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REVIEW

Effects of selective attention on the C1 ERP component: A systematic review and meta-analysis

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Abstract

The C1 event-related potential (ERP) captures the earliest stage of feedforward processing in the primary visual cortex (V1). An ongoing debate is whether top-down selective attention can modulate the C1. One side of the debate pointed out that null findings appear to outnumber positive findings; thus, selective attention does not seem to influence the C1. However, this suggestion is not based on a valid approach to summarizing evidence across studies. Therefore, we conducted a systematic review and meta-analysis investigating the effects of selective attention on the C1, involving 47 experiments and 794 subjects in total. Despite heterogeneity across studies, results suggested that attention has a moderate effect on the C1 (Cohen's $d_z = 0.33$, $p < .0001$); that is, C1 amplitude is larger for visual stimuli that are attended than unattended. These results suggest that C1 is affected by top-down selective attention.

KEYWORDS

attentional load, C1, ERP, primary visual cortex, selective attention, spatial attention

1 | INTRODUCTION

The C1 is the earliest visual evoked potential derived from scalp EEG in humans (Clark et al., 1994; Di Russo et al., 2002, 2003; Foxe & Simpson, 2002; Jeffreys & Axford, 1972). Characterized by an early onset at about 40–50 ms and a peak latency at about 60–90 ms after stimulus onset (Rauss et al., 2011), it originates in the primary visual cortex (V1, Brodmann area 17), which encompasses the calcarine fissure and its surround on the medial surface of the occipital lobe, and thus corresponds to the first wave of feedforward activation in the occipital cortex along the retino-geniculate-striate pathway (Clark et al., 1994). It is a retinotopic component (Engel et al., 1997; Gilbert & Li, 2013), showing systematic variation in amplitude, polarity, and topography depending on the position of the stimulus in the visual field. As shown in Figure 1, the C1 reverses its polarity between upper and lower visual field stimulation. For lower visual field stimuli, the C1 is a positive-going phasic wave over occipito-parietal electrodes along the midline, while for upper visual field stimuli, it is characterized by a negative wave with similar latency over the same electrode locations. This property is consistent with the activation of pyramidal neurons situated in the upper bank of the calcarine cortex for lower visual field stimulations and in the lower bank of the calcarine cortex for upper visual field stimulations (Clark et al., 1994; Jeffreys & Axford, 1972; Kelly, Schroeder, & Lalor, 2013; Kelly, Vanegas, et al., 2013). Note that this polarity inversion alone is not sufficient to identify V1 generators, as polarity inversion has also been shown for later, extrastriate components, such as the P1

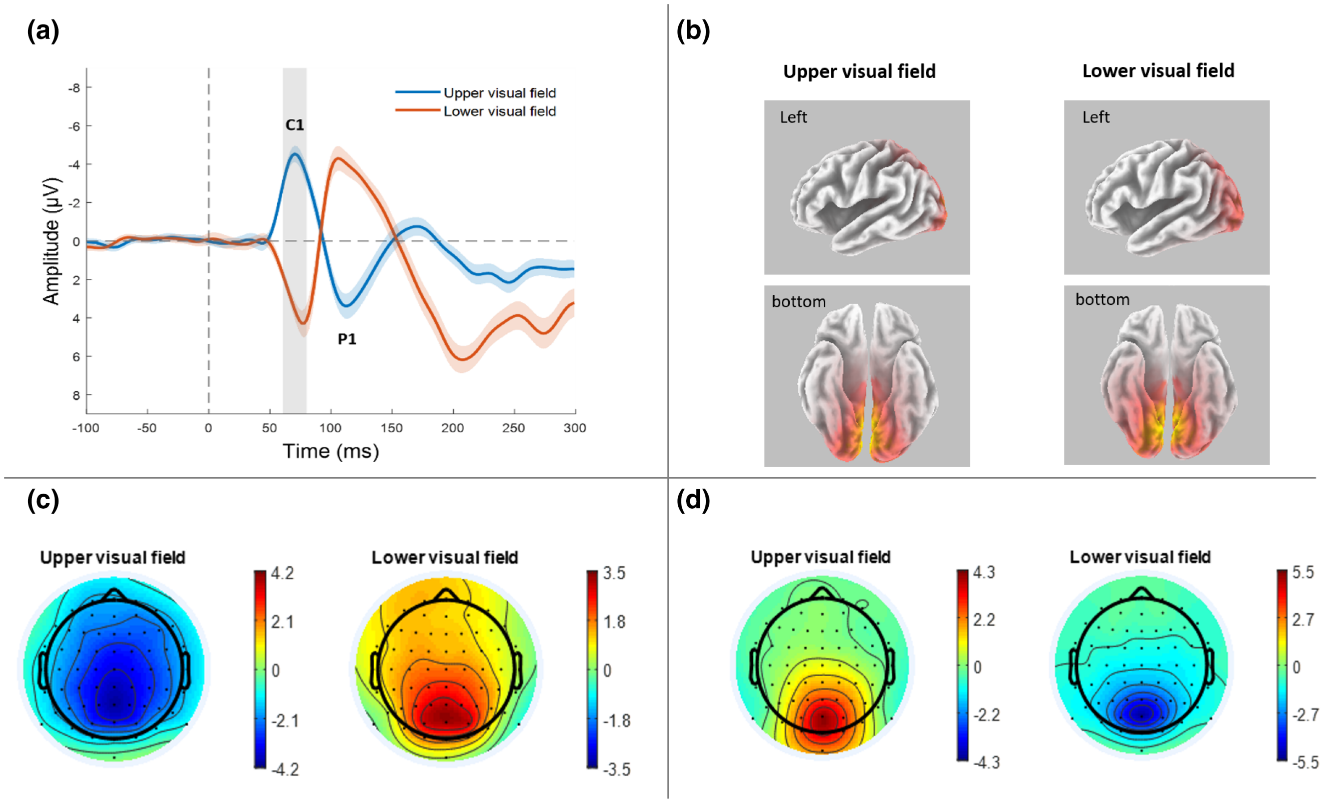

FIGURE 1 Results from an EEG study (Qin & Pourtois, in preparation) in which 33 adult healthy participants kept fixation on a central location where a visual detection task had to be performed, while task-irrelevant peripheral stimuli (i.e., textures composed of line bars) were shown either in the upper or lower visual field. (a) Grand average (±1 SEM) visual ERPs (electrodes CPz, Pz and POz pooled together) for peripheral stimuli revealed a clear polarity reversal peaking at around 70 ms (C1), followed by a second one (P1), peaking at 110 ms after stimulus onset. (b) the corresponding distributed inverse solution for the C1 at 70 ms after stimulus onset for upper visual field presentation and 76 ms for lower visual field presentation. It suggests activation of the occipital pole, extending to the medial side of the occipital lobe, including V1. These solutions were obtained using sLORETA. (c) the corresponding topographical voltage maps for the C1, for each hemifield separately. (d) Topographical voltage maps of the subsequent P1, indicating a polarity reversal in the opposite direction for this extrastriate component.

(Ales et al., 2010, 2013). However, as laid out by Kelly, Schroeder, and Lalor (2013); Kelly, Vanegas, et al. (2013), polarity inversion for extrastriate components is opposite to that of the C1, and the empirically observed shifts in C1 topography as a function of stimulus location are uniquely consistent with neural generators in V1 (see also Clark et al., 1994; Di Russo et al., 2002; Jeffreys & Axford, 1972; Kelly et al., 2008). Taken together, the C1 component is considered a reliable electrophysiological correlate of the earliest cortical activation in V1 following stimulus onset in humans (Foxe et al., 2008; Foxe & Simpson, 2002; Gomez Gonzalez et al., 1994).

Previous ERP research explored whether the C1 component is influenced by selective attention. Selective attention allows selecting relevant information in the environment while filtering out or attenuating irrelevant stimuli (Desimone & Duncan, 1995; Posner, 1980; Posner & Petersen, 1990). As a result, a particular location (spatial attention), feature (feature-based), or object (object-based attention) can be selected for in-depth processing (Carrasco, 2011; Olson, 2001; Scholl, 2001). Early investigations used the Posner cueing task (Posner, 1980). Results suggested that the C1 component was unaffected by spatial attention (e.g., Anllo-Vento & Hillyard, 1996; Clark & Hillyard, 1996). In those studies, participants processed a visual stimulus shown at a specific location in the visual field, which was either cued or not beforehand. This led to valid and invalid trials, corresponding to *more* and *less attention* to the C1-eliciting stimulus, respectively. Results showed that the C1 elicited by this stimulus was comparable for the two conditions, while the subsequent P1 component was larger for valid than invalid trials. Moreover, the complementing source-localization results showed that V1 could be influenced by spatial attention, yet via delayed feedback effects arising from the extrastriate cortex (where the P1 is mainly generated) onto the striate cortex (see Di Russo et al., 2003; Martínez et al., 1999).

These ERP results were compatible with single-cell recordings in non-human primates (Ito & Gilbert, 1999; Roelfsema et al., 1998), as well as fMRI studies in human subjects (Gandhi et al., 1999; Hopf et al., 2004; Somers et al., 1999), showing that spatial attention can modify sensory responses in V1 via reentrant processing. Hence, the picture emerged that spatial attention can create early gain control effects in V1, but that these do not result from an initial feedforward effect but rather from a delayed feedback effect (Hillyard & Anllo-Vento, 1998).

More recently, this view has been challenged by a growing number of ERP studies suggesting that the C1 is influenced by selective attention (Fu et al., 2008, 2009, 2010a, 2010b). Using a spatial cueing task, Kelly et al. (2008) instructed subjects to covertly attend to a certain location and to detect a target at the cued location only. Results showed that the C1 was larger for attended compared to unattended locations. This early spatial attention effect was most clearly revealed when subject-specific characteristics regarding the topography of the C1 were taken into account to score this early visual ERP component. Besides spatial attention, several studies reported that object-based and feature-based attention influence the C1 too (Khoe et al., 2005; Proverbio et al., 2010; Zani & Proverbio, 2005). Proverbio et al. (2010) instructed participants to concurrently pay attention to a specific location in the visual field and a particular spatial frequency. Results showed that the C1 was larger for frequency-relevant (*more attention*) than frequency-irrelevant (*less attention*) stimuli, suggesting a modulatory effect of feature-based attention on early visual processing. Moreover, several studies based on the load theory of selective attention (Lavie, 1995, 2005) suggested that increasing load of the task at fixation reduces the amplitude of the C1 to a peripheral visual stimulus (Rauss et al., 2009; Rossi & Pourtois, 2012, 2014, 2017). In this framework, the effects of selective attention are explored by changing the amount of attentional resources needed to perform the task at hand, which in turn influences distractor processing. More specifically, it is assumed that the peripheral distractors get more attention if the demands associated with the central task are low compared to high. In line with this reasoning, Rauss et al. (2009) manipulated attentional load at fixation, which was either low or high in different blocks, while the same irrelevant distractors were shown in the periphery at an unpredictable time point. Using this paradigm, the authors were able to compare the C1 elicited by the distractors between low load (*more attention*) and high load (*less attention*). Results showed that the C1 was larger for low than high load, a result compatible with earlier fMRI data showing a decrease of V1 activity evoked by peripheral distractors (as well as activity in the ventral extrastriate visual cortex) by high attentional load allocated elsewhere (Schwartz et al., 2005). Taken together, recent ERP studies have documented modulations of the C1 by selective attention.

In contrast, other studies have not found statistically significant effects, and studies reporting null findings seem at first glance to outnumber those that have reported significant C1 effects. These observations have led some authors to argue that overall, the empirical evidence favors a *majority view* in line with the classic notion that the C1 is impermeable to selective attention (see Baumgartner et al., 2018). This position is in sharp contrast with the *minority view* according to which selective attention can influence the C1 component under specific circumstances (see Slotnick, 2018 for a recent review of this debate). However, this division is mostly based on a narrative review of the literature, as opposed to a quantitative integration of the available evidence in form of a systematic review and meta-analysis (Baumgartner et al., 2018; Rauss et al., 2011). Moreover, the proponents of the *majority view* (see Baumgartner et al., 2018) mostly base their conclusions on counting the number of (nonsignificant) *p*-values, which is not a mathematically rigorous way of synthesizing the results from different studies. Meta-analysis, however, with larger sample size and higher statistical power, allows us to combine the effects and evaluate the statistical significance of the summary effect (Borenstein et al., 2009). Furthermore, null hypothesis significance testing per se cannot be used to prove the absence of an effect, particularly so if there is no a priori power analysis. Thus, a nonsignificant effect of attention on the C1 does not provide conclusive statistical evidence in favor of the null hypothesis (Dienes, 2008; Halsey et al., 2015; Makin & Orban de Xivry, 2019). Furthermore, although a *p*-value may be indicative of whether an effect exists, it cannot reveal the size of this effect. These issues can be reduced by performing a meta-analysis (Cumming, 2013a, 2013b; Lipsey & Wilson, 2001; Borenstein et al., 2009), which combines effect sizes across existing studies to determine whether there is an overall effect. Moreover, a meta-analysis can quantify how much published effect sizes differ from the variability that would be expected by chance alone. More specifically, a test of heterogeneity can be performed to examine whether the studies included in the meta-analysis measure a common underlying effect (Higgins et al., 2003; Higgins & Thompson, 2002). If heterogeneity is found, then the potential cause can be explored with a moderator analysis (e.g., Wiens et al., 2016).

To decide between the *majority* and *minority* views, and to overcome the limitations mentioned above, we performed a systematic review and meta-analysis of published C1 studies. The goal of our study was to assess whether the evidence available in the literature favors the notion that the C1 is modulated by selective attention (*minority view*), or not (*majority view*). Since we performed an exhaustive evaluation of the existing evidence, we were able to consider different types of selective attention (i.e., *spatial attention*, *attentional load*, and *other*). Hence, we could also explore, through a moderator analysis, whether modulatory effects on the C1 may be specific to particular classes of selective attention, which is an important question at the theoretical level (Nienborg & Cumming, 2014; Nobre et al., 2014; Petersen & Posner, 2012; Posner & Petersen, 1990).

2 | METHOD

2.1 | Literature search strategy

The literature search (run on March 3, 2021) was conducted in PubMed and in Web of Science using the following terms: (*C1* OR *VEP* OR *polarity reversal* OR *first sweep* OR *early wave* OR *earliest component* OR first deflection OR *initial evoked response*) AND (*V1* OR *Brodmann area 17* OR *striate cortex* OR *laminar* OR *calcarine* OR *upper bank* OR *lower bank* OR *feedforward* OR *feedback* OR *early visual processing* OR *primary visual cortex* OR *retinotopic* OR *cruciform* OR *occipital*) AND (*perception* OR *attention* OR *plasticity* OR *learning* OR *perceptual* OR *cognition* OR *cognitive*). This search resulted in 374 hits in Web of Science and 241 hits in PubMed. We also performed a forward search based on the few review articles that were retrieved (Rauss et al., 2011; Slotnick, 2018). During the process of effect size calculation (see below), we also contacted several authors of C1 studies and asked them to share newly published results for this ERP component with us. After removing duplicates, 444 articles were retained.

2.2 | Selection criteria

1. The study had to be reported in English and published in a peer-reviewed journal.
2. Participants had to be healthy adults.
3. Since we aimed to systematically review EEG studies assessing the attentional effect on the C1, scalp EEG had to be used as the main investigation technique, and the amplitude of the C1 had to be compared between at least two conditions. Non-EEG studies (i.e., MEG) were not included.
4. The meta-analysis considered only studies examining attention in the visual modality. Studies examining modulatory effects of auditory attention or intermodal attention (e.g., Karns & Knight, 2009) were excluded.
5. To review studies which directly manipulated selective attention, C1 studies were excluded if they focused on perceptual learning (e.g., Bao et al., 2010; Pourtois et al., 2008) or specific emotional or motivational processes (e.g., Acunzo et al., 2019; Pourtois et al., 2004; Rossi et al., 2017; Vanlessen et al., 2013, 2014), because previous studies have shown that emotion and motivation may operate via mechanisms distinct from those of attention (Baldassi & Simoncini, 2011; Chelazzi et al., 2013; Pourtois et al., 2013). Likewise, a more recent ERP study exploring the effects of prediction (or predictability) on early visual processing (Jabar et al., 2017) was also excluded.
6. When the same data were reported in multiple publications (Martínez et al., 1999, 2001), only the first published study was included.
7. One ERP study (Fu et al., 2008) was excluded because the C1 results were mostly explained by overlapping and confounding P1 effects.

2.3 | Study selection

Two authors (NQ and GP) screened these 444 articles, using the abovementioned inclusion criteria and following the PRISMA checklist (see Figure 2; Moher et al., 2009). Each coder inspected the title and abstract of each article. This led to a selection of 81 articles. When there was disagreement between raters, the article was marked for full-text reading in a later phase. After excluding 36 articles, 45 articles were eventually retained. Given that some of them included several experiments, 51 experiments could finally be analyzed. For one of them (Zani & Proverbio, 2006) only the mean C1 amplitude and the *p*-value for the interaction term (i.e., location relevance × frequency relevance) were reported, making it impossible to calculate the effect size. Hence, 50 experiments extracted from 44 articles were eventually identified as eligible for the meta-analysis. The total number of participants was 846 (mean number of subjects per study = 17; *SD* = 6.18; minimum = 8; maximum = 38).

2.4 | Data extraction

Full information on data extraction and coding is provided and can be consulted online (see MA_Data.xlxs at https://osf.io/rydc5/). Even though the included experiments varied in how attention was manipulated (e.g., attended vs. unattended; valid vs. invalid cue; cued vs. un-cued location; low vs. high load; relevant vs. irrelevant feature), we combined them by always coding one condition as *more attention* on the C1-eliciting stimulus (*CondMoreAtt*) and the other as *less attention* on the C1-eliciting stimulus (*CondLessAtt*). In this way, for each

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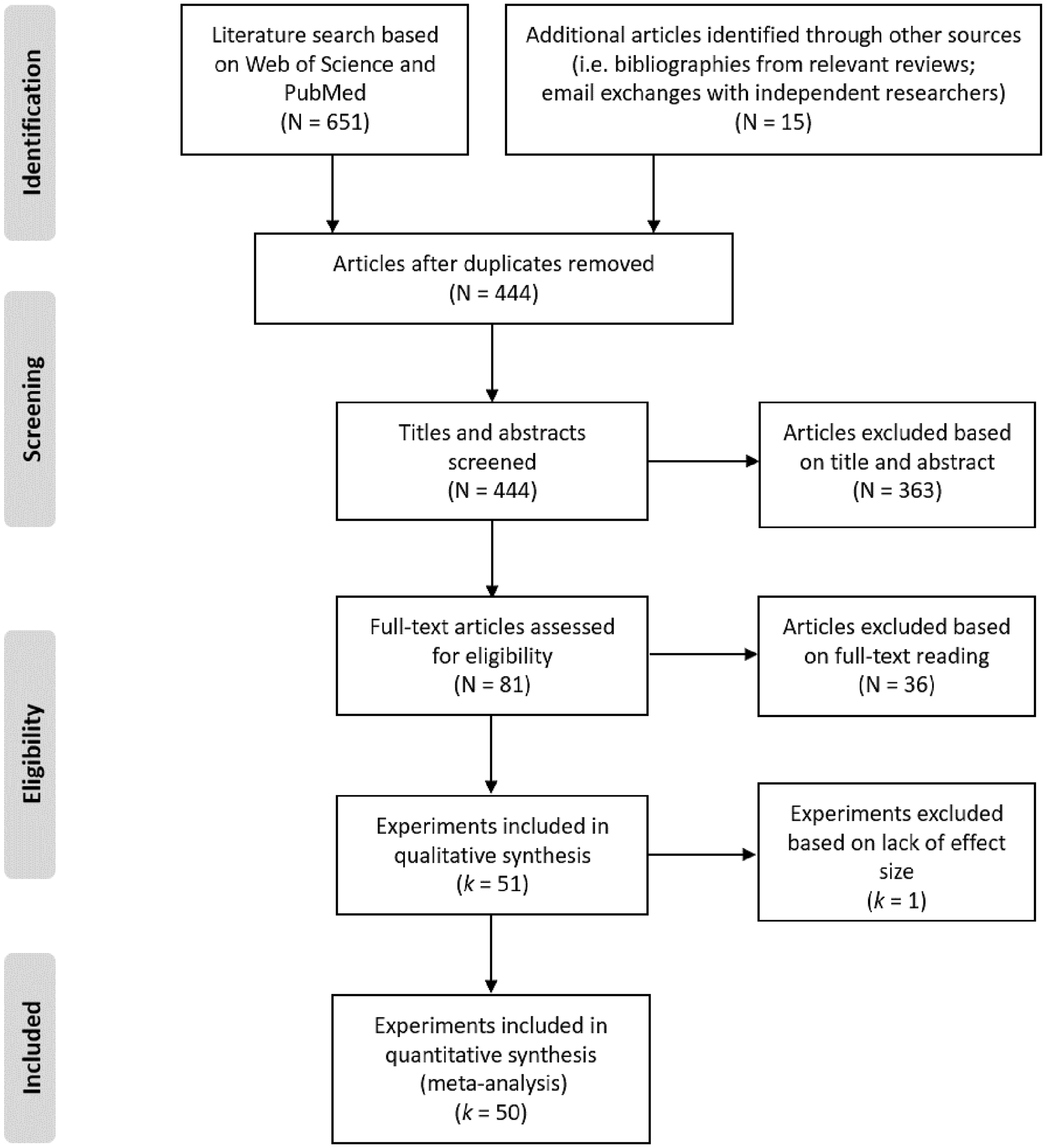

FIGURE 2 PRISMA flow diagram showing the selection steps carried out to identify eligible C1 ERP experiments for the meta-analysis. Here *N* refers to the number of articles and *k* refers to the number of experiments included in each step.

experiment, we could directly compare the more to the less attention condition (see MA_Data.xlxs in the OSF project).

To make sure that effect sizes were calculated for independent samples of participants (see Bar-Haim et al., 2007; Moran et al., 2017; Pool et al., 2016), only one effect size was calculated per experiment according to the following criteria:

1. If more than two levels of attention were used (e.g., low, medium, and high load), the two extreme levels were compared to each other (i.e., low vs. high load in this example).
2. If an experiment tested more than one type of attention for the same participants, we selected the one that was tested and mentioned explicitly in the article (title and/or abstract) as the main factor. If all attention types were tested without different weights, we selected the one for which fewer experiments were available in the meta-analysis, so as to improve the balance between categories in the moderator analysis. For instance, for an experiment manipulating both spatial attention and

feature-based attention without any hierarchy suggested between them (Proverbio et al., 2007), we focused on the effect of feature-based attention on the C1.
3. When multiple time windows were used and reported to score the C1, only the earliest one was selected. For instance, Alilović et al. (2019) reported both early C1 and peak C1, and we only selected the former.
4. When the C1 was measured at several electrodes, the effect size was calculated for the average of all electrodes.

2.5 | Data analysis

As recommended (Lakens et al., 2016), all analytical steps and the results of the meta-analysis are shared openly (see the OSF project at https://osf.io/rydc5/). Analyses were conducted with RStudio (Rstudio Team, 2020) in *R* (R Core Team, 2013) together with selected packages (Auguie & Antonov, 2017; Viechtbauer, 2010; Wickham et al., 2019; Zhu, 2019). The main file is saved as a Rmarkdown (Allaire et al., 2018) file and is called C1meta.html, which can be opened with any browser.

Our analysis is based on the assumption that if attention modulates the C1 component, its amplitude should be larger for *more attention* to the C1-eliciting stimulus compared to *less attention* to the C1-eliciting stimulus. This is in accordance with classic gain-control models of attention (Hillyard et al., 1995, Hillyard & Anllo-Vento, 1998; Luck et al., 1997). Consequently, we used a measure of effect size that captures the difference of *more attention* minus *less attention*. If attention has an effect on the C1, this effect size should yield a positive value.[1] This is complicated by the fact that the C1 can be positive or negative depending on the study design. For example, if a study used upper visual field stimuli, a negative mean difference (e.g., −0.8 minus −0.2 = −0.6) would support the idea that C1 amplitudes were larger during more attention than less attention. In contrast, if the study used lower visual field stimuli, then a positive mean difference (e.g., 0.8 minus 0.2 = 0.6) would support the idea that C1 amplitudes were larger during more attention than less attention. We firstly defined *PredSignForMoreMinusLess* to capture if the *more attention* minus *less attention* difference was predicted a priori to be negative or positive. However, for some studies, the effect size had to be calculated from two-tailed statistical values (*F*, *t*, or *p*), which are uninformative as to the actual direction of the effect. Therefore, we additionally defined *ActualSignForMoreMinusLess*, to capture if the actual difference retrieved from each experiment was negative, positive, or unclear. If unclear, we explicitly defined it as either positive or negative (yielding liberal or conservative results, as explained below). For the meta-analysis, both variables are needed to calculate the effect size for each experiment correctly (e.g., also for the rare cases where the sign of the C1 was not compatible with the upper/negative C1 vs. lower/positive C1 a priori division). More details about effect size calculation per experiment and across experiments are available in the Excel file MA_Data.xlxs in the corresponding OSF project.

For each of the 50 experiments retained, the effect size of attention on the C1 was expressed in terms of Cohen's d_z (Cohen, 1988). Here, Cohen's d_z refers to the mean difference of more attention to the C1-eliciting stimulus minus less attention to the C1-eliciting stimulus in experiments in which each subject participated in both attention conditions; thus, the level of analysis is intra-individual rather than inter-individual (Lakens, 2013). This approach was used because all experiments except one had a within-subject design (Proverbio & Adorni, 2009; see below for the calculation of effect size for that experiment).

As suggested by Lakens (2013) for intra-individual analyses, we took the correlation between the scores of the two attention conditions (i.e., *more* vs. *less attention*) into account. Cohen's d_z was computed from the mean amplitudes and the standard deviation of the difference scores between two experimental conditions ($k = 10$) (Cohen, 1988; Lakens, 2013). Here, k refers to the experiments included in this step. $M1$ and $M2$ refer to the means for the more attention condition and the less attention condition, respectively. SD_{diff} is the standard deviation of the difference for the two conditions. $V(d_z)$ refers to the variance of Cohen's d_z and n is the sample size:

$$\text{Cohen's } d_z = \frac{M1 - M2}{SD_{diff}}$$

$$V(d_z) = \frac{1}{n} + \frac{d_z^2}{2n}$$

When it was not possible to retrieve the standard deviation of the difference scores, we calculated it from the standard deviation of both groups and the correlation (r) between the scores ($k = 5$) where SD_1 and SD_2 are the standard deviation of the scores for the two conditions, respectively, and r is the Pearson's correlation coefficient between conditions (Cohen, 1988):

$$SD_{diff} = \sqrt{SD_1^2 + SD_2^2 - 2 \times r \times SD_1 \times SD_2}$$

[1]In two studies, the experimental conditions could not easily be labeled as *more* versus *less* attention. In Noesselt et al. (2002), attention was directed to either left or right side. In Proverbio and Adorni (2009), participants were instructed to attend either to orthographic or lexical features. For these studies, the condition showing a larger C1 was arbitrarily coded as *more attention*.

If the correlation was not retrievable but the t-value of a paired t test was reported ($k = 1$), Cohen's d_z was calculated as follows (Lakens, 2013; Rosenthal, 1991):

$$\text{Cohen's } d_z = \frac{t}{\sqrt{n}}$$

In case only the p-value of the t test was reported, we retrieved t using the inverse t distribution (Moran et al., 2017) and then obtain Cohen's d_z as above ($k = 1$). When only F values of ANOVAs were reported ($k = 19$), we obtained the corresponding t values according to:

$$t = \sqrt{F}$$

provided that the underlying analysis was a one-way, two-level ANOVA (Brožek & Howard, 1950; Rosenthal, 1991).

During this process, we contacted the corresponding authors (and/or senior authors) of some of the included studies to ask for missing values or raw data, and many of them responded positively to our request. However, for cases in which data were unavailable (marked in the excel file MA_Data.xlxs in the OSF project to ease their detection), we computed effect sizes by assuming a hypothetical correlation or p-value, based on either a liberal or a conservative mode of calculation, as follows:

1. When mean amplitudes and standard deviations (or standard errors) for the two conditions were reported but not the correlation between the scores ($k = 6$), the effect size was calculated by assuming a correlation of .75 (Dunlap et al., 1996). For the more conservative calculation, the correlation was assumed to be .5.

2. When the results were nonsignificant and reported incompletely (e.g., $p > .05$ or n.s.), we obviously could not compute the exact effect size ($k = 7$). However, removing nonsignificant effects can be problematic, as it could inflate the global effect size. To overcome this problem, we computed the effect size of these experiments by assuming $p = .5$ (Moran et al., 2017). Accordingly, these nonsignificant results could be included in the meta-analysis and contribute to the pooled effect size (summary statistic). The p-value was assumed to be .99 with a conservative calculation. In some cases, the actual direction of the effect could not be retrieved from the original article ($k = 10$). For the liberal calculation, we computed the effect size according to the coding scheme outlined here above (i.e., higher C1 assumed for the *more* attention condition), while for the conservative calculation, the opposite direction for the effect of attention was used (i.e., lower C1 assumed for the *more* attention condition).

Given that the liberal way is more widely used in the existing literature, we focus on this procedure in the main text but mention it where relevant. However, details about the conservative calculation and the results can be found in the file C1meta_conservative.html in the OSF project.

As mentioned above, only one experiment ($k = 1$) was based on a between-subjects design, with two groups of participants of equal size (Proverbio & Adorni, 2009). For that experiment, Cohen's d_z and its variance was calculated using the following formulas (Cohen, 1988):

$$\text{Cohen's } d_z = \frac{M_1 - M_2}{\sqrt{\frac{SD_1^2 + SD_2^2}{2}}}$$

$$\mathrm{V}\left(d_z\right) = \frac{2n}{n^2} + \frac{d_z^2}{4n}$$

In an initial analysis including all 50 experiments, potential outliers were identified using the Baujat plot (Baujat et al., 2002), which allows visualizing the contribution of each experiment to the Q-test statistic for heterogeneity and describing the influence of each experiment on the overall effect. Next, the standardized residual of each experiment was calculated. As suggested by Lipsey and Wilson (2001), an experiment whose standardized residual z score of the effect size exceeds three can be regarded as an outlier.

After excluding outliers, a random-effect model was fitted to estimate the pooled effect, which is a weighted average of all individual effect sizes, using the *metafor* package (Viechtbauer, 2010). Experiments with larger sample sizes have a heavier weight than those with smaller ones. The unstandardized effect size, namely the mean C1 amplitude difference was also calculated. Moreover, heterogeneity was assessed with the Q test and the I^2 statistic. The Q test (Borenstein et al., 2009) examines whether there is statistically significant heterogeneity between experiments, while the I^2 statistic further quantifies heterogeneity as the percentage of the total variance due to the true between-study difference (Higgins et al., 2003). Separate meta-analyses were also conducted for upper and lower visual field presentations.

A moderator analysis assessing the type of attention was subsequently conducted by fitting a mixed-effects model. Publication bias refers to the fact that studies with significant results are usually more likely to be published (Borenstein et al., 2009; Rosenthal, 1979, 1991). Publication bias was examined with a funnel plot and the trim-and-fill procedure. This involves plotting the effect size of each experiment along the X-axis and the standard error for that effect size on the Y-axis. The standard error indicates the precision of the effect size as an estimate of the population parameter. This precision increases with increasing sample size, such that results from small studies will scatter widely at the bottom of the plot, with the spread narrowing for the larger studies plotted toward the top. If the resulting plot is a symmetrical inverted funnel, this indicates no bias. If publication bias is present, then the plot will be asymmetrical (Sterne & Egger, 2001). Egger's test can be used to quantify the extent of publication bias (Egger et al., 1997), by regressing the standardized effect sizes on their precision (i.e., the inverse of the standard error). If a publication bias exists, then the regression intercept is expected to deviate significantly from zero. If there is a publication bias, then funnel plots usually show more studies with small samples and large effect sizes falling toward the right, where significant effects are displayed, and fewer toward the left. To correct for this bias, a trim-and-fill procedure (Duval, 2005; Duval & Tweedie, 2000a, 2000b) can be applied that imputes the 'missing' studies on the left and adds their (negative) effect sizes to the analysis. More specifically, studies with large effect sizes and small sample sizes are removed until symmetry is achieved. Next, the trimmed studies are added back and virtual mirror studies across the mean are also inputted to recalculate a new overall effect size.

A cumulative meta-analysis was also conducted where we could order all studies chronologically, and eventually assess how the global effect size evolved between the first published experiment in 1994 and the last one in 2021.

3 | RESULTS

3.1 | Overall effect

Before computing the pooled effect size, the Baujat plot was used to detect possible outliers (Figure 3). Based on the standardized residual of each experiment, Zani and Proverbio 2009 ($Z = 4.17$), Zani and Proverbio 2012 ($Z = 3.09$), and Zhang et al., 2012 ($Z = 3.06$) were identified as outliers and removed from further analyses.

Figure 4 shows the Forest plot of the remaining 47 experiments involving a total of 794 subjects. The meta-analysis revealed an effect of attention on the C1 ($d_z = 0.33$, 95% CI: [0.23, 0.43], $p < .0001$). Notably, results of the main analysis suggested that the between-study heterogeneity ($I^2 = 42.60\%$, $Q = 81.28$, $p = .001$) was moderate (Higgins et al., 2003). Also, when we used a conservative approach to compute effect sizes (see Method), the meta-analysis continued to suggest an effect of attention on the C1 ($d_z = 0.21$, 95% CI [0.10, 0.33], $p = .0004$, $k = 48$).

Because most studies (23 out of 50 experiments) did not report mean amplitudes at all, we had to conduct the meta-analysis on standardized effect sizes. To get some idea of the possible size of the unstandardized effect size (in μV), we used the estimated *SD* of the difference scores to convert d_z to an effect size in μV; note however, that the estimate *SD* is based on only 10 studies (see C1meta_liberal.html in the OSF project for details). Results suggested that the values for the unstandardized effect size are similar to those for the standardized effect: The mean

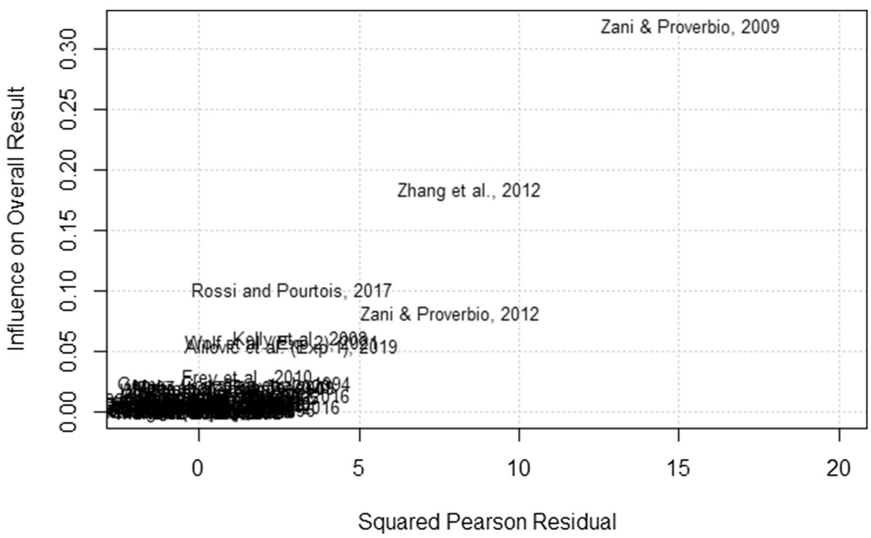

FIGURE 3 The Baujat plot illustrates the contribution of each experiment to the pooled effect size. The *X*-axis corresponds to the squared Pearson residual of an experiment, which is the individual contribution to the test for residual heterogeneity. The *Y*-axis corresponds to the standardized squared difference between the predicted/fitted value for the experiment with and without the experiment included in the model fitting. Hence, an experiment that has the greatest variation from the overall effect size estimate and the most substantial contribution to the estimate is located in the top right part of the plot. Three experiments appear to contribute substantially to the observed heterogeneity and were identified as outliers.

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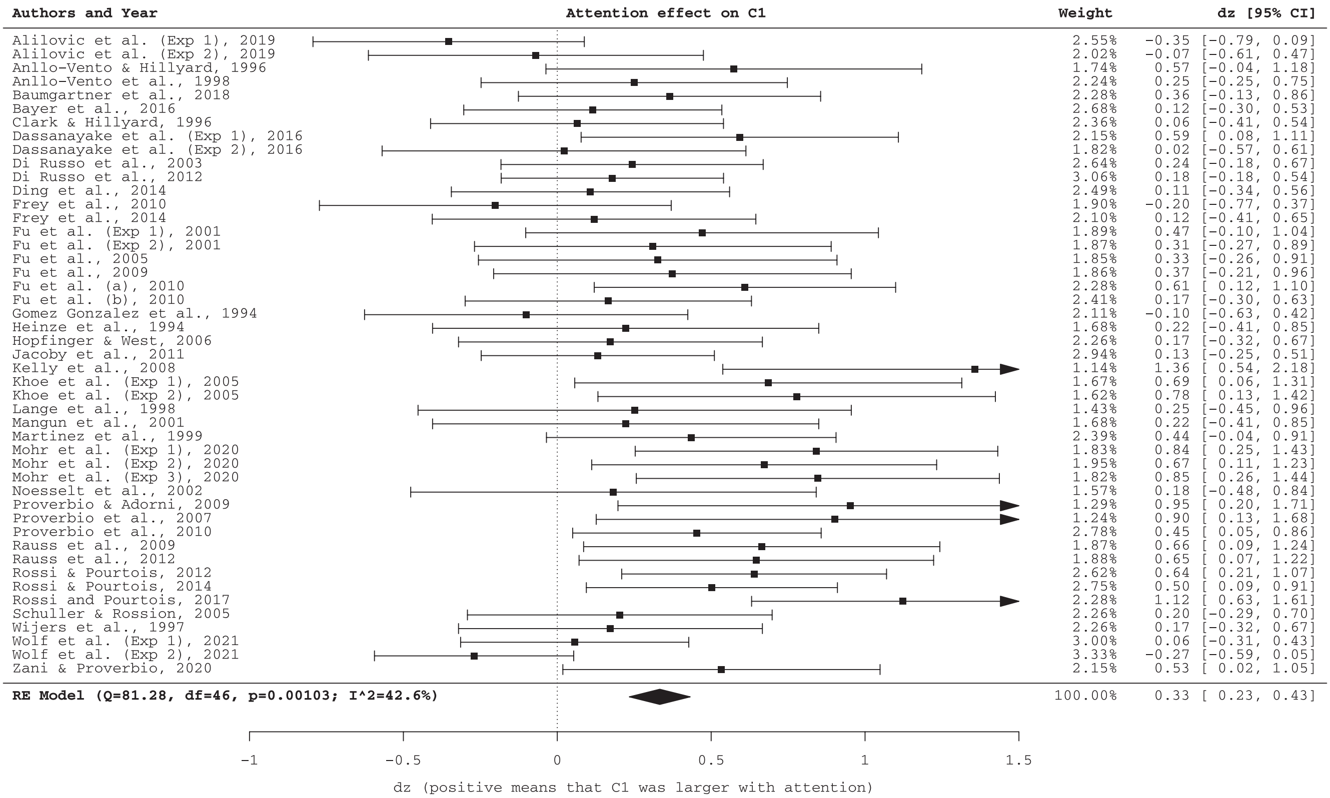

FIGURE 4 Forest plot of the meta-analysis including 47 experiments exploring the effect of attention on the C1 ERP component. A positive effect size refers to a larger C1 amplitude found for *more attention* to the C1-eliciting stimulus compared to *less attention* to the C1-eliciting stimulus.

amplitude difference of the C1 between more and less attention to the C1-eliciting stimulus was estimated to be 0.32 μV, 95% CI [0.23, 0.42].

When assessing the effect of attention on the C1 for upper (Figure 5a) and lower visual field stimulation (Figure 5b) separately, it turned out to be stronger in the former compared to the latter case. Note that in several experiments, only combined effects for both visual fields were reported; thus, they could not be included in this analysis. In the meta-analysis for upper visual field stimulation ($k = 40$), Rossi and Pourtois (2017) was identified as an outlier, hence 39 experiments were finally included. The pooled effect size turned out to be significant and moderate ($d_z = 0.28$, 95% CI: [0.19, 0.37], $p < .0001$), and the between-study heterogeneity ($I^2 = 12.95\%$, $Q = 43.70$, $p = .24$) was not significant. For lower visual field stimulation ($k = 8$), Zhang et al. (2012) was identified as an outlier. The remaining 7 experiments showed a marginally significant effect on the C1 ($d_z = 0.23$ (95% CI: [−0.01, 0.46], $p = .06$) and the heterogeneity across these experiments was significant ($I^2 = 52.47\%$, $Q = 12.90$, $p = .04$). When the conservative approach was used, the effect sizes were 0.20 (95% CI: [0.08, 0.31], $p = .0007$) and 0.19 (95% CI: [−0.03, 0.40], $p = .09$), for upper and lower visual field stimulation respectively.

3.2 | Moderator analysis

To explore whether the type of attention contributes to the observed heterogeneity of effect sizes, a mixed-effects model was fitted. Results revealed no significant effect ($Q(2) = 3.35$, $p = .19$); thus, results did not suggest that type of attention modulated effects of selective attention on the C1 (Figure 6). After accounting for the type of attention in the model, heterogeneity decreased but remained moderate ($I^2 = 39.80\%$, $Q = 73.84$, $p = .003$). Notably, within each type of attention, the pooled effect sizes suggested a significant effect of attention: d_z for *spatial attention* ($k = 31$) was 0.28, 95% CI [0.17, 0.39], $p < .0001$; d_z for *load* ($k = 7$) was 0.54, 95% CI [0.28, 0.79], $p < .0001$; and d_z for *other* ($k = 9$) was 0.35, 95% CI [0.07, 0.62], $p = .0126$.

3.3 | Publication bias

The funnel plot showed that the experiments ($k = 47$) were not symmetrically distributed around the pooled effect size, suggesting the presence of a publication bias (Figure 7a). This observation was further supported by Egger's regression test ($Z = 3.71$, $p = .0002$). The subsequent trim-and-fill procedure indicated that, to achieve

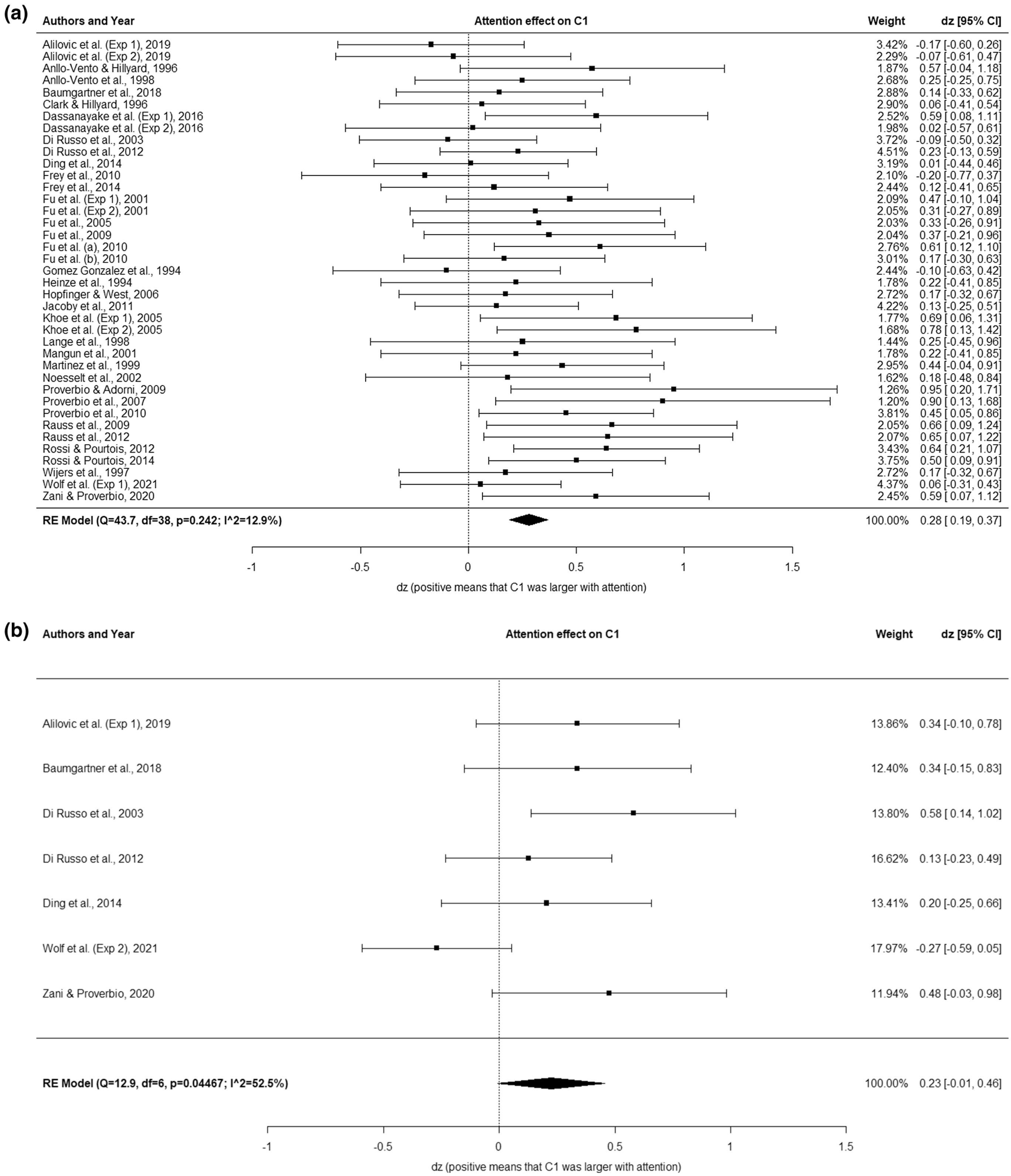

FIGURE 5 Forest plots of the meta-analysis for experiments exploring the effect of attention on the C1 ERP component using (a) upper ($k = 39$) and (b) lower visual field presentations ($k = 7$), respectively.

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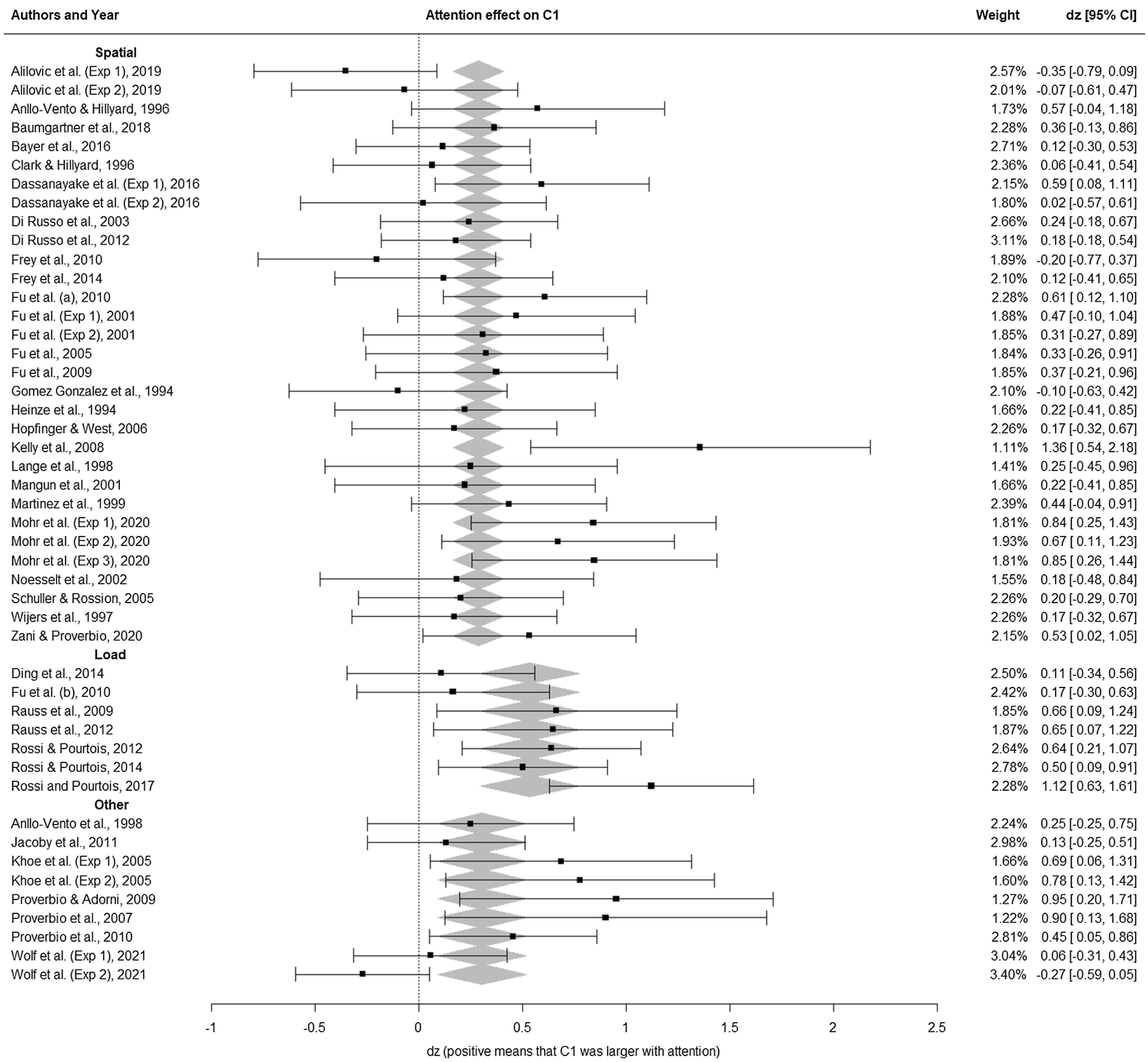

FIGURE 6 The Forest plot shows the effect size for each experiment when the category was included as moderator. The gray diamond indicates the modeled effect size per category.

symmetry, nine additional experiments needed to be added to the left side of the funnel plot (Figure 7b). After those values were added, there remained a significant overall effect of attention ($d_z = 0.24$, 95% CI [0.13, 0.34], $p < .0001$).

3.4 | Cumulative analysis

In the cumulative meta-analysis (Figure 8), the combined evidence of early experiments was very impressive. However, the combined evidence already suggested an effect in 2001 that has remained stable since 2007. Note that the present analysis used a liberal approach to compute effect sizes (see Method). When a conservative approach was used, results suggested that an overall effect was present no later than 2014.

4 | DISCUSSION

This meta-analysis suggests that selective attention has a moderate effect on the C1, suggesting that selective attention to a visual stimulus is associated with a larger C1 component. Moreover, the results of the cumulative meta-analysis suggest that this effect was already present about

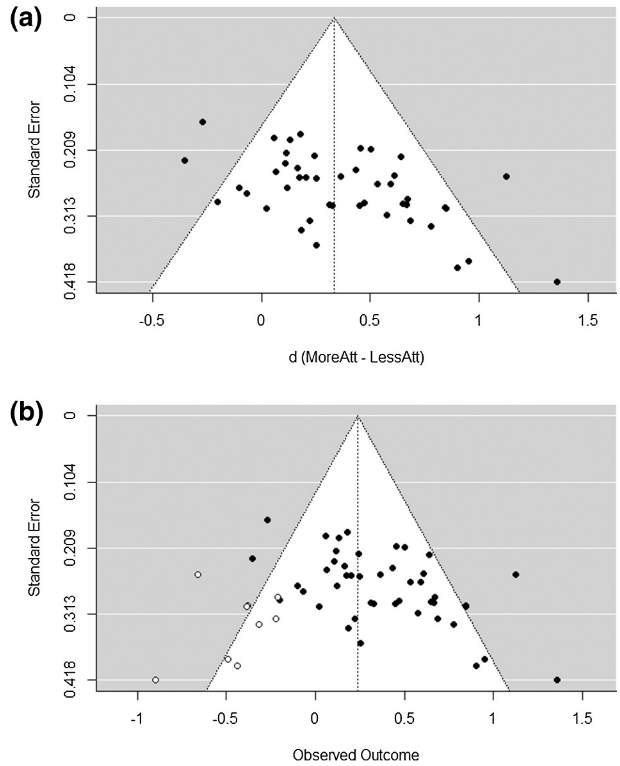

FIGURE 7 (a) Funnel plot showing the distribution of experiments in the meta-analysis. The effect size of each experiment is plotted on the *X*-axis and its standard error is plotted on the *Y*-axis. This plot is asymmetrical, suggesting the presence of a publication bias. (b) Funnel plot obtained after performing the trim-and-fill procedure (see methods). Nine experiments on the right with large effect sizes and small sample sizes were first removed to reach a symmetry. Then these studies as well as their virtual mirror studies which were assumed to be suppressed due to the publication bias were added back to the left side for the re-calculation of the effect size. The unfilled/white circles represent the 9 added fictional experiments in an attempt to compensate for the publication bias.

7–15 years ago, hence long before the notion of a *majority view* vs. *minority view* was introduced for the first time in the literature (see Baumgartner et al., 2018). Furthermore, the explorative moderator analysis did not suggest that the effect of attention on the C1 differed with type of attention; thus, effects of attention were similar for spatial attention, attentional load, or the other type of attention under consideration.

The main outcome of this meta-analysis is not compatible with the *majority view* (Baumgartner et al., 2018), according to which the C1 is impermeable to selective attention. If the *majority view* were true, no effect of selective attention on the C1 should be found when considering the cumulative evidence. In contrast with this view, our meta-analysis suggests that selective attention exerts a modulatory effect on the C1, and this effect is best described as being of moderate size. Thus, results support the *minority view*. It could be argued that the most recent ERP studies on the C1 published in the literature actually contributed the most to support the *minority view* and/or invalidate the *majority view*. However, the results of the cumulative meta-analysis (Lau et al., 1995) clearly show that since 2007 (or at least 2014), the pooled effect size remained stable over the years (Figure 8), suggesting that the significant influence of selective attention on the C1 could already be established about 7 to 15 years ago.

While we found an effect of attention on the C1 in this meta-analysis, the results showed that this effect was likely inflated by a publication bias (see Figure 7). Publication biases are commonly observed in meta-analyses simply because significant results are more likely to be published than null findings (Borenstein et al., 2009). In an attempt to correct this publication bias, we performed a trim-and-fill procedure (Duval, 2005; Duval & Tweedie, 2000a, 2000b). Even though it has been argued that funnel plots and trim and fill procedures are somewhat inappropriate for meta-analyses with heterogeneous data (Terrin et al., 2003), they remain valuable and can be regarded as sensitivity analyses, enabling the identification of experiments that have an excessively large impact on the mean effect size. After correcting for this publication bias, results showed that there remained an effect of attention on the C1.

Moreover, to deal with nonsignificant results or missing data, we contacted each corresponding author of the articles in question, asking them to share their raw data to estimate missing effect sizes. Despite that, there were still problematic cases where the data could not be retrieved by the authors of the original studies, mostly because they were carried out a long time ago and the data were no longer accessible. In these cases, we had to calculate effect sizes by using hypothetical correlation coefficients or *p*-values, and we note that some variability exists in the way effect sizes can or should be calculated in this situation. While some researchers compute the effect size of a nonsignificant result as zero (e.g., Rosenthal, 1991; Voyer et al., 2017), others assume a *p*-value of .50 (e.g., Coll, 2018; Moran et al., 2017; Rosenthal et al., 1994) to obtain the corresponding effect size. Whereas setting the effect size as zero is a conservative approach (i.e., against the presence of an effect), setting the *p*-value to .50 is a liberal approach (i.e., for the presence of an effect). Moreover, different authors use correlations between measures of .75 (Dunlap et al., 1996), .70 (Coll, 2018), or .50 (Rosenthal, 1991). In our meta-analysis, we assumed a liberal approach and set the correlation to be .75 for the missing values (see Pool et al., 2016 for a similar approach), which may have led to a slight overestimation of missing effect sizes. To address this concern, we also performed the meta-analysis using more conservative values, namely .50 for the correlation and .99 for *p*-value (see C1meta_conservative.

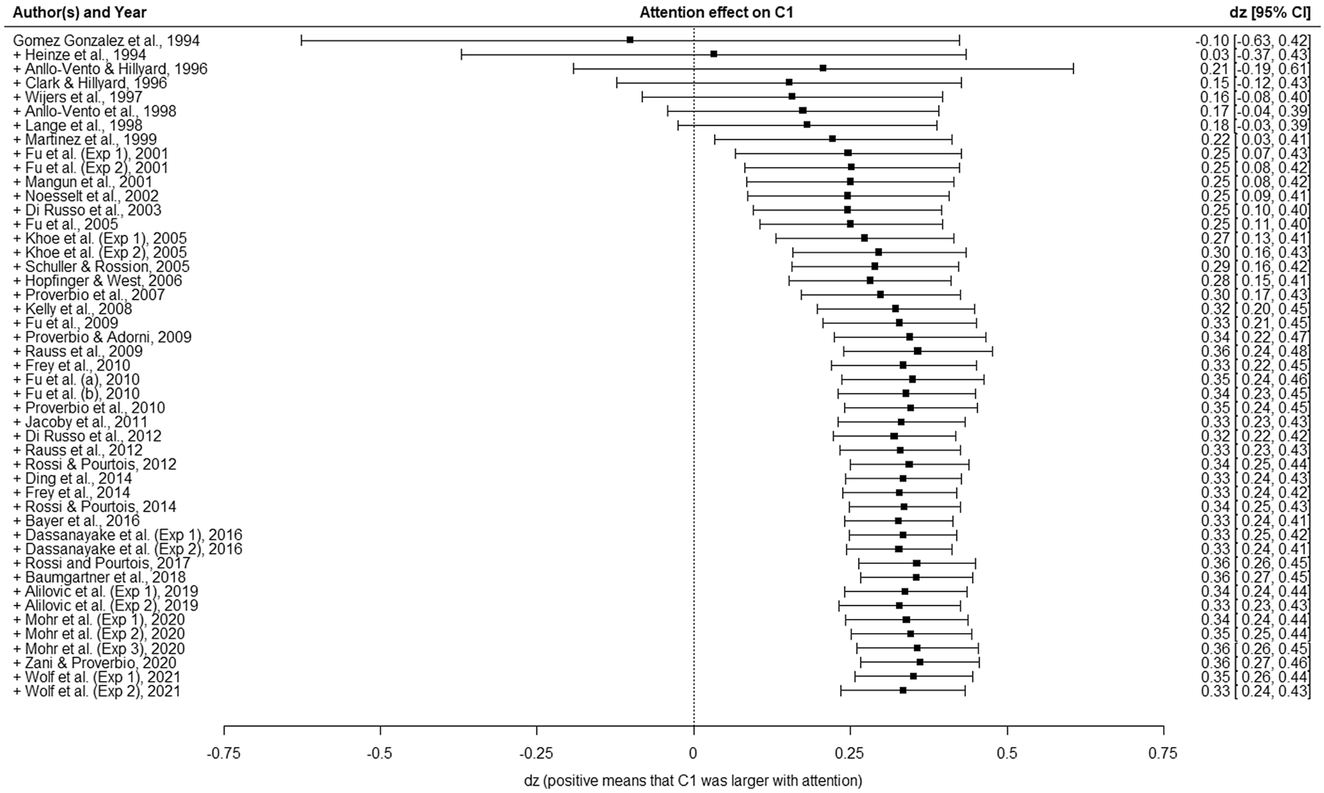

FIGURE 8 Forest plot of the cumulative analysis. It shows the cumulative effect of selective attention on the C1 and includes all experiments up to that time (i.e., year of publication).

html in the OSF project). This analysis yielded a small effect size ($d_z = 0.21$, 95% CI [0.10, 0.33], $k = 48$). To avoid these issues in the future, we recommend reporting open data (Munafò et al., 2017). Specifically, we suggest to report more systematically and exhaustively the results in C1 studies—e.g., to report effect sizes as well as standard deviations of differences, or correlations, between the experimental conditions, in addition to means and *SD*s, and to do so even in cases of nonsignificant results. These efforts would greatly help to more easily integrate new findings with existing C1 studies, and to eventually obtain a better estimate of the true effect of selective attention on this first visual ERP component.

A second main contribution of our study pertains to the moderator analysis. Results did not suggest differences between the three types of attention under consideration. As a matter of fact, the *majority view* mainly refers to studies of spatial attention but not attentional load (see Slotnick, 2018 for a discussion). These two effects are probably subserved by different attention control processes (Handy & Mangun, 2000; Lavie, 1995, 2005, 2010; Torralbo & Beck, 2008). However, the results of our meta-analysis do not suggest that their implementation at the level of the C1 differs. Even though the effect size of attentional load on the C1 appears to be numerically larger than that of spatial attention (or the *other* category; see Figure 6), a direct statistical comparison between them revealed no significant differences. These results need to be interpreted with caution, for statistical as well as conceptual reasons: statistically, a nonsignificant finding is not proof of the absence of an effect (Dienes, 2008; Makin & Orban de Xivry, 2019; Wiens & Nilsson, 2017); and the numbers of experiments for each of the three types of attention were not comparable. Perhaps more importantly, on a conceptual level, it remains unclear whether spatial attention, attentional load, and other attentional mechanisms can be directly compared in terms of their effects on behavior or neural processing.

In addition to the type of attention, whether the stimuli were presented in upper or lower visual field was also considered in separate analyses. Results showed a significant effect of attention in the upper visual field, which was weaker in the lower visual field. However, these results need to be interpreted with caution, because not all experiments examined and reported effects of attention on the C1 separately for the two hemifields. Thus, after removing outliers, 39 experiments could be included in the meta-analysis for upper visual field presentation, whereas only 7 actually presented results for the lower visual field. Hence, the results for the lower visual field are based on a very limited number of studies. Although there are a number of anisotropies between the upper and lower visual fields (Karim & Kojima, 2010; Previc, 1990; Skrandies, 1987), our results do not provide strong evidence for a clear-cut difference in terms of their permeability for attentional modulations.

Notably, the between-study heterogeneity was of medium size and remained so even after modeling the type of attention in a moderator analysis. Thus, it could be argued that our division of the existing C1 ERP studies according to this moderator did not explain a sufficient amount of variance between different protocols. This is an important caveat and additional research is needed to examine the potential role of other potential moderators.

The results of this meta-analysis have important implications at the theoretical level. Selective attention can operate at different stages and involves the modulation of several cortical as well as subcortical systems (Nobre et al., 2014; Scolari et al., 2015; Szczepanski & Kastner, 2013). Although the visual system is hierarchically organized, many areas in the occipital and temporal lobes contribute to visual processing in parallel (Van Essen & Maunsell, 1983). By computing priority maps, perceptual features encoded in early visual cortex (e.g., spatial frequency, intensity, color) and observer-dependent biases presumably generated in prefrontal and parietal areas (e.g., goals, expectations) can be integrated in a dynamic and flexible way, thereby supporting an efficient selection of the most relevant bits of information (Bisley & Goldberg, 2010; Itti & Koch, 2001; Serences & Yantis, 2006). The results of our meta-analysis indirectly suggest that one can trace the downstream consequences of such priority maps early on following stimulus onset in the primary visual cortex, where the C1 component is mainly generated.

Moreover, although priority maps could be different and non-overlapping for spatial attention and attentional load in fronto-parietal networks, our results suggest that their top-down influences on the earliest stage of cortical processing in V1 in humans could be similar, suggesting a common fate for them in this area. Alternatively, the comparable gain-control effects found at the C1 level for spatial attention and attentional load could reflect the involvement of a multidimensional and domain-general central executive network (Scolari et al., 2015; Shomstein & Gottlieb, 2016). In line with this assumption, recent studies have reported common neural effects within distributed fronto-parietal networks for the processing of spatial and non-spatial information (Hou & Liu, 2012; Liu et al., 2011; Scolari et al., 2015; Szczepanski et al., 2010; Szczepanski & Kastner, 2013). While our meta-analysis does not provide information on the organization of priority maps within fronto-parietal networks, our results suggest a modulation of the earliest stage of cortical processing in V1, irrespective of the type of top-down attention control signal generated in those networks.

In conclusion, this meta-analysis suggests that the C1 ERP component is influenced by selective attention and this top-down effect on initial processing in V1 is of moderate size (Cohen, 1988). When we estimated the effect size in terms of microvolts (i.e., when calculating the unstandardized effect size), the effect of attention corresponds to a 0.32 μV difference between the more and less attention condition in within-subject designs. As such, these results do not support the *majority view* according to which the C1 would be impermeable to top-down attention control. Moreover, despite heterogeneity across experiments, results did not provide evidence that the type of attention manipulated (i.e., spatial attention, attentional load, and other) moderates this effect or reduces this heterogeneity. Tentatively, these results suggest that whereas different top-down control signals for spatial attention or attentional load may originate from non-overlapping regions of the frontal and parietal cortex (Corbetta & Shulman, 2002), they could exert comparable gating effects in V1 early on following stimulus onset.

This meta-analysis is a first attempt to provide a systematic, statistical evaluation of the combined effect of selective attention on the C1. Despite our finding of a moderate global effect, we believe that more empirical work in this area is needed. Recent research suggests that because of publication biases and selective reporting, meta-analyses may overestimate actual effect sizes. For example, when results of meta-analyses were compared with those of large-scale, preregistered replications, effect size estimates were three times larger in meta-analyses than replications, and common statistical correction procedures could not remove this bias (Kvarven et al., 2020). Therefore, we advise to conduct a large-scale preregistered study, preferably in the form of an adversarial registered report. In this framework, researchers with different views have to agree on a feasible study design, method, and analysis; also, they have to agree that the results will be informative no matter their outcome (Nosek & Errington, 2020a, 2020b). Preregistering the study will clarify differences between prediction and postdiction, and minimize biases (Baldwin, 2017; Nosek et al., 2018; Wagenmakers et al., 2018). Although our moderator analyses cannot account for the substantial heterogeneity among experiments, they emphasize that the various experiments may not examine the same effect. An important task for the future is to identify important moderators. Additionally, the effects of prediction (Alilović et al., 2019), motivation (Bayer et al., 2018; Rossi et al., 2017) and task demands (Mohr et al., 2020; Wolf et al., 2021) should also be taken into account in future investigations, as they might moderate the early effects of attention on V1, as captured by the C1 component.

AUTHOR CONTRIBUTIONS

Nan Qin: Conceptualization; data curation; formal analysis; investigation; methodology; project administration; software; visualization; writing – original draft; writing – review and editing. **Stefan Wiens:** Formal analysis; software; writing – review and editing. **Karsten Rauss:** Writing – review and editing. **Gilles Pourtois:** Conceptualization; data curation; funding acquisition; project administration; supervision; writing – review and editing.

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