The redundant target paradigm and its use as a blindsight-test: A meta-analytic study

Doris Schmid[a]*, Constanze Hesse[b], Thomas Schenk[a]

[a] Clinical Neuropsychology

Department of Psychology

Ludwig-Maximilians-Universität München

Leopoldstr. 13, 80802 Munich

Germany

Email: Doris.Schmid@psy.lmu.de *(corresponding author), +49 (0)89 2180-9679

Thomas.Schenk@psy.lmu.de

[b] School of Psychology

University of Aberdeen

King's College

William Guild Building

Aberdeen AB24 3FX

United Kingdom

Email: c.hesse@abdn.ac.uk

Highlights

- Reaction times are shorter for two targets compared to a single target
- This redundant target effect has been used to diagnose blindsight
- Our meta-analysis shows that this effect is replicable in healthy participants
- Only 14 healthy participants are required to achieve the effect with a power of 95%
- For a single patient absence of the effect does not indicate absence of blindsight

Abstract

The redundant target effect (RTE) is the well-known effect whereby a single target is detected faster when a second, redundant target is presented simultaneously. The RTE was shown in different experimental designs and applied in various clinical contexts. However, there are also studies showing non-effects or effects in the opposite direction. Our meta-analysis aims to investigate the replicability of the RTE. Herein, we focused on the clinical context within which the RTE has been applied most often and for which it gained particular prominence: The research on blindsight and other forms of residual vision in patients with damage to the neuronal visual system. The application of the RTE in clinical contexts assumes that whenever vision is present, an RTE will be found. Put differently, the RTE as a tool to uncover residual vision presumes that the RTE is a consistent feature of vision in the healthy population. We found a significant summary effect size of the RTE in healthy participants. The effect size depended on certain experimental features: task type, target configuration in the redundant condition, and how reaction times were computed in the single condition. A specific feature combination is typically used in blindsight research. Analysing studies with this feature combination revealed a significant summary effect size in healthy participants, predicting positive RTEs for future studies. A power-analysis revealed a required sample size of 14 participants to obtain an RTE with high reliability. However, the required sample size is rarely reached in blindsight research. Rather, blindsight research is mostly based on single-case studies. In summary, the RTE is a robust effect on group level but does not occur in every single individual. This means failure to obtain an RTE in a single patient should not be interpreted as evidence for the absence of residual vision in this patient.

1. Introduction

One of the simplest versions of the redundant target paradigm (RTP, Marzi et al., 1986; Schmid & Schenk, 2022) is described as follows: A visual target is presented and the observer needs to press a button as soon as they detect the target. Reaction time in this condition is then compared with a second condition. In the second condition, reaction times are measured in response to two visual targets presented at the same time. It turns out that the presence of a second, redundant target leads to shortened reaction times (Raab, 1962). This is typically called the *redundant target effect* (e.g., Kinchla, 1974). The quantitative reduction in reaction times is termed *redundancy gain* (e.g., Reuter-Lorenz et al., 1995).

In the current study, we applied meta-analytic procedures to quantify the average effect size of the redundant target effect (RTE). Herein, we focused on one specific variant of the RTP: the RTP as used in testing residual vision, e.g., blindsight, in patients suffering from homonymous visual field defects (HVFD; Marzi et al., 1986). In the following sections, we will explain how the RTP has been used in clinical contexts, why we focused on blindsight-testing, and why it is important to quantify the effect size of the RTE for this application.

The RTE is an interesting effect in its own right. Its discovery stimulated research and led to the creation of two theoretical models trying to explain the effect. Raab (1962) postulated the so-called *horse-race model* which is based on simple probability summation. Herein, the reaction to redundant stimuli is faster because there is a race between (at least) two independent stimulus processes. The faster process wins the race and elicits the response. This leads to a higher probability for fast reaction times compared to a single stimulus (Raab, 1962). However, subsequent research revealed that the RTE can be significantly larger than expected by probability summation (e.g., Miller, 1978). Hence, Miller (1982) suggested the mechanism of neuronal summation as an explanation. His model posits a convergence of redundant stimuli leading to faster reaction times and was termed the *coactivation model*.

In many studies, stimuli in the single condition are presented to one hemisphere, e.g., a visual target displayed within the left hemifield. In the redundant condition, both hemispheres are stimulated, e.g., by visual targets displayed within the left and right hemifields (bilateral-redundant condition). This allowed researchers to study inter-hemispheric integration in the healthy but also in the diseased brain. One clinical context investigated inter-hemispheric communication in split-brain patients (for a recent meta-analysis see Westerhausen, 2022; for a list of studies see table 1). In more recent years, the application of the RTP extended to psychiatric disorders already associated with disturbed inter-hemispheric communication like schizophrenia (Florio et al., 2008; Florio et al., 2013) or bipolar disease (Florio et al., 2013).

Table 1

Literature on the RTP in different clinical groups

Clinical group	N studies	References
Homonymous visual field defect	15	Celeghin, Savazzi, et al. (2015); Corbetta et al. (1990); de Gelder et al. (2001); Georgy et al. (2016); Leh et al. (2006); Marzi et al. (2009); Marzi et al. (1986); Müller-Oehring et al. (2009*); Ross et al. (2018); Schärli et al. (1999); Striemer et al. (2009); Tamietto et al. (2010); Tomaiuolo et al. (1997); Whitwell et al. (2011); Wüst et al. (2002)
Split-brain patients	11	Corballis (1998); Corballis et al. (2004); Corballis et al. (2005*); Corballis et al. (2002); Iacoboni et al. (2000); Ouimet et al. (2009); Pollmann and Zaidel (1999); Reuter-Lorenz et al. (1995); Roser and Corballis (2002, 2003); Savazzi and Marzi (2004).
Spatial neglect	3	Corballis et al. (2005*); Müller-Oehring et al. (2009*); Ogourtsova et al. (2011)
Visual extinction	2	Marzi et al. (2000); Marzi et al. (1996)
Schizophrenia	2	Florio et al. (2008); Florio et al. (2013*)
Bipolar disease	1	Florio et al. (2013*)

Note. N studies = Number of studies for the clinical group. * Studies are listed in two categories: Müller-Oehring et al. (2009) and Florio et al. (2013) investigate two patient groups; Corballis et al. (2005) investigate a patient with complete callosotomy showing symptoms of spatial neglect.

In other clinical contexts, the RTP was used as a method to investigate patients with an acquired brain lesion affecting one hemisphere, in particular spatial neglect (Corballis et al., 2005; Müller-Oehring et al., 2009; Ogourtsova et al., 2011), visual extinction (Marzi et al., 2000; Marzi et al., 1996), and HVFD (e.g., Marzi et al., 1986; for further studies see table 1). Spatial neglect, visual extinction, and HVFD share the issue that processing of stimuli is imbalanced between left and right hemifields. In spatial neglect and extinction, the underlying attentional deficit impairs the processing of stimuli in the contralesional hemifield (for a review see Driver & Vuilleumier, 2001). In HVFD, patients are (partially) blind in the contralesional visual field. Patients suffering from one of these three disabilities might respond to stimuli within the affected hemifield under certain conditions. In the following article, we will refer to this phenomenon as *residual visual capacities* (RVCs). In rehabilitation, patients are trained to improve the processing of contralesional stimuli to, ultimately, reduce impairments in daily life (for a review about the rehabilitation in HVFD-patients, see Melnick et al., 2016).

The RTP is a tool used to investigate RVCs. It is worth noting that the RTP used in these three clinical groups is similar to that used in split-brain patients. In the single condition, a target is presented to the ipsilesional hemifield. In the bilateral-redundant condition, a second target is

presented simultaneously to the contralesional hemifield. The comparison between single and bilateral-redundant condition is used to make a conclusion about the processing of the contralesional target. As an example, if HVFD-patients are completely blind within the affected contralesional hemifield, the contralesional target cannot be processed visually. Hence, HVFD-patients perceive only the ipsilesional target, i.e., one target, in the bilateral-redundant and in the single condition. Consequently, there should be no RTE. However, if results show a significant RTE, authors conclude that there must have been residual processing of the contralesional target in the bilateral-redundant condition reducing reaction times (e.g., Marzi et al., 1986).

For the meta-analysis, we focused on the clinical context in which the RTP was used most frequently (see table 1) and for which the RTP gained particular prominence: the diagnosis of RVCs in HVFD-patients (Leh et al., 2006; Striemer et al., 2009). When asked about their visual experience, such patients report that they see nothing within the HVFD. However, some patients perform above chance level in response to visual targets presented within their blind visual field. These RVCs in HVFD-patients have been called *blindsight* (Weiskrantz et al., 1974). The dissociation between intact visual performance and impaired visual consciousness led to influential theoretical claims about the origin of visual awareness and its neuronal correlates (e.g., Weiskrantz, 1999). It was conjectured that if certain brain lesions can leave some visual capacities intact while destroying all awareness of the visual input, that the brain processes underlying visual awareness and visual capacities must be to some extent distinct and independent of each other. Specifically, it was assumed that the very brain region whose damage destroys visual awareness but not visual function, namely the primary visual cortex, must also be the region that is most critically linked to the emergence of visual consciousness (Weiskrantz, 1999). Put differently, the phenomenon of blindsight offered the promise of providing a better understanding of the neuronal processes that lead to visual consciousness. This association of blindsight with the quest to identify the neural basis of consciousness explains why blindsight attracted interest from scientists of various fields including philosophy, neuroscience, cognition, and medicine (for a review see Cowey, 2010).

From a clinical perspective, blindsight offered a promising starting point for treating HVFD. As some patients with HVFD show reliable signs of above-chance performance for targets presented to their blind visual field, it should be possible to use this residual vision to improve the lives of patients with HVFD. Particularly promising is a treatment that uses dynamic visual targets. For example, Huxlin et al. (2009) found that extensive training with random-dot motion targets in the patients' blind visual field improved their ability to identify the direction of the presented movement pattern. This was found even in patients who did not show signs of above-chance movement discrimination prior to the training (Huxlin et al., 2009; Saionz et al.,

2020). Furthermore, the training leads to an expansion of the sighted visual field (Cavanaugh & Huxlin, 2017; Saionz et al., 2020) and to an increase in awareness for moving targets presented within the HVFD (Huxlin et al., 2009; Saionz et al., 2020). Thus, it seems that restoring RVCs offers a promising approach to treat patients with HVFD (for studies of other research groups see for example Ajina et al., 2021; Sahraie et al., 2013; Vaina et al., 2014). Still, the question remains of how many patients possess RVCs in their blind field and hence, how many patients could benefit from such a treatment.

An impressive range of different methods to explore potential RVCs have been employed in past studies leading to the categorization of different types of RVCs (Danckert et al., 2019). *Motion blindsight*, also called the *Riddoch's phenomenon* (Riddoch, 1917), has been tested for example as the temporal detection of a drifting Gabor patch (Ajina et al., 2015). *Action blindsight* describes the capacity of some patients to localize targets within their blind visual field by eye or hand movements (for a review see Danckert & Rossetti, 2005). In *affective blindsight* paradigms, the discrimination or influence of emotional stimuli, like fearful faces, is tested within the blind visual field (for a review see Celeghin, de Gelder, et al., 2015). The term *agnosopsia* (Zeki & Ffytche, 1998) describes RVCs for specific perceptual features like shapes (e.g., Overgaard et al., 2008) or color (e.g., Morland et al., 1999). In the context of this categorization, the RTP most likely measures *attention blindsight*, i.e., stimuli presented to the blind visual field change processing of stimuli within the sighted visual field (Danckert et al., 2019). To account for this variety of RVCs, researchers often run series of tests with a given patient, e.g., there are multiple studies on patient *GY* (see for example Weiskrantz et al. (1974) and de Gelder et al. (2001)). It is possible that a given patient shows one type of RVC but not the other. This reflects the fact that different types of RVCs are related to different neuronal pathways. For instance, affective blindsight is likely mediated by the pathway from the superior colliculi to the amygdala (Ajina et al., 2020). The pathway between the lateral geniculate nucleus and the motion area hMT+ is likely responsible for motion blindsight (Ajina & Bridge, 2018). Depending on the lesion location, the neuronal pathway for motion blindsight might be preserved and the neuronal pathway for affective blindsight might be damaged. This was true for patient *P13* in Ajina et al., 2020. In line with this, patient *P13* showed RVCs for motion stimuli but not for affective stimuli in behavioral testing. Importantly, results for patient *P7* revealed the opposite pattern suggesting a double dissociation for these types of RVCs (Ajina et al., 2020).

For the application of the RTP, this means that the presence of an RTE might be a sign for attention blindsight. However, it is not necessarily predictive for other types of blindsight. The absence of an RTE is a sign for the potential absence of attention blindsight leaving the possibility that other types of blindsight are present.

The RTP could be just another paradigm within the range of different RVC-tests. However, the RTP is worth a closer look because it offers several advantages.

Firstly, the RTP avoids the problem of measuring awareness. The original claim for the description of blindsight, i.e., a patient's ability to use visual information despite complete absence of visual awareness, was questioned in several studies showing that findings depend on how awareness is measured (for a review see Overgaard, 2011). As an example, Mazzi et al. (2016) demonstrated that awareness measures varied between a dichotomous scale and a 4-level scale. The RTP is not affected by this issue because patients do not need to rate their awareness.

In addition, the outcome of the RTP is not affected by biased response criteria (Cowey, 2010; Cowey & Weiskrantz, 1963). In other paradigms, participants are often required to classify their perception by making implicit boundaries between response alternatives. This is especially critical for yes/no decisions about the detection of targets. Patients with a liberal criterion for a '*yes*' response have a lower chance for '*no*' responses for the same visual perception (Cowey, 2010). Hence, for such paradigms differences in blindsight results might reflect differences in response criteria and not necessarily differences in blindsight capacity (e.g., Azzopardi & Cowey, 1998). During the RTP, patients always perceive the target within the sighted visual field. RVCs are measured as the influence of the second, redundant target on reaction times. Hence, the RTP does not require an explicit decision on the presence or nature of targets presented in the blind visual field. As a result, the RTP remains unaffected by differences in response criteria.

Secondly, experimental settings for other RVC-paradigms often have high technical requirements like precise measurement of eye and hand movements in localization tasks (e.g., Ross et al., 2018). In contrast, measuring the RTP requires minimal experimental effort. Experimenters have to present two static, simple, visual targets in two conditions (single vs. redundant) and reaction times should be measured precisely. Moreover, even though fixation behavior has to be monitored closely, a precise recording of saccade characteristics is not necessary. It is sufficient to reliably detect deviations from fixation. Following this, the implementation of an RTP is possible in standard experimental set-ups.

Thirdly, the simple nature of the RTP also markedly reduces the demands on participants. Patients only need to understand and memorize a very simple instruction: "Press a button as fast as possible whenever you see a target". Hence, the task can be conducted in patients having impairments in memory or executive functions. As targets can be big, high contrast, and achromatic, visual acuity can be low and color vision is not necessary. To accomplish a button-press, the demands on the motor system are low.

Fourthly, the fact that patients respond to targets within their sighted visual field solves a further problem: frustration. If patients are urged to respond to targets they do not consciously perceive, they face a seemingly intractable task. As experiments often consist of hundreds of trials, patients tend to lose motivation, get frustrated and tired. Consequently, patients' ability to maintain attention is reduced and RVCs could be underestimated.

These advantages could make the RTP the ideal test for RVCs.

However, another important property of such a test would be that it has very high diagnostic reliability, meaning that the test has a high hit rate and a low false alarm rate. The reliability of the RTP to detect the presence of visual capacities can be investigated in the unaffected hemifield of patients with HVFD, neglect or extinction and also in observers with healthy brains. The visual capacity in the unaffected hemifield was tested in six studies investigating HVFD-patients. Thereof, some studies showed a significant RTE (Celeghin, Savazzi, et al., 2015; Corbetta et al., 1990; Marzi et al., 1986; Tomaiuolo et al., 1997) other studies did not (Müller-Oehring et al., 2009; Wüst et al., 2002). The RTE was also absent for the unilateral-redundant condition in neglect patients (Müller-Oehring et al., 2009). One limitation of this approach is that in these patient groups, visual perception of the ipsilesional hemifield could also be impaired by the lesion, for example, in the sense of '*sightblindness*' in HVFD-patients (Bola et al., 2013).

This limitation does not apply to the RTE in healthy participants. We, therefore, turn now to our second prediction: RTEs should be reliably found in brain-healthy observers. There is, to our knowledge, only one meta-analysis examining the RTE in healthy participants (Westerhausen, 2022). Results showed a significant RTE between 13.9 to 19.1ms. This is a promising result. However, Westerhausen (2022) only included RTE-studies that examined healthy observers *and* split-brain patients. Thereby, seriously limiting the selection of available studies. In addition, Westerhausen (2022) focused on one specific version of the RTP, the bilateral-redundant condition, leaving other variations of the RTP uninvestigated. Hence, the generalizability of the results from this meta-analysis is limited.

In fact, there are RTP-studies, not included in the study of Westerhausen (2022), showing non-significant results (e.g., Omura et al., 2004) or even effects in the wrong direction, i.e., longer reaction times in the redundant condition (Grice et al., 1984).

To summarize, the interpretation of results in clinical samples depends on the consistency of the effect within the underlying healthy population. This consistency has not yet been quantified systematically, leaving conclusions about interhemispheric communication or RVCs in question.

With the current study, we like to address this open research question by applying meta-analytical measures using the entire available literature on the RTP in healthy participants. We will examine the statistical evidence for the robustness of the RTE and thereby test whether the RTP is a reliable clinical tool to investigate residual capacities in the lesioned brain.

In the first step, we defined relevant RTE-paradigms for our meta-analysis. We only wanted to investigate paradigms that were also relevant for the detection of RVCs in HVFD-patients. To achieve this, we analyzed all RVC-studies using RTP and extracted experimental parameters typically used in those studies. In the process of collecting relevant literature for our meta-analysis, we initially considered all studies that tested healthy participants with an RTP irrespective of which clinical group, if any, was tested. Due to our more liberal inclusion criteria, effect size estimates from our meta-analysis were based on a considerably larger set of selected studies than the set used in Westerhausen (2022).

In HVFD-patients, the RTP should detect remnants of visual functions. Thus, the RTE should be as strong as possible to increase the probability to detect RVCs. To see which experimental parameters lead to the strongest effect size, we calculated separate analyses for subgroups of specific test configurations. In this regard, we focused on three experimental characteristics relevant for RVC-research: target configuration, task paradigm, and calculation of the reaction time measure for the single condition. As an add-on, we calculated a supplementary analysis on the following four additional target characteristics: shape, presentation duration, size, and eccentricity (see appendix F). Lastly, we calculated an additional meta-analysis on the subset of studies using the exact combination of experimental features necessary to measure RVCs in HVFD-patients.

Applying these methods, we hope to contribute to RVC-research by quantifying the replicability of the RTE and by providing a recommendation for the experimental design leading to the strongest effect size.

2. Methods

We report how we determined our sample size, all data exclusions, all data inclusion/exclusion criteria, whether inclusion/exclusion criteria were established prior to data analysis, all manipulations, and all measures in the study. Neither the procedure nor the analysis plan of the meta-analysis were pre-registered. Data extracted from the included studies and all analysis codes are available under: https://osf.io/knx28/.

2.1. Systematic literature research and inclusion criteria

To find all relevant, published studies using the RTP, we ran a systematic literature review between 16.06.2020 and 24.09.2020 using databases of Web of Science, Psyndex, PsycInfo, PsyArticles, and PubMed. Search words were: (1) *redundant target effect*, *redundant signal effect*, *redundancy gain*, *spatial summation,* and *hemispheric summation* together with (2) *reaction time*, *hemianop**, *Blindsight*, and *cerebral blindness*. Results of all possible combinations were screened in the following order: by title, by abstract, by full text. Screening was done by two persons. In the case of disagreement, record was kept for further inspection. We aimed to include studies presenting experimental data of the RTE that matched the design of previously known RVC-studies.

In the following, we describe our inclusion criteria. The description of those criteria is divided into two sections: (1) RTP-configurations; (2) stimulus types and stimulus arrangements.

(1) RTP-configurations

We were specifically interested in examining the assumptions concerning the reliability of the RTE for the interpretation of RVCs in HVFD-patients. Therefore, we focused on those studies investigating the RTE in healthy participants with paradigms typically used in previous HVFD-samples (see table 1). It turned out that the following test configurations and methods were used in RVC-studies, i.e., in studies investigating RVCs in HVFD-patients using an RTP:

The typical experimental design of the RTP used in these RVC-studies consisted of two conditions with either one (single condition) or two visual stimuli (redundant condition; exception: four stimuli in Celeghin, Savazzi, et al. (2015) and Georgy et al. (2016)). In all RVC-studies, the outcome measure was reaction time. In some studies, there was an additional experimental condition, e.g., the eccentricity of targets (Roser & Corballis, 2002). We only included these studies, if the RTE was calculated across these conditions, e.g., as a main effect in an analysis of variance (ANOVA). Hence, studies in healthy participants had to provide inference statistics comparing reaction times between the two conditions (one stimulus versus two stimuli). In the meta-analysis, we estimated the average effect size of this specific difference. Regarding the redundant condition of RVC-studies, visual stimuli were identical and presented simultaneously. Redundant stimuli were presented in different configurations. Specifically, the following configurations were found in RVC-studies and included in our meta-analysis. *Bilateral-redundant*: One target in the sighted and one target in the blind visual field. If reaction times are significantly reduced in the bilateral-redundant condition compared to a single target in the sighted visual field, we speak of RVCs. *Unilateral-redundant*: Two targets in the sighted visual field. This condition serves as a control condition to make sure that a given patient shows the RTE independent of the HVFD. In addition, we also included studies that

used the *vertical-redundant configuration (i.e.,* one target above and one target below the fixation symbol). While this configuration was not used in RVC-studies (but only in split-brain studies, e.g., Ouimet et al. (2009)), we included it nevertheless since such a configuration could be usefully employed as a control condition in studies on RVC. As these target configurations serve different purposes, we estimated the average effect size separately for each target configuration. Hence, studies were excluded if the RTE was calculated with mixed target configurations. In summary, studies investigating healthy participants were included if redundant stimuli were identical, presented simultaneously, and presented in one of the three mentioned configurations. Importantly, while the selection of paradigms to be examined was based on RVC-studies, data from all studies measuring healthy participants using those specific RTP-configurations were included irrespective of whether the study investigated an additional clinical group and, if applicable, irrespective of the type of clinical group, e.g., HVFD, neglect, or split-brain patients.

(2) Stimulus type and stimulus arrangements

Different stimulus types and different combinations of stimuli have been used in RTP-research. In our meta-analysis, we wish to focus on the most basic type of redundancy gain, namely the gain that can be mostly attributed to neuronal summation. Neuronal summation refers to the assumption that, by increasing the number of stimulated receptors and activated neurons, the detection of visual events is enhanced, thereby leading to improved processing and faster detection responses. This mechanism is assumed to underly redundancy gains found in HVFD-patients. Evidence for this assumption comes from a study by Tomaiuolo et al. (1997). They examined patients with their cortex removed within one brain hemisphere in an RTP using green light flashes. Significant redundancy gains were found for redundant targets presented to the contralesional visual hemifield. Clearly, in these patients, perceptual processing in visual cortex cannot be responsible for the redundancy gains. It is thus assumed that fairly simple, neuronal summation processes, presumably implemented in preserved subcortical structures, such as the superior colliculi, account for the RTE in these cases. RTP-configurations exploiting such simple mechanisms are presumably the most prevalent, resilient and thus most sensitive measures of RVC and will therefore form the basis of our meta-analysis. Besides, there are RTP-versions that use more complex stimuli (such as words, faces, or emotionally-evocative pictures) or use stimuli arrangements that include distractors, noise signals, or illusory contours. RTEs observed with such complex stimuli presumably depend on emotional processes, Gestalt mechanisms, mechanisms involved in object and pattern recognition as well as other high-level perceptual processes. Success in such complex versions of the RTP will, thus, require the intactness of specific higher areas of the visual cortex. The RTE in such paradigms can therefore only be expected in patients where those

areas and input to those areas have been spared. Since we were primarily interested in paradigms that could detect the most universal and most basic form of RVCs, we decided to focus on paradigms with simple stimuli, such as simple geometric figures (circles, rectangles), gratings, checkerboard patterns and excluded studies that used more complex stimuli and more complex stimulus arrangements.

Furthermore, we only included English-language publications. Apart from that no further exclusion criteria were employed. We accepted all types of behavioral tasks and response instructions, e.g. we included detection-, two-choice- and go/ no-go paradigms (for a more detailed description of typical paradigms used in RTP-research, see Schmid and Schenk (2022)).

An overview of the systematic literature search can be found in figure A.1 in the appendix.

2.2. Meta-analytical procedure

On the basis of the systematic literature research, we selected studies that tested the RTP in healthy participants. For the meta-analysis, we extracted the statistical test analyzing the RTE. These statistical tests will be called *RTE-tests* (redundant target effect - tests) in the remainder of this article. There were 31 studies in which at least one RTE-test fulfilled the inclusion criteria.

In some studies, there were multiple RTE-tests that fulfilled the inclusion criteria. Importantly, effects included in a meta-analysis should be independent of each other (Gleser & Olkin, 2009). However, if statistical tests are calculated on data of the same participants, RTE-tests are dependent. This is true whenever authors ran multiple experiments in the same sample or whenever authors calculated multiple RTE-tests between certain conditions of one experiment. In appendix B, we describe how we dealt with those ambivalent cases.

For each selected RTE-test, we extracted the following information: experimental paradigm, descriptive statistics, and inference statistics. Furthermore, we extracted or calculated the size of the redundancy gain (RG), i.e., reaction time of the single minus reaction time of the redundant condition.

The calculations for the meta-analysis were run in R (version 4.0.3, 2020-10-10). For the meta-analytic procedure, we used the R packages *meta* (Schwarzer et al., 2015), *metafor* (Viechtbauer, 2010), and *dmetar* (Harrer et al., 2019a). Initially, we extracted the effect sizes reported in the studies. If Cohen's *d* was not provided, we estimated Cohen's *d* based on the reported test statistic (see appendix C for formulas based on Rosenthal (1993) and Cooper et al. (2009)). To correct for the population bias, we applied Hedges' *g* correction (Cooper et al.,

2009). If the experiment reported a negative RTE, meaning longer reaction times in the redundant compared to the single condition, we defined Hedges' *g* to be negative.

Next, we estimated the standard error (*SE*) of each effect size. For within-subject designs, to calculate the *SE* we need to know the coefficient *r*, i.e., the correlation between reaction time values from the different conditions (Cooper et al., 2009). However, this correlation was never reported in the included studies. To solve this problem, we relied on the RTP-dataset of our previous study in healthy participants (Schmid et al., 2022, N = 19). In the single condition, we presented one target in the sighted visual field. In the redundant condition, we presented two targets in the sighted visual field either bilaterally or unilaterally. Reaction times in the redundant condition were significantly faster than reaction times in the single condition. Re-analyzing this dataset, we found that the Pearson's product-moment correlation, calculated between the reaction times for the single and redundant condition, was high and significantly different from 0: $r_p(17) = .94$, p <.001. As it is unlikely that all studies showed such a high correlation, we ran the meta-analysis once with $r = 0.94$, once with $r = 0.74$, and once with $r = 0.54$ to get a range of probable results.

As we expected differences in the effect size between RTP-studies using different experimental features, we ran a random effects model across all included experiments and reported the relevant parameters of heterogeneity (parameter τ^2, I^2 statistic, χ^2 Q-statistic). If the Q-statistic is significant, the true effect size is not the same in each study. In this case, the effect size estimated by the meta-analytical model does not reflect one common effect size underlying all studies but the average of a number of different true effect sizes. The parameter τ^2 is the variance of the true effect sizes which is estimated using the Restricted Maximum Likelihood method (REML, Viechtbauer, 2005). Please consult appendix D for an explanation on interpreting the statistics from the random effects model including the measures of heterogeneity.

In the meta-analytic model, each study is assigned a weight indicating how well they estimate the underlying effect size (Harrer et al., 2019b). In general, the precision of the estimate is increased by bigger sample sizes. Hence, studies with bigger sample sizes are assigned higher weights. This is also true for the random-effects model of the current meta-analysis. However, the differences between weights are considerably smaller than in fixed-effects models. This is based on the fact that for a random-effects model, it is necessary to calculate adjusted random-effects weights. The adjusted random-effects weight is calculated on the basis of the SE of each study and the parameter τ^2.

$$weight = \frac{1}{SE^2 + \tau^2}$$

If the parameter r^2, i.e., variance of the true effect size, is high, the weights become similar. Hence, the similarity of weights is a characteristic of random-effects models with a high amount of between-study heterogeneity (Viechtbauer, 2021).

Regarding the summary effect size, we reported the 95% confidence interval (CIs) and the 95% prediction interval (PIs). Both CI and PI are interesting but provide answers to different questions (IntHout et al., 2016). The CI shows the range within which the true mean of effect sizes can be expected. The PI shows within which range an effect size of similar RTP-studies can be expected. Since we wish to establish the reliability with which a significant RTE can be expected in individual studies, the PI-estimate is the more relevant one (IntHout et al., 2016).

To determine which version of the RTP elicits the strongest effect size, we ran a subgroup analysis for three experimental features: paradigm, target configuration, and analysis procedure.

Firstly, different RTE-tests used different paradigms: two choice, go/ no-go, and detection. In two-choice paradigms, participants press one button for the first target type and another button for the second target type. In go/ no-go paradigms, participants respond only to one of two target types. In detection paradigms, there is only one target type.

Secondly, RTE-tests applied different target configurations in the redundant condition: Bilateral, unilateral, and vertical. In the bilateral-redundant configuration, one target is presented to the right side and the other target to the left side. In the unilateral-redundant configuration, two targets are presented either to the left or to the right side. In the vertical-redundant configuration, one target is presented above and one target below the fixation symbol.

Thirdly, there were differences in determining the reaction time of the single condition. The single condition consists of two configurations, e.g., target on the left side and target on the right side. Most RTE-tests were based on the average reaction time of these single target configurations. We will call this the *average-procedure*. Other studies used a more conservative measure of the RTE. Authors argued that reaction times of the redundant condition should be compared with the best performance in response to a single target (e.g., Van der Heijden et al., 1984). As some participants might have a preferred target location, e.g., they respond faster to targets on the right side, reaction times of these trials should be used to calculate the RTE. This way of analyzing the RTE will be called the *faster-procedure*. Regarding the faster-procedure, authors either calculated the RTE-tests on the basis of the single condition with the faster mean reaction time on group level (indicated by F^* in figure 1; e.g., experiment 5 in Grice and Gwynne (1987)) or they calculated the RTE-test selecting the

single condition with the faster mean reaction time per participant and then averaging these values for the group analysis (indicated by *F* in figure 1; e.g., Van der Heijden et al. (1984)).

For the subgroup analysis, we calculated the Q-G statistic (Harrer et al., 2019b). If the Q-G statistic is significant, effect sizes are significantly different between subgroups. In addition, we report measures of heterogeneity per subgroup (same as for the general random effects model: parameter τ^2, I^2 statistic, χ^2 Q-statistic). A detailed description is provided in appendix D.

From all possible combinations of experimental features (paradigm, target configuration, analysis procedure), we are particularly interested in the specific combination used to measure RVCs in HVFD-patients. The RVC-combination is a detection paradigm with a bilateral-redundant condition analyzing the single reaction time with the average-procedure. To estimate the summary effect size of this specific RVC-combination, we selected the RTE-tests from the studies included in the meta-analysis matching this feature combination. Importantly, these RTE-tests were all examined in healthy participants. For the remainder of this article, we will refer to this selection of RTE-tests as the *RVC-combination subset*.

To test the RTE for the RVC-combination subset most precisely, we applied an influence analysis based on the Leave-One-Out method (Viechtbauer & Cheung, 2010). This method provides us with an estimate for those RTE-tests that have the greatest impact on the summary effect size and the heterogeneity between RTE-tests. Influence measures were DIFFITS, Cook's distance, and the covariance ratio (Harrer et al., 2019b). Additionally, we used the Baujat plot to reveal RTE-tests contributing most to the heterogeneity (Baujat et al., 2002). Furthermore, we estimated the publication bias for the RVC-combination subset using the small-sample method (funnel plot; Egger's test; Egger et al., 1997). A funnel plot is based on the dependence of sample size and SE. If a sample size is small, the SE is big and thus the probability is high that a statistical test is not significant. As a consequence, these studies have a lower chance to get published. This pattern is visualized in a funnel plot showing the effect size on the x-axis and the SE on the y-axis. If there is no publication bias, the effect sizes should form a symmetric funnel plot. If there is a publication bias, the funnel plot is asymmetric. The Egger's test quantifies the asymmetry of the funnel plot, i.e., the small-sample effect (Egger et al., 1997). If the Egger's test is significant, the funnel plot is asymmetric indicating the presence of the small-sample effect. Lastly, we used Duval & Tweedie's trim-and-fill procedure to estimate the true summary effect size taking into account the small-sample effect (Duval & Tweedie, 2000). This procedure is also based on the assumption that, in the absence of a publication bias, the funnel plot is symmetric. In the first step, the procedure trims the dataset by removing outlying effect sizes. Then the summary effect size is recalculated. In the second step, the new summary effect size is used as the center of all effect sizes. Furthermore,

the procedure adds one new effect size for each trimmed effect size. However, this new effect size is now mirrored, i.e., at the opposite side of the center. This results in a roughly symmetric funnel plot. Last, the summary effect size is calculated on the trimmed and filled set of effect sizes. This final value is interpreted as an estimate for the true summary effect size corrected for the small-sample effect (Duval & Tweedie, 2000).

3. Results

3.1. Overall meta-analysis

We extracted 39 RTE-tests from 31 studies registered in the systematic literature research. The most easily interpretable outcome variable for the meta-analysis would be the mean reaction time difference between single and redundant condition. This redundancy gain was directly reported or calculable for 35 RTE-tests. For a meta-analysis, it is necessary to additionally calculate the SE of the redundancy gain based on the standard deviations of both conditions (Harrer et al., 2019b). However, the standard deviations were only available for five RTE-tests.

It was, however, possible to estimate Hedges' g and its corresponding SE for 33 RTE-test. Hence, we decided to use Hedges' g as the outcome variable for the meta-analysis. For six RTE-tests, we could not estimate Hedges' g because no test statistic was reported. Hence, these RTE-tests were not included in the meta-analysis (Donkin et al., 2014, Exp.1; Leh et al., 2006; Ridgway et al., 2008; Savazzi & Marzi, 2008, Exp. 1 and Exp. 2; Turatto et al., 2004).

We also excluded the second experiment of Donkin et al. (2014) because of the unusual task instruction. Results showed a reversed redundancy gain. However, this result was based on a main effect across a speed and an accuracy emphasis condition. Herein, the negative RTE was likely driven by the impact of the accuracy emphasis (see also results of the first experiment). Importantly, all the other included RTE-tests in the meta-analysis and all patient studies emphasized speed (respond as fast as possible). Given the unusual manipulation and instruction, the RTE-tests of Donkin et al. (2014) are unrepresentative. Consequently, the RTE-test of the second experiment of Donkin et al. (2014) was also excluded from analysis (post-hoc decision).

From the remaining 32 RTE-tests, two were not significant (Grice & Gwynne, 1987, Exp. 5; Omura et al., 2004) and two RTE-tests were reversed showing longer reaction times in the redundant condition (Grice et al., 1984, Exp. 1 & Exp. 2; see also figure 1).

Results of the meta-analysis with $r = 0.94$ across all 32 RTE-test can be seen in figure 1 (results for $r = 0.74$ and r = 0.54 see table E.1 in the appendix). All models ($r = 0.94$, $r = 0.74$, $r = 0.54$) were significant and showed a summary effect size of $g \geq 1.17$. Thereby the 95%-CIs never

included zero. However, the lower borders of the 95%-PIs were negative in all three models showing that the predicted range of true effect sizes in similar studies included zero. Measures of heterogeneity were high. All Q-statistics were significant and all I^2 were above 75% showing that studies did not share a common effect size. Rather, there is a high amount of variance in the underlying true effect sizes (Borenstein et al., 2009; Higgins et al., 2003). Importantly, the SEs increased with decreasing r ($r = 0.94$: $SE = 0.13\pm0.09$; $r = 0.74$: $SE = 0.28\pm0.18$; $r = 0.54$: $SE = 0.37\pm0.24$). Hence, the heterogeneity within the random effects models decreased with decreasing r. This was also true for the subsequent subgroup analysis and for the later analysis of the RVC-combination subset.

Figure 1 [please insert figure 1 as 2-column fitting image]

Results of the random effects model (r = 0.94) across all studies

Study	N	RG	Config	Task	RT		g	95% CI	weight
Donkin et al., 2014 (Exp. 2)	8	-20.00	V	T	A		-0.99	[-1.29; -0.70]	0.0%
Grice et al., 1984 (Exp. 2)	28	-10.00	V	T	F		-0.71	[-0.86; -0.57]	3.2%
Grice et al., 1984 (Exp. 1)	28	-16.00	B	T	F*		-0.56	[-0.70; -0.42]	3.2%
Yu et al., 2014 (Exp. 2)	128	21.88	V	T	A		0.23	[0.17; 0.29]	3.2%
Omura et al., 2004	21	18.22	B	D	A		0.31	[0.16; 0.46]	3.2%
Grice et al., 1987 (Exp. 5)	28	7.00	V	T	F*		0.35	[0.22; 0.48]	3.2%
Yu et al., 2014 (Exp. 1)	57	25.90	V	D	A		0.38	[0.28; 0.47]	3.2%
Grice et al., 1992 (Exp. 1 & 2)	30	16.50	V	G	A		0.54	[0.41; 0.68]	3.2%
Mooshagian et al., 2008	15	17.18	B	D	A		0.58	[0.39; 0.77]	3.2%
Grice et al., 1990 (Exp. 1)	28	13.00	V	G	F		0.67	[0.53; 0.82]	3.2%
Fischer et al., 2008	32	17.00	B	D	A		0.75	[0.61; 0.88]	3.2%
Schröter et al., 2011	16	7.00	B	D	A		0.86	[0.66; 1.06]	3.2%
Ben-David et al., 2014	44	6.00	V	T	F		0.89	[0.77; 1.01]	3.2%
Tamietto et al., 2010 (Exp. 1)	11		B	D	A		1.04	[0.79; 1.30]	3.1%
Railo et al., 2014	11		B	G	A		1.04	[0.79; 1.30]	3.1%
Mordkoff et al., 1996 (Exp. 2)	12	15.00	V	G	F		1.08	[0.83; 1.32]	3.1%
Van der Heijden et al., 1984	24	11.00	B	G	F		1.09	[0.91; 1.26]	3.2%
Tomaiuolo et al., 1997 (Exp. 2)	4	12.40	B	D	A		1.41	[0.93; 1.89]	3.0%
Miller & Van Nes, 2007 (Exp. 1)	40	25.75	B	D	A		1.45	[1.30; 1.61]	3.2%
Van Koningsbruggen et al., 2017 (1. CG)	11	24.30	B	D	A		1.48	[1.18; 1.77]	3.1%
Murray et al., 2001	15	8.75	B	D	A		1.56	[1.30; 1.82]	3.1%
Van Koningsbruggen et al., 2017 (2. CG)	12	33.00	B	D	A		1.68	[1.37; 1.98]	3.1%
Roser et al., 2002	14	13.60	B	D	A		1.75	[1.46; 2.04]	3.1%
Miller, 2007	16	20.00	B	D	F*		1.78	[1.51; 2.05]	3.1%
Miniussi et al., 1998	12	9.00	B	D	A		1.99	[1.65; 2.33]	3.1%
Miller et al., 2006	14	27.00	B	D	A		2.04	[1.72; 2.36]	3.1%
Murray et al., 2001	15	11.75	U	D	A		2.07	[1.76; 2.38]	3.1%
Schärli et al., 1999	22	15.00	U	D	A		2.08	[1.83; 2.34]	3.1%
Savazzi et al., 2004 (Exp. 2)	8	29.10	B	D	A		2.28	[1.82; 2.73]	3.0%
Florio et al., 2008 (Exp. 2)	18	16.92	B	D	A		2.30	[2.00; 2.61]	3.1%
Tomaiuolo et al., 1997 (Exp. 2)	4	13.40	U	D	A		2.47	[1.79; 3.15]	2.8%
Corballis, 2002	58	16.40	B	D	A		2.48	[2.30; 2.66]	3.2%
Savazzi et al., 2004 (Exp. 1)	8	22.70	B	D	A		4.88	[4.02; 5.74]	2.7%
Overall effect							1.29	[0.95; 1.64]	100.0%
Prediction interval								[-0.76; 3.35]	

Heterogeneity: I^2 = 99%, *p* = 0

-5   0   5   10

Note. The forest plot shows the effect size *g* with its associated 95% confidence interval for each included RTE-test. The summary effect size is represented by the diamond with the corresponding dashed line. The black bar at the bottom represents the 95% prediction interval of the summary effect size. N = Number of participants; RG = redundancy gain in milliseconds; Config = redundant target configuration; V = vertical; B = bilateral; U = unilateral; Task = experimental paradigm of study; T = two-choice, D = detection, G = go/ no-go; RT = reaction time measure of single target condition; A = average reaction time across all single target configurations, F = selected faster single target configuration per participant, F* = selected faster single target configuration on group level; g = Hedges' *g*; 95% CI = 95% confidence interval of Hedges' *g* based on the calculation of the standard error with *r* = 0.94; weight = relative weight of each included RTE-test; Studies that are excluded have a weight of 0.0%. Exp. = experiment, CG = control group.

3.2. Subgroup analysis

As the pattern of results in subgroup analysis was similar for all three values of *r*, we present only the results of *r* = 0.94 in table 2 (results of *r* = 0.74 and *r* = 0.54 see table E.2 in the appendix). As expected, studies using the mean reaction time for the single condition had a higher summary effect size than studies using the condition with the faster reaction time. The 95%-CIs of the summary effect size based on the faster single reaction time included zero. Regarding target configurations, unilateral stimulation led to the highest summary effect size, followed by bilateral stimulation. Vertical stimulation led to a very small summary effect size. All three 95%-CI for the target configurations did not include zero. Comparing paradigms across studies, results showed that two-choice tasks led to the smallest summary effect size (95%-CI includes zero). Detection paradigms had the highest summary effect size, followed by go/no-go designs. Importantly, the number of included RTE-tests (*k*) varied considerably between subgroups, for example, *k* = 5 for go/ no-go and *k* = 22 for detection paradigms.

Table 2

Results of subgroup analysis for random effects model with r = 0.94

model	*k*	*g*	95% CI	95% PI	Q	τ^2	I^2	Q-G
faster	8	0.57	[-0.02, 1.16]	[-1.62, 2.76]	679.2	*0.71*	*99.0*	*7.4***
average	24	1.54***	[1.16, 1,91]	[-0.42, 3.49]	1522.0	0.85	98.5	
bilateral	21	1.50***	[1.08, 1.93]	[-0.59, 3.60]	1254.3	0.96	98.4	
unilateral	3	2.11***	[1.92, 2.30]	**[0.87, 3.34]**	1.16	0.00	0.0	63.3***
vertical	8	0.43*	[0.05, 0.80]	[-0.96, 1.81]	362.2	0.28	98.1	
two-choice	5	0.04	[-0.54, 0.62]	[-2.27, 2.35]	397.8	0.44	99.0	
go/ no-go	5	0.87***	[0.64, 1.11]	[0.00, 1.74]	36.2	0.06	89.0	24.0***
detection	22	1.68***	[1.30, 2.05]	[-0.20, 3.55]	1047.2	0.77	98.0	

Note. Model = Model of subgroup analysis for different experimental features: (1) reaction times for single condition: faster- or average-procedure; (2) target configuration in redundant condition: bilateral, unilateral, or vertical; (3) experimental paradigm: two-choice, go/ no-go, or detection. *k* = number of included effects; *g* = estimate of summary effect size based on Hedges' *g*; 95% CI = 95% confidence interval of the summary effect size; 95% PI = 95% prediction interval of the summary effect size; 95% PIs excluding zero are highlighted in bold; Q = Q statistic for statistical heterogeneity; τ^2 = estimate of the variance between RTE-tests; I^2 = percentage of the observed variance which is due to real differences in effect sizes; Q-G = Q statistic for subgroup differences; * p<.05. ** p<.01. ***p<.001.

These differences and the correlation coefficients *r* affected the measures of heterogeneity and hence the 95%-PIs. When considering the PIs (rather than the CI) the effects were less robust. Most 95%-PIs included zero. The 95%-PI of the unilateral-redundant target configuration did not include zero, but only for *r* = 0.94. For *r* = 0.74 and *r* = 0.54, the 95%-PIs of the go/ no-go paradigm and the detection paradigm did not include zero.

3.3. Subset of studies using the RVC-paradigm

To assess the summary effect size specifically for RVC-tests, we calculated a random effects model for the RVC-combination subset. Importantly, RTE-tests investigated the RTE in healthy participants and we selected the experimental paradigms fitting to RVC-research. Regarding outliers, the RTE reported in Savazzi and Marzi (2004, Exp. 1) has an effect size apparently far above all other studies (g = 4.88; compare to other values in figure 2). This is confirmed by the influence analysis showing extreme values for DIFFITS, Cook's distance, and the covariance ratio (see table E.3 in the appendix). For *r* = 0.94, the values are even above the cutoff proposed by Viechtbauer and Cheung (2010). In the Baujat plot, the RTE-tests of Omura et al. (2004) and Corballis (2002) were contributing most to the heterogeneity for all three values of *r*. Additionally, these two RTE-tests were conspicuous for some of the measures of influence, e.g., a covariance ratio below 1. However, they did not exceed the cutoff values. Furthermore, Omura et al. (2004, N = 21) and Corballis (2002, N = 58) had a comparatively high number of participants compared to Savazzi and Marzi (2004, Exp. 1, N = 8). Consequently, we decided to keep Omura et al. (2004) and Corballis (2002) and only exclude Savazzi and Marzi (2004) from the RVC-combination subset for all values of *r*.

The meta-analysis for all three levels of *r* yielded significant summary effect sizes of *g* ≥ 1.43 (see figure 2 for *r* = 0.94; see table 3 for all values of *r*). All 95%-CIs and 95%-PIs were above zero. Again, heterogeneity was high for all values of *r* with I^2 >75% and significant *Q*-tests (Borenstein et al., 2009; Higgins et al., 2003).

Funnel plots showed an asymmetric distribution of effect sizes and sample sizes indicating the presence of a publication bias for all values of *r* (see figure E.1 in the appendix). This asymmetry in the data was confirmed by significant Egger's tests (table 4) showing that there is a lack of RTE-tests having small sample sizes and small effect sizes. With Duval & Tweedie's trim-and-fill procedure compensating for a publication bias, the summary effect sizes were still significant. However, with that correction, the 95%-PIs included zero. This means that if all studies about the RTP fitting to our RVC-combination subset would be published, the predicted range of true effect sizes in similar studies could include zero.

Figure 2 [please insert figure 2 as 2-column fitting image]

Results of the random effects model (r = 0.94) for the RVC-combination subset with outlier correction

Study	N	RG		g	95% CI	weight
Omura et al., 2004	21	18.22		0.31	[0.16; 0.46]	6.5%
Mooshagian et al., 2008	15	17.18		0.58	[0.39; 0.77]	6.4%
Fischer et al., 2008	32	17.00		0.75	[0.61; 0.88]	6.5%
Schröter et al., 2011	16	7.00		0.86	[0.66; 1.06]	6.4%
Tamietto et al., 2010 (Exp. 1)	11			1.04	[0.79; 1.30]	6.3%
Tomaiuolo et al., 1997 (Exp. 2)	4	12.40		1.41	[0.93; 1.89]	5.7%
Miller & Van Nes, 2007 (Exp. 1)	40	25.75		1.45	[1.30; 1.61]	6.5%
Van Koningsbruggen et al., 2017 (1. CG)	11	24.30		1.48	[1.18; 1.77]	6.2%
Murray et al., 2001	15	8.75		1.56	[1.30; 1.82]	6.3%
Van Koningsbruggen et al., 2017 (2. CG)	12	33.00		1.68	[1.37; 1.98]	6.2%
Roser et al., 2002	14	13.60		1.75	[1.46; 2.04]	6.2%
Miniussi et al., 1998	12	9.00		1.99	[1.65; 2.33]	6.1%
Miller et al., 2006	14	27.00		2.04	[1.72; 2.36]	6.2%
Savazzi et al., 2004 (Exp. 2)	8	29.10		2.28	[1.82; 2.73]	5.8%
Florio et al., 2008 (Exp. 2)	18	16.92		2.30	[2.00; 2.61]	6.2%
Corballis, 2002	58	16.40		2.48	[2.30; 2.66]	6.4%
Savazzi et al., 2004 (Exp. 1)	8	22.70		4.88	[4.02; 5.74]	0.0%
Overall effect				**1.49**	**[1.17; 1.81]**	**100.0%**
Prediction interval					**[0.07; 2.91]**	

Heterogeneity: I^2 = 97%, $p < 0.01$

-2   0   2   4   6

Note. The forest plot shows the effect size *g* with its associated 95% confidence interval for each included RTE-test. The summary effect size is represented by the diamond with the corresponding dashed line. The black bar at the bottom represents the 95% prediction interval of the summary effect size. The RVC-combination subset included all RTE-tests using the mean single reaction time of a detection paradigm with a bilateral-redundant condition. N = Number of participants; RG = redundancy gain in milliseconds; redundancy gain is the difference in reaction times between redundant and single condition; *g* = estimation of effect size based on Hedges' *g*; 95% CI = 95% confidence interval of Hedges' *g* based on the calculation of the standard error with *r* = 0.94; weight = relative weight of each included effect; Outliers that are excluded have a weight of 0.0%; Exp. = experiment, CG = control group.

Based on the results of the RVC-combination subset, we conducted a power analysis to compute the sample size that is necessary for future studies to obtain an RTE with high probability. The summary effect size estimated in the meta-analysis, namely Hedges' *g,* is based on Cohens' d_z (see appendix C for formulas). Hence, summary effect size, *g* = 1.49 (for *r* = 0.94), can be used to conduct an a-priori power analysis using G*Power 3.1.9.7 (Faul et al., 2009). The statistical test used is a one-sided paired t-test with a significance level of 5%. In the case of RTP-studies in HVFD-patients, we should aim for a very high probability for detecting RVCs. This can be achieved by conducting studies with high power. The higher the power, the higher the probability that a significant result reflects a true effect (true positive, see calculation of positive predictive value) and the lower the probability to miss true effects (false negative) (Button et al., 2013). Hence, we decided to use a power-value of 95%. As a result,

the required sample size was N = 7 (non-centrality parameter δ = 3.94, critical t = 1.94, df = 6, actual power = 0.96). The resulting value of 7 participants was the same for $r = 0.74$ ($g = 1.46$) and $r = 0.54$ ($g = 1.43$). Using the more common power-level of 80%, the required sample sizes were N = 5 for all values of r. However, this calculation is based on the effect size estimated on the basis of published studies. It is known that this estimate is often an overestimation of the true effect, since studies with lower effect sizes will frequently remain unpublished. When taking this publication bias into account we arrive at substantially higher value for the required sample size.

Based on Duval & Tweedie's trim-and-fill correction, the summary effect size is estimated to be $g = 0.96$ (for $r = 0.94$), resulting in a required sample size of N = 14 (α = .05, power = 0.95, non-centrality parameter δ = 3.59, critical t = 1.77, df = 13, actual power = 0.96), irrespective of the r-value chosen (i.e., for $r = 0.74$, $g = 0.96$, N = 14; for $r = 0.54$, $g = 0.95$, N = 14). Using the more common power-level of 80%, the required sample sizes compute to N = 9 for all values of r.

Table 3

Results of the random effects model for the RVC-combination subset with outlier correction

r	k	g	95% CI	95% PI	Q	τ^2	I^2
0.94	16	1.49***	[1.17, 1.81]	**[0.07, 2.91]**	585.4***	0.41	97.4
0.74	16	1.46***	[1.12, 1.79]	**[0.09, 2.83]**	135.1***	0.38	88.9
0.54	16	1.43***	[1.09, 1.78]	**[0.10, 2.76]**	76.4***	0.35	80.4

Note. r = correlation coefficient used to estimate the standard error of the effect size for each RTE-test; k = number of included RTE-tests; g = estimate of summary effect size based on Hedges' g; 95% CI = 95% confidence interval of the summary effect size; 95% PI = 95% prediction interval of the summary effect size; 95% PIs excluding zero are highlighted in bold; Q = Q statistic for statistical heterogeneity; τ^2 = estimate of the variance between RTE-tests; I^2 = percentage of the observed variance which is due to real differences in effect sizes; * p<.05. ** p<.01. ***p<.001.

Table 4

Results of the publication bias tests in the RVC-combination subset

r	*Egger's test*			Duval & Tweedie's trim-and-fill			
	Intercept	95% CI	t	add k	*g*	95% CI	95% PI
0.94	8.88	[0.97, 16.79]	2.20*	7	0.96***	[0.55, 1.37]	[-1.14, 3.06]
0.74	4.27	[0.47, 8.07]	2.20*	7	0.96***	[0.55, 1.36]	[-1.05, 2.96]
0.54	3.21	[0.35, 6.06]	2.20*	7	0.95***	[0.55, 1.36]	[-0.97, 2.87]

Note. r = correlation coefficient used to estimate the standard error of the effect size for each RTE-test; Egger's test for asymmetry due to publication bias; Intercept = intercept of asymmetry in the data; 95% CI = 95% confidence interval of the intercept; t = t-statistic of Egger's test; Duval & Tweedie's trim-and-fill = procedure to estimate the summary effect size without the publication bias; add k = number of added studies; *g* = estimate of summary effect size based on Hedges' *g*; 95% CI = 95% confidence interval of the summary effect size; 95% PI = 95% prediction interval of the summary effect size; * p<.05. ** p<.01. *** p<.001.

4. Discussion

4.1. Summary of results from meta-analysis

The redundant target paradigm (RTP) is based on the redundant target effect (RTE) which describes reduced reaction times in response to two simultaneously presented targets compared to reaction times in response to a single target. In our meta-analysis, we found a significant positive summary effect size for this phenomenon in healthy participants. The selection of studies for the meta-analysis was tuned to fit those experimental designs typically applied to investigate visual functions in patients suffering from homonymous visual field defects (HVFD). Hence, we conclude that overall this version of the RTP shows a robust RTE.

However, it should be noted that the prediction intervals of the effect sizes included zero. This means, that when conducting a new study with this paradigm, we also have to expect a null effect. This broad range of predicted effect sizes might be due to the considerable variability in the experimental design of included studies. In fact, the summary effect size varied significantly depending on experimental features. In particular, we compared the summary effect size between different types of tasks. The RTE was strongest if participants simply had to detect targets. Go/ no-go tasks led to a considerably smaller but still positive summary effect size. In contrast, the summary effect size was low for two-choice paradigms. Hence, we recommend avoiding two-choice paradigms in future studies. Next, we compared the effect size between target configurations in the redundant condition. Most studies tested a bilateral-

redundant target configuration. This means that one target was presented in the left and the other target in the right hemifield. In this condition, we found a strong summary effect size. A unilateral-redundant target configuration showed an even higher summary effect size. In such conditions, both redundant targets were presented within one hemifield, i.e., either left or right. Last, the summary effect size was low for studies showing targets in a vertical-redundant configuration. Thus, for future studies we suggest to avoid a vertical-redundant configuration. Studies did not only differ in the paradigm itself, but also in the calculation of the outcome variable. This concerned in particular the reaction times of the single condition. There are two possible target locations, e.g., in the left and right hemifield. Hence, there are trials in which a single target is presented in the left hemifield and trials in which a single target is presented in the right hemifield. Most studies calculated the average reaction times across both target locations, i.e., using the average-procedure, and compared these values to the redundant condition. Our results showed a strong summary effect size for this calculation. Some studies, however, chose the target location with the faster reaction time (either on group- or on participant-level; i.e., the faster-procedure) and compared them to the redundant condition. As expected, this is a more conservative estimate of the RTE. As the summary effect size was low, we do not recommend using the faster-procedure in future studies.

Combining the results of the subgroup analysis, the following experimental features should lead to the strongest RTE: a detection task with a unilateral-redundant condition analyzing the average single reaction times.

In the investigation of HVFD-patients, one combination of experimental features is particularly interesting: a detection task with a bilateral-redundant condition using the average-procedure to determine reaction times of the single condition. With this combination, studies investigated residual visual capacities (RVCs), e.g., blindsight (Weiskrantz et al., 1974), within the blind visual field. Herein, one target is presented within the blind visual field and the other target is presented within the sighted visual field. It is a sign for RVCs if reaction times in the bilateral-redundant condition are significantly reduced compared to the single condition (e.g., Marzi et al., 1986). To estimate the summary effect size of the RTP for this specific case, we selected a subset of studies that used this exact combination of experimental features in healthy participants. On the basis of the findings from our meta-analysis for this subset, we can predict a range of true effects above zero for similar studies. This means that one can expect a significant RTE in subsequent studies applying these experimental features (detection task, bilateral-redundant condition, average-procedure for reaction times of single condition) in groups of healthy participants. This is in principle good news for RVC-research.

However, the publication bias might have led to an overestimation of the summary effect size for this subset analysis. The publication bias is a general problem in scientific literature as

already stated by Rosenthal (1979). It occurs because non-significant but valid results are published with a lower probability than significant results. To get a more valid estimate of the true effect size, researchers should conduct pre-registered studies in the future (Munro & Prendergast, 2019). Applying registered reports, non-significant or weak effects would be published with a higher probability. Subsequent meta-analysis could then be based on a less biased set of studies leading to a more valid estimate of the true effect size.

As a practical outcome, the sample size for future studies can be planned on the basis of the current meta-analysis. Accounting for a potential publication bias, we recommend testing 14 participants to obtain an RTE with high reliability (α = 0.05, power = 0.95). With healthy participants, this is easily feasible. For patients-studies, a sample size of 14 might be a bigger challenge but not unsurmountable. As a minimum, future researchers should test at least 5 patients (α = 0.05, power = 0.80, neglecting the potential publication bias).

4.2. The role of different target configurations in RVC-research

In HVFD-patients, we like to measure RVCs. Hence, we should use the combination of experimental features providing the strongest summary effect size to increase the chances to find RVCs. In this respect, one experimental feature, the target configuration in the redundant condition, could still be improved. In particular, results of the meta-analysis showed that the summary effect size for unilateral-redundant configurations was even higher than for bilateral-redundant configurations. Testing RVCs with a unilateral-redundant configuration is not possible in patients suffering from a complete homonymous hemianopia. Both targets would be presented within the blind visual field and hence the patients might, most likely, not respond at all. However, it would be possible in patients in which the HVFD does not affect the whole hemifield. As an example, a patient has a HVFD in the upper right quadrant. To test RVCs, it would be possible to present one target in the upper, i.e., blind, quadrant and simultaneously one target in the lower, i.e., sighted, quadrant. It would be a sign for RVCs, if reaction times are significantly faster in this unilateral-redundant condition compared to a single target in the lower quadrant.

Having said that, it is important to note that the advantage of the unilateral-redundant condition is not as clear as results suggest. In particular, the two studies, included in the meta-analysis, that tested both target configurations require a closer look (Murray et al., 2001; Tomaiuolo et al., 1997). In Tomaiuolo et al. (1997), t-values differed considerably between unilateral (6.79) and bilateral RTE (3.88). However, descriptively, the unilateral redundancy gain differed from the bilateral redundancy gain by only 1 millisecond. In Murray et al. (2001), the RTE was present for unilateral (left and right) and bilateral (upper and lower) redundant conditions. Comparing redundancy gains, the left unilateral redundancy gain was significantly bigger than

the redundancy gains of all other redundant conditions. Descriptively, the right unilateral redundancy gain was smaller than in both bilateral conditions. This means that the advantage of the unilateral target configuration was driven by the left unilateral condition. In Ouimet et al. (2009), there was no main effect of target presentation including bilateral, unilateral, and vertical configurations in the redundant condition. Taken together, further research is required to investigate the specific effects of target configuration on the RTE.

Still, the influence of different target configurations on the RTE is relevant for the interpretation of results in HVFD-patients. In studies on RVCs, performance in the sighted visual field is often taken as a reference for the performance in the blind visual field, e.g., in measuring contrast sensitivity (Mikellidou et al., 2019). The same holds true for rehabilitation. As an example, Huxlin et al. (2009) measured motion discrimination within the blind and sighted visual field prior to training. The performance level found in the sighted visual field was then taken as the training goal. This means that patients trained to discriminate motion direction within the blind visual field until they reached the pre-test performance of the sighted visual field (Huxlin et al., 2009). Regarding the RTP, five studies tested a unilateral-redundant condition as a control condition in HVFD-patients (Corbetta et al., 1990; Marzi et al., 1986; Müller-Oehring et al., 2009; Tomaiuolo et al., 1997; Wüst et al., 2002). Hence, they used the unilateral redundancy gain as a reference for the to-be-expected redundancy gain with bilateral-redundant stimulation. Put simply, it is often assumed that the redundancy gains in the unilateral-redundant and the bilateral-redundant condition are the same. It is then tempting to interpret differences in redundancy gains in the two conditions as a measure of the difference in strength of the signals coming from the blind versus the sighted visual field. However, the findings of the meta-analysis show that such differences in redundancy gains might also be present in healthy participants with two sighted hemifields. This suggests that unilateral versus bilateral differences in the RTE should not be used as a measure for assessing the relative strengths of the visual signals coming from the sighted versus blind visual hemifields in HVFD-patients.

4.3. Limitations of the current meta-analysis

There are two limitations of the current meta-analysis: the unknown correlation coefficient r as well as the high level of heterogeneity.

Regarding the first limitation, the correlation between reaction times of the redundant and of the single condition is necessary to calculate the standard error of the effect size per study which in turn is necessary to calculate meta-analytic models (Cooper et al., 2009). However, no RTP-study included in our meta-analysis reported this correlation coefficient r. As an approximation, we used the value from the RTP-dataset of one of our previous studies (Schmid et al., 2022, $r = 0.94$). This value was taken for each of the included studies. To see whether

results were robust, we additionally calculated the meta-analytic models twice more with lower values of r ($r = 0.74$, $r = 0.54$). Comparing results between values of r, there were only slight differences in terms of which prediction intervals were above zero in the subgroup analysis. Importantly, the pattern of results was consistent for the general model including all studies and for the RVC-combination subset. In summary, the main conclusions of the current study are consistent across varying values of r. Still, we recommend to report the correlation coefficient r for the RTE in future studies. Knowing the true value of r for each study, it would be possible to verify the validity of our results.

The second limitation concerns the high level of heterogeneity in the meta-analytic models, not only in the full set of studies, but also in the RVC-combination subset. When calculating a meta-analysis, it is conventionally assumed that all included studies share a common effect size (Borenstein et al., 2009). However, our results show that this assumption is violated in all our meta-analytic models (significant Q-tests; high values of I^2) indicating that the true effect size varies between studies. In the subgroup analysis, we determined that certain experimental features significantly influence the effect size: the task type, the target configuration in the redundant condition, and the way to calculate the reaction time value for the single condition. However, even in the RVC-combination subset, for which we selected studies with a specific combination of these experimental features, measures for heterogeneity remained high.

The experimental features discussed so far are not the only features contributing to the heterogeneity of RTP-studies. In fact, clinical and neuroscientific research questions led to further variations of stimuli and procedure. As an example, studies in healthy participants (Mohr et al., 2002) as well as in HVFD-patients (Bertini et al., 2013) presented faces to investigate higher-order visual functions. Other studies used colored stimuli to test the contribution of specific brain regions (superior colliculi; Leh et al., 2006; Marzi et al., 2009). Regarding the procedure, there are several variations with unknown effect on the RTE, for instance, an acoustic warning signal at the start of a trial which was implemented in some (e.g., Grice et al., 1984) but not in all studies (e.g., Yu et al., 2014). Another experimental variation of the procedure concerns the duration of stimulus presentation. In some studies, the stimulus was presented until the response (e.g., Grice et al., 1984, Exp. 1). In other studies, stimulus duration was as short as 32 ms (Savazzi & Marzi, 2008, Exp. 2). In the supplementary analysis (see appendix F), we analyzed the most important of these parameters: target shape, presentation duration, target size, and target eccentricity. Results showed that target shape and presentation duration significantly affected the size of the RTE.

4.4. Considerations on sample size and single-case analysis

Besides the limitations of our meta-analysis, our literature research has revealed one important limitation regarding our current knowledge on the RTP. This limitation has consequences for the interpretation of the RTP in the context of RVC-research and for the way in which the RTP can and cannot be used in making judgments about the existence of RVCs in patients suffering from HVFD.

Early on, we gave an overview on different types of RVCs (Danckert et al., 2019) and emphasized that a patient might show one type of RVCs but not the other (e.g., Ajina et al., 2020). In this framework, the RTP can be seen as a measure of attention blindsight (Danckert et al., 2019). In other words, a significant RTE is said to be indicative for this type of RVC. The question that we addressed in this study is whether the RTP can be a sensitive task to measure attention blindness. Assuming that the RTP is a sensitive measure for detecting residual vision leads to the expectation that the RTP should be seen consistently in the intact visual fields of healthy observers. Our meta-analysis presents evidence in favor of this assumption. While there is consensus that the presence of an RTE in a patient's blind field constitutes evidence for residual vision in that blind field, it is less clear how to interpret the absence of an RTE in patients with visual field defects. One might assume that the absence of an RTE indicates impaired or absent attention blindsight. However, this is only valid if the sample size of HVFD-patients has sufficient power to elicit an RTE with high reliability. Regarding the literature on RVCs, there are five studies calculating the RTE on a group level. Thereof, one study showed a significant result which indicates that RVCs were present in this group of HVFD-patients (Celeghin, Savazzi, et al., 2015, N = 6). The RTE was non-significant in the other four studies which indicates the absence of RVCs in their HVFD-samples (Marzi et al., 1986, N = 20; Müller-Oehring et al., 2009, N = 11; Ross et al., 2018, N = 6; Tomaiuolo et al., 1997, N = 4).

Above, we recommended a sample size of at least 14 patients on the basis of our power analysis (95% power). This value was reached only by Marzi et al. (1986). Three studies (Celeghin, Savazzi, et al., 2015; Müller-Oehring et al., 2009; Ross et al., 2018) reached the minimum sample size required for a power of 80% (5 patients).

It is important to note that with a sample size of 5 patients, we still accept a 20% risk of failing to find RVCs even if residual vision is present in the group. This means that even in those few studies that examined RVCs in groups of patients, there is a high risk, that residual vision may have remained undetected. In actual fact, the risk is even higher than what the above considerations imply, because with the exception of one study (Celeghin, Savazzi, et al., 2015) all other studies looking at groups of patients based their interpretations on the analysis of single cases.

More generally, the single-case approach is the preferred approach in RVC-research and the single-case approach is also obviously the method used in clinical diagnosis. Our literature research revealed ten more RTP-studies investigating HVFD-patients all exclusively conducting single-case analysis (Corbetta et al., 1990; de Gelder et al., 2001; Georgy et al., 2016; Leh et al., 2006; Marzi et al., 2009; Schärli et al., 1999; Striemer et al., 2009; Tamietto et al., 2010; Whitwell et al., 2011; Wüst et al., 2002).

It is worth noting that effect size and power estimates based on group-level analysis cannot be transferred to single-case analysis. In particular, it is possible to obtain a significant group effect while a considerable number of participants does not show the effect on single-case level. One study tested the RTE on group- and on single-case level in healthy participants (Schärli et al., 1999). On a group level, the effect was significant with a high effect size (see figure 1). On a single-case level, only 17 out of 22 participants showed a significant effect (77%). This means that 5 healthy participants did not show a significant RTE even though they had full visual functions and no neurological issues. Most likely, these 5 healthy participants did not show the effect due to random fluctuations in reaction times. This result could be used as an estimate for the sensitivity of the RTP.

The study by Schärli et al. (1999) was not meant as a psychometric study for establishing the diagnostic reliability of the RTP and the size and characteristics of their sample are therefore not suited to provide robust estimates of the diagnostic qualities of the RTP as a tool for detecting RVC. Nevertheless, the study shows clearly, that employing the RTP as a test for RVC entails a considerable risk that residual vision might go undetected in relevant patients.

This is relevant for two reasons. First, because the RTP is a very popular test used to diagnose RVC. Secondly, because diagnosing RVC is of considerable clinical and scientific importance. In the clinical context the presence or absence of RVC might determine whether a patient is suited for a given training program and might also influence the evaluation of a given therapeutic intervention. In the context of neuroscientific research, the presence or absence of RVC in patients with damage to certain brain structures might influence our assessment of the functional role of this brain structure in vision. As an example, Leh et al. (2006) investigated the functional role of the superior colliculi (SC) for attention blindsight as measured with the RTP in five hemispherectomized patients suffering from HVFD. The RTP was administered in two versions, one with black/white stimuli and one with blue/yellow stimuli. Colors were chosen based on earlier literature showing that black/white stimuli but not blue/yellow stimuli are processed via the SC. Two of the patients that did not show RVCs in earlier studies also showed no RTE in either condition. In contrast, those three patients with previously established RVCs showed a significant RTE with the black/white stimuli but not with the blue/yellow stimuli. The authors concluded that the absent RTE for blue/yellow stimuli in these three patients was

because such blue/yellow stimuli were not processed via the SC. This dissociation was taken as evidence for the claim that effective sensory processing within the collicular pathways is a necessary precondition for attention blindsight. Our meta-analysis presented evidence that the presence of the RTE is an indicator for RVCs. Hence, the positive findings for black/white stimuli can be interpreted validly as signs for attention blindsight. However, we have also seen that interpreting the lack of the RTE as evidence for the absence of attention blindsight might be questionable if the sample size requirements are not met. This condition was not met in the study by Leh et al. (2006) in which the RTE was calculated on single-case level. Furthermore, a later study showed that Leh et al.'s initial assumption of SC being blind to blue/yellow stimuli is incorrect (Hall & Colby, 2014). This example illustrates the danger of interpreting the absence or condition-specific absence of an RTE in small-sample studies as evidence for absence or condition-specific absence of blindsight.

Given our current knowledge, it is clear that clinical decision based on the RTP are problematic and scientific claims derived from the RTP studies may be potentially flawed. Here, we need to distinguish between claims that are based on positive findings and those based on negative findings. We have currently no reason to assume that the RTP produces many false positives. Thus, claims based on positive findings are not in doubt. In contrast claims based on the absence of redundancy gains have to be treated with great caution given that we know that such gains will regularly fail to materialize even in observers with perfectly intact vision. Our findings and the examination of the relevant scientific literature point to two goals for future studies on the RTP. First, we should aim to establish reliable estimates for the specificity and sensitivity of the RTP. Secondly, we should develop RTPs that are both easy to implement and highly sensitive to the presence of RVCs.

4.5. Conclusion

To summarize, we conclude that the RTE, in particular the version of the RTP relevant for RVC-research, shows a positive and robust summary effect size. For healthy participants, a sample size of only 14 participants is required to reliably obtain an RTE using the specific experimental features. However, this sample size is rarely reached in HVFD-research. Rather, most studies employ single-case analysis to diagnose RVCs for individual patients. We have seen that this approach is problematic in particular when clinical decisions or scientific claims are based on negative findings, i.e., the absence of the RTE. Such negative findings can occur in observers with full visual capacities. This means that the RTP is associated with a considerable risk of false negatives. Thus, absence of the RTE might mistakenly be interpreted as absence of residual vision. In future studies, we should aim to provide robust estimates of

the diagnostic accuracy of the RTP for the diagnosis of RVC and work towards versions of the RTP that can detect RVC reliably.

Acknowledgements

We would like to thank Malte Kühn for his invaluable help in conducting and rating the systematic literature research.

Funding

Appendix A. Overview of the systematic literature research

Figure A.1 [please insert figure A.1 as 2-column fitting image]

Overview of the systematic literature research

```
┌─────────────────────────────────────────────────────┐
│ 1038 records identified by search word combination   │
│                                                       │
│   • 366 records identified by Web OfScience          │
│   • 40 records identified by Psyndex                 │
│   • 153 records identified by PsycInfo               │
│   • 14 records identified by PsyArticles             │
│   • 465 records identified by PubMed                 │
└─────────────────────────────────────────────────────┘
                    │
                    │──────────▶ ┌──────────────────────────┐
                    │            │ 561 duplicates removed   │
                    ▼            └──────────────────────────┘
┌─────────────────────────────────────────────────────┐
│ 477 records screened by title                         │
└─────────────────────────────────────────────────────┘
                    │
                    │──────────▶ ┌──────────────────────────┐
                    │            │ 254 records excluded     │
                    ▼            └──────────────────────────┘
┌─────────────────────────────────────────────────────┐
│ 223 records screened by abstract                      │
└─────────────────────────────────────────────────────┘
                    │
                    │──────────▶ ┌──────────────────────────┐
                    │            │ 5 duplicates removed     │
                    │            │ 87 records excluded      │
                    ▼            └──────────────────────────┘
┌─────────────────────────────────────────────────────┐
│ 131 full text articles accessed                       │
└─────────────────────────────────────────────────────┘
                    │
                    │──────────▶ ┌──────────────────────────┐
                    │            │ 109 records excluded     │
                    ▼            └──────────────────────────┘
┌─────────────────────────────────────────────────────┐
│ 22 records included                                   │
└─────────────────────────────────────────────────────┘
                    │
                    │──────────▶ ┌────────────────────────────────┐
                    │            │ 9 records added by screening of │
                    │            │ reference lists or other sources│
                    ▼            └────────────────────────────────┘
┌─────────────────────────────────────────────────────┐
│ 31 records included in systematic literature review  │
└─────────────────────────────────────────────────────┘
```

Appendix B. Multiple RTE-Tests in a single study

In this appendix we describe how we dealt with studies that reported several RTE-tests that matched our inclusion criteria but for which the independence assumption was actually or potentially violated. The independence assumption is crucial for conducting a meta-analysis (Gleser & Olkin, 2009). A typical problem for independence are data-sets that stem from a sample with identical or overlapping members. In some studies, it was unclear whether the RTE-tests came from identical or overlapping samples of participants (Grice et al., 1984; Grice & Reed, 1992); in other cases it was clear that the different RTE-findings came from the same set of participants (Murray et al., 2001; Ridgway et al., 2008; Savazzi & Marzi, 2008; Tamietto et al., 2010; van Koningsbruggen et al., 2017). Depending on the specific circumstances different options were pursued to resolved those issues.

Firstly, in case of multiple RTE-tests based on one sample, we only included more than one RTE-test if those multiple RTE-tests related to different target configurations. The reason being that different target configurations are used for different research questions in RVC-research and should, therefore, be evaluated separately. The bilateral-redundant configuration tests RVCs. The unilateral-redundant configuration and vertical-redundant configuration could serve as control conditions. Hence, we kept the RTE-tests separate to examine if the summary effect sizes are comparable across these different configurations. This option was applied to Grice et al. (1984) and Murray et al. (2001) (for details, see table B.1).

Secondly, if there were no a-priori reasons to favor one RTE-test over the other, we calculated the average statistical value for this experiment (e.g., t- or F-statistic). As an example, this applied to studies calculating separate RTE-tests for different single target positions, for instance, a single target configuration with a target on the left side compared to the redundant configuration and a single target configuration with a target on the right side compared to the redundant configuration. The same applied to those experiments in which statistical tests were calculated separately for certain condition that are not relevant for the current meta-analysis, e.g., for different colors. Such averaging of statistical tests was done for six studies (Grice & Reed, 1992; Murray et al., 2001; Ridgway et al., 2008; Savazzi & Marzi, 2008; Tamietto et al., 2010; van Koningsbruggen et al., 2017). The average statistical values were then used as the RTE-test (for details, see table B.2).

Table B.1

Inclusion of studies: Option 1 for dependent statistical tests

Study	Exp.	Target configuration
Grice et al. (1984)	1	Bilateral
	2	Vertical
Murray et al. (2001)		Unilateral
		Bilateral

Note. The table shows studies meeting the inclusion criteria but containing multiple statistical tests based on the same sample of participants. For these studies, we applied option 1 meaning that we kept the statistical tests as separate RTE-tests because they examined different target configurations. The reason being that different target configurations are used for different research questions in RVC-research and should, therefore, be evaluated separately. Exp. = Experiment; RTE = redundant target effect; RVC = residual visual capacities.

Table B.2

Inclusion of studies: Option 2 for dependent statistical tests

Study	Exp.	Averaging procedure
Grice and Reed (1992)	1 & 2	We averaged the statistical values of experiment 1 and 2. Experiments differed in the choice of letter stimuli (Exp. 1: A, a, E, e, & Y; Exp. 2: A, D, E, & R) as well as in the stimulus duration (Exp. 1: 200ms; Exp. 2: 150ms).
Murray et al. (2001)	Unilateral	We averaged the statistical values of the tests single vs. unilateral-left and single vs. unilateral-right.
	Bilateral	We averaged the statistical values of the tests single vs. bilateral-upper and single vs. bilateral-lower.
Ridgway et al. (2008)		We averaged the statistical values for the conditions with or without random luminance modulation.
Savazzi and Marzi (2008)	1	We averaged the statistical values for the conditions with short (32ms) or long (96ms) stimulus duration.
	2	We averaged the statistical values for the conditions with low (4.2 cd/m^2) or high (14.2 cd/m^2) target luminance.
Tamietto et al. (2010)		We averaged the statistical values for the color-conditions gray vs. purple and gray vs. red.
van Koningsbruggen et al. (2017)	CG 1	Control group for patient *RE*. We averaged the statistical values for the tests single-left vs. redundant and single-right vs. redundant.
	CG 2	Control group for patient *ML*. We averaged the statistical values for the tests single-left vs. redundant and single-right vs. redundant.

Note. The table shows studies meeting the inclusion criteria but containing multiple statistical tests based on the same sample of participants. For these studies, we applied option 2: As there were no a-prior reasons to favor one statistical test over the other, we averaged statistical values. Exp. = experiment, CG = control group.

Appendix C. Formulas to compute effect size estimates per RTE-test

If Cohen's *d* was not reported for the RTE-tests, we estimated Cohen's *d* based on the reported test statistic using formula (A.1)-(A.3) (Rosenthal, 1993).

$$\text{Cohen's } d_z = \frac{2 \times r}{\sqrt{1-r^2}} \quad \text{with} \quad r = \frac{Z}{\sqrt{n}} \tag{A.1}$$

$$\text{Cohen's } d_z = \frac{t}{\sqrt{n}} \tag{A.2}$$

$$\text{Cohen's } d_z = \sqrt{\frac{F}{n}} \tag{A.3}$$

Where *Z* is a normalized test statistic, *t* is the test statistic of a paired t-Test, *F* is the F-statistic of a repeated measures ANOVA with 1 degree of freedom (df) and *n* is the sample size. These formulas are based on the following relation between *t*- and *F*-values with df = 1 (Rosenthal, 1993):

$$F = t^2 \tag{A.4}$$

Cohen's d_z is also referred to as the standardized mean difference (Lakens, 2013).

The resulting estimates of Cohen's d_z were then corrected for the population bias by using Hedges' *g* correction (Cooper et al., 2009):

$$\text{Hedges'} g = \text{Cohen's } d_z \times \left(1 - \frac{3}{4(n-1)-1}\right) \tag{A.6}$$

Next, we calculated the variance of the effect size v_d (Cooper et al., 2009):

$$v_g = \left(\frac{1}{n} + \frac{g^2}{2n}\right) 2(1-r) \tag{A.7}$$

As the correlation of reaction times between single and redundant target condition was never reported in the included studies, we relied on the RTP-dataset of our previous study in healthy participants (Schmid et al., 2022, N = 19). The result showed a significant Pearson's product moment correlation of $r_p = 0.94$, p <.001. As we cannot expect such a high correlation in all studies included in the meta-analysis, we repeated the analysis for $r = 0.94$, $r = 0.74$, and $r = 0.54$. Like this, we get a range of plausible results.

Derived, from this, we calculated the standard error (*SE*) of the effect size (Cooper et al., 2009).

$$SE_g = \sqrt{v_g} \tag{A.8}$$

The estimated values of Hedges' *g* and *SE* were then input to the meta-analysis functions of the R packages *meta* (Schwarzer et al., 2015) and *metafor* (Viechtbauer, 2010).

Appendix D. Notes on interpreting statistics from the random effects model and the subgroup analysis

Random-effects model statistics

A random effects model allows that the true effect size might vary from study to study. Hence, there are two possible cases. In the first case, studies share a common effect size. In the second case, studies are based on varying effect sizes. To investigate which case applied, we reported measures of heterogeneity for all meta-analytic models: Parameter τ^2, the I^2 statistic and the χ^2 Q-statistic (Borenstein et al., 2009). In the first case, studies share a common effect size and hence heterogeneity is based solely on within-study error. This means that a study's effect size falls within a certain range around the common effect size. In the second case, there is still within-study error but in addition there is a second source of variation namely the variation between the underlying true effect sizes. The ratio of between-studies variation to within-study error indicates which case applies to the current studies. Statistically, we use this ratio to calculate the observed value of Q. This standardized measure is independent of the metric of the effect sizes. The observed value of Q is then compared to an expected value of Q. The expected value of Q assumes that the first case holds true, i.e., there is only within-study error, and is calculated as the degrees of freedom (number of studies minus one). The difference between observed and expected value of Q reflects the between-studies variation. As Q-values follow a χ^2 - distribution, we can test whether the difference is significantly different from zero, i.e., reject the null hypothesis that there is only within-study error (Borenstein et al., 2009). On the basis of the difference between observed and expected value of Q, we can calculate τ^2. τ^2 is the variance of the true effect sizes in the same metric as the effect sizes. The square-root of τ^2, gives us the standard deviation of the true effect sizes τ. If we assume a normal distribution, we can describe the distribution of true effect sizes with τ as the standard deviation and the summary effect size as the mean (Borenstein et al., 2009). The I^2 statistic is also based on the difference between observed and expected value of Q. However, this time, the difference is dived by the observed value of Q and then multiplied with 100%. Hence, I^2 is the proportion of the total variance reflecting between-studies variation (Borenstein et al., 2009). To interpret I^2, we applied the categorization of Higgins et al. (2003) defining a heterogeneity of 25% as low, of 50% as moderate, and of 75% as high.

Subgroup analysis

The Q-G statistic comparing effect sizes of subgroups follows a similar idea as the Q-statistic for heterogeneity (Harrer et al., 2019b). Again, there are two possible cases. In the first case, subgroups share a common effect size meaning that heterogeneity is based solely on the within-subgroup variance. In the second case, there is within-subgroup variance and additional

between-subgroup variance meaning that subgroups have separate true effect sizes. Firstly, we calculate the observed Q-value based on the ratio of between-subgroup variance to within-subgroup variance. Secondly, we calculate the expected Q-value as the degrees of freedom (G - 1) whereas G is the number of subgroups. The difference between the observed and the expected value of Q follows a χ^2 - distribution. Hence, we can test the null hypothesis that the effect sizes of all subgroups are equal (Harrer et al., 2019b).

Appendix E. Results

Table E.1

Results of random effects model across all included RTE-tests

r	k	g	95% CI	95% PI	Q	τ^2	I^2
0.94	32	1.29***	[0.95, 1.64]	[-0.76, 3.35]	2340.0***	0.98	98.7
0.74	32	1.22***	[0.90, 1.53]	[-0.58, 3.01]	540.0***	0.75	94.3
0.54	32	1.17***	[0.86, 1.47]	[-0.52, 2.85]	305.2***	0.65	89.8

Note. r = correlation coefficient used to estimate the standard error of the effect size for each experiment; k = number of included effects; g = estimate of summary effect size based on Hedges' g; 95% CI = 95% confidence interval of the summary effect size; 95% PI = 95% prediction interval of the summary effect size; Q = Q statistic for statistical heterogeneity; τ^2 = estimate of the between-study variance; I^2 = percentage of the observed variance which is due to real differences in effect sizes; * p<.05. ** p<.01. ***p<.001.

Table E.2

Results of subgroup analysis for random effects model with r = 0.74 and r = 0.54

r = 0.74

model	k	g	95% CI	95% PI	Q	τ^2	I^2	Q-G
faster	8	0.56	[-0.03, 1.14]	[-1.59, 2.70]	156.7	0.68	95.5	
average	25	1.44***	[1.11, 1.77]	[-0.18, 3.06]	351.2	0.58	93.5	6.7**
bilateral	21	1.42***	[1.05, 1.80]	[-0.33, 3.17]	289.4	0.66	93.1	
unilateral	3	2.11***	[1.71, 2.51]	[-0.46, 4.68]	0.3	0.00	0.0	38.8***
vertical	8	0.42*	[0.05, 0.78]	[-0.90, 1.73]	83.6	0.26	91.6	
two-choice	5	0.04	[-0.54 0.62]	[-2.23, 2.31]	91.8	0.42	95.6	
go/ no-go	5	0.84***	[0.59, 1.08]	**[0.08, 1.59]**	8.4	0.04	52.2	24.1***
detection	22	1.58***	[1.25, 1.91]	**[0.05, 3.10]**	241.7	0.51	91.3	

r = 0.54

model	k	g	95% CI	95% PI	Q	τ^2	I^2	Q-G
faster	8	0.54	[-0.04, 1.12]	[-1.55, 2.63]	88.6	0.64	*92.1*	
average	24	1.38***	[1.06, 1.70]	[-0.12, 2.88]	198.5	0.50	88.4	6.1*
bilateral	21	1.37***	[1.01, 1.74]	[-0.26, 3.00]	163.6	0.57	87.8	
unilateral	3	2.11***	[1.58, 2.64]	[-1.31, 5.53]	0.2	0.00	0.0	30.5***
vertical	8	0.41*	[0.05, 0.77]	[-0.84, 1.66]	47.2	0.23	85.2	
two-choice	5	0.05	[-0.53, 0.62]	[-2.18, 2.27]	51.9	0.40	92.3	
go/ no-go	5	0.81***	[0.56, 1.06]	**[0.21, 1.40]**	4.7	0.02	15.4	22.6***
detection	22	1.52***	[1.20, 1.85]	**[0.11, 2.94]**	136.6	0.43	84.6	

Note. Model = Model of subgroup analysis for different experimental features: (1) reaction times for single condition: faster- or average-procedure; (2) target configuration in redundant condition: bilateral, unilateral, or vertical; (3) experimental paradigm: two-choice, go/ no-go, or detection. k = number of included effects; g = estimate of summary effect size based on Hedges' g; 95% CI = 95% confidence interval of the summary effect size; 95% PI = 95% prediction interval of the summary effect size; 95% PIs excluding zero are highlighted in bold; Q = Q statistic for statistical heterogeneity; τ^2 = estimate of the between-study variance; I^2 = percentage of the observed variance which is due to real differences in effect sizes; Q-G = Q statistic for subgroup differences; * p<.05. ** p<.01. ***p<.001.

Table E.3

Studies defined as outliers based on the influence analysis within the RVC-combination subset

r	Study	Influence Analysis	DIFFITS	Cook's distance	Covariance Ratio	Baujat plot	Outside 95%-CI
	Savazzi, et al., 2004, (Exp. 1)	X	1.00*	0.57*	0.51*		above
0.94	Omura et al., 2004		-0.38	0.13	0.97	X	below
	Corballis, 2002		0.23	0.05	1.08	X	above
	Savazzi, et al., 2004, (Exp. 1)		0.55	0.28	0.80		above
0.74	Omura et al., 2004		-0.44	0.16	0.89	X	below
	Corballis, 2002		0.42	0.16	0.97	X	above
	Savazzi, et al., 2004, (Exp. 1)		0.38	0.14	0.91		above
0.54	Omura et al., 2004		-0.48	0.18	0.90	X	below
	Corballis, 2002		0.57	0.27	0.92	X	above

Note. r = correlation coefficient used to estimate the standard error of the effect size for each experiment; Influence Analysis = Studies marked as outlier by *InfluenceAnalysis* function (Harrer et al., 2019a) are indicated by *X*. DIFFITS = How much the predicted pooled effect changes after excluding the study in standard deviations. Cook's distance = Distance between the fitted values of all *k* studies by including versus excluding the study. Covariance Ratio = Ratio of the variance-covariance matrix of parameter estimates with excluded versus included study. A value below 1 means that removing this RTE-test leads to a more precise estimate of the summary effect size. Baujat plot = Plot shows the contribution of each study to the heterogeneity in the meta-analysis. Outliers are indicated by *X*. Outside 95%-CI = Indicates if study is above or below the 95% confidence interval of the pooled effect. Exp. = experiment. *outside the cutoff suggested by Viechtbauer and Cheung (2010).

Figure E.1 [please insert figure E.1 as 2-column fitting image]

Funnel plots for the random effects model with outlier correction of the RVC-combination subset
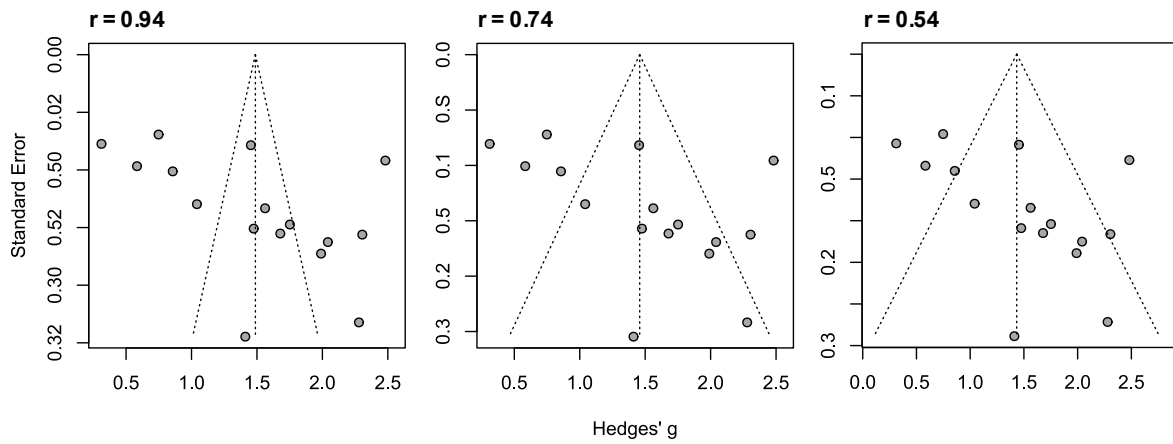
Note. The funnel plots show the standard error on the inverted y-axis and Hedges' *g* on the x-axis. Each dot represents one study. The inverted funnel is centered on the summary effect size of the redundant target effect. The RVC-combination subset includes all effects using the mean single reaction times of a detection paradigm with bilateral-redundant stimulation.

Appendix F. Analysis of target characteristics

One main finding of the meta-analysis is the high amount of heterogeneity between the effect sizes of studies. Hence, we investigated further potential sources of heterogeneity that were independent of our focus on RVC-research. In particular, we analysed the following target characteristics: shape, presentation duration, size, and eccentricity. Studies that did not yield sufficient information on certain target characteristics were excluded from the corresponding analysis.

Target shape and duration were categorical variables analysed via subgroup analysis (see explanations in section 2.2. Meta-analytical procedure and in appendix D).

Target size and eccentricity were continuous variables analysed via meta-regression analysis (Harrer et al., 2019). This was done using the *metareg* function of the R package *meta* (Schwarzer et al., 2015). Analogous to general regression models, meta-regression tries to predict the study's effect size by a certain factor. The regression line is fitted using the weighted least squares method which means that studies with a smaller standard error, i.e., better estimators for the true effect size, are weighted higher. The overall fit of the regression model can be evaluated using R_*^2. R_*^2 indicates how much additional heterogeneity variance (in percent) was explained by the regression model with the predictor compared to the default model containing only the summary effect size. Next, it is tested whether the regression weight of the predictor is significant using a *z*-statistic (Wald-type test; Harrer et al., 2019).

Each analysis was conducted three times, once for each potential value of the correlation coefficient: $r = 0.94$, $r = 0.74$ and $r = 0.54$.

Target shape

Studies used a variety of different target types that were categorized into three groups: `letters` (`falsefont` stimuli of Murray et al. (2001) were also categorized as letters), `rectangular targets` (squares, rectangles), or `circular targets` (disks, LED lights, circles, dots).

Results of the subgroup analysis showed that the summary effect size differed significantly between target shapes (except for $r = 0.54$; see table F.1 and figure F.1). Rectangular targets led to the highest summary effect size followed by circular targets. Letter targets had a considerably lower summary effects size. All 95%-CIs excluded zero. For rectangular targets and $r = 0.54$, the 95%-PI excluded zero. All other 95%-PIs included zero meaning that we have to expect null effects in future studies using similar designs. Importantly, from the ten studies investigating letter targets, five studies were administered by the same research group (Grice and colleagues) and hence might not have been independent from each other.

In conclusion, rectangular and circular targets lead to higher effect sizes than letter targets.

Table F.1

Results of subgroup analysis for target shape

r	model	k	g	95% CI	95% PI	Q	τ^2	I^2	Q-G
	circular	8	1.44***	[0.82, 2.06]	[-0.86, 3.73]	814.2	0.78	99.1	
0.94	letter	10	0.69*	[0.16, 1.22]	[-1.37, 2.76]	770.8	0.73	98.8	6.7*
	rectangular	14	1.65***	[1.11, 2.19]	[-0.63, 3.93]	441.4	1.02	97.1	
	circular	8	1.39***	[0.75, 2.03]	[-0.87, 3.65]	187.9	0.75	96.3	
0.74	letter	10	0.67*	[0.15, 1.20]	[-1.32, 2.67]	177.9	0.68	94.9	6.3*
	rectangular	14	1.52***	[1.10, 1.94]	[-0.15, 3.18]	101.9	0.54	87.2	
	circular	8	1.35***	[0.69, 2.01]	[-0.90, 3.60]	106.2	0.73	93.4	
0.54	letter	10	0.65*	[0.14, 1.17]	[-1.28, 2.59]	100.5	0.63	91.0	5.8
	rectangular	14	1.43***	[1.04, 1.82]	**[0.02, 2.85]**	57.6	0.38	77.4	

Note. Model = Model of subgroup analysis for target shape. k = number of included effects; g = estimate of summary effect size based on Hedges' g; 95% CI = 95% confidence interval of the summary effect size; 95% PI = 95% prediction interval of the summary effect size; 95% PIs excluding zero are highlighted in bold; Q = Q statistic for statistical heterogeneity; τ^2 = estimate of the variance between RTE-tests; I^2 = percentage of the observed variance which is due to real differences in effect sizes; Q-G = Q statistic for subgroup differences; * p<.05. ** p<.01. ***p<.001.

Figure F.1 [please insert figure F.1 as 2-column fitting image]

Results of the random effects model (r = 0.94) for the subset of studies yielding sufficient information on target shape

Study	N	RG	Shape		g	95% CI	weight
Tomaiuolo et al., 1997 (Exp. 2)	4	13.40	Circular		2.47	[1.79; 3.15]	2.8%
Tomaiuolo et al., 1997 (Exp. 2)	4	12.40	Circular		1.41	[0.93; 1.89]	3.0%
Schärli et al., 1999	22	15.00	Circular		2.08	[1.83; 2.34]	3.1%
Corballis, 2002	58	16.40	Circular		2.48	[2.30; 2.66]	3.2%
Roser et al., 2002	14	13.60	Circular		1.75	[1.46; 2.04]	3.1%
Schröter et al., 2011	16	7.00	Circular		0.86	[0.66; 1.06]	3.2%
Yu et al., 2014 (Exp. 1)	57	25.90	Circular		0.38	[0.28; 0.47]	3.2%
Yu et al., 2014 (Exp. 2)	128	21.88	Circular		0.23	[0.17; 0.29]	3.2%
Grice et al., 1984 (Exp. 1)	28	-16.00	Letter		-0.56	[-0.70; -0.42]	3.2%
Grice et al., 1984 (Exp. 2)	28	-10.00	Letter		-0.71	[-0.86; -0.57]	3.2%
Van der Heijden et al., 1984	24	11.00	Letter		1.09	[0.91; 1.26]	3.2%
Grice et al., 1987 (Exp. 5)	28	7.00	Letter		0.35	[0.22; 0.48]	3.2%
Grice et al., 1990 (Exp. 1)	28	13.00	Letter		0.67	[0.53; 0.82]	3.2%
Grice et al., 1992 (Exp. 1 & 2)	30	16.50	Letter		0.54	[0.41; 0.68]	3.2%
Mordkoff et al., 1996 (Exp. 2)	12	15.00	Letter		1.08	[0.83; 1.32]	3.1%
Murray et al., 2001	15	11.75	Letter		2.07	[1.76; 2.38]	3.1%
Murray et al., 2001	15	8.75	Letter		1.56	[1.30; 1.82]	3.1%
Ben-David et al., 2014	44	6.00	Letter		0.89	[0.77; 1.01]	3.2%
Miniussi et al., 1998	12	9.00	Rectangular		1.99	[1.65; 2.33]	3.1%
Omura et al., 2004	21	18.22	Rectangular		0.31	[0.16; 0.46]	3.2%
Savazzi et al., 2004 (Exp. 1)	8	22.70	Rectangular		4.88	[4.02; 5.74]	2.7%
Savazzi et al., 2004 (Exp. 2)	8	29.10	Rectangular		2.28	[1.82; 2.73]	3.0%
Miller et al., 2006	14	27.00	Rectangular		2.04	[1.72; 2.36]	3.1%
Miller, 2007	16	20.00	Rectangular		1.78	[1.51; 2.05]	3.1%
Miller & Van Nes, 2007 (Exp. 1)	40	25.75	Rectangular		1.45	[1.30; 1.61]	3.2%
Fischer et al., 2008	32	17.00	Rectangular		0.75	[0.61; 0.88]	3.2%
Florio et al., 2008 (Exp. 2)	18	16.92	Rectangular		2.30	[2.00; 2.61]	3.1%
Mooshagian et al., 2008	15	17.18	Rectangular		0.58	[0.39; 0.77]	3.2%
Tamietto et al., 2010 (Exp. 1)	11		Rectangular		1.04	[0.79; 1.30]	3.1%
Railo et al., 2014	11		Rectangular		1.04	[0.79; 1.30]	3.1%
Van Koningsbruggen et al., 2017 (1. CG)	11	24.30	Rectangular		1.48	[1.18; 1.77]	3.1%
Van Koningsbruggen et al., 2017 (2. CG)	12	33.00	Rectangular		1.68	[1.37; 1.98]	3.1%
Overall effect					1.29	[0.95; 1.64]	100.0%
Prediction interval						[-0.76; 3.35]	

Heterogeneity: I^2 = 99%, $p = 0$

-5   0   5   10

Note. The forest plot shows the effect size *g* with its associated 95% confidence interval for each included RTE-test. RTE-tests are sorted by target shape. The summary effect size is represented by the diamond with the corresponding dashed line. The black bar at the bottom represents the 95% prediction interval of the summary effect size. N = Number of participants; RG = redundancy gain in milliseconds; redundancy gain is the difference in reaction times between redundant and single condition; *g* = estimation of effect size based on Hedges' *g*; 95% CI = 95% confidence interval of Hedges' *g* based on the calculation of the standard error with *r* = 0.94; weight = relative weight of each included effect; Exp. = experiment, CG = control group.

Target duration

Target duration was analysed in a subgroup analysis consisting of three groups. Most studies used a pre-defined target duration ranging from 32 to 200ms. These were grouped into `Below 100`, i.e., the target was presented for less than 100ms, and `100-200`, i.e., the target was presented for 100 to 200ms. The third group `until response` included studies in which the target was presented until the participant responded or until a maximum time was reached.

Results of the subgroup analysis showed that the summary effect size differed significantly between the categories of target duration (see table F.2 and figure F.2). The summary effect size for studies in which targets were shown for 100-200ms or until the response were similar. If targets were shown for less than 100ms, the summary effect size was considerably higher. The 95%-PI was above zero for targets shown less than 100ms at $r = 0.54$. All other 95%-PIs included zero meaning that we have to expect null effects in future studies using similar designs.

In conclusion, the highest summary effect size was reached with targets presented for less than 100ms.

Table F.2

Results of subgroup analysis for target duration

r	model	*k*	*g*	95% CI	95% PI	Q	τ^2	I^2	Q-G
	below 100	10	1.99***	[1.32, 2.67]	[-0.57, 4.56]	220.9	1.12	95.9	
0.94	**100-200**	10	0.95**	[0.37, 1.54]	[-1.31, 3.22]	947.5	0.88	99.1	8.7*
	until response	7	0.71*	[0.14, 1.29]	[-1.42, 2.85]	574.9	0.61	99.0	
	below 100	10	1.84***	[1.30, 2.37]	[-0.01, 3.68]	51.0	0.56	82.3	
0.74	**100-200**	10	0.94**	[0.35, 1.52]	[-1.29, 3.16]	218.7	0.84	95.9	9.1*
	until response	7	0.70*	[0.13, 1.27]	[-1.38, 2.78]	132.7	0.57	95.5	
	below 100	10	1.73***	[1.23, 2.22]	**[0.22, 3.24]**	28.8	0.37	68.8	
0.54	**100-200**	10	0.92**	[0.34, 1.51]	[-1.25, 3.11]	123.6	0.81	92.7	8.4*
	until response	7	0.68*	[0.12, 1.25]	[-1.33, 2.70]	75.0	0.53	92.0	

Note. Model = Model of subgroup analysis for target duration. k = number of included effects; g = estimate of summary effect size based on Hedges' g; 95% CI = 95% confidence interval of the summary effect size; 95% PI = 95% prediction interval of the summary effect size; 95% PIs excluding zero are highlighted in bold; Q = Q statistic for statistical heterogeneity; τ^2 = estimate of the variance between RTE-tests; I^2 = percentage of the observed variance which is due to real differences in effect sizes; Q-G = Q statistic for subgroup differences; * p<.05. ** p<.01. ***p<.001.

Figure F.2 [please insert figure F.2 as 2-column fitting image]

Results of the random effects model (r = 0.94) for the subset of studies yielding sufficient information on target duration

Study	N	RG	Duration		g	95% CI	weight
Tomaiuolo et al., 1997 (Exp. 2)	4	13.40	Below 100		2.47	[1.79; 3.15]	3.4%
Tomaiuolo et al., 1997 (Exp. 2)	4	12.40	Below 100		1.41	[0.93; 1.89]	3.6%
Miniussi et al., 1998	12	9.00	Below 100		1.99	[1.65; 2.33]	3.7%
Schärli et al., 1999	22	15.00	Below 100		2.08	[1.83; 2.34]	3.7%
Murray et al., 2001	15	11.75	Below 100		2.07	[1.76; 2.38]	3.7%
Murray et al., 2001	15	8.75	Below 100		1.56	[1.30; 1.82]	3.7%
Savazzi et al., 2004 (Exp. 1)	8	22.70	Below 100		4.88	[4.02; 5.74]	3.2%
Savazzi et al., 2004 (Exp. 2)	8	29.10	Below 100		2.28	[1.82; 2.73]	3.6%
Mooshagian et al., 2008	15	17.18	Below 100		0.58	[0.39; 0.77]	3.8%
Railo et al., 2014	11		Below 100		1.04	[0.79; 1.30]	3.7%
Grice et al., 1984 (Exp. 2)	28	-10.00	100-200		-0.71	[-0.86; -0.57]	3.8%
Van der Heijden et al., 1984	24	11.00	100-200		1.09	[0.91; 1.26]	3.8%
Grice et al., 1987 (Exp. 5)	28	7.00	100-200		0.35	[0.22; 0.48]	3.8%
Grice et al., 1990 (Exp. 1)	28	13.00	100-200		0.67	[0.53; 0.82]	3.8%
Grice et al., 1992 (Exp. 1 & 2)	30	16.50	100-200		0.54	[0.41; 0.68]	3.8%
Corballis, 2002	58	16.40	100-200		2.48	[2.30; 2.66]	3.8%
Roser et al., 2002	14	13.60	100-200		1.75	[1.46; 2.04]	3.7%
Omura et al., 2004	21	18.22	100-200		0.31	[0.16; 0.46]	3.8%
Miller et al., 2006	14	27.00	100-200		2.04	[1.72; 2.36]	3.7%
Tamietto et al., 2010 (Exp. 1)	11		100-200		1.04	[0.79; 1.30]	3.7%
Grice et al., 1984 (Exp. 1)	28	-16.00	Until response		-0.56	[-0.70; -0.42]	3.8%
Miller, 2007	16	20.00	Until response		1.78	[1.51; 2.05]	3.7%
Miller & Van Nes, 2007 (Exp. 1)	40	25.75	Until response		1.45	[1.30; 1.61]	3.8%
Schröter et al., 2011	16	7.00	Until response		0.86	[0.66; 1.06]	3.8%
Ben-David et al., 2014	44	6.00	Until response		0.89	[0.77; 1.01]	3.8%
Yu et al., 2014 (Exp. 1)	57	25.90	Until response		0.38	[0.28; 0.47]	3.8%
Yu et al., 2014 (Exp. 2)	128	21.88	Until response		0.23	[0.17; 0.29]	3.8%
Overall effect					1.27	[0.86; 1.67]	100.0%
Prediction interval						[-0.96; 3.49]	
Heterogeneity: I^2 = 99%, p = 0							

-5 0 5 10

Note. The forest plot shows the effect size *g* with its associated 95% confidence interval for each included RTE-test. RTE-tests are sorted by target duration. The summary effect size is represented by the diamond with the corresponding dashed line. The black bar at the bottom represents the 95% prediction interval of the summary effect size. N = Number of participants; RG = redundancy gain in milliseconds; redundancy gain is the difference in reaction times between redundant and single condition; *g* = estimation of effect size based on Hedges' *g*; 95% CI = 95% confidence interval of Hedges' *g* based on the calculation of the standard error with *r* = 0.94; weight = relative weight of each included effect; Exp. = experiment, CG = control group.

Target size

To get a comparable target size for letters, rectangular, and circular targets, we calculated the target area in square degree visual angle (deg^2). For letters and rectangular targets, we calculated the target area as the vertical size multiplied by the horizontal size. For circular targets, the size was indicated as the diameter. Hence, we calculated the target area using the formula: *pi*(diameter/2)^2*.

Target area ranged from 0.05 to 25.00 deg^2. However, there were two outliers with a target area of 10 deg^2 (Miniussi et al., 1998) and 25 deg^2 (Tamietto et al., 2010). All other studies had a target area below 2.6 deg^2. Hence, we excluded Miniussi et al. (1998) as well as Tamietto et al. (2010) for the meta-regression.

Results of the meta-regression showed that target size did not predict the study's effect size. (see table F.3 and figure F.3). The regression weight of the predictor was non-significant and no additional heterogeneity variance (R^2_*) was explained by the regression model compared to the default model without the predictor target size. This was true for all values of *r*.
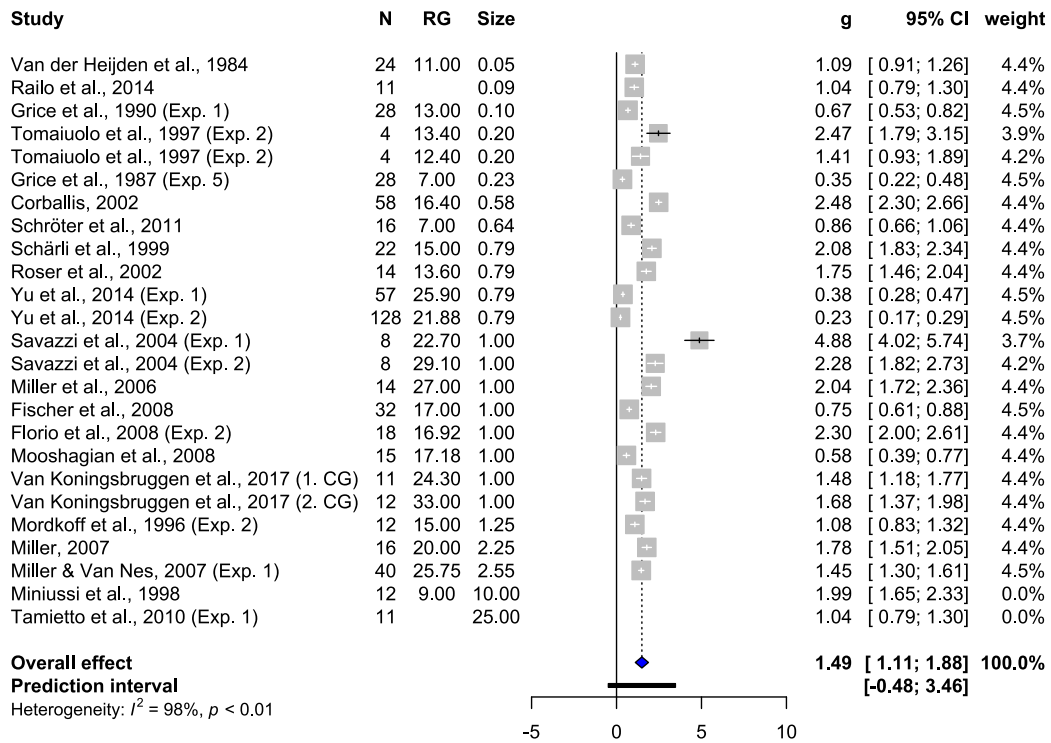
Table F.3

Results of the meta-regression for target size

r	*k*	intercept	β	*z*	*p*	R^2_*	Q	τ^2	I^2
0.94	23	1.28***	0.26	0.79	0.432	0.00	1338.4***	0.88	99.1
0.74	23	1.17***	0.26	0.96	0.339	0.00	308.9***	0.56	94.1
0.54	23	1.10***	0.27	1.04	0.301	1.17	174.6***	0.46	88.1

Note. *r* = correlation coefficient used to estimate the standard error of the effect size for each RTE-test; *k* = number of included RTE-tests; intercept = expected summary effect size (based on Hedges' *g*) when the predictor is zero; β = regression weight of the predictor; *z* = z-value of the significance test for the regression weight; *p* = p-value of the significance test for the regression weight; R^2_* = percentage of heterogeneity variance that is explained by the predictor; Q = Q statistic for statistical heterogeneity; τ^2 = unexplained heterogeneity variance; I^2 = percentage of the observed variance which is due to real differences in effect sizes; * p<.05. ** p<.01. ***p<.001.

Figure F.3 [please insert figure F.3 as 2-column fitting image]

Results of the random effects model (r = 0.94) for the subset of studies yielding sufficient information on target size

Study	N	RG	Size	g	95% CI	weight
Van der Heijden et al., 1984	24	11.00	0.05	1.09	[0.91; 1.26]	4.4%
Railo et al., 2014	11		0.09	1.04	[0.79; 1.30]	4.4%
Grice et al., 1990 (Exp. 1)	28	13.00	0.10	0.67	[0.53; 0.82]	4.5%
Tomaiuolo et al., 1997 (Exp. 2)	4	13.40	0.20	2.47	[1.79; 3.15]	3.9%
Tomaiuolo et al., 1997 (Exp. 2)	4	12.40	0.20	1.41	[0.93; 1.89]	4.2%
Grice et al., 1987 (Exp. 5)	28	7.00	0.23	0.35	[0.22; 0.48]	4.5%
Corballis, 2002	58	16.40	0.58	2.48	[2.30; 2.66]	4.4%
Schröter et al., 2011	16	7.00	0.64	0.86	[0.66; 1.06]	4.4%
Schärli et al., 1999	22	15.00	0.79	2.08	[1.83; 2.34]	4.4%
Roser et al., 2002	14	13.60	0.79	1.75	[1.46; 2.04]	4.4%
Yu et al., 2014 (Exp. 1)	57	25.90	0.79	0.38	[0.28; 0.47]	4.5%
Yu et al., 2014 (Exp. 2)	128	21.88	0.79	0.23	[0.17; 0.29]	4.5%
Savazzi et al., 2004 (Exp. 1)	8	22.70	1.00	4.88	[4.02; 5.74]	3.7%
Savazzi et al., 2004 (Exp. 2)	8	29.10	1.00	2.28	[1.82; 2.73]	4.2%
Miller et al., 2006	14	27.00	1.00	2.04	[1.72; 2.36]	4.4%
Fischer et al., 2008	32	17.00	1.00	0.75	[0.61; 0.88]	4.5%
Florio et al., 2008 (Exp. 2)	18	16.92	1.00	2.30	[2.00; 2.61]	4.4%
Mooshagian et al., 2008	15	17.18	1.00	0.58	[0.39; 0.77]	4.4%
Van Koningsbruggen et al., 2017 (1. CG)	11	24.30	1.00	1.48	[1.18; 1.77]	4.4%
Van Koningsbruggen et al., 2017 (2. CG)	12	33.00	1.00	1.68	[1.37; 1.98]	4.4%
Mordkoff et al., 1996 (Exp. 2)	12	15.00	1.25	1.08	[0.83; 1.32]	4.4%
Miller, 2007	16	20.00	2.25	1.78	[1.51; 2.05]	4.4%
Miller & Van Nes, 2007 (Exp. 1)	40	25.75	2.55	1.45	[1.30; 1.61]	4.5%
Miniussi et al., 1998	12	9.00	10.00	1.99	[1.65; 2.33]	0.0%
Tamietto et al., 2010 (Exp. 1)	11		25.00	1.04	[0.79; 1.30]	0.0%
Overall effect				1.49	[1.11; 1.88]	100.0%
Prediction interval					[-0.48; 3.46]	

Heterogeneity: I^2 = 98%, p < 0.01

Note. The forest plot shows the effect size *g* with its associated 95% confidence interval for each included RTE-test. RTE-tests are sorted by their value of target size. The summary effect size is represented by the diamond with the corresponding dashed line. The black bar at the bottom represents the 95% prediction interval of the summary effect size. N = Number of participants; RG = redundancy gain in milliseconds; redundancy gain is the difference in reaction times between redundant and single condition; *g* = estimation of effect size based on Hedges' *g*; 95% CI = 95% confidence interval of Hedges' *g* based on the calculation of the standard error with *r* = 0.94; weight = relative weight of each included effect; Studies that are excluded have a weight of 0.0%. Exp. = experiment, CG = control group.

Target eccentricity

We defined target eccentricity as the distance between the centre of the target to the centre of the fixation symbol in degree visual angle. If studies gave a value on eccentricity but no detailed description on the exact calculation of the distance, we still kept the stated value as the presumed eccentricity.

Results of the meta-regression showed that target eccentricity did not predict the study's effect size significantly (see table F.4 and figure F.4). There was a tendency (p <.10) that the effect size increased with increasing eccentricity. This was true for all values of *r*. 8.1-10.6% of additional heterogeneity variance (R_*^2) was explained by the regression model compared to the default model without the predictor target eccentricity.

In conclusion, there was a tendency that the summary effect size increased with increasing eccentricity.

Table F.4

Results of the meta-regression for target eccentricity

r	*k*	intercept	β	*z*	*p*	R_*^2	Q	τ^2	I^2
0.94	25	0.69	0.13	1.72	0.085	8.1	2005.1***	1.04	99.3
0.74	25	0.65	0.12	1.85	0.064	10.5	462.7***	0.76	95.9
0.54	25	0.62	0.12	1.87	0.061	10.6	261.5***	0.66	92.5

Note. r = correlation coefficient used to estimate the standard error of the effect size for each RTE-test; *k* = number of included RTE-tests; intercept = expected summary effect size (based on Hedges' *g*) when the predictor is zero; β = regression weight of the predictor; *z* = z-value of the significance test for the regression weight; *p* = p-value of the significance test for the regression weight; R_*^2 = percentage of heterogeneity variance that is explained by the predictor; Q = Q statistic for statistical heterogeneity; τ^2 = unexplained heterogeneity variance; I^2 = percentage of the observed variance which is due to real differences in effect sizes; * p<.05. ** p<.01. ***p<.001.

Figure F.4 [please insert figure F.4 as 2-column fitting image]

Results of the random effects model (r = 0.94) for the subset of studies yielding sufficient information on target eccentricity

Study	N	RG	Eccentricity		g	95% CI	weight
Van der Heijden et al., 1984	24	11.00	0.69		1.09	[0.91; 1.26]	4.0%
Schröter et al., 2011	16	7.00	0.90		0.86	[0.66; 1.06]	4.0%
Grice et al., 1984 (Exp. 2)	28	-10.00	1.50		-0.71	[-0.86; -0.57]	4.1%
Grice et al., 1987 (Exp. 5)	28	7.00	1.50		0.35	[0.22; 0.48]	4.1%
Grice et al., 1990 (Exp. 1)	28	13.00	1.50		0.67	[0.53; 0.82]	4.1%
Grice et al., 1992 (Exp. 1 & 2)	30	16.50	1.50		0.54	[0.41; 0.68]	4.1%
Mordkoff et al., 1996 (Exp. 2)	12	15.00	1.53		1.08	[0.83; 1.32]	4.0%
Murray et al., 2001	15	11.75	2.40		2.07	[1.76; 2.38]	4.0%
Murray et al., 2001	15	8.75	2.40		1.56	[1.30; 1.82]	4.0%
Railo et al., 2014	11		2.50		1.04	[0.79; 1.30]	4.0%
Grice et al., 1984 (Exp. 1)	28	-16.00	3.00		-0.56	[-0.70; -0.42]	4.1%
Corballis, 2002	58	16.40	5.00		2.48	[2.30; 2.66]	4.0%
Savazzi et al., 2004 (Exp. 1)	8	22.70	5.50		4.88	[4.02; 5.74]	3.5%
Savazzi et al., 2004 (Exp. 2)	8	29.10	5.50		2.28	[1.82; 2.73]	3.9%
Miller et al., 2006	14	27.00	6.00		2.04	[1.72; 2.36]	4.0%
Yu et al., 2014 (Exp. 1)	57	25.90	6.00		0.38	[0.28; 0.47]	4.1%
Yu et al., 2014 (Exp. 2)	128	21.88	6.00		0.23	[0.17; 0.29]	4.1%
Miniussi et al., 1998	12	9.00	6.10		1.99	[1.65; 2.33]	4.0%
Miller & Van Nes, 2007 (Exp. 1)	40	25.75	6.80		1.45	[1.30; 1.61]	4.1%
Miller, 2007	16	20.00	7.00		1.78	[1.51; 2.05]	4.0%
Florio et al., 2008 (Exp. 2)	18	16.92	7.50		2.30	[2.00; 2.61]	4.0%
Van Koningsbruggen et al., 2017 (1. CG)	11	24.30	8.00		1.48	[1.18; 1.77]	4.0%
Van Koningsbruggen et al., 2017 (2. CG)	12	33.00	8.00		1.68	[1.37; 1.98]	4.0%
Fischer et al., 2008	32	17.00	8.10		0.75	[0.61; 0.88]	4.1%
Tamietto et al., 2010 (Exp. 1)	11		10.13		1.04	[0.79; 1.30]	4.0%
Overall effect					1.28	[0.87; 1.70]	100.0%
Prediction interval						[-0.95; 3.52]	

Heterogeneity: $I^2 = 99\%$, $p = 0$

(x-axis: -5 0 5 10)

Note. The forest plot shows the effect size *g* with its associated 95% confidence interval for each included RTE-test. RTE-tests are sorted by their value of target eccentricity. The summary effect size is represented by the diamond with the corresponding dashed line. The black bar at the bottom represents the 95% prediction interval of the summary effect size. N = Number of participants; RG = redundancy gain in milliseconds; redundancy gain is the difference in reaction times between redundant and single condition; *g* = estimation of effect size based on Hedges' *g*; 95% CI = 95% confidence interval of Hedges' *g* based on the calculation of the standard error with *r* = 0.94; weight = relative weight of each included effect. Exp. = experiment, CG = control group.

References

Ajina, S., & Bridge, H. (2018). Blindsight relies on a functional connection between hMT+ and the lateral geniculate nucleus, not the pulvinar. *PLOS Biology, 16*(7), e2005769. https://doi.org/10.1371/journal.pbio.2005769

Ajina, S., Jünemann, K., Sahraie, A., & Bridge, H. (2021). Increased Visual Sensitivity and Occipital Activity in Patients With Hemianopia Following Vision Rehabilitation. *Journal of Neuroscience, 41*(28), 5994-6005. https://doi.org/10.1523/JNEUROSCI.2790-20.2021

Ajina, S., Pestilli, F., Rokem, A., Kennard, C., & Bridge, H. (2015). Human blindsight is mediated by an intact geniculo-extrastriate pathway. *Elife, 4*(e08935). https://doi.org/10.7554/eLife.08935

Ajina, S., Pollard, M., & Bridge, H. (2020). The Superior Colliculus and Amygdala Support Evaluation of Face Trait in Blindsight. *Frontiers in Neurology, 11*(769). https://doi.org/10.3389/fneur.2020.00769

Azzopardi, P., & Cowey, A. (1998). Blindsight and visual awareness. *Consciousness and Cognition, 7*(3), 292-311. https://doi.org/10.1006/ccog.1998.0358

Barbur, J. L., Ruddock, K. H., & Waterfield, V. A. (1980). Human visual responses in the absence of the geniculo-calcarine projection. *Brain: a Journal of Neurology, 103*(4), 905-928. https://doi.org/10.1093/brain/103.4.905

Baujat, B., Mahé, C., Pignon, J. P., & Hill, C. (2002). A graphical method for exploring heterogeneity in meta-analyses: application to a meta-analysis of 65 trials. *Statistics in Medicine, 21*(18), 2641-2652. https://doi.org/https://doi.org/10.1002/sim.1221

Ben-David, B. M., Eidels, A., & Donkin, C. (2014). Effects of aging and distractors on detection of redundant visual targets and capacity: Do older adults integrate visual targets differently than younger adults? *PLoS One, 9*(12), e113551. https://doi.org/10.1371/journal.pone.0113551

Bertini, C., Cecere, R., & Làdavas, E. (2013). I am blind, but I "see" fear. *Cortex, 49*(4), 985-993. https://doi.org/10.1016/j.cortex.2012.02.006

Bola, M., Gall, C., & Sabel, B. A. (2013). "Sightblind": perceptual deficits in the "intact" visual field. *Frontiers in Neurology, 4*(80), 1-5. https://doi.org/10.3389/fneur.2013.00080

Borenstein, M., Hedges, L. V., Higgins, J. P. T., & Rothstein, H. R. (2009). *Introduction to meta-analysis* (1 ed.). Wiley.

Button, K. S., Ioannidis, J. P. A., Mokrysz, C., Nosek, B. A., Flint, J., Robinson, E. S. J., & Munafo, M. R. (2013). Power failure: why small sample size undermines the reliability of neuroscience. *Nature Reviews Neuroscience, 14*, 365-376. https://doi.org/10.1038/nrn3475

Cavanaugh, M. R., & Huxlin, K. R. (2017). Visual discrimination training improves Humphrey perimetry in chronic cortically induced blindness. *Neurology, 88*(19), 1856-1864. https://doi.org/10.1212/wnl.0000000000003921

Celeghin, A., de Gelder, B., & Tamietto, M. (2015). From affective blindsight to emotional consciousness. *Consciousness and Cognition, 36*, 414-425. https://doi.org/10.1016/j.concog.2015.05.007

Celeghin, A., Savazzi, S., Barabas, M., Bendini, M., & Marzi, C. A. (2015). Blindsight is sensitive to stimulus numerosity and configuration: evidence from the redundant signal effect. *Experimental Brain Research, 233*(5), 1617-1623. https://doi.org/10.1007/s00221-015-4236-6

Cooper, H. M., Hedges, L. V., & Valentine, J., C. (2009). *Handbook of Research Synthesis and Meta-Analysis* (2 ed.). Russel Sage Foundation.

Corballis, M. C. (1998). Interhemispheric neural summation in the absence of the corpus callosum. *Brain*, *121*(9), 1795-1807. https://doi.org/10.1093/brain/121.9.1795

Corballis, M. C. (2002). Hemispheric interactions in simple reaction time. *Neuropsychologia*, *40*(4), 423-434. https://doi.org/10.1016/S0028-3932(01)00097-5

Corballis, M. C., Corballis, P. M., & Fabri, M. (2004). Redundancy gain in simple reaction time following partial and complete callosotomy. *Neuropsychologia*, *42*(1), 71-81. https://doi.org/10.1016/s0028-3932(03)00152-0

Corballis, M. C., Corballis, P. M., Fabri, M., Paggi, A., & Manzoni, T. (2005). Now you see it, now you don't: Variable hemineglect in a commissurotomized man. *Cognitive Brain Research*, *25*(2), 521-530. https://doi.org/10.1016/j.cogbrainres.2005.08.002

Corballis, M. C., Hamm, J. P., Barnett, K. J., & Corballis, P. M. (2002). Paradoxical interhemispheric summation in the split brain. *Journal of Cognitive Neuroscience*, *14*(8), 1151-1157. https://doi.org/10.1162/089892902760807168.

Corbetta, M., Marzi, C. A., Tassinari, G., & Aglioti, S. (1990). Effectiveness of different task paradigms in revealing blindsight. *Brain*, *113*(3), 603-616. https://doi.org/10.1093/brain/113.3.603

Cowey, A. (2010). The blindsight saga. *Experimental Brain Research*, *200*, 3-24. https://doi.org/10.1007/s00221-009-1914-2

Cowey, A., & Weiskrantz, L. (1963). A perimetric study of visual field defects in monkeys. *Quarterly Journal of Experimental Psychology*, *15*(2), 91-115. https://doi.org/10.1080/17470216308416561

Danckert, J., & Rossetti, Y. (2005). Blindsight in action: what can the different sub-types of blindsight tell us about the control of visually guided actions? *Neuroscience & Biobehavioral Reviews*, *29*(7), 1035-1046. https://doi.org/10.1016/j.neubiorev.2005.02.001

Danckert, J., Tamietto, M., & Rossetti, Y. (2019). Definition: Blindsight. *Cortex*, *119*, 569-570. https://doi.org/10.1016/j.cortex.2019.01.027

de Gelder, B., Pourtois, G., van Raamsdonk, M., Vroomen, J., & Weiskrantz, L. (2001). Unseen stimuli modulate conscious visual experience: evidence from inter-hemispheric summation. *Neuroreport*, *12*(2), 385-391.

Donkin, C., Little, D. R., & Houpt, J. W. (2014). Assessing the speed-accuracy trade-off effect on the capacity of information processing. *Journal of Experimental Psychology: Human Perception and Performance*, *40*(3), 1183-1202. https://doi.org/10.1037/a0035947

Driver, J., & Vuilleumier, P. (2001). Perceptual awareness and its loss in unilateral neglect and extinction. *Cognition*, *79*(1-2), 39-88. https://doi.org/10.1016/S0010-0277(00)00124-4

Duval, S., & Tweedie, R. (2000). Trim and fill: A simple funnel-plot–based method of testing and adjusting for publication bias in meta-analysis. *Biometrics*, *56*(2), 455-463. https://doi.org/10.1111/j.0006-341X.2000.00455.x

Egger, M., Smith, G. D., Schneider, M., & Minder, C. (1997). Bias in meta-analysis detected by a simple, graphical test. *BMJ*, *315*, 629-634. https://doi.org/10.1136/bmj.315.7109.629

Faul, F., Erdfelder, E., Buchner, A., & Lang, A.-G. (2009). Statistical power analyses using G*Power 3.1: tests for correlation and regression analyses. *Behavior Research Methods*, *41*, 1149-1160. https://doi.org/10.3758/BRM.41.4.1149

Fischer, R., & Miller, J. (2008). Differential redundancy gain in onset detection versus offset detection. *Perception & Psychophysics*, *70*(3), 431-436. https://doi.org/10.3758/pp.70.3.431

Florio, V., Marzi, C. A., Girelli, A., & Savazzi, S. (2008). Enhanced redundancy gain in schizophrenics: A correlate of callosal dysfunction? *Neuropsychologia*, *46*(11), 2808-2815. https://doi.org/10.1016/j.neuropsychologia.2008.05.007

Florio, V., Savazzi, S., Conca, A., & Marzi, C. A. (2013). Differential impairment of interhemispheric transmission in bipolar disease. *Experimental Brain Research*, *230*, 175-185. https://doi.org/10.1007/s00221-013-3642-x

Georgy, L., Celeghin, A., Marzi, C. A., Tamietto, M., & Ptito, A. (2016). The superior colliculus is sensitive to gestalt-like stimulus configuration in hemispherectomy patients. *Cortex*, *81*, 151-161. https://doi.org/10.1016/j.cortex.2016.04.018

Gleser, L. J., & Olkin, I. (2009). Stochastically Dependent Effect Sizes. In H. M. Cooper, L. V. Hedges, & J. Valentine, C. (Eds.), *Handbook of Research Synthesis and Meta-Analysis* (2 ed., pp. 357-376). Russel Sage Foundation.

Grasso, P. A., Pietrelli, M., Zanon, M., Làdavas, E., & Bertini, C. (2020). Alpha oscillations reveal implicit visual processing of motion in hemianopia. *Cortex*, *122*, 81-96. https://doi.org/10.1016/j.cortex.2018.08.009

Grice, G. R., & Canham, L. (1990). Redundancy phenomena are affected by response requirements. *Perception & Psychophysics*, *48*(3), 209-213. https://doi.org/10.3758/BF03211520

Grice, G. R., Canham, L., & Gwynne, J. W. (1984). Absence of a redundant-signals effect in a reaction time task with divided attention. *Perception & Psychophysics*, *36*(6), 565-570. https://doi.org/10.3758/BF03207517

Grice, G. R., & Gwynne, J. W. (1987). Dependence of target redundancy effects on noise conditions and number of targets. *Perception & Psychophysics*, *42*(1), 29-36. https://doi.org/10.3758/BF03211511

Grice, G. R., & Reed, J. M. (1992). What makes targets redundant? *Perception & Psychophysics*, *51*(5), 437-442. https://doi.org/10.3758/BF03211639

Hall, N., & Colby, C. (2014). S-cone visual stimuli activate superior colliculus neurons in old world monkeys: implications for understanding blindsight. *Journal of Cognitive Neuroscience*, *26*(6), 1234-1256. https://doi.org/10.1162/jocn_a_00555

Harrer, M., Cuijpers, P., Furukawa, T., & Ebert, D. D. (2019a). *dmetar: Companion R Package For The Guide 'Doing Meta-Analysis in R'*. In (Version R package version 0.0.9000.) http://dmetar.protectlab.org/

Harrer, M., Cuijpers, P., Furukawa, T., & Ebert, D. D. (2019b). *Doing Meta-Analysis in R: A Hands-on Guide*. Retrieved 30.03.2021 from https://bookdown.org/MathiasHarrer/Doing_Meta_Analysis_in_R/

Higgins, J. P., Thompson, S. G., Deeks, J. J., & Altman, D. G. (2003). Measuring inconsistency in meta-analyses. *BMJ*, *327*, 557-560. https://doi.org/10.1136/bmj.327.7414.557

Huxlin, K. R., Martin, T., Kelly, K., Riley, M., Friedman, D. I., Burgin, W. S., & Hayhoe, M. (2009). Perceptual relearning of complex visual motion after V1 damage in humans. *Journal of Neuroscience*, *29*(13), 3981-3991. https://doi.org/10.1523/JNEUROSCI.4882-08.2009

Iacoboni, M., Ptito, A., Weekes, N. Y., & Zaidel, E. (2000). Parallel visuomotor processing in the split brain: cortico-subcortical interactions. *Brain*, *123*(4), 759-769. https://doi.org/10.1093/brain/123.4.759

IntHout, J., Ioannidis, J. P. A., Rovers, M. M., & Goeman, J. J. (2016). Plea for routinely presenting prediction intervals in meta-analysis. *BMJ Open*, *6*(e010247), 1-6. https://doi.org/10.1136/bmjopen-2015-010247

Kasten, E., Wüst, S., Behrens-Baumann, W., & Sabel, B. A. (1998). Computer-based training for the treatment of partial blindness. *Nature medicine*, *4*(9), 1083-1087. https://doi.org/10.1038/2079

Kinchla, R. (1974). Detecting target elements in multielement arrays: A confusability model. *Perception & Psychophysics*, *15*(1), 149-158. https://doi.org/10.3758/BF03205843

Lakens, D. (2013). Calculating and reporting effect sizes to facilitate cumulative science: a practical primer for t-tests and ANOVAs. *Frontiers in Psychology*, *4*(863), 1-12. https://doi.org/10.3389/fpsyg.2013.00863

Leh, S. E., Mullen, K. T., & Ptito, A. (2006). Absence of S-cone input in human blindsight following hemispherectomy. *European Journal of Neuroscience*, *24*, 2954-2960. https://doi.org/10.1111/j.1460-9568.2006.05178.x

Marzi, C. A., Girelli, M., Miniussi, C., Smania, N., & Maravita, A. (2000). Electrophysiological correlates of conscious vision: evidence from unilateral extinction. *Journal of Cognitive Neuroscience*, *12*(5), 869-877. https://doi.org/10.1162/089892900562471

Marzi, C. A., Mancini, F., Metitieri, T., & Savazzi, S. (2009). Blindsight following visual cortex deafferentation disappears with purple and red stimuli: A case study. *Neuropsychologia*, *47*(5), 1382-1385. https://doi.org/10.1016/j.neuropsychologia.2009.01.023

Marzi, C. A., Smania, N., Martini, M. C., Gambina, G., Tomelleri, G., Palamara, A., Alessandrini, F., & Prior, M. (1996). Implicit redundant-targets effect in visual extinction. *Neuropsychologia*, *34*(1), 9-22. https://doi.org/10.1016/0028-3932(95)00059-3

Marzi, C. A., Tassinari, G., Aglioti, S., & Lutzemberger, L. (1986). Spatial summation across the vertical meridian in hemianopics: A test of blindsight. *Neuropsychologia*, *24*(6), 749-758. https://doi.org/10.1016/0028-3932(86)90074-6

Mazzi, C., Bagattini, C., & Savazzi, S. (2016). Blind-sight vs. degraded-sight: Different measures tell a different story. *Frontiers in Psychology*, *7*(901), 1-11. https://doi.org/10.3389/fpsyg.2016.00901

Melnick, M. D., Tadin, D., & Huxlin, K. R. (2016). Relearning to see in cortical blindness. *Neuroscientist*, *22*(2), 199-212. https://doi.org/10.1177/1073858415621035

Mikellidou, K., Arrighi, R., Aghakhanyan, G., Tinelli, F., Frijia, F., Crespi, S., De Masi, F., Montanaro, D., & Morrone, M. C. (2019). Plasticity of the human visual brain after an early cortical lesion. *Neuropsychologia*, *128*, 166-177. https://doi.org/10.1016/j.neuropsychologia.2017.10.033

Miller, J. (1978). Multidimensional same–different judgments: Evidence against independent comparisons of dimensions. *Journal of Experimental Psychology: Human Perception and Performance*, *4*(3), 411-422. https://doi.org/10.1037/0096-1523.4.3.411

Miller, J. (1982). Divided attention: Evidence for coactivation with redundant signals. *Cognitive Psychology*, *14*(2), 247-279. https://doi.org/10.1016/0010-0285(82)90010-X

Miller, J. (2007). Contralateral and ipsilateral motor activation in visual simple reaction time: A test of the hemispheric coactivation model. *Experimental Brain Research*, *176*(4), 539-558. https://doi.org/10.1007/s00221-006-0641-1

Miller, J., & Adam, J. J. (2006). Redundancy gain with static versus moving hands: A test of the hemispheric coactivation model. *Acta psychologica*, *122*(1), 1-10. https://doi.org/10.1016/j.actpsy.2005.08.003

Miller, J., & Van Nes, F. (2007). Effects of response task and accessory stimuli on redundancy gain: Tests of the hemispheric coactivation model. *Journal of Experimental Psychology: Human Perception and Performance*, *33*(4), 829. https://doi.org/10.1037/0096-1523.33.4.829

Miniussi, C., Girelli, M., & Marzi, C. A. (1998). Neural site of the redundant target effect: Electrophysiological evidence. *Journal of Cognitive Neuroscience*, *10*(2), 216-230. https://doi.org/10.1162/089892998562663

Mohr, B., Landgrebe, A., & Schweinberger, S. R. (2002). Interhemispheric cooperation for familiar but not unfamiliar face processing. *Neuropsychologia*, *40*(11), 1841-1848. https://doi.org/10.1016/S0028-3932(02)00040-4

Mooshagian, E., Kaplan, J., Zaidel, E., & Iacoboni, M. (2008). Fast visuomotor processing of redundant targets: the role of the right temporo-parietal junction. *PLoS One*, *3*(6), e2348. https://doi.org/10.1371/journal.pone.0002348

Mordkoff, J. T., Miller, J., & Roch, A.-C. (1996). Absence of coactivation in the motor component: Evidence from psychophysiological measures of target detection. *Journal of Experimental Psychology: Human Perception and Performance*, *22*(1), 25. https://doi.org/10.1037/0096-1523.22.1.25

Morland, A. B., Jones, S. R., Finlay, A. L., Deyzac, E., Lê, S., & Kemp, S. (1999). Visual perception of motion, luminance and colour in a human hemianope. *Brain*, *122*(6), 1183-1198. https://doi.org/10.1093/brain/122.6.1183

Müller-Oehring, E. M., Schulte, T., Kasten, E., Poggel, D. A., Müller, I., Wüstenberg, T., & Sabel, B. A. (2009). Parallel interhemispheric processing in hemineglect: Relation to visual field defects. *Neuropsychologia*, *47*(12), 2397-2408. https://doi.org/10.1016/j.neuropsychologia.2009.04.014

Munro, K. J., & Prendergast, G. (2019). Encouraging pre-registration of research studies. *International Journal of Audiology*, *58*(3), 123-124. https://doi.org/10.1080/14992027.2019.1574405

Murray, M. M., Foxe, J. J., Higgins, B. A., Javitt, D. C., & Schroeder, C. E. (2001). Visuo-spatial neural response interactions in early cortical processing during a simple reaction time task: a high-density electrical mapping study. *Neuropsychologia*, *39*(8), 828-844. https://doi.org/10.1016/S0028-3932(01)00004-5

Ogourtsova, T., Korner-Bitensky, N., Leh, S. E., Eskes, G., & Ptito, A. (2011). Superior colliculi involvement in poststroke unilateral spatial neglect: A pilot study. *Topics in stroke rehabilitation*, *18*(6), 770-785. https://doi.org/10.1310/tsr1806-770

Omura, K., Tsukamoto, T., Kotani, Y., Ohgami, Y., Minami, M., & Inoue, Y. (2004). Different mechanisms involved in interhemispheric transfer of visuomotor information. *Neuroreport*, *15*(18), 2707-2711.

Ouimet, C., Jolicœur, P., Miller, J., Ptito, A., Paggi, A., Foschi, N., Ortenzi, A., & Lassonde, M. (2009). Sensory and motor involvement in the enhanced redundant target effect: A study comparing anterior- and totally split-brain individuals. *Neuropsychologia*, *47*(3), 684-692. https://doi.org/10.1016/j.neuropsychologia.2008.11.023

Overgaard, M. (2011). Visual experience and blindsight: a methodological review. *Experimental Brain Research*, *209*, 473-479. https://doi.org/10.1007/s00221-011-2578-2

Overgaard, M., Fehl, K., Mouridsen, K., Bergholt, B., & Cleeremans, A. (2008). Seeing without seeing? Degraded conscious vision in a blindsight patient. *PLoS One, 3*(8), e3028. https://doi.org/10.1371/journal.pone.0003028

Pollmann, S., & Zaidel, E. (1999). Redundancy gains for visual search after complete commissurotomy. *Neuropsychology, 13*(2), 246–258. https://doi.org/10.1037/0894-4105.13.2.246

Raab, D. H. (1962). Statistical facilitation of simple reaction times. *Transactions of the New York Academy of Sciences, 24*, 574-590. https://doi.org/10.1111/j.2164-0947.1962.tb01433.x

Railo, H., Andersson, E., Kaasinen, V., Laine, T., & Koivisto, M. (2014). Unlike in clinical blindsight patients, unconscious processing of chromatic information depends on early visual cortex in healthy humans. *Brain Stimulation, 7*(3), 415-420. https://doi.org/10.1016/j.brs.2014.01.060

Reinhard, J., Schreiber, A., Schiefer, U., Kasten, E., Sabel, B. A., Kenkel, S., Vonthein, R., & Trauzettel-Klosinski, S. (2005). Does visual restitution training change absolute homonymous visual field defects? A fundus controlled study. *British Journal of Ophthalmology, 89*, 30-35. https://doi.org/10.1136/bjo.2003.040543

Reuter-Lorenz, P. A., Nozawa, G., Gazzaniga, M., & Hughes, H. (1995). Fate of neglected targets: A chronometric analysis of redundant target effects in the bisected brain. *Journal of Experimental Psychology: Human Perception and Performance, 21*(2), 211–230. https://doi.org/10.1037/0096-1523.21.2.211

Richards, W. (1973). Visual processing in scotomata. *Experimental Brain Research, 17*, 333-347. https://doi.org/10.1007/BF00234098

Riddoch, G. (1917). Dissociation of visual perceptions due to occipital injuries, with especial reference to appreciation of movement. *Brain, 40*(1), 15-57. https://doi.org/10.1093/brain/40.1.15

Ridgway, N., Milders, M., & Sahraie, A. (2008). Redundant target effect and the processing of colour and luminance. *Experimental Brain Research, 187*, 153-160. https://doi.org/10.1007/s00221-008-1293-0

Rosenthal, R. (1979). The file drawer problem and tolerance for null results. *Psychological Bulletin, 86*(3), 638–641. https://doi.org/10.1037/0033-2909.86.3.638

Rosenthal, R. (1993). *Meta-Analytic Procedures For Social Research* (6 ed.). Sage Publications.

Roser, M., & Corballis, M. C. (2002). Interhemispheric neural summation in the split brain with symmetrical and asymmetrical displays. *Neuropsychologia, 40*(8), 1300-1312. https://doi.org/10.1016/S0028-3932(01)00219-6

Roser, M., & Corballis, M. C. (2003). Interhemispheric neural summation in the split brain: effects of stimulus colour and task. *Neuropsychologia, 41*(7), 830-846. https://doi.org/10.1016/S0028-3932(02)00290-7

Ross, A. I., Schenk, T., Billino, J., MacLeod, M.-J., & Hesse, C. (2018). Avoiding unseen obstacles: Subcortical vision is not sufficient to maintain normal obstacle avoidance behaviour during reaching. *Cortex, 98*, 177-193. https://doi.org/10.1016/j.cortex.2016.09.010

Sahraie, A., Trevethan, C. T., MacLeod, M.-J., Weiskrantz, L., & Hunt, A. R. (2013). The continuum of detection and awareness of visual stimuli within the blindfield: From blindsight to the sighted-sight. *Investigative Ophthalmology & Visual Science, 54*(5), 3579-3585. https://doi.org/10.1167/iovs.12-11231

Saionz, E. L., Tadin, D., Melnick, M. D., & Huxlin, K. R. (2020). Functional preservation and enhanced capacity for visual restoration in subacute occipital stroke. *Brain, 143*(6), 1857-1872. https://doi.org/10.1093/brain/awaa128

Savazzi, S., & Marzi, C. A. (2004). The superior colliculus subserves interhemispheric neural summation in both normals and patients with a total section or agenesis of the corpus callosum. *Neuropsychologia*, *42*(12), 1608-1618. https://doi.org/10.1016/j.neuropsychologia.2004.04.011

Savazzi, S., & Marzi, C. A. (2008). Does the redundant signal effect occur at an early visual stage? *Experimental Brain Research*, *184*, 275-281. https://doi.org/10.1007/s00221-007-1182-y

Schärli, H., Harman, A. M., & Hogben, J. H. (1999). Blindsight in subjects with homonymous visual field defects. *Journal of Cognitive Neuroscience*, *11*(1), 52-66. https://doi.org/10.1162/089892999563247

Schmid, D., & Schenk, T. (2022). Definition: Redundant target paradigm. *Cortex, 146*, 271-272. https://doi.org/10.1016/j.cortex.2021.11.002

Schmid, D., Schneider, S., & Schenk, T. (2022). How to test blindsight without light scatter artefacts? *Neuropsychologia*, *173*, 108308. https://doi.org/10.1016/j.neuropsychologia.2022.108308

Schröter, H., Fiedler, A., Miller, J., & Ulrich, R. (2011). Fusion prevents the redundant signals effect: Evidence from stereoscopically presented stimuli. *Journal of Experimental Psychology: Human Perception and Performance*, *37*(5), 1361-1368. https://doi.org/10.1037/a0024280

Schwarzer, G., Carpenter, J. R., & Rücker, G. (2015). *Meta-analysis with R*. Springer.

Striemer, C. L., Chapman, C. S., & Goodale, M. A. (2009). "Real-time" obstacle avoidance in the absence of primary visual cortex. *Proceedings of the National Academy of Sciences*, *106*(37), 15996-16001. https://doi.org/10.1073/pnas.0905549106

Tamietto, M., Cauda, F., Corazzini, L. L., Savazzi, S., Marzi, C. A., Goebel, R., Weiskrantz, L., & de Gelder, B. (2010). Collicular vision guides nonconscious behavior. *Journal of Cognitive Neuroscience*, *22*(5), 888-902. https://doi.org/10.1162/jocn.2009.21225

Tomaiuolo, F., Ptito, M., Marzi, C. A., Paus, T., & Ptito, A. (1997). Blindsight in hemispherectomized patients as revealed by spatial summation across the vertical meridian. *Brain*, *120*(5), 795-803. https://doi.org/10.1093/brain/120.5.795

Turatto, M., Mazza, V., Savazzi, S., & Marzi, C. A. (2004). The role of the magnocellular and parvocellular systems in the redundant target effect. *Experimental Brain Research*, *158*, 141-150. https://doi.org/10.1007/s00221-004-1884-3

Vaina, L. M., Soloviev, S., Calabro, F. J., Buonanno, F., Passingham, R., & Cowey, A. (2014). Reorganization of retinotopic maps after occipital lobe infarction. *Journal of Cognitive Neuroscience*, *26*(6), 1266-1282. https://doi.org/10.1162/jocn_a_00538

Van der Heijden, A. H. C., Schreuder, R., Maris, L., & Neerincx, M. (1984). Some evidence for correlated separate activation in a simple letter-detection task. *Perception & Psychophysics*, *36*(6), 577-585. https://doi.org/10.3758/BF03207519

van Koningsbruggen, M., Koller, K., & Rafal, R. D. (2017). Deafferentation of the superior colliculus abolishes spatial summation of redundant visual signals. *Frontiers in Systems Neuroscience*, *11*(9), 1-8. https://doi.org/10.3389/fnsys.2017.00009

Viechtbauer, W. (2005). Bias and Efficiency of Meta-Analytic Variance Estimators in the Random-Effects Model. *Journal of Educational and Behavioral Statistics*, *30*(3), 261-293. https://doi.org/10.3102/10769986030003261

Viechtbauer, W. (2010). Conducting meta-analyses in R with the metafor package. *Journal of Statistical Software*, *36*(3), 1-48. https://doi.org/10.18637/jss.v036.i03

Viechtbauer, W. (2021). Weights in Models Fitted with the rma.mv() Function. Retrieved 07.07.2023 from https://web.archive.org/web/20230514233245/https://www.metafor-project.org/doku.php/tips:weights_in_rma.mv_models

Viechtbauer, W., & Cheung, M. W.-L. (2010). Outlier and influence diagnostics for meta-analysis. *Research Synthesis Methods*, *1*, 112-125. https://doi.org/https://doi.org/10.1002/jrsm.11

Weiskrantz, L. (1999). *Consciousness lost and found: A neuropsychological exploration*. Oxford University Press. https://doi.org/10.1093/acprof:oso/9780198524588.001.0001

Weiskrantz, L., Warrington, E. K., Sanders, M. D., & Marshall, J. (1974). Visual capacity in the hemianopic field following a restricted occipital ablation. *Brain*, *97*(1), 709-728. https://doi.org/10.1093/brain/97.1.709

Westerhausen, R. (2022). Interhemispheric integration after callosotomy: A meta-analysis of poffenberger and redundant-target paradigms. *Neuropsychology Review*. https://doi.org/10.1007/s11065-022-09569-w

Whitwell, R. L., Striemer, C. L., Nicolle, D. A., & Goodale, M. A. (2011). Grasping the non-conscious: Preserved grip scaling to unseen objects for immediate but not delayed grasping following a unilateral lesion to primary visual cortex. *Vision Research*, *51*(8), 908-924. https://doi.org/10.1016/j.visres.2011.02.005

Wüst, S., Kasten, E., & Sabel, B. A. (2002). Blindsight after optic nerve injury indicates functionality of spared fibers. *Journal of Cognitive Neuroscience*, *14*(2), 243-253. https://doi.org/10.1162/089892902317236876

Yu, J.-C., Chang, T.-Y., & Yang, C.-T. (2014). Individual differences in working memory capacity and workload capacity. *Frontiers in Psychology*, *5*(1465), 1-15. https://doi.org/10.3389/fpsyg.2014.01465

Zeki, S., & Ffytche, D. H. (1998). The Riddoch syndrome: insights into the neurobiology of conscious vision. *Brain*, *121,* 25-45. https://www.ncbi.nlm.nih.gov/pubmed/9549486

Zihl, J., & von Cramon, D. (1985). Visual field recovery from scotoma in patients with postgeniculate damage. A review of 55 cases. *Brain*, *108*(2), 335-365. https://doi.org/10.1093/brain/108.2.335