1 (*never*) to 7 (*always*). This scale was highly reliable at both T1 (α = .94) and T2 (α = .94).

Depression. We used the Patient Health Questionnaire-4 (PHQ-4; Löwe et al., 2010) to measure depression. Participants were asked to report the frequency of several depressive symptoms during the last two weeks using a scale ranging from 1 (*never*) to 7 (*always*). This scale was highly reliable at both T1 (α = .89) and T2 (α = .90).

Health status. We adopted 2 items from DeSalvo et al. (2006) to assess health status. Participants rated their own health status (*How would you say your health is?*) and their health status compared with other of their age (*Compared to others your age, how would you say your health is?*) using a scale ranging from 1 (*very poor*) to 7 (*excellent*). This scale was highly reliable at both T1 (α = .93) and T2 (α = .95).

Control variables. In all our models, we controlled for age, gender (0 = *male*, 1 = *female*), and subjective social status (1 = *low status*, 10 = *high status*; adapted from Adler et al., 2000). We also included a measure of household income, which was available in the Digital Influence Project coded in percentiles (1 = *0–10 percentile*, 2 = *11–30 percentile*, 3 = *31–70 percentile*, 4 = *71–90 percentile*, and 5 = *91–100 percentile*). Given that the nonresponse rate in income was higher than for the rest of the variables (15.3%), we included this control variable in separate models.

Results and Discussion

Descriptive statistics and correlations matrix are shown in Table 1. Descriptive statistics by country can be found in the Online Supplemental Material (OSM; Table S1).

Before conducting the main analyses, we ran a series of confirmatory factor analyses to test measurement invariance both at the longitudinal level of analysis and between countries (see OSM Table S2). Overall, these analyses indicate that the measurement models showed metric invariance across countries and strict invariance across both waves.

We used cross-lagged panel models (CLPMs) with latent variables to test our hypotheses (Raykov & Marcoulides, 2006) considering the multilevel structure of the data in which individuals are nested within countries (Heck, 2009). In a first step, these models specified that, at the individual level, conservatism (T1) would predict well-being (T2) over time, and that well-being (T1) would predict conservatism (T2) over time, while controlling for the autoregressors. In a second step, we also included the control variables (i.e., gender, age, and social status). In a third step, we added income as an additional control variable. We specified equivalent yet separate models for life satisfaction, anxiety, depression, and health status. We also specified covariances between the residuals of observed variables at T1 and T2 in these models (Biesanz, 2012). At the country level, we did not have specific hypotheses; therefore, we only included the measurement models, but we excluded covariances between the residuals of observed variables at T1 and T2 given that we only had 19 countries. Missing values were treated using full information maximum likelihood and all models were conducted through Mplus v.6.12 (Muthen & Muthen, 2012).[2] Following, Newman's (2014) recommendations, we estimated the missing data patterns at the item-level (i.e., an individual not responding to an individual item), construct-level (i.e., an individual not responding to all the items measuring the same construct), and person-level (i.e., an individual not responding to all the items measuring all the constructs in each of the main analyses). We found that the most important source of missing data was income, and the percentages of partial respondents (i.e., individuals without responses for all the constructs) were all below 30%, which is the threshold suggested by Newman (2014) to use full information maximum likelihood as an appropriate form to handle missing data (see OSM Table S77). Furthermore, we evaluated the goodness of fit in our models using three indexes (Raykov & Macrolides', 2006): Values higher than .95 for the Comparative Fit Index (CFI) and the Tucker Lewis Index (TLI) and lower than .06 for the Root Mean Square Error of Approximation (RMSEA) were treated as evidence supporting appropriate goodness of fit (Hu & Bertler, 1999). Importantly, given that we had only two waves in Study 1, these goodness-of-fit statistics provide information about the fit of the measurement model but not the structural model as we modeled all the associations between the latent constructs (Hamaker et al., 2015). Here, we only present the relevant information to our theoretical discussion, which is related to the individual level. All the relevant parameters of our models can be found in the OSM (Tables S7-S18). An example of this analytic approach at the individual level can be found in Figure 1.

Life satisfaction. The model for life satisfaction without covariates showed an appropriate goodness of fit, χ^2 (192) = 1,455.97, p < .001, CFI = .988, TLI = .985, RMSEA = .027. Conservatism positively predicted life satisfaction over time, b = .01, β = .02, p = .003, 95% confidence interval (CI): [.01, .02], and life satisfaction positively predicted conservatism over time, b = .05, β = .03, p < .001, 95% CI [.03, .08]. When we included the control variables, we also found an appropriate goodness of fit, χ^2 (234) = 5,226.17, p < .001, CFI = .952, TLI = .941, RMSEA = .050, but conservatism did not predict life satisfaction over time, b = <.01, β = .01, p = .409, 95% CI [−.01, .01], and life satisfaction predicted conservatism over time only at α = .05 (i.e., p < .05), b = .04, β = .02, p = .040, 95% CI [<.01, .07].

Regarding the control variables, subjective social status predicted life satisfaction at T2, b = .11, β = .17, p < .001, 95% CI [.09, .12], but not conservatism, b = .02, β = .02, p = .091, 95% CI [−<.00, .04]; gender did not predict life satisfaction, b = .03, β = .01, p = .107, 95% CI [−.01, .07], or conservatism, b = −.06, β = −.02, p = .076, 95% CI [−.12, .01]; and age did not predict life satisfaction, b = <.01, β = .01, p = .130, 95% CI [<.01, <.01], or conservatism, b = <.01, β = .02, p = .068, 95% CI [<.01, <.01]. When including income as an additional covariate, the model showed an appropriate goodness of fit, χ^2 (248) = 4551.18, p < .001, CFI = .952, TLI = .941, RMSEA = .049, and income predicted life satisfaction, b = .03, β = .03, p = .007, 95% CI [.01, .05], but not conservatism, b = <.01, β = <.01, p = .828, 95% CI [−.04, .03]. More importantly, in this model conservatism did not predict life satisfaction over time, b = <.01, β = .01, p = .584, 95% CI [−.01, .01], and life satisfaction did not predict conservatism over time either, b = .03, β = .02, p = .098, 95% CI [<.01, <.01].

Anxiety. The model for life satisfaction without covariates showed an appropriate goodness of fit, χ^2 (322) = 4,206.65, p < .001, CFI = .975, TLI = .970, RMSEA = .037. Conservatism did not predict anxiety over time, b = −.01, β = −.01, p = .291, 95% CI [−.02, .01], and anxiety did not predict conservatism over time, b = −.01, β = <.01, p = .622, 95% CI [−.03, −.02]. When we included the control variables, we also found an appropriate goodness of fit, χ^2 (376) = 5,476.97, p < .001, CFI = .967, TLI = .961, RMSEA = .040. In this

Table 1. Descriptive Statistics and Correlations Matrix (Study 1).

Variable	*M*	*SD*	1	2	3	4	5	6	7	8	9	10	11	12	13
1. Conservatism (T1)	4.97	2.25	1												
2. Conservatism (T2)	5.13	2.10	.70***	1											
3. Life satisfaction (T1)	4.83	1.14	.07***	.08***	1										
4. Life satisfaction (T2)	4.79	1.15	.08***	.10***	.76***	1									
5. Anxiety (T1)	3.17	1.38	−.04***	−.04**	−.39***	−.33***	1								
6. Anxiety (T2)	3.15	1.40	−.04***	−.02	−.33***	−.36***	.70***	1							
7. Depression (T1)	3.10	1.42	−.04***	−.04***	−.44***	−.37***	.93***	.67***	1						
8. Depression (T2)	3.09	1.42	−.05***	−.03*	−.38***	−.41***	.68***	.94***	.70***	1					
9. Health status (T1)	4.76	1.32	.06***	.07***	.63***	.52***	−.29***	−.24***	−.32***	−.27***	1				
10. Health status (T2)	4.70	1.31	.06***	.08***	.53***	.61***	−.25***	−.26***	−.28***	−.29***	.77***	1			
11. Social status	5.42	1.81	.15***	.13***	.51***	.47***	−.16***	−.13***	−.19***	−.17***	.32***	.32***	1		
12. Income	3.01	1.07	.03*	.03**	.23***	.23***	−.08***	−.06***	−.09***	−.08***	.15***	.16***	.38***	1	
13. Age	45.57	14.61	.07***	.06***	.08***	.07***	−.27***	−.27***	−.25***	−.26***	−.02*	−.02	−.03**	−.05***	1
14. Gender (female)	51.10%		−.04***	−.04***	.06***	.05***	.12***	.11***	.09***	.09***	.04***	.04***	−.02	−.06***	−.11***

*$p < .05$. **$p < .01$. ***$p < .001$.

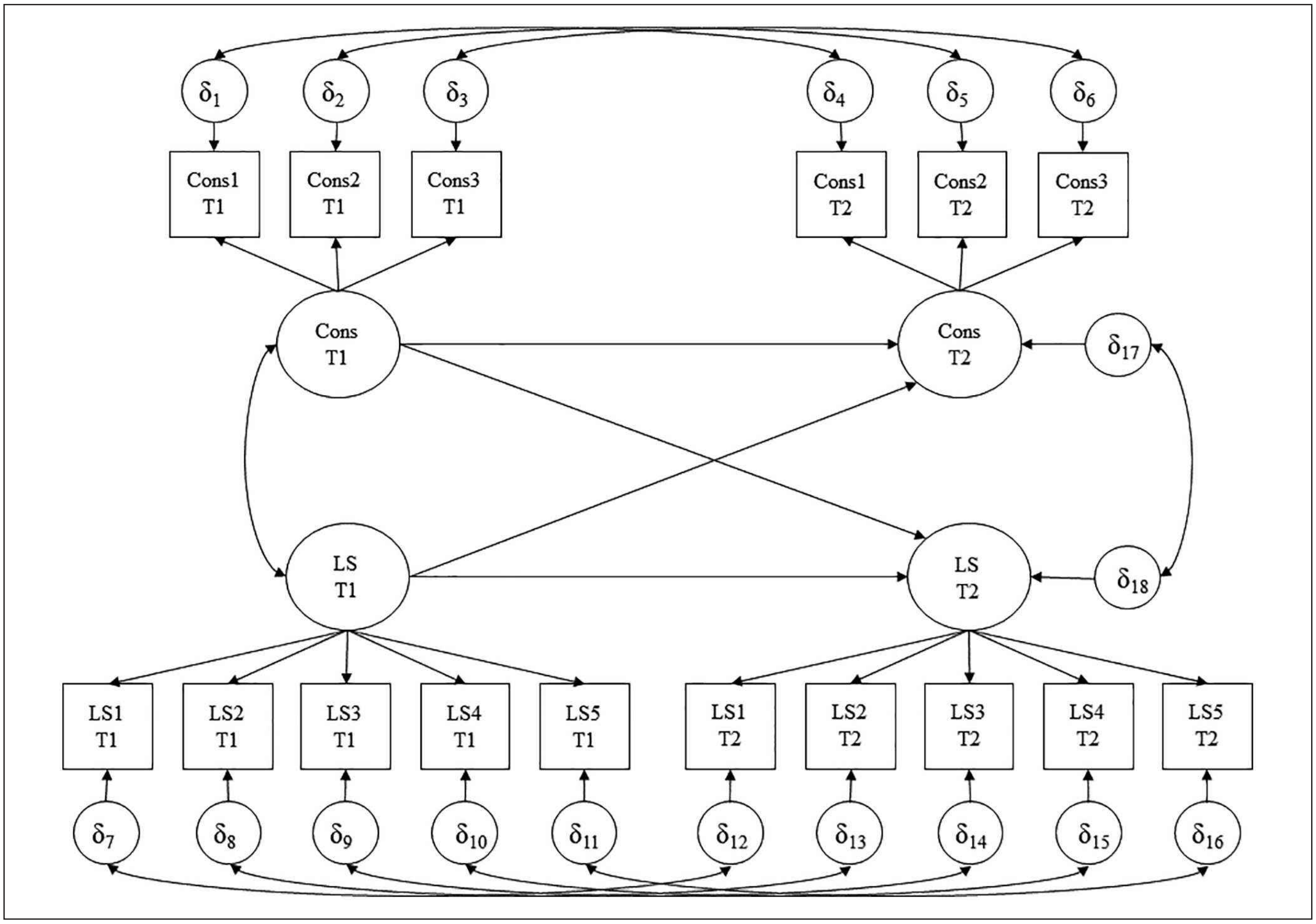

Figure 1. Example of cross-lagged panel model with life satisfaction.
Note. Variables in this example are conservatism (Cons) and Life satisfaction (LS) measured at time 1 (T1) and time 2 (T2). Residuals for each observed and latent variable are indicated by δ.

model, conservatism did not predict anxiety over time, $b = <.01$, $\beta = <.01$, $p = .867$, 95% CI [−.01, .01], and the reciprocal path was not significant, $b = .01$, $\beta = .01$, $p = .362$, 95% CI [−.01, .04].

Regarding the control variables, subjective social status predicted anxiety at T2, $b = -.03$, $\beta = -.04$, $p < .001$, 95% CI [−.04, −.01], and conservatism, $b = .03$, $\beta = .03$, $p < .001$, 95% CI [.02, .05]; gender predicted anxiety, $b = .07$, $\beta = .03$, $p = .001$, 95% CI [.03, .11], but not conservatism, $b = -.05$, $\beta = -.01$, $p = .099$, 95% CI [−.12, .01]; and age did predicted both anxiety, $b = -.01$, $\beta = -.08$, $p < .001$, 95% CI [−.01, −.01], and conservatism, $b = <.01$, $\beta = .02$, $p = .027$, 95% CI [<.01, .01]. When including income as an additional covariate, the model showed appropriate goodness of fit, χ^2 (394) = 4,773.79, $p < .001$, CFI = .967, TLI = .961, RMSEA = .039, and income did not predict anxiety, $b = <.01$, $\beta = <.01$, $p = .903$, 95% CI [−.02, .02], or conservatism, $b = <.01$, $\beta = <.01$, $p = .963$, 95% CI [−.04, .03]. More importantly, in this model conservatism did not predict anxiety over time, $b = <.01$, $\beta = <.01$, $p = .820$, 95% CI [−.01, .01], and anxiety did not predict conservatism over time either, $b = .02$, $\beta = .01$, $p = .196$, 95% CI [−.01, .05].

Depression. The model for depression without covariates showed an appropriate goodness of fit, χ^2 (139) = 4,533.85, $p < .001$, CFI = .955, TLI = .941, RMSEA = .060. Conservatism did not predict depression over time, $b = -.01$, $\beta = -.01$, $p = .185$, 95% CI [−.02, <.01], and depression did not predict conservatism over time, $b = -.01$, $\beta = -.01$, $p = .574$, 95% CI [−.03, .02]. When we included the control variables, we observed a goodness of fit slightly below the thresholds suggested by the literature, χ^2 (175) = 5,858.18, $p < .001$, CFI = .941, TLI = .925, RMSEA = .062. In this model, conservatism did not predict depression over time, $b = <.01$, $\beta = <.01$, $p = .925$, 95% CI [−.01, .01], and the reciprocal path was not significant, $b = .01$, $\beta = .01$, $p = .421$, 95% CI [−.02, .04].

Regarding the control variables, subjective social status predicted depression at T2, $b = -.04$, $\beta = -.06$, $p < .001$, 95% CI [−.05, −.03], and conservatism, $b = .03$, $\beta = .03$,

$p < .001$, 95% CI [.02, .05]; gender predicted depression, $b = .05$, $\beta = .02$, $p = .015$, 95% CI [.01, .10], but not conservatism, $b = -.05$, $\beta = -.01$, $p = .104$, 95% CI [−.12, .01]; and age did predicted both depression, $b = -.01$, $\beta = -.09$, $p < .001$, 95% CI [−.01, −.01], and conservatism, $b = <.01$, $\beta = .02$, $p = .029$, 95% CI [<.01, .01]. When including income as an additional covariate, the model showed a similar goodness of fit than the previous model, χ^2 (187) = 5116.60, $p < .001$, CFI = .41, TLI = .924, RMSEA = .060, and income did not predict depression, $b = -.01$, $\beta = -.01$, $p = .587$, 95% CI [−.03, .02], or conservatism, $b = <.01$, $\beta = <.01$, $p = .956$, 95% CI [−.04, .03]. More importantly, in this model conservatism did not predict depression over time, $b = <.01$, $\beta = <.01$, $p = .736$, 95% CI [−.01, .01], and depression did not predict conservatism over time either, $b = .02$, $\beta = .02$, $p = .275$, 95% CI [−.01, .05].

Health status. The model for health status without covariates showed an appropriate goodness of fit, χ^2 (36) = 291.95, $p < .001$, CFI = .995, TLI = .992, RMSEA = .029. Conservatism positively predicted health status over time, $b = .01$, $\beta = .02$, $p = .048$, 95% CI [−.01, .04], but health status did not predict conservatism over time, $b = .01$, $\beta = .01$, $p = .119$, 95% CI [−.01, .04]. When we included the control variables, we also found an appropriate goodness of fit, χ^2 (54) = 1,428.37, $p < .001$, CFI = .974, TLI = .961, RMSEA = .055. In this model, conservatism did not predict health status over time, $b = <.01$, $\beta = <01$, $p = .385$, 95% CI [−.01, .01], and the reciprocal path was not significant, $b = .01$, $\beta = <.01$, $p = .621$, 95% CI [−.02, .03].

Regarding the control variables, subjective social status predicted health status at T2, $b = .07$, $\beta = .10$, $p < .001$, 95% CI [.06, .08], and conservatism, $b = .03$, $\beta = .03$, $p = .002$, 95% CI [.01, .05]; gender predicted health status, $b = .04$, $\beta = .02$, $p = .025$, 95% CI [.01, .08], but not conservatism, $b = -.05$, $\beta = -.01$, $p = .111$, 95% CI [−.11, .01]; and age predicted conservatism, $b = <.01$, $\beta = .02$, $p = .038$, 95% CI [<.01, .01], but not health status, $b = <.01$, $\beta = -.01$, $p = .473$, 95% CI [<.01, <.01]. When including income as an additional covariate, the model showed an appropriate goodness of fit, χ^2 (60) = 1,224.75, $p < .001$, CFI = .974, TLI = .962, RMSEA = .052, and income did not predict health status, $b = .02$, $\beta = <.02$, $p = .147$, 95% CI [−.01, .04], or conservatism, $b = <.01$, $\beta = <.01$, $p = .926$, 95% CI [−.04, .03]. More importantly, in this model conservatism did not predict health status over time, $b = <.01$, $\beta = <.01$, $p = .528$, 95% CI [−.01, .01], and health status did not predict conservatism over time either, $b = <.01$, $\beta = <.01$, $p = .793$, 95% CI [−.03, .03].

Supplementary analyses. Based on these results, we ran three supplementary analyses. First, we compared the countries in the dataset using multigroup structural equation modeling. In these analyses, we controlled for multiple comparisons (i.e., 19 countries) using Bonferroni correction (i.e., $\alpha/19 = .05/19 = .003$; see OSM Tables S19-S30). Second, we conducted the main analyses using only the item of political conservatism, which is consistent with previous studies (e.g., Napier & Just, 2008; Schlenker et al., 2012; see OSM Tables S31-S46). And third, we conducted cross-sectional results using both waves to compare our results with those previously shown in the literature (e.g., Napier & Jost, 2008; Stavrova & Luhmann, 2016; see OSM Tables S47-S76). Overall, we found little support for the hypothesis that conservatism would predict well-being over time across countries. We only found the following significant associations (i.e., $p \leq .003$): Life satisfaction predicted conservatism in Estonia (both with and without covariates), anxiety negatively predicted conservatism in Estonia (only without covariates), depression negatively predicted conservatism in Estonia (only without covariates), conservatism positively predicted depression in Korea (only with covariates), and health status predicted conservatism in Estonia (both without and with covariates). In addition, the main results were similar when using a single-item measure of conservatism. Finally, cross-sectional results were consistent with previous literature when not including control variables. Conservatism predicted greater life satisfaction and health status, lower depression, and did not predict anxiety. However, conservatism did not predict life satisfaction or health status after including the covariates (i.e., social status, gender, age, and income) and positively predicted anxiety and depression when including the covariates.

In summary, Study 1 showed no support for our main hypothesis. Contrary to the hypothesis, conservatism did not positively predict well-being over time. However, the lack of additional measures over time (i.e., three or more waves) prevented us from conducting more sophisticated analyses. This is relevant because CLPMs do not disentangle temporal stability and individual-level stability and also assume the absence of trait stability (Hamaker et al., 2015). Therefore, we tested our hypotheses in Study 2 using a four-wave longitudinal design in a representative sample from Chile through Random Intercept Cross Lagged Panel Models (RI-CLPM; Hamaker et al., 2015). We selected this country because the left-right continuum is associated with acceptance of inequality and resistance to change—with rightists scoring higher in both dimensions—in a similar way than people from other countries where the association between conservatism and well-being has been tested (e.g., US; Solano Silva, 2018).

Study 2

Method

Sample. We used data from the Chilean Social Longitudinal Study in Study 2. This is a longitudinal survey conducted yearly since 2016 among a nationally representative sample of Chilean adults. Participants were selected using a multistage probabilistic sampling design and the surveys were conducted by trained interviewers at participants' homes with a time lag of approximately 1 year (Centro de Estudios de Conflicto y Cohesión Social, 2018). The research team conducting this survey (Centro de Estudios de Conflicto y Cohesión Social, 2020) created a variable that identified the existence of inconsistencies in demographic information across waves for each participant (e.g., changes in age higher than 2 years when comparing two consecutive waves). This variable was created as a proxy to identify whether a case in the dataset was indeed the same person that responded to all the waves of this survey. From those participants that completed the first wave (N = 2,928), we selected those without any inconsistencies, leading to a sample size of 2,554 participants. The mean age at T1 was 45.57 (SD = 15.28) and 60.18% of participants were female.

Measures. In Study 2, we included all measures available in the dataset to test our hypotheses.

Conservatism. *We used a single-item measure of conservatism: Traditionally in our country, people define his or her political positions as being closer to the left, to the center, or to the right. Using a scale from 0 to 10 where 0 is to be from the "left," 5 is to be from the "center," and 10 is to be from the "right," where do you place yourself on this scale?*

Life satisfaction. We used a single-item measure of life satisfaction whose answers ranged from 1 (*totally dissatisfied*) to 5 (*totally satisfied*).

Depression. We used a Spanish version of the Patient Health Questionnaire-9 (PHQ-9; Kroenke et al., 2001) to measure depression. Participants were asked to report the frequency of several depressive symptoms during the last 2 weeks using a scale ranging from 1 (*never*) to 7 (*every day*). This scale was highly reliable at T1 (α = .85), T2 (α = .88), T3 (α = .88), and T4 (α = .86). Given the large number of items, to handle missing data in this scale, we computed a depression score for each participant with the items available for them (Newman, 2014).

Health status. We took an item from DeSalvo et al. (2006) to assess health status. Participants rated their own health status (*How would you say your health is?*) using a scale ranging from 1 (*poor*) to 5 (*excellent*).

Control variables. We included age and gender (0 = *male*, 1 = *female*) as 2 time-invariant variables predicting well-being scores at each wave.[3] We also included subjective social status as a longitudinal control variable (0 = *low status*, 10 = *high status*; adapted from Adler et al., 2000). We used a measure of monthly household income measured at each wave of the survey. Participants were asked to indicate their monthly household net income in Chilean pesos (CLP). If they refused to respond to this question, they were asked to indicate their household income in ranges, ranging from 1 (*<270,000*) to 20 (*>2,700,000*).[4] To use data from all participants that indicated their income in any of these two questions, we recoded raw income into ranges and used this variable in Study 2. As in Study 1, we included this control variable in different models.

Results and Discussion

Descriptive statistics and correlations matrix are presented in Table 2. Overall, these correlations show that conservatism was not significantly related to our measures of well-being at the bivariate level at each wave.

We tested our hypotheses using RI-CLPM (Hamaker et al., 2015). We used these models to examine the longitudinal association between conservatism and well-being while considering individual-level stability in these constructs. These models decompose the scores of the longitudinal variables in the grand means (i.e., means across all participants in the same wave), between-individuals components (i.e., random intercepts that represent time-invariant deviations from the grand mean), and within-individuals components (i.e., differences for each participant between their scores at each wave and their expected scores computed through the grand mean and random intercept; Mulder & Hamaker, 2021). Then, two main results are of relevance to test our hypotheses. First, the association between the random intercepts of conservatism and well-being (i.e., between-individuals components) indicates whether overall participants high in conservatism also score high in well-being. And second, the cross-lagged associations (within-individuals level) indicate whether the individuals' deviations from their expected means in one of the constructs are associated with subsequent individual's deviations from their expected means in the other construct (i.e., whether increments in conservatism at the individual level are associated with increments in well-being at the individual level). We constrained the autoregressive paths and the longitudinal associations to be equal across waves because we did not have specific hypotheses regarding differences in these coefficients (Orth et al., 2021). For this reason, we also report here unstandardized coefficients for our main results. As in Study 1, missing values were treated using full information maximum likelihood, and all models were conducted through Mplus v.6.12 (Muthen & Muthen, 2012). Missing data at the construct level for all the variables included in the analyses were higher in later than in earlier waves (see OSM Table S178). Indeed, 62.5% of participants responded to all four waves (see OSM Table S177). This led to a percentage of partial respondents above 80% for all the main analyses. This percentage is mainly driven by missing data in conservatism (45.7% - 56.3%). For this reason, we tested whether, for participants identifying in the left/liberal-right/conservative continuum, this variable was associated with proxies of two general orientations

Table 2. Descriptive Statistics and Correlations Matrix (Study 2).

Variable	*M*	*SD*	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16	17	18	19	20	21	22	23	24	25	26
1. Conservatism (t1)	4.69	2.48	1																									
2. Conservatism (t2)	4.84	2.84	.60	1																								
3. Conservatism (t3)	5.01	2.53	.52	.64	1																							
4. Conservatism (t4)	4.60	2.44	.53	.55	.60	1																						
5. Life satisfaction (t1)	3.73	0.88	.06	.02	.03	.01	1																					
6. Life satisfaction (t2)	3.78	0.86	−.02	−.02	.04	.02	.27	1																				
7. Life satisfaction (t3)	3.87	0.84	.04	−.05	.04	.09	.24	.25	1																			
8. Life satisfaction (t4)	3.77	0.85	.01	.03	.04	.05	.26	.27	.29	1																		
9. Depression (t1)	1.67	0.64	.00	−.03	−.05	−.05	−.27	−.18	−.19	−.21	1																	
10. Depression (t2)	1.66	0.66	.00	.03	−.05	−.05	−.16	−.22	−.19	−.21	.40	1																
11. Depression (t3)	1.76	0.67	−.02	.00	−.07	−.07	−.18	−.17	−.24	−.20	.42	.44	1															
12. Depression (t4)	1.80	0.65	.05	.00	.01	−.05	−.18	−.19	−.19	−.30	.34	.31	.40	1														
13. Health status (t1)	2.69	0.89	−.03	−.04	.01	.01	.21	.16	.14	.14	−.34	−.25	−.23	−.24	1													
14. Health status (t2)	2.68	0.88	.00	−.02	.00	.04	.15	.23	.15	.18	−.28	−.35	−.27	−.22	.46	1												
15. Health status (t3)	2.63	0.88	−.01	−.06	.00	.03	.17	.16	.24	.17	−.26	−.30	−.33	−.25	.45	.49	1											
16. Health status (t4)	2.59	0.84	−.02	.00	.03	.03	.12	.16	.16	.23	−.23	−.27	−.23	−.32	.47	.49	.53	1										
17. Social status (t1)	4.32	1.56	.16	.11	.08	.09	.22	.17	.08	.12	−.14	−.09	−.05	−.05	.17	.16	.10	.10	1									
18. Social status (t2)	4.52	1.62	.10	.05	.02	.02	.12	.19	.09	.10	−.12	−.07	−.03	−.04	.12	.16	.07	.06	.34	1								
19. Social status (t3)	4.43	1.53	.12	.08	.10	.09	.13	.18	.12	.14	−.10	−.08	−.10	−.09	.15	.16	.14	.11	.31	.38	1							
20. Social status (t4)	4.23	1.56	.11	.12	.02	.16	.13	.15	.13	.18	−.10	−.05	−.08	−.13	.13	.13	.13	.14	.32	.33	.38	1						
21. Income (t1)	7.04	5.45	.07	.03	−.03	−.06	.19	.17	.14	.12	−.13	−.06	−.09	−.08	.25	.24	.23	.25	.33	.23	.24	.27	1					
22. Income (t2)	8.58	5.40	.01	.06	.01	−.02	.16	.17	.16	.10	−.11	−.06	−.03	−.06	.20	.22	.20	.20	.26	.26	.28	.27	.68	1				
23. Income (t3)	7.57	5.62	.04	.01	−.02	−.02	.16	.19	.19	.14	−.14	−.10	−.10	−.10	.26	.28	.28	.28	.30	.24	.29	.30	.69	.69	1			
24. Income (t4)	7.93	5.76	.03	−.01	−.05	−.06	.17	.22	.16	.18	−.14	−.12	−.10	−.12	.27	.29	.29	.29	.29	.26	.27	.32	.64	.68	.74	1		
25. Age	45.57	15.28	.05	.08	.10	.11	.03	−.03	.04	<.01	.02	.02	−.07	−.04	−.30	−.29	−.27	−.29	−.07	−.07	−.07	−.08	−.20	−.15	−.29	−.31	1	
26. Gender (1 female)	0.60		.03	.02	<.01	.02	<.01	−.01	−.04	−.05	.17	.18	.21	.16	−.16	−.18	−.19	−.18	−.04	−.02	−.05	−.04	−.17	−.12	−.21	−.20	.07	1

Note. All correlations $|r| \geq .05$ were significant at $\alpha = .05$ (i.e., $p < .05$).

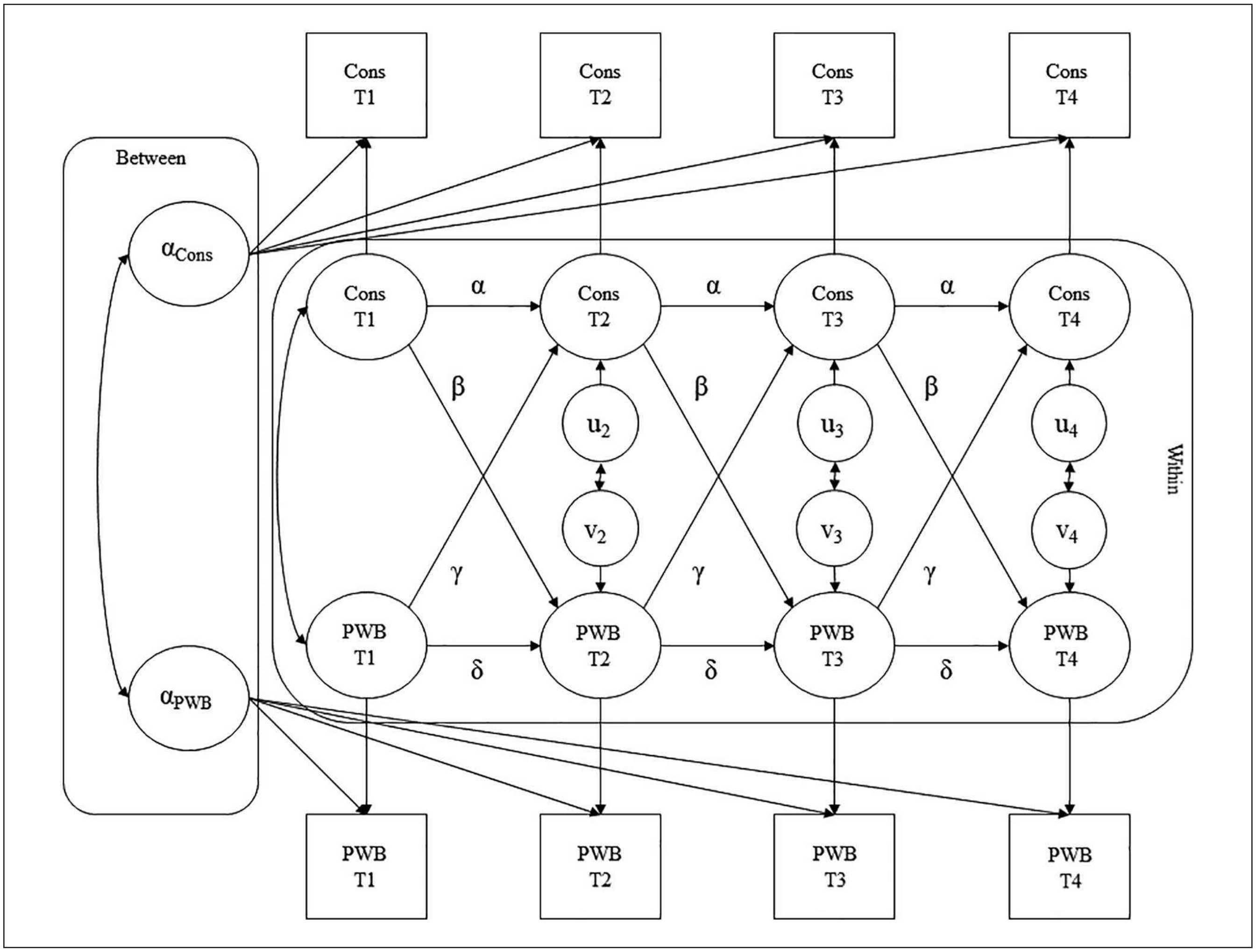

Figure 2. Random-intercept cross-lagged panel models.

Note. Variables were conservatism (Cons) and psychological well-being (PWB; i.e., life satisfaction, depression, or health status). Autoregressive paths (i.e., α and δ) and longitudinal associations (i.e., β and γ) were constrained to be equal across time (e.g., $\alpha = \alpha_1 = \alpha_2 = \alpha_3$; $\beta = \beta_1 = \beta_2 = \beta_3$). Random intercepts are indicated by α_{Cons} (conservatism) and α_{PWB} (psychological well-being). Residuals are indicated by u and v.

commonly related to conservatism such as social dominance orientation and right-wing authoritarianism. Overall, we found that at all waves, conservatism was positively related to both variables (*r*s from .07 to .30, *p*s $< .05$; see OSM Table S200). We also compared missing data in conservatism at all waves regarding subjective well-being measures at T1. We only found that participants without missing data (vs. with missing data) in conservatism at T1 scored higher in health status (T1), participants without missing data in conservatism at T2 scored higher in life satisfaction (T1), and participants without missing data in conservatism at T3 scored higher in depression (T1). No other comparisons were significant (see OSM Table S179). The conceptual model tested in Study 2 is presented in Figure 2.[5] In this section, we report only information relevant to our theoretical discussion. All the relevant parameters of our models can be found in the OSM (Tables S81-S96).

Life satisfaction. The model for life satisfaction without covariates showed an appropriate goodness of fit, χ^2 (17) = 26.45, $p = .067$, CFI = .995, TLI = .992, RMSEA = .015. At the within-individuals level, conservatism predicted life satisfaction over time, but the coefficient was negative, $b = -.02$, $p = .026$, 95% CI [−.04, <.01], and the reciprocal path did not reach conventional levels of significance, $b = .10$, $p = .098$, 95% CI [−.02, .22]. Unexpectedly, the autoregressive paths explaining life satisfaction were not significant, $b = .02$, $p = .461$, 95% CI [−.03, .05]. We did not find evidence that participants high in conservatism also scored higher in life satisfaction (i.e., between-level), given that the covariance between both random intercepts was not significant, $b = .06$, $p = .096$, 95% CI [−.01, .13]. When we included control variables, we also found an appropriate goodness of fit, χ^2 (59) = 113.91, $p < .001$, CFI = .983, TLI = .975, RMSEA = .019. In this model, conservatism also negatively predicted life satisfaction over time, $b = -.02$, $p = .027$, 95% CI [−.04, <.01], but life satisfaction did not predict conservatism over time, $b = .11$, $p = .075$, 95% CI [−.01, .23].

Regarding the control variables, gender did not predict life satisfaction, $b = -.02$, $p = .375$, 95% CI [−.07, .03], or conservatism, $b = .13$, $p = .189$, 95% CI [−.06, .32]; age predicted both life satisfaction, $b = <.01$, $p = .030$, 95% CI [<.01, <.01], and conservatism, $b = .02$, $p < .001$, 95% CI [.01, .02]; and social status did not predict life satisfaction, $b = .01$, $p = .489$, 95% CI [−.01, .03], or conservatism over time, $b = -.07$, $p = .081$, 95% CI [−.14, .01]. Importantly, we found that at the between-individuals level, social status was associated with conservatism, $b = .44$, $p < .001$, 95% CI [.14, .20], and life satisfaction, $b = .21$, $p < .001$, 95% CI [.30, .58]. In other words, participants high in social status scored higher in conservatism and life satisfaction, but individual deviations from participants' expected means in social status did not predict individual deviations from expected means in conservatism or life satisfaction. When including income as an additional longitudinal control variable, we observed an appropriate goodness of fit, χ^2 (98) = 482.93, $p < .001$, CFI = .954, TLI = .929, RMSEA = .039, and the results were similar than in the previous model: conservatism negatively predicted life satisfaction over time, $b = -.03$, $p = .016$, 95% CI [−.05, −.01], but the reciprocal path was not significant, $b = .11$, $p = .083$, 95% CI [−.01, .22]. Income did not predict life satisfaction, $b = <.01$, $p = .471$, 95% CI [−.01, .01], or conservatism over time, $b = <.01$, $p = .965$, 95% CI [−.03, .04].[6]

Depression. The model for depression without covariates showed an appropriate goodness of fit, χ^2 (17) = 41.09, $p < .001$, CFI = .991, TLI = .985, RMSEA = .024. At the within-individuals level, conservatism did not predict depression over time, $b = .01$, $p = .519$, 95% CI [−.01, .02], and depression did not predict conservatism over time, $b = -.11$, $p = .245$, 95% CI [−.29, .08]. At the between-individuals level, we found that conservatism was not associated with depression, $b = -.03$, $p = .365$, 95% CI [−.08, .03]. When we included control variables, we also found an appropriate goodness of fit, χ^2 (59) = 129.48, $p < .001$, CFI = .982, TLI = .973, RMSEA = .022. In this model, conservatism did not predict depression over time, $b = <.01$, $p = .712$, 95% CI [−.01, .02], and the reciprocal path was not significant either, $b = -.14$, $p = .127$, 95% CI [−.32, −.04].

Regarding the control variables, gender predicted depression, $b = .23$, $p < .001$, 95% CI [.20, .27], but not conservatism, $b = .13$, $p = .172$, 95% CI [−.06, .32]; age predicted conservatism, $b = .02$, $p < .001$, 95% CI [.01, .02], but not depression, $b = <.01$, $p = .091$, 95% CI [<.01, <.01]; and social status did not predict depression, $b = <.01$, $p = .695$, 95% CI [−.02, .01], or conservatism, $b = -.07$, $p = .080$, 95% CI [−.14, .01]. However, at the between-individuals level, social status was associated with conservatism, $b = .44$, $p < .001$, 95% CI [.30, .58], and depression, $b = -.07$, $p < .001$, 95% CI [−.09, −.04]. When including income as an additional longitudinal control variable, we observed an appropriate goodness of fit, χ^2 (98) = 491.10, $p < .001$, CFI = .956, TLI = .932, RMSEA = .040, and the results were similar than in the previous model: conservatism did not predict depression, $b = <.01$, $p = .607$, 95% CI [−.01, .02], and depression did not predict conservatism, $b = -.13$, $p = .146$, 95% CI [−.31, .05]. Income predicted depression, $b = .01$, $p = .012$, 95% CI [<.01, .02], but not conservatism over time, $b = <.01$, $p = .909$, 95% CI [−.04, .03].[7]

Health status. Finally, the model for health status without covariates showed an appropriate goodness of fit, χ^2 (17) = 21.41, $p = .208$, CFI = .999, TLI = .998, RMSEA = .010. At the within-individuals level, conservatism did not predict health status over time, $b = -.01$, $p = .136$, 95% CI [−.03, <.01], and health status did not predict conservatism over time, $b = -.09$, $p = .242$, 95% CI [−.23, .06]. At the between-individuals level, we found that conservatism was not associated with health status, $b = .02$, $p = .607$, 95% CI [−.06, .10]. When we included control variables, we also found an appropriate goodness of fit, χ^2 (59) = 105.06, $p < .001$, CFI = .991, TLI = .986, RMSEA = .017. In this model, conservatism did not predict health status, $b = -.01$, $p = .155$, 95% CI [−.03, .01], and health status did not predict conservatism over time, $b = -.06$, $p = .378$, 95% CI [−.21,.08]. Regarding the control variables, gender predicted health status, $b = -.26$, $p < .001$, 95% CI [−.31, −.21], but not conservatism, $b = .14$, $p = .161$, 95% CI [−.05, .31]; age predicted conservatism, $b = .02$, $p < .001$, 95% CI [.01, .02], and health status, $b = -.02$, $p < .001$, 95% CI [−.02, −.01]; and social status did not predict health status, $b = .01$, $p = .465$, 95% CI [−.01, .03], or conservatism, $b = -.06$, $p = .096$, 95% CI [−.14, .01]. However, at the between-individuals level, social status was associated with conservatism, $b = .45$, $p < .001$, 95% CI [.31, .59], and health status, $b = .12$, $p < .001$, 95% CI [.09, .04]. When including income as an additional longitudinal control variable, we observed an appropriate goodness of fit, χ^2 (98) = 491.51, $p < .001$, CFI = .964, TLI = .944, RMSEA = .038. Results were similar than in the previous model: conservatism did not predict health status, $b = -.01$, $p = .133$, 95% CI [−.03, <.01], and health status did not predict conservatism, $b = -.07$, $p = .368$, 95% CI [−.21, .08]. Income did not predict health status, $b = <.01$, $p = .877$, 95% CI [−.01, .01], or conservatism over time, $b = <.01$, $p = .927$, 95% CI [−.03, .04].[8]

Supplementary analyses. We ran three sets of supplementary analyses. First, the RI-CLPM has been criticized for using a different form of causal inference (i.e., individual deviations from expected means over time) than other models such as the CLPM (i.e., group differences over time; Ludtke & Robitzsch, 2021; Orth et al., 2021). For this reason, we compared the RI-CLPM with models constraining the variances and covariances between the random intercepts to zero, which is equivalent to a CLPM (Hamaker et al., 2015). We found all these comparisons as significant (all $p_s < .001$; see OSM Tables S138-S140), with CLPMs with the goodness of fit below the thresholds suggested by the literature (see OSM Tables S108-S137). These models showed that life satisfaction *positively* predicted conservatism over time and depression negatively predicted conservatism over time, both when including and not including relevant covariates. Second, we freely estimated the cross-lagged coefficients from T3 to T4 (i.e., not constrained as equal versus the rest of the cross-lagged coefficients) because the last wave of the survey (i.e., T4) was collected after the beginning of a massive leftist social movement in Chile, which might confound our results (see OSM Tables S99-S107). Results from these models showed that life satisfaction positively predicted conservatism over time from T3 to T4 (both when not including covariates and including covariates) and from T to T+1 (i.e., T1 → T2 and T2 → T3) only when not including covariates. The reciprocal path from T3 to T4 was not significant in any of our models, and from T to T+1 was negative and significant in all models. For the models with depression, none of the cross-lagged coefficients were significant. For the models with health status, conservatism negatively predicted health status from T to T+1 (but not from T3 to T4) only when including covariates.[9] And third, conducted cross-sectional analyses at all waves. We found that conservatism positively predicted psychological well-being only at T1 (life satisfaction only without covariates) and T3 (depression without and with covariates). None of the rest of the results were consistent with our hypotheses (see OSM Tables S141-S176).

In summary, and consistent with Study 1, Study 2 showed no significant cross-lagged correlations where conservatism predicted a higher level of well-being using three indicators in a four-wave longitudinal dataset. Specifically, we did not find associations between conservatism and depression, and health status over time. Furthermore, conservatism predicted life satisfaction, but this association was negative—and not positive, as we had predicted. Nevertheless, when using the same analytical approach as in Study 1 (i.e., CLPM), we found that life satisfaction longitudinally predicted more conservatism, and depression predicted less conservatism, which is not consistent with the order we had predicted (i.e., conservatism did not longitudinally predict higher subjective psychological well-being).

General Discussion

Previous research has shown that conservatives exhibit higher levels of psychological well-being than liberals (e.g., Napier & Jost, 2008; Schlenker et al., 2012; Subramanian & Perkins, 2010). Different theoretical accounts have explained such association assuming that conservatism antecedes well-being (e.g., Jetten et al., 2013; Jost & Hunyady, 2002; Schlenker et al., 2012). However, this literature has failed to provide evidence supporting this theoretical assumption. Indeed, researchers have argued that psychological well-being could antecede conservatism because people high in well-being might seek to protect their positive status (Subramanian et al., 2009) or they could be less attentive to injustice cues (Napier et al., 2020).

In this research, we sought to contribute to filling this gap in the literature using two longitudinal datasets. Importantly, our analyses captured multiple dimensions of subjective well-being, including life satisfaction, anxiety, depression, and health status. Contrary to previous findings (e.g., Napier & Jost, 2008; Schlenker et al., 2012; Subramanian & Perkins, 2010), our results showed that conservatism did not positively predict well-being over time. Most of these associations were nonsignificant, especially when considering anxiety, depression, and health status. Those associations that reached conventional levels of significance indicated that life satisfaction predicted conservatism (Study 1) and that conservatism *negatively* predicted life satisfaction (Study 2). Importantly, these results differed from those obtained when using the first waves in both studies. Cross-sectionally, conservatism predicted greater life satisfaction, lower depression, and higher health status in Study 1 and not in Study 2—but only when not including covariates. These results are different from those in the extant literature that relies on large datasets such as the World Values Survey, which includes a wider array of countries and higher sample sizes, but also includes fewer and less elaborate indicators of psychological well-being (e.g., Napier & Jost, 2008; Stavrova & Luhmann, 2016).

Another contribution of our research is that the analyses showed different results when using different analytical approaches to test longitudinal hypotheses, at least when treating life satisfaction and depression as the operationalization of well-being. For instance, in Study 1, using CLPMs (Raykov & Marcoulides, 2006), we found that life satisfaction positively predicted conservatism. These results were consistent with Study 2 when using the same analytical approach, as shown in the supplementary analyses. However, when taking in consideration individual-level stability through RI-CLPMs (Hamaker et al., 2015), we found better goodness-of-fit statistics, and the results showed that conservatism negatively predicted life satisfaction. In other words, both studies showed similar results when using similar analytical approaches for life satisfaction, but they were different when we used a different test of longitudinal hypotheses in Study 2. Also, in Study 2, depression did not predict conservatism over time when using RI-CLPM, but it did show a negative longitudinal association when using CLPM. More importantly, none of these results were consistent with the idea that conservatism positively predicts well-being over time. Some of these associations might support the idea that people high in well-being seek to protect their positive status

(Subramanian et al., 2009) or are less attentive to injustice cues (Napier et al., 2020). Nevertheless, these results were found only in some of the analyses we conducted and should be cautiously interpreted.

The results from our research are relevant because they provide important caveats to claims about psychological differences between liberals and conservatives. Theory and research on this issue have argued that liberals and conservatives differ in terms of their psychological profiles (i.e., personality traits; e.g., Jost et al., 2003; Schlenker et al., 2012) and/or psychological motivations (i.e., system justification; Jost & Hunyady, 2002). These differences are hypothesized to underlie further differences, such as cultural consumption (Rogers, 2020). In the specific case of well-being, research has assumed that those psychological differences lead to enhanced levels of well-being among conservatives (e.g., Napier & Jost, 2008; Schlenker et al., 2012). In this research, however, we showed that despite the plausibility of the theory stating that liberals and conservatives have different psychological profiles, conservatism does not predict well-being over time. The longitudinal results reported here are consistent with studies arguing that the association between conservatism and well-being can be explained through the influence of third variables such as socioeconomic status (Jetten et al., 2013). Indeed, the few longitudinal associations we found between conservatism and subjective well-being consistent with our hypotheses became nonsignificant when controlling for subjective social status.

A related field of research has analyzed the specific case of system justification as a predictor of well-being in both longitudinal (Vargas-Salfate et al., 2018) and experimental research (Li et al., 2020). However, system justification (i.e., support for the status quo) only represents one component of the broader ideology of conservatism. Future research should be dedicated to disentangling conservatism from system justification and account for different meanings ascribed to conservatism and its association with well-being across cultures.

Limitations and Future Research Directions

Our research has caveats that need to be acknowledged. We used a restricted sample of countries comprising mainly Western and developed countries, excluding samples from Africa. This low number of countries prevented us from developing and testing further hypotheses at the country level (e.g., Stavrova & Luhmann, 2016). For instance, it would be of interest to test whether some country-level variables such as economic wealth, economic inequality, or the degree of aggregated conservatism are directly related to well-being or interact with conservatism at the individual level when predicting psychological well-being (Stavrova & Luhmann, 2016). Furthermore, we do not have direct evidence that conservatism was understood in the same form across all the countries included in our two studies. We only had indirect evidence in Study 1 based on measurement invariance analyses and in Study 2 by associating conservatism with social dominance orientation and right-wing authoritarianism. In addition, brief measures of the constructs examined were administered, mainly because of the cost associated with administering surveys across countries (Study 1) and large representative questionnaires at participants' homes (Study 2). Taken as together, these three limitations imply constraints on the generalizability of our research that need to be addressed in future studies.

Our research also has several limitations related to longitudinal models. In Study 1, we used CLPMs, which have been criticized for not accounting for trait stability (Hamaker et al., 2015). In Study 2, we could account for such stability through RI-CLPMs. However, researchers have raised concerns about the underlying conceptual differences between these models when testing longitudinal associations (Orth et al., 2021). CLMPs focus on between-person differences, and RI-CLPMs focus on within-person differences: These model different causal processes/pathways. In our research, CLPMs allowed us to test whether conservatives would show higher levels of well-being when compared with liberals, while RI-CLPMs allowed us to test whether increments in conservatism at the individual level would be associated with increments in well-being. However, extant theory on the association between conservatism and well-being (e.g., Jost & Hunyady, 2002; Schlenker et al., 2012; Stavrova & Luhmann, 2016) does not provide strong theoretical arguments to disentangle which process might be occurring when analyzing this association—as it has also been suggested for other psychological subfields (Orth et al., 2021). In that sense, future research should provide more detailed arguments to disentangle the specific causal processes leading (or not leading) conservatism to be associated with psychological well-being.

In addition, we used arbitrary time lags of 6 months (Study 1) and 1 year (Study 2). We relied on secondary data in both studies; therefore, we could not vary these lags between waves. This is an important limitation given that different time lags or intervals can have consequences for the hypothesis testing of longitudinal data through CLPMs and RI-CLPMs (Kuiper & Ryan, 2018). However, this also reflects a broader concern in the literature on the association between conservatism and well-being. None of the explanations in the literature (e.g., Jetten et al., 2013; Jost et al., 2003; Schlenker et al., 2012) provide theoretical insights to propose an appropriate time lag to test this association in a longitudinal design. Future research should consider providing more accurate descriptions of these associations that could allow researchers to propose specific time intervals. We doubt there is a canonical lag time that would be ideal across cultures, but rather suspect that historical events occurring between measurement points (e.g., 9-11 happening in the United States, the GFC around the world in 2008) may affect the relationship between conservatism and well-being in specific cultural contexts (for theory on the impact of historical events on psychological relationships, see Liu & Khan, 2021).

Finally, in this research, we equated conservatism with ideological self-placement in the left-right continuum. This is a common practice in psychological research (e.g., Federico & Malka, 2018), but there is no strong evidence showing this equivalence. We suspect this might be a relevant issue for future research given that the nonresponse for the left-right self-placement was higher in Chile than in other contexts where researchers have argued that not all individuals self-identify as liberals or conservatives (e.g., Kalmoe, 2020). Despite this high nonresponse rate, we found that among people identifying in the left-right continuum, this ideological variable was positively related to social dominance orientation and right-wing authoritarianism. This confirms previous studies showing that in Chile, an important number of people do not identify as leftist or rightist, but among those that self-identify along this continuum, there are differences in key ideological variables (Solano Silva, 2018). Importantly, this limitation should also be extended to the analytic procedure to handle missing data. We used full-information maximum likelihood, which requires a lower percentage of partial respondents than those observed in Study 2. For this reason, all the results from Study 2 should be confirmed by future research. Although our study had all these limitations, this is one of the few longitudinal tests of the association between conservatism and well-being and shows evidence contradicting previous studies.

Conclusion

Research has found a small positive association between conservatism and subjective well-being. Most of the theoretical accounts have proposed a causal association such that conservatism enhances well-being, through the endorsement of ideological beliefs or differences in personality traits. In this research, we used longitudinal data to test the hypothesis that conservatism antecedes well-being. In two studies across 20 countries, and using different methods to treat the longitudinal data, we did not find evidence consistent with this hypothesis when considering 6-month and 1-year time lags. We look forward to seeing future research attempting to replicate these results using more measures of conservatism and subjective well-being as well as different time intervals or lags.

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Open Practices

This study was not preregistered. All materials (database for Study 1, scripts, and measures) are publicly available at https://osf.io/z53cq

Database for Study 2 is publicly available at https://doi.org/10.7910/DVN/SOQJ0N

Supplemental Material

Supplemental material is available online with this article.

Notes

1. A list of all publications based on this dataset can be found at https://www.dropbox.com/s/oko40j9uzzh1i8j/Digital%20Influence%20World%20Project%20Research%20Output.docx?dl=0
2. We also estimated a series of models without cross-lagged coefficients to measure the stability of the constructs. These models showed coefficients of $\beta = .72$ for conservatism and β_s ranging from .71 to .78 for our four dependent variables (see OSM Tables S3-S6).
3. These two time-invariant variables function as control variables in the CLPMs but not in the RI-CLPMs (Hamaker et al., 2015; Mulder & Hamaker, 2021). We decided to include them in the RI-CLPMs to keep comparability with the results from the CLPMs.
4. 1 CLP = .0012 USD (October 6, 2021). Median job income was 401,000 CLP (October 2021) (https://www.ine.cl/prensa/2020/10/26/ingreso-laboral-promedio-mensual-en-chile-fue-de-$620.528-en-2019).
5. As in Study 1, we also estimated a series of models without cross-lagged coefficients to measure the stability of the constructs. Results showed that conservatism was stable over time at the individual level (b_s ranging from .23 to .24, all $p_s < .001$), as well as depression ($b = .09$, $p < .001$) and health status ($b = .05$, $p = .025$). However, life satisfaction at T was not associated with life satisfaction at T+1 ($b = .02$, $p = .430$), which indicates that individual's deviations from their expected means in life satisfaction do not predict subsequent deviations in the same variable (see OSM Tables S78-S80).
6. Further support for the differences between the significant path from life satisfaction to conservatism and the reciprocal path can be found when comparing these models with models constraining the cross-lagged coefficients as equal. Results from these comparisons suggest significant differences ($p_s \leq .032$) but below the thresholds suggested by the literature for CFI (>.01; Milfont & Fischer, 2010) and RMSEA (>.015; Putnick & Bornstein, 2016). These results indicate that the path from life satisfaction to conservatism was stronger than the path from conservatism to life satisfaction, but this difference is rather small (see OSM Tables S84, S90, S193, and S138).
7. When comparing these models with those models constraining the cross-lagged coefficients as equal, we found nonsignificant

differences (all $p_s > .100$). This suggest that there might not be a reciprocal association between depression and conservatism (see OSM Tables S85, S91, S97, and S139).

8. When comparing these models with those models constraining the cross-lagged coefficients as equal, we found nonsignificant differences (all $p_s \geq .299$). This suggest that there might not be a reciprocal association between health status and conservatism (see OSM Tables S86, S92, S98, and S140).
9. We also conducted exploratory analyses by unconstraining all parameters in the models (see OSM Tables S180-S198). Overall, when not including covariates, both the RI-CLPMs and CLPMs did not show significantly different goodness of fit than the models only unconstraining the paths from T3 to T4 (see OSM Tables S189 and S199). For the models that included covariates, there were significant differences but none of the results were consistent with the idea that conservatism predicted life satisfaction (or vice versa) from T1 to T2 and T2 to T3.

References

Adler, N. E., Epel, E. S., Castellazzo, G., & Ickovics, J. R. (2000). Relationship of subjective and objective social status with psychological and physiological functioning: Preliminary data in healthy, White women. *Health Psychology*, *19*(6), 586–592. https://doi.org/10.1037/0278-6133.19.6.586

Biesanz, J. (2012). Autoregressive longitudinal models. In R. H. Hoyle (Ed.), *Handbook of structural equation modeling* (pp. 459–471). Guilford Press.

Briki, W., & Dagot, L. (2020). Conservatives are happier than liberals: The mediating role of perceived goal progress and flow experience—A pilot study. *Current Psychology*, *41*, 1267–1278. https://doi.org/10.1007/s12144-020-00652-0

Butz, S., Kieslich, P., & Bless, H. (2017). Why are conservatives happier than liberals? Comparing different explanations based on system justification, multiple group membership, and positive adjustment. *European Journal of Social Psychology*, *47*(3), 362–372. https://doi.org/10.1002/ejsp.2283

Centro de Estudios de Conflicto y Cohesión Social. (2018). *Manual de Usuario de Estudio Longitudinal Social de Chile, Primera Ola (2016). Chilean Social Longitudinal Study User Manual, First Wave (2016). Cross-sectional Corte Transversal* (Versión 2.0).

Centro de Estudios de Conflicto y Cohesión Social. (2020). *Estudio Longitudinal Social de Chile, Versión Panel Combinada 2016-2019* (ELSOC_Wide_2016_2019_v1.00) [Archivo de datos]. Chilean Social Longitudinal Study, Merged Longitudinal Version 2016-2019. www.coes.cl

DeSalvo, K. B., Fisher, W. P., Tran, K., Bloser, N., Merrill, W., & Peabody, J. (2006). Assessing measurement properties of two single-item general health measures. *Quality of Life Research*, *15*(2), 191–201. https://doi.org/10.1007/s11136-005-0887-2

Diener, E. (2009). Subjective well-being. In E. Diener (Ed.), *The science of well-being* (pp. 11–58). Springer.

Federico, C., & Malka, A. (2018). The contingent, contextual nature of the relationship between needs for security and certainty and political preferences: Evidence and implications. *Advances in Political Psychology*, *39*(1), 3–48. https://doi.org/10.1111/pops.12477

Gil de Zúñiga, H., & Liu, J. (2017). Second screening politics in the social media sphere: Advancing research on dual screen use in political communication with evidence from 20 countries. *Journal of Broadcasting & Electronic Media*, *61*(2), 193–219. https://doi.org/10.108/08838151.2017.1309420

Godfrey, E., Santos, C., & Burson, E. (2017). For better or worse? System-justifying beliefs in sixth-grade predict trajectories of self-esteem and behavior across early adolescence. *Child Development*, *90*(1), 180–195. https://doi.org/10.1111/cdev.12854

Hamaker, E. L., Kuiper, R. M., & Grasman, R. P. (2015). A critique of the cross-lagged panel model. *Psychological Methods*, *20*(1), 102–116. https://doi.org/10.1037/a0038889

Harding, J., & Sibley, C. (2013). The palliative function of system justification: Concurrent benefits versus longer-term costs to wellbeing. *Social Indicators Research*, *113*(1), 401–418. https://doi.org/10.1007/s11205-012-0101-1

Heck, R. (2009). Multilevel modeling with SEM. In G. A. Marcoulides & R. E. Schumacker (Eds.), *New developments and techniques in structural equation modeling* (pp. 89–127). Lawrence Erlbaum.

Hu, L., & Bentler, P. (1999). Cutoff criteria for fit indexes in covariance structure analysis: Conventional criteria versus new alternatives. *Structural Equation Modeling: A Multidisciplinary Journal*, *6*(1), 1–55. https://doi.org/10.1080/10705519909540118

Jetten, J., Haslam, S., & Barlow, F. (2013). Bringing back the system: One reason why conservatives are happier than liberals is that higher socioeconomic status gives them access to more group memberships. *Social Psychological and Personality Science*, *4*(1), 6–13. https://doi.org/10.1177/1948550612439721

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., & Hunyady, O. (2002). The psychology of system justification and the palliative function of ideology. *European Review of Social Psychology*, *13*(1), 111–153. https://doi.org/10.1080/10463280240000046

Kalmoe, N. (2020). Uses and abuses of ideology in political psychology. *Political Psychology*, *41*(4), 771–793. https://doi.org/10.1111/pops.12650

Kroenke, K., Spitzer, R. L., & Williams, J. B. (2001). The PHQ-9. Validity of a brief depression severity measure. *Journal of General Internal Medicine*, *16*(9), 606–613. https://doi.org/10.1046/j.1525-1497.2001.016009606.x

Kuiper, R., & Ryan, O. (2018). Drawing conclusions from cross-lagged relationships: Re-considering the role of the time-interval. *Structural Equation Modeling: A Multidisciplinary Journal*, *25*(5), 809–823. https://doi.org/10.1080/10705511.2018.1431046

Lau, A., Cummins, R., & McPherson, W. (2005). An investigation into the cross-cultural equivalence of the personal well-being index. *Social Indicators Research*, *72*(3), 403–430. https://doi.org/10.1007/s11205-004-0561-z

Li, W., Wu, J., & Kou, Y. (2020). System justification enhances life satisfaction of high- and low-status people in China. *Social Psychological and Personality Science*, *11*(5), 588–596. https://doi.org/10.1177/1948550619866182

Liu, J., & Khan, S. (2021). Implications of a psychological approach to collective remembering: Social representations as cultural ground for interpreting survey and experimental results. *Journal of Pacific Rim Psychology*, *15*, 1–11. https://doi.org/10.1177/18344909211007938

Löwe, B., Wahl, I., Rose, M., Spitzer, C., Glaesmer, H., Wingenfeld, K., Schneider, A., & Brähler, E. (2010). A 4-item measure of depression and anxiety; Validation and standardization of the Patient Health Questionnaire-4 (PHQ-4) in the general population. *Journal of Affective Disorders*, *122*(1–2), 86–95. https://doi.org/10.1016/j.jad.2009.06.019

Ludtke, O., & Robitzsch, A. (2021). A critique of the random intercept cross-lagged panel model. *Preprint*. https://doi.org/10.31234/osf.io/6f85c

Milfont, T., & Fischer, R. (2010). Testing measurement invariance across groups: Applications in cross-cultural research. *International Journal of Psychological Research*, *3*(1), 111–121.

Mulder, J., & Hamaker, E. (2021). Three extensions of the random intercept cross-lagged panel model. *Structural Equation Modeling: A Multidisciplinary Journal*, *28*(4), 638–348. https://doi.org/10.1080/10705511.2020.1784739

Muthen, L., & Muthen, B. (2012). *Mplus user's guid* (7th ed.).

Napier, J. L., Bettinsoli, M., & Suppes, A. (2020). The palliative function of system-justifying ideologies. *Current Opinion in Behavioral Sciences*, *34*, 129–134. https://doi.org/10.1016/j.cobeha.2020.03.002

Napier, J. L., & Jost, J. T. (2008). Why are conservatives happier than liberals? *Psychological Science*, *19*(6), 565–572. https://doi.org/10.1111/j.1467-9280.2008.02124.x

Newman, D. (2014). Missing data: Five practical guidelines. *Organizational Research Methods*, *17*(4), 372–411. https://doi.org/10.1177/1094428114548590

Orth, U., Clark, D. A., Donnellan, M. B., & Robins, R. W. (2021). Testing prospective effects in longitudinal research: Comparing seven competing cross-lagged models. *Journal of Personality and Social Psychology*, *120*(4), 1013–1034. http://doi.org/10.1037/pspp0000358

Pratto, F., Stallworth, L. M., & Sidanius, J. (1997). The gender gap: Differences in political attitudes and social dominance orientation. *The British Journal of Social Psychology*, *36*(Pt. 1), 49–68. https://doi.org/10.1111/j.2044-8309.1997.tb01118.x

Putnick, D. L., & Bornstein, M. H. (2016). Measurement invariance conventions and reporting: The state of the art and future directions for psychological research. *Developmental Review: DR*, *41*, 71–90. https://doi.org/10.1016/j.dr.2016.06.004

Raykov, T., & Marcoulides, G. (2006). *A first course in structural equation modeling*. Lawrence Erlbaum.

Rogers, N. (2020). Split screens: A content analysis of American liberals' and conservatives' respective television favorites. *Psychology of Popular Media*, *9*(1), 45–58. https://doi.org/10.1037/ppm0000197

Schlenker, B., Chambers, J., & Le, B. (2012). Conservatives are happier than liberals, but why? Political ideology, personality, and life satisfaction. *Journal of Research in Personality*, *46*(2), 127–146. https://doi.org/10.1016/j.jrp.2011.12.009

Solano Silva, D. (2018). Conservatism and political orientation, Is their relationship similar in Latin-America and Western societies? *Psicoperspectivas*, *17*(1). https://doi.org/10.5027/psicoperspectivas-vol17-issue1-fulltext-1098

Spitzer, R., Kroenke, K., Williams, J., & Löwe, B. (2006). A brief measure for assessing generalized anxiety disorder: The GAD-7. *Archives of Internal Medicine*, *166*(10), 1092–1097. https://doi.org/10.1001/archinte.166.10.1092

Stavrova, O., & Luhmann, M. (2016). Are conservatives happier than liberals? Not always and not everywhere. *Journal of Research in Personality*, *63*, 29–35. https://doi.org/10.1016/j.jrp.2016.04.011

Subramanian, S. V., Huijts, T., & Perkins, J. M. (2009). Association between political ideology and health in Europe. *European Journal of Public Health*, *19*(5), 455–457. https://doi.org/10.1093/eurpub/ckp077

Subramanian, S. V., & Perkins, J. M. (2010) . Are republicans healthier than democrats? *International Journal of Epidemiology*, *39*(3), 930–931. https://doi.org/10.1093/ije/dyp152

Vargas-Salfate, S., Paez, D., Khan, S., Liu, J., & Gil de Zúñiga, H. (2018). System justification enhances well-being: A longitudinal analysis of system justification in 18 countries. *British Journal of Social Psychology*, *57*(3), 567–590. https://doi.org/10.1111/bjso.12254