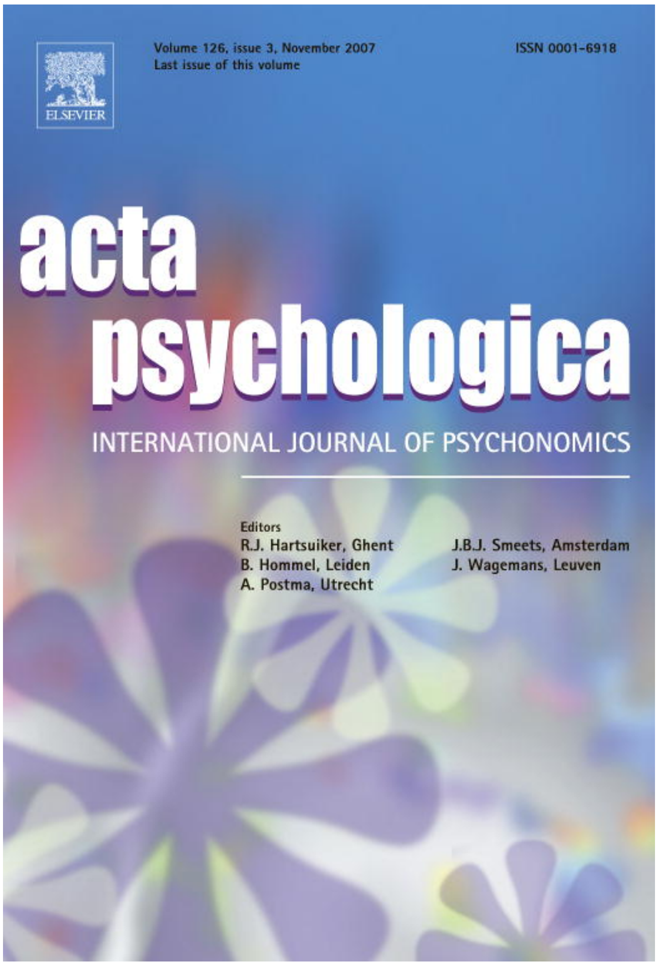
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Is there a general task switching ability?

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Abstract

Participants were tested on two analogous task switching paradigms involving Shape/Size tasks and Vertical/Horizontal tasks, respectively, and three measures of psychometric intelligence, tapping fluid, crystallized and perceptual speed abilities. The paradigms produced similar patterns of group mean reaction times (RTs) and the vast majority of the participants showed switching cost (switch RT minus repeat RT), mixing cost (repeat RT minus single-task RT) and congruency effects. The shared intra-individual variance across paradigms and with psychometric intelligence served as criteria for general ability. Structural equations modeling indicated that switching cost with ample preparation ("residual cost") and mixing cost met these criteria. However, switching cost with little preparation and congruency effects were predominantly paradigm specific.

PsycINFO classification: 2340; 2300; 2346; 3297

Keywords: Task switching; Individual differences; Intelligence; Structural equations modeling

1. Introduction

The concept of executive control refers to a broad class of skills required to ensure context appropriate action in conditions in which inborn tendencies or overlearned skills and habits provide inappropriate guidance. These self control skills include shifting mental sets, inhibition of prepotent tendencies, and working memory updating (Miyake et al., 2000), as well as the ability to choose between alternative strategies, monitoring

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performance and its consequences, and so forth (Logan, 1985; Logan & Gordon, 2001; Smith & Jonides, 1999). Although there is no agreed-upon listing of functions, the ability to flexibly switch mental sets is regarded as a core executive function by most theories.

Following influential conceptions by Baddeley (1986) and Norman and Shallice (1986), it is still widely believed that executive functions represent, at least at some level, the operation of a single entity, the ''central executive'', presumably involving the prefrontal cortex. However, individual difference examinations challenge this notion in showing that different functions of this central system tap nearly independent abilities. In Miyake et al.'s (2000) work, for example, three mildly correlated functions were identified: inhibition, working memory updating, and shifting. Dual task coordination, which was another function studied, was not correlated with any of these three functions, suggesting that a fourth dimension may be included in this taxonomy. A more recent work from this group (Friedman & Miyake, 2004) indicates that Inhibition, which Miyake et al. suggested to be a single ability, comprises two nearly independent inhibition abilities: Response/Distractor inhibition and Resistance to proactive interference.

In the present work we addressed a similar question: Is there a general task switching ability? A related question has been asked in the neuroimaging field by Dosenbach et al. (2006), who conducted a meta-analysis on the results of 10 fMRI studies on task switching. These authors identified two loci which were activated in all the 10 paradigms that were investigated, including an area in the dorsal anterior cingulate gyrus and an area in the anterior insular cortex, bilaterally. Derrfuss, Brass, Neumann, and von Cramon (2005) who also performed a meta-analysis on imaging studies, identified a region in the lateral prefrontal cortex, the Inferior Frontal Junction, which was commonly activated by task switching and the Stroop task. While these results support the notion of a central ability controlling task switching in general, they remain inconclusive. The reason is that the areas which were commonly activated may be involved in functions that do not necessarily link directly to performance, and therefore do not indicate a common ability to perform the tasks. Such functions include monitoring and assessing the degree of response conflict (e.g., Botvinick, Braver, Carter, Barch, & Cohen, 2001) and emotional reactions (Critchley et al., 2003; Jennings, van der Molen, van der Veen, & Debski, 2002). A performance-based individual differences investigation seems to be a more promising approach to demonstrate an across domain task switching ability.

A number of published studies found that the performance costs associated with task switching show shared variance across paradigm content (Friedman & Miyake, 2004; Friedman et al., 2006; Kray & Lindenberger, 2000; Meiran, Israeli, Levi, & Grafi, 1994; Miyake et al., 2000; Salthouse, Fristoe, McGurthy, & Hambrick, 1998; see also Ward, Roberts, & Phillips, 2001). The present study addresses two potential shortcomings of these investigations, one relating to the statistical approach and the other relating to the type of tasks studied.

1.1. Statistical approach

The first potential shortcoming concerns the statistical approach. These studies fractionated the variance of the switch cost estimates into two components: a common component and a unique component. This was done by explaining each observed cost estimate as the sum of the latent ability influence and an unexplained variance. Formally, $\sigma^2_{\mathrm{Observed}} = \sigma^2_{\mathrm{Ability}} + \sigma^2_{\mathrm{Unexplained}}$. However, this separation does not acknowledge the fact that the unexplained variance component has in principle two sub-components: unreliable variance, and reliable variance that is truly unique to that measure. Formally, $\sigma^2_{\mathrm{Observed}} = \sigma^2_{\mathrm{Ability}} + \sigma^2_{\mathrm{Unique}} + \sigma^2_{\mathrm{Error}}$. Consequently, it was impossible to estimate, based on these previous modeling attempts, what is the proportion of reliable variance that is shared across task contents. Formally, it was impossible to estimate $\sigma^2_{\mathrm{Ability}}/(\sigma^2_{\mathrm{Ability}} + \sigma^2_{\mathrm{Unique}})$.

1.2. Task choice

The second potential shortcoming of the previous studies refers to the distinction between perceptual processing streams within the associative cortex. These include the dorsal stream, involved in computing locations, possibly in the service of visually guided actions (the WHERE system), and the ventral stream, involved in computing object identity information, the WHAT system (Goodale & Milner, 1992; Mishkin, Ungerleider, & Macko, 1983).

From this perspective, all the previous studies included WHAT tasks and did not include any WHERE tasks. Thus, it is conceivable that the common switching variance found in the previous investigation may be partly or totally due to the shared perceptual processing rather than the shared central control functioning. These studies either focused on tasks involving making a decision based on object identity or on the ones requiring semantic decisions. Importantly, none of these investigations included spatial location tasks. Indeed, some paradigms that were used required shifting visual attention between right and left locations (Salthouse et al., 1998), between the global and local features of objects (Friedman & Miyake, 2004; Meiran et al., 1994; Miyake et al., 2000; Ward et al., 2001), and between two visually separated elements in the display (Miyake et al.; Friedman and Miyake; Meiran et al.). Nonetheless, the decisions made in the tasks did not require taking into account the object's location in space.

There are at least two reasons why the WHAT/WHERE distinction may be important in the present context. First, there are suggestions that the WHERE and WHAT streams continue into the prefrontal cortex that presumably plays a critical role in executive functioning (e.g., Smith & Jonides, 1999). Additionally, there is a growing appreciation that the posterior cortex (e.g., Braver, Reynolds, & Donaldson, 2003; Dosenbach et al., 2006; Sohn, Ursu, Anderson, Stenger, & Carter, 2000) and sub-cortical brain structures (e.g., Berger et al., 2005; Cools, Barker, Sahakian, & Robbins, 2001; Meiran, Friedman, & Yehene, 2004) are also critically involved in task switching performance, and some of these regions, especially the posterior cortex, are clearly differentiated according to WHAT and WHERE processing. Based on these considerations, we studied paradigms that involved both WHAT and WHERE tasks, something that had not been done beforehand.

2. The present study

In the present study, we used two different task switching paradigms, one emphasizing WHERE processing, involving UP–DOWN and RIGHT–LEFT judgments and the other paradigm in which at least one of the tasks (shape) is likely to involve WHAT processing. Aside from this difference, the two paradigms had the same structure in terms of number of tasks, number of alternative responses per task, block sequence and trial sequence. We thought that this aspect about our design would maximize the chances of identifying common control processes, if present (e.g., see Altmann, 2004, for the importance of these factors in shaping performance strategies). We used an age-homogenous sample, ruling out any age-related common variance that could account for some of the previous results such as Kray and Lindenberger's (2000) and Salthouse et al.'s (1998). Additionally, the paradigms involved cued task switching, meaning that the tasks varied randomly and each trial began with the presentation of an instructional cue. This was done because previous studies suggest that, when the task sequence is instructed in advance, the ability to retain the sequence in memory influences performance (Luria & Meiran, 2003; Rubinstein, Meyer, & Evans, 2001) and shows individual differences (Logan, 2004). Moreover, there is evidence that switching cost results in part from the processing of the task cue rather than from task switching (Arbuthnott & Woodward, 2002; Logan & Bundesen, 2003; Mayr & Kliegl, 2003). In order to minimize the role of any shared variance that is due to cue processing, the sets of task cues that we used in the two tasks were quite different from one another. One set was spatial, pictorial and composed of two independent elements. The other set was verbal and was made of a single element (a word). Finally, since the type of stimulus–response mapping as arbitrary as opposed to natural could also be a factor driving the common variance, as one paradigm involved a natural, S–R mapping (UP = UP, LEFT = LEFT) while the other paradigm involved an arbitrary mapping (CIRCLE = RIGHT).

Even if common variance is found between the two paradigms, it still may be due, in part, to non-executive processes. With regard the Shape–Size paradigm, while shape is clearly related to WHAT processing, size judgments may involve the dorsal stream because objects' size is an important piece of information for guided grasping, which is believed to be a dorsal function. For example, Oliver and Thompson-Schill (2003) found dorsal activation when participants had to recall the size of named objects.

In addition, all the tasks required manual key-presses. Because the keys are identified according to their location and because the action is manual, responding probably involved the dorsal pathway. Furthermore, both tasks required response selection, which according to cognitive theorizing (Pashler, 1998) constitutes a central system that is independent of perception. (However, Cohen & Shoup, 1997, provide behavioral evidence and Jiang & Kanwisher, 2003, provide neuroimaging evidence that response selection and perception are not independent of one another.)

2.1. *Indices of executive functioning in the task switching paradigm*

The task switching paradigm yields a plethora of experimental effects, many of which could, in principle, serve as indices of executive functioning. In the present work we focused on three effects: switching cost, mixing cost, and the task rule congruency effect, or "congruency effect", for short. Switching cost is defined as the difference in performance between two types of trials present in experimental blocks in which two (or more) tasks are intermixed: task switch trials (or "switch trials", for short) and task repetition trials (or "repeat trials", for short). The typical finding is that switch trials are associated with an increased error rate and prolonged response times compared to repeat trials. Switching cost presumably reflects the cost of having just switched to a new task, although the nature of the exact mechanism is still debated (Monsell, 2003, for review). Mixing cost is defined conceptually as the difference in performance between mixed tasks trials and the single-task baseline. Typically, even repetition trials are associated with slower and more error prone responses compared with single-task trials. Mixing cost probably reflects both the increased demand for ongoing monitoring and control in mixed tasks conditions (Braver et al., 2003), and the need to make a task decision, which is common to switch and non-switch trials (e.g., Koch, Prinz, & Allport, 2005; Rubin & Meiran, 2005). The costs have been shown to be empirically dissociable. For example, mixing cost is more strongly affected by aging (Kray & Lindenberger, 2000), while switching cost is more strongly affected by attention deficit (Cepeda, Cepeda, & Kramer, 2000). The costs have also been shown to activate different brain regions (Braver et al., 2003) and to be differentially affected by experimental manipulations (e.g., Steinhauser & Hübner, 2005). Additionally, mixing cost is present even without task switching (Yehene, Meiran, & Soroker, 2005), whereas switching cost seems to depend on actual switching (Schuch & Koch, 2003).

Finally, the congruency effect is a common finding (Kiesel, Wendt, & Peters, in press; Meiran, 1996, 2000, 2005; Sudevan & Tylor, 1987; but see Arbuthnott, 2005, for an important exception). It indicates that performance is better for targets in which both attributes are associated with the correct response (e.g., both features LARGE and SQUARE should result in pressing the left key) as compared to targets in which the two stimulus attributes are associated with different correct responses. This effect may index the ability to overcome conflict, as exemplified in the task switching paradigm. We are unaware of any individual differences studies that incorporated this measure beforehand. There are several studies which incorporated analogous measures. For example, Ward et al. (2001) studied the Stroop effect: slower color naming of incongruent stimuli such as the word RED written in green ink. Their results show that the effect correlated negligibly with an analogous effect in a numerical-judgment task. Perhaps the most relevant study is Friedman and Miyake's (2004). In this study, they identified a latent variable which explained the common variance of various interference tasks, including the Stroop effect and the flanker compatibility effect, which may be analogous to the congruency effect. Peterson et al.'s (2002) imaging study is especially relevant here because they found that the Simon task, involving overcoming spatial distracting information, and the Stroop task, which involves overcoming object-based distracting information, activated very similar brain regions. This evidence suggests that the congruency effect found in the Vertical/Horizontal paradigm should correlate with the congruency effect in the Shape/Size paradigm.

2.2. Statistical approach

In the present work we employed structural equations modelling (SEM), which is an approach to modeling correlational data. A notable aspect of SEM is that it is confirmatory in nature. Unlike exploratory factor analysis, SEM makes it possible to compare competing models, and also to estimate the degree to which a given model fits the data.

The novel aspect of our analytic design is the fractionation of variance of each observed variable into three components, as opposed to just two components in previous studies on task switching. Our design resembles that used by Cunningham, Preacher, and Banaji (2001), who studied implicit attitudes. Like them, we used multiple estimates (two, in our case) for each measure. The approach is depicted schematically in Fig. 1. For example, in modeling switch costs, one measure was the cost found in the Shape/Size paradigm given a short cue-target interval. In the relevant model, we computed this cost twice, once from the results of the odd-numbered blocks of the experiment and once from the even numbered blocks. We also defined a latent variable explaining the common variance of these two estimates. As a result, the latent variable just described represents the reliable

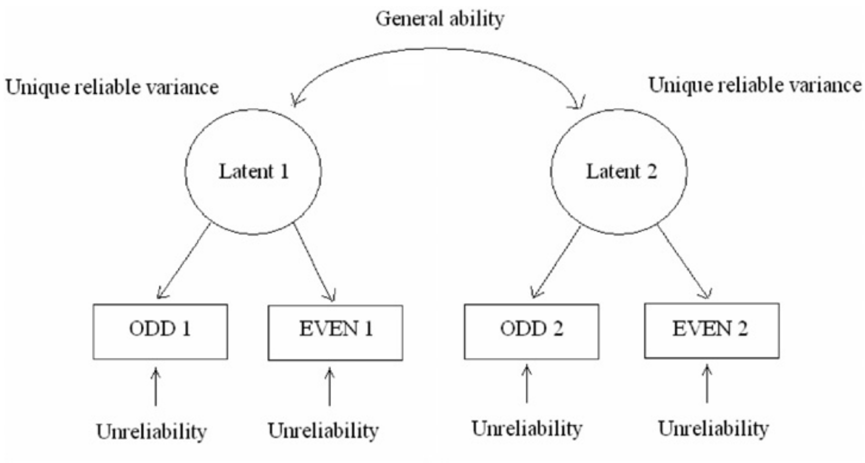

Fig. 1. Schematic description of the statistical approach. Latent variables represent the shared (reliable) variance across odd-numbered and even-numbered experimental blocks. The (squared) correlation among them represents the across content common reliable variance. Their unshared variance represents paradigm specific reliable variance.

variance of switching cost in the Shape/Size paradigm, given a short cue-target interval. To show evidence for a general switching ability, we estimated the correlation between the aforementioned latent variable and an equivalent latent variable related to the Vertical/Horizontal paradigm. The common variance component is represented by the correlation between latent variables. The reliable but unique component is represented by the proportion of the latent variable's variance not shared with the equivalent latent variable from the other paradigm. The unreliable variance is the proportion of variance in the observed indices not explained by the latent variable. To the best of our knowledge, this is the first time such an approach has been applied in the task switching literature.

2.3. Criteria for general ability

Our core question can, in fact, be separated into two related questions: "Is there a general task switching ability?" and "is task switching ability general?" To answer these questions, we adopted the following two criteria regarding the correlations between latent variables. Our first criterion refers to the correlation between equivalent measures of switching efficiency based on the two paradigms. This shared variance indicates a switching ability that is common to these paradigms. A significant and non-trivial degree of shared variance between the indices of switching ability taken from the two paradigms would provide a positive answer to the question, is there a general task switching ability? Our second criterion related to the question, is task switching ability general? It relates to the correlation between switching ability and psychometric intelligence based on standard paper and pencil tests. We intentionally use the term "psychometric intelligence" instead of "general intelligence" or "fluid intelligence" because the sample of abilities we used was deliberately biased. This correlation interested us in part because of Friedman et al.'s (2006) recent and quite provocative results. These authors showed that switching ability and the ability to inhibit prepotent responses showed negligible correlations with general fluid intelligence. To create a measure of psychometric intelligence, we defined a latent variable that represented the shared variance among Raven matrices, representing general fluid ability; a Vocabulary test, representing crystallized ability; and a content-balanced composite measure of perceptual speed tests. This latent variable was intentionally biased towards perceptual speed, to take into account the fact that our measures of switching ability were based on reaction time (RT) effects, and because previous authors have argued that the measures of executive functions do not tap much additional individual variance beyond the variance that is already picked up by the measures of fluid abilities and perceptual speed (Salthouse, 2005).

3. Method

3.1. Participants

Ninety eight undergraduate students (51 women and 47 men) from Ben Gurion University of the Negev and two affiliated colleges took part in the study. They received a partial course credit in return for their participation. In both task switching paradigms we fully counterbalanced across participants the following three conditions: (1) key assignment in the Vertical/Horizontal paradigm (up-left and down-right for one half of the participants and up-right and down-left for the remaining participants, see Fig. 2), (2) the task performed as a single-task in the Vertical/Horizontal paradigm, and (3) the task performed as a single-task in the Shape/Size paradigm. Due to an experimental error, we lost the Vertical/Horizontal paradigm data of two participants and the Shape/Size paradigm data of one participant. Therefore, we report the results from those ninety five participants who had complete data.

3.2. Apparatus and paper and pencil tests

The task switching testing was performed using an IBM clone with 14″ monitors and controlled by software written in MEL 2.0 (Schneider, 1988). Responses were collected with a standard keyboard, and the claimed RT recording accuracy was to the nearest 1 ms.

The paper and pencil tests included Raven's (1965) advanced progressive matrices, with 30 min administration time, the Hebrew Vocabulary subset from the Hebrew battery, AKA (Fischman, 1982), and three perceptual speed tests including a verbal test (Letter Cancellation, from AKA) a numerical test (Digit Cancellation, from AKA) and a pictorial test (Identical Pictures from the ETS kit, Ekstrom, French, Harman, & Dermen, 1976). The description and psychometric properties of the AKA tests may be found in Meiran and Fischman (1989). The perceptual speed composite measure was the mean of the sample Z-scores of the three perceptual speed tests.

3.3. The task switching paradigms

One paradigm involved spatial location (Fig. 2), where the tasks were UP–DOWN and RIGHT–LEFT. The other involved required CIRCLE–SQUARE and SMALL–LARGE decisions (Fig. 3). In both paradigms, the single-task block came last in order not to cause the tasks to be differentially practiced, or to have one task be more recently practiced than the other. As commonly done in studies of individual differences, we used a uniform testing order. However, the fact that the participants were always tested on the

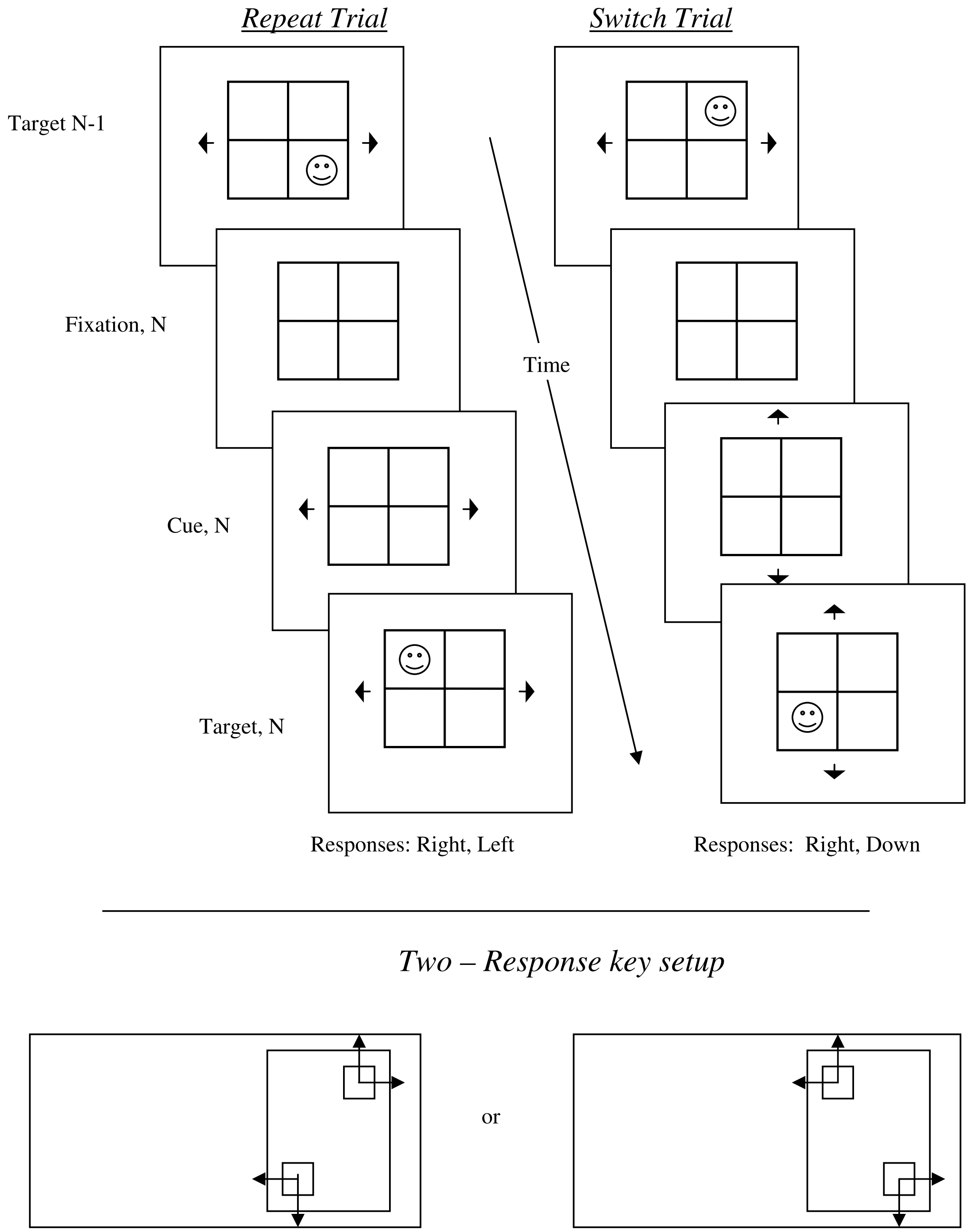

Fig. 2. Schematic illustration of the Vertical/Horizontal paradigm. The left side of the figure displays a sequence of events in a repeat trial in which a RIGHT–LEFT task trial was followed by another RIGHT–LEFT task trial. The right side of the figure displays the events for a switch trial in which a LEFT–RIGHT task trial was followed by a UP–DOWN task trial.

Vertical/Horizontal experiment after they were tested on the Vertical/Horizontal paradigm is not problematic given that training does not transfer between these two task switching paradigms in particular (Sosna, 2001), and between paradigms employing different target stimuli in general (Armony-Shimoni, 2001).

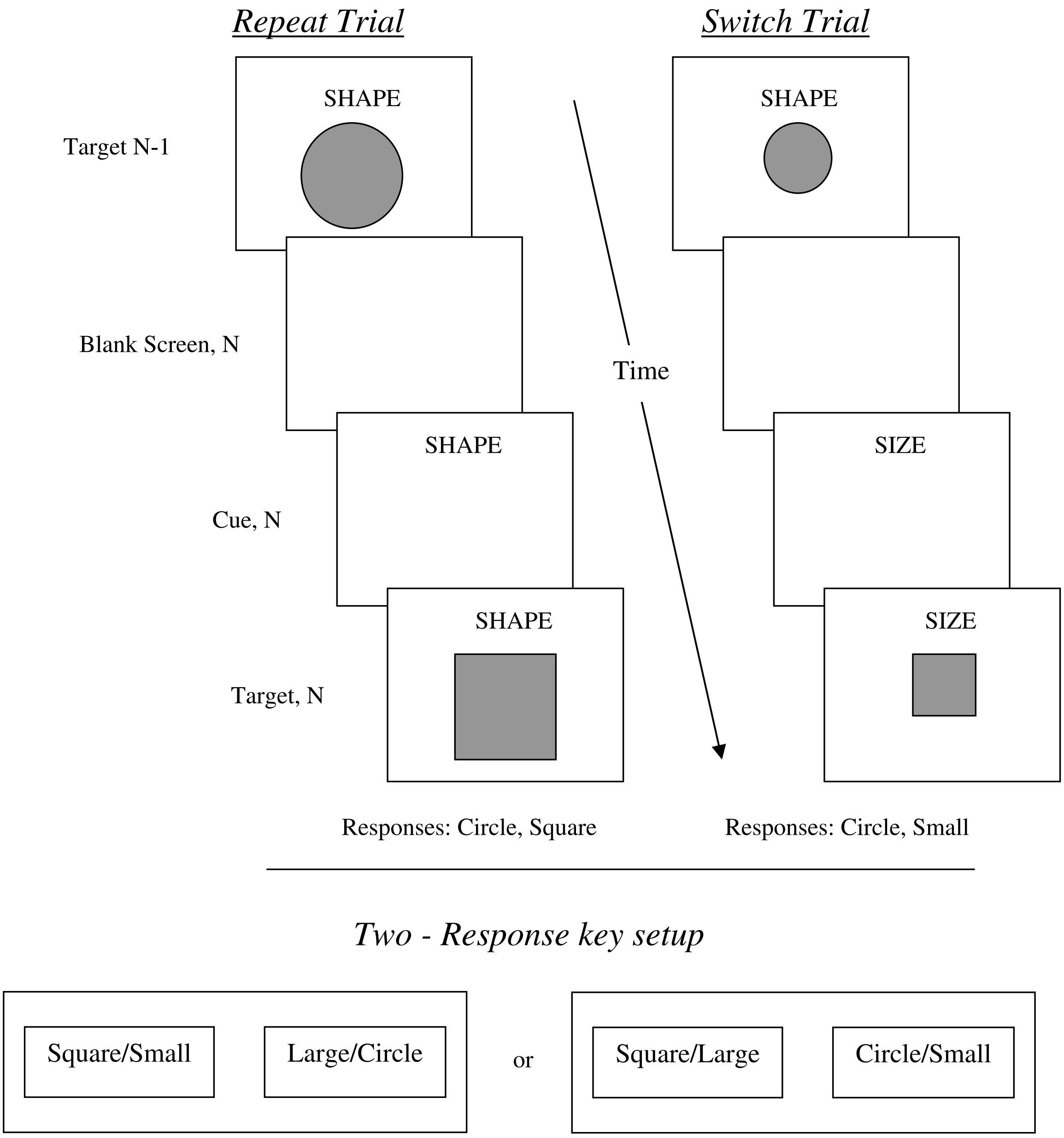

Fig. 3. Schematic illustration of the Shape/Size paradigm. The left side of the figure displays a sequence of events in a repeat trial in which a SHAPE task trial was followed by another SHAPE task trial. The right side of the figure displays the events for a switch trial in which a SIZE task trial came after a SHAPE task trial.

3.3.1. Vertical/horizontal paradigm

(This is the exact same paradigm as in Meiran, Gotler, & Perlman, 2001, Experiment 1). The stimuli were drawn in white on black using graphic symbols in the extended ASCII code and included a 2×2 grid that was presented at the screen center and subtended a visual angle of approximately 3.4° (width) × 2.9° (height) (These values were calculated assuming an observation distance of 60 cm). The target stimulus was the smiley-face character (ASCII code 1), which subtended approximately .3° (width) × .5° (height). The task cues were arrow heads (ASCII codes 16, 17, 30, and 31) subtended approximately .3° × .3° and were positioned .7° from the end of the grid. Responses were given by pressing keys on the keyboard. Half of the participants were assigned to a response key combination in which "UP" and "LEFT" responses were mapped to the upper left key (the number

"3" in the keyboard), and "DOWN" and "RIGHT" responses were mapped to the lower right key (the number "7" in the keyboard). The other half of the participants were assigned to a reversed response key combination in which "UP" and "RIGHT" responses were mapped to the upper right response key (the number "1" in the keyboard) key, while "DOWN" and "LEFT" responses were mapped to the lower left response key (the number "9"in the keyboard).

3.3.2. Shape/Size paradigm

This paradigm is the same as that used by Yehene et al. (2005). The stimuli were filled in gray color and were either a small/large circle (with a diameter subtending a visual angle of approximately 1.4° or 3.0°) or a small/large square (each side subtending 1.4° or 3.0°). The task cues were the Hebrew equivalents of the words SIZE and SHAPE. They were presented 1 degree above the position where the target stimulus would have been presented. Their height was .5° and their width varied between 2° and 2.5°. Responses were given by pressing on the keyboard. Half of the participants were assigned to a response key combination in which "SMALL" and "CIRCLE" responses were mapped to the right key (the letter "*L*" in the keyboard), and "LARGE" and "SQUARE" responses were mapped to the left key (the letter "*S*" in the keyboard). The other half of the participants were assigned to a reversed response key combination in which "LARGE" and "CIRCLE" responses were mapped to the right response key ("*L*") key while "SMALL" and "SQUARE" responses were mapped to the left response key ("*S*"). Small stickers with the first letter of the attribute were placed on the relevant buttons pointing out the appropriate attributes to the participants.

3.4. Procedure

The study was run in three sessions, separated by 2–4 days. The first session, of approximately 1.5 h, was devoted to intelligence testing and was conducted in small groups of up to 20 participants. Participants were examined individually on the Vertical/Horizontal and Size/Shape task switching paradigms in Sessions 2 and 3, respectively. To ensure that the monitor was positioned above the center of the response keys, in the Vertical/Horizontal paradigm the center of the keypad was aligned with the center of the monitor by shifting the entire keyboard to the left, and in the Shape/Size paradigm the center of the entire keyboard was aligned with the center of the monitor. In every other respect, the paradigms had the same procedure. Each of them began with 20 mixed tasks trials serving for warm-up, followed by five experimental blocks (80 trials, each). The first four blocks were mixed tasks blocks, and the last block was a single-task block. Participants were asked to respond as quickly and as accurately as possible. Also, participants were informed about the transition to the single-task block and were told that from now on they would be requested to perform only one task. Each trial began after the response in the preceding trial and consisted of (1) a response-cue interval of 2032 ms. In the Vertical/Horizontal paradigm there was an empty grid that was presented during this interval serving for fixation. This was followed by (2) the presentation of the instructional cue for a cue-target interval of either 116 or 1016 ms, and (3) the presentation of the target stimulus until the response was given. A 400 Hz beep was presented for 100 ms after an error. The task (in Blocks 1–4), target location, and CTI were selected randomly on each trial. Hence, the instructional cue did not indicate the upcoming target location, key-press, or the precise target onset.

The potential problem of using a fixed response-cue interval is that it confounds set preparation time (operationalized by the cue-target interval) and the time allowed for the previous task-set to decay (Meiran, 1996). However, Meiran, Chorev, and Sapir (2000), who used the present Vertical/Horizontal paradigm, showed that set decay became asymptotic after about 1 s or less. Therefore, set decay probably contributed negligibly to the presently reported effects.

4. Results and discussion

Responses immediately following an error were omitted from all the analyses. In addition, responses slower than 3000 ms or quicker than 100 ms were analyzed for accuracy, only. The first 16 single-task trials were also discarded in order to minimize fadeout effects (Mayr & Liebscher, 2001; Meiran et al., 2001). We adopted an $\alpha = .05$ in all the comparisons. Errors were rare (2.4% and 2.9 % for the Vertical/Horizontal and Shape/Size paradigms, respectively) and there was no evidence for speed accuracy tradeoff. We therefore do not report the error analyses.

4.1. Analysis of means

The RT cell means of correct responses within the 100–3000 ms range were submitted to a five-way analysis of variance (ANOVA). The independent variables were Paradigm, CTI (short, long), task switch (switch, repeat, single-task), Congruency (congruent, incongruent), and response-repetition (same response, different response). Note that we included response-repetition in the present analysis in spite of not considering it in the individual differences analyses. We chose to do so in order to demonstrate even more convincingly that the two paradigms produced very similar patterns of means. response-repetition is an important variable to add here because it produces large and highly replicable interactions with switch (e.g., Rogers & Monsell, 1995, and many others since then). Table 1 reports the ANOVA results.

4.1.1. Effects which were found in both of the paradigms

Two high order interactions were statistically equivalent in both paradigms. That is, they were not involved in a significant higher order interaction with paradigm. One involved cue-target interval, switch and response-repetition and the other involved congruency and switch. Although the interaction between congruency, response-repetition and cue-target interval was significant and did not enter into a significant four-way interaction with paradigm (meaning that statistically it was the same in both paradigms), separate analyses indicated that this triple interaction reached significance in the Vertical/Horizontal paradigm, $F(1,94) = 5.26$, $\mathrm{MSe} = 4309.58$, but not in the Shape/Size paradigm, $F(1,94) = 1.83$, $\mathrm{MSe} = 6609.89$, $p = .18$. Moreover, as can be seen in Fig. 4, this interaction was very subtle. For these reasons, we do not discuss it any further.

4.1.1.1. Switch $\times$ cue-target interval $\times$ response-repetition (Fig. 5). We explored this interaction by a series of planned contrasts of the simple–simple effects of response-repetition.

Table 1
Analysis of variance of mean reaction times

Effect	df	F	MSe	p
Paradigm (P)	1, 94	15.16	245,562.57	[a]
Congruency (C)	1, 94	338.78	17,552.93	[a]
Switch (S)	2, 188	461.60	78,319.40	[a]
Response-Repetition (R)	1, 94	2.11	9,123.74	.15
Cue-target interval (CTI)	1, 94	1,388.77	21,713.88	[a]
$P \times C$	1, 94	4.88	19,428.31	[a] (.016[b])
$P \times S$	2, 188	7.96	39,199.82	[a] (.009[b])
$C \times S$	2, 188	61.27	6,746.79	[a]
$P \times R$	1, 94	60.41	6,841.18	[a]
$C \times R$	1, 94	0.89	4,312.48	.35
$S \times R$	2, 188	60.49	6,225.50	[a]
$P \times$ CTI	1, 94	161.20	16,640.94	[a] (.089[b])
$C \times$ CTI	1, 94	1.56	4.842.24	.21
$S \times$ CTI	2, 188	293.42	11,846.45	[a]
$R \times$ CTI	1, 94	3.25	4,695.98	.07
$P \times C \times S$	2, 188	1.09	6,310.81	.33
$P \times C \times R$	1, 94	0.20	5.743.96	.65
$P \times S \times R$	2, 188	0.72	5,720.48	.49
$C \times S \times R$	2, 188	0.57	7,293.72	.56
$P \times C \times$ CTI	1, 94	0.21	4,710.96	.64
$P \times S \times$ CTI	2, 188	5.60	8,508.41	[a] (.014[b])
$C \times S \times$ CTI	2, 188	2.74	6,072.97	.07
$P \times R \times$ CTI	1, 94	3.19	5,924.92	.08
$C \times R \times$ CTI	1, 94	8.05	4,212.54	[a]
$S \times R \times$ CTI	2, 188	31.32	5.955.44	[a]
$P \times C \times S \times R$	2, 188	1.10	5.672.94	.33
$P \times C \times S \times$ CTI	2, 188	0.59	3,733.49	.55
$P \times C \times R \times$ CTI	1, 94	0.12	6,706.94	.73
$P \times S \times R \times$ CTI	2, 188	2.37	6,747.82	.10
$C \times S \times R \times$ CTI	2, 188	0.42	4,855.01	.66
5-Way Interaction	2, 188	0.92	5,691.73	.40

[a] $p < .05$.
[b] Proportion of the effect variance explained by Paradigm (The presented values were computed as $SS_{\text{Interaction with paradigm}}/SS_{\text{Overall effect}}$).

When the cue-target interval was short, this effect was significant in switch trials, $F(1,94) = 44.04$, repeat trials, $F(1,94) = 124.88$, but not in single-task trials, $F < 1.0$. When the cue-target interval was long, this effect was non-significant in switch trials, $F = 1.02$, and in single-task trials, $F < 1.0$, but was significant in repeat trials, $F(1,94) = 4.06$. In conclusion, response-repetition resulted in response speeding in repeat trials and it resulted in response slowing in switch trials with a short cue-target interval.

4.1.1.2. Switch × congruency (Fig. 6). We explored this interaction by testing the simple effects of congruency. This effect was significant in all, switch, $F(1,94) = 243.82$, repeat, $F(1,94) = 209.00$, and single-task trials, $F(1,94) = 135.07$. In conclusion, the congruency effect was largest in size for switch trials, intermediate for repeat trials, and relatively small (37 ms) but significant in single-task trials. The presence of congruency effects in the single-task condition indicates some carryover from the mixed tasks condition which preceded it.

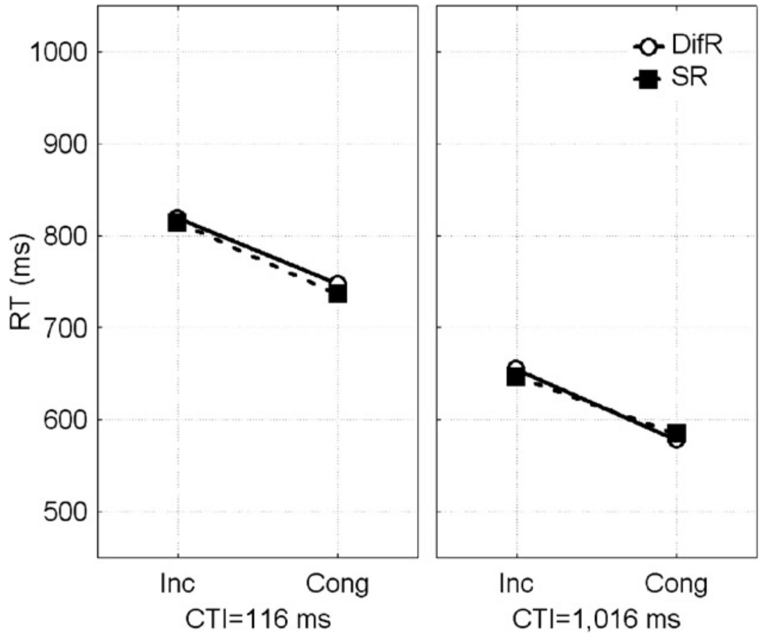

Fig. 4. Mean RT (ms) according to congruency, response-repetition and cue-target interval (CTI)*. *Inc = incongruent, Cong = congruent; DifR = different response; SR = same response.

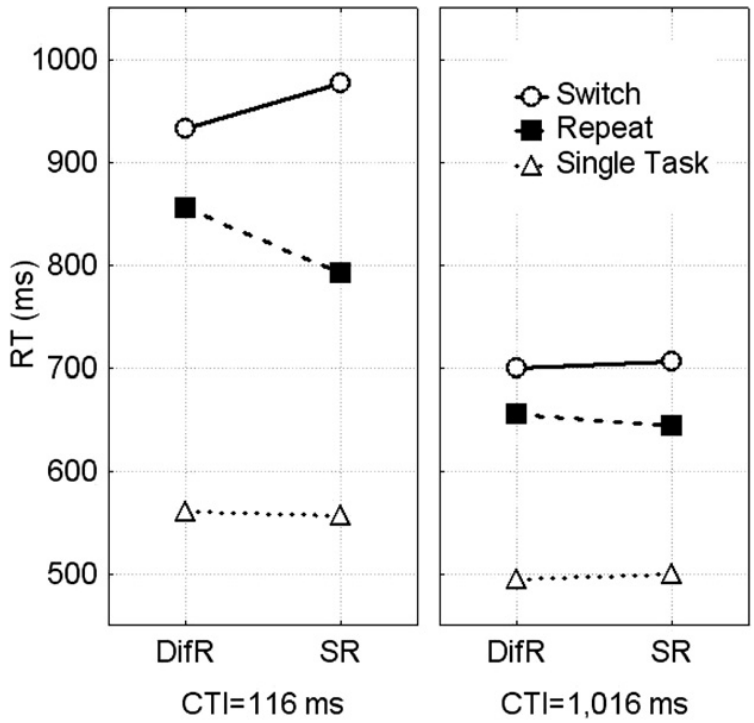

Fig. 5. Mean RT (ms) according to task switch, response-repetition and cue-target interval (CTI)*. *DifR = different response; SR = same response.

4.1.2. Paradigm effects

There was a significant main effect for Paradigm, showing that responses were slower overall in the Shape/Size paradigm (726 ms) as compared with the Vertical/Horizontal paradigm (669 ms). Paradigm also entered into significant interactions with switch, cue-target interval and with the switch by cue-target interval interaction. Nonetheless, in these

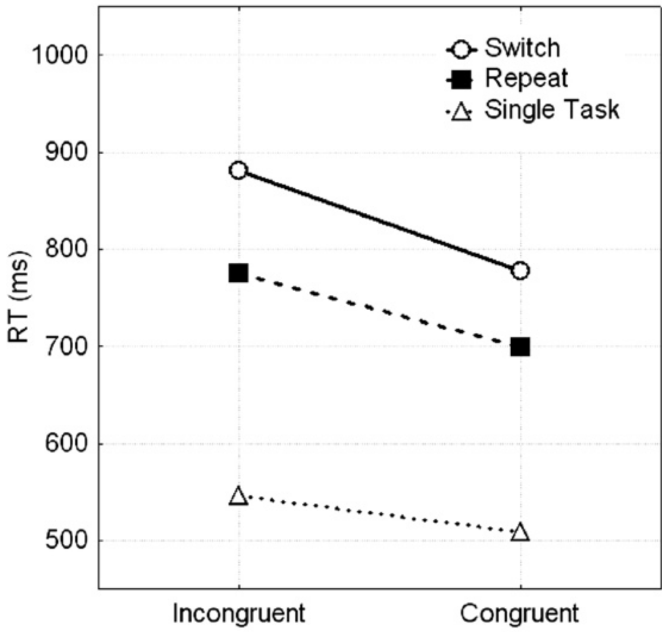

Fig. 6. Mean RT (ms) according to task switch and congruency.

cases, the qualitative trends were the same in the two paradigms. Moreover, in most instances, the size of the interactions involving paradigm was relatively small, around 1% of the overall effect (proportion computed as $SS_{\text{Interaction with paradigm}}/SS_{\text{Overall effect}}$). Paradigm also modulated the effects of congruency and response-repetition.

4.1.2.1. Paradigm × switch × cue-target interval (Fig. 7). We explored this interaction by a series of planned contrasts on switch cost (switch vs. repeat) and on mixing cost as it is usually defined (repeat vs. single-task). Switch cost was significant in the Vertical/Horizontal paradigm, $F(1,94) = 260.23$, 60.40, in the short and long cue-target intervals, respectively, and in the Shape/Size paradigm, $F(1,94) = 158.67$, 44.54, in the two intervals, respectively. Mixing cost was significant in the Vertical/Horizontal paradigm, $F(1,94) = 247.45$, 152.28, in the two intervals, respectively, and in the Shape/Size paradigm, $F(1,94) = 417.40$, 140.48, in the two intervals, respectively. We also ran a series of simple-interaction contrast analyses examining the difference between the paradigms in switching cost and mixing cost. Separate analyses were conducted in each cue-target interval. There was only one significant difference between the paradigms. Mixing cost in the short cue-target interval was larger for Shape/Size than for Vertical/Horizontal, $F(1,94) = 11.32$.

4.1.2.2. Paradigm × congruence. This comparison is of particular interest because the congruency effect in the Vertical/Horizontal paradigm could involve pre-experimental tendencies in the form of a Simon-like effect. Specifically, congruent trials in the Vertical/Horizontal paradigm occupied the same relative position as their response key. Thus, when performing the UP–DOWN task for example, the target's irrelevant positioning along the right–left axis could have influenced performance because the position of the response keys also differed along this axis (see Fig. 2). The overall congruence effect turned out to be somewhat larger in the Vertical/Horizontal paradigm (82 ms) than in the Shape/Size paradigm (63 ms), and it was significant in both cases $F(1,94) = 162.39$, and 165.00, respectively. The fact that the difference between paradigms was relatively slight is in

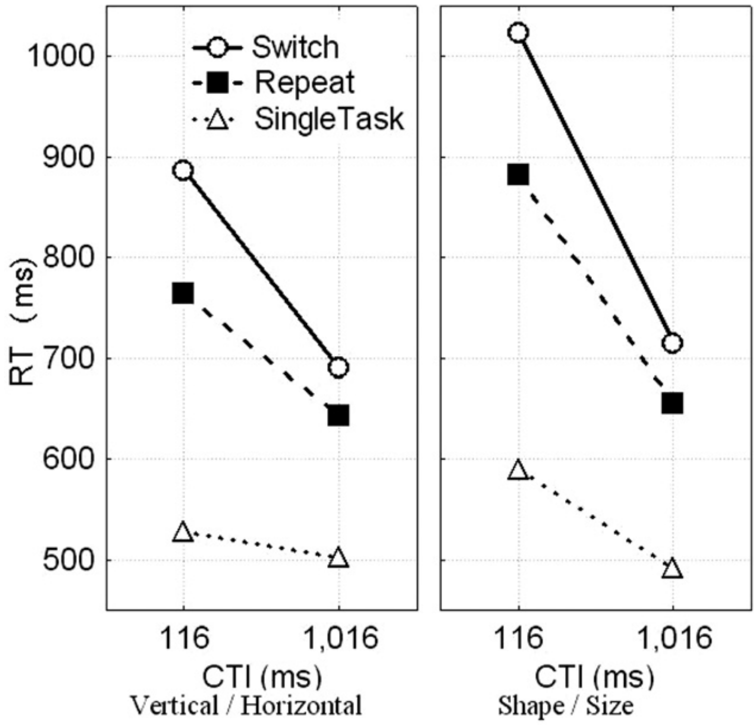

Fig. 7. Mean RT (ms) according to paradigm, task switch and cue-target interval (CTI).

agreement with Meiran's (2005) results showing that the congruency effect observed in the Vertical/Horizontal paradigm was additive with respect to (and, hence, independent of) Simon-like effects.

4.1.2.3. Paradigm × response-repetition. This interaction is the only one which corresponded to different qualitative trends in the two paradigms. The response-repetition effect collapsed over the various conditions was facilitatory in the Shape/Size paradigm (23 ms), $F(1,94) = 32.89$, while it retarded responses in the Vertical/Horizontal paradigm (15 ms), $F(1,94) = 19.03$.

4.1.3. Conclusions

The most notable aspect of the results for the present focus is that we obtained very similar effects in the two paradigms. Because the focus of the present paper is on individual differences in task switching abilities and because the findings are quite standard, we do not discuss their other implications (but see Meiran, 2005, for the discussion of very similar findings).

Although there were some significant differences between the paradigms, in all but one case these differences were quantitative and not qualitative. Moreover, the quantitative differences were usually relatively small as compared to the trend observed across paradigms, with the exception of a larger effect of cue-target interval seen in the Shape/Size paradigm.

4.2. Individual differences

4.2.1. Process scores

Table 2 provides the descriptive statistics for the variables we analyzed in this section. Switch cost was defined as RT in switch trials minus RT in repeat trials. Mixing cost was defined as the averaged RT in switch and repeat trials minus the RT in single-task trials.

Table 2
Descriptive statistics for the variables included in the structural equations models[*]

Measure	Mean (SD)	Range	Reliability[a]	% Participants who had above zero effect
Switch cost, Vertical/Horizontal, short CTI	118 ms (70)	−61 to 348	.44	95
Switch cost, Shape/Size, short CTI	134 ms (104)	−109 to 539	.60	96
Switch cost, Vertical/Horizontal, long CTI	47 ms (60)	−64 to 296	.46	96
Switch cost, Shape/Size, long CTI	58 ms (85)	−76 to 561	.73	83
Mixing cost, Vertical/Horizontal, short CTI	305 ms (161)	−76 to 798	.88	98
Mixing cost, Shape/Size, short CTI	406 ms (164)	97–964	.88	100
Mixing cost, Vertical/Horizontal, long CTI	171 ms (124)	−133 to 678	.76	98
Mixing cost, Shape/Size, long CTI	220 ms (150)	−33 to 812	.83	99
Congruency effect, Vertical/Horizontal, switch	113 ms (102)	−82 to 460	.68	90
Congruency effect, Shape/Size, switch	86 ms (79)	−198 to 333	.50	89
Congruency effect, Vertical/Horizontal, repeat	88 ms (77)	−48 to 387	.62	94
Congruency effect, Shape/Size, repeat	65 ms (68)	−88 to 269	.27	96
Single-task, Vertical/Horizontal	514 ms (123)	380–1239	.97	–
Single-task, Shape/Size	522 ms (113)	371–868	.95	–
Raven's	20 items (5)	5–31	.77[d]	–
Vocabulary	28 items (6)	7–38	.89[b]	–
PS – letter cancellation	238 items (45)	183–376	0.75[b]	–
PS – identical pictures	70 items (14)	33–96	0.81[c]	–
PS – number comparison	24 items (5)	12–35	0.55[c]	–

[a] Spearman-Brown boosted split half reliability (odd vs. even blocks).
[b] From Fischman (1982).
[c] From Ekstrom et al. (1976).
[d] From Carpenter et al. (1990).
[*] CTI = cue-target interval; PS = perceptual speed composite.

Congruency effects were defined as RT in incongruent trials minus RT in congruent trials, both taken from the mixed tasks blocks. The reason why we defined mixing cost differently than it is usually defined (repeat RT minus single-task RT) is that the present definition ensures that the effects represent orthogonal contrasts (see Kray & Lindenberger, 2000). That is, the contrast vector used to define mixing cost in the usual way correlates with the contrast vector used to define switching cost. Finally, internal reliabilities were calculated as Spearman-Brown boosted correlations between the estimates taken from odd and even numbered blocks. To this end, the single-task block was divided for the purpose of analysis into mini-blocks of eight trials each.

Two findings are noteworthy. First, the vast majority of the participants, often 95% or more, exhibited the effects we studied. Second, many of the processing variables exhibited reasonably high reliabilities, especially given the fact that these estimates are based on difference scores, and such scores tend to show depressingly low reliabilities (Cronbach & Furby, 1970). For example, the range of reliabilities for the switching cost measures falls within the same range as that reported by Friedman and Miyake (2004, range = .43–.82), who used comparable procedures.

4.2.2. Structural equations modeling

We used SAS™ CALIS procedure to perform a maximum likelihood estimation of the structural equations models of the covariance matrix. Initial parameter values were taken from the least square solution.

4.2.2.1. Switching cost

4.2.2.1.1. Measurement model. This analysis was conducted on eight variables, representing switching cost in the short/long cue-target intervals, the two paradigms, and in odd and even-numbered blocks. We took cue-target interval into account because Friedman and Miyake (2004) have shown that switching costs taken from short and long cue-target intervals are highly correlated and we wanted to see if this remains true when considering paradigm content.

A model with a single latent variable that explains the common variance was deemed inappropriate, $\chi^2(20) = 51.82$, $p < .0001$, $\chi^2/\text{df} = 2.59$, RMSD = .093 RMSE = .13, NFI = .675. A model with two latent variables defined according to cue-target interval (short vs. long) was slightly better but still inappropriate, $\chi^2(19) = 43.05$, $p < .005$, $\chi^2/\text{df} = 2.26$, RMSEA = .116, RMSD = .087, NFI = .728. A model with two latent variables defined according to paradigm content was considerably better, yet still inappropriate, $\chi^2(19) = 34.53$, $p < .05$, $\chi^2/\text{df} = 1.82$, RMSD = .074, RMSEA = .093, NFI = .782. Only a model in which four latent variables were defined according to both cue-target interval and paradigm content showed reasonably satisfactory fit, $\chi^2(14) = 21.99$, $p = .078$, $\chi^2/\text{df} = 1.57$, RMSD = .056, RMSEA = .078, NFI = .861 (Table 3). The last measurement model indicated that the correlations among the latent variables were numerically very different from one another. A way to test if these differences between correlations were reliable is to compare the last model with an equivalent model in which we imposed an equality constraint on the correlations. This model provided a significantly worse fit than the model in which the correlations were free to vary, $\chi^2(5) = 12.27$, $p = .031$, for the difference between the models. Thus, we can conclude that the correlations between latent variables were significantly different from one another. As one can see in Table 3, three correlations among the latent switch cost variables were high. This includes the two correlations between short and long cue-target intervals, each computed within the same paradigm. The third high correlation was between the latent variables of the two paradigms

Table 3
Maximum likelihood standardized estimates, switching cost model[a]

	Vertical/ Horizontal, short	Vertical/ Horizontal, long	Shape/ Size, short	Shape/ Size, long	Psychometric intelligence
Vertical/Horizontal short odd	.71				
Vertical/Horizontal short even	.39				
Vertical/Horizontal long odd		.60			
Vertical/Horizontal long even		.50			
Shape/Size short odd			.61		
Shape/Size short even			.70		
Shape/Size long odd				.87	
Shape/Size long even				.67	
Raven's					.43
Vocabulary					.25 $p = .07$
Perceptual speed					.71
Correlations among latent variables					
Vertical/Horizontal short		.80	.46	.30 $p = .05$	−.30 $p = .12$
Vertical/Horizontal long			.41	.61	−.37 $p = .06$
Shape/Size short				.70	−.13 ns
Shape/Size long					−.39

[a] All the parameters are statistically significant by a two-sided t-test, except as noted. "Long" and "Short" refer to the duration of the cue-target interval.

given a long cue-target interval. These relatively high correlations contrast with the remaining correlations, which were all low. We therefore tested a model in which we imposed an equality constraint on all the ‘‘high’’ correlations and imposed a different equality constraint on all the ‘‘low’’ correlations, so that the only difference between correlations was among these two groups. This last model provided a fit ($\chi^2(18) = 23.60$, $p = .168$, $\chi^2/\mathrm{df} = 1.31$, RMSD = .060, RMSEA = .058, NFI = .851) that was statistically no worse than the fit provided by the model in which the correlations were all allowed to vary, $\chi^2(4) = 1.61$, ns, for the difference between the models. Thus, we can conclude that the differences among the ‘‘high’’ correlations are non-significant and the same is true for the differences among the ‘‘low’’ correlations. However, the difference between ‘‘high’’ and ‘‘low’’ correlations was significant.

4.2.2.1.2. Correlations with psychometric intelligence. In this analysis we included the three paper-and-pencil measures of intellectual functioning and defined a latent variable that represented their common variance. We fit a model including the four switching cost latent variables and a latent variable representing psychometric intelligence. This model fits the data reasonably well, $\chi^2(34) = 49.92$, $p = .038$, $\chi^2/\mathrm{df} = 1.47$, RMSD = .069, RMSE = .071, NFI = .760. This is presented in Table 3.

The results can be summarized as follows. First, like Friedman and Miyake (2004), we found that, within paradigm content, the correlations between switching costs measures with a long and a short cue-target interval were high. The present results therefore extend Friedman and Miyake's results to WHERE processing. Second, without considering paradigm content, one could reach the conclusion that the cue-target interval is unimportant to consider for individual differences. Our results are novel in showing that the cue-target interval is important from this perspective. The reason is that, when the interval was short, the across-paradigm correlation was low, while when this interval was long, the across paradigm correlation was high. Thus, switching cost given a short cue-target interval does not meet our criteria for a general ability. In contrast, the present analysis shows that switching cost given a long cue-target interval meets the across-content criterion. Third, regarding the second criterion of generality, one cannot speak indiscriminately about switching cost as a general ability, because some switching cost variables showed very low correlations with psychometric intelligence, especially when the cue-target interval was short or the paradigm involved WHAT processing. This finding closely replicates Friedman et al. (2006), who also used a short cue-target interval with WHAT paradigms. In contrast, switching cost given a long cue-target interval also meets the second criterion for generality because it showed moderate but significant correlations with psychometric intelligence.

The high degree of generality of the switching cost measures given a long cue-target interval supports the special status of the famous ‘‘residual SC’’ (Rogers & Monsell, 1995). The reason why this residual switching cost was found to be general according to both of our criteria is not completely clear, and different theories may provide different answers. For example, some theories suggest that, when the cue-target interval is short, cue priming effects contribute to the switching cost (Logan & Bundesen, 2003; Mayr & Kliegl, 2003), and these processes may differ between the two paradigms, especially because the cue types were deliberately designed to differ among our paradigms. Other theories suggest that the switching cost in the short cue-target interval partly represents the time devoted to adjusting the perceptual attentional system(s) in an effort to filter out irrelevant target stimulus information (Meiran, 2000). Such filtering is performed according to the theory because it allows task appropriate responding. Importantly, because these system(s) possibly deal with some form of perceptual attention, they are likely to be affected by the type of perceptual pathway being recruited. Lastly, De Jong's (2000) theory suggests that switching cost given long cue-target intervals represents the proportion of trials in which advance task preparation took place. This proportion does not affect switching cost when the cue-target interval is short according to that theory. One could therefore make the case that the proportion of prepared trials reflects a general ability, which is why switching cost given a long cue-target interval was found to be general.

4.2.2.2. Mixing cost.

4.2.2.2.1. Measurement model. This analysis was conducted on eight variables, defined in the same way as the switching cost variables. A model with a single latent variable yielded a clearly unsatisfactory fit, $\chi^2(20) = 264.28$, $p < .0001$, $\chi^2/\mathrm{df} = 13.21$, RMSD = .191, RMSEA = .360, NFI = .529. A model with two latent variables defined according to cue-target interval (short vs. long) was also unsatisfactory, $\chi^2(19) = 262.61$, $p < .0001$, $\chi^2/\mathrm{df} = 13.82$, RMSD = .190, RMSEA = .369, NFI = .532. A model with two latent variables defined according to paradigm content yielded much better but still not very satisfactory fit, $\chi^2(19) = 83.78$, $p < .0001$, $\chi^2/\mathrm{df} = 4.41$, RMSD = .050, RMSEA = .190, NFI = .851. This fit was not significantly improved when the model included four latent variables defined by both cue-target interval and paradigm content, $\chi^2(14) = 73.64$, $p < .0001$, $\chi^2/\mathrm{df} = 5.26$, RMSD = .044, RMSEA = .213, NFI = .869. The improvement of fit relative to the preceding model was non-significant, $\chi^2(4) = 10.14$, $p = .07$, suggesting that adding two latent variables was unjustified. Although the model with two latent variables defined by paradigm content did not provide a good fit, we did not try to improve it any further because any improvement would need to take into account differences between odd-numbered and even-numbered blocks, which does not seem to be theoretically meaningful. We therefore proceeded with this model to estimate the correlations with psychometric intelligence.

4.2.2.2.2. Correlations with psychometric intelligence. The procedure was the same as that used beforehand. The resultant model, now accounting also for the three paper and pencil measures, yielded poor but a tolerable fit, $\chi^2(41) = 104.39$, $p < .0001$, $\chi^2/\mathrm{df} = 2.54$, RMSD = .052, RMSEA = .128, NFI = .827. This model is presented in Table 4.

Here there was some evidence for a general, across domain ability, accounting for a significant ∼15% of the reliable variance ($r = .39$). Additionally, the correlations with general ability were moderate but significant. The fact that mixing cost showed a moderate degree of generality may be explained by the fact that it represents strategic sustained (Braver et al., 2003) and intentional control (Yehene et al., 2005) as well as the general ability to quickly make a task decision (Koch et al., 2005; Rubin & Meiran, 2005).

4.2.2.3. Congruency effects. Eight congruency-effect variables were analyzed. The estimates were taken from the mixed blocks only (where the congruency effect was large). In the present analyses, we did not define variables according to the cue-target interval but according to switching (switch vs. repeat). The considerations were driven by the mean RT results showing that switching cost and mixing cost were strongly affected by the cue-target interval, whereas the congruency effect was not. Instead, the congruency effect was relatively strongly affected by switching. As before, we also used paradigm content and the odd/even status of the block number to define the variables.

Table 4
Maximum likelihood standardized estimates, mixing cost model[a]

	Vertical/Horizontal	Shape/Size	Psychometric intelligence
Vertical/Horizontal short odd	.87		
Vertical/Horizontal short even	.90		
Vertical/Horizontal long odd	.72		
Vertical/Horizontal long even	.74		
Shape/Size short odd		.87	
Shape/Size short even		.87	
Shape/Size long odd		.78	
Shape/Size long even		.88	
Raven's			.59
Vocabulary			.31
Perceptual speed			.51
Correlations among latent variables			
Vertical/Horizontal		.39	−.38
Shape/Size			−.38

[a] All the parameters are statistically significant by a two-sided *t*-test. "Long" and "Short" refer to the duration of the cue-target interval.

4.2.2.3.1. Measurement model. A model with a single latent variable yielded an unsatisfactory fit,. $\chi^2(20) = 41.35$, $p < .005$, $\chi^2/\mathrm{df} = 2.07$, RMSD = .099, RMSEA = .107, NFI = .704. A model with two latent variables defined by switching (switch vs. repeat) did not improve the fit much, $\chi^2(19) = 40.49$, p < .005, $\chi^2/\mathrm{df} = 2.13$, RMSD = .098, RMSE = .110, NFI = 710. However, a model with latent variables defined according to paradigm content provided a good fit, $\chi^2(19) = 21.68$, $p = .30$, $\chi^2/\mathrm{df} = 1.14$, RMSD = .071, RMSEA = .039. This fit was not improved significantly in a four latent-variable model in which the factors were defined by both switching and paradigm content, $\chi^2(14) = 15.90$, $p = .32$, $\chi^2/\mathrm{df} = 1.14$, RMSD = .063, RMSEA = .038, NFI = .886. The difference in fit between the last two models was non-significant, $\chi^2(5) = 5.78$, $p = .33$. Thus, the chosen model was the two factor model with factors defined by paradigm content. Because the next model had convergence problems (see below) we report the inter-factor correlation here. (Convergence problems indicate that the search algorithm used to solve the set of equations failed identifying an optimal solution). This correlation was $r = .12$ and non-significant, mirroring a similar finding by Ward et al. (2001) who studied Stroop and Stroop-like congruency effects.

4.2.2.3.2. Correlations with psychometric intelligence. A single model such as that used beforehand had convergence problems. We therefore estimated the correlations in two separate models, each including only one of the two congruency-effect latent variables (Table 5). The results indicate low and non-significant correlations between the congruency effect and psychometric intelligence. The model fit indices were, $\chi^2(13) = 12.63$, 10.45, $p > .5$, $\chi^2/\mathrm{df} = .97$, .80, RMSD = .053, .063, RMSEA = .000, .000, NFI = .899, .783, for the model with the Vertical/Horizontal and Shape/Size variables, respectively.

4.2.2.4. Single-task performance. An important baseline to which we could compare the results reported so far is the performance in the single-task block. This performance presumably reflects choice RT, a measure that arguably does not involve executive control as

Table 5
Maximum likelihood standardized estimates, task rule congruency model[a]

	Vertical/ Horizontal	Shape/ Size	Psychometric intelligence (Vertical/Horizontal model above, Shape/Size model, below)
Vertical/Horizontal switch odd	.73		
Vertical/Horizontal switch even	.70		
Vertical/Horizontal repeat odd	.67		
Vertical/Horizontal repeat even	.62		
Shape/Size switch odd		.99	
Shape/Size switch even		.34	
Shape/Size repeat odd		.18, ns	
Shape/Size repeat even		.29	
Raven's			.67
			.51
Vocabulary			.28 $p = .06$
			.26 $p = .08$
Perceptual speed			.46
			.62
Correlations among latent variables[b]			
Vertical/Horizontal		.12 ns	−.17 ns
Shape/Size			.20 ns

[a] All the parameters are statistically significant by a two-sided t-test, except as noted.
[b] Two models were fit separately: a model with Shape/Size variables and a model with Vertical/Horizontal variables, see text for details.

much as do switching cost, mixing cost, and the congruency effect. Of course, choice RT has a choice component, and a goal memory component, both of which may involve executive functioning. Moreover, our single-task condition appeared after the mixed tasks condition. As a result, the stimuli could remind the participants of the wrong task and impair performance. Despite all that, we believe that very few people would dispute the fact that the single-task condition demanded much less cognitive control than the mixed tasks conditions.

4.2.2.4.1. Measurement model. Given the relatively small number of trials (only the single-task block was analyzed) and the lack of cue-target interval effects, only four variables were included in this analysis, representing the odd and even numbered mini-blocks of the two paradigms, respectively. (Each consecutive 10 trials defined a mini-block). The first model included one latent variable and it failed, $\chi^2(2) = 146.71$, $p < .0001$, χ^2/df = NFI = .661, RMSEA = .877, RMSD = .159. The second model had two latent variables defined according to paradigm content and it was more succeessful, $\chi^2(1) = 6.13$, $p = .013$, $\chi^2/\mathrm{df} = 6.13$, RMSD = .006, RMSEA = .233, NFI = .986 see Table 6.

4.2.2.4.2. Correlations with psychometric intelligence. The final model estimated the correlations between the two single-task latent variables and psychometric intelligence. This model (Table 6) fit the data well, $\chi^2(11) = 11.36$, $p = .41$, $\chi^2/\mathrm{df} = 1.03$, RMSD = .043, RMSEA = .019, NFI = .975. The results show that single-task performance represents a general ability according to one criterion (high correlation across paradigm content) but not according to the other criterion. The correlation with psychometric intelligence was low and non-significant for the Shape/Size task.

Table 6
Maximum likelihood standardized estimates, single-task model[a]

	Vertical/Horizontal	Shape/Size	Psychometric intelligence
Vertical/Horizontal odd	.95		
Vertical/Horizontal even	1.00		
Shape/Size odd		1.00	
Shape/Size even		.92	
Raven's			.43
Vocabulary			.29
Perceptual speed			.68
Correlations among latent variables			
Vertical/Horizontal		.54	−.41
Shape/Size			−.20 ns

[a] All the parameters are statistically significant by a two-sided *t*-test, except as noted.

5. General discussion

Our question in the present work was whether there is a general task switching ability. In answering this question, we adopted two criteria for generality: shared variance across paradigm content and shared variance with psychometric intelligence. Our results can be summarized as follows. At the level of the group means, the Vertical/Horizontal and Shape/Size paradigms yielded very similar results. Additionally, a vast majority of the participants showed the three effects that we measured: switching cost, mixing cost and congruency effects. The most important results concerning the correlations between latent variables are summarized in Table 7. Before discussing these results, we remind the reader that the mixing cost model that we ended up endorsing did not provide a good fit and therefore the results concerning this index should be treated with caution.

The results suggest that the most general ability according to our criteria is tapped by switching cost given a long cue-target interval, followed by mixing cost. In both cases, there was a significant amount of shared reliable variance across task content, 37% and 15%, respectively. The single-task performance results, while showing a high generality in one respect, show a low degree of generality in another respect. Specifically, there was a relatively large proportion of shared variance across paradigm content (29%) and a significant amount of shared variance between Vertical/Horizontal processing and psychometric intelligence (17%). Nonetheless, the shared variance between psychometric

Table 7
Correlations among latent variables: summary[**]

	Cross content correlation	Correlations with psychometric intelligence	
		Vertical/Horizontal	Shape/Size
Switch cost-short CTI	.46[*]	−.30	−.13
Switch cost-long CTI	.61[*]	−.37 $p = .06$	−.39[*]
Mixing cost	.39[*]	−.38[*]	−.38[*]
Congruency effect	.12	−.17	.20
Single-task performance	.54[*]	−.41[*]	−.20

[*] $p < .05$ by a two sided *t*-test.
[**] CTI = cue-target interval.

intelligence and Shape/Size processing was small (4%) and non-significant. The difference between the Shape/Size and Vertical/Horizontal results could possibly be explained by the hypothesized underlying functional neuroanatomy. Specifically, general fluid abilities seem to involve dorsal aspects of the PFC (e.g., Duncan et al., 2000), as presumably does the Vertical/Horizontal task. Furthermore, the Shape/Size task probably involves the dorsal aspects of the cortex to a lesser degree.

There were two surprising results. One result concerns switching cost given a short cue-target interval. According to some authors (e.g., De Jong's (2000); Meiran, 1996, 2000; Rogers & Monsell, 1995), the ability to prepare for a task switch is a hallmark of executive functioning. These theories therefore focus on the reduction in switching cost as a result of extending the cue-target interval. Accordingly, one may predict a greater degree of executive involvement in cases in which the cue-target interval is short, when the duration of executive processing presumably adds to RT. The present results challenge this line of reasoning. However, it is unclear in exactly what way they challenge it. Several explanations come to mind. First, the assumption regarding preparation being a hallmark of executive functioning might be maintained. It is only the notion that this ability is general that needs to be abandoned. Rather, one may assume analogous, yet separate executive systems operating in their respective pathways, which extend into the prefrontal cortex (in agreement with Smith and Jonides's, 1999, results). Second, one may argue that performance in the short CTI is reactive because there is insufficient time to generate an appropriate anticipatory set in that condition. In contrast, performance with a long cue-target interval is proactive. Accordingly, the more reactive processing (short cue-target interval) is not general in nature because it does not (yet) involve executive functioning. This line of reasoning accords with De Jong's (2000) theory. According to this theory, the within participant variability in switching cost given long preparation intervals reflects the proportion of trials in which preparation took place. Whether one prepared or did not prepare (which presumably reflects executive functioning) does not contribute to the switching cost when the cue-target interval is short, according to this theory. Another somewhat related idea is that responding is always reactive in the sense that participants use whatever control strategy is currently available to them rather than wait until a proper preparatory set is established. When the cue-target interval is short, this preparatory set is perceptual in nature, and thus pathway specific. When the cue-target interval is long, the preparatory set focuses on action and action representation (what pressing a given key means), and as such is more general in nature (Meiran, Kessler, & Adi-Japha, submitted for publication). Finally, Logan and Bundesen's (2003) and Mayr and Kliegl's (2003) results are relevant here. These authors showed that switching cost in the short cue-target interval partly reflects the time taken to process the task cue, and this processing may not be accomplished by executive processing.

The other surprising result concerns the congruency effect, which did not show any generality. One reason why this finding is surprising is that the congruency effect seems to be the best indicator for response conflict within the task switching paradigm and that resolving response conflict is considered by many authors to be a core executive function (e.g., Botvinick et al., 2001). It is important to note that an analogous result was obtained by Ward et al. (2001), who used Stroop and Stroop-like tasks, considered to be more canonical conflict measures as compared with the congruency effect.

The results concerning the congruency effect are important in ruling out an alternative explanation for the other results based on an intriguing hypothesis by Garlic (2002).

According to Garlic, the correlations between the measures of cognitive functioning may not be due to a shared of cognitive operation or even shared brain regions. Instead, the correlation may be driven by a common biological process, such as the ability to develop dense dendrites. Had such a factor been responsible for the correlations we observed, it should have produced high correlations across the board. The fact that some correlations were relatively high while others were negligible suggests that the correlations were indeed driven by shared cognitive processing.

5.1. Comparison to previous studies

The present study has several noteworthy limitations, including a relatively moderate sample size and the use of only one paradigm to represent a domain. The fact that, in spite of these limitations, we were able to replicate key findings in the literature, reassures us in our interpretation of the novel results. The findings that we replicated include the following:

1. Within paradigm content, there was a strong correlation between switching cost measured with a short and a long cue-target interval (Friedman & Miyake, 2004).
2. Like Ward et al. (2001) who studied Stroop and Stroop-like effects, we found negligible correlation between the measures of conflict (the congruency effect) across paradigm content.
3. Like Friedman et al. (2006), we found negligible correlations between switching cost given a short cue-target interval and psychometric intelligence.
4. Since old age is associated with decreased fluid abilities (e.g., Cattell, 1971), the correlations with aging may indirectly reflect on the correlations with psychometric intelligence among young adults. In that respect, switching cost given a long cue-target interval and mixing cost showed the greatest proportion of age-related variance in Meiran et al.'s (2001) study, a pattern which resembles the pattern of correlations with psychometric intelligence is reported in the present paper.

6. Conclusions and implications

In the present work, we were concerned with the question of whether there is a general task switching ability. The results of the congruency effect and switching cost given a short cue-target interval clearly refute the notion of a general ability. These results show that if such a general ability exists, it explains a negligible proportion of the reliable variance. In contrast, the residual switching cost (switching cost given a long cue-target interval) and mixing cost represent a general ability, operating across paradigm content. It is important to note that even in these cases the latent variables we identified showed a considerable degree of paradigm specificity. Moreover, we cannot completely rule out the possibility that this shared variance was due to the fact that the "non-executive" processes in our paradigms may have involved similar brain mechanisms. The reason is that, while the paradigms probably engaged the dorsal and ventral routes differentially, this differentiation is unlikely to have been complete because in the Shape/Size paradigm, the size task potentially involves the WHERE pathway. However, if the paradigm commonalities we report are due to shared non-executive processing, it is difficult to understand why some measures of control functioning showed substantial across-paradigm shared variance while others did not.

The fact that only some functions show generality whereas others do not is not easy to reconcile with the idea of a central system. To do so, one must assume that if such a central system exists, it is related neither to the congruency effect nor to the switching cost, measured without advance preparation. Additionally, what we found with these two effects has implications for neuroimaging and the assessment of special populations. In such cases, the use of converging evidence from multiple paradigms seems to be essential.

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