The partisan mind: Is extreme political partisanship related to cognitive inflexibility?

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Abstract:

The rise of partisan animosity, ideological polarization, and political dogmatism has reignited important questions about the relationship between psychological rigidity and political partisanship. Two competing hypotheses have been proposed: one hypothesis argues that mental rigidity is related to a conservative political orientation, and the other suggests that it reflects partisan extremity across the political spectrum. In a sample of over 700 US citizens, partisan extremity was related to lower levels of cognitive flexibility, regardless of political orientation, across three independent cognitive assessments of cognitive flexibility. This was evident across multiple statistical analyses, including quadratic regressions, Bayes factor analysis, and interrupted regressions. These findings suggest that the rigidity with which individuals process and respond to non-political information may be related to the extremity of their partisan identities.

Keywords:

Main Text:

In *The True Believer* (1951), the thinker Eric Hoffer wrote about crowds and mass movements, "*All movements, however different in doctrine and aspiration, draw their early adherents from the same types of humanity; they all appeal to the same types of mind.*" Hoffer tapped into an important idea: there may be a certain "*type of mind*" that is particularly drawn to adopting an ideology or doctrine, almost regardless of its content or ambition. Indeed, scholars and scientists have long sought to identify the psychological underpinnings of rigid ideological adherence, particularly since the atrocities committed in the name of political ideologies at the start of the 20th century.

Many psychologists have theorized that *rigid adherence to a political ideology* may reflect underlying *cognitive rigidity* (Adorno et al., 1950; Rokeach, 1954). Nonetheless, there have been two competing hypotheses regarding the precise relationship between cognitive rigidity and political identity. One hypothesis, the *ideological extremity hypothesis*, has suggested that individuals on the political extremes – of both liberal and conservative ideologies – are more cognitively rigid than moderates (Greenberg & Jonas, 2003). This hypothesis emerges from the notion that extreme political partisanship may be linked to inflexible belief systems that capture the world in black-and-white terms and offer certainty and simplicity (Fernbach, Rogers, Fox, & Sloman, 2013; van Prooijen & Krouwel, 2017; see also Crawford, 2019). Indirect evidence has illustrated that political extremists on both the political left and right are more dogmatically intolerant (van Prooijen & Krouwel, 2017), more likely to feel superior about their beliefs (Toner, Leary, Asher, & Jongman-Sereno, 2013) and to avoid exposure to each other's opinions (Frimer, Skitka, & Motyl, 2017) relative to political moderates. Consequently, the *rigidity-of-the-extreme* hypothesis suggests that political extremism is underpinned by a cognitive rigidity that facilitates the attitudinal rigidity that often characterizes political extremists.

The second hypothesis, the so-called *rigidity-of-the-right hypothesis*, has argued that conservatives perceive the world in a more inflexible, categorical way than liberals (Altemeyer, 1998; Jost, Glaser, Kruglanski, & Sulloway, 2003a). This hypothesis has received empirical support through studies demonstrating a relationship between political conservatism and psychological preferences for traditionalism, familiarity, and certainty, and between political liberalism and acceptance of uncertainty and ambiguity (for reviews: Hibbing, Smith, & Alford, 2014; Jost, 2017; Van Hiel, Onraet, & De Pauw, 2010).

Notably, empirical research on this debate has almost exclusively relied on self-report questionnaires as proxies for cognitive rigidity, as opposed to the use of objective cognitive assessments to quantify cognitive inflexibility directly. Self-report measures of cognitive style tend to suffer from considerable limitations, such as failure in self-assessment of cognitive abilities (Kruger & Dunning, 1999), inclusion of political content in non-political measures and non-political content in political measures (Malka, Lelkes, & Holzer, 2017), and poor psychometric properties (Kirton, 1981). Additionally, a recent meta-analysis has shown that correlations between cognitive style and political attitudes are inflated when relying on self-report questionnaires to quantify cognitive rigidity rather than behavioural assessments (Van Hiel, Onraet, Crowson, & Roets, 2016). Furthermore, most measures of "cognitive and perceptual rigidity" used thus far in political psychology have not corresponded to those used in the cognitive psychology literature. Consequently, there is a need to test these hypotheses using established objective and ideologically-neutral measures of cognitive rigidity.

Defining Cognitive Flexibility and Rigidity

In the neuropsychological literature, cognitive flexibility is defined as the ability to adapt to novel or changing environments and a capacity to switch between modes of thinking (Cools & Robbins, 2004). Specifically, it can be defined as "the ability to flexibly switch

perspectives, focus of attention, or response mappings" (Diamond, 2006, p. 70). Cognitive inflexibility is therefore represented by perseveration, "the tendency of an individual *not* to change" (Schultz & Searleman, 2002, p. 166), and maladaptive continuation of unrewarded behaviours and thought patterns.

There is significant individual variation in cognitive flexibility within the general population (Braver, Cole, & Yarkoni, 2010; Miyake & Friedman, 2012), which has been linked to dopaminergic functioning (Barnett, Jones, Robbins, & Müller, 2007). Furthermore, from a clinical perspective, some populations manifest a deficit in cognitive flexibility by persisting with previously-established rules or behavioural patterns even when this is maladaptive, as evident in patients of obsessive-compulsive disorder (Chamberlain et al., 2006), addiction (Verdejo-Garcia, Pérez-García, & Bechara, 2006), and frontal lobe damage (Anderson, Damasio, Jones, & Tranel, 1991).

As reviewed by Ionescu (2012), there are several behavioural tasks that are classically used to operationalize cognitive flexibility in adults, including the Wisconsin Card Sorting Test (Grant & Berg, 1948), task switching and optional shift paradigms (Miyake & Friedman, 2012; Monsell, 2003), the Alternative Uses Task (AUT; Guilford, 1967), insight problems, and induction tasks (Shafto, Coley, & Vitkin, 2007). In these tasks, performance is measured via participants' accuracy rates, reaction times, or the number and variety of provided responses to open-ended problems. Given the extensive use and validation of behavioural tasks for assessing cognitive flexibility in neuropsychology and cognitive science, the present study will rely on three validated and independent behavioural measures of cognitive flexibility in order to investigate the relationship between inflexibility and political ideology.

Rigidity-of-the-Extreme: The Psychology of Extreme Partisanship

The two hypotheses – the *ideological extremity hypothesis* and the *rigidity-of-the-right hypothesis* – tap into two orthogonal aspects of political identity. As noted by Settle, Dawes, and Fowler (2009), "Partisanship is typically evaluated along two dimensions – the strength of reported partisan attachment and the direction of that attachment" (p. 601). While a rich literature has emerged on the psychology of partisan *direction* (i.e. left versus right political orientation), there has been little empirical investigation of the cognitive origins of partisan *intensity*, that is, the strength of a person's partisan identity and attachment (Van Bavel & Pereira, 2018).

Nonetheless, new findings have demonstrated several ways in which political extremists, on both sides of the political spectrum, differ psychologically from political moderates. Relative to moderates, political extremists tend to experience more negative emotions about politics and to derogate outgroups (van Prooijen, Krouwel, Boiten, & Eendebak, 2015) and view politics in more simplistic terms (Lammers, Koch, Conway, & Brandt, 2017). Political extremists' tendency to believe in simple political solutions to complex societal challenges also predicts their greater likelihood of believing in conspiracy theories (van Prooijen, Krouwel, & Pollet, 2015). Furthermore, research with cognitive anchoring tasks has suggested that political extremists exhibit greater belief superiority and are more likely to reject external information than moderates (Brandt, Evans, & Crawford, 2015). In the case of the EU refugee crisis, political extremists possessed greater judgmental certainty about their knowledge of the crisis, independently of their actual knowledge, than political moderates (van Prooijen, Krouwel, & Emmer, 2017). This demonstrates that both partisan direction and intensity matter for how individuals evaluate ideological arguments and intergroup conflict.

Recently, Jost (2017) pointed out several key questions currently faced by political psychology, including: (1) *What does political psychology have to say about left-wing rigidity?* and (2) *Are ideological differences only evident in subjective, self-report measures and not behaviourally?* With regards to the question of left-wing rigidity, while recent evidence has shown that left-wing authoritarianism exists and is predictive of voting behaviour (Conway III, Houck, Gornick, & Repke, 2018; Conway III & McFarland, 2019), there is a lack of research directly examining the relationship between partisan intensity and cognitive rigidity independently of partisan direction. When the rigidity-of-the-right hypothesis and the ideological extremity hypothesis have been pitted against each other, the psychological variables of interest have consisted of intolerance of ambiguity, need for cognitive closure, dogmatism, and integrative complexity (Jost, Glaser, Kruglanski, & Sulloway, 2003b); none have assessed cognitive rigidity directly and objectively. This is related to the second question posed by Jost (2017) about the contrast between self-report and behavioural methods – perhaps the use of behavioural methodologies has the potential to illuminate ideological differences (and similarities) which are absent or obscured in self-report questionnaires.

Rigidity-of-the-Right: Evidence and Measurement Problems

Two recent meta-analyses have evaluated the state of the evidence in favour of the rigidity-of-the-right hypothesis. Jost (2017) identified 16 studies which investigated whether conservatives score more highly on tests of "perceptual or cognitive rigidity" than liberals. Nine out of 16 studies supported the rigidity-of-the-right hypothesis, such that the overall unweighted (r=.32) and weighted (r=.38) effect sizes were large and significant. Notably, however, out of the nine studies demonstrating a significant effect, six studies used self-report rather than behavioural measures: five studies used a self-report measure of cognitive rigidity (Gough & Sanford, 1952), and one study used a self-report measure of intolerance of trait inconsistency (Steiner & Johnson, 1963). Two studies used unvalidated behavioural measures

which do not clearly tap into cognitive flexibility: Kidd & Kidd's (1972) study asked participants to identify changes between visual objects drawn on cards. However, recognizing changes or discrepancies between visual stimuli does not imply adaptability to change or flexibility of thought, and the author-designed task was not validated or used as a measure of cognitive flexibility in later studies. The second study by Neuringer (1964) used the "Rokeach Map Test" as a test of rigidity, in which participants were presented with sequential street maps and asked to find the shortest distance between two points on the map. After being presented with seven maps in which the shortest route was identical across maps, participants were shown five maps in which a new shortcut was possible. If participants chose the shortcut on any of the five test maps, they were characterized as "nonrigid". This study's small number of test trials (i.e. five), low threshold for being classified as "nonrigid" (scoring above 0 out of 5), small sample (N=15), along with the problematic nature of binary assessments of continuous cognitive constructs, should raise serious doubts regarding the validity of Neuringer's (1964) findings. Consequently, the significant findings from this meta-analysis of "cognitive and perceptual rigidity" were based largely on self-report measures and a couple of problematic behavioural tests.

The second meta-analysis, by Van Hiel and colleagues (2016), identified a larger number of studies in which a measure of rigidity was administered (N=46), but separated the studies according to whether rigidity was operationalized with a behavioural task (N=31) or a self-report scale (N=15). Overall, the findings corroborate Jost's (2017) meta-analysis, as conservatism was significantly related to self-reported rigidity (r=.47) and, to a lesser extent, behaviourally-assessed rigidity (r=.11). Even in this expanded sample of behavioural studies, there is significant variation in how cognitive rigidity was operationalized: for instance, while Rokeach (1954) utilized a set of perceptual problem solving tasks (i.e. the Einstellung problems; which have received criticisms: Levitt, 1956), Sidanius (1985) used the "Political

Prediction Test" in which participants are asked to make judgements within the political domain and cognitive flexibility is extracted as an index from how these judgments are made. Consequently, some behavioural measures of cognitive rigidity possess ideological content and so are not inherently politically-neutral.

These two meta-analyses clearly demonstrate that no behavioural study has been conducted for over 20 years on disentangling the rigidity-of-the-right and ideological extremity hypotheses (since 1997; see Van Hiel et al., 2016), and the publication year for a behavioural study up to 1997 is on average 1959 (range: 1948 to 1997). Moreover, the debate has been plagued by relatively small sample sizes (mean=65, range=15 to 225; calculated from Van Hiel et al., 2016) and diverse methodologies and conceptualizations of both conservatism and rigidity. The present study therefore sought to rely on established, ideologically-neutral cognitive tasks from the neuropsychological literature and recruit a large sample that would be adequately powered to detect these effects.

Materials and Methods
Participants and Procedure

We sought to recruit 648 participants via Amazon's Mechanical Turk (www.mturk.com) to achieve greater than 90% power to detect a small effect of Cohen's f=.1 at α=.005 in our primary one-way ANCOVAs of the cognitive variables and three groups of political affiliation and conservatism. We anticipated a small effect due to previous work done by Sidanius (1985) and reviewed by Jost and colleagues (2003), which found an average correlation of .15 between general conservatism and conceptually similar cognitive flexibility constructs. We oversampled by 15% and recruited 750 participants, 7 of which were excluded due to providing incomplete responses, yielding a total sample of 743 participants. Data collection was not continued after data analysis. All participants were American citizens. Participants were redirected from Amazon's Mechanical Turk to an online survey hosted by Qualtrics Survey Software for completion of all the self-reported items and the RAT, and later redirected again to Inquisit 5 by Millisecond Software in order to temporarily download software that allows for accurate measure of performance and reaction times in the WCST. Participants were asked about demographic variables such as age, gender, and educational attainment (ranging from high school graduate, some college but no degree, associate degree (2-year), bachelor's degree (4-year), Master's degree, professional degree (JD, MD), or doctoral degree). The research was conducted with the ethical approval of the institution's Department of Psychology Ethics Committee.

Analysis was conducted using the R packages *BayesFactor* (Morey et al., 2015)*, ggstatsplot* (Patil & Powell, 2018), *visreg* (Breheny & Burchett, 2013), and *jtools* (Long, 2019).

Measures

Political Identity Measures

Political Party Affiliation

To indicate their political party affiliation, participants were asked: "Generally speaking, do you usually think of yourself as a Republican, a Democrat, an Independent, or something else?", with the response options being "Republican", "Democrat", "Independent", "No preference", or the option to indicate another affiliation.

Political Partisanship

To measure participants' feeling of attachment to their political party, participants were presented with a validated measure of identity fusion, the Dynamic Identity Fusion Index (DIFI; Jimenez et al., 2016), which consists of a continuous pictorial representation that allows participants to move a small circle representing "the self" by clicking and dragging it toward or away from a large circle representing "the group/ideology". The amount of overlap between the two circles has been shown to indicate the extent to which individuals feel their personal identity is fused with a collective identity (e.g. Jimenez et al., 2016; Jong, Whitehouse, Kavanagh, & Lane, 2015; Kapitány, Kavanagh, Buhrmester, Newson, & Whitehouse, 2019; Misch, Fergusson, & Dunham, 2018). It has temporal stability, as well as convergent and discriminant validity (Jimenez et al., 2016). In this study, participants were presented with the DIFI twice; once where the group was the "Republican Party" and another when the group was "Democratic Party". Hence, to compute Republican Party Identity Fusion, the size of the overlapping area between the "Me" circle and the "Republican Party" circle was calculated. The same methodology was applied to the calculation of Democratic Party Identity Fusion.

To calculate political partisanship, we computed each participant's maximum identity fusion to either political party (Democratic/Republican Party), and retrieved the value for the party

with which the participant was most fused. To capture the left-right spectrum, if participants' maximum fusion was with the Democratic Party, their fusion score was multiplied by -1, to create a spectrum from -100 to 100. This was taken as an index of participants' political partisanship, weighted by partisan direction.

$$PP_{max} = \begin{cases} -Fusion_D & if\ Fusion_D > Fusion_R \\ Fusion_R & otherwise \end{cases}$$

Where $Fusion_R$ reflects Republican Party Identity Fusion and $Fusion_D$ reflects Democratic Party Identity Fusion.

As a robustness check, we also computed the difference in participants' fusion to the two parties:

$$PP_{diff} = Fusion_R - Fusion_D$$

Furthermore, in order to obtain an unweighted measure of political partisanship, for cases when it is necessary to examine participants' fusion with their favoured party regardless of the party's political direction, we collapsed PP_{max} into:

$$PP_{|max|} = \begin{cases} Fusion_D & if\ Fusion_D > Fusion_R \\ Fusion_R & otherwise \end{cases}$$

Consequently, the $PP_{|max|}$ scale reflects political party identity fusion between 0 and 100, rather than between -100 and 100 as in PP_{max}.

Political Conservatism

The 12 Item Social and Economic Conservatism Scale (SECS; Everett, 2013) was administered, which asks participants about how positively or negatively they feel towards 12 social and economic political issues on a scale of 0-100 (with increments of 10). As detailed by Everett (2013), the social issues consisted of abortion, religion, the family unit, traditional marriage, traditional values, patriotism, and military and national security. The economic issues included welfare benefits, limited government, business, gun ownership, and fiscal

responsibility. By summing the participant's score in response to social issues and separately in response to economic issues, we could compute a score of issue-specific social conservatism and economic conservatism respectively. The reliability of the SECS was high (Cronbach's alpha for overall scale = .883, social conservatism subscale = .872, economic conservatism subscale = .736).

Cognitive Flexibility Tests

Remote Associates Test (RAT)

The Remote Associates Test (RAT; Mednick, 1962) consisted of 15 compound remote associate problems, in which participants are presented with three cue words (e.g. *cottage, swiss, and cake*), and must generate the compound word solution that connects these three words (e.g. *cheese*). Items of varying difficulty levels were selected from established remote associate problems (Bowden & Jung-Beeman, 2003). Participants were given 20 seconds to respond to each item. RAT performance was computed in terms of the proportion of correct words.

Wisconsin Card Sorting Test (WCST)

The WCST (Grant & Berg, 1948) was administered with Inquisit 5 by Millisecond Software in standard fashion (Heaton, 1981). Participants are presented with four key cards and a deck of response cards that vary on three dimensions (colour, shape, and number of geometric figures) and are asked to match a fifth card from the sequentially-presented response cards to one of the four key cards. Participants need to identify the correct classification rule (out of three potential rules: matching by colour, shape, or number) according to the feedback they receive after each trial. They are informed that the classification rule may change without warning, and indeed the rule alternates after participants correctly respond to ten consecutive trials, requiring a flexible set shift. The task ends after participants complete six categories (twice for each of the

three rules) or after 128 trials. To index participants' performance, the number of categories completed and the accuracy rate were computed.

Alternative Uses Test (AUT)

In the AUT (Guilford, 1967), participants are asked to generate as many possible uses for two common household items (brick and newspaper) for 2 min. Participants' responses were recorded and scored along four components by two independent raters in accordance with previous guidelines (Cronbach's $\alpha = 0.994$; Chermahini & Hommel, 2010; Madore, Addis, & Schacter, 2015; Roberts et al., 2017). *Flexibility* was scored according to the number of distinct categories that participants' responses for a given item could be clustered into (e.g. using a newspaper for making origami and making paper dolls are uses that would fall under the same category of arts and crafts, while using a brick for swatting a fly would fall under a separate category). The total flexibility score comprised the sum from all trials. *Fluency* constituted the total number of appropriate responses. *Elaboration* reflected the amount of detail provided by the participants (for brick, "build" would receive a score of 0; "build a house" would receive a score of 1; and "a weapon to protect family when your house is robbed" would be awarded 2 points for specifying detailed use and context). To score *originality*, each response was compared to the responses from the rest of the participants, such that responses to a given object that were only provided by 5% of the sample received an originality point. The total originality score reflected the sum of original scores per participant across all trials. In accordance with convention, nonsensical uses were excluded prior to coding of responses. To establish inter-rater reliability for appropriate categories, level of detail for the elaboration scoring, and originality, the raters separately scored 25 random participants' responses, and once high inter-rater reliability was established with this set on all AUT measures (Cronbach's $\alpha > 0.91$ on all measures); the raters independently scored the rest of the participants. Each AUT measure reflects the mean score given by the two independent raters.

Results:

The sample consisted of 743 participants (Age: *M*=36.58, *SD*=13.46; 55.6% female), including 323 self-affiliated Democrats, 161 Republicans, and 203 Independents (56 participants with responded as "Other" and so were excluded for the analyses of political affiliation but were included in all other analyses). We first validated known associations between political affiliation, political partisanship, and conservatism. Univariate ANOVAs on participants' conservatism with respect to their party affiliation demonstrated significant differences between Democrats, Independents, and Republicans in social conservatism, $F(2,596)=111.01$, $p=1.046\times10^{-41}$, $\eta_p^2=.271$, and economic conservatism, $F(2,621)=169.268$, $p=2.113\times10^{-59}$, $\eta_p^2=.353$. After post-hoc Bonferroni correction, across both conservatism measures, Democrats (social conservatism: *M*=338.85, *SD*=135.65; economic conservatism: *M*=225.73, *SD*=67.17) were significantly more liberal than Independents (*p*s<.001; social conservatism: *M*=392.64, *SD*=148.57; economic conservatism: *M*=284.56, *SD*=83.49), and Independents were significantly more liberal than Conservatives (*p*s<.001; social conservatism: *M*=540.52, *SD*=117.09; economic conservatism: *M*=361.88, *SD*=76.12).

Democrats were no more fused to their favoured party – i.e. were no more partisan – than Republicans, $t(451)=-.237$, $p=.813$ (see Fig 1). Participants who identified as Independents were significantly less partisan (to the party with which they felt greater identity fusion) than either Democrats or Republicans, $F(2,617)=38.582$, $p<.001$, $\eta_p^2=.111$.

Pearson's correlations among the cognitive flexibility variables were modest, replicating past research: AUT Flexibility was positively related to RAT ($r=.20$, $p<.001$) and WCST ($r=.19$, $p<.001$), and the correlation between RAT and WCST approached significance ($r=.10$, $p=.066$). There was manifest individual variability in each of the cognitive flexibility measures

(RAT: M=68.02%, SD=23.85%, range=0 to 100%; WCST: M=71.61%, SD=15.31%, range=20.31%; AUT Flexibility: M=3.745, SD=1.636, range=0 to 9).

In terms of the demographic variables, there were no gender differences in performance on any of the cognitive flexibility measures: RAT (t(639)=1.346, p=.176), WCST (t(396)=-.172, p=.863), and AUT Flexibility (t(716)=.416, p=.678). There was a significant gender difference in absolute political partisanship ($PP_{|max|}$) (t(636)=2.410, p=.016), with women (M=56.09, SD=31.99) exhibiting heightened identity fusion to their favoured political party relative to men (M=49.94, SD=31.95). With regards to conservatism, there was no significant gender difference in social conservatism (t(622)=.112, p=.911) but there was a significant difference in economic conservatism (t(659)=-3.777, p<.001), with men (M=293.24, SD=91.80) reporting heightened economic conservatism relative to women (M=266.13, SD=90.69). Educational attainment was not significantly related to any of the variables of interest (ps>.124). Furthermore, age was positively related to RAT performance (r=.136, p=.001), but not WCST (r=-.042, p=.403) or AUT Flexibility (r=-.021, p=.573). Age was also positively related to social conservatism (r=.217, p<.001) and economic conservatism (r=.083, p=.034) but not to absolute political partisanship ($PP_{|max|}$, r=-.003, p=.943). However, there was no significant age difference between Republicans (M=38.32, SD=12.831), Democrats (M=35.69, SD=14.059), and Independents (M=36.23, SD=12.913; F(2,676)=2.076, p=.126), nor a significant difference in educational attainment (F(2,682)=.884, p=.414; Republicans: M=2.24, SD=1.373; Democrats: M=2.41, SD=1.289; Independents: M=2.38, SD=1.261). Consequently, and in accordance with convention, the variables of age, educational attainment, and gender were consistently included as covariates in all statistical analyses unless otherwise specified.

Fig 1. Political partisanship according to political affiliation. Democrats and Republicans were equally partisan in their identity fusion with their favoured political party ($PP_{|max|}$).
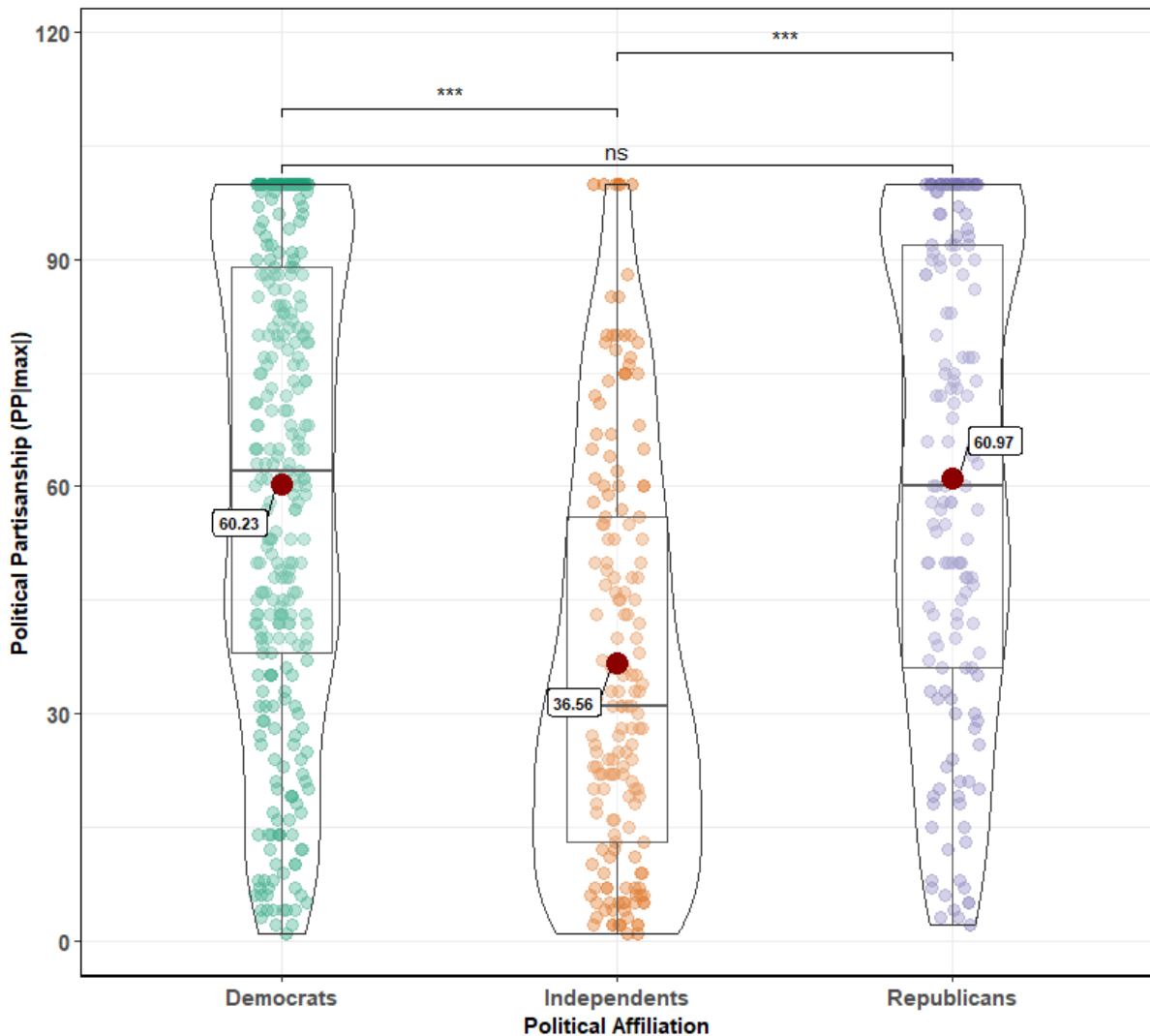
Political Affiliation and Cognitive Flexibility

A univariate ANCOVA on RAT performance according to self-reported political affiliation (Democrats, Republicans, versus Independents) was computed, with age, gender, and educational attainment as covariates. This revealed a main effect of political affiliation, $F(2,595)=5.345, p=.005, \eta_p^2=.018$, such that Independents performed significantly better on the RAT ($M=72.38\%$, SD$=22.26\%$) than Democrats ($M=66.67\%$, $SD=24.19\%$, Independent-Democrat estimated marginal mean difference$=.065$, $p=.016$, 95% CI = [.009, .121]) and

Republicans (M=65.16%, SD=25.38%, Independent-Republican estimated marginal mean difference=.078, p=.010, 95% CI = [.014, .142]) after post-hoc Bonferroni correction. There was no significant difference between Democrats and Republicans. There were no effects of gender or educational attainment, but there was a significant effect of age, $F(1,595)$=10.131, p=.002, η_p^2=.017 (Fig 2).

Similarly, a univariate ANCOVA on WCST performance with age, gender, and educational attainment as covariates demonstrated a main effect of political affiliation, $F(2,363)$=5.794, p=.003, η_p^2=.031, whereby Independents performed significantly better on the WCST (M=75.53%, SD=11.83%) than Democrats (M=71.27%, SD=15.67%, Independent-Democrat estimated marginal mean difference=.045, p=.047, 95% CI = [.000, .089]) and Republicans (M=68.28%, SD=17.06%, Independent-Republican estimated marginal mean difference=.072, p=.003, 95% CI = [.019, .125]) after post-hoc Bonferroni correction. There was no effect of age, gender, or educational attainment. There were no significant differences between Democrats and Republicans (Fig 2).

Finally, a univariate ANCOVA on the AUT Flexibility score, with age, gender, and educational attainment as covariates revealed a main effect of political affiliation, $F(2,664)$=5.872, p=.003, η_p^2=.017 (see Fig 2), such that Democrats (M=3.903, SD=1.533) and Independents (M=3.987, SD=1.481) performed significantly better than Republicans (M=3.456, SD=1.637), after post-hoc Bonferroni correction (Republican-Democrat estimated marginal mean difference = -.443, p=.014, 95% CI = [-.797, -.068]; Republican-Independent estimated marginal mean difference = -.535, p=.004, 95% CI = [-.932, -.137]). There was no difference in AUT Flexibility between Democrats and Independents, and there was no effect of age, gender, or educational attainment.
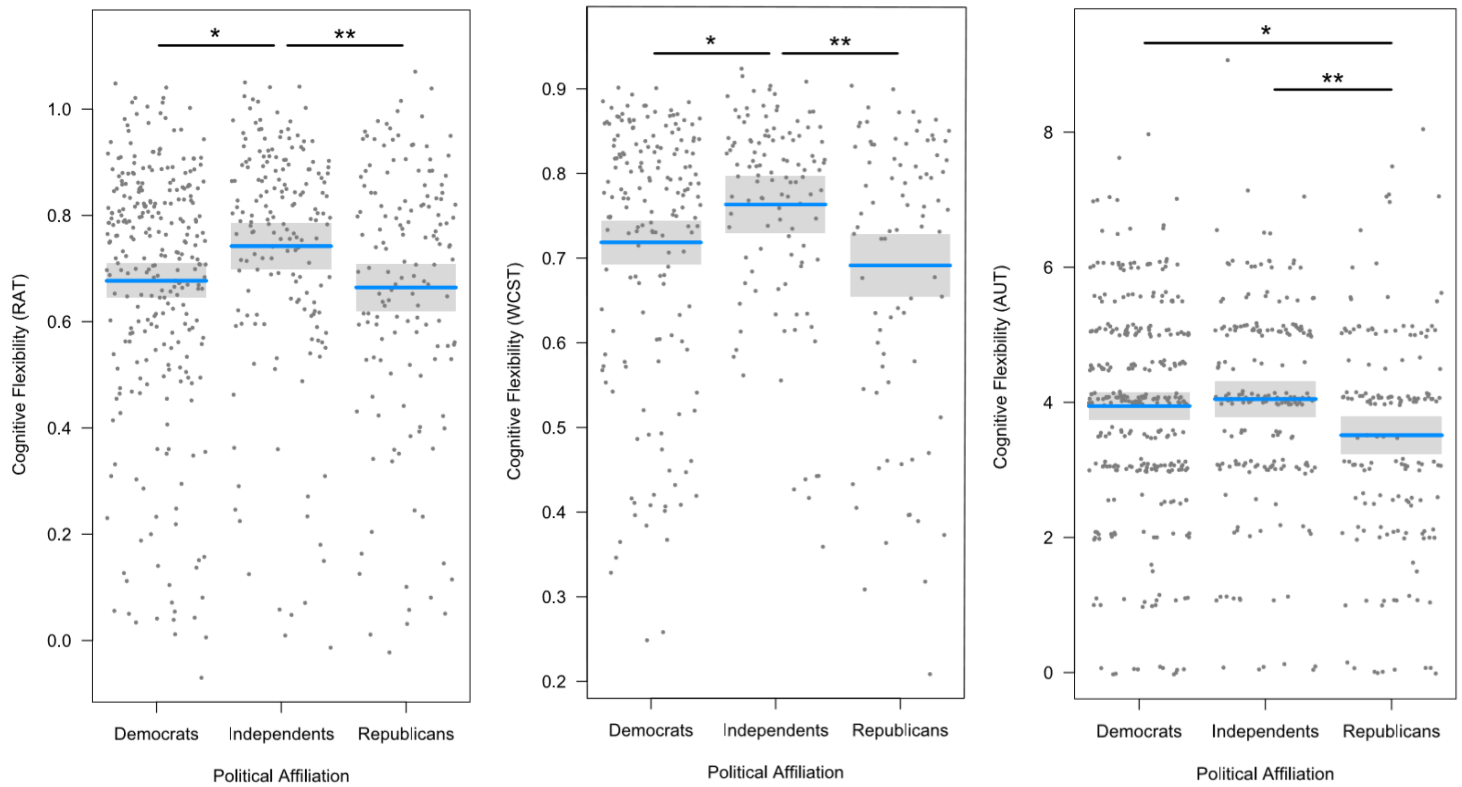
Fig 2. Cognitive flexibility according to political affiliation for the three cognitive flexibility tests: Remote Associates Task, Wisconsin Card Sorting Test, and Alternative Uses Test Flexibility score. The shaded areas reflect 95% confidence intervals. Comparisons indicate significant differences after Bonferroni correction, accounting for age, gender, and educational attainment. Data points reflect partial residuals. *p<.05, **p<.01.

Partisanship and Cognitive Flexibility

Bayesian Quadratic Regressions

Given the inverted-U-shaped relationship evident in Fig 2, we tested for quadratic associations between cognitive flexibility and political partisanship across the political spectrum (PP_{max}). For each cognitive flexibility measure, we conducted two regression models: one that assumed a linear relationship and another that assumed a quadratic relationship between cognitive flexibility and political partisanship. We computed Bayes factors for the linear and quadratic

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regressions to compare the strength of the evidence for a quadratic (subscript q) over a linear (subscript l) model specification (BF_{ql}).

The Bayes factors in this case express the relative likelihood of a quadratic model versus a linear model given the data and certain prior expectations. To calculate Bayes factors using Bayesian regression, we relied on a default Bayesian approach promoted by Wetzels and colleagues (2011), Rouder and Morey (2012) and Liang and colleagues (2008), and computationally specified in the R package *BayesFactor* (Morey & Rouder, 2015). As evident in Table 1, across all three measures, there was evidence in favour of quadratic regressions over linear regressions. This was especially pronounced for the RAT and AUT Flexibility. For instance, in the case of the RAT results, the data are more than 2000 times more likely to have occurred under a quadratic model than a linear model, reflecting decisive evidence (for more details on evidence categories see: Wetzels et al., 2011; Jerffreys, 1961) in favour of a quadratic relative to a linear model. Moreover, the coefficients of the quadratic regressions, provided in Table 1, demonstrate a significant quadratic effect of political partisanship after adjusting for age, gender, and educational attainment, for all three tests of cognitive flexibility.

Table 1. Standardized coefficients of quadratic regressions predicting cognitive flexibility. BF_{ql} indicates the evidence in favour of a quadratic regression over a linear regression model. Interpretation reflects evidence categories for Bayes factors as suggested by Jeffreys (1961) and Wetzels and colleagues (2011). *p<.05, ***p<.001

Cognitive Flexibility Test	RAT	WCST	AUT Flexibility
Political Partisanship (PP_{max})			
Linear Effect	.00007	-.047	-.091*
Quadratic Effect	-.189***	-.112*	-.186***
Age	.149***	-.051	-.018
Gender	-.076	-.005	.006
Educational Attainment	.060	.022	.022
BF$_{ql}$	2117.664	1.782	3712.614
BF Interpretation	*Decisive evidence*	*Anecdotal evidence*	*Decisive evidence*

This was corroborated by a robustness check with the alternative measure of political partisanship, PP_{diff}. Regressing political partisanship (PP_{diff}) on RAT revealed a significant quadratic term (β=-.125, t=02.861, p=.004) and greater evidence for a quadratic than a linear association, BF_{ql} = 10.329. Similarly, there was a quadratic effect when regressing political partisanship (PP_{diff}) on AUT Flexibility (β =-.114, t=-2.731, p=.007), with strong evidence for a quadratic regression model BF_{ql} = 7.493. The quadratic effect on WCST was not statistically significant (β =-.097, t=-1.665, p=.097), with anecdotal evidence for a quadratic regression model BF_{ql} =1.018.

Interrupted Regressions

To validate the existence of inverted U-shaped relationships, we use an interrupted regression model (the two-lines test; Simonsohn, 2018). This method simultaneously estimates two regression lines – one for low *x*-values and one for high *x*-values. This facilitates the testing of a sign change, that is, whether the average effect of *x* on *y* is of opposite sign for high versus

low values of x. The two-lines test also identifies a data-driven change point where the two lines split using the "Robin Hood" algorithm, which has been demonstrated to obtain higher statistical power for detecting U-shaped relationships than other algorithmic alternatives (Simonsohn, 2018). An inverted-U-shaped relationship exists if the explanatory variable positively correlates with the outcome variable at low values, and negatively correlates with the outcome variable at high values.

The results are presented in Fig 3 (see also Supplementary Fig S2 for figures with data points). The slope between cognitive flexibility and political partisan identity on the left of the political spectrum is positive across all three measures of cognitive flexibility, suggesting that greater party partisanship on the political left is related to reduced cognitive flexibility. Symmetrically, the slope between cognitive flexibility and political partisan identity on the right of the political spectrum is negative, indicating that greater political partisanship on the political right is also related to reduced cognitive flexibility on the RAT and AUT. In the case of the WCST, the results indicated that the relationship between flexibility and partisanship was present on the political left but not political right.

Fig 3. Interrupted regression results according to cognitive flexibility test. Political partisanship is operationalized according to PP_{max}, such that partisan intensity is weighted by partisan direction (below 0 reflects left-leaning partisanship and above 0 reflects right-leaning partisanship). The shaded area reflects 95% confidence intervals.
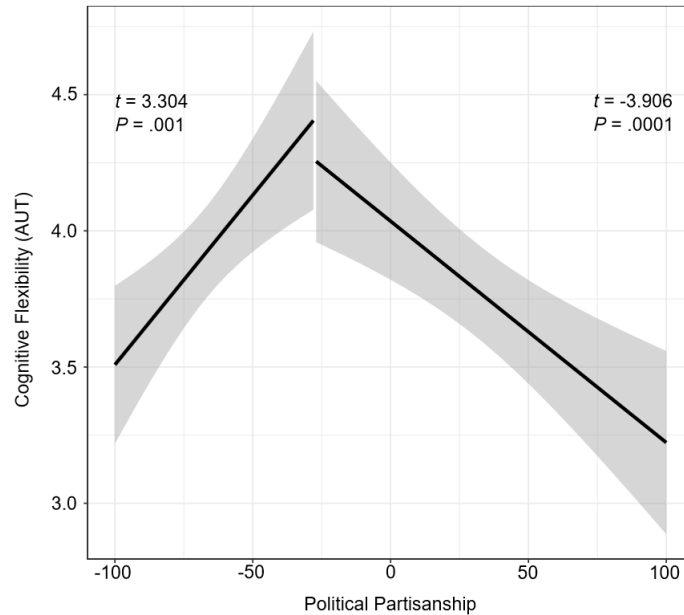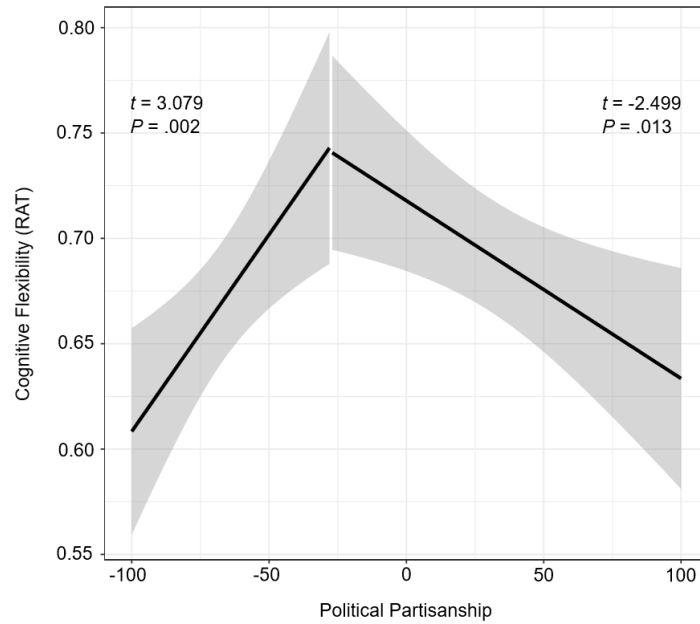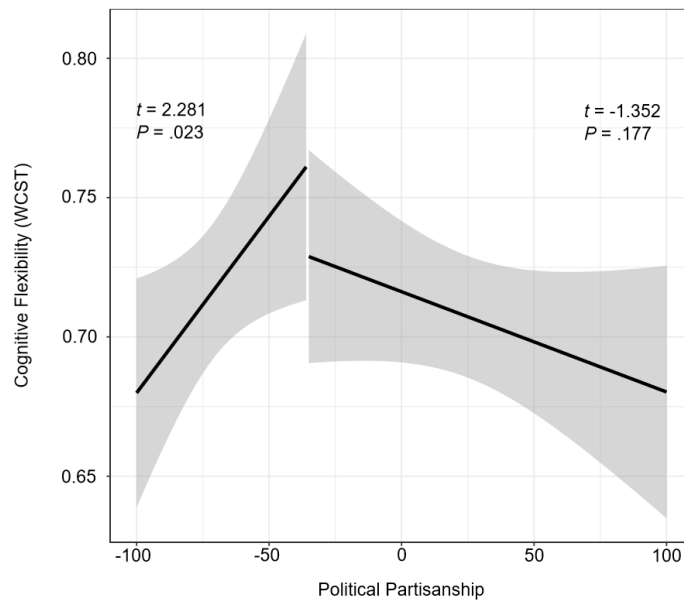
The Specificity of Flexibility

Additionally, since the AUT assesses multiple aspects of cognition, including flexibility, fluency, elaboration, and originality, we were able to investigate whether cognitive flexibility is specifically implicated in political partisanship relative to other facets of cognition. We conducted a multiple linear regression predicting political partisanship, as operationalized in $PP_{|max|}$, with the four AUT measures as predictors. This regression was significant, $F(4,619)=5.353$, $p=.0003$, and revealed that cognitive flexibility was the only significant predictor out of the four measurements ($\beta = -.166$, $t=-3.284$, $p=.001$), and elaboration ($\beta = -.043$, $p=.319$), fluency ($\beta = .054$, $p=.492$), and originality ($\beta = -.056$, $p=.440$) were not related to political partisanship.

Bayes factor analysis corroborated these findings (see Supplementary Fig S1). Computing the Bayesian regressions and associated Bayes factors for all possible combinations of AUT sub-measure predictors allows us to balance predictive power and parsimony (by removing redundant predictors). This analysis revealed that the best model is the one that predicts political partisanship only with cognitive flexibility, and no other AUT predictors. The Bayes factor for this model was $BF_{10}=1100.362$, indicating that the data are 1100 times more likely

under this model than an intercept-only null model (H_0). Furthermore, the flexibility-only model (subscript x) was over 100 times more likely than the full model (subscript f) with all AUT sub-measures for these data ($BF_{xf}=109.551$).

The Role of Political Conservatism

The analyses above demonstrate that the relationship between political partisanship and cognitive inflexibility is evident on both the political right and left, thus giving support for the rigidity-of-the-extreme hypothesis. The rigidity-of-the-right hypothesis also makes a prediction regarding political conservatism specifically: that there would be a linear relationship between conservatism and cognitive inflexibility. Correlational and Bayesian analyses offered inconclusive support for this hypothesis. Social conservatism was significantly correlated with AUT Flexibility ($r=-.14$, $p=.001$, $BF_{10}=3.09$ (in favour of relationship)) but not with WCST ($r=-.03$, $p=.567$, $BF_{01}=2.34$ (in favour of null)) or RAT ($r=-.08$, $p=.055$, but approached significance, $BF_{01}=0.91$ (in favour of null)). Economic conservatism was not significantly correlated with AUT Flexibility ($r=-.07$, $p=.067$, but approached significance, $BF_{01}=1.15$ (in favour of null)), WCST ($r=.04$, $p=.451$, $BF_{01}=2.24$ (in favour of null)), or RAT ($r=-.03$, $p=.535$, $BF_{01}=2.57$ (in favour of null)).

To ensure that relationships were not attenuated due to the influence of covariates, we conducted a set of multiple hierarchical linear regressions, in which social and economic conservatism were simultaneously included as predictors of cognitive flexibility, while adjusting for the demographic covariates of age, gender, and educational attainment. Social conservatism, economic conservatism, and the demographic variables were therefore included in the first step of the hierarchical linear regression for each cognitive flexibility measure. In the second step, we tested whether the relationship demonstrated above between cognitive inflexibility and political partisanship (unweighted by partisan direction, as operationalized

with the collapsed $PP_{|max|}$ measure) would persist after accounting for any relationships between cognitive inflexibility and political conservatism. Consequently, in the second step of the hierarchical regressions, we included $PP_{|max|}$ as a predictor of cognitive flexibility.

With respect to predicting flexibility on the RAT, social conservatism was a significant negative predictor (β =-.170, t=-2.842, p =.005), while economic conservatism was not (β =.085, t=1.452, p =.147) in the first step. In the second step, political partisanship was a significant negative predictor of flexibility (β =-.207, t=-4.665, p <.001) as was social conservatism (β =-.123, t=-2.076, p =.038). Economic conservatism was still not a significant predictor in the second step (β =.051, t=.884, p =.377). Furthermore, predicting flexibility on the AUT demonstrated that economic conservatism was not a significant predictor in the first step (β =-.024, t=-.418, p =.676) and social conservatism approached significance (β =-.108, t =-1.892, p =.059). In the second step, political partisanship was a significant negative predictor (β =-.162, t =-3.762, p<.001) while neither economic conservatism (β =-.048, t =-.852, p =.395) or social conservatism (β =-.073, t =-1.273, p =.204) were statistically significant predictors. WCST was not linearly predicted by either social conservatism (β =-.086, t =-1.096, p =.274) or economic conservatism (β =.090, t =1.178, p =.240) in the first step, and was also not predicted by political partisanship in the second step (β =-.057, t =-.996, p=.320). This may be due to the asymmetry evident in the interrupted regressions (Fig 3; Supplementary Fig S2) whereby there is a significant relationship between WCST flexibility and partisanship with regards to the Democratic Party, but not the Republican Party.

Lastly, we also tested whether quadratic relationships between social conservatism and cognitive flexibility would better reflect the data than linear relationships. Bayesian analyses suggested that a linear relationship was either better than a quadratic relationship or there was insufficient data to conclude (WCST: BF_{lq}=3.433, AUT: BF_{lq}=1.64, RAT: BF_{lq}=.501), after controlling for the demographic covariates. Notably, the rigidity-of-the-extreme hypothesis

does not predict a curvilinear relationship with regards to conservatism specifically – it predicts that extreme ideologues or partisans would exhibit cognitive rigidity relative to moderates (as shown in the sections above). Consequently, we would not expect a curvilinear relationship with respect to conservatism, because it is not a measure that was designed to measure strength of ideology – rather it focuses on particular policy attitudes.

Discussion:

The present investigation sought to address the question: *Does mental rigidity reflect one's partisan intensity or political orientation?* The results reveal that strong partisan intensity predicts reduced cognitive flexibility, regardless of the political party's orientation and doctrine. Quadratic regressions revealed that strong partisan intensity, on both the political left and right, was related to reduced flexibility across all three tests of cognitive flexibility (Table 1). This was corroborated by Bayes factor analysis, which demonstrated that the relationship between political partisanship across the political spectrum was quadratically – rather than linearly – related to cognitive flexibility (Table 1). Bayes factor analysis illustrated that the data were over 2000 times more likely to occur under a quadratic model than a linear regression model for the RAT and AUT (Table 1). The inverted U-shaped relationship between flexibility and partisanship was further validated with interrupted regressions (the two-lines test; Simohnson, 2017; Fig 3), underscoring that lower cognitive flexibility is evident amongst strong political partisans of both liberal and conservative ideologies. This was further corroborated by the finding that participants who self-described as political Independents exhibited heightened cognitive flexibility relative to Democrats and Republicans on the WCST and RAT (Fig 2). Consequently, investigating the roots of partisan intensity uncovers important psychological similarities between adherents to opposing political ideologies.

Analysis of participants' performance on the AUT, which measures multiple aspects of cognition, highlighted that flexibility was specifically implicated as a psychological correlate of partisanship. Other cognitive traits, such as fluency, elaboration, or originality, were not significantly related to partisan intensity. Moreover, when political partisanship was regressed on all possible predictor combinations of the four AUT cognitive measures, the model consisting of cognitive flexibility as the only predictor was the best model (i.e. with the greatest

evidential strength). The data were more than 1000 times more likely to occur under a model predicted only by cognitive flexibility than a null hypothesis model (Supplementary Fig S1).

These results have implications for the two dominant hypotheses about the nature of mental flexibility and political ideology. To the best of our knowledge, these findings constitute the first direct objective testing of the *ideological extremity hypothesis* using behavioural assessments of cognitive flexibility rather than self-report questionnaires. The data here support the essential claim of the ideological extremity hypothesis: political extremists were more cognitively rigid than political moderates, across multiple tests of cognitive flexibility (Table 1, Figs 2 and 3). These results suggest that the *rigidity-of-the-right hypothesis* may be incomplete, as it does not account for the presence of the "rigidity-of-the-left".

When partisan intensity and partisan direction (i.e. conservatism) were simultaneously entered into a linear multiple regression, the relationship between partisan intensity and cognitive inflexibility persisted after controlling for conservatism. This adds to an emerging literature illustrating that political extremists across the political spectrum tend to possess and exhibit similar levels of dogmatism and prejudice (for review see: Brandt, Reyna, Chambers, Crawford, & Wetherell, 2014), partisan bias and selective exposure to opposing opinions (Ditto et al., 2018; Frimer, Skitka, & Motyl, 2017), moral motives (Frimer, Biesanz, Walker, & MacKinlay, 2013; Frimer, Tell, & Motyl, 2017), simplicity of political categorization (Lammers et al., 2017), and belief in conspiracy theories (Krouwel, Kutiyski, van Prooijen, Martinsson, & Markstedt, 2017). Notably, social conservatism, but not economic conservatism, was a significant predictor of cognitive inflexibility, indicating that there may still be ideological asymmetries that need to be empirically evaluated (Jost, 2017). Note, however, that the regression coefficient of partisan intensity was consistently larger and more statistically significant than that of social conservatism in the multiple linear regression predicting cognitive flexibility (see Results section). Partisan intensity therefore appears to be more

predictive of mental flexibility than partisan direction. Economic conservatism was consistently statistically insignificant as a predictor of flexibility. This offers nuanced directions for future research on the rigidity-of-the-right hypothesis – perhaps it is specific to social, as opposed to economic, right-wing conservatism (see also Crowson, 2009; Feldman & Johnson, 2014; Malka & Soto, 2015). Moreover, this gives rise to the question: given the cognitive similarity between individuals on both political extremes, what factors determine their political orientation? Future studies should seek to replicate and expand these results, as well as explore ways in which the two hypotheses can be combined and empirically negotiated. Jost and colleagues (2003b) proposed a way in which the ideological extremity and rigidity-of-the-right hypotheses may be combined (see Figure 2 in their original paper), whereby there would be a U-shaped relationship between mental rigidity and political partisanship, but there would be a weaker relationship between partisan intensity and mental rigidity on the political left relative to the political right. This merits future examination.

The present investigation is relevant to other studies in political psychology which have sought to use indicators of cognitive processing that do not explicitly rely on self-reports. These studies have not directly assessed cognitive flexibility, instead focusing on other cognitive domains, such as confidence and meta-cognition (Brandt et al., 2015; Rollwage et al., 2018), exploratory behaviour (Shook & Fazio, 2009), integrative complexity (Tetlock et al., 1994; Van Hiel & Marvielde, 2003), information evaluation (Ditto et al., 2018), political categorization tendencies (Lammers et al., 2017), and threat sensitivity (Hibbing et al., 2014). Moreover, many of these tasks inherently possess political content (e.g. Ditto et al., 2018; Tetlock et al., 1994; Lammers et al., 2017) and so are not reflections of ideologically-neutral cognitive tendencies. Future research will benefit from examining multiple cognitive domains in tandem in order to evaluate the relative contributions of these individual differences to political ideology and how they may interact (for a review see: van Prooijen & Krouwel, 2019).

Furthermore, it is important to acknowledge the limitations of obtaining representative samples using online participant samples, such as Amazon Mechanical Turk, and so replicating these effects in nationally representative samples in the US and other political systems will constitute a valuable validation.

In sum, the present findings signify that the way individuals process neutral stimuli and react to the environment can reveal how they process social and political information, and consequently how they form their ideological attitudes and political identities. Moreover, the findings highlight that investigating *processes* such as partisan intensity, attachment, and extremism across opposing ideological orientations may be as scientifically fruitful (if not more so) as studying the *content* of those ideologies. This is consistent with recent research across multiple studies demonstrating that objectively-assessed cognitive inflexibility is related to greater ideological thinking in the realms of nationalism (Zmigrod, Rentfrow, & Robbins, 2018), religiosity (Zmigrod, Rentfrow, Zmigrod, & Robbins, 2018), extremist attitudes (Zmigrod, Rentfrow, & Robbins, 2019), and general dogmatism (Zmigrod, Zmigrod, Rentfrow, & Robbins, 2019). The intensity and strictness of one's adherence to an ideology therefore appears to be more relevant to – and revealing of – one's mental flexibility than the content of one's favoured ideology. The results of this study, together with those across studies, suggest that the cognitively inflexible mind may be especially susceptible to the clarity, certainty, and safety frequently offered by strong loyalty to collective ideologies and doctrines, regardless of their subject matter and motivation. This is in line with Rokeach's (1954) argument that adherents of both extreme left-wing and right-wing ideologies would exhibit tendencies towards rigidity. The cognitively flexible mind may be more likely to formulate socio-political opinions in a way that does not fully conform with the particular constellation of beliefs advocated by a political party. These findings nicely echo Hoffer's early theoretical

writings which suggested that "*all movements, however different in doctrine and aspiration…*

all appeal to the same types of mind" (Hoffer, 1951).

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