MEASUREMENT AND INTERPRETATION OF ATTENTIONAL BIAS

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MEASUREMENT AND INTERPRETATION OF ATTENTIONAL BIAS

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Is Sham Training Still Training? An Alternative Control Group for Attentional Bias Modification

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Keywords: control group, appetitive stimuli, modification, attentional bias, sham training

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The tendency to selectively attend to environmental stimuli congruent with self-relevant concerns has been documented across a wide range of mental and physical health domains. In particular, such attentional biases have now been demonstrated for a number of appetitive and/or addictive substances, including cigarettes among smokers (e.g., Waters et al., 2003), alcohol in heavy drinkers (e.g., Townshend and Duka, 2001) and high-calorie food in obese individuals (e.g., Kemps et al., 2014). The most common way to demonstrate attentional bias is implicitly, via the dot probe task (Posner et al., 1980), in one version of which pairs of words or pictures are presented briefly, followed by a small dot in the spatial location of one of the stimuli. The participant's task is simply to determine the location of the dot probe as quickly as possible. When the pairs consist of one self-relevant stimulus (e.g., a picture of beer) and one neutral stimulus (e.g., a picture of a glass of water), attentional bias is demonstrated by speeded detection of probes replacing the self-relevant stimulus relative to the neutral stimulus.

More recently, research has extended the protocol initially used to successfully modify attentional bias for threat-related stimuli in anxiety (MacLeod et al., 2002) to addictive and craved substances. In this modified dot probe task, a contingency is introduced whereby the dot probe appears disproportionately (90–100%) in place of the neutral word or picture, thereby training attention away from the substance-relevant cue. The central idea is that over time the repeated practice of responding to probes in the spatial location of the neutral cue induces a shift of attention (as indicated by relative response latencies) away from the substance-relevant cue and toward the neutral cue. This is seen as an implicit and gradual process (MacLeod et al., 2002; Koster et al., 2009; Kemps et al., 2014), although there is still considerable uncertainty as to the precise mechanisms mediating the effect (Heeren et al., 2013). A number of reviews and meta-analyses have now shown some effectiveness for attentional bias modification in appetitive domains, but effects are small and conclusions limited by methodological weaknesses (Beard et al., 2012; Turton et al., 2016; Jones et al., 2018; Boffo et al., 2019). In addition, some reviews have questioned the clinical utility of attentional bias modification, concluding that there is insufficient evidence that positive effects on attentional bias translate into any effect on addiction outcomes (Christiansen et al., 2015; Cristea et al., 2016).

One identified limitation lies in the use of different and suboptimal control conditions (Turton et al., 2016; Jones et al., 2018; Boffo et al., 2019). Following MacLeod et al.'s protocol, some studies have implemented comparison conditions where the contingency is reversed, i.e., the dot probe disproportionately replaces the substance-relevant stimulus (training attention toward the substance-relevant cue). Such a protocol both maximizes observed differences between "attend" and "avoid" conditions and means that they cannot be attributed unambiguously to a reduction in attentional bias in the "avoid" group. In addition, an "attend" condition is often not viable in studies of addictive substances for ethical reasons. Other studies have used no training (treatment as usual), wait-list, or unrelated tasks as control conditions. Over time, however, researchers have settled on what has come to be called "sham training" as the optimal control. In sham training, the dot probe replaces the substance-relevant stimuli and neutral stimuli with equal frequency (50/50), meaning that attention is not particularly directed to either. Sham training has become the "gold standard" control condition because it is so well-matched to the attentional re-training experimental condition in both stimulus exposure and response requirements; only the contingency is different. Perhaps not surprisingly, fewer significant effects of attentional bias modification are observed when contrasted against the more stringent sham training than against other control conditions (Beard et al., 2012).

While widely accepted as the best available control condition, there have recently been some questions raised about the nature of sham training. In particular, a pattern noted across multiple domains is that the lack of significant difference between experimental and control groups often comes about because both groups improve (Cristea et al., 2015). This seems especially the case with clinical samples (Blackwell et al., 2017). Interestingly, a similar pattern has been noted in studies of approach bias modification (Kakoschke et al., 2018). This observation has led some researchers to suggest that sham training may in fact have an active component, rather than offering, as assumed, an "inert," "neutral," or "placebo" training, that should not of itself change cognitive biases in any specific direction. Cristea et al. (2015) suggest the operation of various demand characteristics, while Boffo et al. (2019) suggest a more general exposure or desensitization process, whereby continued exposure to substance-related stimuli (irrespective of any contingency or response) may result in participants becoming less sensitive over time to the motivational meaning of the substance-related cue.

When used as a control condition, sham training is typically characterized as "no contingency" training. However, in the dot probe (and other tasks), contingency is a continuous variable, running from 100/0 (dot replaces neutral stimulus 100% of the time) to 0/100 (dot replaces substance-relevant stimulus 100% of the time). Thus, when probes replace the neutral and substance-related stimuli with equal frequency, this is better described as a 50/50 (rather than "no") contingency. Accordingly, a few authors have suggested that sham training actively trains equal attention to substance-relevant and neutral stimuli and thereby may affect control over attention for substance-related stimuli (e.g., Schoenmakers et al., 2010; Badura-Brack et al., 2015; Khanna et al., 2016). Others have suggested that sham training serves to train participants to ignore emotional stimuli when confronted with them (Gladwin, 2017; Gladwin et al., 2019). In line with these ideas, the sham training protocol has sometimes been reconceptualized and renamed as "attentional control training," and viewed as a more top-down goal-directed process (Gladwin, 2017). In combat veterans, such attentional control training has been shown to be more effective for PTSD symptom reduction than traditional attentional bias modification (Badura-Brack et al., 2015; Khanna et al., 2016). These latter authors view attentional control training as particularly effective at normalizing attention allocation, although the exact cognitive mechanism(s) underlying the effect of exposure to a 50/50 contingency (conceptualized either as sham training or attentional control training) is yet to be clarified.

To the extent that exposure to the 50/50 contingency can be seen as balancing attentional allocation between relevant and neutral stimuli (Badura-Brack et al., 2015), the size of effect of sham training will logically depend upon an individual's initial level of attentional bias for the target stimulus category, be it alcohol, cigarettes, or food. To elaborate, for people who have no initial attentional bias (approximately equal attention to substance-relevant and neutral stimuli), sham training should produce little change. But for individuals with a strong initial attentional bias toward a particular substance, the gradual training of equal attention to substance-relevant and neutral stimuli involved in sham training represents a substantial shift in relative attention away from the substance-relevant stimuli (where the majority of their initial attention was directed). Thus, sham training with its 50/50 contingency will serve to decrease attentional bias for these individuals. This account is able to explain the observation that sham training tends to lead to greater improvements in clinical samples (Blackwell et al., 2017). This trend, in turn, accounts for the general conclusion that attentional bias modification is more successful (when sham training is the control) in unselected or analog samples than it is in clinical samples (Cristea et al., 2015).

Of course, it is difficult to conceptualize a better alternative control task than sham training. Schoenmakers et al. (2010) developed a novel categorization task that avoided the 50/50 contingency. Ideally, what is required is a neutral control protocol that does not manipulate contingencies between stimulus categories and responses (a truly "no contingency" training condition), but otherwise matches the stimulus exposure and response requirements of attentional bias retraining. We have attempted to devise such a protocol and put it forward here for scrutiny. In our version of sham training, which we call sham-n training ("n" for "neutral," or "no contingency"), instead of stimulus pairs consisting of one substance-relevant and one neutral picture, they are constructed to be of either two substance-relevant pictures or two neutral pictures. On half the trials, participants are presented with two substance-relevant pictures, and on the other half of trials with two neutral pictures (in random order and appropriately counter-balanced). As before, the probe is set to replace the left and right pictures with equal frequency (50/50) and participants need to determine the location of the probe. Thus, sham-n training directs attention equally to stimuli within a category, but not across categories. In a nutshell, participants receive the same number of trials and amount of stimulus exposure and are required to make the same judgements (probe location) and associated motor responses in sham-n training as in the original task. But in sham-n training, there is absolutely no relation between stimulus category and response.

The critical test of the sham-n training protocol is that, as befits an inert control condition, it should not lead to any change in attentional bias. In our first trial with an undergraduate sample, we found that this control condition resulted in no change in attentional bias for chocolate. The mean of 6.64 after sham-n training was very similar to the mean of 6.61 before sham-n training. This contrasted with results for the "avoid chocolate" group who experienced a significant reduction in attentional bias for chocolate (and "attend chocolate" group who experienced a significant increase). We also have preliminary data from a small field study of people trying to lose weight, where sham-n training (administered multiple times via smartphone app) produced no significant change in attentional bias for unhealthy foods, whereas the experimental training did. Although these results would carry more weight if they were contrasted with traditional sham training or some other control condition, they can be taken as preliminary proof-of-concept for our task. An important next step would be to test the protocol with individuals who demonstrate elevated levels of attentional bias toward any substance, such as clinical samples.

In sum, we present sham-n training as a potential control protocol in studies of attentional bias modification. The same logic could potentially be applied to developing control protocols for the modification of other cognitive biases (e.g., approach and interpretation biases). As pointed out by Blackwell et al. (2017), control conditions rarely generate the interest or excitement of active training conditions but are nevertheless critically important to the interpretation and value of their results. We welcome feedback on the sham-n training protocol and, of course, we welcome other investigators trialing it in their own settings.

AUTHOR CONTRIBUTIONS

MT and EK conceptualized all aspects of the opinion piece together. All authors contributed to the article and approved the submitted version.

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A Promising Candidate to Reliably Index Attentional Bias Toward Alcohol Cues–An Adapted Odd-One-Out Visual Search Task

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Attentional bias (AB) has been suggested to contribute to the persistence of substance use behavior. However, the empirical evidence for its proposed role in addiction is inconsistent. This might be due to the inability of commonly used measures to differentiate between attentional engagement and attentional disengagement. Attesting to the importance of differentiating between both components of AB, a recent study using the odd-one-out task (OOOT) showed that substance use was differentially related to engagement and disengagement bias. However, the AB measures derived from the OOOT showed insufficient reliability to be used as a solid measure of individual differences. Therefore, the current study aimed to improve the reliability of the AB measures derived from the OOOT by using more distinct contrast stimuli, adding practice trials, increasing the number of trials, and by having participants perform the task in an alcohol-relevant context. We contrasted the original OOOT with the adapted OOOT (i.e., OOOT-adapt) and assessed AB in low- and high-drinking individuals. Participants were 245 undergraduate students who typically tend to drink either low or high amounts of alcohol. In one condition, AB was measured with the original OOOT in a typical laboratory context, whereas in the other condition, AB was measured with the OOOT-adapt in a bar (i.e., alcohol-relevant) context. The OOOT-adapt showed superior internal consistency, especially for the high-drinking group. Further, specifically the OOOT-adapt differentiated between low- and high-drinking participants showing that high drinkers engaged faster with alcohol cues than did low drinkers. Thus, the OOOT-adapt was found to be a promising candidate to reliably index AB in the context of alcohol use. The OOOT-adapt further showed superior criterion validity as it could differentiate between low- and high-drinking individuals, thereby adding to the evidence that AB might be involved in substance use behavior.

Keywords: attentional bias, alcohol use, addiction, reliability, internal consistency, visual search

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INTRODUCTION

Dual process models of addiction attribute an important role to automatic processes when explaining the development and persistence of addiction (Wiers et al., 2007; Stacy and Wiers, 2010). One of these processes is biased selective attention, also referred to as attentional bias (AB). AB can be expressed by a relatively strong tendency to direct attention to substance-relevant cues in the environment (i.e., engagement bias) and/or by a difficulty to redirect attention away from these cues (i.e., disengagement bias; Posner, 1980; Posner and Petersen, 1990). Although, in general, the contributing role of AB to the persistence of addictive behavior has been extensively studied throughout the past 15 years, only little is known about the specific role of engagement and disengagement bias. Directly distinguishing between engagement and disengagement bias might not only help to improve the general understanding of the disorder, but might also deliver knowledge that can be used to improve treatment (see, for example, Rinck et al., 2005; Hollitt et al., 2010).

One important reason for the limited knowledge about the role of engagement and disengagement bias relates to the fact that most measures of AB, such as the visual probe task (MacLeod et al., 1986), the addiction Stroop task (Cox et al., 2006), the flicker-induced change blindness task (Jones et al., 2002), or more recently developed tasks (e.g., Pennington et al., 2020), are not configured to differentiate between these two underlying processes of attention (Field and Cox, 2008; Grafton and MacLeod, 2014). That is, these assessment tasks deliver one overall index for AB. There are studies using for example the visual probe task, which aimed to disentangle engagement and disengagement bias by the use of different stimulus presentation durations (i.e., brief durations to index engagement bias and longer durations to index disengagement bias; e.g., Bradley et al., 2003; Field et al., 2006; Noël et al., 2006). Although this approach provided relevant information about initial and maintained attention, it has been pointed out that the use of different stimulus presentation intervals in the visual probe task does not allow disentangling engagement and disengagement processes (Grafton and MacLeod, 2014). One task that is configured to deliver separate indices for engagement bias and disengagement bias is the so-called odd-one-out task (OOOT; Hansen and Hansen, 1988; Rinck et al., 2005), which has been successfully used in previous research including studies on anxiety (De Voogd et al., 2017), sexual pain disorders (Melles et al., 2016), and eating behavior (Jonker et al., 2019). In the OOOT, participants are presented with an array of multiple stimuli identifying whether these stimuli are from the same category of images or whether one stimulus is defiant from the others (i.e., an odd-one-out). The task includes trials in which (1) all images are either disorder-relevant or disorder-irrelevant; (2) a disorder-relevant image is presented among disorder-irrelevant distractors; (3) a disorder-irrelevant image is presented among disorder-relevant distractors; and (4) a disorder-irrelevant image is presented among disorder-irrelevant distractors. The last trial type allows calculating a baseline of how long it generally takes to identify an odd-one-out among distractors allowing to calculate separate indices for engagement and disengagement bias by contrasting the reaction time of this neutral trial type with the other two trial types including disorder-relevant images. That is, engagement bias is expressed by the difference between trials in which a disorder-relevant image is presented among disorder-irrelevant distractors and the neutral trial type, whereas disengagement bias is expressed by the difference between trials in which a disorder-irrelevant image is presented among disorder-relevant distractors and the neutral trial type.

First indication that the OOOT also seems useful in examining engagement and disengagement bias in the context of alcohol use comes from a previous study from our laboratory in which, in a student sample, it was found that the disengagement index of the OOOT, but not the engagement index, was related to alcohol consumption, meaning that consuming higher amounts of alcohol was related with more difficulty to disengage attention from alcohol cues (Heitmann et al., 2020). However, the robustness of these findings may be questioned as the results indicated unacceptably low internal consistency of the AB indices. Yet, especially when being used as a measure of individual differences, it is critical that indices of AB show adequate reliability (e.g., McNally, 2019), and the commonly found low reliability of popular AB measures has been highlighted as a major threat for progress within this field of research (Rodebaugh et al., 2016). Therefore, the current study was designed to take up the challenge to modify the OOOT in a way to reach an acceptable level of reliability.

There are several aspects that might explain the low internal consistency of the AB indices as calculated from the original OOOT in the previous study (Heitmann et al., 2020). That is why we made several improvements to the design in the current study. First, the OOOT was improved by using more distinct contrast categories. That is, in the previous study, the neutral contrast categories (i.e., soft drinks and flowerpots) of the OOOT might have been insufficiently distinct from the target stimuli (i.e., alcoholic drinks), as we found that participants tend to make a substantial number of mistakes when following the task instruction to indicate whether a trial included an odd-one-out. Other studies, using more distinct contrast categories that were visually as well as content-wise less similar to the target category, have reported lower error rates and better internal consistency (e.g., Jonker et al., 2019). Second, the OOOT was further improved by adding practice trials including feedback, as well as adding more trials of trial types that are crucial to compute the AB indices (i.e., trial types including an odd-one-out). This seemed relevant as, in the previous study, only a limited number of trials of the OOOT were available to compute the AB indices (i.e., due to its configuration and high error rate), and a sufficient number of trials are necessary to reliably measure AB (Ataya et al., 2012). Third, the current study assessed AB in an alcohol-relevant context, as it has been shown that contextual factors might influence the stability of AB indices (Field et al., 2014; Christiansen et al., 2015). Fourth, AB was assessed in two groups, namely, low-drinking participants (i.e., low-alcohol group; 1–7 standard units a week) and high-drinking participants (i.e., high alcohol group; at least 14 standard units of alcohol a week). Thereby, we could test whether AB measures are more stable when assessing individuals for which alcohol cues are relatively salient/motivationally relevant—more likely individuals who drink higher amounts of alcohol (Field and Christiansen, 2012). Given that the previous study included a student sample in which the amount of used alcohol varied from little to high, the task might not have measured the processes of interest as, at least for the participants drinking little alcohol, the alcohol cues might have been less motivationally relevant (i.e., no AB for alcohol cues; Heitmann et al., 2020).

To follow up on the previous study and to investigate whether the internal consistency could be improved by using more distinct non-alcohol contrast stimulus categories, adding practice trials and increasing the number of trials, by having participants perform the task in a relevant context, and by assessing AB in low- and high-drinking individuals, we compared the internal consistency of this new and improved task, called the OOOT-adapt, with the original OOOT. First, we hypothesized that the OOOT-adapt would show better internal consistency than the OOOT, which would be especially evident in the high-alcohol group. Second, we expected students in the high-alcohol group to show stronger AB to alcohol cues than students in the low-alcohol group. And finally, we hypothesized that if indeed internal consistency of the OOOT-adapt is superior compared to the internal consistency of the OOOT, the difference between the low- and high-alcohol group would be more pronounced when AB was measured with the OOOT-adapt.

MATERIALS AND METHODS

This study was preregistered with OSF and can be accessed via the following link[1].

Participants

Participants signed up for the study via an online participant platform. There were two advertisements on this platform, one recruiting individuals who drink low amounts of alcohol (low-alcohol group; 1–7 units per week) and one recruiting individuals who drink high amounts of alcohol (high-alcohol group; 14 units or more per week). Based on power analyses on the main analyses, a medium effect size of 0.6, power of 95%, and an α level of 0.05, we aimed for a sample size of 122 participants in each group. This was in line with previous studies showing a medium effect size when differentiating between groups using an AB task (e.g., Grafton and MacLeod, 2014). Eventually, 245 undergraduate students (46% male, $mean_{age} = 20.3$, $SD_{age} = 2.08$) from the psychology bachelor program of the University of Groningen participated in the study.

Materials

Alcohol Use and Craving

The Measurements in Addiction for Triage and Evaluation Questionnaire (MATE-Q; Schippers and Broekman, 2014) was used to assess the quantity and frequency of alcohol use in the past 30 days, as well as craving for alcohol in the past 7 days. Quantity of use was indexed by summing the amount of standard glasses of alcohol consumed on a typical Monday, Tuesday, etc. This sum score was then multiplied by four to represent the amount of alcohol consumed in a typical month. Frequency of use was indexed by the question: "How often in the last 30 days have you used alcohol?" Alcohol craving was indexed by the Obsessive–Compulsive Drinking Scale (OCDS5) of the MATE-Q. The OCDS5 consists of five items measuring the desire for alcohol in the past 7 days, answered on a 5-point Likert scale. Alcohol craving was calculated by the sum of all items. Internal consistency of the OCDS5 was poor (Cronbach α of 0.51). This seemed to be related to item 4 of this questionnaire (i.e., "How much of an effort do you make to resist these thoughts or try to disregard or turn your attention away from these thoughts as they enter your mind?"). In line with our previous study (Heitmann et al., 2020), this item was therefore excluded, resulting in an acceptable internal consistency of the sum score of the remaining four items (Cronbach α of 0.70).

Alcohol Use Problems

Alcohol use–related problems were indexed with the shorted version of the Rutgers Alcohol Problem Index (RAPI-18; White and Labouvie, 1989). Participants had to indicate how often they experienced the 18 described situations in the past, using a 5-point Likert scale ranging from "never" (1) to "very often" (5). Per participant, a sum of scores was calculated. Internal consistency of the RAPI-18 was good (Cronbach α of 0.86).

Attentional Bias to Alcohol

Attentional bias to alcohol cues was measured with the original OOOT, as used in Heitmann et al. (2020), or the adapted version of the OOOT (OOOT-adapt). During the original OOOT, participants focused their attention on a red fixation cross in the center of the screen for 500 ms after which they had to indicate as quickly and correctly as possible whether there was an odd-one-out image within a 5 × 4 image matrix (500 × 500 pixels) by pressing the "0" (no odd-one-out) or "1" (yes, odd-one-out present) button on the keyboard. The task consisted of 54 trials with an odd-one-out and 18 trials without an odd-one-out (72 trials in total). The task was divided into three blocks of 24 trials. There were no practice trials in this task. The task consisted of three types of odd-one-out-present trials: alcohol target trials, with an alcohol odd-one-out and neutral (soft drinks or flower pots) distractors; alcohol distractors trial, with alcohol distractors and a soft drink or flower pot odd-one-out; neutral target in neutral distractors trial, with a soft drink odd-one-out in flower pot distractors; or a flower pot odd-one-out in soft drinks distractors. The three trial types without an odd-one-out consisted of either 20 alcohol images, 20 soft drink images, or 20 flower pot images. All trial types were randomly presented, and odd-one-out images randomly appeared over the possible positions, with the exception of directly above or below the fixation cross. Attentional engagement and attentional disengagement were inferred from trials in which an odd-one-out was present. Engagement bias was calculated by subtracting the mean response latency of alcohol target trials from the mean response latency of neutral target in neutral distractors trials. More attentional engagement with alcohol cues is then reflected in higher (more positive) scores. Disengagement bias was calculated by subtracting the mean response latency of neutral target in neutral distractors trials from the mean response latency of alcohol distractors trials. More difficulty to disengage

[1]https://osf.io/yfm25

attention from alcohol cues is reflected in higher positive scores. See **Figure 1** for an example of a trial from the OOOT.

The OOOT-adapt differed from the original OOOT in the following aspects: (1) the OOOT-adapt included at least 12 practice trials during which participants received feedback. If necessary, the number of practice trials was increased by one until a participant correctly responded to at least nine trials; (2) the OOOT-adapt consisted of 162 trials with 126 odd-one-out trials and 36 trials without an odd-one-out, divided into three blocks of 54 trials each; (3) the neutral distractors were images of office supplies and flowers instead of soft drinks and flower pots (see Heitmann et al., 2020). See **Figure 2** for an example of a trial of the OOOT-adapt.

Procedure

This study was approved by the ethical committee of the psychology department of the University of Groningen (PSY-1819-S-0081 and PSY-1819-S-0082). From the low-alcohol group and the high-alcohol group, half of the participants were assigned to the original OOOT and half to the OOOT-adapt. Participants were not aware that there were two different versions of the task (i.e., two conditions). On top of the adaptations to the task (see materials), also the location in which the OOOT-adapt was performed was different from the location in which the OOOT was performed. That is, the original version of the task was performed in a laboratory where assessment took place throughout the whole day similar to the study of Heitmann et al. (2020), whereas the adapted version of the task was performed in an alcohol-relevant context after 3 PM in the afternoon, i.e., a bar. To ensure that the location would not reveal the two different versions of the task, or bias the participants who would sign up, information about the location was given only 12 h prior to participants' appointment. At that time, the online participant platform no longer accepted switching time slot.

The procedures of both versions of the study were similar. On entry to the laboratory or bar, participants signed informed consent. Then they indicated their gender and age. They also reported on their state alcohol craving by answering how much they currently craved alcohol on a 7-point Likert scale ranging from "no craving" (1) to "a lot of craving" (7). Hereafter participants completed the OOOT or OOOT-adapt, followed by the MATE-Q and the RAPI-18. Given the difference in the

FIGURE 1 | Example trial of the OOOT—an alcohol distractors trial.

FIGURE 2 | Example trial of the OOOT-adapt—an alcohol distractors trial.

number of trials between the OOOT and the OOOT-adapt, participants completing the OOOT needed approximately 20 min to complete the study, whereas participants completing the OOOT-adapt needed approximately 30 min. All participants received course credits in return for their participation. For the first 16 participants, the RAPI-18 was erroneously not included in the study.

Analyses Plan

Data Reduction of OOOT and OOOT-Adapt

Data reduction was performed separately for both conditions (OOOT and OOOT-adapt) and both groups (low- and high-alcohol groups). Participants who fell more than three SDs below the mean accuracy of their condition and group were excluded. As a next step, trials with incorrect responses were deleted. Further, trials in which participants respond faster than 200 ms (i.e., expected anticipation errors) or fall more than three SDs below or above their mean response latency of that trial type were excluded.

Hypothesis 1

To examine whether the OOOT-adapt showed better internal consistency than the original version of the OOOT, internal consistency of the OOOT and the OOOT-adapt was calculated, per group (i.e., low- and high-alcohol groups), in two different ways: (1) a split-half Spearman–Brown coefficient was calculated from the outcomes of the tasks based on the trials of the first half and the second half of the tasks; and (2) a second method in which Spearman–Brown coefficients were calculated from outcome measures based on half of the trials where we distributed the trials alternately to one of two subsets. The first trial of one particular trial type was randomly allocated to either of the two subsets. Internal consistency was calculated for the engagement and disengagement indices of both tasks. The Fisher Z test was used to statistically compare the internal consistency coefficients of the engagement and disengagement indices as calculated from the OOOT and the OOOT-adapt.

Hypothesis 2

To examine whether students of the high-alcohol group showed a stronger AB to alcohol cues than students in the low-alcohol group, we performed one-tailed independent t tests comparing students drinking low amounts of alcohol, with students drinking high amounts of alcohol. We examined group differences for the OOOT and OOOT-adapt separately. Per condition, two independent t tests were performed, one on attentional engagement and one on attentional disengagement. Given multiple comparisons per group (engagement and disengagement bias), for the one-tailed independent t tests, we used an adjusted α of 0.025, to reduce the likelihood of incorrectly rejecting the null hypothesis (i.e., making a Type I error).

To increase confidence in our results delivered by the t tests following the frequentist approach, we also reported results following the Bayesian approach. Therefore, Bayesian independent-samples t tests with Cauchy priors were calculated, which are set at the recommended default $r = 0.707$. BF_{10}, which quantifies the evidence for the alternative hypotheses over the null hypotheses, was reported. A Bayes factor of 1 is considered no evidence, between 1 and 3 anecdotal, between 3 and 10 moderate, between 10 and 30 strong, between 30 and 100 very strong, and more than 100 extreme evidence that the data are in line with the alternative hypothesis. Conversely, a Bayes factor between 1/3 and 1 will be considered anecdotal; between 1/3 and 1/10, moderate evidence; between 1/10 and 1/30 strong evidence; between 1/30 and 1/100, very strong evidence; and less than 1/100, extremely strong evidence that the data are more likely under the null hypothesis (Wagenmakers et al., 2017).

Hypothesis 3

To examine if the difference between students in the low-alcohol group and high-alcohol group was more pronounced when AB was measured with the OOOT-adapt, we compared the confidence interval of the effect size comparing students who drink low vs. high amounts of alcohol derived from the OOOT-adapt with the confidence interval of the effect size derived from the OOOT.

RESULTS

Descriptives

Of the 245 participants signing up for the study, four reported no alcohol consumption in the past month and were therefore excluded from the study. All four belonged to the low-drinking group. Of the remaining 241 participants, 157 identified themselves as drinking low amounts of alcohol (1–7 units a week), and 84 as drinking high amounts of alcohol (>14 units a week). However, there seemed to be anomalies in these self-identified groups and the quantity of alcohol use reported during the study (**Table 1**). We therefore decided to test our hypotheses based on the self-identified groups, as well as on groups based on the quantity of alcohol use reported during the study, that is, a group of low drinkers (1–10 units a week) and a group of high drinkers (11 or more units a week). In the following, we will refer to these two approaches as self-identified groups and reported groups. The numbers of participants per group and per condition for both approaches as well as the group characteristics are provided in **Table 2**.

Data Reduction of OOOT and OOOT-Adapt

Participants who fell more than three SDs below the mean accuracy were excluded. In the OOOT-adapt, five participants (two in the low- and three in the high-alcohol group) were

TABLE 1 | Use per week per group based on self-identification.

Quantity	Low (n = 157)	High (n = 84)
1–7	77	0
8–10	32	3
11–13	14	6
≥14	34	75

Quantity, units of alcohol consumed in an average week in the past 30 days.

excluded for this reason. These numbers were identical for the data reduction based on self-identified and reported alcohol use. Mean percentage of correct responses after exclusion per group is reported in **Table 3**. Trials with incorrect responses were deleted. Further, trials in which participants responded faster than 200 ms (i.e., expected anticipation errors) or fell more than three SDs below or above the mean response latency of that trial type were excluded. For the OOOT, no too slow or too fast responses were found. Also for the OOOT-adapt, there were no too slow responses, but there were too fast responses. In the self-identified low-alcohol group, one response was faster than 200 ms; in the self-identified high-alcohol group, six; in the reported low-alcohol group, none; and in the reported high-alcohol group, seven responses were faster than 200 ms and therefore deleted.

Hypothesis 1: Does the OOOT-Adapt Have Better Internal Consistency Than the Original Version of the OOOT?

Internal consistency calculated with the split-half method and the alternating method and the related confidence intervals are reported in **Table 4**. The Fisher Z test was used to statistically compare the internal consistency coefficients of the OOOT and the OOOT-adapt. Internal consistency as calculated via the split-half method showed that the internal consistency of the OOOT-adapt was indeed higher than that of the OOOT. This was not consistently the case for the internal consistency as measured with the alternating method.

Hypothesis 2: Do Student Who Drink High Amounts of Alcohol Have a Stronger AB to Alcohol Cues Than Students Who Drink Low Amounts of Alcohol?

Mean AB scores and outcomes of the one-tailed independent t tests are reported in **Table 5**. Taking into account the adjusted α of 0.025, the results showed that only the OOOT-adapt was able to differentiate between the low- and high-alcohol group. Specifically, individuals in the high-alcohol group have more attentional engagement with alcohol cues than individuals in the low-alcohol group. This was the case when the groups were assigned based on self-identified alcohol use, as well as the reported amount of used alcohol. Bayes factors showed moderate to strong evidence that there are no differences between the groups on engagement and disengagement bias as measured with the OOOT or on the disengagement bias as measured with the OOOT-adapt.

TABLE 2 | Group characteristics.

Group			Frequency		Quantity		State craving		Craving		Problems	
			Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD
OOOT	Low	Self-identified (n = 96)	5.88	3.39	43.29	54.69	1.83	1.12	5.64	1.37	29.54	8.82
		Reported (n = 68)	4.93	2.84	22.94	10.50	1.71	1.05	5.37	1.27	26.93	6.64
	High	Self-identified (n = 31)	11.48	6.20	132.26	93.20	2.16	1.55	6.74	1.39	36.89	7.68
		Reported (n = 59)	9.93	5.33	113.49	89.43	2.15	1.40	6.53	1.41	36.38	8.94
OOOT-adapt	Low	Self-identified (n = 61)	6.82	3.95	41.25	30.80	1.79	1.10	5.69	1.38	29.24	7.52
		Reported (n = 44)	5.55	2.77	25.00	8.68	1.73	1.00	5.50	1.44	27.54	6.80
	High	Self-identified (n = 53)	12.04	5.44	117.66	57.23	2.94	1.63	7.13	1.96	37.69	9.08
		Reported (n = 70)	11.57	5.29	109.31	53.50	2.70	1.62	6.90	1.83	36.92	8.83

TABLE 3 | Percentage correct per group.

Group		OOOT		OOOT-adapt	
		Mean	SD	Mean	SD
Low	Self-identified	69.05	22.61	86.04	8.36
	Reported	69.98	21.83	86.90	8.14
High	Self-identified	70.93	23.29	85.62	9.15
	Reported	68.69	23.81	85.19	9.02

Hypothesis 3: Is the Difference Between the Low- and the High-Alcohol Group More Pronounced When AB Was Measured With the OOOT-Adapt When Compared With the OOOT?

It was originally planned to compare the confidence intervals of the analyses of the OOOT and the OOOT-adapt. However, given the findings, this became redundant. That is, these analyses were planned on the premises that the tasks would provide relatively similar outcomes and group differences, yet one might be more pronounced than the other. However, the OOOT gives a negative attentional engagement score, and the OOOT-adapt, a positive attentional engagement score. For the attentional disengagement scores, this is reversed, but also here the tasks provide very different outcomes. Furthermore, only the OOOT-adapt showed a significant difference between the low- and high-alcohol group.

DISCUSSION

The current study showed that using more distinct non-alcohol contrast categories, adding practice trials and increasing the number of trials, having participants perform the AB assessment task in an alcohol-relevant context, and assessing AB in high-drinking individuals resulted in increased internal consistency of the alcohol AB measure. The updated version of the task, called the OOOT-adapt, was also able to differentiate between participants who drank low amounts of alcohol and those who drank a high amount of alcohol.

In accordance with our first hypothesis, we found the internal consistency of the AB indices to be higher when measured with the OOOT-adapt than when measured with the original OOOT. This was especially true when the internal consistency of the tasks was calculated using the split-half method. When calculating the internal consistency with the alternating method, the internal consistency of the OOOT-adapt was significantly higher for the disengagement bias compared with the OOOT. Although the same tendency was evident for the engagement bias, the difference between the OOOT-adapt and OOOT did not reach significance. Similar results in which the split-half method revealed higher internal consistency were found in a previous study (Jonker et al., 2019). One explanation for this apparently consistent difference between both ways of allocating trials to one or the other half could be that the split-half method is less sensitive to variable carryover effects of individual trials and reflects therefore a more stable reflection of the process of interest. In addition, the findings indicated that the internal consistency of the OOOT-adapt was most favorable in the group of participants who drank high amounts of alcohol. This is in line with the idea that AB measures are more stable in individuals where the salience/motivational relevance of the cues is higher (Field and Christiansen, 2012), generally individuals who drink more frequent and higher amounts. Based on the current findings, one can expect the reliability of the OOOT-adapt to be even better when assessing AB in a clinical sample. Therefore, the reliability of the OOOT-adapt might further be tested in future research including treatment-seeking individuals diagnosed with substance use disorder. Especially, as the current sample was restricted to a homogenous sample of university students, it seems important to test the generalizability of results in the clinical range.

TABLE 4 | Internal consistency (Spearman–Brown).

			Low				High			
			OOOT	**OOOT-adapt**	***Z***	***P***	**OOOT**	**OOOT-adapt**	***Z***	***p***
Self-identified	Eng	Split-half	0.07 (−0.14; 0.28)	0.40 (0.16; 0.60)	2.113	0.017	−0.21 (−0.54; 0.17)	0.44 (0.18; 0.64)	2.904	0.002
		Alternating	0.19 (−0.02; 0.38)	0.24 (−0.02; 0.47)	0.313	0.377	0.48 (0.14; 0.72)	0.61 (0.40; 0.76)	0.788	0.215
	Dis	Split-half	−0.15 (−0.35; 0.06)	0.26 (0.00; 0.48)	2.494	0.006	0.19 (−0.19; 0.52)	0.66 (0.47; 0.79)	2.544	0.005
		Alternating	0.01 (−0.20; 0.22)	0.47 (0.24; 0.65)	2.989	0.001	0.53 (0.20; 0.75)	0.74 (0.58; 0.84)	1.866	0.031
Reported	Eng	Split-half	0.01 (−0.24; 0.26)	0.35 (0.05; 0.59)	1.782	0.037	−0.06 (−0.32; 0.20)	0.47 (0.26; 0.64)	3.149	0.001
		Alternating	0.16 (−0.09; 0.39)	0.22 (−0.09; 0.49)	0.312	0.377	0.48 (0.25; 0.66)	0.57 (0.38; 0.71)	0.688	0.246
	Dis	Split-half	−0.37 (−0.57; −0.14)	0.49 (0.22; 0.69)	4.635	<0.001	0.14 (−0.13; 0.39)	0.54 (0.34; 0.69)	2.558	0.005
		Alternating	0.19 (−0.06; 0.42)	0.53 (0.27; 0.72)	1.995	0.023	0.33 (0.08; 0.54)	0.66 (0.50; 0.78)	2.458	0.006

Eng, engagement; dis, disengagement. Internal consistency 95% CI around Spearman–Brown coefficients are given in parentheses.

In accordance with our second hypotheses, participants drinking high amounts of alcohol showed a stronger AB for alcohol cues than participants drinking low amounts of alcohol. This was only the case when assessing AB with the OOOT-adapt (making our third hypotheses redundant). That is, the OOOT-adapt successfully differentiated between low- and high-drinking individuals, and results showed that high-drinking individuals engage faster their attention with alcohol cues than low-drinking individuals. This difference was even more pronounced when the calculation was based on participants' reported amount of used alcohol in the past month when compared with the self-identified average amount of used alcohol. With regard to the disengagement bias, there was no difference between low- and high-drinking individuals when measured with the OOOT-adapt, which seemed in contrast with the findings of the previous study in which alcohol use was related with disengagement bias but not engagement bias when measured with the OOOT (Heitmann et al., 2020). Looking more closely, a similar trend was evident in the current study, but remained non-significant (after the correction of α for multiple comparisons), and also the Bayes factor indicated no clear difference of disengagement bias between groups when AB was assessed with the original OOOT (i.e., when the calculation was based on the reported amount of consumed alcohol in the past month). In addition, taking the low internal consistency of the OOOT in the previous study, as well as in the current study into account, the meaning of this finding remains inconclusive. In contrast, the OOOT-adapt revealed itself as a promising task to be used as a measure of individual differences as it was able to differentiate between low- and high-drinking individuals and at the same time showed improved internal consistency (e.g., McNally, 2019). Future research might want to investigate the predictive validity of the AB indices as derived from the OOOT-adapt regarding alcohol use and craving.

Strengths and Limitations

This study has several strengths, such as the high number of participants and the blinded allocation to one of the two conditions. There are also some limitations to bear in mind when interpreting the results of the current study. First, although the administration of AB using the OOOT-adapt in an alcohol-relevant context was a relevant adaptation, it also entails some disadvantages. That is, the current design of the study, in which the OOOT was administered in the laboratory and the OOOT-adapt in the bar, does not allow disentangling whether the adaptation to the task itself or the context lead to increased internal consistency of the OOOT-adapt. However, this approach allowed increasing the chance of a reliable and valid measure. Knowing that the adaptations indeed improved the reliability of the AB measure, a next step for future research could be to test to what extent the increased internal consistency can be attributed to the optimization of the task, or the context, or whether they have both summatively contributed to the improvement. It is also conceivable that the adaptation to the context might have reduced the reliability of the OOOT-adapt. That is, participants might have been more distracted in the bar context than when completing the task in the laboratory. Although the administration of the task took place in the afternoon when (almost) no visitors were present, we cannot rule out that participants were distracted at any point from the task. However, based on the percentage of correct responses, there is no indication that participants in the OOOT-adapt condition who completed the task in the bar made more mistakes. Second, the current study design does not allow disentangling which adaptations of the OOOT-adapt lead to improved internal consistency. For example, there is evidence that the number of trials from a task can influence its reliability (e.g., Tavakol and Dennick, 2011; Ataya et al., 2012). It might therefore be that the larger number of trials of the OOOT-adapt might partially

TABLE 5 | Mean attentional bias scores and one-tailed independent *t* tests.

	OOOT					
	Self-identified					
	Low (n = 88)	High (n = 29)	t	p	BF_{10}	Cohen d
Engagement	−391.70 (470.60)	−373.75 (560.16)	−0.170	0.433	0.255	−0.036
Disengagement	728.75 (624.34)	819.08 (658.80)	−0.666	0.254	0.399	−0.143
	Reported					
	Low (n = 63)	High (n = 57)	t	p	BF_{10}	Cohen d
Engagement	−389.11 (470.66)	−507.46 (727.23)	1.068	0.856	0.102	0.195
Disengagement	688.41 (549.49)	934.39 (854.01)	−1.894	0.030	1.871	−0.346
	OOOT-adapt					
	Self-identified					
	Low (n = 59)	High (n = 50)	t	p	BF_{10}	Cohen d
Engagement	133.73 (227.35)	257.26 (313.77)	−2.377	0.010	2.463	−0.457
Disengagement	−241.36 (212.45)	−266.53 (350.95)	0.461	0.677	0.224	0.089
	Reported					
	Low (n = 42)	High (n = 67)	t	p	BF_{10}	Cohen d
Engagement	101.15 (199.92)	246.33 (301.82)	−3.014	0.003	5.693	−0.542
Disengagement	−227.19 (222.21)	−269.02 (316.24)	0.749	0.772	0.267	0.147

explain its improvement regarding internal consistency. The number of trials from the OOOT-adapt was actually comparable with other AB measures (e.g., Townshend and Duka, 2007; Pennington et al., 2020). Nevertheless, future research might want to disentangle which adaptations of the OOOT-adapt are relevant regarding its reliability, for example, the influence of the number of trials and in particular the number of trials necessary to reliably measure the process of interest (i.e., AB). Third, although the OOOT-adapt showed improved internal consistency, it did not reach a value that is considered as a "good" reliability coefficient (≤0.8) based on commonly reported thresholds (Clark and Watson, 1995). This might relate to the fact that the task follows an unblocked task design in which trials are randomly presented, the use of divers images, and/or the fact that the task was assessed in a non-clinical sample (see above; Ataya et al., 2012). Furthermore, it has been argued that the commonly used thresholds as defined to assess reliability of questionnaires might not hold for measuring processes such as AB based on reaction times (e.g., Elgersma et al., 2019). Fourth, it might be important to consider that there was a difference in the number of alcohol stimuli on the screen between the alcohol distractors trials that are critical to compute the disengagement bias and the alcohol target trials that are critical to calculate engagement bias (19 vs. 1). The presentation of multiple alcohol images in the alcohol distractors trials was necessary to ensure that the initial attention would be typically directed on an alcohol image, thereby allowing to test how much difficulty participants would experience to redirect their attention to find the single neutral odd-one-out stimulus. However, this difference in the number of alcohol images on the screen between both types of trials might have differentially affected participants' response times, for example, by eliciting stronger craving or distraction from the task when responding to alcohol distractors trials showing multiple alcohol images. In addition, one could also speculate that the multitude of alcohol images elicited multiple instances of engagement next to a difficulty to disengage. Future research might want to investigate to what extent slowed responding to alcohol distractors trials indeed reflects disengagement bias, for example, by using eye-tracking during task performance. Fifth, there were discrepancies between individuals' self-identified average amount of used alcohol and what was later reported during the study about the past month. As indicated, we therefore completed all analyses based on self-identification prior to the study and based on the reported amount of consumed alcohol. Generally, results seem to point in the same direction, and we therefore do not expect that group allocation influenced the results in a relevant way.

Conclusion

Adapting the original OOOT by using more distinct contrast stimulus categories and adding practice trials and more relevant trials, as well as assessing this task in an alcohol-relevant context and in high-drinking individuals, indeed improved the internal consistency of the AB measure. This improved task also showed superior criterion validity as the engagement bias index of the OOOT-adapt could differentiate between low- and high-drinking individuals, thereby adding to the evidence that AB might be involved in substance use. To further test the utility of the OOOT-adapt to index AB, a critical next step would be to evaluate whether the promising psychometric properties also hold in the clinical range, and whether the AB measure not only remains consistent within one assessment procedure but also shows stability over time (test–retest reliability). If proven to be a reliable measure, the OOOT-adapt can enhance the field of research by serving as a task to further test the causal role of AB in addiction.

DATA AVAILABILITY STATEMENT

The raw data supporting the conclusions of this article will be made available by the authors, without undue reservation.

ETHICS STATEMENT

The studies involving human participants were reviewed and approved by Ethical Committee of Psychology (ECP) of the University of Groningen. The patients/participants provided their written informed consent to participate in this study.

AUTHOR CONTRIBUTIONS

JH and NJ were responsible for the data collection. NJ conducted the analysis. JH drafted this manuscript. All authors contributed to the design of the study and further contributed to the writing process and approved this final manuscript.

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Psychometric Properties of the Suicide Stroop Task in a Chinese College Population

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Objective: This study aimed to test the psychometric properties of the suicide stroop task in a Chinese college population.

Methods: College students (n = 121) who were in the 1st–4th grade, fluent in Chinese, and without color blindness were recruited from a university in Guangzhou, China from September to December 2019. Participants were administered the suicide stroop task at baseline and 1-month follow-up.

Results: The suicide stroop task showed excellent internal reliability (Cronbach's α ranged from 0.940 to 0.953). However, the suicide stroop task did not reveal suicide-related attentional biases among current suicide ideators and was not significantly associated with the severity of suicidal ideation, depression, hopelessness, nor anhedonia (all p values $>$ 0.05), indicating a lack of concurrent validity for the task. Additionally, the two-time data of interference scores could not generate intraclass correlation coefficients (ICCs) due to a negative average covariance among data, which indicated poor test–retest consistency for the task.

Conclusion: The results of this study did not support the use of the suicide stroop task on the identification of suicidal risk among Chinese college students. It is crucial to assess the psychometric properties of behavioral measures rigorously as self-report measures before large applications in clinical and community settings.

Keywords: suicide, suicide stroop task, attentional bias, reliability, validity

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INTRODUCTION

Suicide is a major public health issue in young people, with suicide being the second leading cause of death in people between the ages of 15 and 29 years worldwide (Turecki and Brent, 2016). Additionally, suicide has received increasing attention among subgroups of these young people including college students. A meta-analysis showed that pooled prevalence estimates of lifetime suicidal ideation, plans, and attempt were 22.3, 6.1, and 3.2% among college students, and higher estimates were found in samples from Asia (Mortier et al., 2018). It is important to effectively identify people at risk for suicide behaviors to prevent fatal attempt, but the prediction of suicide continues to be a critical challenge (Franklin et al., 2017).

Currently, the screening of suicide risk commonly relies on self-report. However, self-report assessments are limited by the individuals' willingness (e.g., to avoid hospitalization) and ability to report suicidal thoughts (i.e., not aware of suicidal thoughts/suicidal risk) (Glenn et al., 2019). Moreover, a systematic review found that the Beck Hopelessness Scale and the Beck Suicide Intent Scale, two commonly used self-report suicide risk scales, did not have sufficient evidence to support their use on predicting suicide in high-risk samples (Chan et al., 2016). Thus, it seems insufficient to identify suicide risk by self-report alone, and there are increasing arguments on the need of more objective tools on suicide risk determination.

According to the cognitive model of suicidal behavior, suicide-specific attentional bias leads to a fixation on suicide as the sole escape solution, and combined with a state of hopelessness, it would ultimately result in a suicide attempt (Wenzel and Beck, 2008). Previous research found that suicide-specific attentional bias is relevant to previous suicidal attempts in clinical samples (Williams and Broadbent, 1986; Becker et al., 1999; Cha et al., 2010). Specifically, the study conducted by Cha et al. (2010) suggested that suicide-specific attentional bias can be used as a potential behavioral marker to predict future suicide attempt. As these results were very promising, researchers in different countries tried to generalize the measure used in Cha et al.'s study (Cha et al., 2010), the suicide stroop task, into different samples including college students, patients with mood disorders, and community-based samples reporting past-month suicidal ideation (Chung and Jeglic, 2016; Richard-Devantoy et al., 2016; Cha et al., 2017). However, mixed findings were reported. Additionally, a systematic review of the existing seven studies found that the suicide stroop task had excellent internal reliability, but poor classification accuracy to classify suicide attempter from non-attempters (Wilson et al., 2019).

The validity of the suicide stroop task has not been tested in the Chinese context. In this current study, we made a Chinese-language adaption of the suicide stroop task and tested its internal reliability, concurrent validity, and test–retest reliability in Chinese college students. This study aimed to provide more evidence whether the suicide stroop task could be used in a community-based sample in which the majority would not report suicidal ideation and have never made a serious suicidal attempt before. Based on previous research, we hypothesized that (1) those who reported current suicidal ideation (current SI) would also have slower reaction times to suicide-related words than those without current suicidal ideation (non-ideator) and (2) the performance of the suicide stroop task would be significantly associated with suicidal ideation severity, depression, hopelessness, and anhedonia.

METHODS

Participants and Procedures

College students who were in the 1st–4th grade, fluent in Chinese, and without color blindness were recruited from a university in Guangzhou, China from September to December 2019. Participants were recruited online (e.g., WeChat group). Interested participants were invited to a computer laboratory. All participants were asked to provide written informed consent and then to complete the baseline survey and the suicide stroop task. One month later, participants were invited to complete the retest survey and the suicide stroop task in the same laboratory.

This study was approved by the institutional review boards of the Affiliated Brain Hospital, Guangzhou Medical University. Written informed consent has been obtained from all participants.

The Suicide Stroop Task

The suicide stroop task is a computer-based behavior task that uses response latencies of how quickly participants identify the color of different words presented on a computer screen. The test material and test conditions were replicated based on the methodology used in Cha et al. (2010). In this study, stimuli for the task were presented, and response latencies were recorded using E-prime 2.0 software.

After reading the instructions, participants were asked to complete eight practice trial, followed by 48 critical trials. Each trial started with a blank white screen for 4 s, followed by a

TABLE 1 | Sample characteristics and attrition analysis.

	Baseline (*n* = 121)	1-month retest			
		Loss (*n* = 18)	Non-loss (*n* = 103)	χ^2/t	*p*
Age (mean ± SD)	19.0 ± 4.1	19.7 ± 5.0	18.9 ± 4.0	0.729	0.468
Gender				0.007	0.934
Male	46 (38.0)	7 (38.9)	39 (37.9)		
Female	75 (62.0)	11 (61.1)	64 (62.1)		
Residence				0.476	0.490
Urban	83 (68.6)	13 (72.2)	70 (68.0)		
Rural	38 (31.4)	5 (27.8)	33 (32.0)		
Single child				0.129	0.719
Yes	45 (37.2)	8 (44.4)	37 (35.9)		
No	76 (62.8)	10 (55.6)	66 (64.1)		
Relationship status				1.235	0.266
Single	93 (76.9)	12 (66.7)	81 (78.6)		
In a relationship	28 (23.1)	6 (33.3)	22 (21.4)		
Previous suicidal attempt				4.030	0.045
No	117 (96.7)	16 (88.9)	101 (98.1)		
Yes	4 (3.3)	2 (11.1)	2 (1.9)		
Current suicidal ideation				0.462	0.497
No	86 (71.1)	14 (77.8)	72 (6939)		
Yes	35 (28.9)	4 (22.2)	31 (30.1)		
Severity of SI (mean ± SD)	2.5 ± 4.2	1.7 ± 3.3	2.7 ± 4.4	0.907	0.366
Depression (mean ± SD)	5.0 ± 3.5	4.1 ± 3.7	5.2 ± 3.5	1.227	0.222
Hopelessness (mean ± SD)	7.6 ± 2.3	7.9 ± 1.8	7.6 ± 2.3	0.526	0.600
Anhedonia (mean ± SD)	24.1 ± 5.3	22.6 ± 4.7	24.3 ± 5.4	1.242	0.217

TABLE 2 | Comparing the suicide stroop performance across groups of suicidal behaviors (n = 121).

Score*	Overall (n = 121)	Current suicidal ideation			
		Yes (n = 35)	No (n = 86)	t	p
Mean RT_{Sui}	513.03 ± 142.39	530.00 ± 163.72	506.13 ± 133.17	0.835	0.405
Mean RT_{Neg}	513.33 ± 151.64	529.20 ± 177.64	506.87 ± 140.33	0.733	0.465
Mean RT_{Pos}	499.55 ± 144.47	518.46 ± 170.60	491.85 ± 132.73	0.918	0.361
Mean RT_{Neu}	507.55 ± 151.70	529.42 ± 176.81	498.65 ± 140.38	1.012	0.314
$Interference_{Sui}$	5.48 ± 42.23	0.58 ± 44.72	7.48 ± 41.27	0.814	0.417
$Interference_{Neg}$	5.78 ± 44.31	−0.22 ± 32.66	8.22 ± 48.21	0.950	0.344
$Interference_{Pos}$	−8.00 ± 50.58	−10.97 ± 41.03	−6.80 ± 54.16	0.410	0.683

*All suicide stroop score means and standard deviations are reported in milliseconds (ms). *Mean RT, mean response time; interference, suicide/negative/positive word RT–neutral word RT.*

centered "+" in red for 1 s, another blank screen for 1 s, and then the word either in blue or in red color; the words remained on the screen until either a blue or a red key was pressed.

During the critical trial, neutral [house (*fangwu*), paper (*baizhi*), and car (*qiche*)], positive [happy (*kaixin*), success (*chenggong*), and pleasure (*kuaile*)], negative [alone (*gudu*), rejected (*jujue*), and stupid (*yuchun*)], and suicide-related [funeral (*zangli*), dead (*siwang*), and suicide (*zisha*)] words in Chinese characters were presented. After discussion with psychologists, museum, and engine, which were used as neutral words by Cha et al. (2010), were replaced by house and car (in Chinese characters) based on the Chinese contexts in this study. Each of these words was presented four times in random order during the 48 critical trials. The interference score for each category was calculated by subtracting the mean response time (RT) for neutral words from the mean RT for positive, negative, or suicide-related words.

Measures

Socio-Demographics

Socio-demographic information including age, gender, residence, single child or not, and relationship status was collected.

History of Suicidal Attempts

In this study, we used the introduction interview part of the Pathway to Suicidal Action Interview (PSAI) to collect data on previous suicidal behaviors. Approved and assisted by the first author of the PSAI [Millner, A.J. (Millner et al., 2017)], a panel of three bilingual public health researchers, who were also trained in psychiatry and suicide prevention, translated the original English version of the PSAI into simplified Chinese. For an action to be considered as a suicidal attempt, an individual must have had engaged in a potentially deadly behavior with some intention to die (Millner et al., 2017).

Current Suicidal Ideation

The Beck Sale for Suicidal Ideation (BSSI) (Beck et al., 1979) was used to assess the severity of suicidal ideation in the past week. Each item is rated on a 0–2-point scale, with higher scores reflecting more severe suicidal ideation. If one rated either item four or five with a score of one or greater, the person was considered as having current suicidal ideation.

Depression

The degree of depression was assessed by the Patient Health Questionnaire Depression Scale (PHQ-9) (Bian et al., 2009). It consists of nine items related to the diagnostic criteria of major depressive disorder based on the Diagnostic and Statistical Manual of Mental Disorders, Fourth Edition (DSM-IV). The total score ranges from 0 to 27, with higher scores indicating higher levels of depression.

Hopelessness

Hopelessness was measured by the 4-item Beck's Hopelessness Scale (BHS-4) (Yip and Cheung, 2006; Ma et al., 2020). It consists of four items relevant to success, dark future, breaks, and faith. Item responses range from 1 (strongly agree) to 5 (strongly disagree). The possible score ranges from four to 20, and a higher score represents a higher level of hopelessness.

Anhedonia

Anhedonia was measured by the Snaith–Hamilton Pleasure Scale (SHAPS) (Snaith et al., 1995). It is a validated and reliable scale that was developed to assess the ability to experience pleasure in normally pleasurable activities in the past few days. It consists of 14 items, and each item is rated on a 4-point Likert format, ranging from 1 (strongly agree) to 4 (strongly disagree) (Hu et al., 2017). The total score ranges from 14 to 56, with higher scores indicating lower ability to experience pleasure.

TABLE 3 | Partial correlations of the suicide stroop performance and other psychosocial variables (n = 121).

	1	2	3	4	5	6	7
$Interference_{Sui}$	1.000	–	–	–	–	–	–
$Interference_{Neg}$	0.663**	1.000	–	–	–	–	–
$Interference_{Pos}$	0.402**	0.299**	1.000	–	–	–	–
Severity of SI	−0.082	−0.094	−0.022	1.000	–	–	–
Depression	0.011	0.024	0.051	0.307**	1.000	–	–
Hopelessness	−0.010	0.038	0.085	0.390**	0.321**	1.000	–
Anhedonia	0.035	−0.003	−0.015	0.225*	0.325**	0.363**	1.000

All suicide stroop score means and standard deviations are reported in milliseconds (ms). Interference, suicide/negative/positive word RT–neutral word RT. $^{*}p < 0.05$; $^{**}p < 0.01$.

Statistical Analysis

Regarding the suicide stroop task, we included trials with correct responses in the analysis. For all participants, the rate of correct response was 97.7%, and the correct response rates for suicide-related (97.5%), negatively-valenced (97.0%), positive-valenced (97.9%), and neutral words (98.3%) did not significantly differ from one another ($\chi^2 = 5.301$, $p = 0.151$). Additionally, we eliminated trials with response latencies ±2 SD from each participant's mean response latency.

Internal reliability was evaluated using the criterion of Cronbach's alpha $\geq$ 0.70. Regarding concurrent validity, we firstly performed independent sample *t*-tests to assess the group differences in mean RTs or interference scores (suicide/negative/positive word RT–neutral word RT) for each valence word, and then we conducted Group × Valence (repeated measures analysis) ANOVAs. Group comparisons included current ideators vs. non-ideators. For the within-subjects factor, valence had four levels in mean RT analyses (i.e., suicide-related, negative, positive, and neutral) and three levels in interference scores (i.e., suicide-related, negative, and positive). Additionally, Pearson correlation analysis was used to evaluate the correlations between suicide stroop task performance (mean RTs and interference scores) and the severity of current suicidal ideation, depression, hopelessness, and anhedonia. Test–retest reliability was assessed by the paired-sample *t*-test and intraclass correlation coefficients (ICCs). All analyses were performed by SPSS version 23.0 (SPSS Inc., Chicago, IL, United States). The level of significance was set at 0.05.

RESULTS

Demographic Characteristics

As presented in **Table 1**, a total of 121 college students participated in this study. Among them, 62.0% were female, and the mean age was 19.0 years (SD = 4.1). There were 3.3% of participants reporting previous suicidal attempts and 28.9% reporting current suicidal ideation. One month after baseline, 103 students (85%) completed the retest. Except for previous suicidal attempts, no significant differences were found in socio-demographic or psychosocial characteristics at baseline between lost samples and those who finished the retest (**Table 1**).

Internal Reliability

The mean RTs for each valence word demonstrated excellent internal reliability (Cronbach's α ranged from 0.940 to 0.953).

Concurrent Validity

Across the sample, a significant difference was found from the mean RT for suicide-related words, $M = 513.03$ (SD = 142.39 ms); negative valenced words, $M = 513.33$ (SD = 151.64 ms); positive valenced words, $M = 499.55$ (SD = 144.47 ms); and neutral words, $M = 507.55$ (SD = 151.70 ms), $F = 5.139$, $p = 0.025$. A least significant difference (LSD) analysis was conducted for multiple comparisons. The results of LSD indicated that the mean RT for suicide-related words and negative valenced words was significantly longer than the mean RT for positive valence words (ds = −13.486, −13.782, ps < 0.05).

As shown in **Table 2**, the results of independent sample *t*-tests revealed that no group difference in mean RTs or interference scores for each valence word was related to current suicidal ideation (t = 0.410–1.012, p = 0.314–0.683). Group × Valence interactions (repeated measures analysis) were also not significant when testing mean RTs or interference scores for two-group comparison (current SI vs. non-ideators, $F = 0.795$, $p = 0.374$).

As shown in **Table 3**, the results of Pearson correlation analysis showed that the interference score for each valence word was not significantly associated with the scores of suicidal ideation severity, depression, hopelessness, or anhedonia (rs = −0.094–0.085, ps > 0.05).

Test–Retest Reliability

As shown in **Table 4**, the paired-sample *t*-test showed no significant differences for mean RTs or interference scores of each

TABLE 4 | Comparing the suicide stroop performance and other psychosocial variables between baseline and retest (n = 103).

	Baseline	Retest	*t*	*p*	ICC*	*p*
Mean RT_{Sui}	516.36 ± 147.35	507.78 ± 124.34	0.856	0.394	0.838	0.000
Mean RT_{Neg}	517.07 ± 157.07	502.37 ± 119.66	1.445	0.151	0.842	0.000
Mean RT_{Pos}	504.21 ± 151.28	501.66 ± 126.56	0.227	0.821	0.799	0.000
Mean RT_{Neu}	511.62 ± 158.19	497.06 ± 106.35	1.394	0.166	0.817	0.000
$Interference_{Sui}$	4.74 ± 42.85	10.72 ± 47.01	0.842	0.402	−0.801	0.998
$Interference_{Neg}$	5.45 ± 45.64	5.30 ± 43.74	0.023	0.982	−0.081	0.652
$Interference_{Pos}$	−7.41 ± 52.37	4.60 ± 50.03	1.530	0.129	−0.533	0.984
Severity of SI	2.65 ± 4.38	2.43 ± 4.37	0.736	0.464	0.859	0.000
Depression	5.16 ± 3.48	4.67 ± 3.45	1.473	0.144	0.696	0.000
Hopelessness	7.58 ± 2.35	7.75 ± 2.45	0.796	0.428	0.762	0.000
Anhedonia	24.30 ± 5.42	23.70 ± 5.90	1.343	0.182	0.808	0.000

*All suicide stroop score means and standard deviations are reported in milliseconds (ms). Mean RT, mean response time; interference, suicide/negative/positive word RT–neutral word RT. *The ICC value was negative due to a negative average covariance among data collected at baseline and retest, which violated reliability model assumptions.*

valence word between baseline and retest. However, the two-time data of interference scores could not generate ICC values due to a negative average covariance among data, which violated reliability model assumptions.

DISCUSSION

The goal of the current study was to test the psychometric properties of the suicide stroop task. Consistent with previous research (Wilson et al., 2019), the mean RTs for all valence words demonstrated good internal reliability. However, the suicide stroop task performance lacked concurrent validity, as the suicide stroop task did not reveal suicide-related attentional biases among current suicide ideators. We also found that the suicide stroop task performance was not significantly associated with the severity of suicidal ideation, depression, hopelessness, nor anhedonia, which indicated a lack of concurrent validity for the task as well. Additionally, the interference scores of all stimuli showed poor test–retest consistency, whereas other self-report measures (i.e., BSSI, PHQ-9, BHS-4, and SHAPS) showed moderate-to-good test–retest consistency. Thus, the results of this study did not support the use of the suicide stroop task on the identification of suicidal risk among Chinese college students.

There might be some reasons for these results. First, the general reaction time is associated with age-related differences in cognitive ability. Our samples were much younger than those in studies with positive results (Williams and Broadbent, 1986; Becker et al., 1999; Cha et al., 2010). Second, the suicide stroop task might be more sensitive in depressive people with recent suicidal attempts (Chung and Jeglic, 2016), whereas in this study, the majority were not depressive and did not have a recent suicidal attempt. Third, as the suicide stroop task uses manual reaction times (i.e., press a key) in responding to the stimuli as a measure, other paradigms, such as voice and eye movements in responding to the stimuli, might perform better.

This study is limited by a small convenience sample. Among 121 participants, 35 participants had current suicidal ideation, and four participants reported previous suicidal attempts. However, in the community, most people will not report suicidal ideation, and the majority will have never made a serious suicide attempt before. That is the reason why we need more sensitive measures with high accuracy on screening suicidal risk.

Over the past 50 years, there was a surge of research designed to identify the risk factors for suicidal behaviors, and many different theories of suicide have been proposed (Wenzel and Beck, 2008; Franklin et al., 2017). It is still a critical challenge on the identification of suicide risk and the prediction of suicide. We believe it is of great meaning to explore more objective measures or behavior markers related to suicidal behaviors. However, it is crucial to assess the psychometric properties of behavioral measures rigorously as self-report measures before large applications in clinical and community settings.

DATA AVAILABILITY STATEMENT

The raw data supporting the conclusions of this article will be made available by the authors, without undue reservation.

ETHICS STATEMENT

The studies involving human participants were reviewed and approved by The Affiliated Brain Hospital of Guangzhou Medical University. The patients/participants provided their written informed consent to participate in this study.

AUTHOR CONTRIBUTIONS

LN conducted the data analysis and drafted the manuscript. All authors contributed to the study design, provided substantial editorial input in the drafting of the manuscript, and read and approved the final manuscript.

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A Pictorial Dot Probe Task to Assess Food-Related Attentional Bias in Youth With and Without Obesity: Overview of Indices and Evaluation of Their Reliability

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Several versions of the dot probe detection task are frequently used to assess maladaptive attentional processes associated with a broad range of psychopathology and health behavior, including eating behavior and weight. However, there are serious concerns about the reliability of the indices derived from the paradigm as measurement of attentional bias toward or away from salient stimuli. The present paper gives an overview of different attentional bias indices used in psychopathology research and scrutinizes three types of indices (the traditional attentional bias score, the dynamic trial-level base scores, and the probability index) calculated from a pictorial version of the dot probe task to assess food-related attentional biases in children and youngsters with and without obesity. Correlational analyses reveal that dynamic scores (but not the traditional and probability indices) are dependent on general response speed. Reliability estimates are low for the traditional and probability indices. The higher reliability for the dynamic indices is at least partially explained by general response speed. No significant group differences between youth with and without obesity are found, and correlations with weight are also non-significant. Taken together, results cast doubt on the applicability of this specific task for both experimental and individual differences research on food-related attentional biases in youth. However, researchers are encouraged to make and test adaptations to the procedure or computational algorithm in an effort to increase psychometric quality of the task and to report psychometric characteristics of their version of the task for their specific sample.

Keywords: attentional bias, dot probe paradigm, reliability, children and adolescent, obesity

INTRODUCTION

Different theoretical accounts on problematic eating, overweight, and obesity propose that food stimuli automatically attract visual attention, particularly in individuals with overweight and weight concerns (e.g., Appelhans, 2009; Berridge, 2009; Appelhans et al., 2016; Tanofsky-Kraff et al., 2020). An attentional preference for food is thought to have been evolutionary adaptive in ancient times since it facilitated finding scarce food in the harsh environment, allowing food intake whenever possible, and thus increasing chances for survival. However, in the present obesogenic environment, where energy-dense food is abundantly available, this same mechanism might trigger overeating and increase the risk for overweight and obesity (Robinson and Berridge, 1993, 2008; Blundell and Cooling, 2000; Paquet et al., 2017).

Neurophysiological studies, using brain imaging techniques or eye tracking procedures, indeed support the prediction that individuals with overweight and obesity show stronger attentional biases toward food than individuals with healthy weight, in adults (Hendrikse et al., 2015) and youth (van Meer et al., 2016; Biehl et al., 2020) alike. Evidence stemming from behavioral paradigms, however, is equivocal and shows small to moderate effect sizes at best, especially in youth populations (van Meer et al., 2016; Brand et al., 2020; Hagan et al., 2020; Kemps et al., 2020; Hardman et al., 2021). A possible explanation for these ambiguous results precluding clear conclusions on the role of attentional processes in eating behavior and weight, relates to the considerable methodological inconsistency between studies. Several reaction time tasks are used to measure attentional bias toward food in youth samples, with among others adapted versions of the Stroop task (Stroop, 1935; Braet and Crombez, 2003) and visual search paradigms (Verghese, 2001; Brand et al., 2020). A dot probe detection task (Macleod et al., 1986) with pictures of unhealthy food as targets is the most widely used behavioral paradigm to investigate attentional processes toward food in youth (Kemps et al., 2020).

Several versions of the dot probe procedure have been used to investigate attentional processes in a broad range of problems, like anxiety, depression, addiction, obesity, and problematic eating behavior (Puliafico and Kendall, 2006; Field et al., 2016; Starzomska, 2017; Jiang and Vartanian, 2018; Burris et al., 2019; Rojo-Bofill et al., 2019; Kemps et al., 2020), encompassing a large variety in task characteristics like presentation times, stimulus types (verbal, pictorial), stimulus alignment (vertical, horizontal), instruction (detect probe, categorize probe orientation), and number of trials. Despite this procedural variability, the basic set-up of a trial in the dot probe is rather straight-forward. In general, a pair of stimuli is presented simultaneously on the screen. In target trials, one of these stimuli is considered relevant (e.g., a picture of food in obesity research), while the other is neutral (e.g., a picture of household appliances). Presentation duration of the stimulus pair is typically short (e.g., 200–250 ms: Godijn and Theeuwes, 2002; Born et al., 2011) in an attempt to trigger and capture fast, automatic or uncontrolled processes. In research with youth samples, presentation duration is generally 500 ms (Shechner et al., 2012; Dudeney et al., 2015). Immediately after the stimulus pair is removed from the screen, a probe (e.g., a white dot) is presented at one of the two stimulus locations. If this probe appears at the location of the relevant target stimulus, the trial is considered "congruent." If the probe appears at the opposite location of the target stimulus, the trial is considered "incongruent." Participants are asked to indicate as fast as possible in which location the probe appears using one of two response keys. Irrespective of the procedural variability, all studies start from the same theoretical assumptions that participants will consistently react faster to a stimulus that appears in a location where their attention was already focused on than to a stimulus appearing in an unattended location (Posner, 1980). This assumption has given rise to the calculation of an attentional bias score (ABS) as the mean difference score between reaction times on incongruent and congruent trials (RT incongruent–RT congruent, e.g., Macleod et al., 1986). Applied to attentional biases in eating behavior and weight problems, an attentional bias toward food would therefore be visible in positive attentional bias scores, while attentional avoidance of food would be visible in negative attentional bias scores. This "traditional" ABS is still the most frequently used behavioral index of biased attention. However, the index shows considerably problematic psychometric characteristics in terms of reliability and validity (Schmukle, 2005; Ataya et al., 2012a,b; Field and Christiansen, 2012; Rodebaugh et al., 2016; Chapman et al., 2019; McNally, 2019).

While the dot probe paradigm is frequently used tool in experimental research allowing to test group differences, its adequacy to assess individual differences in correlational clinical research has been questioned repeatedly. One major objection against its use is driven by its unacceptably low levels of reliability (Schmukle, 2005; Ataya et al., 2012a,b; De Schryver, 2018; Parsons et al., 2018; Chapman et al., 2019; MacLeod et al., 2019; Van Bockstaele et al., 2020). This observation is related to the so-called Reliability Paradox (De Schryver et al., 2016; Hedge et al., 2018; Goodhew and Edwards, 2019). Experimental research seeks to minimize differences between individuals in experimental conditions aiming to maximize between-group differences following experimental manipulation. A reliable task in experimental research is a task with low measurement error that yields the most homogeneous performance in one group compared to the homogeneous performance in the other group. Between-group differences can then be attributed to the experimental manipulation rather than to individual differences. In contrast, correlational research seeks to maximize interindividual differences in heterogeneous samples. A reliable task in correlational research depends on the extent to which the instrument consistently ranks individuals based on the variance in their true-score variance (Cronbach, 1957). These diverging takes on reliability in experimental vs. correlational research, based on the different aims of both research domains, has its implications on the transfer of experimental paradigms to correlational studies. Reliabilities in correlational research generally do not reach the cut-off guidelines used in experimental research, let alone the 0.90 that is recommended for making inferences about individuals (Rodebaugh et al., 2016). However, if low reliability is mainly due to a lack of true-score variance rather than a large amount of error-score variance, a (correct) use of the instrument might not be problematic (De Schryver et al., 2016, 2018a).

Generally, in individual differences research, reliability, and validity of reaction time-based indices of attentional bias, including the ABS calculated from the dot probe, are not routinely reported (Green et al., 2016; Rodebaugh et al., 2016; Parsons et al., 2018; Goodhew and Edwards, 2019). This practice stands in stark contrast with the strict psychometric requirements posed to variables based on questionnaire scores (Vasey et al., 2003; Parsons et al., 2018) and discords with the prerequisite of reliable instruments for effective research (Lebel and Paunonen, 2011; De Schryver et al., 2016, 2018a). The scarce psychometric reports of tasks based on the dot probe procedure that have been published, have repeatedly shown unacceptable low levels of reliability of the traditional ABS, both in adult populations (Schmukle, 2005; Lebel and Paunonen, 2011; Rodebaugh et al., 2016; Parsons et al., 2018; Chapman et al., 2019; Hagan et al., 2020; Molloy and Anderson, 2020) and in youth (Britton et al., 2013; Brown et al., 2014; Waechter et al., 2014; Fu and Perez-Edgar, 2019). Furthermore, instead of reporting reliability of the ABS, several reliability reports (e.g., Vervoort et al., 2011; Haft et al., 2019) are often limited to the reliability of the unprocessed reaction times (RT) which are rather stable and consistent among participants. However, general reaction time, thus being (relatively) fast or slow, does not provide information of attentional bias: to infer about attentional bias, comparison between responses on congruent vs. incongruent trials is needed. Therefore, since reliability indices should be referring to the outcome of interest (i.e., ABS as index of attentional bias, not general RT as index of processing speed), this practice is non-committal (Kruijt et al., 2016; Parsons et al., 2018).

Aiming to improve psychometric properties of the dot probe task, researchers have been considering adaptations in the task design and the computation of the attentional bias index (Price et al., 2015). Procedural adaptations include using idiosyncratic, personally relevant rather than general stimuli (Christiansen et al., 2015; van Ens et al., 2019), or prolonging stimulus presentation up to 3000–5000 ms (Waechter et al., 2014; van Ens et al., 2019). However, none of these procedural adaptations managed to establish adequate reliability (Jones et al., 2018). In addition to such procedural adaptations, several scholars examined different computational methods to calculate alternative indices of biased attention, exploring their impact on validity and reliability (Price et al., 2019). Reliability might be improved when traditional ABS are calculated by using only bottom-target trials (instead of both top- and bottom-target trials) in vertically oriented dot probe tasks (Price et al., 2015; Aday and Carlson, 2019), although this approach is not always successful (Jones et al., 2018). Simply distinguishing between vigilance (difference between congruent relevant trials and neutral trials) and disengagement (difference between incongruent relevant trials and neutral trials) aspects of attention also failed to improve reliability (Koster et al., 2004; Waechter et al., 2014). However, adopting a response-based approach to vigilance and disengagement might result in higher reliability scores (Evans and Britton, 2018).

Alternatively, researchers challenged the assumption of attentional bias as a stable concept underlying the traditional calculation of the ABS, and suggest that attentional bias is a dynamic process fluctuating over time, with attention being switched back and forth between relevant and neutral stimuli (Iacoviello et al., 2014; Zvielli et al., 2014, 2015; Rodebaugh et al., 2016; McNally, 2019; Hardman et al., 2021). To account for this dynamic in attention, Iacoviello and colleagues proposed the attention-bias variability score (ABVS, Iacoviello et al., 2014). The ABVS is computed by grouping the dot-probe trials in sequential bins, calculating the ABS for each bin, and dividing the SD of ABS across all bins by the mean RT of the total task. The resulting ABVS is an index of stability of attention biases, with increasing ABVS thus suggesting more fluctuation in attentional biases toward and away from relevant stimuli over time. The ABVS, however, does not allow to differentiate between the approach and avoidance aspects of these dynamics, which might be of particular interest in eating behavior (Liu et al., 2019a,b; Hardman et al., 2021). Zvielli et al. (2015) proposed the trial-based bias-score (TL-BS) as a way of simultaneously distinguishing direction of attention and dynamic variability over time. The TL-BS is computed by forming pairs of congruent and incongruent trials on the basis of temporal proximity and subtracting the RT of the congruent trial from RT of the incongruent trial for each pair. From the resulting time-series of TL-BS's, five indices of biased attention can then be derived for each participant: The mean and peak of all the positive TL-BS's in the series (TL-BSpos), the mean and peak of all negative TL-BS's in the series (TL-BSpos), and a TL-BS variability index, computed as the mean absolute distance across the whole series of TL-BS's. Applied to attentional biases in eating behavior and weight problems, mean TL-BSpos is considered to reflect the amount of approach bias toward food, mean TL-BSneg the amount of avoidance bias away from food, peak TL-BSpos and peak TL-BSneg, the maximum expression of bias toward vs. away from food, respectively and TL-BS variability the amount of fluctuation between bias toward and away from food over time (Liu et al., 2019a,b). ABV and TL-BS scores are thought to show a cyclic pattern, reflecting one's attention switching toward and away from the relevant stimuli over time (Iacoviello et al., 2014; Zvielli et al., 2015). In adult samples, reliability of these dynamic indices of attentional bias is superior compared to traditional indices (Zvielli et al., 2015; Rodebaugh et al., 2016; Molloy and Anderson, 2020). On top of the general theoretical assumption underlying the traditional approach to the dot probe (faster RT to a stimulus appearing in the already attended location, i.e., faster RT in congruent than in incongruent trials), the dynamic approach adds the assumption that the RT differences between congruent and incongruent trials may vary meaningfully over time. Higher variability over time is thought to reflect pathological attention switching while a more stable pattern of attention orienting is thought to be adaptive (Zvielli et al., 2015). However, when accounting for general variability in RT, these assumptions may not hold (Zvielli et al., 2015; Kruijt et al., 2016; Carlson and Fang, 2020). In a monte-carlo simulation it was shown that the dynamic indices are likely to capture not only information of attentional bias, but also of measurement error (Kruijt et al., 2016; McNally, 2019). Furthermore, when accounting for general variability in RT in adult samples, the superior reliability of the dynamic indices is also lost (Carlson and Fang, 2020).

Research on reaction-time paradigms has illustrated that RT-based indices are largely influenced by general response speed (Fazio, 1990; Faust et al., 1999; Greenwald et al., 2003; Glashouwer et al., 2013; De Schryver et al., 2018b), with larger indices (independent from direction) for individuals with slower reaction times across the task. General response speed and RT-variability are found to decrease from childhood over adolescence to adulthood, while increasing from then on (Dykiert et al., 2012; Adleman et al., 2016). This developmental trajectory might typically be associated with even smaller indices in youth samples compared to adult samples. An innovative approach to compute meaningful indices based on RT based data while accounting for differences in general response speed, has been proposed recently by De Schryver and de Neve (2018). They suggested the Probability Index (PI) as an index for the Implicit Associations Test (I.A.T., Greenwald et al., 1998), with enhanced reliability over traditional I.A.T.-indices (De Schryver et al., 2018b). The PI reflects the probability that a randomly chosen response on a congruent trial is faster than a randomly chosen response on an incongruent trial. Although not earlier used to index attentional processes, this approach can easily be transferred to the dot probe paradigm, with higher PI's reflecting stronger attentional bias toward the relevant stimuli. Applied to attentional biases in eating behavior and weight problems, an attentional bias toward food would therefore be visible in higher PI's (PI > 0.05), while attentional avoidance of food would be visible in lower PI's (PI < 0.05).

Irrespective of the decennia-old abundance of literature discussing the limitations of behavioral reaction time paradigms to assess individual differences in biased attention and the static or dynamic nature of attentional processes (Schmukle, 2005; Field et al., 2016; Rodebaugh et al., 2016; Goodhew and Edwards, 2019), scholars investigating attention bias to food and developers of innovative theory-based interventions targeting these processes (Eichen et al., 2017; Kemps et al., 2020) nevertheless keep on using reaction time tasks, including the dot probe task, in their work, often without evaluating psychometric properties of the specific test in the specific study sample. This practice urged the effort to establish evidence for the use of a pictorial dot probe task to investigate food-related attentional biases in youth with and without obesity. The present study will take on this challenge, by scrutinizing psychometric properties of traditional as well as innovative indices of the dot probe: the traditional ABS, the dynamic TL-BS, and the probabilistic PI. It will be examined whether food-related attentional biases can be meaningfully and reliably assessed using the different bias indices computed from responses on a pictorial dot probe, in a sample of youth with and without obesity. The applicability of this specific task procedure will be evaluated for experimental research, by testing group differences, and for individual differences research, by calculating correlations of the indices with weight (Greenwald et al., 2003). Reliability (in terms of performance stability) of the indices will be evaluated by comparing performance in the first part of the task with performance in the second part.

METHODS

Sample

Participants of the present study were 337 children and adolescents (65% girls), aged between 7 and 19 ($M = 14$, $SD = 2.59$). 59.64% of the participants were recruited in the WELCOME-project (ISRCTN14722584, Naets et al., 2018), a RCT evaluating executive functions training for weight control in youth. Children and adolescents (age $M = 14$, $SD = 2.45$) in this subsample were all obese (adjusted BMI: $M = 183.47$, $SD = 35.17$). The remaining participants (age $M = 13$, $SD = 3.36$) were recruited in convenience samples by Master students at Ghent University under supervision of LV and TN. They were all normal weight (adjusted BMI: $M = 99.18$, $SD = 7.95$). This sample size was justified by data availability: all data that were collected at Ghent University, using this particular dot probe procedure between 2017 and 2020 were used. As such, the sample size is sufficiently large to detect group differences of $d = 0.4$, which is considered the smallest effect size of interest in psychology (Lakens et al., 2018), and reach 80% power for alpha $= 0.05$ (Brysbaert, 2019). Both data collection procedures were approved by the IRB (UZGent 2017/0305 and UGent FPPW 2019/79).

Weight

To index weight status in a developmentally appropriate way, age and sex adjusted Body Mass Index (adjBMI) was calculated by dividing measured BMI (weight in kg/squared length in cm) by norm BMI for age and sex, and multiplying this by 100. Norm BMI for age and sex was determined as the 50th percentiles of the BMI for age and sex based normative data. An adjBMI equal to or smaller than 85% is considered underweight, equal to or $>$120% as overweight, equal to or $>$140% as moderate obesity, equal to or $>$160% as extreme obesity.

Dot Probe Task

Attentional bias toward food-related stimuli was measured using a pictorial version of the dot probe task (Macleod et al., 1986) with food and neutral stimuli selected from the Food-pics database (www.food-pics.sbg.ac.at Blechert et al., 2014). Picture pairs were matched for visual complexity, brightness, and contrast. The data were collected by means of a dedicated JavaScript web application that runs in the browser. Stimulus presentation routines were handled by a custom Python-based backend. The software was developed by ImplicitMeasures.com, a spin-off company of Ghent University (Belgium). After presentation of a white fixation cross in the middle of the screen, a picture pair is presented for 500 ms, one to the left and one to the right of the center. This procedure was chosen to match earlier work on food-related attentional bias in adult samples (Kemps et al., 2014). Immediately following the pictures, a white dot appears on one of the locations (either left or right). Participants are asked to react to the dot by pressing "e" on a keyboard when the dot appears on the left side, and pressing "i" when it appears on the right side. In total, 140 trials are presented, of which 10 neutral-neutral trials as practice trials, 16 food-neutral pairs each presented four times, resulting in 64 experimental trials. The remaining trials are filler trials presenting two neutral pictures (neutral-neutral trials) (Naets et al., 2018).

Indices for Attentional Bias

Trials with RT outliers (trial RT $<$ 200 ms or $>$ 1.500 ms) or incorrect responses were excluded.

Traditional Attentional Bias Score (ABS) (Macleod et al., 1986) was calculated by subtracting RT of congruent trials from RT of incongruent trials, such that ABS>0 are indicative of bias toward food and ABS<0 of bias away from food. Additionally, the absolute value of ABS is taken, so higher values indicate stronger effects, either toward or away from food.

Dynamic indices (Zvielli et al., 2015) were conceptualized as TL-BS parameters. The TL-BS is computed by subtracting the RT of a congruent trial from RT of its incongruent counterpart for pairs of trials that were in close temporal proximity (not further than five trials apart). Mean and peak values of all the positive TL-BS's in the series are indicative of bias toward food. Mean and peak values of all the negative TL-BS's are indicative of bias away from food. The TL-BS variability value indexes the amount of fluctuation between bias toward and away from food over time.

To accommodate for the expected response speed artifact, the *Probability Index* (PI) (De Schryver and de Neve, 2018) is calculated using the following formula so that higher PI is indicative of attentional bias toward food:

$$PI = \frac{U}{(N_{CT} \times N_{IT})},$$

with U being the Wilcoxon test statistic for two samples. To ignore direction of the effect, the absolute value of PI, abs(PI-0.5), is taken. Again, higher values indicate stronger effects, either toward or away from food.

Criteria for Evaluating the Indices and Analytic Plan

Evaluating the different indices was done stepwise, vis-à-vis the considerations below.

Independence of General Response Speed

RT-based effects, as the ABS and the TL-BS are known to be inflated for individuals responding slowly (Fazio, 1990; Faust et al., 1999; Greenwald et al., 2003; Glashouwer et al., 2013; De Schryver et al., 2018b). Since general response speed gets faster from childhood to adolescence (Dykiert et al., 2012; Adleman et al., 2016), RT-based effects are expected to be negatively correlated with age. To maximize the independency of the different indices and the measure for general response speed, the average RT of the neutral trials as a measure of general response speed was chosen. A positive correlation between these indices and general response speed on neutral trials can therefore be expected. Such a correlation is expected to be non-significant when using the PI. It would be preferable for an index of attentional bias to minimize the correlation with general response speed and age.

Reliability of the Indices

Split-half reliability of these indices (ABS, TL-BS, and PI) will be estimated by Pearson correlations between index scores calculated in both test halves, for the total group and for both weight status groups separately. To test if reliability was influenced by general response speed, linear models predicting performance in first test half by reliability in second test half and RT were computed.

The Dot Probe in an Experimental Context

Based on the theoretical assumptions on problematic eating and overweight and obesity, a significant difference can be expected between youth with obesity and youth with normal weight on their reaction to food vs. neutral stimuli. This will be tested using linear mixed models (LMM, Field, 2012) with RT as dependent variable, fixed factors weight status, trial type (both effect coded), and the interaction term weight status x trial type and with participant as random factor (Model 1). In addition to raw RT, logRT will also be tested to account for the typical skewness of the raw RT distribution (Model 2). If the pictorial dot probe would be suitable for use in experimental research in this youth sample, a significant interaction effect between trialtype and weight status would emerge.

In order to evaluate whether the attentional biases indices are capable of predicting weight status, nine separate Linear Probabilistic Models (LPMs) with weight status group as dependent variable (dummy coded), and the indices as independent variables will be reported. To control for general response speed, mean RT on neutral-neutral trials will also be added as between-subject variable. If an index would be a meaningful measurement of food-related attentional bias in experimental research for this youth sample, weight status would be significantly predicted by the index, with no significant effect of general response speed.

The Dot Probe Indices in An Individual Differences Context

Because attentional bias for food is thought to be stronger in individuals with higher weight, the linear association between the attention bias indices and adjusted BMI will be estimated by Pearson correlations. If an index would be a meaningful individual differences variable, significant positive correlations would emerge.

RESULTS

Descriptives

Table 1 shows the descriptives of the seven attention bias indices, for the total sample ($n = 337$), and both weight groups separately (obesity: $n = 201$, normal weight, $n = 136$).

The Dot Probe in an Experimental Context: Group Differences Between Obesity and Normal Weight

Table 2 shows the results of the LMMs. The main effects of congruency and weight status were not significant when predicting raw RT. Also, the crucial interaction term between

TABLE 1 | Descriptives of the different attention bias indices.

	Total sample		Obesity		Normal Weight	
	M	(SD)	M	(SD)	M	(SD)
ABS	0.96	(35.48)	0.08	(36.78)	2.24	(33.59)
mean TL-BSpos	120.01	(58.98)	122.13	(60.50)	116.94	(56.80)
peak TL-Bspos	428.73	(232.34)	423.49	(229.50)	436.35	(237.06)
mean TL-Bsneg	−119.07	(55.13)	−124.08	(55.40)	−111.81	(54.11)
peak TL-Bsneg	−442.88	(229.87)	−461.05	(227.11)	−416.50	(232.13)
TL-BS variability	169.86	(72.46)	173.22	(71.99)	164.97	(73.13)
PI	0.50	(0.07)	0.51	(0.08)	0.50	(0.07)

TABLE 2 | Regression coefficients of the Linear mixed models.

Fixed effects*	Model 1: raw RT			Model 2: logRT		
	Estimate	Std. error	*t*-value**	Estimate	Std. error	*t*-value
(Intercept)	522.44	5.90	88.62	6.21	0.01	612.06
Trial type	0.71	0.98	0.72	0.00	0.00	1.02
Weight status group	−6.59	5.90	−1.12	−0.01	0.01	−0.94
Trial type × weight status group	0.34	0.98	0.35	0.00	0.00	−0.21

**Trial type and Weight status group are effect coded. **No p-values are reported by the lmer-package: absolute t-values > 2 are considered significant for alpha = 0.05.*

TABLE 3 | Correlation of attention bias indices with age and mean neutral-neutral RT.

	Age	RT (neutral-neutral)
ABS	0.17*	−0.09
Absolute value ABS	−0.20**	0.31**
mean TL-BSpos	−0.32**	0.65**
peak TL-BSpos	−0.22**	0.46**
mean TL-BSneg	0.39**	−0.68**
peak TL-BSneg	0.22**	−0.41**
TL-BS variability	−0.35**	0.65**
PI	0.12	−0.08
Abs(PI-0.5)	0.08	−0.01

**p < 0.05; **p < 0.01.*

TABLE 4 | Split-half correlation as index of reliability for the different attentional bias indices.

	Total sample	Obesity	Normal weight
ABS	0.05	0.03	0.09
Absolute value ABS	−0.02	0.01	−0.07
Mean TL-Bspos	0.29**	0.26*	0.34**
Peak TL-Bspos	0.24**	0.20	0.32*
Mean TL-Bsneg	0.28**	0.29*	0.30
Peak TL-Bsneg	0.13	0.10	0.19
TL-BS variability	0.43**	0.39**	0.45**
PI	0.08	0.02	0.16
Abs(PI-0.5)	0.01	−0.02	0.07

**p < 0.05; **p < 0.01.*

those two factors was not significant. The same observation was made when predicting log RT. In other words, there is no evidence that RT depends on the congruency of the trials, not even for the obese weight group.

Independence of Mean Response Speed and Age

Table 3 shows Pearson correlations of the attention bias indices with age and mean reaction time on neutral-neutral trials. Traditional ABS, PI and abs(PI-0.5) were not significantly related to mean RT; absolute ABS and TL-BS indices, however, were, with correlations indicating medium effect sizes for absolute value ABS and peak TL-BS, and large effect sizes for mean TL-BS and TL-BS variability. PI and abs(PI-0.5) were not significantly related to age. The other indices, however, were, with correlations indicating small to medium effect sizes.

Reliability

Table 4 shows the split-half reliability estimates for the nine attention bias indices, for the total sample, and the weight status groups separately. Split-half reliability was only significant for mean and peak TL-BSpos, mean TL-BSneg, and TL-BS variability. **Table 5** shows, however, that when the association between the scores on both test-halves is controlled for general response speed (on neutral trials), no significant associations between the two halves remain.

TABLE 5 | Linear models predicting performance in the first task half by performance in the second task half and general response speed.

	mean TL-BSpos		Peak TL-BSpos		Mean TL-BSneg		TL-BS variability	
	Std coeff	**0.01 CI**	**Std coeff**	**0.01 CI**	**Std coeff**	**0.01 CI**	**Std coeff**	**0.01 CI**
(Intercept)	0.00	[−0.15, 0.15]	0.00	[−0.16, 0.16]	0.00	[−0.16, 0.16]	0.00	[−0.14, 0.14]
Performance 2nd half	0.05	[−0.12, 0.22]	0.09	[−0.08, 0.27]	0.08	[−0.11, 0.26]	0.20	[0.02, 0.37]
RT (neutral-neutral)	0.50	[0.33, 0.67]	0.36	[0.19, 0.54]	−0.38	[−0.56, −0.20]	0.41	[0.24, 0.59]

The Dot Probe in an Experimental Context: Predicting Weight Status

Table 6 shows the linear models predicting weight status. Although mean TL-BSneg was found to be a significant predictor of weight status ($B = -0.001$, $SE = 0.001$, $t(328) = -2.03$, $p = 0.04$) none of the models predicting weight status by attention and general response speed reached significance (all Fs < 1, all ps > 0.05), with multiple R-squared and adjusted R-squared ranging between $R^2 = 0.002$ (for mean TL-BSpos) and $R^2 = 0.01$ [for mean TL-BSneg and Abs(PI-0.5)], and between adj$R^2 = -0.004$ (for mean TL-BSpos) and adj$R^2 = 0.008$ [for mean TL-BSneg and Abs(PI-0.5)], respectively.

The Dot Probe in an Individual Differences Context: Associations Adjusted BMI

Table 7 shows Pearson correlations of the attention bias indices with adjusted BMI. None of the correlations reached significance.

DISCUSSION

The present study investigated whether attentional bias toward food could be meaningfully assessed in a youth sample of children with and without obesity, using a pictorial version of the dot probe task. The rationale for this study, was grounded in the widespread practice to use the dot probe procedure to measure and modify food-related attentional bias, both in experimental laboratory studies and in clinical intervention studies (Kemps et al., 2020), despite ample reports of debatable psychometric properties of different dot probe tasks (Schmukle, 2005; Ataya et al., 2012a,b; Parsons et al., 2018; Chapman et al., 2019; MacLeod et al., 2019). Attempting to save the case for the dot probe, we sought to examine the psychometric properties of the task in a comprehensive manner by scrutinizing different indices of attentional bias that could be calculated from our version of the task, with pictures selected from the Food-Pics database (Blechert et al., 2014) administered to children and adolescents aged 7–19, with and without obesity (for a complete description of the task, see Naets et al., 2018). We evaluated the indices meticulously and thoroughly, by testing whether they would be independent of general response speed, whether they would lead to reliable scores in the present sample, whether they could differentiate between different groups for whom we expected differential performance based on theory (i.e., normal weight vs. obesity), and whether they would meaningfully be associated with individual differences in weight. We will discuss the findings on each of these domains.

Because it is known that RT-based scores are often inflated in individuals who are slower in responding (Fazio, 1990; Faust et al., 1999; Greenwald et al., 2003; Glashouwer et al., 2013; De Schryver et al., 2018b), and response speed increases from childhood to adolescence (Dykiert et al., 2012; Adleman et al., 2016), it is imperative to estimate the association of the attentional bias index with general response speed and age. The probabilistic indices aim to account for differences in general response speed, and in this study, they achieved this aim: PI and abs(PI-0.5) were not significantly correlated with mean RT on neutral trials, nor with age. The traditional ABS showed no significant correlation with RT either, but correlated significantly with age: traditional ABS scores were higher for older than for younger participants. However, both abs(ABS) and all dynamic indices, were strongly correlated with both RT and age, with correlations indicative of medium to large effect sizes for RT and small to medium effect sizes for age. The linear models predicting TL-BS in the first test half by mean RT and performance in the second test half, supported the conclusion that TL-BS indices are significantly determined by response speed. The criterium of independence from general response speed was only met by the traditional ABS and the PI scores. Only the PI scores showed independence from age.

Reliability was estimated by comparing the indices calculated on the first half of the task with the indices calculated on the other half. Both traditional and probabilistic indices showed near-zero correlations. The dynamic indices (except the peak TL-BSneg) showed higher and significant correlations, comparable to those reported by their developers (Zvielli et al., 2016). However, the estimates of reliability still did not reach conventional cut-off guidelines (Cronbach, 1951), let alone the recommended 0.90 for individual design research (Rodebaugh et al., 2016). Furthermore, since the linear models predicted that performance in the first test half was largely and significantly determined by reaction time, it can be concluded that these inflated correlations reflect stability in general response speed rather than stability in the attentional process. The criterium of acceptable reliability was met by none of the indices.

Based on theoretical assumptions that food-related attentional processes differ between people with and without eating and weight problems (Appelhans, 2009; Berridge, 2009; Appelhans et al., 2016; Tanofsky-Kraff et al., 2020), significant between-group differences would need to emerge on meaningful indices of attentional bias. However, in the LMMs, there was no support for differential performance on the dot probe task between youth with and without obesity in the present study, irrespective of the index. Furthermore, in the linear models predicting weight status, only mean TL-BSneg emerged as

TABLE 6 | Linear models predicting weight status.

	Index of attentional bias							
	ABS				**Absolute value ABS**			
	Estimate	**St. error**	***t*-value**	***p***	**Estimate**	**St. error**	***t*-value**	***p***
(intercept)	0.47	0.14	3.40	<0.001	0.48	0.14	3.44	<0.001
Index	0.00	0.00	−0.46	0.64	0.00	0.00	1.07	0.29
RT on neutral trials	0.00	0.00	0.90	0.37	0.00	0.00	0.57	0.57
	R^2 0.003		adjR^2 −0.003		R^2 0.006		adjR^2 −0.0001	
	$F(2, 331) = 0.55$	$P = 0.58$			$F(2, 331) = 1.02$	$P = 0.36$		
	Mean TL-BSpos				**Peak TL-BSpos**			
	Estimate	**St. error**	***t*-value**	***p***	**Estimate**	**St. error**	***t*-value**	***P***
(intercept)	0.51	0.16	3.25	0.001	0.47	0.15	3.16	0.002
Index	0.00	0.00	0.40	0.69	0.00	0.00	−0.94	0.35
RT on neutral trials	0.00	0.00	0.30	0.76	0.00	0.00	1.09	0.28
	R^2 0.002		adjR^2 −0.004		R^2 0.004		adjR^2 −0.002	
	$F(2, 328) = 0.35$	$p = 0.70$			$F(2, 328) = 0.71$	$p = 0.49$		
	Mean TL-BSneg				**Peak TL-BSneg**			
	Estimate	**St. error**	***t*-value**	***p***	**Estimate**	**St. error**	***t*-value**	***p***
(intercept)	0.59	0.16	3.81	<0.001	0.50	0.15	3.38	<0.001
Index	−0.001	0.00	−2.03	0.04	0.00	0.00	−1.57	0.12
RT on neutral trials	0.00	0.00	−0.83	0.41	0.00	0.00	0.04	0.97
	R^2 0.01		adj*R*2 0.008		R^2 0.009		adjR^2 0.003	
	$F(2, 328) = 2.33$	$p = 0.09$			$F(2, 328) = 1.51$	$p = 0.22$		
	TL-BS variability							
	Estimate	**St. error**	***t*-value**	***p***				
(intercept)	0.51	0.15	3.37	<0.001				
Index	0.00	0.00	0.70	0.48				
RT on neutral trials	0.00	0.00	0.11	0.91				
	R^2 0.003		adjR^2 −0.003					
	$F(2, 328) = 0.52$	$p = 0.59$						
	PI				**Abs(PI-0.5)**			
	Estimate	**St. error**	***t*-value**	***p***	**Estimate**	**St. error**	***t*-value**	***p***
(intercept)	0.33	0.24	1.40	0.16	0.39	0.14	2.77	<0.01
Index	0.25	0.36	0.68	0.50	1.17	0.60	1.96	0.05
RT on neutral trials	0.00	0.00	1.00	0.32	0.00	0.00	0.96	0.34
	R^2	0.004	adjR^2 −0.002		R^2 0.01		adjR^2 0.008	
	$F(2, 331) = 0.67$	$p = 0.51$			$F(2, 331) = 2.37$	$p = 0.10$		

a significant predictor. However, the model in question (as the other models), did not reach significance, with no more than 1% of variance explained. The criterium of differential performance between groups or predictive validity in terms of group membership was met by none of the indices. As such, there was no evidence that the dot probe task as administered in the present study, could meaningfully be used to assess group differences in experimental research with youth with and without obesity.

Theory (Appelhans, 2009; Berridge, 2009; Appelhans et al., 2016; Tanofsky-Kraff et al., 2020) also states that food-related attentional biases would get stronger in individuals with increasing weight. Although there is some debate on whether this attentional bias would reflect increased approach or increased avoidance (Liu et al., 2019a,b; Hardman et al., 2021), effects are predicted to be significantly correlated with weight parameters. However, none of the indices correlated significantly with adjusted BMI. As such, there was no evidence

TABLE 7 | Correlation of attention bias indices with adjusted BMI.

	Adjusted BMI
ABS	0.03
Absolute value ABS	0.02
mean TL-BSpos	−0.03
peak TL-BSpos	−0.04
mean TL-BSneg	−0.03
peak TL-BSneg	−0.05
TL-BS variability	0.01
PI	0.06
Abs(PI-0.5)	0.05

that the dot probe task as administered in the present study, could meaningfully be used to assess individual differences in food-related attentional bias.

The sobering results of the present study cast doubt on the use of the dot probe procedure as an instrument for assessing maladaptive attentional processes in problematic behavior or psychopathology. However, this need not be the deathblow of the dot probe procedure, since several issues need to be taken into account. Here, the results only pertain to this specific version of the test, with these specific procedural characteristics (e.g., stimuli, presentation times, . . .), When using this test set up, administered to this specific sample (children and adolescents with and without obesity, aged 7–19), to compute these specific indices (ABS, TL-BS, PI), we were unable to provide evidence for the task's applicability to assess food-related attentional biases. However, these conclusions pertain only to this test version, in this sample in this context (De Schryver et al., 2018a), and preclude generalization to other versions of the task in other samples and contexts. Adaptations to the task, that might be worth trying, could be, among others, the use of personally relevant stimuli (Christiansen et al., 2015) or prolonging presentation time (Waechter et al., 2014). Although these adaptations did not result in increased reliability in adult samples (Jones et al., 2018), they were not evaluated in younger samples. Given the impact of test length on reliability (Gulliksen, 1950; Morera and Stokes, 2016; McNally, 2019), one might consider administering more trials. However, the boredom which might be triggered by long repetitive tasks, could potentially be detrimental to attention (Eastwood et al., 2012; Hunter and Eastwood, 2018), especially in younger samples (Hsu et al., 2020). The optimal number of trials, balancing effects on reliability and boredom, still needs to be determined, and would undoubtably depend on the population one is interested in (e.g., age, problem domain). The present study evaluated three indices of attentional bias that are based on differences scores between or differential probability of responding in congruent vs. incongruent trials. Alternative computational methods, like drift-diffusion modeling, are found to yield improved reliability estimates for a verbal dot probe test in adults with clinical anxiety. The index computed following this approach is considered by the authors to be a more precise measure of attentional bias than the traditional ABS (Price et al., 2019). However, this approach has not been evaluated with a pictorial food-related dot probe test, nor in a sample of youth.

To conclude, the present study could not provide evidence for the use of this particular version of the dot probe test to assess food-related attentional bias in youth with and without obesity. These results warn against the ill-considered and casual use of a dot probe task in experimental or correlational research, and again display the need to carefully scrutinize the psychometric properties of the test in the same meticulous way they would evaluate the psychometric properties of other measures (i.e., questionnaires) (Rodebaugh et al., 2016; De Schryver et al., 2018a; Parsons et al., 2018). If researchers would decide on reporting results of the dot probe task, they are urgently and insistently encouraged to also report, evaluate and discuss the psychometric characteristics (e.g., reliability of indices, correlations between general RT and indices) of their test version for their sample.

DATA AVAILABILITY STATEMENT

The raw data supporting the conclusions of this article will be made available by the authors, without undue reservation.

ETHICS STATEMENT

The studies involving human participants were reviewed and approved by Ghent University FPPW. Written informed consent to participate in this study was provided by the participants' legal guardian/next of kin.

AUTHOR CONTRIBUTIONS

LV and TN share project administration. LV and MD contributed to conception and design of the study. LV and MB did the literature search. TN collected data and wrote down the data management plan. MB organized and prepared the database. LV, MB, and MD performed the statistical analysis and wrote sections of the manuscript. LV wrote the first draft of the manuscript. All authors contributed to manuscript revision, read, and approved the submitted version.

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A Dual-Pathway Perspective on Food Choices in Adolescents: The Role of Loss of Control Over Eating

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Introduction: One in three adolescents frequently consume unhealthy snacks, which is associated with negative developmental outcomes. To date, it remains unclear how intrapersonal factors account for food choices in adolescents. Guided by the dual-pathway model, the current study aimed to: (1) examine the joint contribution of inhibitory control and attentional bias in predicting unhealthy food choices in adolescents, and (2) determine whether this mechanism is more pronounced in adolescents who experience loss of control over eating (LOC).

Materials and Methods: A community sample of 80 adolescents (65% female; 10–17 years old, M_{age} = 13.28, *SD* = 1.94) was recruited. Based on a self-report questionnaire, 28.7% of this sample reported at least one episode of LOC over the past month. Food choice was assessed using a computerized food choice task. Both inhibitory control and attentional bias were measured with behavioral tasks (go/no-go and dot probe task, respectively). Binary logistic regressions were conducted to address the research questions.

Results: Inhibitory control and attentional bias did not significantly interact to predict unhealthy food choices. However, there was a significant three-way interaction between inhibitory control, attentional bias and LOC. For adolescents without LOC, the combination of poor inhibitory control and low attentional bias was significantly associated with unhealthy food choice. Surprisingly, for adolescents with LOC, there was no significant association between unhealthy food choice and inhibitory control or attentional bias.

Discussion: Dual-pathway processes do not seem to add to the explanation of food choice behavior for adolescents with LOC. For adolescents who do not experience LOC, those with poor inhibitory control combined with low attentional bias might be at particular risk for making unhealthy food choices.

Keywords: adolescents, food choices, dual-pathway, inhibitory control, attentional bias, loss of control over eating

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INTRODUCTION

Food Choices in Adolescents

The daily consumption of unhealthy snacks is common among adolescents, with prevalence rates up to 23% in Europe (Inchley et al., 2007) and 27% in Flanders (Matthys et al., 2003). Importantly, snacking is responsible for 20–24% of the total energy intake in this age group (De Cock et al., 2016). Unhealthy snacking has been found to be associated with negative physical (e.g., increased risk of overweight and obesity and related medical morbidities such as cardiovascular diseases) and psychosocial outcomes (e.g., depression, poor academic performance) (WHO, 2003; Gopinath et al., 2014; Khalid et al., 2016; Chikwere, 2019). On the contrary, healthy eating can be considered a protective factor due to its associations with a wide range of positive health outcomes (Haines et al., 2019). Consequently, tackling unhealthy eating and improving healthy eating are key public health priorities. Because adolescence is a period when individuals gain more autonomy from parents, and thus assume greater responsibility for their own food choices and behavior, this developmental period is of particular importance to study eating behavior in general, and food choices in particular (Steinberg, 2005). Moreover, adolescence is a period of increasing cognitive maturation, characterized by high reactivity to the environment (e.g., attention for rewarding stimuli) and further development of behavioral regulatory skills (e.g., inhibiting responses) (Crone et al., 2016). Therefore, a more thorough understanding of the underlying mechanisms that drive food choices in adolescents is warranted.

Dual-Pathway Perspective on Food Choices

Transactional models assume that our everyday behavior, such as our eating behavior, is driven by a complex interplay of intrapersonal (e.g., self-regulation) and interpersonal (e.g., food environment) factors (Braet et al., 2014; Lewis, 2014). Notwithstanding the impact of interpersonal factors on eating behavior (for example see Downs and Demmler, 2020), intrapersonal factors are an important target of eating behavior interventions (for example see Köster, 2009; Luis-Ruiz et al., 2020).

A comprehensive theoretical account of the intrapersonal determinants of eating behavior is the dual-pathway perspective (Strack and Deutsch, 2004). This perspective proposes that eating behavior is governed by two interacting systems: *regulatory processes* which are slow and deliberate (e.g., inhibitory control) and *reactive processes* which are fast and effortless (e.g., attentional bias). According to this perspective, unhealthy food choices may be the result of an imbalance between immature regulatory processes (e.g., poor inhibitory control when confronted with palatable food) coupled with strong reactive processes (e.g., automatic attention toward palatable food in the environment) (Steinberg et al., 2018).

Guided by the dual-pathway perspective, researchers have recently found evidence for the *joint contribution* of regulatory and reactive processes to eating behavior. For example, Kakoschke et al. (2015) reported that the combination of poor regulatory processes and strong reactive processes predicted unhealthy food intake from a taste test in adults. Similarly in adolescents, poor regulatory processes coupled with strong reactive processes have been associated with self-reported unhealthy food intake (Stok et al., 2015). In the same vein, both Van Malderen et al. (2020) and Booth et al. (2018) provided evidence in support of a dual-pathway account of self-reported uncontrolled eating among adolescents.

Although the dual-pathway perspective states that both types of processes interact to predict eating behavior, there is already plenty of evidence for the *independent role* of either *poor inhibitory control* (e.g., Nederkoorn et al., 2012; Byrne et al., 2020a) or *high attentional bias* toward food (e.g., Werthmann et al., 2011; Yokum et al., 2011; Folkvord et al., 2015) in predicting eating behavior among adolescents (e.g., uncontrolled eating, unhealthy snacking), whereas studies that investigate the combination of these processes are scarce.

Furthermore, studies have generally not specifically focused on *food choices* as an outcome variable, but rather on broader outcome variables such as food consumption or uncontrolled eating. However, to gain a more comprehensive understanding of the determinants of unhealthy eating behavior, a crucial first step is to identify the factors that contribute to unhealthy food choices in adolescents.

Role of Loss of Control Over Eating in Food Choice

It is unclear whether the central assumptions of the dual-pathway perspective apply to food choices in all adolescents or whether these are particularly pronounced in those who experience early signs of eating-disordered behavior. Specifically, it has been shown that one in three adolescents report *loss of control over eating (LOC)*, which can be defined as the experience of lack of control while eating (He et al., 2016; Van Malderen et al., 2020). LOC is a central feature of binge eating and research has demonstrated that adolescents who experience LOC are at an increased risk for developing negative health outcomes such as overweight and obesity (Goossens et al., 2009a; Shomaker et al., 2010; Tanofsky-Kraff et al., 2011). Moreover, longitudinal research has shown that one episode of LOC in adolescents may be considered an early sign of eating disordered behavior given its prospective value for the development of clinical eating disorders (e.g., Bulimia Nervosa) and other types of psychopathology (e.g., depression, addiction) (Tanofsky-Kraff et al., 2011; Herpertz-Dahlmann et al., 2015), thereby emphasizing its clinical significance. Importantly, previous research has provided evidence for the dual-pathway perspective in predicting LOC among adolescents (Booth et al., 2018; Van Malderen et al., 2020). In addition, it has been shown that adolescents who experience LOC eat more palatable food and make more unhealthy food choices (Dalton et al., 2013; Ng and Davis, 2013; Byrne et al., 2020b). Both findings highlight the importance of taking into account *how* one feels while eating (i.e., food experience) alongside *what* one chooses to eat (i.e., food choices). Thus, in investigating food choice behavior in adolescents from a dual-pathway perspective, it may be important to include LOC as a moderator to distinguish adolescents with LOC from those without LOC.

Current Study

The current study aimed to investigate the underlying mechanisms of food choice behavior in adolescents. To this end, the study addressed two main research questions. First, based on the dual-pathway perspective, we examined the interaction between regulatory (i.e., inhibitory control) and reactive (i.e., attentional bias) processes in predicting food choice in adolescents. Guided by previous empirical evidence (e.g., Kakoschke et al., 2015; Stok et al., 2015; Booth et al., 2018; Van Malderen et al., 2020), we expected that the combination of poor inhibitory control and high attentional bias would be associated with the greatest risk of unhealthy food choices among adolescents. Second, to determine whether dual-pathway assumptions apply to food choice behavior of all adolescents or may be more pronounced in those with early signs of eating-disordered behavior (for example see Herpertz-Dahlmann et al., 2015), LOC was included as an additional moderator. As LOC has previously been associated with the dual-pathway processes (Booth et al., 2018; Van Malderen et al., 2020), as well as with unhealthy food choices (Dalton et al., 2013; Ng and Davis, 2013; Byrne et al., 2020b), it was hypothesized that the interaction between regulatory and reactive processes would be more pronounced in adolescents with LOC compared to those without LOC.

MATERIALS AND METHODS

Participants and Procedure

The sample consisted of 80 participants, recruited from the general population. Participants were contacted by 3rd year psychology students (in the context of a practical course). Each student was instructed to recruit two participants between 10 and 18 years old (there were no other in- or exclusion criteria). We based this age range on the commonly used definition of adolescence in the literature (i.e., the transitional period between childhood and adulthood) (Sawyer et al., 2018). In the final sample, participants were between 10 and 17 years old (M_{age} = 13.28, *SD* = 1.94) and 65% (*N* = 52) of the sample was female. Data collection occurred during a home visit, and consisted of two parts. First, participants were presented with several online questionnaires. Second, participants completed two computer tasks the order of which was counterbalanced (i.e., go/no-go task and dot probe task), followed by a computerized food choice task (see section "Materials"). The total duration of each home visit was approximately 2 h. All adolescents and their parents signed an active informed consent and the entire study protocol was approved by the Faculty Ethics Committee. In the informed consent, the study was described as investigating risk and protective factors for the development of psychological problems during adolescence. No incentives were provided for participation. The study was part of a broader project on eating behavior among adolescents and some of the data have been reported previously (Van Malderen et al., 2019, 2020). The focus of the current study was on the dual-pathway predictors of food choice among adolescents, whereas the other studies focused on the role of affectivity (Van Malderen et al., 2019) and self-regulation (Van Malderen et al., 2020) in loss of control over eating in adolescents. Consequently, only participants who had completed the food choice task (*N* = 80) were included in the current sample.

Materials

Control Variables

Participants self-reported their age and gender. During the home visit, height and weight were objectively measured (using a tape measure and scales). An adjusted body mass index was calculated {[actual body mass index (kg/m^2)/percentile 50 of body mass index for age and gender] × 100} (Roelants and Hauspie, 2004; Rolland-Cachera et al., 2015). Because food choice may be influenced by age, gender, and adjusted body mass index, these were included as control variables in all analyses (Manippa et al., 2017; Andrade et al., 2019; Perrar et al., 2020).

Food Choice

Food choice was assessed with a computerized *food choice task* (see Veling et al., 2013). In this task, participants were presented with a 4 × 4 square grid with 16 pictures of snacks on a computer screen and asked to select eight items that they would like to take home. There were eight healthy snacks (i.e., carrots, gingerbread, health bars, fruit salad, apple, muesli bars, crackers, rice cake) and eight unhealthy snacks (i.e., potato chips, chocolate, muffin, salted nuts, cheese balls, M&M's, chocolate chip cookies, cookies). The pictures for this task were derived from Veling et al. (2013) who validated these in terms of palatability and healthiness in an independent sample of participants. Importantly, the food pictures in this task represented the same broad categories as the food pictures of the tasks that capture dual-pathway processes (see below). The time limit for making food selections was 15 s. Following previous research (e.g., Furst et al., 1996; Kakoschke et al., 2017), the outcome measure of interest was the first snack item chosen (0 = healthy snack, 1 = unhealthy snack). This ensured that an "automatic" decision was captured.

Dual-Pathway Processes

The *"go/no-go task" (GNG)* was used as a measure of *regulatory processing*, and more specifically *inhibitory control* (see Kakoschke et al., 2015). In this task, participants were presented with two blocks of 160 trials. Each trial had a duration of 1500 ms in which a picture was shown coupled with either a "go" cue (e.g., the letter "p") to which participants responded by pressing the space bar, or a "no-go" cue (e.g., the letter "f") which signaled that participants should withhold their response. The "go" and "no-go" cues appeared randomly at one of the four corners of the picture. Both types of cues appeared equally often during the task and were counterbalanced (i.e., for some participants the "go" cue was the letter "f" and for others it was the letter "p"). The pictures were taken from the food-pics database of Blechert et al. (2014). Specifically, the pictures consisted of images of 20 palatable foods (e.g., chips, chocolate) and 20 non-foods (i.e., animals rated to be of similar appeal) (e.g., giraffe, butterfly). The outcome measure was the number of commission errors (i.e., CE; space bar pressed in response to a "no-go" cue) (e.g., see Meule and Kübler, 2014). A higher number of commission errors on food pictures (CE_{food}) reflects poorer inhibitory control toward food.

To measure *reactive processing*, and specifically *attentional bias*, the *"dot probe task" (DP)* was used (see Kakoschke et al., 2014). The task consisted of 258 trials. Each trial commenced with a fixation cross presented in the middle of the screen (for 500 ms), followed by two pictures presented simultaneously on the left and right hand sides of the screen (also for 500 ms). Next, a dot (probe) was presented in the location of one of the two pictures. The participant's task was to indicate as quickly as possible whether the dot (probe) appeared on the left or right hand side of the screen by pressing a key on an AZERTY-keyboard ("W" and "N," respectively). Again, the pictures were sourced from the food-pics database of Blechert et al. (2014). Two pairs of stimuli were used: food versus neutral non-food (32 experimental pairs) and neutral non-food versus neutral non-food (16 control pairs). For the experimental pairs, household objects were chosen as the neutral non-food category. Only these pairs were used to calculate an attentional bias score for inclusion in the analyses. The stimuli for the control pairs consisted of animals, because like food, animals are overall appealing. Household objects and animals are commonly used neutral non-food categories in attentional bias research (e.g., Kemps et al., 2014; Liu et al., 2019). The pictures in each pair were matched on color and shape. Picture pairs were presented in a new random order for each participant. The dot (probe) appeared equally often on both sides of the screen. The outcome measure was reaction time (RT; in milliseconds). Attentional bias scores (AB) were computed from the experimental trials by subtracting the RT on trials where the probes replaced the food pictures from the RT on trials where the probes replaced the neutral non-food pictures. A positive score reflects an attentional bias toward food pictures and a negative score an attentional bias away from food pictures.

Loss of Control Over Eating (LOC)

The experience of loss of control over eating (LOC) was assessed with the Dutch translation and adaptation of the "Children's Eating Disorder Examination Questionnaire" (ChEDE-Q; Fairburn and Beglin, 1994; Decaluwé and Braet, 1999). This self-report questionnaire consists of four underlying subscales (i.e., restrictive eating, concerns about eating, weight, and shape), and in addition assesses different types of uncontrolled eating episodes (i.e., objective binge eating, subjective binge eating). For the current study, only the questions assessing uncontrolled eating episodes were used. Participants were first asked if they had experienced that type of eating episode over the past month (yes/no). If yes, the total number of such episodes over the past month was determined. The variable of interest was whether or not participants had experienced at least one episode of uncontrolled eating over the past month (0 = no LOC episode over the past month, 1 = at least one LOC episode over the past month). This operationalization (i.e., one episode over a 1-month time frame) is in line with other such studies among adolescents in the general population (e.g., Tanofsky-Kraff et al., 2011, 2020; Kelly et al., 2016). Previous research has shown that the ChEDE-Q is a valid and reliable measure of LOC in adolescents (Decaluwé et al., 2003; Van Durme et al., 2015).

Statistical Analysis

To test the interaction between regulatory (i.e., inhibitory control) and reactive (i.e., attentional bias) processing in predicting food choice, a binary logistic regression was conducted. First, food choice was entered as a categorical dependent variable (0 = healthy snack, 1 = unhealthy snack). Second, age, gender and adjusted body mass index were entered as control variables. Third, the main effects of inhibitory control (i.e., CE_{food}), attentional bias (i.e., AB), and their interaction (i.e., $CE_{food} \times AB$) were included as independent variables.

To investigate whether the dual-pathway perspective applies to food choice in adolescents in general or is more pronounced in those with existing disturbed eating behavior, an additional binary logistic regression analysis was performed. This analysis was identical to the first regression, but with LOC included as an additional categorical moderator (0 = no LOC episode over the past month, 1 = at least one LOC episode over the past month) in the interaction term that was added in the last step (i.e., $CE_{food} \times AB \times LOC$). To ascertain the robustness of any interaction effect, the analyses were also performed without covariates (i.e., age, gender, and adjusted body mass index).

Only the full logistic regression models (including the control variables and all independent variables) are displayed (see **Tables 2**, **3**) in the Results. Significant interactions were interpreted by comparing the means between the different groups using independent sample *t*-tests. An alpha value of $p \leq 0.05$ was used to determine statistically significant effects and odds ratios (OR) were reported as effect sizes for all analyses. The analyses were conducted with SPSS version 24.0.

RESULTS

Descriptive Statistics

The mean adjusted body mass index of the sample was 100.32 (SD = 17.41), and ranged from adolescents having underweight (minimum = 58.62) to adolescents having obesity (maximum = 163.47). Specifically, 11% of the sample was classified as having underweight (adjusted body mass index ≤85), 79% as having a normal weight (85 <adjusted body mass index <120), 5% as having overweight (120 <adjusted body mass index <140), and 5% as having obesity (adjusted body mass index ≥140).

Mean number of commission errors on the food pictures of the go/no-go task was low (M = 3.10, SD = 4.67), reflecting good overall inhibitory control capacities toward food pictures. However, there was a large degree of variability in the number of commission errors, ranging from 0 to 37 (also see Kakoschke et al., 2015; Van Malderen et al., 2020). The mean attentional bias score from the dot probe task was positive ($M = 1.07$, $SD = 21.51$). However, a one sample t-test showed that this was not significantly different from zero, $t(79) = 0.44$, $p = 0.658$. The standard deviation was again large, indicating substantial variability across participants (ranging from −50.75 to 61.96). In total, 28.7% ($N = 23$) of participants reported at least one episode of LOC over the past month (according the ChEDE-Q; Fairburn and Beglin, 1994; Decaluwé and Braet, 1999). Among those, the number of episodes ranged from 1 to 20 (*Median* = 2.00, *Mean* = 4.04, $SD = 4.80$). Most adolescents in that group reported 1 (21.7% or $N = 5$) or 2 episodes of LOC (34.9% or $N = 8$). Furthermore, 13.1% adolescents ($N = 3$) reported 4 episodes of LOC, 13.1% ($N = 3$) 3 episodes, and 4.3% ($N = 1$) 5, 11, 15, and 20 episodes of LOC over the past month. In the food choice task, 48.8% ($N = 39$) of participants chose an unhealthy food first. **Table 1** gives an overview of all sample characteristics and correlations between the variables of interest.

Main Analyses

The first binary logistic regression analysis which tested the interaction between regulatory (i.e., inhibitory control) and reactive (i.e., attentional bias) processing in predicting food choice was not significant [$\chi^2(6) = 7.16$, $p = 0.307$], and revealed no significant main or interaction effects (see **Table 2**). Without the covariates, the analysis was also not significant [$\chi^2(3) = 1.15$, $p = 0.766$], and revealed no significant main or interaction effects.

The second binary logistic regression analysis which investigated whether the dual-pathway perspective may be more pronounced in adolescents with LOC was significant [$\chi^2(7) = 14.27$, $p = 0.047$], and revealed a significant three-way interaction ($p = 0.043$) (see **Table 3**). Without the covariates, the analysis was trend significant [$\chi^2(4) = 8.85$, $p = 0.065$], and again revealed a significant three-way interaction ($p = 0.035$). Specifically, inhibitory control (CE_{food}), attentional bias (AB) and LOC significantly interacted to predict unhealthy food choice. **Figure 1** shows this three-way interaction. As can be seen, two different patterns emerged for the LOC-group and the NoLOC-group.

In the *NoLOC-group* (left panel), participants with low AB scores (white bars) were significantly more likely to choose an unhealthy snack first when they also had high levels of CE_{food} (i.e., weaker inhibitory control toward food) ($M = 0.714$, $SD = 0.202$) compared to when they had low levels of CE_{food} (i.e., stronger inhibitory control toward food) ($M = 0.493$, $SD = 0.152$), [$t(29) = 3.46$, $p = 0.002$]. For participants with high AB scores (gray bars), there was no significant association between level of CE_{food} and the likelihood of choosing an unhealthy snack first [$t(24) = -1.84$, $p = 0.078$]. Thus, for participants with high levels of CE_{food} (i.e., weaker inhibitory control toward food), the likelihood of choosing an unhealthy food first was significantly

TABLE 1 | Descriptive statistics and correlations.

Total sample	*N*	*M* (*SD*) or %	Min – Max	LOC[a]	Age	AdjBMI	CE_{food}	AB
Gender	80	65% female						
Food choice	80	48.8% unhealthy						
LOC	80	28.7% LOC	1 – 20	1				
Age	80	13.28 (1.94)	10 – 17	0.15	1			
AdjBMI	80	100.32 (17.41)	58.62 – 163.47	0.18	−0.07	1		
CE_{food}	80	3.10 (4.67)	0 – 37	−0.02	−0.21	0.04	1	
AB	80	1.07 (21.51)	−50.75 – 61.96	−0.07	0.01	−0.14	−0.08	1
LOC-Group[b]	*N*	*M* (*SD*) or %	Min – Max		Age	AdjBMI	CE_{food}	AB
Gender	23	69.6% female						
Food choice	23	47.8% unhealthy						
Age	23	13.74 (1.96)	10 – 17		1			
AdjBMI	23	105.16 (18.44)	86.05 – 163.47		−0.09	1		
CE_{food}	23	3.78 (7.54)	0 – 37		−0.22	0.10	1	
AB	23	−0.05 (25.39)	−45.06 – 61.96		−0.13	−0.10	0.05	1
NoLOC-Group[b]	*N*	*M* (*SD*) or %	Min – Max		Age	AdjBMI	CE_{food}	AB
Gender	57	63.2%						
Food choice	57	49.1% unhealthy						
Age	57	13.09 (1.91)	10 – 17		1			
AdjBMI	57	98.37 (16.75)	58.62 – 152.24		−0.10	1		
CE_{food}	57	2.82 (2.85)	0 – 12		−0.28*	−0.06	1	
AB	57	1.52 (19.97)	−50.75 – 51.24		−0.09	−0.16	−0.26	1

*LOC, Loss of Control over Eating; AdjBMI, Adjusted Body Mass Index; CE_{food}, Commission Errors on Food Pictures; AB, Attentional Bias Score. [a] These correlations are Spearman's correlations; all other correlations are Pearson's correlations. [b] There were no significant group differences between the LOC-Group and the NoLOC-Group on any of these variables. *$p \leq 0.050$.*

TABLE 2 | Logistic regression analysis: Inhibitory control × attentional bias in predicting unhealthy food choice.

	Wald χ^2	*B (SE)*	*p*	*OR*
Covariates:				
Gender	2.09	0.75 (0.52)	0.149	2.11
Age	2.93	0.22 (0.13)	0.087	1.24
AdjBMI	0.45	−0.01 (0.01)	0.504	0.99
CE_{food}	0.09	−0.02 (0.06)	0.759	0.98
AB	0.02	−0.00 (0.02)	0.889	1.00
CE_{food} × AB	0.02	0.00 (0.00)	0.892	1.00
Model test	χ^2 (6) = 7.16, p = 0.307			
−2LL (Nagelkerke R^2)	103.70 (0.11)			

OR, Odds Ratio; AdjBMI, Adjusted Body Mass Index; CE_{food}, Commission Errors on Food Pictures; AB, Attentional Bias Score.

TABLE 3 | Logistic regression analysis: Inhibitory control × attentional bias × LOC in predicting unhealthy food choice.

	Wald χ^2	*B (SE)*	*p*	*OR*
Covariates:				
Gender	1.74	0.70 (0.53)	0.187	2.02
Age	2.38	0.20 (0.13)	0.123	1.22
AdjBMI	0.66	−0.01 (0.02)	0.418	1.01
CE_{food}	0.29	−0.04 (0.07)	0.592	1.01
AB	0.15	−0.01 (0.02)	0.703	1.00
CE_{food} × AB	2.00	0.01 (0.01)	0.158	1.01
CE_{food} × AB × LOC	4.09	−0.02 (0.01)	0.043*	1.00
Model test	χ^2 (7) = 14.27, p = 0.047*			
−2LL (Nagelkerke R^2)	96.59 (0.22)			

*OR, Odds Ratio; AdjBMI, Adjusted Body Mass Index; CE_{food}, Commission Errors on Food Pictures; AB, Attentional Bias Score; LOC, Loss of Control over Eating. *$p \leq 0.050$.*

greater for those with low (M = 0.714, SD = 0.202) compared to high (M = 0.370, SD = 0.140) AB scores [$t(22) = -4.49$, $p \leq 0.001$].

In the *LOC-group* (right panel), there was no significant association between level of CE_{food} and the likelihood of choosing an unhealthy snack first, neither for participants with low AB scores (white bars) [$t(10) = 0.60$, $p = 0.561$], nor for those with high AB scores (gray bars) [$t(9) = 1.61$, $p = 0.142$].

DISCUSSION

Guided by the dual-pathway model, the current study aimed to investigate the interaction between inhibitory control and attentional bias in predicting unhealthy food choices in adolescents. An additional goal was to determine whether this dual-pathway perspective was more pronounced in adolescents with early signs of eating-disordered behavior and specifically those who experience LOC (for example see Herpertz-Dahlmann et al., 2015). By addressing these two research questions, this study sought to contribute to the underlying mechanisms that drive food choice behavior in adolescents.

Based on previous studies and the theoretical dual-pathway perspective (e.g., Kakoschke et al., 2015; Stok et al., 2015; Booth et al., 2018; Van Malderen et al., 2020), we expected a significant interaction between poor inhibitory control and strong attentional bias in predicting unhealthy food choice in adolescents. Contrary to expectation, there was no significant interaction between inhibitory control and attentional bias in predicting food choice (research question 1). However, the inclusion of LOC as an additional moderator revealed a significant three-way interaction between inhibitory control, attentional bias and LOC (research question 2). This shows that the relationship between inhibitory control and attentional bias in predicting unhealthy food choice depends on whether adolescents experienced LOC over the past month. Surprisingly, the direction of this three-way interaction was not in line with dual-pathway predictions and – contrary to our expectations – it was not more pronounced in adolescents with LOC compared to those without LOC (for example see Booth et al., 2018; Byrne et al., 2020b).

In particular, adolescents *without LOC* were more likely to choose an unhealthy food first when they exhibited a combination of *poor inhibitory* control and *low attentional bias*. This result is at odds with dual-pathway predictions that the combination of *poor inhibitory control* and *high attentional bias* would be associated with unhealthy food choices. It also contradicts previous empirical evidence for this dual-pathway perspective in the context of overweight or unhealthy eating in children (e.g., Kemps et al., 2020), adolescents (e.g., Stok et al., 2015), and adults (e.g., Kakoschke et al., 2015). Nevertheless, the current result is in line with other previous observations in adults (Manasse et al., 2015) as well as in adolescents (Van Malderen et al., 2018) which have also found an interaction between poor inhibitory control and low attentional bias. However, it should be noted that all these previous studies focused on overweight or unhealthy eating as the outcome variable and not food choice specifically. In particular, these results seem to indicate that, in adolescents without LOC, the level of attentional bias determines the extent to which inhibitory control capacities contribute to unhealthy food choices. One possible explanation might be that adolescents with *low attentional bias* are less preoccupied with food in general (for example see Brignell et al., 2009; Hardman et al., 2020). When those adolescents also have good inhibitory control capacities, they are then able to go for a healthy food option. However, when this low attentional bias is coupled with poor inhibitory control, their poor regulatory abilities increase their risk of choosing unhealthy food. On the other hand, adolescents with *high attentional bias* are generally more preoccupied with food in the environment, regardless of their inhibitory control capacities (Hendrikse et al., 2015; Kemps et al., 2020). As a result, they are not more nor less likely to make unhealthy food choices depending on their level of inhibitory control. An alternative explanation might be that adolescents with *high attentional bias* deliberately avoid palatable food in the environment and intentionally prefer a healthy snack over an unhealthy one. This adaptive avoidant strategy may explain

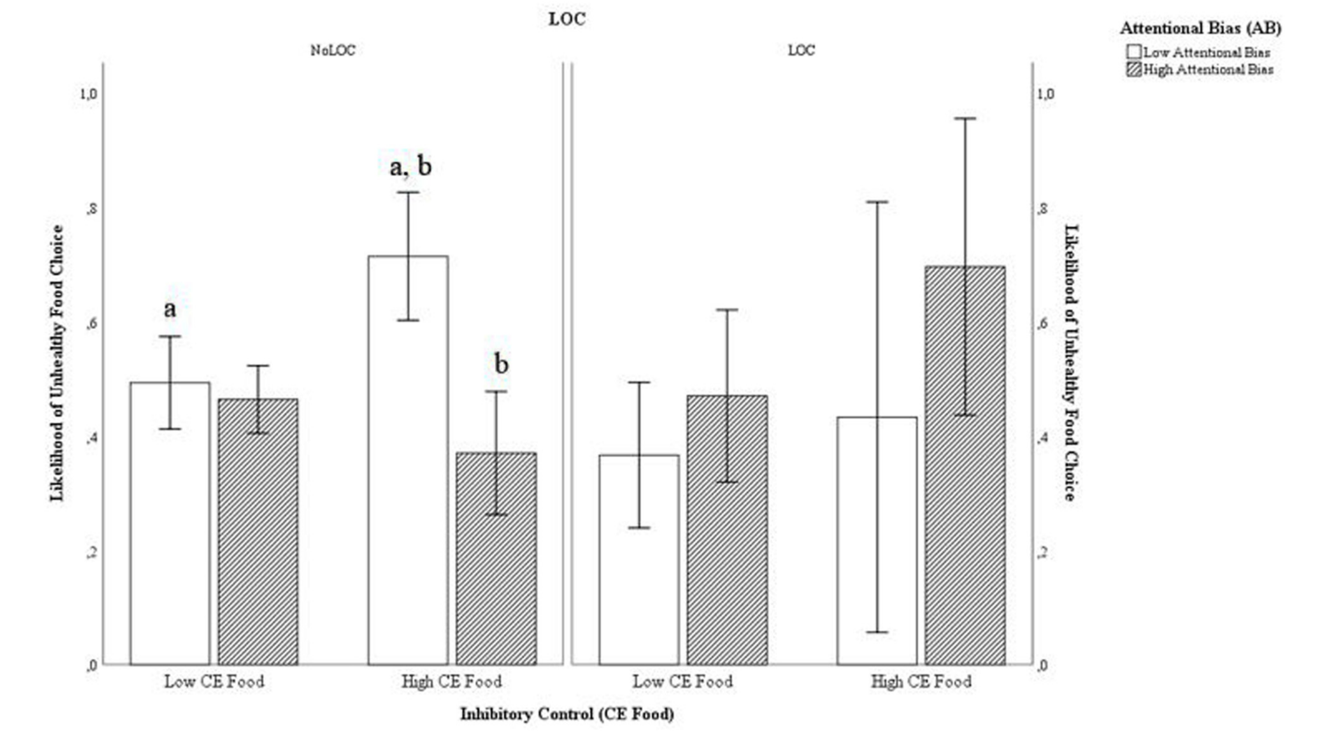

FIGURE 1 | Three-way interaction between inhibitory control (CE_{food}), attentional bias (AB), and LOC in predicting unhealthy food choice. CE_{food}, Commission Errors on Food Pictures; AB, Attentional Bias Score; LOC, Loss of Control over Eating. a: $t(29) = 3.46$, $p = 0.002$**; b: $t(22) = -4.49$, $p \leq 0.001$***. Error Bars: 95% confidence interval.

the finding that, in the NoLOC-group, adolescents with poor inhibitory control are significantly less likely to choose an unhealthy food when they have a strong attentional bias for food. Yet another explanation for the unexpected direction of the interaction in the NoLOC-group, might be that this group is not homogeneous and consists of several important subtypes (e.g., depending on adjusted BMI, temperament, environmental factors) which may have contributed the results (for example see Kubik et al., 2003; Davis and Fox, 2008; De Cock et al., 2016). Future research may distinguish between these possible subtypes by including them as additional moderating variables.

In addition, we hypothesized that poor inhibitory control combined with a strong attentional bias (i.e., dual-pathway perspective) would be particularly associated with unhealthy food choices in adolescents who experience LOC. However, adolescents *with LOC* were not significantly more nor less likely to choose an unhealthy food regardless of their levels of inhibitory control or attentional bias. As an extension of previous evidence for a dual-pathway account of LOC among adolescents (e.g., Booth et al., 2018; Van Malderen et al., 2020), the current findings seem to indicate that this vulnerability does not have additional explanatory value in the context of food choices. A possible explanation might be that the LOC-group is characterized by considerable variation regarding the types of uncontrolled eating episodes. For example, at one particular time one can experience LOC while eating objectively large amounts of food (i.e., objective binge eating episodes) during which it would be expected that unhealthy food would be preferred (for example see Ng and Davis, 2013; Byrne et al., 2020b). In contrast, at other times one can experience LOC while eating an amount of food that is considered to be large only according to the individual but not to others (i.e., subjective binge eating episodes). This subjective type of uncontrolled eating behavior is often accompanied by eating subjectively large amounts of (healthy or unhealthy) foods (Fitzsimmons-Craft et al., 2014). This variation in types of uncontrolled eating behavior might have obscured any effect of inhibitory control and attentional bias on LOC in general.

Strengths, Limitations and Directions for Future Research

The current study has several important strengths. First, the sample consisted of adolescents. This is a crucial population for studying eating behavior in general and food choices in particular because of the well-known risk of developing eating problems during that age period (Steinberg, 2005). Second, although the dual-pathway perspective emphasizes the importance of investigating the interaction between inhibitory control and attentional bias, most research has focused on these processes individually. The current study adds to the literature specifically by examining their combined contribution. Third, this study focused specifically on food choice instead of broader eating-related outcome variables (e.g., food consumption). This approach enables the further disentanglement of the underlying mechanisms that drive unhealthy food choices, which in turn, contribute to our understanding of unhealthy eating behavior. Finally, the assumptions of the dual-pathway perspective were tested in adolescents with versus those without early signs of eating-disordered behavior, namely LOC (for example see Tanofsky-Kraff et al., 2007; Goossens et al., 2009b; Van Malderen et al., 2020). In making this distinction, the present results may help shed light on whether the theoretical dual-pathway perspective applies to adolescents in general or to particular subgroups.

There are also some noteworthy limitations that should be acknowledged. First, the sample size of the LOC-group was quite small. In addition, the current study did not distinguish between different types of uncontrolled eating episodes (e.g., objective versus subjective binge eating episodes). A larger sample may afford the ability to distinguish between several types of uncontrolled eating episodes and ascertain any clinically relevant effects that were precluded in the current study due to a lack of power. In addition, a larger sample size would allow for the inclusion of other important control variables in the context of food choice behavior, such as educational level or household income (Baumann et al., 2019). Furthermore, the percentage of adolescents with overweight (5%) and obesity (5%) in the present study was lower compared to prevalence rates in the general population (World Health Organization, 2016), limiting the generalizability of the findings.

Second, the study design was cross-sectional, precluding any causal inferences. Moreover, the current design carries the risk of limited ecological validity, making replication in a real-life setting an important goal for future research (e.g., assessing food choices from a real-life food buffet, conducting ecological momentary assessment).

Third, the food choice task was limited in the number and types of food presented. Because food preferences are known to be personal (Birch, 1999; Giese et al., 2015; Robino et al., 2019), a next step would be to personalize the food choice task or use a real-life food buffet.

Fourth, the dual-pathway distinction between regulatory and reactive processes is not entirely clear-cut because these two types of processing cannot be strictly separated (i.e., slow and deliberate regulatory processes versus fast and effortless reactive processes). In particular, it could be argued that the behavioral task that was used to measure regulatory processing (i.e., go/no-go task) not only assesses slow and deliberate processes but also captures fast and effortless processing. However, this task has been widely used to measure regulatory processing in eating behavior research (e.g., Ames et al., 2014; Kakoschke et al., 2015; Veling et al., 2017; Jones et al., 2018). Nevertheless, future research could usefully consider other tasks to capture slow and deliberative processes. Relatedly, it should be noted that the attentional bias scores derived from the dot probe task included both negative (indicating an attentional avoidance) and positive (indicating attentional bias) scores. Although these were roughly equally distributed across our sample, future research in larger samples could seek to distinguish these two types of attentional processes and their relationships with food choice behavior.

Finally, the current study did not measure hunger. As hunger level could have influenced both the cognitive processing of food items (in the go/no-go and dot probe tasks) and the selection of energy-dense foods (in the food choice task), future research should endeavor to measure hunger at the start of the testing session.

Theoretical and Clinical Implications

From a *theoretical* perspective, the current results may add to the dual-pathway model as a way of understanding intrapersonal determinants of eating behavior in adolescents. The current results should be considered preliminary and replicated in larger samples to substantiate *clinical implications*. In this context, the different results for the LOC-group and the NoLOC-group seem to indicate that the dual-pathway vulnerability does not apply to all adolescents. Specifically, for adolescents who experience LOC, the dual-pathway processes do not seem to have an added explanatory value in the context of food choices. However, for adolescents who do not experience LOC, screening efforts based on dual-pathway processes might be valuable. Adolescents with poor inhibitory control combined with low attentional bias might be particularly at risk of making unhealthy food choices.

CONCLUSION

In conclusion, the current study did not find evidence for the dual-pathway perspective in predicting food choice behavior in adolescents. However, dual-pathway processes interacted with LOC in predicting food choice behavior. In adolescents with LOC, dual-pathway processes do not seem to have additional explanatory value. However, for adolescents who do not experience LOC, those with poor inhibitory control combined with low attentional bias might be particularly at risk of making unhealthy food choices. These findings provide an important next step in understanding food choice behavior in adolescents.

DATA AVAILABILITY STATEMENT

The raw data supporting the conclusions of this article will be made available by the authors, upon request.

ETHICS STATEMENT

The studies involving human participants were reviewed and approved by the Ghent University Ethics Committee. Written informed consent to participate in this study was provided by the participants' legal guardian/next of kin. Written informed consent was obtained from the minor(s)' legal guardian/next of kin for the publication of any potentially identifiable images or data included in this article.

AUTHOR CONTRIBUTIONS

LG, EK, SV, LC, and EV designed the study and wrote the protocol. EV was responsible for data collection, under the supervision of LG. EV conducted the statistical analyses and wrote the first draft of the manuscript. All authors edited subsequent drafts and have approved the final manuscript.

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Pushing or Pulling Your "Poison": Clinical Correlates of Alcohol Approach and Avoidance Bias Among Inpatients Undergoing Alcohol Withdrawal Treatment

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Introduction: Alcohol approach bias, the tendency to automatically move toward alcohol cues, has been observed in people who drink heavily. However, surprisingly, some alcohol-dependent patients demonstrate an alcohol avoidance bias. This inconsistency could be explained by the clinical or demographic profile of the population studied, yet this has not been examined in approach bias modification (ABM) trials to date. We aimed to determine the proportion of patients with an approach or avoidance bias, assess whether they differ on demographic and drinking measures, and to examine the clinical correlates of approach bias.

Method: These research questions were addressed using baseline data from 268 alcohol-dependent patients undergoing inpatient withdrawal treatment who then went on to participate in a trial of ABM.

Results: At trial entry (day 3 or 4 of inpatient withdrawal), 155 (57.8%) had an alcohol approach bias and 113 (42.2%) had an avoidance bias. These two groups did not differ on any demographic or relevant drinking measures. Approach bias was significantly and moderately associated with total standard drinks consumed in the past 30 days ($r = 0.277$, $p = 0.001$) but no other indices of alcohol consumption or problem severity.

Conclusion: Whilst the majority of alcohol-dependent patients showed an alcohol approach bias, those with an avoidance bias did not differ in demographic or clinical characteristics, and the strength of approach bias related only to recent consumption. Further research is needed to develop more accurate and personally tailored measures of approach bias, as these findings likely reflect the poor reliability of standard approach bias measures.

Keywords: approach bias, avoidance bias, cognitive bias modification, approach bias modification, alcohol use disorder, measurement

INTRODUCTION

Contemporary neurocognitive models of addiction posit that automatic cognitive processes play a critical role in the maintenance of alcohol use disorders (AUDs) (Stacy and Wiers, 2010). Research has demonstrated that alcohol-related cues can automatically capture attention (i.e., known as an "attentional bias"; Field and Cox, 2008), and that these cues can trigger automatic action tendencies to approach alcohol (i.e., known as an "approach bias"; Field et al., 2008). Alcohol-related attentional and approach biases are thought to develop through extended periods of frequent drinking, which involve numerous associative learning experiences in which the rewarding effects of alcohol are paired with various alcohol-related cues (Field and Cox, 2008). These associations become sensitised (i.e., very easily, rapidly, and strongly activated) such that re-exposure to these cues (or even memory of them) is quickly and easily able to activate mental representations of alcohol's desired effects, influencing attentional and approach biases (Stacy and Wiers, 2010) and leading to alcohol consumption (Field and Cox, 2008; Martin-Braunstein et al., 2016). Whilst alcohol-dependent individuals are theorised to demonstrate an approach bias for alcohol-related cues (Field et al., 2008), there is inconsistent evidence about the proportion who actually do (Spruyt et al., 2013; Ernst et al., 2014), and studies have failed to investigate whether the presence of approach bias is limited to individuals with certain clinical or demographic profiles.

Approach bias is typically measured through behavioural reaction tasks which assess whether individuals are faster to respond to substance-related stimuli displayed on a computer screen compared to their response to neutral stimuli. Commonly used tasks are the approach avoidance task (AAT; Rinck and Becker, 2007), in which participants move stimuli towards (approach) and away (avoid) from themselves using a joystick, and the stimulus–response compatibility task (SRC; De Houwer et al., 2001), where participants perform a symbolic movement by making a manikin walk towards (approach) or away (avoid) from stimuli (Kersbergen et al., 2015). Faster RTs for approaching vs. avoiding alcohol stimuli relative to non-alcohol stimuli indicate an approach bias (positive score), and the opposite indicates an avoidance bias (negative score; Kersbergen et al., 2015). Both relevant-feature (R) and irrelevant-feature (IR) versions of the AAT and SRC exist, where relevant-feature tasks instruct participants to respond to the explicit contents of the presented stimuli (e.g., avoid images containing alcohol, approach images containing soft-drinks), whilst irrelevant-feature tasks provide a more implicit assessment of approach bias by requiring participants to respond to an extraneous image feature (e.g., avoid images presented in landscape orientation, approach images presented in portrait orientation; De Houwer, 2003; Wiers et al., 2009, 2017).

People experiencing alcohol problems have been shown to demonstrate stronger alcohol approach bias compared to those without problems (Sharbanee et al., 2013; Ernst et al., 2014; Wiers et al., 2014). For example, Ernst et al. (2014) compared 21 alcohol-dependent inpatients to 21 matched controls and found that patients demonstrated a stronger alcohol approach bias compared to controls, and that patients demonstrated stronger neural activations of reward circuitry when approaching rather than avoiding alcohol pictures, where the reverse was found for healthy controls (Ernst et al., 2014). Alcohol approach bias has also been associated with higher levels of self-reported craving and weekly alcohol consumption (Field et al., 2008). However, contrary to predictions, some studies have found that alcohol-dependent patients in residential settings report an overall avoidance bias at baseline (Spruyt et al., 2013; Snelleman et al., 2015; Field et al., 2017; Rinck et al., 2018). For example, one study with 40 abstaining alcohol-dependent patients (18–21 days after drinking) found that they demonstrated a relative alcohol avoidance bias at baseline (using the R-SRC; Spruyt et al., 2013), while another study found that participants (120 recently detoxified alcohol-dependent patients) demonstrated no overall mean approach or avoidance bias at baseline (Field et al., 2017). However, it is worth noting that performance on the R-SRC is not correlated with performance on the IR-AAT (Wiers et al., 2013), which could explain the differences in findings.

We recently conducted a large trial of approach bias modification (ABM) where we trained alcohol-dependent patients to push away alcohol cues and hence reduce their alcohol approach bias (Manning et al., 2021). We found that these participants had a mean approach bias on the IR-AAT at baseline, which is consistent with some other large ABM trials of alcohol-dependent patients (Wiers et al., 2011; Eberl et al., 2013). In none of these studies do we know what proportion of participants had an avoidance or approach bias on treatment entry (since only mean approach bias scores are reported), nor do we know whether those with an approach or avoidance bias exhibited any clinical or demographic differences. This would be useful for identifying suitable targets for ABM. Given the growing adoption of ABM into standard alcohol treatment practise guidelines (Mann et al., 2017; Haber, 2021), addressing these knowledge gaps is warranted.

This study aimed to determine the proportion of alcohol-dependent patients with an alcohol approach bias and an alcohol avoidance bias (using the IR-AAT) in a large inpatient sample and to examine whether these two groups exhibit any demographic or clinical differences in the early stages of treatment. Additionally, we aimed to examine whether approach/avoidance bias scores on the IR-AAT were associated with any clinical indices of AUD problem severity. Specifically, we hypothesised that a greater proportion of participants would demonstrate an alcohol approach bias at baseline, and that these participants would demonstrate greater indices of alcohol use (i.e., drinking days, heavy drinking days and standard drinks in past month) and problem severity (i.e., number of previous withdrawal treatment episodes, duration of problematic alcohol use, craving, severity of dependence) compared to those with an avoidance bias. Additionally, we hypothesised that the strength of approach bias (i.e., higher scores on the IR-AAT) would be significantly correlated with indices of use and problem severity among those with a baseline approach bias.

TABLE 1 | Baseline sample characteristics and differences between participants with a baseline alcohol approach bias relative to an alcohol avoidance bias (*N* = 268).

Variable	*N*[a]	Whole sample	Alcohol approach bias (*n* = 155)	Alcohol avoidance bias (*n* = 113)	*p*[b]
Age, mean (SD)	268	43.3 (10.4)	43.0 (11.0)	43.8 (9.6)	0.534
Years of problematic alcohol use, mean (SD)[c]	267	17.0 (11.0)	16.6 (11.2)	17.7 (10.8)	0.418
Number of previous withdrawal treatment episodes, mean (SD)	268	2.5 (3.9)	2.6 (4.2)	2.4 (3.5)	0.622
Years of education, mean (SD)	266	12.5 (2.6)	12.5 (2.4)	12.5 (2.8)	0.921
Drinking days in past month, mean (SD)[d]	268	27.2 (5.1)	27.0 (5.3)	27.6 (4.6)	0.310
Heavy drinking days in past month, mean (SD)	259	26.5 (5.9	26.2 (6.3)	27.0 (5.3)	0.308
Standard drinks in past month, mean (SD)	265	575.3 (325.3)	584.4 (338.4)	562.9 (307.5)	0.595
VAS craving score, mean (SD)	268	31.4 (26.6)	31.3 (26.8)	31.7 (26.5)	0.906
ACQ total score, mean (SD)	268	3.9 (1.4)	3.8 (1.4)	4.0 (1.4)	0.209
SADQ total score, mean (SD)	242	32.3 (11.8)	32.7 (11.7)	31.7 (12.0)	0.509
Gender, male %	268	56.3	54.8	58.4	0.605
Born in Australia, %	268	83.6	85.2	81.4	0.414
Aboriginal or Torres Strait Islander, %	268	6.3	5.2	8.0	0.352
Unemployed, %	263	74.5	75.5	73.2	0.674
Unstable housing, %	267	16.5	14.2	19.6	0.236
PDOC was alcohol, %	268	95.5	97.4	92.9	0.176
Family history of SUD, %	264	56.4	56.2	56.8	0.929
Psychiatric disorder, %	268	29.9	80.6	78.8	0.704

[a] *N values are displayed due to missing data for some variables.*
[b] *Statistical difference between those with an alcohol approach bias vs. alcohol avoidance bias.*
[c] *This was calculated as the difference between participants' current age and self-reported age of onset of problematic alcohol use.*
[d] *Median score was 30, where 58.2% of participants consumed alcohol on 30 of the past 30 days.*
VAS, Visual Analogue Scale; ACQ, Alcohol Craving Questionnaire; SADQ, Severity of Alcohol Dependence Questionnaire; PDOC, Primary Drug of Concern; SUD, Substance Use Disorder.

METHOD

Participants

The full sample from the existing RCT, conducted in alcohol residential withdrawal units (Manning et al., 2021), consisted of 300 participants. However, for the purposes of these analyses, we included only the 268 participants who completed a baseline assessment of alcohol approach bias. Participants were recruited from four inpatient withdrawal treatment units in Melbourne, Australia between 2017 and 2019. Participants were required to be aged between 18 and 65; to meet Diagnostic and Statistical Manual of Mental Disorders, Fifth Edition (DSM-5) criteria for moderate or severe AUD (American Psychiatric Association, 2013); to have used alcohol at least weekly in the month prior to admission to inpatient withdrawal treatment; and to be planning to stay in treatment long enough to complete the 4-day ABM (treatment) or Sham (control) training protocol of the larger RCT (Manning et al., 2021). Patients were excluded if they had a diagnosed history of neurological illness or injury, concussion resulting in loss of consciousness longer than 30 min or any diagnosed intellectual disability, or if they were assessed by clinical staff to be too acutely unwell to provide informed consent or participate. Participants provided written informed consent. Summary statistics are provided in **Table 1**.

Measures

Clinical and Demographic Questionnaire

Prior to completing the IR-AAT, researchers administered a questionnaire at baseline assessing participants' age, gender, employment status, housing status, country of birth, Aboriginal or Torres Strait Islander status, education history, years of problematic alcohol use, number of previous withdrawal treatment episodes, primary and secondary drugs of concern, family history of substance use disorder, and psychiatric history.

Timeline Follow-Back

The timeline follow-back (TLFB; Sobell and Sobell, 1996) was used to measure alcohol consumption and captured drinking days, heavy drinking days (defined as consuming $\geq$5 standard drinks for females or $\geq$6 standard drinks for males) and total standard drinks at baseline (covering the 30 days preceding inpatient admission). The TLFB is a widely used interview method for estimating alcohol use and has been shown to concur well with other measures of alcohol use in previous research (Simons et al., 2015).

Alcohol Craving Questionnaire—Short Form—Revised

The Alcohol Craving Questionnaire—Short Form—Revised (ACQ-SF-R) was used to assess current cravings for alcohol.

The scale contains 12 items from the 47-item Alcohol Craving Questionnaire (ACQ-NOW; Singleton et al., 1995), which are strongly correlated with the ACQ-NOW and its four subscales (compulsivity, expectancy, purposefulness, and emotionality). Scores are summed to give a total score or can be summed to yield individual subscale scores, where higher scores indicate stronger cravings (Tiffany et al., 2000).

Severity of Alcohol Dependence Questionnaire

The Severity of Alcohol Dependence Questionnaire (SADQ; Stockwell et al., 1994) is a 20-item questionnaire which assesses symptoms of alcohol dependence, including physical withdrawal, affective withdrawal and drinking to relieve withdrawal symptoms. Higher scores indicate greater severity of alcohol dependence, and the SADQ has been shown to demonstrate good concurrent validity and test–retest reliability (Stockwell et al., 1983).

Irrelevant-Feature Alcohol Approach Avoidance Task

An assessment version of the IR-AAT was used to assess alcohol approach bias (Wiers et al., 2009). Using a laptop and a joystick, participants were presented with a series of images in landscape or portrait orientation (10 alcohol-related images and 10 non-alcohol-related images, each repeated 2 times, yielding a total of 40 image presentations) and instructed to push away (avoid) landscape images and pull (approach) portrait images. Each image type (alcoholic or non-alcoholic) appeared in landscape and portrait orientation 50% of the time. Pushing and pulling the joystick caused the images to decrease and increase in size, respectively. Incorrect responses were followed by a red "X," and participants were required to correct their response in order to proceed with the task.

Trials were considered valid if the initial joystick response was correct and the reaction time was between 300 and 3,000 ms. If at least 70% of the trials were valid (i.e., at least seven of the 10 trials for any picture-response category were correct), median reaction times were calculated separately for each of the four picture-response categories (alcohol-pull, alcohol-push, non-alcohol-pull, non-alcohol-push). If <70% of the trials were valid, the median for that picture-response category was considered missing. Approach bias was calculated separately for alcohol-related trials and non-alcohol-related trials by subtracting the median reaction time for pull responses from the median reaction time for push responses. The non-alcohol approach bias scores were then subtracted from the alcohol approach bias scores to provide an index of alcohol approach bias relative to non-alcohol-related images. Internal consistency of the IR-AAT was calculated using the method reported by Kersbergen et al. (2015). The internal consistency was low (Cronbach's $\alpha = 0.35$ for alcohol-related items and 0.34 for non-alcohol-related items), and the test–retest reliability (calculated only for the 117 participants in the sham training control condition who completed both baseline and post-test AAT assessments in the larger RCT) was poor ($r = 0.027$, $p = 0.774$; see Manning et al., 2021).

Procedure

Participants were screened on admission for eligibility by clinicians at the participating withdrawal treatment sites and referred to a member of the research team if interested in participating. Participants provided consent and completed baseline assessments on day 3 of their 7-day inpatient admission ($M = 7.3$ days, $SD = 2.6$). The baseline assessments included a baseline questionnaire which assessed eligibility, demographic and clinical characteristics, TLFB, SADQ, ACQ, and the AAT. The first session of ABM was typically conducted either on the same day as the baseline questionnaires or the following day. This study was approved by the St. Vincent's Hospital Melbourne Human Research Ethics Committee (HREC; reference number 030/17) and the Monash University HREC (project number 8447).

RESULTS

Of the 268 participants who completed the baseline AAT, 155 (57.8%) had an alcohol approach bias compared to 113 (42.2%) who had an avoidance bias. Overall, the sample had an approach bias for alcohol cues relative to non-alcohol cues ($M = 34.56$, 95% bootstrapped CI = 5.71–62.22; $SD = 243.77$). There were no significant differences between those with an alcohol approach bias compared to those with an avoidance bias on any of the clinical or demographic variables analysed (see **Table 1**).

Pearson's bivariate correlations were used to analyse whether alcohol approach bias was associated with indices of AUD severity among participants with an alcohol approach bias. See **Table 2** for the correlation matrix. There was a significant, moderate association between alcohol approach bias and the number of standard drinks consumed in the past 30 days ($r = 0.277$, $p = 0.001$, $n = 153$); however, associations with all other indices of AUD severity did not reach significance. Among those participants who possessed an avoidance bias at baseline ($n = 113$), the association between bias score and standard drinks was non-significant ($r = 0.033$, $p = 0.729$), and there were no associations with other indices of AUD severity that approached significance. Finally, the relationship between standard drinks and relative alcohol approach bias among those with a baseline approach bias was further confirmed in a multiple regression analysis, where standard drinks emerged as the only significant predictor of approach bias ($B = 0.182$, $t = 3.05$, $p = 0.003$), model summary: $F_{(7, 124)} = 1.99$, $p = 0.062$, $R^2 = 0.05$, $f^2 = 0.112$.

DISCUSSION

The present study sought to examine the presence of alcohol approach biases and their clinical correlates in a large sample of alcohol-dependent inpatients undergoing withdrawal treatment. In support of our first hypothesis, we found that just over half of the participants (57.8%) had an IR-AAT score indicative of an alcohol approach bias, and that the overall sample mean was also indicative of an alcohol approach bias. However, participants with an approach bias did not significantly differ on any of the demographic or clinical variables analysed from

TABLE 2 | Pearson correlation matrix for indices of alcohol consumption and problem severity among those with an approach bias (*N* = 155).

	1	2	3[a]	4[a]	5[a]	6	7	8
1. Alcohol approach bias	-							
2. Standard drinks in past month	0.277**	-						
3. Drinking days[a]	0.048	0.377**	-					
4. Heavy drinking days[a]	0.107	0.524**	0.877**	-				
5. Number of previous withdrawal treatment episodes[a]	−0.096	0.124	−0.128	−0.045	-			
6. Years of problematic alcohol use	0.073	0.137	0.010	0.065	0.129*	-		
7. SADQ total score	0.015	0.382**	0.060	0.191**	0.299**	0.080	-	
8. ACQ total score	0.120	0.117	0.023	0.024	0.089	−0.094	0.187*	-

[a]*Spearman's rho is displayed due to skewed data; SADQ, Severity of Alcohol Dependence Questionnaire; ACQ, Alcohol Craving Questionnaire.*
***Correlation is significant at the 0.01 level.*
**Correlation is significant at the 0.05 level.*

those with an avoidance bias. In contrast to our prediction for the second hypothesis, we found that the alcohol approach bias score was significantly associated with only past-month alcohol consumption and no other indices of consumption or problem severity among those with a baseline approach bias.

The finding that majority of participants and the overall sample mean demonstrated an approach bias supports a number of patient studies (Wiers et al., 2011, 2014; Eberl et al., 2013; Ernst et al., 2014) but contrasts with those reporting a mean avoidance bias (Spruyt et al., 2013) or an absence of an approach or avoidance bias altogether (Field et al., 2017). This inconsistency could be explained by the different approach bias measures used in these studies, where Field et al. (2017) and Spruyt et al. (2013) used the R-SRC rather than the IR-AAT (Spruyt et al., 2013; Field et al., 2017). As noted by Wiers et al. (2013), it is possible that the type of avoidance associations measured by the R-SRC is distinct from those captured by the IR-AAT and may be more strongly related to predicting relapse rather than those which are assessed and retrained using the IR-AAT (Wiers et al., 2013).

The absence of any group differences in demographic or clinical variables was unexpected, as was the finding that past-month alcohol consumption and no other indices of consumption or problem severity were correlated with the strength of approach bias. Whilst this supports early work by Field et al. (2008), where approach bias (on the R-SRC) was associated with weekly alcohol consumption, we failed to replicate their finding that approach bias is associated with alcohol cravings (Field et al., 2008). This was surprising, given the theoretical mechanism through which approach bias arises (i.e., through repeated associative learning experiences), whereby those who drink more frequently, in larger quantities, with longer durations of problem drinking and more extensive treatment involvement would be expected to demonstrate stronger approach bias. This may suggest that the presence of approach bias is state-dependent or varies depending on recent alcohol consumption (even among individuals with severe AUD), and future research would benefit from investigating how approach bias fluctuates over time and what factors influence this process.

These inconsistencies could also be due to the notoriously poor reliability of approach bias measures themselves (Kersbergen et al., 2015; Rinck et al., 2018), where we also found poor internal consistency and test–retest reliability in the version of the IR-AAT used in the present study. This measurement error may be at least partly due to the task comprising only half the number of trials used in standard IR-AAT tasks (in order to minimise excessive participant exposure to alcohol images during the vulnerable early withdrawal phase). The poor internal consistency could also be because individuals might only demonstrate an approach bias to alcohol cues that reflect the drinks they regularly consume. Since approach bias is posited to arise through repeated associative learning experiences, it is possible that individuals will only show an approach bias to certain alcoholic beverages that they mostly commonly consume (or at least show a much stronger bias towards these beverages). In our research, we have found that participants typically consume only a narrow range of alcoholic beverages, and given that measures of approach bias use a standard set of alcohol-related images, only a minority of the presented images may elicit responses that are indicative of an approach bias. For example, a participant who only consumes wine may demonstrate a stronger approach bias towards wine-related cues compared to other alcohol-related cues for beverages that they do not drink (e.g., beer or spirits). Indeed, studies have shown that craving and associated psychophysiological indices are stronger in response to alcohol cues that strongly resemble the most commonly consumed beverage (Staiger and White, 1991).

Researchers have begun trialling personalisation in ABM (Manning et al., 2020; Garfield et al., 2021) where participants select or rate alcohol/drug cues that best represent the substances they frequently consume, which subsequently comprise the avoidance stimuli in the training task. Personalising approach bias assessment may therefore provide a more accurate assessment of an individual's alcohol approach bias. Future research may also profit from exploring new ways of measuring approach bias (e.g., virtual reality; Eiler et al., 2019) and should explore how conventional measures of approach bias correlate with objective biomarkers of psychophysiological arousal.

Our findings should be considered in light of the study limitations. Firstly, as previously mentioned, our measure of approach bias contained only half the recommended number of trials and demonstrated poor reliability. Secondly, it is important to note that baseline approach bias was typically assessed on day 3 of inpatient withdrawal treatment when participants were receiving much higher doses of benzodiazepines than at post-training assessment, affecting reaction times, which could explain the poor test–retest reliability observed. A third limitation is that alcohol approach bias was calculated relative to non-alcohol approach bias (thereby indicating the strength of an individual's bias towards alcohol beverage cues compared to non-alcohol beverage cues). It was assumed that the non-alcohol beverage cues represent truly neutral stimuli; however, it is possible that some of the beverage types and brands may still have elicited appetitive responses, particularly when these soft drinks are commonly associated with alcoholic beverages (e.g., bottles of Coke for a whiskey and Coke drinker).

Despite these limitations, this paper addresses important questions regarding the measurement of approach bias in alcohol-dependent patients and identifies key considerations for future research. Whilst it may be desirable to administer ABM only to those patients with an alcohol approach bias, the absence of mediation effects in most clinical trials to date (Wiers et al., 2011; Rinck et al., 2018; Manning et al., 2021), coupled with its clear effects on relapse reduction (Wiers et al., 2011; Eberl et al., 2013; Rinck et al., 2018; Manning et al., 2021), suggests that ABM should be routinely offered to all alcohol-dependent patients—particularly given its low-cost and short administration time. Most importantly, this paper raises some important questions regarding the measurement and clinical relevance of approach bias and highlights the need to develop more accurate measures (e.g., through personally tailored measures or virtual reality paradigms).

DATA AVAILABILITY STATEMENT

The datasets presented in this article are not readily available because longer term outcomes of the broader trial have yet to be completed, hence the full dataset is currently unavailable. Requests to access the datasets should be directed to Victoria.manning@monash.edu.

ETHICS STATEMENT

The studies involving human participants were reviewed and approved by the St. Vincent's Hospital Melbourne Human Research Ethics Committee (reference number 030/17) and the Monash University HREC (project number 8447). The patients/participants provided their written informed consent to participate in this study.

AUTHOR CONTRIBUTIONS

HP wrote the first draft with VM and conducted the analyses. PS provided input and overall guidance regarding the structure of the paper. All authors conceptualised the current paper.

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Food Palatability Directs Our Eyes Across Contexts

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It is often believed that attentional bias (AB) for food is a stable trait of certain groups, like restrained eaters. However, empirical evidence from this domain is inconsistent. High-calorie foods are double-faceted, as they are both a source of reward and of weight/health concern. Their meaning might depend on the food-related context (i.e., focus on health or on enjoyment), which in turn could affect AB for food. This study primed 85 females with hedonic, healthy, and neutral contexts successively and examined whether food-related context affected AB for food and if effects were moderated by dietary restraint. Both the mean tendencies of AB for food and variability of AB for food were assessed in a food dot-probe task with a recording of both reaction times and eye movements. Contrary to our hypotheses, AB for food was not significantly affected by either context or the interaction between context and dietary restraint. Instead, liking of the presented food stimuli was related to longer initial fixations and longer dwell time on the food stimuli. In addition, in line with prior research, body mass index (BMI) was correlated with variability of AB for food instead of mean AB for food. In conclusion, this study did not find any support that AB for food is dependent on food-related context, but interestingly, reaction time-based variability of AB for food seems to relate to BMI, and eye movement-based mean AB seems to relate to appetitive motivation.

Keywords: attentional bias, dynamics, context, priming, restrained eating

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INTRODUCTION

In general, people are naturally attracted by high-calorie foods (e.g., McSorley et al., 2017). In the Western food-rich environment, the abundant presence of high-calorie palatable foods represents salient cues that can induce food craving (Hill and Peters, 1998), subsequent food intake, and ultimately weight gain (Boswell and Kober, 2016). However, in daily life, there are also moments that weight control thoughts or weight/health-related cues can lead to food avoidance. High-calorie foods are frequently craved but are also often a source of worry and weight concern. This is also referred to as the double-facetted nature of food (Roefs et al., 2018), in other words, a conflict between food enjoyment and weight concern. The current study investigates if inducing a hedonic vs. a health context affects attentional bias (AB) for food and if this effect is moderated by dietary restraint.

AB refers to an enhanced attention to salient or relevant stimuli (Drobes et al., 2019). In previous studies, AB for food was mostly considered as a trait-like characteristic of both restrained eaters (REs) and people with overweight/obese. This popular belief is based on the incentive salience model: a reward stimulus, like palatable food, can lead to a craving for it, which reflects on a biased attention to the rewarding food and such attention-grabbing should be especially true for REs and people with overweight/obese (e.g., Berridge, 2009; Nijs and Franken, 2012). However, the empirical evidence for food-related AB in REs and people with overweight/obesity is inconsistent, which has been repeatedly revealed by reviews and meta-analyses (e.g., Doolan et al., 2015; Roefs et al., 2015; Werthmann et al., 2015; Hagan et al., 2020; Hardman et al., 2020). Briefly, some studies indeed found that people high in body mass index (BMI) or dietary restraint biased their attention more to food stimuli than the control group (e.g., Meule et al., 2012; Kemps et al., 2014; Hume et al., 2015), while other studies found that people high in BMI or dietary restraint showed equal (e.g., Werthmann et al., 2013; Doolan et al., 2014; Hodge et al., 2020) or even less (e.g., Nathan et al., 2012; Fang et al., 2019) attention to food cues than their counterparts. Interestingly, some studies even revealed an attention approach-avoidance pattern in people high in BMI and dietary restraint (e.g., Hollitt et al., 2010; Werthmann et al., 2011). This inconsistency might partly be explained by the fact that diverse measurements were used to capture individuals' AB for food (e.g., different paradigms and materials used to assess AB for food). It could also be due to some studies being underpowered, which possibly could have led to biased results. However, we believe that AB for food is not a stable trait-like feature in people with overweight/obesity and REs.

Apart from the inconsistent empirical evidence, linking enhanced AB for food to people high in BMI or dietary restraint also conveys two incorrect beliefs: (1) food is only a source of reward; (2) attention is stable over time, either towards or away from food. As we stated before, palatable foods can be a source of both reward and weight gain/health concern (Roefs et al., 2018). Relevant for the second incorrect belief is that attention is also a reflection of the current motivation (top-down; Connor et al., 2004). Therefore, theoretically, whether an individual focuses on the enjoyment facet or the health facet of food could direct attention towards or away from food, which might have contributed to the inconsistent results across studies. Previous studies indeed provided evidence that it is possible to influence individuals' AB for high-calorie foods and subsequent food intake by manipulating the context. For example, in an online supermarket eye-tracking study, priming health/diet cues (recipe banners containing health and dieting words were presented during food choice) increased low-calorie food choices, decreased high-calorie food choices, and increased total dwell times on low-calorie products (van der Laan et al., 2017). Papies and Hamstra (2010) found that priming dieting cues (a weekly recipe that was "good for a slim figure" and low in calories was attached to the door of the butcher's store) decreased food consumption in REs, but not in unrestrained in unrestrained eaters (uREs). In line with this, a dieting context (participants were asked to choose a 1-day menu from a healthy menu card to their best friend who wants to lose weight) decreased mean AB scores for high-calorie food only in participants with higher dietary restraint scores (Werthmann et al., 2016). All in all, evidence supports the idea that individuals' AB for food is not stable and is influenced by food-related contexts, especially in high REs.

In the tasks used to measure AB, an aggregated mean AB score has been widely adopted to characterize AB for food, which reflects an overall, stable tendency of AB during the task. However, this mean AB score does not do justice to the potentially dynamic nature of AB for food, as attention to food might alternate between approach and avoidance, even within one study within one participant. Interestingly, another method of computing AB was introduced, trial-level-bias scores (TL-BS; Zvielli et al., 2015), which specifically acknowledges that AB might not be a fixed characteristic but may instead fluctuate over the course of an experiment. This method focuses on the degree of fluctuation in AB and how this fluctuation is related to certain traits. TL-BS has shown added value in several domains. That is, studies revealed that TL-BS variability, one of the parameters to measure the variability of AB for critical stimuli based on the TL-BS, could better predict BMI, depression, and spider phobia than corresponding mean AB scores (Zvielli et al., 2015, 2016; Liu et al., 2019a,b). Specifically, for AB for food, it has been shown that people with a higher BMI do not have significantly more AB for food than healthy-weight people but are characterized by more TL-BS variability for food (Liu et al., 2019a,b).

Altogether, both empirical and theoretical evidences support that individuals' AB for food is not a stable trait but fluctuates over time. Attention fluctuations might reflect the momentary inner conflict between food enjoyment and weight/health concern when confronted with palatable food, which possibly can be shaped by manipulating food-related contexts. Moreover, individuals with higher dietary restraint are more sensitive to both food-related reward and punishment (Ahern et al., 2010), and as we mentioned before, it was found that the influence of context on food intake and AB for food only happened in participants high in dietary restraint (Papies and Hamstra, 2010; Werthmann et al., 2016). Therefore, it might be that high REs who frequently experience conflicts between food enjoyment and weight concern (Stroebe et al., 2013) have more fluctuations in AB for food when in a "neutral state," reflecting what might typically occur in daily life, fluctuating between craving and weight concern when confronted with palatable food. These fluctuations might reduce when the context strongly emphasizes either enjoyment or health. Specifically, people may have more and more consistent AB towards high-calorie food when in a hedonic context and more and more consistent AB away from high-calorie food when in a health context. Relevant to mention here is that the mindset or context was not manipulated in previous studies that used TL-BS AB scores. It would therefore be valuable to see how the food-related contexts influence the variability of AB for food.

The current study examined the effect of context priming on both the average AB for food and the variability of AB for food. Context (hedonic, health, and neutral) was manipulated in a within-subject design. It was hypothesized that (1) in the hedonic context as compared with the health context, participants would show a larger mean AB towards food and have longer first fixations and dwell time on food stimuli, whereas the neutral context was expected to fall in between the hedonic and health contexts; (2) compared with the neutral context, participants would show fewer reaction time (RT)-based and eye movement (EM)-based fluctuations on food in both the hedonic and health contexts; (3) contexts would affect both the average and the fluctuations of AB for food more for participants scoring higher on dietary restraint.

MATERIALS AND METHODS

The current study was approved by the Ethical Committee of the Faculty of Psychology and Neuroscience, Maastricht University, and all participants signed the informed consent, in which they were informed about the procedure, storage of data, and their right to withdraw from participation without any consequence. This study was pre-registered at AsPredicted.[1]

Participants

A total of 91 female participants were recruited *via* posters on the university campus (Maastricht University, the Netherlands) or on the Facebook, or *via* an online recruitment system. To disguise the true purpose of this study and avoid influencing participants' behavior, it was informed on the poster that this study aims to investigate "attention patterns for different objects." People who signed up for the study received a screening questionnaire with questions on sex, age, and vision. Females with a normal or corrected-to-normal vision and above 16 years old were invited to participate. Because food pictures depicting meat were included in the study, after the participant finished the experiment, she was asked whether she adheres to a vegetarian or vegan diet. Six participants were excluded from analyses because of either a vegetarian or vegan diet. So 85 participants were included in the analyses. Participants' characteristics can be found in **Table 1**. The sample size is adequate to detect a medium effect size for the main aim of the study, which is to test the effect of priming condition (within-subject) on all measures of AB scores; thus, repeated-measures MANOVAs were conducted. When using repeated-measures MANOVAs (number of groups = 1, number of measurements = 3, medium effect size f = 0.25, alpha = 0.05, power = 0.95), the estimated required sample size was 45.

[1]https://aspredicted.org/96j53.pdf

TABLE 1 | Participant characteristics.

Variables	*M*	*SD*	*Range*
RS score	13.80	5.20	1.00–28.00
Age	21.48	2.96	17.00–31.00
BMI	22.02	3.01	16.81–31.25
Hunger	33.40	23.74	0.00–80.00
Food liking	57.12	19.91	6.50–92.50

RS, restraint scale (Herman and Polivy, 1980); BMI, body mass index; hunger, hunger level of participants (0–100 VAS); food liking, average food liking score for the high-calorie food stimuli presented in the food-dot probe task (0–100 VAS).

Measurements

Dietary Restraint

The Restraint Scale (RS, Herman and Polivy, 1980) was used to measure restrained eating. The RS is an 11-item self-report scale that is used to assess chronic dieting. In this study, Cronbach's alpha was 0.76.

Hunger Level

Participants' hunger level was measured on an online 100-mm visual analog scale (VAS) by asking "How hungry are you right now?," with 0 indicating absolute lack of hunger and 100 indicating extreme hunger.

Liking of the Food Images

Liking of the high-calorie food images that were presented in the food-dot probe task (20 images) was measured on an online 100-ms VAS, with 0 indicating a lack of liking and 100 indicating extreme liking.

Food Dot-Probe Task

AB for high-calorie foods was measured using the food dot-probe task with a recording of both RTs and EMs. The task was presented using Experiment Builder (SR Research, Ontario, Canada).

Trial Procedure

The priming context was manipulated in a blocked fashion, and in each block, one type of priming picture (hedonic or health or neutral) was presented in each trial. Each trial began with a central fixation dot, which disappeared directly after participants fixated on it. Subsequently, a prime image was presented for 1,000 ms. After presentation of another central fixation dot (500 ms), two images were simultaneously presented side by side for 2,000 ms. Next, one of the images was replaced by the probe (*), which randomly and equally often appeared on the left or right side of the screen. The probe was presented until the participant's response or for a maximum of 2,000 ms. Participants were instructed to focus on the central fixation and to respond to the probe as quickly as possible.

Trial Types

Three different types of trials were included: food-incongruent trials (ITs), food-congruent trials (CTs), and neutral-neutral trials (filler trials). On the ITs and CTs, a high-calorie food and a musical instrument picture were presented, whereas on the filler trials, two neutral nonfood pictures (e.g., office supplies), were presented. During ITs, the probe appeared in the location of the musical instrument picture, whereas during CTs, the probe appeared in the location of the food picture. In the filler trials, the probe appeared randomly and equally often on the left and right sides.

Block Types

The task included one practice block, two buffer blocks, and three priming blocks. The three different trial types were evenly and randomly distributed across the priming blocks (40 CTs, 40 ITs, and 40 filler trials for each priming block). Buffer blocks (40 filler trials for each buffer block) only included filler trials and served to neutralize the participants' mindset. Each priming block included 20 different prime pictures, 10 different food-musical instrument pairs, and 10 different neutral-neutral pairs. Each buffer block included five different neutral prime pictures and 10 different neutral-neutral pairs. The buffer blocks were presented between two priming blocks to reduce the interference from the prime pictures of the previous block. The order of the priming blocks was counterbalanced across participants.

Stimuli in the Food Dot-Probe Task

Most of the palatable food stimuli and musical instrument stimuli in the current study were from Werthmann et al. (2011). All image pairs were subjectively matched as closely as possible with regard to the shape, color, brightness, and size of the depicted object. Each picture appeared equally often on the left and right sides of the screen. All the food stimuli were rated by participants in the final online questionnaire according to how much they like the food.

Prime Pictures

Ninety candidate prime pictures, with 30 depicting the eating enjoyment context (e.g., wedding), 30 depicting a health-related context (e.g., sports), and 30 depicting a food-unrelated context (e.g., street view), were rated on how much food indulgence and how many healthy food choices they elicited on a 100-mm VAS by asking "How much would you like to indulge in tasty food after viewing the above picture?" and "How inclined are you to choose healthy food after viewing the above picture?," with 0 reflecting "not at all" and 100 reflecting "very much." Ratings were provided by 37 women (M_{age} = 21, SD_{age} = 3.20; no participants of the current study). The 20 eating enjoyment-related pictures with the highest ratings on food indulgence were selected as the hedonic primes; the 20 health-related pictures with the highest ratings on healthy food choices were selected as the health primes; the 20 food-unrelated pictures, with lower ratings on both food indulgence [hedonic primes vs. neutral primes: $t(19) = 41.26$, $p < 0.001$] and healthy food choices [healthy primes vs. neutral primes: $t(19) = 21.49$, $p < 0.001$], were selected as the neutral primes. The prime pictures in the buffer blocks were landscape pictures. For the average priming picture rating scores per priming condition, see **Table 2**.

Eye-Movement Measurements

Participant's EM data were collected *via* a desktop mounted EyeLink 1000 system. All stimuli were presented on a 24-inch computer screen at a viewing distance of about 57 cm. With the use of DataViewer software (SR Research, Canada), saccades and fixations were extracted. The display screen was divided into three interest areas: the middle section (represented the location of the fixation cross) and the left and right sections (represented the location of the stimuli). The width of the middle-interest area was decided by a given visual angle: 2 horizontal degrees (Amir et al., 2016). Fixations located in the middle-interest area and fixation durations below 60 ms were discarded (Amir et al., 2016).

Manipulation Check

The manipulation check was conducted at the end of the experiment, to test whether the three types of prime pictures influenced participants' desire to indulge in palatable foods differently. All prime pictures used in the task were successively and randomly displayed on the screen in three blocks with the same category of prime pictures in one block. Each picture was presented for 1,000 ms (same display time as in the food dot-probe task). After viewing one block, participants were asked to indicate how much they would like to indulge in tasty food right now on a 100-mm VAS. The three blocks were presented in random order.

Procedure

All participants were tested individually in a laboratory at Maastricht University after 10 a.m. The order of the priming context was counterbalanced across participants. First, the participant's hunger level was measured on a 100-mm VAS together with several filler questions[2] *via* a Qualtrics online survey. Then after a 9-point calibration (calibration for proper gaze recording by the system) with subsequent validation (validation for gaze position accuracy achieved by the current calibration) procedure, the food dot-probe task was administered. Next, the manipulation check, the RS, liking of the food pictures, and self-reported height and weight were assessed in another Qualtrics online survey. Finally, we explained the real purpose of the current study to the participant, and the participant was thanked for her participation and received a small reward (either 1.5-h course credits or 10 euros in gift voucher). The whole procedure lasted about 1.5 h.

Data Reduction and Analysis

All AB scores were computed separately for each priming condition for each participant.

Reaction Time-Based Attentional Bias for Food

Mean AB for food was calculated by subtracting the mean RTs of CTs from the mean RTs of ITs. So a positive value reflects an AB towards food, and a negative value an AB away from food. To obtain sequential TL-BSs, first, each CT was paired with an IT that was as close as possible in time and no further than five trials away. Next, the CT in each pair was subtracted from the IT in that pair. TL-BS variability for

[2]Filler questions: How many glasses of alcohol did you drink yesterday? How many cups of coffee did you drink yesterday? How many cigarettes did you smoke yesterday? How many cups of water did you drink yesterday?

TABLE 2 | Priming pictures rating scores per priming condition.

	Hedonic primes		Health primes		Neutral primes	
Variables	***M***	***SD***	***M***	***SD***	***M***	***SD***
Indulgence	64.24	2.32	36.05	2.60	37.59	2.17
Healthy food choices	39.58	6.15	60.33	3.21	36.57	4.11

food was computed using the sum of absolute distances between sequential TL-BSs divided by the total number of TL-BSs (Zvielli et al., 2015).

Eye Movement-Based Attentional Bias for Food

Based on EM data, two average AB for food scores were computed: mean initial fixation duration bias on food and mean dwell time bias on food. The initial fixation duration represents the duration of the first fixation that remains on one of the picture stimuli. Initial fixation durations were firstly averaged across trials per participant, separately for the two categories (food/nonfood). Then, per participant, the mean initial fixation duration bias on food was calculated by subtracting the averaged initial fixation duration in the interest area containing a musical instrument stimulus from the averaged initial fixation duration in the interest area containing a food stimulus.

The dwell time is the total time that a gaze remained at each stimulus during the 2,000-ms presentation time. Mean dwell time bias was calculated by subtracting the mean dwell time in the interest area with a musical instrument stimulus from the mean dwell time in the interest area with a food stimulus. So positive values reflect more attention for food than for musical instruments. The EM-based dynamic changes of AB for food were operationalized as the standard deviation (*SD*) of the initial fixation duration bias on food, the *SD* of dwell time bias on food, and the number of switches between the food and nonfood stimuli within each trial.

Data Reduction

Firstly, buffer blocks were excluded from analyses, and then trials were excluded from analyses if they contained error responses, were faster than 200 ms, and slower than 2,000 ms, and after that if they deviated more or less than 3 *SD*s from each participant's mean RT (Werthmann et al., 2011; 2.20% of the RT data were excluded). In addition, it was checked if participants moved their eyes on a sufficient proportion (50%) of trials (Bradley et al., 2000), and this led to no participant exclusion. Then both RT-based and EM-based AB scores per priming context were calculated separately, after that under each priming context, it was checked whether any AB score deviated more or less than three *SD*s from the respective mean of the whole sample (outliers). Finally, 29 outlier AB scores (the percentage of these outliers was 1.62%) were replaced by the respective nearest score of the whole sample (Wilcox, 2001).[3]

[3]The analyses with unchanged values or exclusion of these values showed the same pattern of results and led to the same conclusions.

Analysis Plan

First, the manipulation check regarding a priming context was conducted by comparing the scores on the manipulation question between a hedonic/health context and a neutral context in a repeated-measures ANOVA. Then, two repeated-measures MANOVAs were conducted to test the effects of different priming contexts and the interaction between priming contexts and dietary restraint (mean-centered) on both AB fluctuation scores (TL-BS variability, *SD* of initial fixation duration bias, *SD* of dwell time bias, and number of switches between the food and nonfood stimuli) and mean AB scores (mean AB scores, mean initial fixation duration bias, and mean dwell time bias). According to previous studies (e.g., Tapper et al., 2010; Hardman et al., 2020), BMI, hunger level, and food liking scores might influence AB scores. Last, correlations between AB fluctuation scores/ mean AB scores and these variables were conducted as exploratory analyses. Because of the large number of statistical tests, alphas were adjusted using the Bonferroni method.

RESULTS

Manipulation Check

Data were analyzed in a repeated-measures ANOVA with priming condition as the factor and scores on the manipulation check question "how much would you like to indulge in tasty food after viewing the above video" as the dependent variable. Scores on the manipulation check question differed significantly between priming conditions, $F(2, 168) = 83.50$, $p < 0.001$. *Post hoc* tests using the Bonferroni correction (adjusted alpha = 0.025) revealed that ratings were higher for the hedonic condition, $M = 63.88$, $SD = 2.80$, than for both the health, $M = 36.44$, $SD = 2.71$, and the neutral condition, $M = 35.04$, $SD = 2.84$; $ps < 0.001$. The health condition did not differ significantly from the neutral condition, $p = 1.00$. So the hedonic context was successfully induced, but the health context was not.

Average Performance on Tasks

Participants' average performance on the task is displayed in **Table 3**.

Effects of Priming Condition and Restrained Eating on Attentional Bias for Food

Mean Attentional Bias for Food

The results from MANOVA showed that there was no statistically significant main effect of priming contexts, $F(6, 78) = 0.81$,

TABLE 3 | Measures of AB per priming condition.

Variables	Hedonic priming		Healthy priming		Neutral priming	
	M	*SD*	*M*	*SD*	*M*	*SD*
TL-BS variability	132.15	49.65	128.77	46.88	136.74	55.97
Mean AB	9.64	25.75	7.83	24.54	8.83	25.65
Mean initial fixation duration bias	39.08	73.51	35.98	67.11	46.39	73.26
Mean dwell time bias	48.80	295.95	15.56	242.53	50.89	211.06
SD of initial fixation duration bias	359.02	152.65	375.95	173.08	376.39	186.10
SD of dwell time bias	894.42	345.52	911.58	373.72	894.07	363.02
Number of switches	1.84	0.54	1.80	0.66	1.83	0.59

TL-BS, trial-level bias scores; AB, attentional bias; SD, standard deviation.

TABLE 4 | Correlations between BMI, hunger level, food liking, and mean attentional bias indexes.

N = 85	1	2	3	4	5	6
1. Mean AB	1	0.35*	0.50*	0.001	−0.03	0.21
2. Mean initial fixation duration bias		1	0.85*	−0.0003	0.25	0.35*
3. Mean dwell time bias			1	0.01	0.19	0.46*
4. BMI				1	−0.08	−0.23
5. Hunger					1	0.29*
6. Food liking						1

*AB, attentional bias; BMI, body mass index; hunger, hunger level of participants (0–100 VAS); food liking, average food liking score for the high-calorie food stimuli presented in the food-dot probe task (0–100 VAS). *p < 0.0167.*

p = 0.57, Wilks' Λ = 0.94, partial η^2 = 0.06, on the combined dependent variables, mean AB scores. The interaction effect between priming contexts and dietary restraint on the combined dependent variables was also not statistically significant, $F(6, 78) = 1.87$, $p = 0.10$, Wilks' Λ = 0.87, partial η^2 = 0.13.

Fluctuations in Attentional Bias for Food

There was no statistically significant difference between the priming contexts on the combined dependent variables, AB fluctuation scores, $F(8, 76) = 1.20$, $p = 0.31$, Wilks' Λ = 0.89, partial η^2 = 0.11. The interaction effect between priming contexts and dietary restraint on the combined dependent variables was also not statistically significant, $F(8, 76) = 0.73$, $p = 0.66$, Wilks' Λ = 0.93, partial η^2 = 0.07.

Correlational Analyses

Mean Attentional Bias for Food

The results of the correlations (adjusted alpha = 0.0167) showed that food liking was positively related to mean initial fixation duration bias and mean dwell time bias. Apart from that, BMI, hunger level, and food liking were not related to any other mean measures of AB scores; see **Table 4**.

Fluctuations in Attentional Bias for Food

As for the AB fluctuation scores, only BMI was significantly related to the TL-BS variability: participants with a higher BMI showed more variability in AB for food (adjusted alpha = 0.0125). For more information, see **Table 5**.

DISCUSSION

This study examined whether hedonic and health priming conditions influence AB for food and how this effect would be moderated by dietary restraint. Both RT-based and EM-based mean tendencies and variability of AB for high-calorie foods were measured. Unexpectedly, it was found that both the priming condition and the interaction between the priming condition and dietary restraint did not significantly affect AB for food. However, BMI was positively related to TL-BS variability, which is in line with previous studies (Liu et al., 2019a,b). In addition, food liking was positively related to initial fixation duration bias and dwell time bias. So participants who reported higher liking of the presented food stimuli looked at the food stimuli longer.

The results of the manipulation check demonstrated that after experiencing hedonic priming pictures, participants reported that they wanted to indulge in high-calorie foods more as compared with both the neutral and health priming contexts, which means the hedonic priming context indeed induced a hedonic goal. However, the priming manipulation was not entirely successful, as there was no significant difference between the health and neutral priming contexts. Thus, it can be concluded that even though the hedonic priming was successful, it did not translate to effects on AB for food. It contradicts a previous study (Werthmann et al., 2016), which found that a health mindset as compared with a palatability mindset decreased RT-based AB for high-calorie foods in participants with higher dietary restraint. However, it should be noted that in Werthmann et al. (2016), the two mindset conditions (health vs. hedonic) only differed significantly on the rated importance of health, but not on the rated importance of palatability, and the design did not include a neutral mindset condition. So less AB for high-calorie foods in a health mindset than in a palatability mindset should likely be attributed to an increased importance of health in the health mindset condition. In addition, to the best of our knowledge, a previous study only found a significant influence of context on AB for food primed participants with a dieting-related context instead of a hedonic context (van der Laan et al., 2017). So maybe health/weight concerns more easily reduce AB for high-calorie foods than that a hedonic focus increases AB for food. Therefore, the unsuccessful manipulation in the health condition in our study might explain why our results are not in line with previous studies. To induce weight concern, more salient cues should be included in future studies, like the scales and weight/dieting-related information instead of the exercise and healthy food pictures used in the current study. It also suggests that mild health cues in real life might not be enough to influence attention, especially when hedonic cues are presented at the same time. Moreover, the current study used a within-subjects design, whereas previous studies (e.g., Werthmann et al., 2016)

TABLE 5 | Correlations between BMI, hunger level, food liking, and attentional bias fluctuation indexes.

***N* = 85**	**1**	**2**	**3**	**4**	**5**	**6**	**7**
1. TL-BS variability	1	−0.12	−0.12	0.02	0.27*	−0.0002	0.14
2. *SD* of initial fixation duration bias		1	0.66*	−0.57*	0.001	0.14	0.13
3. *SD* of dwell time bias			1	−0.65*	−0.06	0.05	0.10
4. Number of switches				1	0.07	−0.11	−0.17
5. BMI					1	−0.08	−0.23
6. Hunger						1	0.29*
7. Food liking							1

*TL-BS, trial-level bias scores; SD, standard deviation; BMI, body mass index; hunger, hunger level of participants (0–100 VAS); food liking, average food liking score for the high-calorie food stimuli presented in the food-dot probe task (0–100 VAS). *p < 0.0125.*

used a between-subjects design. So it might be that the buffer blocks were not sufficient to avoid spillover between the different priming conditions. The current study only investigated the effect of context priming on AB for palatable foods. It would be interesting to investigate how different contexts affect AB for healthy food stimuli in future studies.

Dietary restraint and BMI are frequently believed to be related to AB for high-calorie palatable food, and this has been tested in many studies (e.g., Castellanos et al., 2009; Meule et al., 2012). However, the empirical evidence does not support this claim (e.g., Doolan et al., 2015; Roefs et al., 2015; Werthmann et al., 2015; Hagan et al., 2020; Hardman et al., 2020). The current study also did not find any significant relationship between dietary restraint or BMI and mean AB for food, not even in the hedonic priming condition. So the results from the current study question the notion that AB for high-calorie palatable food is a trait-like feature of people with a high BMI or scoring high on dietary restraint.

Interestingly, the current study did show a positive association between BMI and TL-BS variability for food, which is in accordance with previous studies (Liu et al., 2019a,b), which included a range of participants (normal-weight females and children, overweight females, and overweight/obese children). Notably, in one of these studies (Liu et al., 2019b), it was found that this relation between BMI and TL-BS variability does not hold if attention control is high and another study directly found that the relationship between anxiety and the TL-BS variability for anxiety-related stimuli was significantly mediated by attention control (Clarke et al., 2020). Therefore, the variability of AB for food might reflect weaker executive control. In detail, weaker executive control might make people less likely to have a consistent, prolonged attention to food stimuli during the task, therefore causing more variability in AB for food.

The current study also found positive associations between participants' food liking and EM-based mean AB for food (both mean initial fixation duration bias and mean dwell time bias), which is consistent with previous studies (e.g., Kemps and Tiggemann, 2009). The incentive salience model (Berridge, 2009) proposed that AB for food reflects appetitive motivation, and except for the influence of momentary motivation on AB for food, the relatively stable trait, food liking, should be also closely related to AB for food. However, the average food liking score was relatively low in the current study, although the food stimuli used in the current study included only widely liked food items, like chocolate and crisps. Tailoring food stimuli to the individual participant should be considered in future studies.

Except for the within-subject design and relatively low food liking score, there are several other limitations of the current study that need to be mentioned here. First, the participants included in this study were fairly homogeneous. They are all females, and most of them were students. Future studies are needed to confirm the effect of context priming on AB for food in diverse populations. Second, because BMI is not our main focus, we used self-reported BMI instead of measuring BMI, which could have caused inaccuracy and therefore have affected the results regarding BMI.

CONCLUSION

The present study assessed the effect of context on AB for food. The situational context was manipulated by priming participants with hedonic, healthy, and neutral pictures. Contrary to our hypothesis, the current results did not provide evidence on the influence of a hedonic or health-related context on individuals' AB for food. A positive association between BMI and TL-BS variability for food was found, which replicates results from previous studies (Liu et al., 2019a,b). Food liking was positively related to AB for food, which is in line with the idea that AB for food reflects appetitive motivation (Field et al., 2016). Finally, adding to the inconsistency in the field (e.g., Doolan et al., 2015; Roefs et al., 2015; Werthmann et al., 2015; Hagan et al., 2020; Hardman et al., 2020), this study failed to show any significant relationship between dietary restraint or BMI and average AB for food.

DATA AVAILABILITY STATEMENT

The original contributions presented in the study are included in the article/**Supplementary Material**, further inquiries can be directed to the corresponding author.

ETHICS STATEMENT

The studies involving human participants were reviewed and approved by the Ethical Committee of the Faculty of Psychology and Neuroscience, Maastricht University. Written informed consent from the participants' legal guardian/next of kin was not required to participate in this study in accordance with the national legislation and the institutional requirements.

AUTHOR CONTRIBUTIONS

YL, AR, and CN designed the study. YL collected and analyzed the data. YL wrote the manuscript. AR and CN gave feedback on the manuscript. All authors contributed to the article and approved the submitted version.

SUPPLEMENTARY MATERIAL

The Supplementary Material for this article can be found online at: https://www.frontiersin.org/articles/10.3389/fpsyg.2021.664893/full#supplementary-material

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Measuring Food-Related Attentional Bias

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Objective: Food-related attentional bias has been defined as the tendency to give preferential attention to food-related stimuli. Attentional bias is of interest as studies have found that increased attentional bias is associated with obesity; others, however, have not. A possible reason for mixed results may be that there is no agreed upon measure of attentional bias: studies differ in both measurement and scoring of attentional bias. Additionally, little is known about the stability of attentional bias over time. The present study aims to compare attentional bias measures generated from commonly used attentional bias tasks and scoring protocols, and to test re-test reliability.

Methods: As part of a larger study, 69 participants (67% female) completed two food-related visual probe tasks at baseline: lexical (words as stimuli), and pictorial (pictures as stimuli). Reaction time bias scores (attentional bias scores) for each task were calculated in three different ways: by subtracting the reaction times for the trials where probes replaced (1) neutral stimuli from the trials where the probes replaced all food stimuli, (2) neutral stimuli from the trials where probes replaced high caloric food stimuli, and (3) neutral stimuli from low caloric food stimuli. This resulted in three separate attentional bias scores for each task. These reaction time results were then correlated. The pictorial visual probe task was administered a second time 14-days later to assess test-retest reliability.

Results: Regardless of the scoring use, lexical attentional bias scores were minimal, suggesting minimal attentional bias. Pictorial task attentional bias scores were larger, suggesting greater attentional bias. The correlation between the various scores was relatively small (r = 0.13–0.20). Similarly, test-retest reliability for the pictorial task was poor regardless of how the test was scored (r = 0.20–0.41).

Conclusion: These results suggest that at least some of the variation in findings across attentional bias studies could be due to differences in the way that attentional bias is measured. Future research may benefit from either combining eye-tracking measurements in addition to reaction times.

Keywords: attentional bias, visual probe, lexical, reaction time, reliability, pictorial

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INTRODUCTION

Attentional bias is the tendency to give preferential attention to stimuli which are personally, motivationally and emotionally relevant (Kuckertz and Amir, 2015). It is of interest to researchers because both theoretical predictions and observational data link attentional bias to important behavioral outcomes, such as food choice/intake (for example, see; Nijs et al., 2010b; van Ens et al., 2019). Attentional bias is commonly assessed using the visual probe task, which was originally developed to study anxiety (MacLeod et al., 1986) and then later used in the addiction field (Mogg et al., 2003). However, this task has increasingly been used to study the link between attentional bias toward food cues and outcomes such as eating patterns and obesity (Field et al., 2016).

Although attentional bias for food has been linked to obesity (Castellanos et al., 2009; Calitri et al., 2010; Nijs et al., 2010b; Kakoschke et al., 2014), numerous studies have reported no such association (Loeber et al., 2012; Garcia-Garcia et al., 2013; Werthmann et al., 2014). It has been suggested that the conflicting findings may be due—at least in part—to differences in the way that attentional bias is assessed from study to study (Nijs and Franken, 2012). For example, various tasks have been used to investigate food-related attentional bias. Tasks such as the Dot Probe (Kemps et al., 2014; van Ens et al., 2019), the Stroop (Nijs et al., 2010a; Phelan et al., 2011), the Flicker (Favieri et al., 2020) and the Go/no-go (Love et al., 2020) share the same underlying goal of assessing food-related attentional bias, but differ in how this is operationalized. Furthermore, variability exists even within-task. For example, the commonly used visual probe paradigm is sometimes conducted with word pairs (Calitri et al., 2010; Kemps et al., 2014) but other times picture pairs are used (Werthmann et al., 2014; Meule and Platte, 2016). Currently, there is no evidence as to whether bias scores obtained from the lexical and pictorial tasks are comparable on an individual level. Adding further variability to the testing procedure, how individual attentional bias tasks are scored also varies across studies. With the visual probe task, a reaction time bias score is generally expressed as a difference score by subtracting the reaction times for the trials where probes replaced neutral stimuli from the trials where the probes replaced target (e.g., food-related) stimuli (van Ens et al., 2019). As such, positive scores indicate a bias toward food related stimuli. However, some studies report bias scores as the difference between; (i) neutral and food stimuli (Ruddock et al., 2018; Fang et al., 2019; Liu et al., 2019; Mas et al., 2019), (ii) neutral and high caloric food stimuli (Favieri et al., 2020; Love et al., 2020), (iii) neutral and low caloric food stimuli (Favieri et al., 2020; Love et al., 2020), or the difference between (iv) low caloric food and high caloric food stimuli (Meule and Platte, 2016; Zhang et al., 2018). It is currently unclear whether these different scoring approaches produce comparable bias scores; as such, we do not know whether these differences in scoring add to the variability in the attentional bias literature. Regardless of the scoring procedure or task used, the results remain varied; findings are inconsistent irrespective of whether studies implement the same procedure or not.

While various tasks have been used to assess attentional bias, the visual probe task is considered "gold standard" by some researchers, remaining the most extensively used in attentional bias research (Kappenman et al., 2014). However, the task's test-retest reliability has been called into question: studies of attentional bias in the addiction and anxiety fields, for example, have reported poor test-retest reliability (Schmukle, 2005; Christiansen et al., 2015; MacLeod et al., 2019). When examining the test-retest of the visual probe for food-related attentional bias, van Ens et al. (2019) reported acceptable test-retest reliability for all reaction time indices. However, van Ens et al. (2019) used a longer stimulus presentation time than is commonly used in attentional bias studies (stimulus presentation time of 3,000 ms vs. the more standard 500 ms). Little is known about the test-retest reliability of the food-related visual probe tasks when using the presentation time of 500 ms. As such, while theorists generally assume that attentional bias is a more or less static variable, with little variation from moment to moment (for example, incentive sensitization theory; Robinson and Berridge, 1993), the test-retest reliability of our most commonly used task has not been extensively studied.

Attentional bias is of theoretical interest to obesity researchers, but to advance the field researchers must understand more about the reliability of the tests being used. Without consistency across measures, it is possible that the observed differences across studies could be due to differences in the way that attentional bias is assessed from study to study. The aim of the present study was to assess the reliability of the visual probe task. As such, the present study was conducted: (i) to compare attentional bias scores obtained from the lexical and pictorial version of the visual probe task using the different methods of scoring, and (ii) to assess the test-retest reliability of the pictorial probe task. We chose to examine the visual probe task as it is the most commonly implemented task assessing food-related attentional bias. Given the exploratory nature of this study, no specific hypotheses were tested.

MATERIALS AND METHODS

Overview

Data for this study are drawing from a larger study designed to investigate the relationship between attentional bias, impulsivity, and real-world eating patterns[1]. The full data set and the code book can be accessed here: https://rdp.utas.edu.au/metadata/2c3122be-fc62-48d0-a42d-41875d21e71b. In addition to completing a series of laboratory attentional bias tasks—the focus of this paper—, participants in the larger study were also required to track their eating and drinking using a smartphone application during a 2-week real-world monitoring period (similar to that described in Schüz et al., 2015). The results from this field-based monitoring will be reported elsewhere. The study was approved by the Tasmanian Social Sciences Human

[1] https://osf.io/b9t2h

Research Ethics Committee prior to the first participant being enrolled (H0018038).

Participants

Potential participants were recruited through a mixture of social media advertisements (see: Frandsen et al., 2013) and flyers placed near the study site. Advertisements called for individuals interested in participating in a study examining eating patterns. Eligibility criteria included being aged 18–75, having no history of eating disorders, not currently dieting, and having a body mass index (BMI) > 18.5. Individuals with concerns regarding body weight, shape and/or eating (as measured by a score of > 20 on the Eating Attitudes Test; Garner and Garfinkel, 1979) were excluded and referred to their general practitioner. We used recruitment targets to ensure that the final sample contained approximately equal numbers of participants in the healthy-weight and high-BMI range. The final sample consisted of 69 participants (67% female) aged 18–71 (M = 30.67, SD = 11.71), of whom 35% were in the healthy weight range (BMI = 18.5–24.9), 29% in the overweight range (BMI = 25.0–29.9), and 36% in the obese range (BMI > 30).

Procedure

The full procedure for the present study has been described elsewhere (Franja et al., 2020) and mirrored the protocol used in earlier studies (Elliston et al., 2016). Briefly, after recording their age, anthropometric measurements and contact details in an online portal, eligible participants were invited to visit the lab and received information about the study and provided consent. In line with previous research (Kemps et al., 2014; Werthmann et al., 2014), participants were instructed to eat a light meal up to 2 h prior coming into the lab visits to ensure they were satiated upon arrival. Hunger was assessed at the beginning of study visits using a 100-point hunger scale (Castellanos et al., 2009; Loeber et al., 2012).

During this initial study visit, participants were asked to complete the pictorial probe task followed by the lexical probe task (described in greater detail below). Participants were then issued with a study-specific electronic diary for a field-based monitoring component of the study (data not reported here). Approximately 14 days after this initial session, participants returned to the lab and completed the pictorial probe task for a second time. Tasks at both sessions were completed seated approximately 50 cm in front of a 21.5 inch monitor using Inqusit 5 (Inquisit 5, 2016). Participants were individually tested in two single sessions of approximately 30 min in duration in a well-lit room in the University of Tasmania's Clinical Research Facility[2]. After testing, participants were thanked and reimbursed AU$60 for their time.

Materials

Lexical Probe Task

The task consisted of 20 food words and 60 animal words. Food words included both high-caloric (e.g., hamburger, brownie), and low-caloric (e.g., broccoli, apple) words. Animal words were made up of species generally not consumed in Western cultures (e.g., cat, hamster). The critical trials were made up of 20 food words paired with animal words, whilst the control trials were made up of animal words paired with other animal words. Based on previous research using this task (Kemps et al., 2014), all word pairs were matched for ratings of valence and arousal, as well as the number of letters and syllables. In addition to the critical (food—animal) and control (animal—animal) trials, an additional 14 word-pairs consisting of stationery items (e.g., pencil, stapler) were used for practice and buffer trials. Participants were asked to place their left index finger on the "T" key and their right index finger on the "B" key. Each trial began with a fixation cross presented in the center of the screen for 500 ms. Following this, word pairs were presented for 500 ms. All words were presented centrally, one above the other, black Arial on a white background, in lower case. After the word presentation, a visual probe ("X") replaced one of the previously presented words (i.e., either top or bottom). Participants were asked to indicate as quickly as possible (by hitting the relevant keys) which word the probe replaced (top or bottom). The probe remained on the screen until a response was made. The intertrial interval was 500 ms. The whole task consisted of 12 practice, 2 buffer, and 160 experimental trials. During the experimental trials, each of the critical (food—animal) and control (animal—animal) trials were presented four times, at each of the word location (top or bottom) and probe location (top or bottom) combinations to ensure that the probes replaced each of the words in each pair equally. The lexical probe task was completed once during the initial study visit.

Pictorial Probe Task

The pictorial task mirrored the lexical task, however, using pictures of food instead of words. All picture pairs were matched for ratings of valence and arousal (the rating based on results of pilot study by Kemps et al., 2014), as well as perceptual characteristics such as brightness and complexity. Unlike the lexical task, picture pairs were presented on either side of the central position. Participants were asked to place their left index finger on the "E" (to signal if the probe was on the left) key, and their right index finger on the "I" key (to signal if the probe was on the right), and to indicate as quickly as possible whether the probe replaced the right or left image. Intertrial interval and picture presentation time mirrored that of the lexical task. This pictorial task was completed once during the initial study visit and then repeated a second time following the field-based monitoring portion of the study.

Hunger Scales

To ensure participants had complied with instructions to eat a light meal up to 2 h prior testing and were satiated upon arrival, a modified version of the hunger scale (Castellanos et al., 2009; Loeber et al., 2012) was administered. Only the two relevant subscales measuring time since last eaten (an estimate to the nearest 15 min), and current level of hunger (rated on a sliding scale from 100 = *Not hungry at all* to 100 = *Extremely hungry*) were included.

[2]https://www.menzies.utas.edu.au/research/the-clinical-research-facility

RESULTS

In accordance with standard protocols (Kemps et al., 2014), mean reaction times (RTs) for critical (food—animal) trials were calculated after deletion of incorrect responses and outliers (i.e., RTs < 150 ms or $> 1,500$ ms, or RTs exceeding the individual's mean + three standard deviations) for both the lexical and pictorial tasks. This resulted in deletion of 1.06% of the trials for the lexical task, and 1.03% (session 1) and 1.04% (session 2) of trials for the pictorial tasks. Control (animal—animal) trials were also discarded. An attentional bias score was calculated for each participant in three ways: "all food" (RT_{animal}—$RT_{allfood}$) "high-caloric" (RT_{animal}—$RT_{high\text{-}caloriefood}$) and "low-caloric" ($RT_{animal}$—$RT_{low\text{-}caloriefood}$). For all three calculations, positive values indicated attentional bias toward food related stimuli. Mean attentional bias scores for both tasks and each of the different stimulus types are shown in **Table 1** and the reaction times are shown in **Table 2**.

Comparison of Attentional Bias Scores Obtained From Different Tasks and Scoring Methods

To address our first aim, we compared attentional bias scores obtained from our two different visual probe tasks (lexical and pictorial), and three common scoring methods (all food, high-caloric and low-caloric), using data gathered during the initial study visit. **Table 3** shows the correlation matrix of attentional bias scores calculated during the initial study visit. The tasks and scoring methods produced significant variation in the measure of attentional bias obtained. Within task correlations were highest when compared to all foods for both the pictorial and lexical tasks. However, comparisons between lexical and pictorial tasks were weak regardless of which scoring method was used. Across participants, the average correlation across the six scores was 0.57. Comparing the two tasks, the three scorings of the pictorial probe task showed slightly higher agreement.

Test-Retest Reliability of the Pictorial Probe Task

To address our second study aim, we compared attentional bias scores obtained from the pictorial probe task at the initial study visit to those obtained when the task was re-administered at the final study visit (∼14-days later). Again, we used the three different attentional bias scoring procedures for the task, yielding attentional bias scores for all food stimuli, high-caloric food stimuli and low-caloric food stimuli. Regardless of the scoring procedure used, the test-retest reliability of the task was poor (low-caloric: $r = 0.41$; high-caloric: $r = 0.20$; all food: $r = 0.40$). As can be seen in **Table 2**, above, participants demonstrated faster response times to both food [$t(68) = 5.62$, $p < 0.001$] and animal [$t(68) = 5.82$, $p < 0.001$] stimuli in session 2 compared to session 1. A paired samples *t*-test confirmed that hunger levels remained consistent across both sessions [Session 1: $M = 26.83$, Session 2: $M = 23.39$; $t(69) = 1.02$, $p = 0.312$].

DISCUSSION

The aim of the present study was to assess the reliability of the visual probe task. We compared attentional bias scores from two of the most commonly used tasks—lexical and pictorial visual probe—using three different scoring methods. Our second aim was to evaluate the test-retest reliability of the pictorial probe task. The correlation between the lexical and pictorial tasks was weak. These findings are consistent with previous research (Freijy et al., 2014), and suggest that task type influences outcome—possibly contributing to the mixed findings within the literature assessing attentional bias using the probe task. The pictorial task yielded a wider range of attentional bias scores, with faster RTs to food compared to animal stimuli. The lexical task yielded similar RTs to both food and animal stimuli. This is in line with the notion that cues presented in picture form are more easily recognized—a phenomenon known as the superiority effect (Shepard, 1967;

TABLE 1 | Attentional Bias scores for all tasks, by stimulus category.

	All food	High-caloric food	Low-caloric food
	Mean (*SD*) [Range]	Mean (*SD*) [Range]	Mean (*SD*) [Range]
Lexical task	0.42 (18.27) [−37.69 −70.66]	0.72 (21.40) [−65.88 −72.89]	0.21 (22.02) [−52.49 −68.67]
Pictorial task 1	5.73 (19.29) [−33.94 −80.05]	5.27 (22.04) [−51.51 −65.25]	6.16 (21.52) [−32.32 −94.07]
Pictorial task 2	4.66 (14.72) [−28.10 −42.29]	4.39 (16.62) [−15.96 −43.93]	4.87 (17.99) [−34.85 −56.45]

All food (RT_{animal}—RT_{food}), high-caloric (RT_{animal}—$RT_{high\text{-}caloric\ food}$) and low-caloric ($RT_{animal}$—$RT_{low\text{-}caloric\ food}$).

TABLE 2 | Mean reaction time for all tasks, by stimulus category.

	All food	Animal	*p*	High-cal food	Low-cal food	*p*
	M (*SD*)	M (*SD*)		M (*SD*)	M (*SD*)	
Lexical task	426.01 (75.82)	426.43 (73.72)	0.848	426.20 (76.32)	426.14 (72.88)	0.855
Pictorial task 1	411.30 (76.36)	417.02 (77.78)	0.016*	411.75 (76.03)	410.86 (77.98)	0.716
Pictorial task 2	386.00 (59.86)	390.66 (61.35)	0.011*	386.27 (62.55)	385.79 (58.71)	0.826

*$p < 0.05$.

TABLE 3 | Correlations between the lexical and pictorial task attentional bias scores, both measured in session 1.

	Pictorial	Pictorial	Pictorial	Lexical	Lexical	Lexical
	All food	**High-cal**	**Low-cal**	**All food**	**High-cal**	**Low-cal**
Pictorial: all food	1.00					
Pictorial: high-cal	0.89	1.00				
Pictorial: low cal	0.88	0.57	1.00			
Lexical: all food	0.20	0.12	0.24	1.00		
Lexical: high-cal	0.23	0.15	0.26	0.84	1.00	
Lexical: low-cal	0.10	0.05	0.13	0.85	0.42	1.00

Pictorial, pictorial task; Lexical, lexical task; high-cal, high caloric food images; low-cal, low caloric food images.

Snodgrass et al., 1972), and suggests that pictorial stimuli may be more useful for capturing attentional engagement.

The test-retest reliability for the pictorial probe task was also poor, regardless of how the attentional bias scores were calculated. This is an important finding, given that the visual probe task is frequently used in measuring food-related attentional bias. Poor test-retest reliability for the visual probe task is in line with previous findings on attentional bias measures in threat/anxiety (for a review, see Schmukle, 2005; MacLeod et al., 2019), and alcohol (Christiansen et al., 2015) research. Aday and Carlson (2019) found that the correlations between test-retest in the first two testing sessions were low, but increased over repeated testing sessions. Additionally, the attentional bias indexes from the later sessions correlated more strongly with participants' trait anxiety scores, suggesting that extended testing may not only improve reliability, but that participants need extensive experience with the tasks in order for such biases to emerge. It is important to note, however, that the task Aday and Carlson (2019) used included personally relevant threatening stimuli. It has been previously demonstrated that using personally relevant stimuli increases internal reliability. For example, Christiansen et al. (2015) found that attentional bias toward personalized alcohol-related stimuli was larger than attentional bias to general alcohol-related stimuli, and, increased the internal reliability of the visual probe task. Future work in this area may like to consider making the food-related visual probe task more personalized to each participant by assessing food preference prior to testing. Additionally, in line with threat research carried out by Aday and Carlson (2019), it may be worthwhile assessing food-related attentional bias over multiple sessions, and correlate these results to participants' more stable trait characteristics such as eating styles (Newman et al., 2008). However, it is important to consider that this approach may increase the risk of inflating assessment reactivity—potentially altering an individual's attentional response style (MacLeod et al., 2019).

When calculating different attentional bias scores, the low-caloric and all food attentional bias scores had higher test-retest reliability compared to the high-caloric food attentional bias score. This is partially in line with previous research which found that all food attentional bias scores had the highest test-retest reliability (van Ens et al., 2019). Given the differences in stimulus presentation times between the present study and van Ens et al. (2019), further research is required to determine the influence of high-vs. low-caloric images on the reliability of the visual probe task. It is also possible that the improved reliability observed with the all food measure was simply due to it having a greater number of trials.

The present findings are in contrast to a recent study which found high test-retest reliability of attentional bias for food using the visual probe task (van Ens et al., 2019). van Ens et al. (2019) reported acceptable test-retest results for all food ($r = 0.835$) and high-caloric food ($r = 0.611$) RT indices. It is possible that the improved reliability was due to the longer stimulus presentation time of 3,000 ms, as it has been suggested that longer presentation times can improve reliability of time-reaction tasks (Waechter et al., 2014). However, it is important to note that longer stimulus presentation time (such as that used by van Ens et al., 2019) reflects the maintenance of attention rather than automatic attentional engagement (Mogg et al., 2004; Nijs and Franken, 2012). Theoretical accounts (such as incentive sensitization theory; Berridge, 2009) regarding food-related attentional bias suggest that this bias is driven by an automatic processing of food-related cues, which is why shorter stimulus presentation times (500 ms) are more common—unless specifically examining sustained attention (for example, see Nijs et al., 2010b).

Limitations

The task parameters used in the present study were based on previous research with a stimulus presentation time of 500 ms (Kemps et al., 2014). Although this presentation time is commonly used (for example, see Kemps and Tiggemann, 2009; Ahern et al., 2010; Calitri et al., 2010; Nijs et al., 2010b; Kakoschke et al., 2015; Meule and Platte, 2016), using only the one presentation time is a limitation of the present study. While this presentation time has been used to measure "initial orientation" (Calitri et al., 2010), it has been suggested that attentional orienting occurs anywhere between 30 and 500 ms, disagreement at 500–1,000 ms, and avoidance at presentation times above 1,000 ms (Ouimet et al., 2009). Therefore, 500 ms presentation time could be tapping into either attentional orienting or disengagement. It is possible that during the 500 ms presentation time where two images are presented simultaneously, that multiple shifts of attention (i.e., attentional disengagement, shift, and engagement with new object) may occur (Doolan et al., 2015). As such, it has been argued that 500 ms presentation time does not reflect automatic orientation, but rather represents the cost of information processing by the attentional control mechanism (Starzomska, 2017). It has therefore been suggested that only very short presentation times (<500 ms) can provide insight into initial orientation of attention (Starzomska, 2017). Future studies could compare test-retest of both <500 ms and >500 ms presentation time to see which of these attentional processes may be more stable. Moving on, when examining the relationship between the lexical and pictorial probe tasks, it is important to note that although there was a high level of comparability between words and images, the stimuli were not 100% identical. Future studies may want to ensure that the stimuli are identical across tasks to minimize any confounding variables.

Contrary to previous findings (e.g., Doolan et al., 2015), participants were generally faster at responding to probes replacing low-caloric food items (i.e., fruits, vegetables, salads) than high-caloric food items (i.e., brownie, waffle, chips) at both testing sessions (see **Table 2**). This highlights another potential limitation; the current sample was made up of healthy participants with low/non-existent rates of disordered eating. Given that we expect food-related attentional bias to work in similar ways to alcohol-related attentional bias (i.e., based on theoretical models underlying addiction), it is possible that food-related attentional bias is more prevalent in those with pathological eating habits. Attentional bias scores may have higher reliability with individuals with underling eating pathology (who in turn are more likely to demonstrate higher levels of attentional bias toward palatable foods), as a higher range of true scores results in higher reliability (Waechter et al., 2014). However, studies assessing attentional bias have found increased attentional bias toward food in healthy individuals with obesity (i.e., Nijs et al., 2010b; Kakoschke et al., 2014; Kemps et al., 2014). As such, we should still expect that the attentional bias score obtained would be consistent across measures (particularly given that two thirds of the present sample were made up of individuals with overweight and obesity); something that we did not observe in our study. Also, findings suggest that variables such as affect and self-exertion also impact attentional bias toward food-related cues (Frayn et al., 2016; Pollert and Veilleux, 2018). It is plausible that some of these variables may have affected task performance between the two testing sessions. However, it is unlikely that such states would have varied enough between testing sessions to account for the poor test-retest reliability observed. Nonetheless, future studies should consider measuring and controlling for such state-like variables when assessing test-retest reliability. It is important to note that the effect of state-like variables on food-related attentional bias challenge the theoretical underpinning of attentional biases, which suggests that attentional bias should be relatively stable. This contrast between underlying theory and published findings on the effects of differing variables on attentional bias require closer examination.

The fact that this is a secondary analysis of a larger study examining real-world eating patterns also leads to limitations. As part of the larger study, participants underwent an intensive ∼14 day monitoring period during which they recorded all food and drink intake. It is possible that the monitoring may have influenced participants' performance at the final attentional bias assessment. **Table 2** shows that participants did generally have faster response times in session 2. However, given that response times shortened for both food and animal stimuli, this highlights the possibility that performance may have been affected by practice effects rather than cue reactivity. Another possible limitation regarding practice effects is that the participants were presented with the lexical and pictorial tasks in the same order. This lack of counterbalancing may have influenced response times. As the lexical probe task was always completed second, fatigue may have also influenced performance on the lexical task. Future research should replicate these findings using a counterbalanced design. Additionally, it is possible this 14-day period may have affected the results in other ways. It may be useful for future work to compare task performance following shorter periods to get a clearer picture of the effect of time on task performance.

CONCLUSION

To conclude, the present study found correlations between the lexical and pictorial probe tasks to be weak. Furthermore, the test-retest reliability of the pictorial task was poor—regardless of how the attentional bias scores were calculated. Going forward, alternate measures of attentional bias should be explored [e.g., electrophysiological monitoring; findings suggest that event-related potentials capturing early attentional engagement have good reliability (Hagan et al., 2020)]. Finally, for attentional bias measures to be of any practical use, it would be useful to assess whether attentional bias is associated with real-world eating patterns.

DATA AVAILABILITY STATEMENT

The full dataset and the code book can be accessed here: https://rdp.utas.edu.au/metadata/2c3122be-fc62-48d0-a42d-41875d21e71b.

ETHICS STATEMENT

The studies involving human participants were reviewed and approved by Tasmanian Social Sciences Human Research Ethics Committee. The patients/participants provided their written informed consent to participate in this study.

AUTHOR CONTRIBUTIONS

SFr and SFe developed the study concept and contributed to the study design. SFr and AM collected the data. SFr performed the data analysis and interpretation under the supervision of SFe. SFr drafted the manuscript. All authors provided revisions and approved the final version of the manuscript for submission.

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A Randomised Controlled Trial of Inhibitory Control Training for Smoking Cessation: Outcomes, Mediators and Methodological Considerations

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Objective: Inhibitory control training (ICT) has shown promise for improving health behaviours, however, less is known about its mediators of effectiveness. The current paper reports whether ICT reduces smoking-related outcomes such as craving and nicotine dependence, increases motivation to quit and whether reductions in smoking or craving are mediated by response inhibition or a devaluation of smoking stimuli.

Method: Adult smokers (minimum 10 cigarettes per day; $N = 107$, $M_{age} = 46.15$ years, 57 female) were randomly allocated to receive 14 days of smoking-specific ICT (named INST; a go/no-go task where participants were trained to not respond to smoking stimuli) or active control training (participants inhibited responding toward neutral stimuli). Participants were followed up to 3-months post-intervention. This trial was preregistered (Australian and New Zealand Clinical Trials Registry ID: ACTRN12617000252314; URL: https://www.anzctr.org.au/Trial/Registration/TrialReview.aspx?id=370204).

Results: There were no significant differences between ICT and active control training groups. Specifically, participants in both groups showed significant reductions in craving, nicotine dependence, motivation and a devaluation (reduced evaluation) of smoking-stimuli up to 3-months follow-up compared to baseline. Inhibition and devaluation of smoking stimuli did not act as mediators. Devaluation of smoking stimuli was an independent predictor of smoking and craving at follow-up.

Conclusion: Inhibitory control training (ICT) was no more effective at reducing smoking-related outcomes compared to the active control group, however, significant improvements in craving, dependence indicators and evaluation of smoking stimuli were observed across both groups. A return to basic experimental research may be required to understand the most effective ICT approach to support smoking cessation.

Keywords: smoking, inhibitory control, craving, cognitive training, e-health, devaluation, response inhibition

INTRODUCTION

Growing literature has suggested that inhibitory control, which is the ability to inhibit automatic prepotent responses, is impaired in smokers (Smith et al., 2014) and this may contribute to difficulties with quitting. Inhibitory control training (ICT) using the go/no-go task (GNG) has been found to reduce unhealthy food and alcohol consumption more effectively than stop signal tasks (SST) (Allom et al., 2015; Jones et al., 2015). Inhibitory control training using the GNG aims to improve inhibition by training participants to refrain from initiating a response toward salient stimuli (e.g., alcohol or unhealthy food). However, less is known about ICT's effectiveness for assisting with smoking cessation. Adams et al. (2017) found that one session of lab-based ICT did not lead to greater reductions in smoking at 1-week post-training in smokers who were not focused on quitting compared to a control group. Our research group (Staiger et al., 2018) recently reported on the smoking outcomes of a 2-week online smoking-specific ICT program with nicotine-dependent individuals. No intervention effect was found for cigarette consumption or cessation; however, exploratory analyses provided initial evidence that ICT may assist with smoking reduction for individuals aged under 36 years (Bos et al., 2019).

The current paper reports on the pre-registered secondary outcomes and mediation analyses to complement the above-mentioned primary intervention outcomes (Bos et al., 2019). Although current data indicates that smoking-specific ICT has no significant effect on smoking cessation, it is important to consider that ICT may have improved other important smoking-related outcomes such as craving (Gass et al., 2014). Furthermore, higher nicotine dependence has been shown to be a predictor of failed cessation attempts (Vangeli et al., 2011), and has been associated with poor inhibition (Smith et al., 2014). In contrast, higher motivation to quit is associated with making a quit attempt (Vangeli et al., 2011). To date, no study has assessed the effects of ICT on these important smoking-related outcomes and doing so may clarify whether ICT helps to facilitate smoking cessation.

With respect to mediation, two potential mechanisms have been proposed for ICT. Firstly, that the consistent pairing of target stimuli with not responding (like in GNG-based ICT) results in "learnt" (associative) inhibition (Jones et al., 2015). Indeed, a meta-analysis found that higher percentages of successful inhibitions during ICT resulted in a greater effect size of ICT for reducing alcohol and food consumption (Jones et al., 2015). However, it was not reported whether improvements in individuals' stimulus-specific inhibition was maintained post-ICT or whether this inhibition acted as a mediator of smoking-specific outcomes. Secondly, another potential mechanism is that ICT devalues (i.e., reduces the positive valence) target stimuli. Evidence for this has been reported in three ICT studies: one food-related (Lawrence et al., 2015), one related to smartphone applications (Johannes et al., 2020) and one targeting smoking (Scholten et al., 2019), although the latter did not report on smoking outcomes. The present study tested both of these potential mechanisms.

For secondary outcomes we hypothesised that, after completing the intervention, smokers who were randomised to the smoking-specific ICT intervention called INST would report significantly less: (1) craving for cigarettes; and (2) nicotine dependence compared to those in the active control group. We also examined two potential mediators of changes in smoking frequency and craving over time: (1) a devaluation of smoking stimuli; and (2) improved response inhibition to smoking stimuli. We also report findings for one pre-registered exploratory hypothesis, which was that smokers who received ICT would report significantly higher levels of motivation to quit smoking compared to smokers in the active control group.

METHOD

A detailed protocol for this parallel, two-group, double-blind block randomised controlled trial (RCT; Staiger et al., 2018) and smoking frequency and cessation primary outcomes (Bos et al., 2019) have been recently published. This study received ethics approval (DU-HREC Project Number 2015-298) and was pre-registered (Australian New Zealand Clinical Trials Registry ID: ACTRN12617000252314). Key details are provided below.

Participants

Eligible participants (see **Table 1**) were traditional tobacco cigarette (tailored or hand-rolled) smokers (n = 107) aged 18 to 60 years (M = 46.15, SD = 9.38, range = 20–60) who over the past 12-months smoked at least 10 cigarettes per day (M = 18.79, SD = 6.93, range = 10–44) and met criteria for moderate (n = 41) or severe (n = 66) Tobacco Use Disorder according to the *Diagnostic and Statistical Manual of Mental Disorders* (5th ed.; DSM-5; American Psychiatric Association, 2013). Participants had: completed at least Year 9 (or equivalent) schooling; a desire to quit smoking; motivation to make a quit attempt during the training stage of the intervention; and regular computer and internet access.

Participants were excluded if they primarily used e-cigarettes; had not smoked for 2 weeks or more in the past 3 months; were using psychotropic (e.g., antidepressant, antipsychotic or anxiolytic) or anti-craving medications (e.g., varenicline or bupropion); used nicotine replacement therapy (NRT) during the training phase; engaged in problematic alcohol and/or drug(s) use other than tobacco; or had a history of traumatic or acquired brain injury or a loss of consciousness of over 30 min.

Go/No-Go Training Tasks

The online intervention was based on a modified smoking-specific GNG ICT task (therein referred to as ICT; see **Figure 1**) designed originally by Lawrence et al. (2015). Inhibitory control training consisted of nine smoking-related images (100% no-go) and nine images of relaxing activities such as sitting by a river or lying in a hammock (100% go) and 18 neutral images of clothing (50:50 go/no-go). The control GNG task was identical to ICT except that the stimuli consisted of 18 neutral images only (e.g., household items). For the ICT, images of relaxing activities were chosen for go trials as compared to alternatives for

TABLE 1 | Demographics of participants at baseline.

Variable	Intervention (*n* = 54) M (SD) or %	Control (*n* = 53) M(SD) or %
Age*	46.20 (9.73)	46.09 (9.10)
Gender (% female)*	55.55	50.94
Age commenced smoking*	16.69 (2.41)	15.75 (2.43)
Relationship status (% in a relationship)	74.07	75.47
Country of birth (% Australia)[a]	79.63	88.68
Education (% tertiary educated)*[b]	70.37	62.26
Employed (% yes)*	81.48	79.25
Cigarettes consumed per day*	18.12 (7.12)	19.48 (6.74)
Household with other regular smokers (% yes)	40.74	37.74
Parents who were regular smokers (%)		
Both	27.28	33.96
One	55.56	45.28
Neither	16.67	20.75
In the past 12 months, number of:		
Quit attempts	1.41 (1.69)	1.81 (3.10)
Different types of quit aids used (%):		
None	35.19	20.75
1 type	31.48	45.28
2 types	18.52	28.30
3+ types	14.81	5.66
DSM-5 Tobacco use disorder symptoms*	6.59 (2.11)	6.57 (1.86)
No. ICT/Control training sessions completed*	10.50 (2.91)	10.89 (3.20)

**Published in Bos et al. (2019).*
[a]One person (total) also identified as Australian Aboriginal and/or Torres-Strait Islander.
[b]Five people (total) also identified as students.

appetitive behaviours like alcohol (i.e., non-alcoholic beverages) or high calorie/high fat foods (e.g., fruit and vegetables) as there are no clear health alternative behaviours for smoking (see Guo, 2018). Neutral images of clothing acted as filler images to increase engagement and difficulty, and to reduce the likelihood of participants identifying the associative patterns within the task (see Lawrence et al., 2015).

Each 10-min computer training session consisted of six blocks of 36 trials (50:50 go/no-go). Each image was displayed once only. At the end of each block, participants were provided with task performance feedback (accuracy and go reaction time) and were encouraged to try to beat their own scores.

Measures

The following psychometrically reliable and valid measures were used: smoking-related stop signal task (SST; Logan et al., 1997), Fagerström Test of Nicotine Dependence (FTND; Heatherton et al., 1991), Timeline Follow-Back (TLFB; Robinson et al., 2014), and visual analog scales for craving, motivation and the stimulus evaluation test (also see Staiger et al., 2018). Stop signal reaction time (SSRT), a measure of response inhibition from the SST (see **Figure 2**) was calculated using the mean method (Logan et al., 1997). Data of participants who exhibited accuracy outside of 40%–60% on stop trials and/or less than 70% on go trials during the SST was excluded [based on Congdon et al. (2012)]. The FTND was used as a more phasic measure as our interest was in changes over time.[1] Smoking frequency measured using the TLFB was defined as mean number of cigarettes consumed per day. The stimulus evaluation test (adapted from Lawrence et al., 2015) involved participants rating the valence of each image of smoking and relaxing activities from the ICT intervention on a 100 mm visual analog scale. Craving for cigarettes and motivation to quit smoking were also measured using visual analog scales. For the abovementioned measures, higher scores indicated higher severity of nicotine dependence or frequency of smoking, poorer inhibition or stronger behavior (e.g., more motivation or more strongly valued visual stimuli). Time (in hours) since last cigarette smoked prior to each training (ICT or active control) session was also collected.

Procedure

Smokers were screened for eligibility via phone or online survey. Eligible participants were instructed to abstain from smoking for 2 h prior to meeting with a researcher at the university, where they provided informed consent, completed the baseline assessment (TLFB, questionnaires, SST) and were reminded that they were required to make a quit or cessation attempt during the 2-week training period. Participants were told the aim of the study was to “investigate which of the two tasks was more effective” to minimise unblinding. They were then randomised by the computer program’s inbuilt algorithm, and completed their first ICT or control GNG training session with the researcher present to ensure they understood the task.

Following session one, participants were emailed a web link to access the online training and instructed to complete this training once per day for the next 13 days and in a quiet place whilst craving cigarettes. More frequent use or making up for missed sessions was not enabled within the program. SMS reminders to complete training were sent twice per week during the training period.

At all three follow-ups (post-intervention, 1-month and 3-months), participants completed the TLFB via phone with a researcher (naïve to group allocation), then completed the online questionnaires and the SST. After the completion of each follow-up participants received a $20 gift card.

Analytic Strategy

Any methods not reported in this brief report are detailed in the **Supplementary Material**. Multiple Imputation by Chained Equations (MICE; Azur et al., 2011) was used to address missing data. Analyses covaried for age due to a potential age effect (Bos et al., 2019) on the outcomes.

Secondary Outcomes

Analyses of secondary outcomes used mixed effects linear regression models with random intercept (to account for clustering of time points within individuals) in Stata 15 (StataCorp, 2017). Specifically, we regressed each outcome on to

[1]Participants with 100% days abstinent at a follow-up period (Bos et al., 2019) did not complete the FTND (n_{total} = 3 at 1 month, n_{total} = 5 at 3 months post-intervention), and were deemed to have a FTND total score of 0.

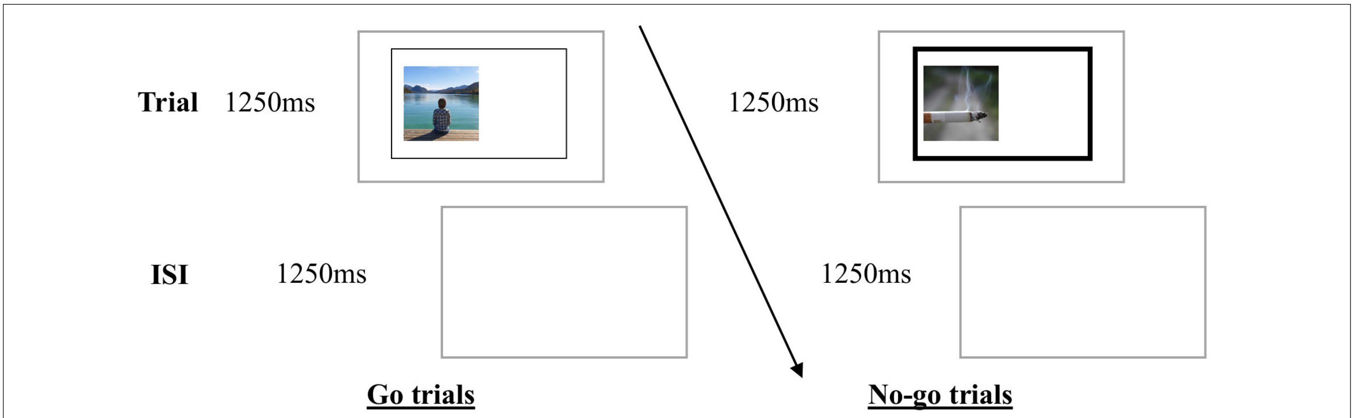

FIGURE 1 | Go/no-go ICT and Active Control Training. Images were presented within a rectangle, followed by an inter-stimulus-interval (ISI). Participants were instructed to indicate as quickly and as accurately as possible the location of an image within the rectangle (left or right) by pressing a computer key (C or M, respectively) when the rectangle was not bolded (go trials). However, when the lines of the rectangle were bold (no-go trials), participants were instructed to refrain from responding.

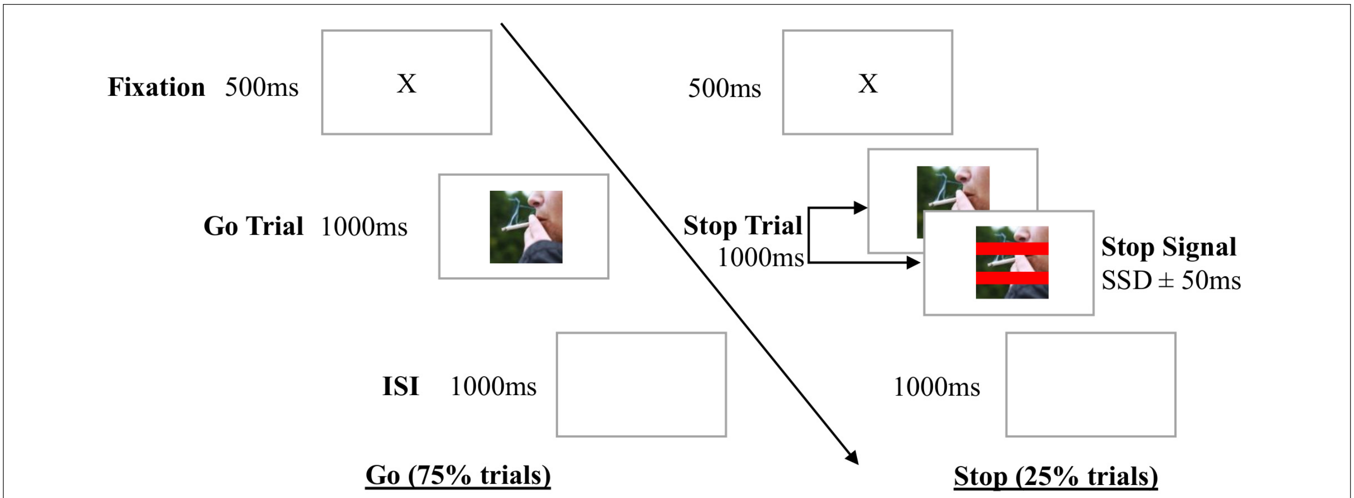

FIGURE 2 | Stop Signal Task (SST). Go stimuli were eight pairs of images of smoking, with one image presenting the cigarette pointing left, and the other its mirror image- with the cigarette pointing right. Each SST begun with a practice block of 10 trials, followed by a test block of 192 trials. After a fixation cross, participants were presented with an image of smoking, followed by an inter-stimulus interval (ISI). Participants were instructed to indicate whether the "lit" or "burnt end" of the cigarette was pointing left or right by pressing a computer key. If red lines (stop signal) appeared across the image, participants were instructed to not respond (stop trials, 25% trials). The stop signal appeared at a short delay (stop signal delay or SSD) after the go stimulus, which began at 250 ms on the first stop trial, and then adjusted by 50 ms in a staircase manner (increased after successful stop trials, or decreased after failed stop trials) so participants had approximately 50% accuracy and converged on a mean SSD.

a variable denoting timepoint (i.e., baseline vs. post-intervention vs. 1-month follow-up vs. 3-month follow-up; note that only baseline and post-intervention available for SST), group (i.e., intervention vs. control), and a timepoint × group interaction. Effect size was measured using Cohen's d_z for paired data and Cohen's d for between group effects. This analysis was then repeated after removing smokers who had abstained at any time after the training period ($n_{total\ removed} = 6$).

Mediation Analyses

Mediation analyses were conducted in Mplus Version 8 (Muthén and Muthén, 2017) using four time point autoregressive longitudinal mediation models (MacKinnon, 2008). Briefly, this involved estimating autoregressive and cross lagged paths between the outcomes (e.g., smoking) and mediation effects over time. We also explored whether (1) evaluation of smoking stimuli and (2) inhibition had direct effects on smoking frequency and craving, independent of group status using mixed effects multiple regression models.

RESULTS

The final intent-to-treat sample was 107 (two participants withdrew, and one was removed for using NRT during the training period [see Bos et al. (2019) for details]. Compliance for smoking no less than 2 h prior to each of the fourteen ICT or control training sessions was 60.64%, with 39.36% smoking less than 2 h before training and 35.30% smoking between 2 and 3 h beforehand. There were no significant differences between groups (ICT vs. active control) in demographics and secondary outcomes at baseline, and no main effects of group for any secondary outcomes (any results not provided within this report were detailed in the **Supplementary Material**). There were no group by timepoint interactions in predicting secondary outcomes; however, we found a significant main effect of timepoint on craving, $F(3, 1009.7) = 8.56$, $p = \leq 0.001$; nicotine dependence, $F(3, 1096.5) = 21.69$, $p < 0.001$; motivation to quit $F(3, 1166.1) = 13.14$, $p < 0.001$; and evaluation of smoking stimuli, $F(3, 1457.3) = 13.43$, $p < 0.001$ (see **Table 2**). Specifically, both groups showed reductions in craving, nicotine dependence and evaluation of smoking stimuli at all follow-up time points compared to baseline; with mostly moderate to large effect sizes (d_z from −0.31 to −0.91). Reductions in motivation to quit smoking showed small to moderate effect sizes (d_z from −0.27 to −0.67).

Analyses repeated with abstinent participants removed ($n_{\text{intervention}} = 49$; $n_{\text{control}} = 52$) showed that reductions in craving, nicotine dependence, evaluation of smoking stimuli and motivation to quit were not different to the main analyses using the full dataset. There were no changes in smoking-specific inhibition across time points observed in either group.

Analysis of the longitudinal autoregressive mediation models showed no evidence to suggest that changes in smoking frequency or craving were mediated by changes in inhibition or evaluation of smoking stimuli ($p = 0.44$ to 0.99 for paths of interest: see **Supplementary Figures 1–4** in **Supplementary Material** for further details). Mixed effects regression models revealed that the change in evaluation of smoking stimuli between baseline and post-intervention significantly predicted smoking and craving at all follow-ups, independent of group (see **Supplementary Material**). Changes in response inhibition did not independently predict smoking and craving at follow-ups. Additional exploratory analyses found that neither age nor the change in motivation to quit over the training period acted as moderators (three-way interaction with time and condition) or independent predictors (interaction with time) for improvements in craving and nicotine dependence (see **Supplementary Material** for detailed method and results).

DISCUSSION

This paper reported on the RCT outcomes of online smoking-specific ICT for heavy dependent smokers: craving, nicotine dependence and motivation to quit. Additionally this paper examined whether (1) evaluation of smoking stimuli and/or (2) inhibition acted as mediators between groups and (a) smoking or (b) craving. ICT was no more effective than the active control group for improving smoking-related outcomes, and both groups showed significant reductions in craving, nicotine dependence, and devaluation of smoking stimuli at all follow-ups compared to baseline. Furthermore, both groups reported *reduced* motivation to quit at 1-month and 3-month follow-ups – opposite to what might have been expected from the decline in smoking-related outcomes. Importantly, inhibition and smoking devaluation did not act as mediators between ICT and observed reductions in smoking or craving, with devaluation instead acting as an independent predictor of reductions in smoking and craving across all follow-ups. This suggests that devaluation in smoking may not be driven by ICT.

While no effect of group was found, the current study observed overall small-moderate significant reductions in craving and moderate-large significant reductions in nicotine dependence across all follow up time points. These effects occurred alongside an overall significant reduction in cigarette consumption, as reported in Bos et al. (2019) and were still present with abstainers removed. While findings from the present study were contrary to the hypotheses, failure to detect differences in craving or dependence between groups may be due to general inhibition training effects from using an active control task and/or self-monitoring of cigarette use (see Bos et al. (2019) for discussion). Significant reductions in craving and nicotine dependence in the absence of abstainers suggests that findings were not solely driven by those who had quit. It is also important to note that sustained reductions in cigarette consumption has also been observed in the absence of quit attempts (Yong et al., 2012). However, a number of other important methodological issues warrant discussion as they have important implications for how we might interpret these non-significant findings.

It is important to consider the potential influence that nicotine satiation may have on the measurement of inhibition. Charles-Walsh et al. (2014) found that smokers at 3-h abstinence did not display deficits in response inhibition, and Tsaur et al. (2015) suggested that deficits may not appear until as late as 72 h nicotine abstinence. Whereas, Grabski et al. (2016) found that smokers who were abstinent for at least 10 h displayed these deficits in inhibition. These findings align with evidence that nicotine improves inhibition (as measured by the SST) in healthy non-smokers (Logemann et al., 2014). Additionally, healthy controls have displayed increased activation in prefrontal regions (measured using fMRI) during successful inhibition on stop trials of the SST after nicotine administration (Kasparbauer et al., 2019). Taken together, this suggests that the neurochemical effects of nicotine may improve inhibition and potentially mask inhibitory deficits, which do not appear until at least 10 h post cigarette consumption. In the current study, the majority of participants smoked 3 h or less prior to each training session. It is possible that when satiated, nicotine may have nullified potential deficits in inhibition, making efforts to improve inhibition redundant. This is a potentially serious limitation in the effectiveness of ICT with smokers unless they have been abstinent for a few days. This demonstrates the critical need to return to laboratory style studies to investigate and understand the relationship between nicotine satiation and inhibition.

Another important consideration specifically relates to the measurement of inhibition. In the current study, training used the GNG whereas inhibition was measured using the SST. Whilst both are measures of inhibition, the SST arguably measures

TABLE 2 | Estimated marginal means (EMM) and standard errors (SE) of each secondary outcome by group over time.

		Baseline	Post-intervention		1-Month follow-up		3-Month follow-up	
	Group × Time interaction on outcome	**EMM (SE)**	**EMM (SE)**	**d_z [95% CI]**	**EMM (SE)**	**d_z [95% CI]**	**EMM (SE)**	**d_z [95% CI]**
Craving	$F(3, 1440.5) = 0.58, p = 0.63$							
Intervention		46.44 (3.71)	32.03 (3.97)	−0.44** [−0.72, −0.16]	32.03 (4.04)	−0.45** [−0.73, −0.17]	34.54 (4.71)	−0.31* [−0.58, −0.03]
Control		47.99 (3.68)	31.87 (4.24)	−0.48** [−0.76, −0.19]	35.67 (4.39)	−0.35* [−0.63, −0.08]	29.71 (5.00)	−0.47** [−0.75, −0.18]
FTND	$F(3, 1392.6) = 0.98, p = 0.40$							
Intervention		5.41 (0.30)	3.80 (0.32)	−0.78*** [−1.08, −0.47]	3.94 (.35)	−0.64*** [−0.93, −0.34]	3.75 (0.35)	−0.72*** [−1.02, −0.42]
Control		5.72 (0.31)	3.79 (0.33)	−0.91*** [−1.22, −0.58]	4.57 (0.35)	−0.50*** [−0.78, −0.21]	4.43 (0.40)	−0.48** [−0.76, −0.19]
SST								
SSRT	$F(1, 578.0) = 0.01, p = 0.92$							
Intervention		264.70 (8.82)	264.72 (10.64)	0 [−0.27, 0.27]	–		–	
Control		258.61 (9.33)	256.84 (12.66)	−0.02 [−0.29, 0.25]	–		–	
Go RT	$F(1, 452.1) = \leq 0.01, p = 0.97$							
Intervention		662.68 (13.63)	652.52 (19.25)	−0.07 [−0.33, 0.20]	–		–	
Control		661.82 (13.82)	652.52 (16.69)	−0.07 [−0.34, 0.20]	–		–	
Evaluation of Images								
Smoking	$F(3, 1638.0) = 0.30, p = 0.83$							
Intervention		51.53 (3.26)	34.32 (3.74)	−0.60*** [−0.89, −0.31]	36.20 (3.75)	−0.53*** [−0.82, −0.25]	33.83 (3.86)	−0.60*** [−0.89, −0.31]
Control		51.76 (3.30)	37.15 (3.86)	−0.50*** [−0.78, −0.21]	41.89 (3.99)	−0.33* [−0.60, −0.05]	36.97 (4.05)	−0.48*** [−0.77, −0.20]
Relaxing activities	$F(3, 2435.6) = 0.98, p = 0.40$							
Intervention		72.68 (2.31)	75.43 (2.56)	0.16 [−0.11, 0.42]	73.21 (2.61)	0.03 [−0.24, 0.30]	72.84 (2.50)	0.01 [−0.26, 0.28]
Control		79.04 (2.33)	76.34 (2.64)	−0.15 [−0.42, 0.12]	78.15 (2.80)	−0.05 [−0.32, 0.22]	78.50 (2.61)	−0.03 [−0.30, 0.24]
Motivation	$F(3,1196.5) = 0.72, p = 0.54$							
Intervention		79.25 (3.93)	70.77 (4.29)	−0.26 [−0.53, 0.01]	65.04 (4.58)	−0.42** [−0.69, −0.14]	61.29 (5.26)	−0.46** [−0.74, −0.18]
Control		82.80 (3.97)	74.03 (4.37)	−0.27*[a] [−0.54, 0.01]	70.12 (4.60)	−0.37** [−0.65, −0.09]	56.94 (5.20)	−0.67*** [−0.97, −0.37]

The presented within groups d_z is that specific time point compared to baseline. Negative d_z and smaller EMM than baseline denote an improvement in secondary outcomes. EMM = Estimated Marginal Means; SE = Standard Error; CI = confidence interval. FTND = Fagerström Test of Nicotine Dependence, SST = stop signal task, SSRT = stop signal reaction time (measure of response inhibition), Go RT = reaction time on go trials. All analyses were adjusted for age.

$^*p < 0.05$; $^{**}p < 0.01$; $^{***}p < 0.001$. $^{*a}p = 0.049$.

top-down inhibition or action cancellation, whereas the GNG is thought to measure automatic bottom-up inhibition or action restraint (Verbruggen and Logan, 2008; Swick et al., 2011; Littman and Takács, 2017). Interestingly, Jones et al. (2018) also found no change in alcohol-specific inhibition following ICT when measured using an alcohol-specific SST. The use of SST to assess inhibition and the effectiveness of GNG ICT may be problematic as the two tasks are thought to assess different aspects of inhibition. Future ICT research should therefore consider employing measures of automatic inhibition, such as slowed response latency to respond to formerly no-go associated stimuli (e.g., Best et al., 2016).

Both groups (ICT and active control) showed significant reductions in the positive evaluation of smoking cues at all follow-ups when compared to baseline. This is similar to findings by Scholten et al. (2019), where smokers showed a devaluation of smoking cues immediately after one ICT session, however, no follow-ups were conducted. Findings of the current study builds upon this by showing that devaluation of smoking cues was maintained long after their last training session, and predicted reductions in smoking frequency and craving. Contrary to study predictions, this effect was also observed in the active control group (not exposed to smoking images) and not just the ICT group. Future studies are needed to clarify whether this may have occurred due to non-specific trial effects (which may have reduced smoking or craving) on devaluation, or a reduction in reactivity to smoking cues alongside cigarette cessation or reduction (Balter et al., 2015). Studies could also consider an active control condition where participants are exposed to smoking images without ICT. Despite the limitations, the observed devaluation of smoking stimuli and its effects on smoking and craving is encouraging and warrants further consideration in future studies.

Other aspects of the study design require consideration. The sample size was powered for the primary outcomes in the expectation of moderate effect sizes (which may have been optimistic), and underpowered for detecting small mediation effects. This limits the interpretation of the non-significant mediation effects on smoking or craving. However, this study was strong in that cigarette consumption was self-report via face-to-face and phone interviews using detailed time line follow-back, which increased accuracy. Although these data collection methods may have been affected by social desirability bias (Latkin et al., 2017; Zhang et al., 2017), there is some suggestion that collecting information regarding consumption (Yeager and Krosnick, 2010) and quit attempts (Persoskie and Nelson, 2013) from smokers is not affected by such bias, and we see no basis to expect differential bias between groups. The observed drop in nicotine dependence post-training, which includes overall estimates of consumption (collected via online survey), was consistent with detailed reported cigarette consumption. This increases our confidence that consumption was measured appropriately, although a biochemical measure for verifying cigarette consumption was not used (Connor Gorber et al., 2009). Future studies should consider including biochemical verification methods of tobacco use, e.g., cotinine, to confirm self-reported cigarette consumption.

An RCT examining alcohol-specific ICT has also reported non-significant findings (Jones et al., 2018). It is possible that unlike the success of ICT for food intake, smoking and alcohol consumption may not be impacted by ICT. Alternatively, further research into intervention design is needed before any conclusions about ICT for reducing smoking and alcohol consumption can be drawn (e.g., types of stimuli for both intervention and control conditions, number of sessions and tailored stimuli type [Staiger and White, 1991)]. It has recently been suggested that the selection of images used as healthy stimuli in contrast to the target behaviour may be important, with Manning et al. (2021) reporting that recipients of cognitive bias modification showed increased approach bias toward non-alcoholic beverages concurrently with increased avoidance bias toward alcohol. Using images of relaxing activities (the current study) and neutral stimuli [e.g., stationary in Adams et al. (2017)] for opposing images to cigarettes have both produced non-significant findings. Future experimental studies could consider trialing various alternative images to cigarettes (e.g., nicotine replacement therapies) to see if these improve the effectiveness of ICT for smoking cessation.

In conclusion, results of the current study suggested that there is no benefit of smoking-specific ICT compared to an active control group, as both groups showed improved craving and nicotine dependence at all follow-ups, and reduced motivation to quit at 1-month and 3-months post-intervention. In addition, no evidence of inhibition or devaluation of smoking stimuli acting as mediators was found, with stimulus devaluation instead independently predicting improvements in smoking and craving. Potentially methodological issues such as non-specific trial effects, nicotine satiation and choice of inhibition measure may have contributed to the reported findings. Therefore, future studies should consider employing an experimental design and addressing these methodological issues.

DATA AVAILABILITY STATEMENT

The datasets presented in this article are not readily available because the authors do not have ethics approval to make the dataset public. Requests to access the datasets should be directed to PS, petra.staiger@deakin.edu.au.

ETHICS STATEMENT

The studies involving human participants were reviewed and approved by Deakin University Human Research Ethics Committee (DUHREC), Deakin University, Geelong, Australia. The patients/participants provided their written informed consent to participate in this study.

AUTHOR CONTRIBUTIONS

LH was responsible for the conceptualisation of the secondary outcomes and mediator variables, data analysis, and preparation of the first draft of the manuscript. PS, MH, and NL conceptualised the study. NL also contributing the model for ICT and active control task development. GY contributed statistical analysis expertise. RB provided input for smoking measures and data interpretation. LH and JB were involved in data collection. PS provided senior oversight of the project. All authors commented on the final manuscript.

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SUPPLEMENTARY MATERIAL

The Supplementary Material for this article can be found online at: https://www.frontiersin.org/articles/10.3389/fpsyg.2021.759270/full#supplementary-material

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Are Emotion Regulation Strategies Associated With Visual Attentional Breadth for Emotional Information in Youth?

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Attentional deployment is currently considered as one of the most central mechanisms in emotion regulation (ER) as it is assumed to be a crucial first step in the selection of emotional information. According to the broaden-and-build theory, positive emotions are associated with attentional broadening and negative emotions with attentional narrowing toward emotional information. Given that ER strategies relying on attentional deployment (i.e., rumination, cognitive reappraisal and distraction) have the possibility to influence positive and negative emotions by (re)directing one's attention, there could be an association with one's attentional scope. The current study investigated the association between the general (trait) use of three specific ER strategies and visual attentional breadth for positive, negative, and neutral information in a selected sample of 56 adolescents ($M = 12.54$, $SD = 1.72$; 49% girls) at risk for developing psychopathology. First, participants self-reported on their overall use of different ER strategies. Next, the previously validated Attentional Breadth Task (ABT) was used to measure visual attention breadth toward emotional information. No evidence was found for the relationship between 2 specific ER strategies (i.e., cognitive reappraisal and rumination) and visual attentional breadth for neutral, positive and negative emotional information. Surprisingly, "distraction" was associated with visual attentional narrowing, which was unrelated to the valence of the emotion. These unexpected results indicate the multifaceted relationship between trait ER, distraction specifically, and visual attentional breadth for emotional information. Future research, especially in younger age groups, could further elaborate on this research domain.

Keywords: emotion regulation, attentional breadth, emotion regulation strategies, children, adolescents

INTRODUCTION

Since the early 1980s, a growing number of studies have been devoted to the role of emotion regulation (ER) in exploring new research avenues within the field of clinical psychology. ER is often defined as *"the processes by which we influence which emotions we have, when we have them, and how we experience and express those emotions"* (Gross, 1998). Given that the experience of both negative and positive emotions occurs on a daily basis, it is indispensable for an individual to effectively regulate these emotions with the aim of sustaining one's emotional wellbeing. It is referred to as emotion *dys*regulation when an individual is unable to adequately regulate one's affect (Izard, 1978; Nezlek and Kuppens, 2008; Grommisch et al., 2019).

The past few decades, emotion *dys*regulation has become a topic of increasing interest as it is linked to 75% of the diagnostic categories of psychopathology described by the Diagnostic and Statistical Manual of Mental Disorders (DSM 5, Kring and Sloan, 2009). Focusing specifically on children and adolescents, emotion *dys*regulation has repeatedly been associated with symptoms of depression, anxiety, aggressive behavior, and eating- and weight related disorders (e.g., obesity) (Shields and Cicchetti, 1998; Southam-Gerow and Kendall, 2000; Sim and Zeman, 2006; Aldao and Nolen-Hoeksema, 2010; Vandewalle et al., 2016). Because of its strong relationship with a plethora of mental health disorders across different age groups, emotion *dys*regulation has been labeled a transdiagnostic mechanism across psychopathology in both children and adolescents, as well as adults (Kring and Sloan, 2009; Aldao and Nolen-Hoeksema, 2010; Fernandez et al., 2016; Schäfer et al., 2017).

From a developmental perspective, ER is a rapidly and constantly evolving process (Thompson and Goodman, 2010). Early in life, children mostly rely on primary caregivers to externally regulate their emotions (e.g., searching comfort with their parents when being sad or looking at the facial expression of a significant other after falling down) (Morris et al., 2007). Still, studies have shown that infants and pre-schoolers are also able to *autonomously* regulate basic emotions (e.g., anger and fear) through the use of behavioral strategies (e.g., playing with toys to distract themselves or cuddle with their favourite bear to comfort themselves) (Kopp, 1982; Eisenberg and Morris, 2002). As cognitive and emotional functioning matures, children additionally start to use more sophisticated cognitive strategies to internally and independently regulate their emotions (e.g., re-evaluation or acceptance of the situation) by the age of eight to nine (Harris, 1989). Although the development of ER appears linear from this point of view, research shows a maladaptive shift in ER during the adolescent phase due to the heightened emotional reactivity and various stressors characterizing this developmental period. This shift is characterized by a general decrease in the use of adaptive ER strategies such as problem solving and cognitive reappraisal and an overall increase in maladaptive strategies such as rumination and aggression (Cracco et al., 2017). Given this remarkable shift in ER, studying ER in adolescence is of key importance. Moreover, such research should pay attention to gender differences in ER as remarkable differences occur from childhood to adolescence under the influence of biological and contextual factors, i.e., girls are generally better at regulating their emotions than boys in childhood, whereas boys appear to do better in adolescence (Morris et al., 2002; Cracco et al., 2017).

Differences can be observed in the specific ways individuals regulate both positive and negative emotions using various ER *strategies* (Weiss et al., 2019). Besides ER strategies for regulating positive emotions (for review see Carl et al., 2013), this study particularly focuses on ER strategies for regulating negative emotions. Next, a distinction is often made between adaptive and maladaptive ER strategies – based on their general (trait) and long-term associations with psychopathology (Aldao and Nolen-Hoeksema, 2012). Adaptive ER strategies lead to a decrease in negative affect (and conversely a restoration or an increase in positive affect), as well as greater levels of psychological well-being in the long-term (Aldao and Nolen-Hoeksema, 2012). Some renowned adaptive ER strategies for regulating negative emotions in children and adolescents are distraction, cognitive reappraisal, problem solving, and acceptance (Cicchetti, 2010; Werner and Gross, 2010; Aldao et al., 2014). On the flipside, maladaptive ER strategies are associated with an increase in negative and a decrease in positive affect and is associated with more overall psychopathology in the long-term (Gross, 2002; Aldao et al., 2010; Werner and Gross, 2010). Examples of maladaptive ER strategies for regulating negative emotions used by both adults, children and adolescents are avoidance, rumination and suppression (Aldao et al., 2010; Aldao and Nolen-Hoeksema, 2010, 2012). Although this categorization has proven its validity in previous research (e.g., Aldao et al., 2014; Schäfer et al., 2017), recent studies point out that when it comes to the use of these strategies in daily life, a more nuanced approach should be considered for the evaluation of the adaptiveness of an ER strategy, taken into account numerous other co-determining factors (e.g., the context in which these strategies are used, the flexibility with which they are employed) (Aldao et al., 2015).

Initially, ER was conceptualized as a one-dimensional process that solely involved the control or elimination of negative emotions. Throughout time, it has evolved toward a multidimensional construct in which different cognitive, attentional and behavioral processes are considered (Cole et al., 2004; Gratz and Roemer, 2004; Rajappa et al., 2012). One of the leading frameworks describing the emotion regulatory processes is the 'process model' of Gross (1998). The model largely distinguishes between five groups of regulation processes that are placed on a temporal dimension being (1) *situation selection* or the ability to influence emotions by selecting or avoiding specific situations in which emotions can occur, (2) *situation modification* or the ability to influence emotions by changing external, physical characteristics of the situation, (3) *attentional deployment* or the process of directing the attention in a particular way *within* a situation in order to influence the situation-specific evoked emotions, which can take many forms including *physical withdrawal of attention* (e.g., covering the eyes), *internal redirection of attention* (e.g., redirecting your thoughts) and *external redirection of attention* (e.g., someone pointing something out), (4) *cognitive change* or the ability to influence emotions by reappraising the emotion-eliciting situation, and lastly (5) *response modulation* or the ability to influence physiological and/or behavioral responses associated with the emotion. It is generally assumed that problems in ER can occur in each of the phases in the ER process (Gross and Thompson, 2007; Koole, 2009).

In contrast to situation selection and/or situation modification, attentional deployment is a regulatory process through which emotions can be controlled without modifying or changing the situation itself. Gross and Thompson (2007) state that attentional processes may be a fundamental mechanism of (a) the emotion generative and (b) the emotion regulation process. More specifically, attentional deployment can be considered a crucial first step in the selection of emotional information and influence the ability to engage in specific ER strategies, but also serves as a central mechanism within certain ER strategies (i.e., cognitive reappraisal) (Gross and Thompson, 2007; Koole, 2009; Wadlinger and Isaacowitz, 2011; Todd et al., 2012).

Competencies and deficits in attentional deployment show many individual differences. Competencies continue to develop throughout childhood and become more prominent with age (Fox and Calkins, 2003; Ochsner and Gross, 2005; Gross and Thompson, 2007; Godara et al., 2020). As early as the age of five, attentional deployment is considered an important regulatory skill (e.g., children distract their selves when being separated from their mother by deploying their attention away trough play) (Sethi et al., 2000; Mischel and Ayduk, 2004). As individuals grow older and reach early adolescence, they become more aware of how they can autonomously manage their emotions by deploying their attention away and will eventually be using even more (complex) strategies that rely on this process (e.g., coping with a fearful situation by taking in the information and consequently deploying attention toward alternative interpretations of the event) (Gross, 2013; Ahmed et al., 2015).

Deficits in attentional deployment can occur at an early age. In one study conducted in 7-year-old clinically anxious children, results showed that there was a significantly larger attentional focus on emotional information (i.e., angry faces) in clinically anxious children compared to children with no or lower levels of anxiety symptoms (Taghavi et al., 1999; Waters et al., 2008). However, evidence on deficits in attentional deployment is largely limited to adults (e.g., Power and Dalgleish, 2015). In general, these studies show that the onset and maintenance of different forms of psychopathology (e.g., depression and anxiety) is associated with the navigation of attention toward negative emotional information, resulting in the maintenance or intensification of concurrent negative emotional states.

Given that both deficits in attentional deployment and ER more broadly are associated with psychopathology, studies unraveling the association between attentional in an emotional context and ER strategy use can provide crucial information for both research and clinical practice (Fox and Calkins, 2003; Ochsner and Gross, 2005; Gross and Thompson, 2007; Godara et al., 2020). Nevertheless, only few studies regarding these associations have been conducted, especially in younger age groups. One study using a visual attentional probe task tentatively suggests that children who are able to control their attention and consequently prevent themselves from focusing on negative information, generally use more adaptive ER strategies and report low levels of psychopathology (Waters et al., 2008). Furthermore, a recent experimental study using a visual attentional breadth task in a selected sample of adolescents (Van Beveren et al., 2020), investigated the link between visual attention across neutral and emotional information contexts and ER in (early) adolescents. Participants were presented various Emoji (i.e., positive-, negative-, and neutrally valenced) and were afterward instructed to correctly identify peripheral information (= dependent variable) on a computer screen (i.e., detect a black circle that was simultaneously presented close or far from the Emoji). Results showed that the general (trait) use of adaptive ER strategies was associated with a broadened attentional scope for neutral (but not for positive or negative) information. This in turn was related to more positive affect and overall emotional well-being (Fredrickson and Branigan, 2005; Van Beveren et al., 2020). However, additional research is pivotal to further invigorate these findings, especially in at risk youth.

EMOTION REGULATION STRATEGIES AND ATTENTIONAL SCOPE

Common trait ER strategies that predominantly rely on attentional deployment, and may therefore be associated with one's attentional scope in an emotional context, are cognitive reappraisal, rumination and distraction (Gross, 2013).

Cognitive Reappraisal

Cognitive Reappraisal is generally categorized as an adaptive ER strategy and refers to the way in which a positive perspective or reinterpretation of the situation can decrease negative affect and/or increase positive affect (McRae et al., 2012; Van Cauwenberge et al., 2017). In general, cognitive reappraisal has been linked to the fourth phase of the ER process (i.e., cognitive change). However, more recent studies (e.g., Strauss et al., 2016; Sanchez-Lopez et al., 2019) show that the quality of reappraisal depends on the earlier process of attentional deployment. More specifically, the ability to focus on negative emotional information, followed by disengaging the attention away from this information and shifting the attention toward other, more positive or neutral, emotional information, is crucial for the successful use of this ER strategy. In both youth and adults, an underutilization or ineffective use of cognitive reappraisal has been associated with higher depressive and anxiety symptoms (Aldao et al., 2010; Dryman and Heimberg, 2018).

Rumination

Rumination is considered a maladaptive ER strategy in which there is a repetitive focus on (negative) emotions, its causes, as well as its consequences (Nolen-Hoeksema et al., 2008; Gross, 2013). Rumination involves directing the attention toward the negative emotion(s) but is, in contrast to cognitive reappraisal, accompanied with negative beliefs (e.g., negative emotions are unacceptable) (Gross, 2013). Rumination has mainly been associated with internalizing problems such as depression and anxiety in both youth and adult populations (Aldao et al., 2010).

Distraction

Distraction represents the intentional shift in attention from one's negative emotions toward an external situation or stimulus (Gross, 1999; Scheibe et al., 2015) and has been found to be an adaptive ER strategy across various studies (Sheppes et al., 2011). However, research shows that distraction can have a rather ambiguous effect on emotions and emotional well-being (Gratz and Tull, 2011; Craske and Barlow, 2014). When distraction is used to completely avoid negative emotions, instead of being used in a flexible and adaptive way to redirect the attention away, it may have limited short-term advantages and even adverse long-term consequences (Gross, 1998; Van Dillen and Papies, 2015; Wolgast and Lundh, 2017).

According to the broaden-and-build theory of positive emotions (Fredrickson, 1998, 2001), attention can be (re)directed. This means that the attentional scope, which refers to a range of thoughts, perceptions, and actions that occur in a certain situation, can be narrowed or broadened (i.e., *attentional breadth*) (Whitmer and Gotlib, 2013). A broadened attentional scope is associated with the intake of peripheral stimuli and global information (e.g., seeing a barking dog but also noticing it is on a leash and the owner is with him) whereas a narrowed attentional scope is associated with processing only central stimuli and local information (e.g., focusing solely on the sharp teeth of the dog) (Sung and Yih, 2016; Wronska et al., 2018).

In this theory it is stated that specifically the attentional scope related to negative and positive emotions has different effects on both cognition and psychophysiology. Studies have shown that positive emotions (e.g., optimism and happiness) are associated with a more broadened attentional scope, which is considered a protective factor against stressful events (Fredrickson and Joiner, 2002; Rowe et al., 2007). Negative emotions (e.g., anger and anxiety) however, are associated with a more narrowed attentional scope which is related to emotional disorders (e.g., anxiety) (Fredrickson and Joiner, 2002). Although the mechanism behind this process is still under study, attentional narrowing is often seen as an evolutionary threat-driven response that is beneficial when it leads to quick and decisive action. Nowadays, feelings of stress-related anger, and anxiety are often not related to real life-threatening stressors. Yet, the response remains the same. As a result, the disproportional and continuous reactivity in response to stress that is characterized by narrowed attention often leads to maladaptive and rigid action (e.g., aggressive actions) and thinking patterns (e.g., ruminating about the stressful event), which eventually can cause emotional problems (Gasper and Clore, 2002; Gasper, 2004; Yoon et al., 2015; Gu et al., 2017).

Based on the broaden-and-built theory of positive emotions (Fredrickson, 2001), it can be assumed that particularly ER strategies that rely on the process of attentional deployment (i.e., cognitive reappraisal, rumination, distraction) will be associated with the breadth of one's attentional scope. As the goal of adaptive ER strategies such as cognitive reappraisal is to restore/increase positive and decrease negative affect, it is plausible to assume an association with visual attentional *broadening* (Aspinwall, 1998; Fredrickson and Joiner, 2002), whereas for maladaptive ER strategies such as rumination an association with visual attentional *narrowing* is theoretically presumed (Nolen-Hoeksema et al., 2008).

To our knowledge, no existing study provides direct evidence for the relationship between the attentional scope and two specific ER strategies (i.e., cognitive reappraisal, and distraction). Based on the aforementioned line of thought and existing literature (see e.g., Sanchez-Lopez et al., 2019), it is to be expected that cognitive reappraisal is associated with a broadened attentional scope as this strategy involves the ability to take in negative emotional information and subsequently shift attention toward more positive or neutral information (Strauss et al., 2016). Given that distraction involves actively shifting the attention away from negative emotional information (Scheibe et al., 2015), one could expect an association with a broadened attentional scope toward emotional information. However, as distraction is considered a rather *ambiguous* ER strategy that becomes *un*helpful when used in the long run to avoid negative emotions, an association with a narrowed attentional scope toward emotional information is also plausible (Förster et al., 2006). Lastly, based on the attentional scope model of rumination (Whitmer and Gotlib, 2013) which states that rumination involves a perseverative focus on negative emotional information, an association between rumination and attentional narrowing is to be expected (Grol et al., 2015; Fang et al., 2017). Unfortunately, evidence for these theoretical propositions is scarce and limited to adults.

THE CURRENT STUDY

The purpose of the current study is to examine the relationship of the general (trait) use of specific ER strategies and the attentional scope in youth (9–15 years) in an experimental study. The broaden-and-build theory of Fredrickson (1998, 2001) states that positive emotions are associated with attentional broadening and negative emotions with attentional narrowing (Aspinwall, 1998; Fredrickson and Joiner, 2002; Rowe et al., 2007). Given that ER has the possibility to influence both positive and negative emotions which could be related to a broadened or narrowed attention scope (Aspinwall, 1998; Fredrickson and Joiner, 2002), and the fact that certain specific ER strategies (i.e., cognitive reappraisal, rumination, and distraction) greatly rely on attentional deployment, it is hypothesized that (1) the use of the adaptive strategy "cognitive reappraisal," in children and adolescents will be associated with a broadened attentional scope, (2) the use of the maladaptive strategy "rumination" will be associated with a more narrowed attentional scope (Grol et al., 2015; Fang et al., 2017), and (3) that the use of the ambiguous ER strategy "distraction," in children and adolescents, will either be associated with narrowed attention or with attentional broadening. Because of age and gender differences in ER, these differences will be considered in all analyses (Hyde et al., 2008).

MATERIALS AND METHODS

Participants

Fifty-six youth between 9 and 15 years (M = 12.54, SD = 1.72; 49% girls) were recruited (see **Table 1**). All youth were admitted to a residential treatment centre for a multidisciplinary obesity treatment [MOT; (Braet and Van Winckel, 2001)]. Although obesity is considered a pathology rather than a

TABLE 1 | Frequency and percentage of age and gender.

Age	N_{Total}	%	N_{boys}	$\%_{boys}$
9	1	2.4	1	100
10	6	14.6	3	50
11	5	12.2	1	20
12	7	17.1	5	71.4
13	7	17.1	6	85.7
14	10	24.4	4	40
15	5	12.	1	20
Total	**41**	**100**	**21**	**51.2**

type of psychopathology, previous research uncovered large co-morbidities with psychological problems such as depressive symptoms, low self-esteem, and behavioral problems (Braet et al., 1997; Strauss, 2000). Recent studies in youth highlight the association between the transdiagnostic mechanism emotion *dys*regulation and emotional and/or external eating (Harrist et al., 2013). Especially an underutilization of adaptive (i.e., reappraisal) as well as a frequent use of maladaptive ER strategies (e.g., rumination) is linked to emotional eating and seems to play a crucial role in eating- and weight-related pathology and related psychological problems (Kubiak et al., 2008; Evers et al., 2010; Vandewalle et al., 2016). Given the association between these strategies and attentional deployment, a more detailed study can give new insight in the development and occurrence of emotional problems within this sample (Gross, 2013). In the current study the Child Depression Inventory (CDI; Kovacs, 1992; Timbremont and Braet, 2002) was used to screen for cognitive, affective, and behavioral symptoms of depression. To interpret the total score of the questionnaire, a cut-off score of 12 was used to indicate the presence of mild to moderate symptoms of depression, a score above the cut-off of 16 indicates severe symptoms of depression. In current sample a mean score of 13.70 (*SD* = 6.54) ranging from 1 to 29 was found. Since the current study is part of a larger project (Debeuf et al., 2020) on studying mechanisms of ER and the effects of training ER in children and adolescents with obesity, this sample can be referred to as a convenient sample of youth at risk for developing psychopathology.

In the current study, inclusion criteria were used. First, youth could not be enrolled in the inpatient treatment for more than two months. Second, youth should master Dutch and/or French. All youth and their parents were informed about the procedure and their right regarding GDPR[1] prior to the study and assented on taking part. Participation was not remunerated.

Procedure

The research protocol was approved by the Committee of Medical Ethics (EC UZG 2018/0101). After obtaining informed consent, participants were asked to fill out a self-reported trait emotion regulation (i.e., FEEL-KJ) and a depression (i.e., CDI) questionnaire using an online platform hosted by the Research Unit on a computer of the treatment center (see measures below) prior to a lab-study that took place on the same day. On the day of testing, adolescents were instructed to first explore the lab with the aim of familiarizing themselves with the lab-setting. As previously mentioned, the current study was part of a larger experimental study, took approximately 45 min per participant and consisted of several phases. Adolescents completed an attentional breadth task (see measures below) after watching a white screen for 3 min, which served as a neutral baseline to prevent experiencing negative affect at the start of the experimental task (Patuzzo et al., 2003). After completing the experimental task for ± 40 min, children and adolescents were verbally instructed to complete an additional 5 min abdominal breathing exercise in order to ensure participants left the lab in a good state of mind (Verbeken et al., 2019). Youth self-reported on fluctuations in affect at 4 different time points during the experiment (see measures below).

Measures

Self-Report Questionnaires

Emotion Regulation Strategies

To assess the general use of ER the *Fragebogen zur Erhebung der Emotionsregulation bei Kindern und Jugendlichen* (FEEL-KJ); (Grob and Smolenski, 2005; Braet et al., 2013) was used. This questionnaire is used in children and adolescents between 8 and 18 years old and measures the way children and adolescents regulate feelings of anger, sadness, and anxiety. In total 90 items measure 15 specific ER strategies that can be categorized as adaptive (e.g., cognitive reappraisal), maladaptive (i.e., rumination), or external ER strategies (i.e., expression). These strategies are measured as trait ER strategies (i.e., the general use of these strategies when experiencing negative affect). For each of the emotions (i.e., anger, sadness, anxiety), the same 30 items are presented on a 5-point Likert scale (1 = almost never, 2 = rarely, 3 = occasionally, 4 = often, 5 = almost always). An example item is: "*When I am sad/angry/anxious, I accept what makes me angry*" A total score can be calculated for each of the strategies (measured with 2 items for each of the 3 emotions; scores ranging between min 0.6 and max 0.30), in which a higher score means the ER strategy is more often used. In addition, a total score can also be calculated for the total use of adaptive, maladaptive, and external ER. The FEEL-KJ is considered reliable and valid with an acceptable to good internal consistency over all subscales (Cronbach's alpha between 0.64 and 0.94). Furthermore, an acceptable to good test-retest reliability was reported with correlation coefficients between 0.76 and 0.90 (Cracco et al., 2015). In the current study, we focused on three strategies relying on attentional deployment: (1) cognitive reappraisal (FEEL-KJ-CR), (2) rumination (FEEL-KJ-RUM) and (3) distraction (FEEL-KJ-DIS). Each of the strategies was assessed with six items (two items per emotion), e.g., "*When I'm sad/angry/anxious, I tell myself that it is not that bad*" to measure cognitive reappraisal; "*When I'm sad/angry/anxious, I can't get this out my mind*" to measure rumination; "*When I'm sad/angry/anxious, I do something I like*" to measure distraction. Cronbach's alphas reveal acceptable to good internal consistency of 0.74, 0.72, and 0.81 for the three specific strategies.

[1]The General Data Protection Regulation (GDPR) implies the protection of natural persons with regard to the processing of personal data and the free movement of such data.

Affective States

To measure fluctuations in affect during the lab study, the intensity of positive and negative affect was obtained through Visual Analog Scales (VAS). Youth scored their feelings of happiness, sadness, anxiety, frustration, boredom, and anger on a scale from zero (not present) to a hundred (very present) (Aitken, 1969; Bond and Lader, 1974). This was measured on four different time points, i.e., before (1) and after (2) the neutral baseline, after (3) the Attentional Breadth Task and after (4) the breathing exercise.

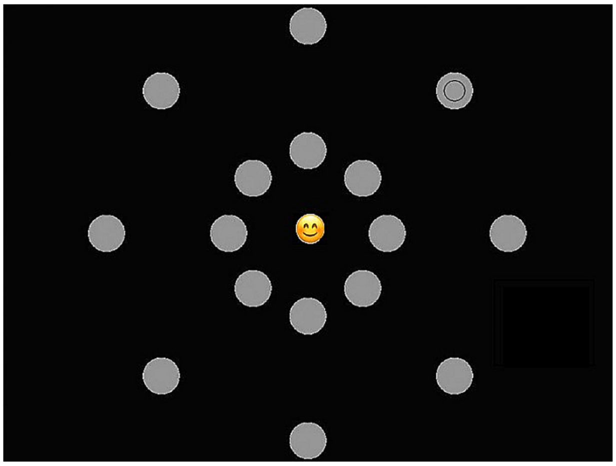

FIGURE 1 | Central Emoji and target stimulus of the attentional breadth task. Note: based on Van Beveren et al. (2020).

Experimental Task

Attentional Breadth Task

Visual attentional breadth in relation to centrally presented emotional stimuli was measured using an experimental task (Bosmans et al., 2009). The task has successfully been used in different studies to measure fluctuations in attentional broadening/narrowing related to centrally presented, relevant information (Bosmans et al., 2009; Grol and Raedt, 2014; Grol et al., 2015). Recently Van Beveren et al. (2020) adapted and evaluated this task using child friendly emotional stimuli (i.e., Emoji) to measure fluctuations in visual attentional breadth for emotional information in (early) adolescents. Participants were seated at a distance of 10.63 inches from a 19″ CRT computer screen, using a chin rest to guarantee correct positioning and distance to the screen. In each trial a picture of an Emoji (82 × 82 pixels) appeared in the centre of the screen (central Emoji; see **Figure 1**). The Emoji could be categorized as negative, positive, or neutral and eight validated Emoji were selected for each category. The selection of Emoji was based on both valence (i.e., negative and positive) and arousal (i.e., low arousal and high arousal) and was evaluated in previous research within a comparable age and gender sample of (early) adolescents (i.e., only Emoji that were clearly identified as either positive, negative or neutral are included) (Vanden Berghe et al., 2020).

When the central Emoji appeared, 16 gray dots with a diameter of 2 cm simultaneously surfaced in two concentric circles around the Emoji (see **Figure 1**). In total 8 imaginary axes appeared around the Emoji with two grays dots on each axe (compare **Figures 1–3**). One of the two gray dots appeared at 4.41 inches (i.e., far) from the central Emoji at 25° of the visual angle, the other gray dot appeared at 1.77 inches (i.e., close) from

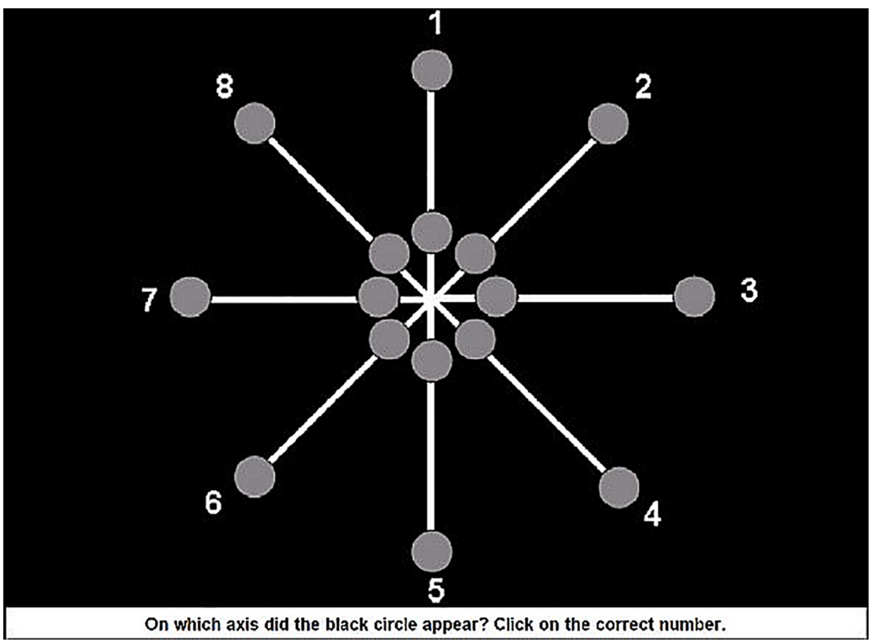

FIGURE 2 | Response screen axe of the target stimuli. Note: based on Van Beveren et al. (2020).

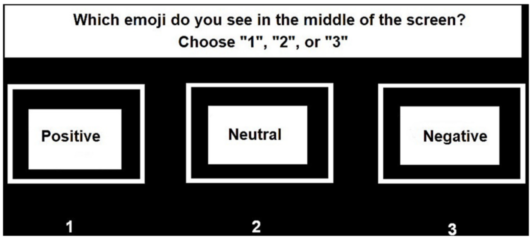

FIGURE 3 | Response screen valence of the central Emoji. Note: based on Van Beveren et al. (2020).

the central Emoji at 10° of the visual angle. At the same time of the presentation of both the emoji and gray dots, a smaller black circle with a diameter of 0.51 inches appeared in one of the 16 gray dots (target stimuli; see **Figure 1**). The black circle could thus be close or far from the Emoji. In order to prevent confounds of saccadic eye movements in search of the peripheral target (Ball et al., 1988), all stimuli were presented for 68 ms.

Participants were instructed to (1) correctly identify the valence of the presented central Emoji (i.e., negative, positive, and neutral), and (2) localize the black circle that appeared in one of the 16 gray dots. For both responses there was no time limit. After the presentation of the central Emoji and target stimuli, a first screen appeared and participants were asked to identify the valence of the central Emoji (see **Figure 2**). After giving this response, a second screen appeared after which participants immediately were asked to identify on which of the eight axes the target stimuli had appeared (see **Figure 3**). In the current study the main dependent variable was the accuracy on localizing the peripheral target stimuli. This was calculated based on trials in which the participants also correctly identified the valence of the central Emoji to make sure the participants maintained attention to the center of the screen during the task.

The task consisted of eight practice trials with a 250 ms presentation time, followed by eight practice trials with a shorter 68 ms presentation time. The test phase itself consisted of 144 trials with six types of trials randomly presented in three blocks of 48 trials each. The type of trials were based on the distance of the target stimuli to the central Emoji, as well as their valence (i.e., positive close, positive far, neutral close, neutral far, negative close, and negative far).

Data-Analytic Plan

For the analyses of the attentional breadth task, all trials were deleted in which the central Emoji was incorrectly identified, to make sure participants also maintained attention to the centre of the screen during the task. The dependent variable was the number of correctly identified target stimuli (black circle). Only trials with correctly identified central Emoji were included. Performance on the attentional breadth task was examined by performing a 3 Valence (Positive vs. Neutral vs. Negative) × 2 Distance (far vs. close) mixed ANOVA with the accuracy rates, i.e., number of correctly localized peripheral targets as the dependent variable.

Thereafter we calculated an Attentional Narrowing Index (ANI) by subtracting the proportion of correctly identified targets in the far trials from the proportion of correctly identified targets in the close trails for positive trials (posANI), neutral trials (neuANI), and negative trials (negANI) separately. Although our primary interest lays in attentional broadening, we calculated ANI scores to keep the outcome variables consistent and to enable comparison with the original task used in previous research (Bosmans et al., 2009; Grol and Raedt, 2014; Grol et al., 2015; Van Beveren et al., 2020). Higher ANI scores imply that more distance from the central picture leads to less correct answers and therefore can considered to be a measure of attentional narrowing/breadth. Next, gender and age effects were added as covariates into the model. Finally, we ran three additional 3 Valence (Positive vs. Neutral vs. Negative) × 2 Distance (Far vs. Close) mixed ANOVAs with FEEL-KJ-DIS, FEEL-KJ-RUM, and FEEL-KJ-CR as continuous predictors to test our research questions.

RESULTS

Preliminary Analyses and Group Characteristics

In total, on average 14.58% of the trials was deleted due to incorrect identification of the central Emoji. Participants ($n = 15$) were excluded from further analysis if the number of deleted trials for any of the different trial types was more than 50% (Grol and Raedt, 2014). Descriptive statistics for all study variables and bivariate correlations amongst these variables are presented in **Table 2**.

A non-parametric Friedman's ANOVA was performed as the data on the percentage of deleted trials was non-normally distributed. The ANOVA test revealed that if the target stimulus was presented close to the central stimulus, there were significant differences in the accuracy for identifying the central stimulus depending on the *emotional valence*, χ^2 (2, N = 41) = 33.90, $p < 0.001$. Follow-up Wilcoxon's signed-rank tests showed that participants made less errors when identifying neutral compared to positive stimuli ($Z = -4.31$, $p < 0.001$), as well as neutral versus negative stimuli ($Z = -4.39$, $p < 0.001$) when the target stimulus was presented close to the central stimulus. No difference between negative and positive stimuli was detected ($Z = -0.29$, $p = 0.769$).

A second ANOVA test revealed that, if the target stimulus was presented far from the central stimulus, significant differences occurred in accuracy for identifying the central stimulus depending on *the emotional valence*, χ^2 (2, N = 41) = 34.85, $p < 0.001$. Again, follow-up Wilcoxon's signed-rank tests showed that participants made less errors when identifying neutral compared to positive stimuli ($Z = -4.29$, $p < 0.001$), as well as neutral versus negative stimuli ($Z = -4.50$, $p < 0.001$) when the target stimulus was presented far from the central stimulus. No difference between negative and positive stimuli was detected ($Z = -0.96$, $p = 0.338$).

TABLE 2 | Descriptive statistics and bivariate correlations in the final sample.

	M (*SD*)	min – max	FEEL-KJ-CR	FEEL-KJ-RUM	FEEL-KJ-AS	FEEL-KJ-MS
Age	12.54 (1.72)	9 – 15				
FEEL-KJ-DIS	43.70 (10.01)	22 – 63	0.391*	0.019	0.786**	–0.230
FEEL-KJ-CR	47.39 (10.45)	28 – 71		0.336*	0.680**	0.369*
FEEL-KJ-RUM	49.98 (12.42)	27 – 80			0.230	0.758**
P-POS-close	0.48 (0.30)	0.00 – 0.96				
P-POS-far	0.22 (0.17)	0.00 – 0.70				
P-NEU-close	0.47 (0.26)	0.04 – 0.96				
P-NEU-far	0.23 (0.14)	0.04 – 0.50				
P-NEG-close	0.46 (0.27)	0.00 – 0.96				
P-NEG-far	0.21 (0.14)	0.00 – 0.71				

*FEEL-KJ-RUM = self-reported emotion regulation strategy "rumination," FEEL-KJ-CR = self-reported emotion regulation strategy "cognitive reappraisal," FEEL-KJ-DIS = self-reported emotion regulation strategy "distraction." P = proportion of correctly localized peripheral targets presented far or close when the central stimulus was positive (POS), neutral (NEU), or negative (NEG). *p < 0.05; **p < 0.01.*

Task Performance

The 3 Valence (Positive vs. Neutral vs. Negative) × 2 Distance (Far vs. Close) mixed ANOVA on the accuracy rates of target detection (see **Figure 4**) yielded a main effect of Distance, F (1, 40) = 67.80, $p < 0.001$, η_p^2 0.63, indicating that the number of correct identifications of the target stimulus was significantly higher when the target stimulus appeared close compared to far from the central stimulus (Mdiff = 5.20, SE = 0.63, p = 0.001), and a main effect of Valence, F (2, 80) = 6.19, p = 0.003, η_p^2 0.13, indicating that the number of correct identifications of the target stimulus was higher when the target stimulus was neutral compared to positive (Mdiff = 0.78, SE = 0.335, p = 0.025) or negative (Mdiff = 1.06, SE = 0.25, $p < 0.001$), but not when the target stimulus was positive compared to negative (Mdiff = 0.280; p = 0.417). Finally, no significant 3 Valence (Positive vs. Neutral vs. Negative) × 2 Distance (Far vs. Close) interaction effect was found, F (2, 80) = 0.096, p = 0.909, η_p^2 0.002.

Next, we added age and gender to the model in order to check whether these variables significantly affected the overall task performance. First, a significant Valence (Positive vs. Neutral vs. Negative) × Gender interaction was found, F (2, 76) = 4.86, p = 0.010, η_p^2 0.113, indicating that accuracy rates were higher for boys (M = 7.69, SE = 1.09) compared to girls (M = 6.147, SE = 1.12) when the target stimulus was positive (see **Figure 5**). Second, a significant Distance (Far vs. Close) × Age interaction occurred, F (1, 38) = 6.45, p = 0.015, η_p^2 0.145, revealing that age moderated the effect of distance on accuracy. Gender or age were implied in the 3 Valence (Positive vs. Neutral vs. Negative) × 2 Distance (Far vs. Close), all ps > 0.188.

Main Analyses

Visual Attentional Breadth and Emotion Regulation Strategies

The effects pertaining to the self-reported ER strategy variables (FEEL-KJ-DIS, FEEL-KJ-CR, FEEL-KJ-RUM) in the 3 Valence (Positive vs. Neutral vs. Negative) × 2 Distance (Far vs. Close) revealed a significant Distance (Far vs. Close) × FEEL-KJ-DIS, F (1, 37) = 5.09, p = 0.030, $\eta_p^2 = 0.121$ interaction. Distraction was associated with visual attentional narrowing when the target stimuli were presented far. No further effects were found.

Post-hoc Analyses

A paired-samples t-test was conducted to examine fluctuations in affect before and after completing the experimental task. There was a significant difference in frustration before (M = 21.88, SD = 32.32) versus after (M = 38.38, SD = 42.80) the task; $t(39)$ = –3,894, $p < 0.05$, in anger before (M = 12, SD = 24.75) versus after (M = 25.50, SD = 37.59) the task; $t(39)$ = –2,996, $p < 0.05$ and in boredom before (M = 50.50, SD = 40.59) versus after (M = 63.00, SD = 39.04) the task; $t(39)$ = –2,492, $p < 0.05$. Suggesting that youth were left significantly more frustrated, angry and bored after completing the ABT.

DISCUSSION

The current study examined the relationship between the general use of specific ER strategies (i.e., cognitive reappraisal, rumination, and distraction) that rely on attentional deployment and visual attentional breadth for negative, positive, and neutral emotional information in youth at risk for developing psychopathology. To measure ones' specific ER strategies on a trait level, the FEEL-KJ was used (Braet et al., 2013). To measure visual attentional breadth for emotional information a previously validated Attentional Breadth Task was included [ABT; (Bosmans et al., 2009; Grol and Raedt, 2014; Grol et al., 2015; Van Beveren et al., 2020)]. Three main findings regarding our proposed aims can be reported. First, no evidence was found for the association between trait "cognitive reappraisal" and attentional *broadening* for emotional information. Second, no evidence was found for the relationship between trait "rumination" and attentional *narrowing* for emotional information. Third, trait "distraction" was associated with overall visual attentional narrowing for emotional information. These rather unexpected findings indicate a multifaceted relationship between ER and visual attentional breadth for emotional information. Future research, especially in younger age groups, is needed to further elaborate on these findings.

In the current study, the role of attention was approached by the broaden-and-build theory in which it is stated that negative and positive emotions have different effects on the attentional scope (Fredrickson, 1998; Fredrickson and Joiner,

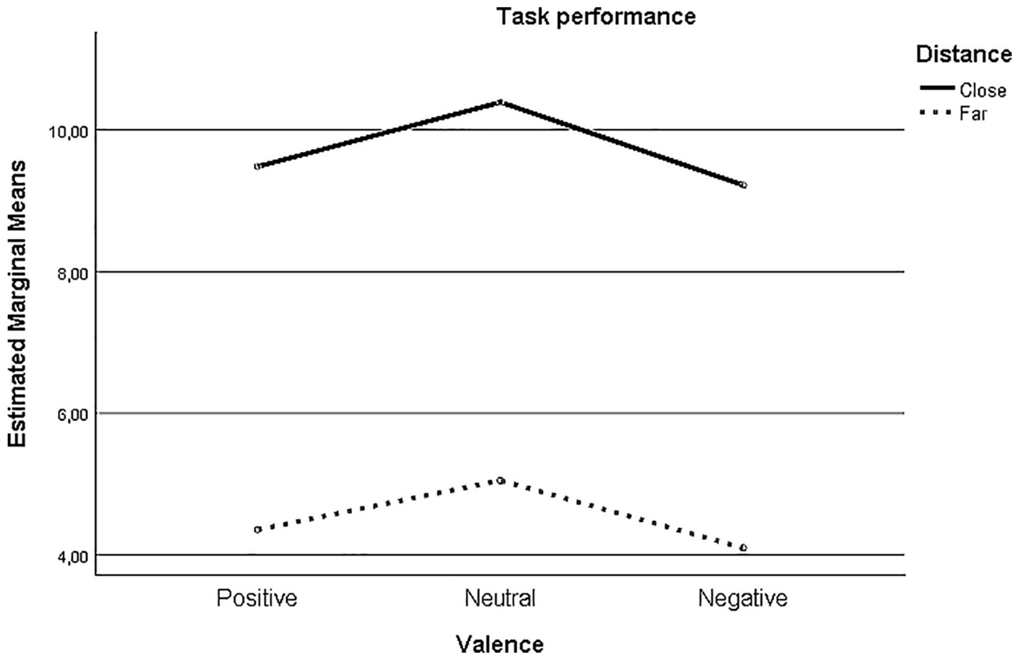

FIGURE 4 | Task performance. Note: estimated marginal means for the 3 Valence (Positive vs Neutral vs Negative) × 2 Distance (Far vs Close) interaction on the number of correctly localized target stimuli are displayed on the *y*-axis.

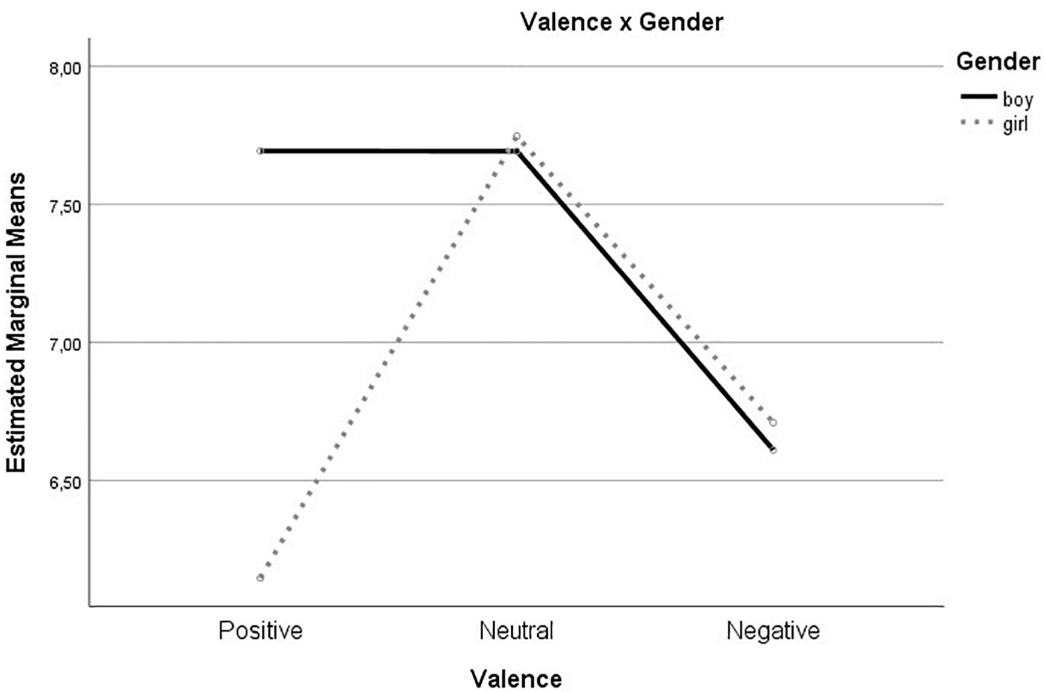

FIGURE 5 | Valence × Gender interaction. Note: estimated marginal means for the valence (Positive vs Neutral vs Negative) × Gender interaction on the number of correctly localized target stimuli are displayed on the *y*-axis.

2002; Fredrickson and Branigan, 2005). Positive emotions are theorized to be related to a broadened attentional scope and greater emotional wellbeing, whereas negative emotions are assumed to be associated with a narrowed attentional scope and risk for psychopathology (Fredrickson and Joiner, 2002). Given that attention is implied in the ER process and ER strategies have the possibility to influence both positive and negative affect, an association between ER and one's visual attentional scope for emotional information was expected (Aspinwall, 1998; Fredrickson and Joiner, 2002).

The first aim of the current study was to examine the relationship between the adaptive ER strategy "cognitive reappraisal" and visual attentional breadth for emotional information. In general, adaptive ER strategies are associated with the upregulation of positive and downregulation of negative affect. Based on the broaden-and-build theory of positive emotions, it was therefore expected that adaptive ER strategies such as cognitive reappraisal are associated with the broadening of the visual attention (Fredrickson, 2001; Fredrickson and Branigan, 2005). Again, no evidence was found for our hypothesis. In contrast to our study, recent studies clearly distinguish between two separate goals of cognitive reappraisal when researching this strategy (i.e., increasing positive versus decreasing negative affect). Although both goals lead to a reduction in negative affect, a significant difference in the increase of *positive* affect is reported by previous studies that examined this ER strategy form a *state* perspective (McRae et al., 2008, 2012). Unfortunately, the distinct goals of cognitive reappraisal were not assessed by the FEEL-KJ questionnaire, which measures "cognitive reappraisal" as a trait. Therefore, it is difficult to understand *if* and *how* youth used this specific strategy during the completion of the ABT lab task.

The second aim of the current study was to investigate the relationship between the maladaptive ER strategy "rumination" and visual attentional breadth for emotional information. Rumination is more likely to occur in a negative mood and can even reinforce the intensity of the emotions experienced (Nolen-Hoeksema et al., 2008). We hypothesized [based on (Fang et al., 2017)] that the use of "rumination" would be associated with a more narrowed attentional scope. In contrast to existing evidence in adults in which this association was confirmed (e.g., Grol et al., 2015), the current study found no significant relationship between rumination and attentional narrowing in youth for either positive, negative of neutral stimuli on the ABT. A potential explanation for our lack in findings could be that narrowed attention is associated with the use of trait rumination for self-related information (i.e., one's own feelings and problems) rather than other-related information (Whitmer and Gotlib, 2013; Grol et al., 2015). Although the current study uses trait rumination as an independent variable, no self-related stimuli were included during the experimental task. This could be an interesting avenue for future research.

Finally, we explored the relationship between distraction and visual attentional breadth for emotional information. Traditionally, distraction was categorized as an adaptive ER strategy in previous research (Sheppes et al., 2011) and also in the FEEL-KJ questionnaire this strategy was allocated to the adaptive ER subscale (Cracco et al., 2015). However, more recently researchers revealed a rather ambiguous relation between distraction and emotional well-being (Gratz and Tull, 2011; Craske and Barlow, 2014). Depending on whether it is used in a flexible way to redirect attention or exclusively to avoid negative affect, this strategy is considered adaptive or maladaptive, respectively (Gross, 1998; Van Dillen and Papies, 2015; Wolgast and Lundh, 2017). Because of this conflicting evidence, we hypothesized the general (trait) use of this strategy could be related to both attentional narrowing and attentional broadening for emotional information. Results revealed that distraction was related to lower accuracy rates for the central Emoji (i.e., positive, negative or neutral emoji) when the target stimuli were presented far versus close. This finding suggests that youth scoring high on trait distraction show a general attentional *narrowing* toward emotional information. Such pattern of overall narrowing can be interpreted as maladaptive in the long term since studies in adults found robust evidence for a relationship between attentional narrowing and negative affect, as well as decreased emotional well-being (Fredrickson and Joiner, 2002; Fredrickson and Branigan, 2005).

In addition to our main findings, preliminary analyses indicated that accuracy rates were higher for boys compared to girls when the central Emoji was positive. A possible explanation for this finding can be found in the fact that the adolescence is a critical developmental period regarding reactivity toward emotional information. Simultaneously, gender differences occur in the way this reactivity is handled (Hyde et al., 2008). More specifically, responses to positive affect decrease in adolescent girls compared to boys within the same age-group (Mezulis et al., 2004). Together with this shift, prevalence rates of adolescent depression become higher in girls compared to boys throughout adolescence.

So far, the only study in youth on the relationship of ER and visual attentional breadth for emotional information with the ABT (Van Beveren et al., 2020) found an association between the general use of adaptive ER and attentional broadening for neutral stimuli *in youth within the general population*, suggesting that this could be an indicator of resilience (Isen, 2000; Wadlinger and Isaacowitz, 2006; Sheppes and Gross, 2011). Although we could not replicate these findings, it extends this work to a sample of children and adolescents at risk for psychopathology that commonly demonstrate emotion dysregulation. The current study therefore provides a first and preliminary step in unraveling the different associations between ER strategy use and attentional breadth in at risk groups, throwing more/better light on underlying processes that contribute toward risk and psychopathology.

STRENGTHS AND LIMITATIONS

To our knowledge, the current study was the first to investigate the relationship between ER and visual attentional breadth for emotional information in youth at risk for developing psychopathology. Despite the innovativeness and specificity of the current study, several limitations should be acknowledged.

First, the lack of significant results for both rumination and cognitive reappraisal could be due to the small sample size (Kline, 2017). Power analyses were specifically determined for the large intervention study in which the current study was embedded and did not anticipate on the proportion of invalid data on the ABT. Out of 56 participating children and adolescents, 15 were excluded from the current study because of high error rates (Grol and Raedt, 2014). One possible explanation for the level of drop out could be due to the nature of the sample (i.e., youth with subclinical depressive symptoms), the duration of the task (± 40 min), task difficulty and the uncomfortable posture the children and adolescents were placed in. We therefore evaluated the performances on the ABT on several secondary parameters. A closer inspection of the fluctuations in affect before and after the task reveals an increase in feelings of frustration, anger, and boredom after completing the task, reflecting the amount of effort and perseverance the completion of the task required. Yet, previous studies with this task evaluated the ABT as valid, reliable, and doable and the analyses pertaining to task performance in the current study affirm that children and adolescents were able to conduct the task. Perhaps the stimuli used in the current study could have somewhat blurred our findings as the ABT (Bosmans et al., 2009; Grol et al., 2015) was originally developed and validated using faces as central stimuli.

Second, research shows that attentional processes for emotional information may be particularly disturbed in emotional disorders (Gasper and Clore, 2002; Yoon et al., 2015; Gu et al., 2017). For a better understanding of the role of visual attentional breadth for emotional information in the development of psychopathology in youth, the current study conducted in a sample of children and adolescents at risk for developing psychopathology can be considered a valuable pilot study. Future studies, also in clinical samples (i.e., depressed and anxious youth) are designated. Furthermore, as the present sample includes youth with obesity, the findings on the maladaptive role of distraction may be disorder specific. We cannot rule out that also other mechanisms (e.g., impaired self-regulation) that are typically observed in children and adolescents with obesity may have driven the results (Graziano et al., 2010; Golan and Bachner-Melman, 2011; Mehl et al., 2017).

Third, as the current study is cross-sectional, future longitudinal studies with multiple measuring points are crucial for determining the causality and directionality of the relationship between ER and visual attentional breadth for emotional information.

Importantly, we only examined whether self-reported use of three trait ER strategies could be related to visual attentional breadth for emotional information. We did not assess *whether* and *how* participants used these ER strategies while conducting the ABT. Therefore, no conclusions can be drawn about the role of attentional processes during the actual employment of these ER strategies.

Finally, the general use of specific ER strategies is measured through self-report. Using a multi-informant approach could counteract biases for social desirability as well as a lack of introspection in youth, by avoiding all outcome variables to depend on the same method and information source (Kline, 2017). The parent-report version of the FEEL-KJ questionnaire has recently been validated (Van Beveren et al., 2020) and could be of value to include in future studies.

FUTURE RESEARCH AND CLINICAL IMPLICATIONS

As mentioned above, no statement can be made about which ER strategies children and adolescents used during the ABT. Yet, such information would have been of great value for a more thorough understanding of the ER process. A growing body of research (e.g., Sheppes, 2014) suggests that, besides individual differences in the selection of certain strategies, there could be significant differences in effectiveness when using one of these strategies in a specific context. This line of reasoning could explain why certain ER strategies can be both adaptive and maladaptive depending on the context in which they are used. Whereas we already discussed the possibility that distraction can be both adaptive and maladaptive, as evidenced by various studies [e.g., (Wolgast and Lundh, 2017)], cognitive reappraisal has also been critically approached in recent research (Ford and Troy, 2019). More specifically, it is suggested that cognitive reappraisal can be successfully used and still be maladaptive when it is not in line with an individual's goals and motivation (e.g., reducing feelings important to your identity neglecting your authentic self) or when it is used in the "wrong" situation (e.g., reducing fear in an actual dangerous situation) (Tamir, 2016). Furthermore, cognitive reappraisal can be used unsuccessfully when emotions are too intense or too unfamiliar to generate an effective reappraisal (e.g., not enough reappraisal sources) (Sheppes and Meiran, 2007). Similar findings are found regarding rumination (Joormann et al., 2006). On the one hand, rumination can be considered maladaptive when it entails "brooding," which refers to making a passive comparison of the current situation with an unachieved standard (e.g., analyzing your own emotions and behavior focusing on the negative) (Treynor et al., 2003). On the other hand, rumination can be adaptive when it contains reflective pondering and thus an analysis of the situation and its causes in a neutral way (e.g., neutrally analysing why you are feeling a certain way in a certain situation) (Cristea et al., 2013). Although the current ER questionnaire (i.e., FEEL-KJ) is a reliable and valid instrument (see Cracco et al., 2015) to measure trait adaptive and maladaptive ER strategies (i.e., based on their factor structure and association with emotional well-being), future research could add experimental studies that start approaching ER strategies using a broader framework for conceptualizing, categorizing and assessing adaptive and maladaptive ER.

Furthermore, besides difficulties in categorizing ER strategies in to maladaptive and adaptive, more recent research suggests it is also too black-and-white to say that all negative emotions narrow and all positive emotions broaden the attentional scope. On the contrary, the impact of positive and negative affect on attentional scope could be more complex and/or flexible than first thought (Huntsinger, 2013). Studies show that not the valence of the affect but rather the motivational intensity (i.e., how important the affect is for an individual) will narrow/broaden the attentional scope (i.e., the higher the motivational intensity, the narrower the attentional scope) (Gable and Harmon-Jones, 2010;

Harmon-Jones et al., 2013). Future research could therefore include how personally relevant emotional information is during an attentional breadth task.

Next, some limitations regarding the experimental task (ABT) have already been mentioned. Future studies could also consider using *a global/local processing task* to measure attentional breadth since this task has some merits and has been evaluated across studies (Navon, 1977; Srinivasan and Hanif, 2010). During this type of task participants identify whether the local (narrow) or global (broad) element of a certain stimulus is more salient to them (Sung and Yih, 2016). Based on the limitations of the current study we suggest integrating the strengths of both tasks as well as to (1) include emotional stimuli, (2) implement a standardized mood induction to evoke the use of state ER, and (3) use a survey that measures both the general (trait) and the actual (state) use of ER strategies (during the task).

From a clinical perspective, upcoming research on training ER provides evidence that learning to use specific adaptive ER strategies improves the ability to influence affect (i.e., increasing positive and decreasing negative emotions) (Southam-Gerow, 2013; Berking and Lukas, 2015; Wante et al., 2018). In the current study we hypothesized that ER and visual attentional breadth for emotional information to be associated (Fredrickson, 2001). If any significant relations would have been found between cognitive reappraisal and attention, this would have indicated that training adaptive ER strategies could also broaden the attentional scope of the individual. As both adaptive ER and attentional broadening are considered protective factors, these two mechanisms could potentially reinforce each other and eventually improve mental health (Fredrickson and Joiner, 2002; Rowe et al., 2007; Aldao and Nolen-Hoeksema, 2010, 2012; Gu et al., 2017). However, based on the results of the current study, no assumptions can be made.

FINAL CONCLUSION

To our knowledge, the current study is the first to investigate the relationship between specific emotion regulation strategies and visual attentional breadth in youth at risk for developing psychopathology. Based on the broaden-and-build theory of positive emotions, it was hypothesized that adaptive ER facilitates positive affect and is therefore related to attentional broadening for emotional information. In contrast, given that maladaptive ER maintains negative affect, it was hypothesized that there would be an association with attentional narrowing (Fredrickson, 1998, 2001; Fredrickson and Branigan, 2005). Three emotion regulation strategies that rely on attentional deployment (i.e., cognitive reappraisal, rumination and distraction) were selected and a validated visual attentional breadth task was used. No evidence was found for the proposed associations between both the adaptive ER strategy cognitive reappraisal and the maladaptive ER strategy rumination and visual attentional breadth for emotional information in at risk children and adolescents. However, a remarkable association between the use of distraction and overall visual attentional narrowing toward negative emotional information was found. These results emphasize the ambiguous character of distraction as an ER strategy (e.g., in some contexts it can be considered adaptive but in others maladaptive) and help to further characterize it by suggesting this strategy is predominantly used maladaptively in children and adolescents at risk for psychopathology. These findings indicate the complex relationship between ER and visual attentional breadth and highlight the need for further research in both clinical and larger (selected) samples of youth.

DATA AVAILABILITY STATEMENT

The raw data supporting the conclusions of this article will be made available by the authors upon reasonable request.

ETHICS STATEMENT

The studies involving human participants were reviewed and approved by Committee of Medical Ethics UZ Gent. Written informed consent to participate in this study was provided by the participants' legal guardian/next of kin.

AUTHOR CONTRIBUTIONS

EB collected all the data. EB and M-LV analyzed the data. All authors were involved in writing the manuscript and had final approval of the submitted and published versions.

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