

Alertness and cognitive control: Testing the spatial grouping hypothesis

Darryl W. Schneider¹

Published online: 23 May 2019
© The Psychonomic Society, Inc. 2019

Abstract

Alertness seems to influence selective attention processes underlying cognitive control in the flanker task, as indicated by previous findings of larger congruency effects on alert trials (in which task stimuli are preceded by alerting cues) than on no-alert trials. One hypothesis for the alerting–congruency interaction is that increased alertness promotes spatial grouping of the target and distractors. In the present study, the author tested the spatial grouping hypothesis in three experiments in which the spatial alignment (collinearity) of the target and distractors was manipulated. Reliable alerting–congruency interactions were obtained, and congruency effects on response times were smaller for misaligned stimuli than for aligned stimuli in all experiments. However, the alerting–congruency interactions were not consistently modulated by alignment, contrary to a prediction derived from the spatial grouping hypothesis. The results suggest that spatial grouping is not a viable mechanism for explaining the alerting–congruency interaction in the flanker task, helping to prune the space of theoretical possibilities for linking alertness to cognitive control.

Keywords Alertness · Cognitive control · Selective attention · Flanker task · Stimulus alignment

The ability to focus attention on targets and ignore distractors in the environment is an important but imperfect part of cognitive control. It is important because selective attention filters out irrelevant information that might impair cognitive processing. It is imperfect because attentional selectivity is limited. This limitation is evident in the flanker task (Eriksen & Eriksen, 1974), in which identification of the central target in a visual stimulus array is influenced by adjacent distractors. Empirical findings (e.g., Gratton, Coles, & Donchin, 1992; Heitz & Engle, 2007) and computational models (e.g., Cohen, Servan-Schreiber, & McClelland, 1992; White, Ratcliff, & Starns, 2011) suggest that attentional focusing on the target can be modulated, motivating the investigation of factors that affect this process. In the present study, I investigated how alertness might influence selective attention processes underlying cognitive control in the flanker task.

Alertness refers to a temporary or sustained state of heightened sensitivity to stimulation (Posner, 2008). One way of increasing alertness is to present a cue or warning signal in

advance of the target stimulus, which often yields a general improvement in performance (e.g., Posner & Boies, 1971). However, a curious effect has been found when manipulating alertness in the flanker task, as shown in variants of the Attention Network Test (ANT; Fan, McCandliss, Sommer, Raz, & Posner, 2002). On each trial of the ANT, subjects perform a flanker task that involves making a spatially compatible response to the left/right direction of a central target arrow flanked by distractor arrows. The distractors can be associated with the same response as the target (congruent) or the opposite response (incongruent). A common finding is a congruency effect—longer response times (RTs) to incongruent stimuli than to congruent stimuli—that is thought to reflect a limitation of cognitive control in focusing attention on the target and filtering out the distractors. When an alerting cue is presented in advance of the arrow stimuli, RTs are generally shorter on these alert trials than on no-alert trials, indicating that a momentary increase in alertness facilitates performance. Surprisingly, despite the overall performance benefit of increased alertness, congruency effects are typically larger on alert trials than on no-alert trials (e.g., Fan et al., 2002; Redick & Engle, 2006; Schneider, 2018a, 2018b; for overviews, see MacLeod et al., 2010; Nieuwenhuis & de Kleijn, 2013). This alerting–congruency interaction suggests that there is an intricate connection between alertness and

✉ Darryl W. Schneider
dws@purdue.edu

¹ Department of Psychological Sciences, Purdue University, 703 Third Street, West Lafayette, IN 47907, USA

cognitive control, and two lines of evidence implicate spatial attention as a critical factor.

First, the alerting–congruency interaction is routinely found in tasks that involve spatial separation of the target and distractors, such as the flanker task or variants of the ANT (e.g., Fan et al., 2002; MacLeod et al., 2010; Redick & Engle, 2006; Schneider, 2018a, 2018b).¹ In contrast, the interaction is often not found in tasks in which the target and distractor features are spatially integrated, such as the color-word Stroop task (Schneider, 2019; Soutschek, Müller, & Schubert, 2013; Weinbach & Henik, 2012).² For example, in a study by Weinbach and Henik (2012), subjects performed a Stroop-like task that involved judging a target color (making a left or right key-press response) while ignoring distractor arrows. In one experiment, the target and distractor features were spatially integrated by presenting colored arrows to subjects, and no alerting–congruency interaction was found. In another experiment, the target and distractor features were spatially separated by presenting colored rectangles flanked by arrows, and a reliable alerting–congruency interaction was found.³ These results led the authors to conclude that “the critical mechanism responsible for the appearance of the interaction is modulation of spatial attention following alerting cues” (Weinbach & Henik, 2012, p. 1537).

Second, the alerting–congruency interaction in the flanker task with arrow stimuli appears to be modulated by spatial variables, but not by nonspatial variables. In Schneider (2018b), I manipulated the horizontal spacing (a spatial variable) or color grouping (a nonspatial variable) of targets and distractors in a variant of the ANT. Replicating previous studies of spacing effects in the flanker task (e.g., Eriksen & Eriksen, 1974; Miller, 1991), congruency effects decreased from narrow to wide spacing of stimuli. Interestingly, the alerting–congruency interaction was present under narrow spacing but absent under wide spacing, producing a significant three-way interaction between spacing, alerting, and congruency. Replicating previous studies of color-grouping effects in the flanker task (e.g., Baylis & Driver, 1992; Kramer & Jacobson, 1991), congruency effects were smaller when

targets and distractors were different colors than when they were the same color. However, the alerting–congruency interaction was unaffected by color grouping, being similar in magnitude in the same-color and different-color conditions.

The results in Schneider (2018b) led me to propose a spatial grouping hypothesis for the alerting–congruency interaction. The hypothesis is a refinement of the global processing hypothesis put forth by Weinbach and Henik (2011, 2012). They suggested that increased alertness yields a bias toward global processing of stimuli, such that attention is less localized to the target and encompasses more distractors. The spatial grouping hypothesis provides an underlying mechanism for a global processing bias by assuming that increased alertness promotes spatial grouping of the target and distractors into a common object that can be selected by attention. In the flanker task, where the instructions are to identify the target and ignore the distractors, spatial grouping of stimuli would impair discrimination between the target and distractors because the latter would be attended and processed, resulting in congruency effects (Logan, 1996). If increased alertness makes spatial grouping more likely, then a larger congruency effect would occur on alert trials than on no-alert trials, consistent with the typical alerting–congruency interaction. In contrast, manipulations that make spatial grouping less likely should attenuate the alerting–congruency interaction, consistent with what happened in Schneider (2018b) when stimulus spacing was manipulated.

I proposed the spatial grouping hypothesis specifically to explain the data in Schneider (2018b).⁴ An important next step is to test the hypothesis in new experiments, ideally by manipulating a variable that differs from stimulus spacing but is still likely to influence spatial grouping. In the present study, I tested the spatial grouping hypothesis by manipulating the spatial alignment of the target and distractors in a variant of the ANT. Subjects performed a flanker task that involved making a spatially compatible key-press response to the left/right direction of a central target arrow flanked by distractor arrows. The distractors were associated with the same response as the target (congruent) or the opposite response (incongruent). On no-alert trials, a variable fixation interval preceded the arrow stimuli. On alert trials, an alerting cue (a filled square at the location of the forthcoming target arrow) was presented briefly during the fixation interval, shortly before the arrow stimuli. Potential alerting–congruency interactions were assessed by comparing the congruency effects on alert and no-alert trials.

Besides congruency and alerting, I manipulated the spatial alignment (collinearity) of the target and distractors without

¹ There is some evidence that the occurrence of the alerting–congruency interaction in the flanker task depends on the type of stimuli used (e.g., Fischer, Plessow, & Kiesel, 2012; Seibold, 2018; for a summary and discussion, see Schneider, 2019, pp. 133–134).

² An exception seems to be the Simon task (for a review, see Lu & Proctor, 1995), which typically involves making a spatial response to a nonspatial feature (e.g., color) of a stimulus presented at an irrelevant spatial location that is either congruent or incongruent with the response. When alertness has been manipulated in the Simon task, alerting–congruency interactions have been found (Böckler, Alpay, & Stürmer, 2011; Fischer, Plessow, & Kiesel, 2010; Klein & Ivanoff, 2011; Soutschek et al., 2013). Considering that the distractor feature in the Simon task is the spatial location of the stimulus, these findings are consistent with my general point that spatial attention seems to be a critical aspect of the link between alertness and cognitive control.

³ This finding was not replicated in a recent study from my laboratory (see Experiment 8 in Schneider, 2019).

⁴ The spatial grouping hypothesis was developed in the context of the flanker task, and it cannot explain why alerting–congruency interactions have been found in the Simon task (see footnote 2), where spatial grouping would not be expected to play a role in task performance.

changing the horizontal spacing of the stimuli (see Fig. 1). In aligned stimulus displays, the target and distractor arrows were collinear because they were located at the same vertical position onscreen, as in Schneider (2018b) and in previous studies using the ANT. In misaligned stimulus displays, all the distractor arrows were shifted up or down relative to the target arrow, such that the distractors were collinear with each other but not with the target. According to the Gestalt laws of perceptual organization (for overviews, see Hartmann, 1935; Palmer, 1999), stimuli that are similar in terms of spatial location, orientation, color, size, or other properties are more likely to be grouped together than are stimuli that are dissimilar. Moreover, linearity has been identified as an emergent feature in the visual system that can produce Gestalt-like grouping of stimuli (Pomerantz & Portillo, 2011). Therefore, it seemed reasonable to assume that the distractors were more likely to be spatially grouped with the target when all the stimuli were aligned (i.e., collinear) than when they were misaligned.

Applying the spatial grouping hypothesis to this context, I predicted a three-way interaction between alignment, alerting, and congruency, analogous to the three-way interaction between spacing, alerting, and congruency obtained in Schneider (2018b). Specifically, if increased alertness makes spatial grouping more likely, then a typical alerting–congruency interaction (i.e., a larger congruency effect on alert trials than on no-alert trials) should occur for aligned stimuli, replicating previous results. In contrast, misaligning the target and distractor arrows should make spatial grouping of the stimuli less likely, mirroring the effect of wider spacing of target and distractor arrows in Schneider (2018b). A decrease in spatial grouping due to misalignment should counteract any increase in spatial grouping due to alerting, thereby attenuating the alerting–congruency interaction. Thus, the predicted three-way interaction is that the difference in

congruency effects between alert and no-alert trials should be smaller for misaligned stimuli than for aligned stimuli.⁵ This prediction was tested in the present experiments, which were identical except for how the alignment manipulation was implemented and the amount of misalignment (as a percentage of the arrow height). In Experiments 1 and 3, aligned and misaligned stimuli were intermixed randomly within every block of trials. In Experiment 2, the alignment manipulation was blocked, such that aligned and misaligned stimuli were presented in separate halves of the experiment. Stimuli were misaligned by 59% in Experiments 1 and 2, whereas they were misaligned by 118% in Experiment 3 (see Fig. 1).⁶

Experiment 1

Method

Subjects A total of 110 students from Purdue University participated for course credit. The preregistered sample size was determined from a simulation-based power analysis that indicated 110 subjects would provide 85% power to detect a three-way interaction between alignment, alerting, and congruency on RTs that was approximately the same effect size (standardized or not) as the three-way interaction between spacing, alerting, and congruency on RTs obtained in Experiment 1 of Schneider (2018b). Data from two additional subjects were excluded: One subject's data were incomplete because of a computer malfunction, and the other subject's mean error rate exceeded the preregistered criterion of 20%. All subjects in the present study reported having normal or corrected-to-normal vision.

Apparatus, task, and stimuli Computers controlled the display of stimuli on monitors and the registration of responses from






Alignment	Direction	Example Stimulus Display
Aligned	—	
Misaligned (59%)	Up	
	Down	
Misaligned (118%)	Up	
	Down	

Fig. 1 Example stimulus displays as a function of alignment. Stimuli misaligned by 59% were used in Experiments 1 and 2, whereas stimuli misaligned by 118% were used in Experiment 3. All experiments also included aligned stimuli. All examples depict incongruent stimuli that require right-button responses

⁵ A reviewer wondered whether one might predict the opposite pattern—a larger difference in congruency effects between alert and no-alert trials for misaligned stimuli than for aligned stimuli. The reasoning was as follows: If misaligning the stimuli reduces the likelihood of spatial grouping, then there is more room for an increase in spatial grouping due to alerting. The result could be an amplified alerting–congruency interaction for misaligned stimuli. The same reasoning can be applied to the spacing manipulation in Schneider (2018b), where wider spacing of stimuli should have reduced the likelihood of spatial grouping. However, instead of the alerting–congruency interaction being amplified under wider spacing, it was eliminated. This surprising result, which was replicated in Schneider (2018b), provides an empirical basis for my predicted three-way interaction in the present study.

⁶ Experiment 1 was a preregistered test of the spatial grouping hypothesis. The protocol is publicly available (<https://aspredicted.org/x5st2.pdf>). Experiment 2 was not preregistered separately, but it adhered closely to the protocol for Experiment 1. The “unpublished experiment” referenced in the protocol is Experiment 1 of Schneider (2018b), which had not yet been published at the time of preregistration. Experiment 3 was conducted in response to an issue raised by a reviewer. It was a preregistered replication of Experiment 1, but with a larger sample size and a greater amount of misalignment. The protocol is publicly available (<https://aspredicted.org/xy6je.pdf>). This article includes all the experiments that were conducted for the present study.

Chronos devices (Psychology Software Tools, Inc.). Stimuli were displayed in white on a black background at an uncontrolled viewing distance of approximately 50 cm. Responses were made by pressing the leftmost and rightmost buttons on the response device with the left and right index fingers, respectively.

The task involved making a spatially compatible response to identify the left/right direction of the central target arrow in a stimulus display consisting of five arrows (see Fig. 1). The target arrow was flanked horizontally by distractor arrows (two on each side), all pointing in either the same direction as the target (congruent) or in the opposite direction (incongruent). Each arrow was 1.15° wide \times 0.46° high. The horizontal spacing of adjacent arrows was 0.11° (edge-to-edge distance), which matches the narrow spacing in Schneider (2018b). In aligned stimulus displays, all arrows were at the same vertical position, forming a collinear arrangement. In misaligned stimulus displays, all the distractor arrows were shifted up or down by 0.27° (59% of the arrow height) relative to the target arrow (see Fig. 1).

Procedure Subjects were tested individually after providing informed consent for a study protocol approved by the Purdue University Institutional Review Board. Instructions were presented on-screen and were read aloud by the experimenter. Subjects were instructed to identify the direction of the central arrow on each trial, ignoring the other arrows. During the instructions, they completed eight example no-alert trials (reflecting all combinations of congruency, alignment, and target direction) with accuracy feedback. Afterward, they completed 13 blocks of 32 trials per block without accuracy feedback, although they were encouraged to respond quickly and accurately.

Each trial began with a fixation cross ($0.23^\circ \times 0.23^\circ$) presented in the center of the screen. On no-alert trials, the fixation cross was replaced by an arrow stimulus display after a randomly selected fixation interval of 900 ms; 1,300 ms; 1,700 ms; or 2,100 ms. Regardless of alignment condition, the target arrow was always presented centrally, at the former location of the fixation cross. The stimuli remained on-screen until a response was registered, then the screen was cleared immediately, and the fixation cross for the next trial appeared. On alert trials, the procedure was identical, except that 500 ms before the arrow stimulus display, an alerting cue (a filled square; $0.46^\circ \times 0.46^\circ$) was presented centrally for 100 ms, temporarily replacing the fixation cross.

Every possible combination of alerting, congruency, alignment, and target direction occurred equally often in random order in each block. For half of the misaligned stimuli, the distractor arrows were shifted up; for the other half, they were shifted down. The form of misalignment was selected randomly, but constrained such that all combinations of the other variables occurred equally often with each form. The fixation

intervals were randomized independently of the other variables, but constrained to occur equally often in each block.

Results

Data trimming and analysis followed the preregistered protocol. The first block of trials was excluded as practice. Trials with RTs exceeding three standard deviations of the mean in each condition for a given subject were excluded (1.7% of trials). Error trials were excluded from the RT analysis. Mean RTs and mean error rates appear as a function of alignment (aligned or misaligned), alerting (no alert or alert), and congruency (incongruent or congruent) in Table 1 and in Fig. 2. The results of repeated-measures analyses of variance (ANOVAs) with those variables as factors are summarized in Table 2. All significant effects ($\alpha = .05$) are described below.

A limitation of the null hypothesis significance testing logic underlying ANOVA is that a nonsignificant effect cannot be interpreted as support for the absence of an effect (i.e., the null hypothesis cannot be accepted). However, establishing the absence of an effect can have important theoretical implications, especially in the case of the three-way interaction that is critical for assessing the spatial grouping hypothesis in the present study. For this reason, the ANOVAs are supplemented by preregistered Bayes-factor analyses of the interactions (scaled JZS Bayes factors with $r = 1$; Rouder, Speckman, Sun, Morey, & Iverson, 2009), enabling me to quantify the strength of evidence in favor of either the null hypothesis or the alternative hypothesis for each interaction effect. Following Schneider (2019), each Bayes factor (BF) is reported with a subscript (null or alt) indicating the statistical hypothesis (null or alternative) under which the data are more likely. For example, $BF_{\text{null}} = 10$ indicates that the data are 10 times more likely under the null hypothesis of no difference than the alternative hypothesis of a difference, whereas $BF_{\text{alt}} = 5$ indicates that the data are 5 times more likely under the alternative hypothesis than the null hypothesis. Consistent with recent proposals (Held & Ott, 2018; Schönbrodt & Wagenmakers, 2018) and the categorization scheme used in

Table 1. Mean response times and mean error rates in Experiment 1

Alignment	Alerting	Congruency	Response time (ms)	Error rate (%)
Aligned	No alert	Incongruent	486 (69)	1.7 (2.8)
		Congruent	440 (56)	0.6 (1.3)
	Alert	Incongruent	465 (57)	2.0 (2.8)
		Congruent	413 (52)	0.4 (1.2)
Misaligned	No alert	Incongruent	476 (59)	2.7 (3.6)
		Congruent	442 (57)	1.0 (1.7)
	Alert	Incongruent	458 (50)	3.8 (4.0)
		Congruent	415 (55)	0.7 (1.8)

Note. Standard deviations appear in parentheses

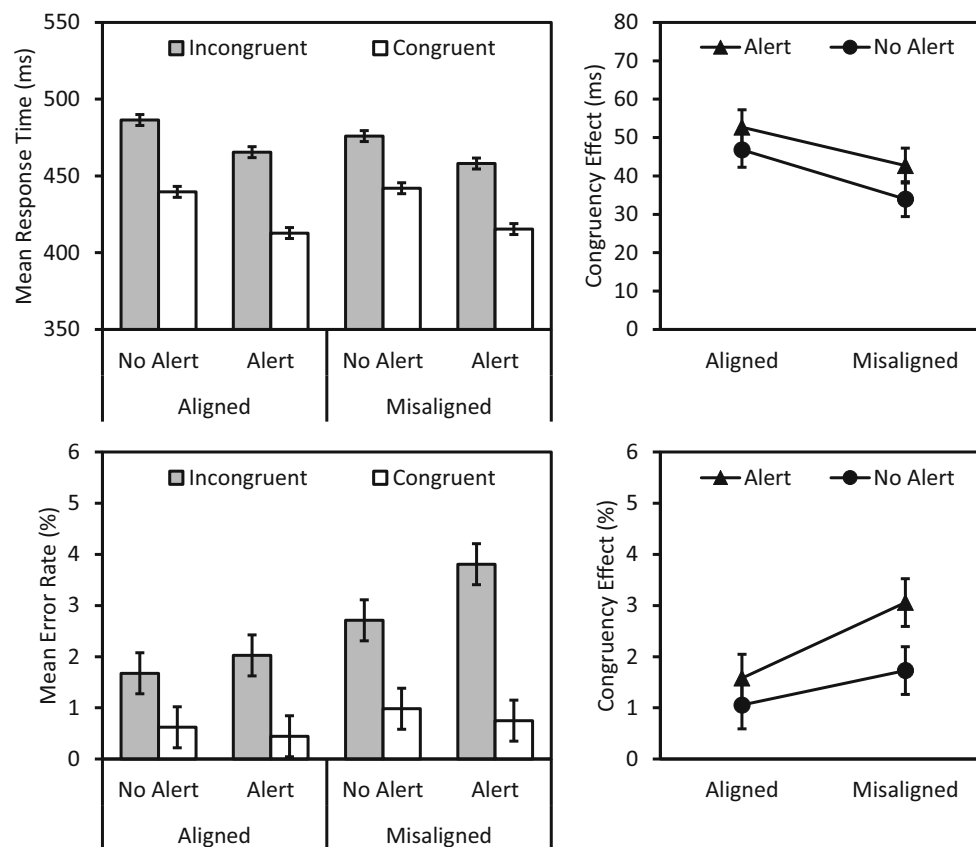


Fig. 2 Experiment 1 data. Mean response times and mean error rates appear in the top-left and bottom-left panels, respectively, as a function of alignment, alerting, and congruency. Congruency effects (performance differences between incongruent and congruent trials) for response times

and error rates appear in the top-right and bottom-right panels, respectively, as a function of alignment and alerting. Error bars represent 95% confidence intervals (Masson & Loftus, 2003)

Schneider (2019), the numerical value of the Bayes factor can be categorized in terms of strength of evidence as follows: 1 = no evidence; 1–3 = weak evidence; 3–10 = moderate evidence; 10–30 = strong evidence; 30–100 = very strong evidence; and >100 = decisive evidence.

RTs Responses were slightly faster to misaligned stimuli (mean RT of 448 ms) than to aligned stimuli (451 ms),

yielding a significant main effect of alignment. Responses were faster on alert trials (438 ms) than on no-alert trials (461 ms), yielding a significant main effect of alerting. Responses were slower to incongruent stimuli (471 ms) than to congruent stimuli (427 ms), yielding a significant main effect of congruency. Two interactions were significant and both reflected modulation of congruency effects (see Fig. 2). First, there was an interaction between alignment and

Table 2. Summary of analyses of variance for Experiment 1

Effect	Response time				Error rate			
	$F(1, 109)$	MSE	p	η_p^2	$F(1, 109)$	MSE	p	η_p^2
Alignment (L)	7.09	324	.009	.061	54.11	3	<.001	.332
Alerting (A)	235.22	497	<.001	.683	2.64	6	.107	.024
Congruency (C)	688.65	620	<.001	.863	76.52	10	<.001	.412
$L \times A$.71	211	.401	.006	1.56	4	.215	.014
$L \times C$	23.89	300	<.001	.180	16.95	4	<.001	.135
$A \times C$	8.06	365	.005	.069	17.99	3	<.001	.142
$L \times A \times C$.52	218	.474	.005	3.08	3	.082	.027

Note. Bolded numbers indicate effects that were significant ($p < .05$) in all experiments

congruency, reflecting a smaller congruency effect for misaligned stimuli (38 ms) than for aligned stimuli (50 ms), $BF_{alt} = 3,377.2$. Second, there was an interaction between alerting and congruency, reflecting a larger congruency effect on alert trials (48 ms) than on no-alert trials (40 ms), $BF_{alt} = 3.6$. The three-way interaction of alignment, alerting, and congruency was nonsignificant, reflecting similar increases in congruency effects from no-alert to alert trials for aligned stimuli (6 ms) and for misaligned stimuli (9 ms), $BF_{null} = 10.3$.

Error rates Responses were incorrect more often to misaligned stimuli (mean error rate of 2.1%) than to aligned stimuli (1.2%), yielding a significant main effect of alignment. Responses were incorrect more often to incongruent stimuli (2.6%) than to congruent stimuli (0.7%), yielding a significant main effect of congruency. Two interactions were significant and both reflected modulation of congruency effects (see Fig. 2). First, there was an interaction between alignment and congruency, reflecting a larger congruency effect for misaligned stimuli (2.4%) than for aligned stimuli (1.3%), $BF_{alt} = 185.6$. Second, there was an interaction between alerting and congruency, reflecting a larger congruency effect on alert trials (2.3%) than on no-alert trials (1.4%), $BF_{alt} = 290.0$. The three-way interaction of alignment, alerting, and congruency was nonsignificant, although the increase in congruency effects from no-alert to alert trials was numerically smaller for aligned stimuli (0.5%) than for misaligned stimuli (1.3%), $BF_{null} = 2.9$.

Discussion

The predicted three-way interaction derived from the spatial grouping hypothesis was that the difference in congruency effects between alert and no-alert trials would be smaller for misaligned stimuli than for aligned stimuli. The results of Experiment 1 do not match the prediction. The three-way interaction between alignment, alerting, and congruency was nonsignificant for RTs and error rates, supported by Bayes-factor analyses favoring null interactions. Moreover, the data pattern was numerically in the opposite direction, with the difference in congruency effects between alert and no-alert trials being numerically larger—not smaller—for misaligned stimuli than for aligned stimuli (see Fig. 2).

Two other aspects of the data indicate that the null three-way interactions were not due to ineffective manipulations. First, there were significant alerting–congruency interactions for both dependent measures, reflecting larger congruency effects on alert trials than on no-alert trials (see Fig. 2), replicating the results of previous studies (e.g., Fan et al., 2002; MacLeod et al., 2010; Redick & Engle, 2006; Schneider, 2018a, 2018b). Second, there were significant alignment–congruency interactions for both dependent measures, but they reflected different data patterns: Congruency effects were

smaller for misaligned stimuli than for aligned stimuli in the RT data, whereas the opposite pattern was present in the error data (see Fig. 2). However, alignment did not modulate the alerting–congruency interaction, contrary to the prediction from the spatial grouping hypothesis.

Before discussing the results further, I deemed it important to try to replicate the findings from Experiment 1, for two reasons. First, given that the null three-way interactions between alignment, alerting, and congruency were in stark contrast to what was predicted, it would be helpful to demonstrate that the observed RT and error data patterns can be replicated before drawing a strong conclusion about the viability of the spatial grouping hypothesis. Second, even though the alignment manipulation affected behavior, its effects differed for RTs and error rates: Performance was slightly faster and congruency effects on RTs were smaller for misaligned stimuli than for aligned stimuli, whereas performance was more error prone and congruency effects on error rates were larger for misaligned stimuli than for aligned stimuli (see Fig. 2). These results raise the possibility that the alignment manipulation induced a form of speed–accuracy trade-off, which is an issue that warrants further investigation.

Experiment 2 was a replication of Experiment 1, but with one primary change: Instead of aligned and misaligned stimuli being intermixed randomly within every block of trials, they were presented in separate halves of the experiment. The rationale behind blocking the alignment manipulation was that if it induces a speed–accuracy trade-off, then the trade-off should be accentuated when each alignment condition is experienced in isolation because subjects can maintain consistent positions on their speed–accuracy trade-off functions for several blocks of trials. Thus, there was a possibility that the methodological change for Experiment 2 would lead to stronger alignment effects than those observed in Experiment 1.

Experiment 2

Method

Subjects A total of 110 students from Purdue University participated for course credit. None of them had participated in Experiment 1. The sample size was determined in the same manner as in Experiment 1. Data from two additional subjects were excluded for exceeding preregistered criteria: One subject's mean error rate exceeded 20%, and the other subject's mean RT exceeded four standard deviations of the group mean.

Apparatus, task, and stimuli These aspects of the experiment were identical to those of Experiment 1.

Procedure The procedure was identical to that of Experiment 1, except for changes related to blocking of the alignment

manipulation. Subjects completed 14 blocks of 32 trials per block. For half of the subjects, the first seven blocks consisted of aligned stimuli, and the last seven blocks consisted of misaligned stimuli. For the other half of the subjects, the order of the alignment conditions was reversed.

Results

Data trimming and analysis were identical to those of Experiment 1 (2.0% of trials were RT outliers), except the first block of each half was excluded as practice, to equate the amount of practice in each alignment condition. Mean RTs and mean error rates appear as a function of alignment, alerting, and congruency in Table 3 and in Fig. 3. The results of repeated-measures ANOVAs with those variables as factors are summarized in Table 4.

RTs Responses were faster on alert trials (434 ms) than on no-alert trials (457 ms), yielding a significant main effect of alerting. Responses were slower to incongruent stimuli (468 ms) than to congruent stimuli (423 ms), yielding a significant main effect of congruency. Two interactions were significant and both reflected modulation of congruency effects (see Fig. 3). First, there was an interaction between alignment and congruency, reflecting a smaller congruency effect for misaligned stimuli (41 ms) than for aligned stimuli (50 ms), $BF_{\text{alt}} = 4.7$. Second, there was an interaction between alerting and congruency, reflecting a larger congruency effect on alert trials (54 ms) than on no-alert trials (37 ms), $BF_{\text{alt}} = 2.8 \times 10^8$. The three-way interaction of alignment, alerting, and congruency was nonsignificant, reflecting similar increases in congruency effects from no-alert to alert trials for aligned stimuli (15 ms) and for misaligned stimuli (17 ms), $BF_{\text{null}} = 12.4$.

Error rates Responses were incorrect more often to incongruent stimuli (3.1%) than to congruent stimuli (0.6%), yielding a significant main effect of congruency. Two interactions were significant and both reflected modulation of congruency effects (see Fig. 3). First, there was an interaction between

alerting and congruency, reflecting a larger congruency effect on alert trials (3.1%) than on no-alert trials (1.9%), $BF_{\text{alt}} = 91.8$. Second, although the interaction between alignment and congruency was nonsignificant, the three-way interaction of alignment, alerting, and congruency was significant (see Table 4), reflecting a smaller increase in congruency effects from no-alert to alert trials for aligned stimuli (0.7%) than for misaligned stimuli (1.7%), $BF_{\text{null}} = 1.6$.

Discussion

The results of Experiment 2 replicate some of the key results of Experiment 1 (compare Figs. 2 and 3) and do not match the prediction derived from the spatial grouping hypothesis. The predicted three-way interaction between alignment, alerting, and congruency was nonsignificant for RTs, supported by a Bayes-factor analysis favoring a null interaction (as in Experiment 1). The interaction was significant for error rates, but a Bayes-factor analysis did not indicate strong evidence for or against an effect. Moreover, the error data pattern was in the opposite direction of the prediction (but consistent with the pattern in Experiment 1), with the difference in congruency effects between alert and no-alert trials being larger—not smaller—for misaligned stimuli than for aligned stimuli (see Fig. 3).

Recall that the rationale for blocking the alignment manipulation in Experiment 2 was that it should accentuate any speed–accuracy trade-off induced by the manipulation. On the one hand, the three-way interaction between alignment, alerting, and congruency that was nonsignificant for error rates in Experiment 1 was significant in Experiment 2 (in both cases, though, Bayes-factor analyses weakly favored null interactions). On the other hand, the significant but divergent main effects of alignment on RTs and error rates in Experiment 1 were both nonsignificant—but still in divergent directions—in Experiment 2. The only effect involving alignment that was significant in both experiments was the interaction between alignment and congruency on RTs (see Tables 2 and 4), reflecting a smaller congruency effect for misaligned stimuli than for aligned stimuli. Thus, the evidence for an alignment-related speed–accuracy trade-off seems somewhat weak.

The replicated finding of significantly smaller congruency effects on RTs for misaligned stimuli than for aligned stimuli indicates that the alignment manipulation was effective in influencing behavior. However, a reviewer noted that the alignment effects were considerably weaker than the spacing effects obtained by Schneider (2018b). More specifically, standardized effect sizes for the interaction between alignment and congruency on RTs ($\eta_p^2 = .180$ and $.073$ in the present Experiments 1 and 2, respectively) were smaller than those for the interaction between spacing and congruency on RTs ($\eta_p^2 = .680$ and $.653$ in Experiments 1 and 4, respectively, of

Table 3. Mean response times and mean error rates in Experiment 2

Alignment	Alerting	Congruency	Response time (ms)	Error rate (%)
Aligned	No alert	Incongruent	480 (71)	2.6 (3.4)
		Congruent	437 (57)	0.7 (1.6)
	Alert	Incongruent	464 (70)	3.1 (4.3)
		Congruent	406 (51)	0.5 (1.1)
Misaligned	No alert	Incongruent	473 (52)	2.7 (3.5)
		Congruent	440 (55)	0.9 (1.6)
	Alert	Incongruent	457 (47)	4.0 (4.2)
		Congruent	408 (50)	0.5 (1.1)

Note. Standard deviations appear in parentheses

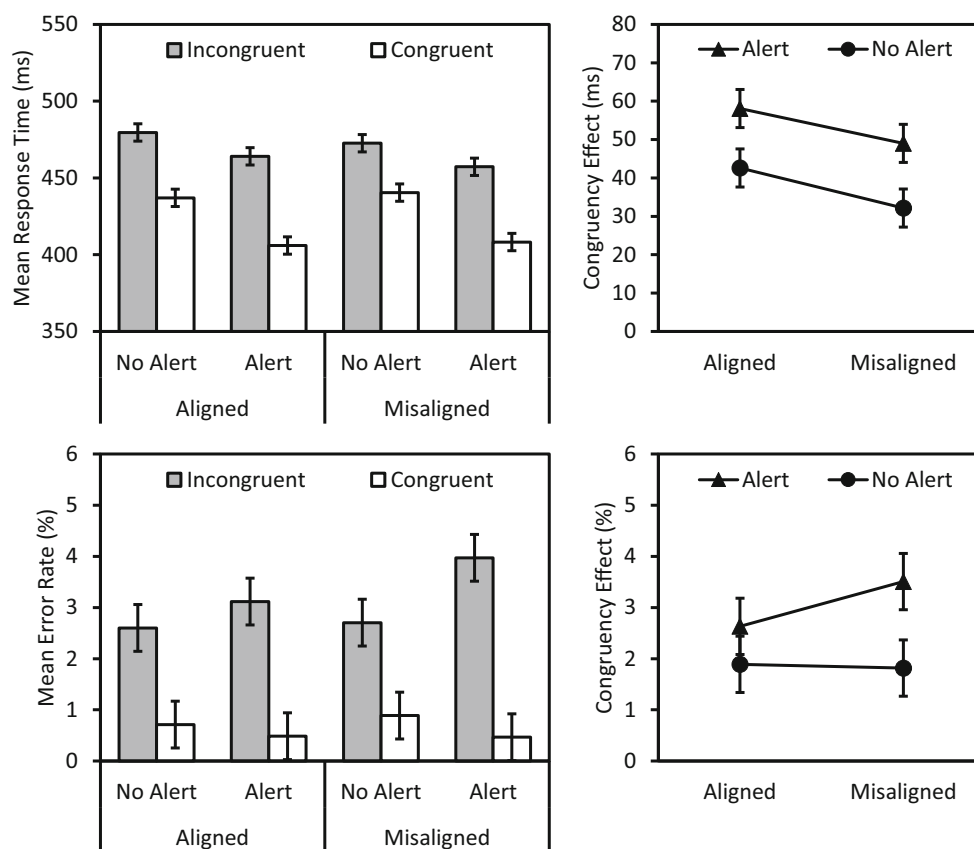


Fig. 3 Experiment 2 data. Mean response times and mean error rates appear in the top-left and bottom-left panels, respectively, as a function of alignment, alerting, and congruency. Congruency effects (performance differences between incongruent and congruent trials) for response times

and error rates appear in the top-right and bottom-right panels, respectively, as a function of alignment and alerting. Error bars represent 95% confidence intervals (Masson & Loftus, 2003)

Schneider, 2018b). This raises the possibility that the alignment manipulation might have been adequate for modulating congruency effects overall, but not potent enough for finer-grained modulation of the alerting–congruency interaction. Recall that the power analysis for Experiments 1 and 2 involved the assumption that the alignment manipulation would yield approximately the same effect size as the spacing manipulation. Given that this assumption appears to be invalid,

an implication is that the experiments might have lacked sufficient power to detect the predicted three-way interaction between alignment, alerting, and congruency, despite having much larger sample sizes ($Ns = 110$) than those typically obtained for studies of the alerting–congruency interaction.

I conducted a third experiment to address this issue. Experiment 3 was a preregistered replication of Experiment 1 (see footnote 6), but with two changes. First, I made the

Table 4. Summary of analyses of variance for Experiment 2

Effect	Response time				Error rate			
	$F(1, 109)$	MSE	p	η_p^2	$F(1, 109)$	MSE	p	η_p^2
Alignment (L)	.23	3,948	.631	.002	3.38	5	.069	.030
Alerting (A)	256.83	475	<.001	.702	3.36	5	.070	.030
Congruency (C)	630.06	722	<.001	.853	90.67	15	<.001	.454
$L \times A$.07	178	.791	.001	1.09	4	.298	.010
$L \times C$	8.64	607	.004	.073	1.86	5	.175	.017
$A \times C$	54.93	262	<.001	.335	15.32	5	<.001	.123
$L \times A \times C$.14	182	.709	.001	4.37	3	.039	.039

Note. Bolded numbers indicate effects that were significant ($p < .05$) in all experiments

alignment manipulation more potent by doubling the amount of misalignment from 59% to 118% of the arrow height (see Fig. 1). Note that any further increase in misalignment would make the alignment manipulation more akin to a vertical spacing manipulation, undermining my aforementioned goal of testing the spatial grouping hypothesis by manipulating a variable that differs from stimulus spacing. Second, I increased the sample size from 110 to 150 subjects. The combination of a stronger alignment manipulation and a larger sample size should have increased the statistical power to detect any three-way interaction between alignment, alerting, and congruency. Moreover, Experiment 3 provided another opportunity to replicate the data patterns observed in Experiments 1 and 2.

Experiment 3

Method

Subjects A total of 150 students from Purdue University participated for course credit. None of them had participated in Experiments 1 or 2. The preregistered sample size was chosen to be larger than that of Experiment 1 while fitting logistical and practical constraints on how many subjects could be run in my laboratory in a timely manner. Data from three additional subjects were excluded for exceeding a preregistered criterion: their mean RTs exceeded four standard deviations of the group mean.

Apparatus, task, and stimuli These aspects of the experiment were identical to those of Experiment 1, except that in misaligned stimulus displays, all the distractor arrows were shifted up or down by 0.54° (118% of the arrow height) relative to the target arrow (see Fig. 1).

Procedure The procedure was identical to that of Experiment 1.

Results

Data trimming and analysis followed the preregistered protocol and were identical to those of Experiment 1 (1.7% of trials were RT outliers). Mean RTs and mean error rates appear as a function of alignment, alerting, and congruency in Table 5 and in Fig. 4. The results of repeated-measures ANOVAs with those variables as factors are summarized in Table 6.

RTs Responses were faster to misaligned stimuli (438 ms) than to aligned stimuli (448 ms), yielding a significant main effect of alignment. Responses were faster on alert trials (433 ms) than on no-alert trials (453 ms), yielding a significant main effect of alerting. Responses were slower to incongruent

Table 5. Mean response times and mean error rates in Experiment 3

Alignment	Alerting	Congruency	Response time (ms)	Error rate (%)
Aligned	No alert	Incongruent	481 (56)	2.6 (3.4)
		Congruent	435 (53)	0.8 (1.5)
	Alert	Incongruent	465 (52)	2.4 (3.6)
		Congruent	411 (49)	0.5 (1.2)
Misaligned	No alert	Incongruent	463 (50)	3.1 (4.3)
		Congruent	432 (51)	0.8 (1.4)
	Alert	Incongruent	449 (49)	3.5 (4.2)
		Congruent	408 (48)	0.5 (1.1)

Note. Standard deviations appear in parentheses

stimuli (465 ms) than to congruent stimuli (422 ms), yielding a significant main effect of congruency. Two interactions were significant and both reflected modulation of congruency effects (see Fig. 4). First, there was an interaction between alignment and congruency, reflecting a smaller congruency effect for misaligned stimuli (36 ms) than for aligned stimuli (50 ms), $BF_{alt} = 9.1 \times 10^{12}$. Second, there was an interaction between alerting and congruency, reflecting a larger congruency effect on alert trials (48 ms) than on no-alert trials (39 ms), $BF_{alt} = 2.8 \times 10^6$. The three-way interaction of alignment, alerting, and congruency was nonsignificant, reflecting similar increases in congruency effects from no-alert to alert trials for aligned stimuli (8 ms) and for misaligned stimuli (10 ms), $BF_{null} = 9.1$.

Error rates Responses were incorrect more often to misaligned stimuli (2.0%) than to aligned stimuli (1.6%), yielding a significant main effect of alignment. Responses were incorrect more often to incongruent stimuli (2.9%) than to congruent stimuli (0.7%), yielding a significant main effect of congruency. One interaction was significant and it reflected modulation of congruency effects (see Fig. 4). There was an interaction between alignment and congruency, reflecting a larger congruency effect for misaligned stimuli (2.7%) than for aligned stimuli (1.8%), $BF_{alt} = 125.6$. The three-way interaction of alignment, alerting, and congruency was nonsignificant, although the increase in congruency effects from no-alert to alert trials was numerically smaller for aligned stimuli (0.1%) than for misaligned stimuli (0.7%), $BF_{null} = 7.0$.

Discussion

The results of Experiment 3 replicate some of the key results of Experiments 1 and 2 (compare Figs. 2–4) and do not match the prediction derived from the spatial grouping hypothesis. The predicted three-way interaction between alignment, alerting, and congruency was nonsignificant for RTs and error rates, supported by Bayes-factor analyses favoring null interactions. Moreover, the data pattern was numerically in the

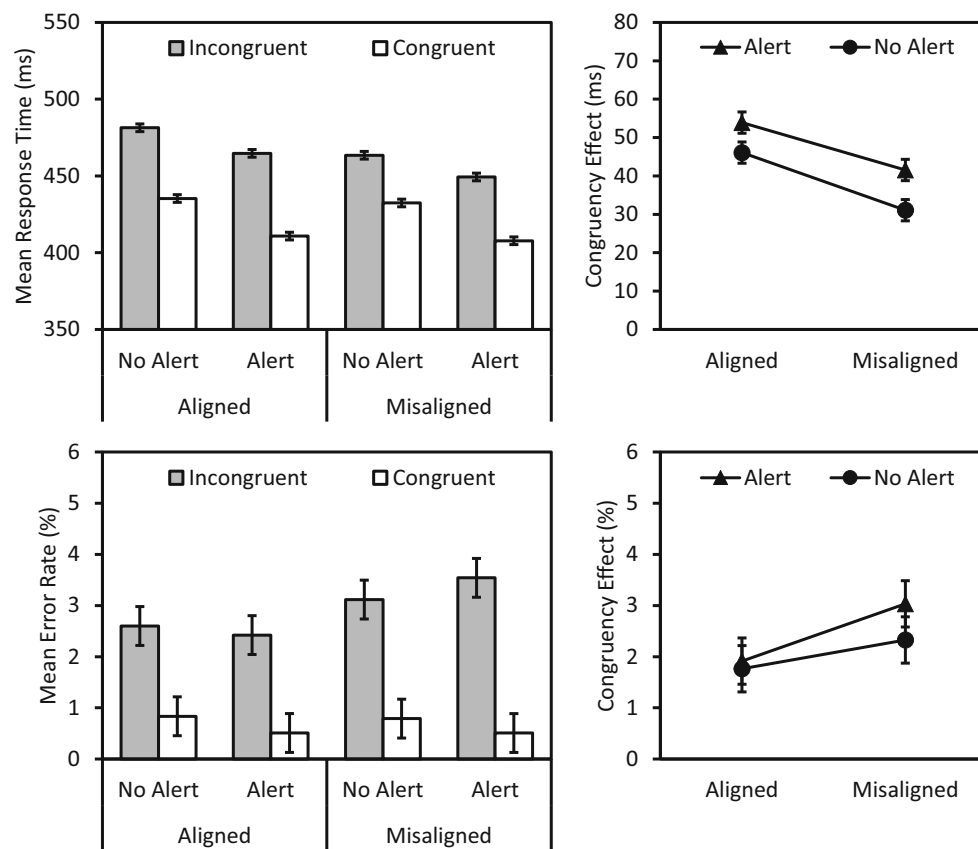


Fig. 4 Experiment 3 data. Mean response times and mean error rates appear in the top-left and bottom-left panels, respectively, as a function of alignment, alerting, and congruency. Congruency effects (performance differences between incongruent and congruent trials) for response times

and error rates appear in the top-right and bottom-right panels, respectively, as a function of alignment and alerting. Error bars represent 95% confidence intervals (Masson & Loftus, 2003)

opposite direction, with the difference in congruency effects between alert and no-alert trials being numerically larger—not smaller—for misaligned stimuli than for aligned stimuli (see Fig. 4).

Recall that the rationale for doubling the amount of misalignment in Experiment 3 was that it should produce stronger alignment effects than those obtained in previous experiments.

The standardized effect size for the interaction between alignment and congruency on RTs was indeed larger in Experiment 3 ($\eta_p^2 = .356$) than in Experiments 1 and 2 ($\eta_p^2 = .180$ and $.073$, respectively), suggesting that I succeeded in making the alignment manipulation more potent. I acknowledge that the alignment manipulation in Experiment 3 still did not yield effect sizes as large as those produced by the spacing

Table 6. Summary of analyses of variance for Experiment 3

Effect	Response time				Error rate			
	<i>F</i> (1, 149)	<i>MSE</i>	<i>p</i>	η_p^2	<i>F</i> (1, 149)	<i>MSE</i>	<i>p</i>	η_p^2
Alignment (L)	134.28	215	<.001	.474	14.79	3	<.001	.090
Alerting (A)	366.68	326	<.001	.711	.50	5	.480	.003
Congruency (C)	958.26	583	<.001	.865	94.65	16	<.001	.388
L × A	.83	131	.363	.006	2.51	3	.115	.017
L × C	82.36	170	<.001	.356	16.04	3	<.001	.097
A × C	39.94	157	<.001	.211	2.75	5	.100	.018
L × A × C	1.08	122	.300	.007	1.60	4	.208	.011

Note. Bolded numbers indicate effects that were significant ($p < .05$) in all experiments

manipulation in Schneider (2018b). Nevertheless, the combination of a strengthened alignment manipulation and an enlarged sample size should have improved detection of any three-way interaction between alignment, alerting, and congruency in Experiment 3. The continued lack of evidence of a three-way interaction in the predicted direction for both RTs and error rates bolsters the case against the spatial grouping hypothesis.

General discussion

Alertness seems to influence selective attention processes underlying cognitive control in the flanker task, as indicated by previous findings of larger congruency effects on alert trials than on no-alert trials (e.g., Fan et al., 2002; MacLeod et al., 2010; Redick & Engle, 2006; Schneider, 2018a, 2018b). Weinbach and Henik (2011, 2012) suggested that increased alertness yields a bias toward global processing of stimuli, and Schneider (2018b) proposed spatial grouping as a mechanism for a global processing bias. In the present study, I tested the spatial grouping hypothesis by manipulating the spatial alignment (collinearity) of the target and distractors in a variant of the ANT. I was interested in whether the predicted three-way interaction between alignment, alerting, and congruency derived from the hypothesis would receive empirical support.

I conducted three experiments that yielded similar data patterns that did not match the prediction from the spatial grouping hypothesis. Instead of the predicted pattern—a smaller difference in congruency effects between alert and no-alert trials for misaligned stimuli than for aligned stimuli—I obtained the numerically opposite pattern in all the experiments (see Figs. 2–4), which was significant only for error rates in Experiment 2. Bayes-factor analyses favored null three-way interactions, indicating that alignment does not modulate the alerting–congruency interaction. These results are difficult to reconcile with the spatial grouping hypothesis.

The only consistently reliable effect involving alignment (see Tables 2, 4, and 6) was the finding of smaller congruency effects on RTs for misaligned stimuli than for aligned stimuli. This result fits well with previous findings of smaller congruency effects from manipulations that disrupt perceptual grouping of the target and distractors in the flanker task (e.g., spacing: Eriksen & Eriksen, 1974; Miller, 1991; color grouping: Baylis & Driver, 1992; Kramer & Jacobson, 1991; shape grouping: Luo & Proctor, 2016). The alerting manipulation also had consistently reliable effects, yielding overall speedups in performance and larger congruency effects on RTs and error rates for alert trials than for no-alert trials, replicating previous findings (e.g., Fan et al., 2002; MacLeod et al., 2010; Redick & Engle, 2006; Schneider, 2018a, 2018b).

The present data lead me to conclude that spatial grouping is not a viable mechanism for explaining the alerting–congruency interaction in the flanker task (cf. Schneider, 2018b). This conclusion does not rule out Weinbach and Henik's (2011, 2012) hypothesis that increased alertness yields a bias toward global processing of stimuli. Instead, it motivates consideration of mechanisms other than spatial grouping for producing a global processing bias.

Weinbach and Henik (2012; see also McConnell & Shore, 2011; Nieuwenhuis & de Kleijn, 2013) suggested that greater alertness is associated with a more diffuse focus of attention, which would allow more distractors to be attended and processed, resulting in larger congruency effects. However, two lines of evidence are inconsistent with this idea. First, if attention is more diffuse on alert trials than on no-alert trials, then it should counteract the effects of spacing the distractors farther from the target in the flanker task, resulting in a larger alerting–congruency interaction for widely spaced stimuli than for narrowly spaced stimuli. Schneider (2018b) found the opposite data pattern when spacing was manipulated in two experiments. Second, it is known that probe stimulus detection is progressively slower as the probe appears farther from the focus of attention, reflecting an attentional gradient that results in a V-shaped probe-RT function (LaBerge, 1983). If attention is more diffuse on alert trials than on no-alert trials, then the probe-RT function should be flatter for the former than for the latter, but Seibold (2018) found no evidence of any flattening in a series of experiments. Thus, recent findings suggest that diffuse attention is unlikely to be the mechanism underlying any global processing bias on alert trials.

Moreover, Weinbach and Henik (2014) provided evidence against the idea of a global processing bias. They reported an experiment in which the stimuli were large arrows composed of several small arrows, and the left/right directions of the large and small arrows were either congruent or incongruent. In separate blocks, subjects identified the direction of either the large arrow (global task) or the small arrows (local task), with the stimuli preceded by alerting cues on 50% of the trials. A key manipulation was whether the relevant or irrelevant stimulus dimension was made more salient by blurring the small arrows (to increase global salience) or presenting the small arrows in different colors (to increase local salience). For each task, a typical alerting–congruency interaction was obtained when the irrelevant dimension was salient, whereas the reverse interaction was obtained (and congruency effects were much smaller overall) when the relevant dimension was salient. These results led Weinbach and Henik to hypothesize that increased alertness is associated with greater sensitivity to salient information, giving rise to a typical alerting–congruency interaction when distractors are salient.

The salience hypothesis can accommodate Schneider's (2018b) finding of no alerting–congruency interaction under wide spacing of stimuli in the flanker task because one could

argue that the distractors became less salient when they were spaced farther from the target.⁷ However, the hypothesis seems inconsistent with Schneider's (2018b) finding in three experiments that presenting target and distractor arrows in different colors—which should have made the distractors more salient—reduced congruency effects without modulating the alerting–congruency interaction, which remained as large as when the target and distractor arrows were the same color. The hypothesis also seems inconsistent with the results of the present study. One could argue that misaligning the target and distractor arrows made the distractors more salient because they were collinear with each other and not with the target. However, misalignment reduced the congruency effects on RTs without modulating the alerting–congruency interaction. Thus, it is unclear whether the salience hypothesis can account for some recent findings, although it is important to note that its application to different stimulus manipulations depends on how salience is defined.

Alternative mechanisms and hypotheses relating alertness to cognitive control also face difficulties in accounting for the available data. Callejas and colleagues (Callejas, Lupiáñez, Funes, & Tudela, 2005; Callejas, Lupiáñez, & Tudela, 2004) hypothesized that cognitive control is inhibited when alertness is increased, but their inhibition hypothesis does not explain why the alerting–congruency interaction is not always found in the flanker task (see the wide-spacing data in Schneider, 2018b) or in some other tasks generally considered to involve cognitive control, such as the color-word Stroop task (Schneider, 2019; Soutschek et al., 2013; Weinbach & Henik, 2012). Fischer et al. (2010, 2012) hypothesized that increased alertness facilitates a direct route for response selection that involves well-learned stimulus–response associations, but this hypothesis also cannot accommodate null alerting–congruency interactions in the color-word Stroop task, where automatic processing of the irrelevant word should be facilitated when alertness is increased. According to the early onset hypothesis of Nieuwenhuis and de Kleijn (2013), stimulus encoding is faster when subjects are more alert, which then allows response selection to start while cognitive control is weak because insufficient time has elapsed for focusing attention. The link between alertness and stimulus encoding is supported by empirical evidence (e.g., Rolke & Hofmann, 2007; Seifried, Ulrich, Bausenhart, Rolke, & Osman, 2010) and computational modeling work (Bundesen, Vangkilde, & Habekost, 2015; Matthias et al., 2010; Petersen, Petersen, Bundesen, Vangkilde, & Habekost, 2017), but the notion that cognitive control is weak because attention is unfocused implies that attention is more diffuse (at least initially) on alert trials than on no-alert trials, and I have already summarized evidence that does not support that idea.

At present, there does not appear to be a clear answer to the question of what cognitive mechanism underlies the alerting–congruency interaction in the flanker task. However, by directly testing a hypothesis for the interaction, the present study contributes to the incremental progress that has been made in pruning the space of theoretical possibilities. As discussed earlier, recent studies by Seibold (2018) and Schneider (2018b) indicate that increased alertness does not seem to be associated with more diffuse attention, a conclusion that has implications for the global processing and early onset hypotheses (see also Weinbach & Henik, 2014). In the present study, I manipulated stimulus alignment and provided evidence that increased alertness does not seem to modulate spatial grouping of the target and distractors, ruling out a potential mechanism for a global processing bias. Future studies in this vein that test other mechanisms and hypotheses will likely provide additional clues for solving the mystery of how alertness is related to cognitive control.

Acknowledgements I thank Paige Asbridge, Haerim Chun, Timothy Durbin, Christine Grosso, Kelsey Lynch, Nicole Magiera, Yijie Peng, Analicia Rios, Sami Wagner, and Julia Woodruff for assistance with data collection. I also thank Sander Nieuwenhuis and an anonymous reviewer for their comments on a previous draft of the manuscript.

References

- Baylis, G. C., & Driver, J. (1992). Visual parsing and response competition: The effect of grouping factors. *Perception & Psychophysics*, 51, 145–162.
- Böckler, A., Alpay, G., & Stürmer, B. (2011). Accessory stimuli affect the emergence of conflict, not conflict control: A Simon-task ERP study. *Experimental Psychology*, 58, 102–109.
- Bundesen, C., Vangkilde, S., & Habekost, T. (2015). Components of visual bias: A multiplicative hypothesis. *Annals of the New York Academy of Sciences*, 1339, 116–124.
- Callejas, A., Lupiáñez, J., Funes, M. J., & Tudela, P. (2005). Modulations among the alerting, orienting, and executive control networks. *Experimental Brain Research*, 167, 27–37.
- Callejas, A., Lupiáñez, J., & Tudela, P. (2004). The three attentional networks: On their independence and interactions. *Brain and Cognition*, 54, 225–227.
- Cohen, J. D., Servan-Schreiber, D., & McClelland, J. L. (1992). A parallel distributed processing approach to automaticity. *American Journal of Psychology*, 105, 239–269.
- Eriksen, B. A., & Eriksen, C. W. (1974). Effects of noise letters upon the identification of a target letter in a nonsearch task. *Perception & Psychophysics*, 16, 143–149.
- Fan, J., McCandliss, B. D., Sommer, T., Raz, A., & Posner, M. I. (2002). Testing the efficiency and independence of attentional networks. *Journal of Cognitive Neuroscience*, 14, 340–347.
- Fischer, R., Plessow, F., & Kiesel, A. (2010). Auditory warning signals affect mechanisms of response selection: Evidence from a Simon task. *Experimental Psychology*, 57, 89–97.
- Fischer, R., Plessow, F., & Kiesel, A. (2012). The effects of alerting signals in action control: Activation of S–R associations or inhibition of executive control processes? *Psychological Research*, 76, 317–328.

⁷ I thank an anonymous reviewer for suggesting this interpretation.

- Gratton, G., Coles, M. G. H., & Donchin, E. (1992). Optimizing the use of information: Strategic control of activation of responses. *Journal of Experimental Psychology: General*, 121, 480–506.
- Hartmann, G. W. (1935). *Gestalt psychology: A survey of facts and principles*. New York: Ronald Press.
- Heitz, R. P., & Engle, R. W. (2007). Focusing the spotlight: Individual differences in visual attention control. *Journal of Experimental Psychology: General*, 136, 217–240.
- Held, L., & Ott, M. (2018). On *p*-values and Bayes factors. *Annual Review of Statistics and Its Application*, 5, 393–419.
- Klein, R. M., & Ivano, J. (2011). The components of visual attention and the ubiquitous Simon effect. *Acta Psychologica*, 136, 225–234.
- Kramer, A. F., & Jacobson, A. (1991). Perceptual organization and focused attention: The role of objects and proximity in visual processing. *Perception & Psychophysics*, 50, 267–284.
- LaBerge, D. (1983). Spatial extent of attention to letters and words. *Journal of Experimental Psychology: Human Perception and Performance*, 9, 371–379.
- Logan, G. D. (1996). The CODE theory of visual attention: An integration of space-based and object-based attention. *Psychological Review*, 103, 603–649.
- Lu, C.-H., & Proctor, R. W. (1995). The influence of irrelevant location information on performance: A review of the Simon and spatial Stroop effects. *Psychonomic Bulletin & Review*, 2, 174–207.
- Luo, C., & Proctor, R. W. (2016). Perceptual grouping of objects occupied by target and flankers affects target-flanker interference. *Attention, Perception, & Psychophysics*, 78, 251–263.
- MacLeod, J. W., Lawrence, M. A., McConnell, M. M., Eskes, G. A., Klein, R. M., & Shore, D. I. (2010). Appraising the ANT: Psychometric and theoretical considerations of the Attention Network Test. *Neuropsychology*, 24, 637–651.
- Masson, M. E. J., & Loftus, G. R. (2003). Using confidence intervals for graphically based data interpretation. *Canadian Journal of Experimental Psychology*, 57, 203–220.
- Matthias, E., Bublak, P., Müller, H. J., Schneider, W. X., Krummenacher, J., & Finke, K. (2010). The influence of alertness on spatial and nonspatial components of visual attention. *Journal of Experimental Psychology: Human Perception and Performance*, 36, 38–56.
- McConnell, M. M., & Shore, D. I. (2011). Mixing measures: Testing an assumption of the attention network test. *Attention, Perception, & Psychophysics*, 73, 1096–1107.
- Miller, J. (1991). The flanker compatibility effect as a function of visual angle, attentional focus, visual transients, and perceptual load: A search for boundary conditions. *Perception & Psychophysics*, 49, 270–288.
- Nieuwenhuis, S., & de Kleijn, R. (2013). The impact of alertness on cognitive control. *Journal of Experimental Psychology: Human Perception and Performance*, 39, 1797–1801.
- Palmer, S. E. (1999). *Vision science: Photons to phenomenology*. Cambridge: MIT Press.
- Petersen, A., Petersen, A. H., Bundesen, C., Vangkilde, S., & Habekost, T. (2017). The effect of phasic auditory alerting on visual perception. *Cognition*, 165, 73–81.
- Pomerantz, J. R., & Portillo, M. C. (2011). Grouping and emergent features in vision: Toward a theory of basic Gestalts. *Journal of Experimental Psychology: Human Perception and Performance*, 37, 1331–1349.
- Posner, M. I. (2008). Measuring alertness. *Annals of the New York Academy of Sciences*, 1129, 193–199.
- Posner, M. I., & Boies, S. J. (1971). Components of attention. *Psychological Review*, 78, 391–408.
- Redick, T. S., & Engle, R. W. (2006). Working memory capacity and attention network test performance. *Applied Cognitive Psychology*, 20, 713–721.
- Rolke, B., & Hofmann, P. (2007). Temporal uncertainty degrades perceptual processing. *Psychonomic Bulletin & Review*, 14, 522–526.
- Rouder, J. N., Speckman, P. L., Sun, D., Morey, R. D., & Iverson, G. (2009). Bayesian *t* tests for accepting and rejecting the null hypothesis. *Psychonomic Bulletin & Review*, 16, 225–237.
- Schneider, D. W. (2018a). Alertness and cognitive control: Testing the early onset hypothesis. *Journal of Experimental Psychology: Human Perception and Performance*, 44, 756–766.
- Schneider, D. W. (2018b). Alertness and cognitive control: Toward a spatial grouping hypothesis. *Attention, Perception, & Psychophysics*, 80, 913–928.
- Schneider, D. W. (2019). Alertness and cognitive control: Is there a spatial attention constraint? *Attention, Perception, & Psychophysics*, 81, 119–136.
- Schönbrodt, F. D., & Wagenmakers, E.-J. (2018). Bayes factor design analysis: Planning for compelling evidence. *Psychonomic Bulletin & Review*, 25, 128–142.
- Seibold, V. C. (2018). Do alerting signals increase the size of the attentional focus? *Attention, Perception, & Psychophysics*, 80, 402–425.
- Seifried, T., Ulrich, R., Bausenhardt, K. M., Rolke, B., & Osman, A. (2010). Temporal preparation decreases perceptual latency: Evidence from a clock paradigm. *Quarterly Journal of Experimental Psychology*, 63, 2432–2451.
- Soutschek, A., Müller, H. J., & Schubert, T. (2013). Conflict-specific effects of accessory stimuli on cognitive control in the Stroop task and the Simon task. *Experimental Psychology*, 60, 140–147.
- Weinbach, N., & Henik, A. (2011). Phasic alertness can modulate executive control by enhancing global processing of visual stimuli. *Cognition*, 121, 454–458.
- Weinbach, N., & Henik, A. (2012). The relationship between alertness and executive control. *Journal of Experimental Psychology: Human Perception and Performance*, 38, 1530–1540.
- Weinbach, N., & Henik, A. (2014). Alerting enhances attentional bias for salient stimuli: Evidence from a global/local processing task. *Cognition*, 133, 414–419.
- White, C. N., Ratcliff, R., & Starns, J. J. (2011). Diffusion models of the flanker task: Discrete versus gradual attentional selection. *Cognitive Psychology*, 63, 210–238.

Open practices statement The data and materials for all experiments are available upon request from the author. Experiments 1 and 3 were preregistered (see footnote 6)..

Publisher's note Springer Nature remains neutral with regard to jurisdictional claims in published maps and institutional affiliations.