

Control processes in voluntary and explicitly cued task switching

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Explicitly cued task switching slows performance relative to performing the same task on consecutive trials. This effect appears to be due partly to more efficient encoding of the task cue when the same cue is used on consecutive trials and partly to an additional task-switching process. These components were examined by comparing explicitly cued and voluntary task switching groups, with external cues presented to both groups. Cue-switch effects varied in predictable ways to dissociate explicitly cued and voluntary task switching, whereas task-switch effects had similar characteristics for both instructional groups. The data were well fitted by a mathematical model of task switching that included a cue-encoding mechanism (whereby cue repetition improves performance) and an additional process that was invoked on task-switch trials. Analyses of response-time distributions suggest that this additional process involves task-set reconfiguration that may or may not be engaged before the target stimulus is presented.

Keywords: Cue encoding; Task switching; Voluntary task switching

Cognitive control operations that govern the ability to shift from one task to another, particularly under conditions that afford multiple tasks simultaneously, can be separated into two kinds. One type of operation involves endogenously controlled shifts in processes that direct attention to the correct aspect of a stimulus display and retrieve and apply the appropriate stimulus–response mapping rules to generate the correct response (e.g., Mayr & Kliegl, 2003; Meiran, 2000; Rogers & Monsell, 1995). An additional class of processes is exogenous in nature and appears to operate without top-down control. These processes may include interference from residual activation of previous task sets, priming of performance when a task

is repeated, and, when the task switches, negative priming of a previously ignored task (e.g., Allport, Styles, & Hsieh, 1994; Arrington & Logan, 2004a; Sohn & Carlson, 2000; Waszak, Hommel, & Allport, 2003). A question of central importance in developing an account of how efficient switching between tasks is accomplished concerns the relative contributions made by these two types of processing.

From one theoretical perspective, proposed by Logan and colleagues (e.g., Logan & Bundesen, 2003; Schneider & Logan, 2005), task switching is orchestrated by purely exogenous processes. They proposed that the cost of switching tasks, when explicit cues are used to indicate which task

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We thank Marnie Jedynak for assistance with data collection and analysis and Gordon Logan for helpful suggestions on model fitting.

This research was supported by the Natural Sciences and Engineering Research Council of Canada through a discovery grant to Michael Masson and a postgraduate scholarship to Sarah Carruthers.

to perform, is entirely due to a benefit arising from repeating the task cue on task-repeat trials, relative to changing the cue to indicate a task switch. Using a procedure in which two different cues were associated with each task, Logan and Bundesen (2003) were able to separate the respective contributions of cue switching and task switching to the overall cost of a task change. When the cue changed between trials but the same task was performed, a large response time increase was found relative to when the cue (and task) remained the same across consecutive trials. Moreover, this *cue-switch* cost dissipated as the interval between the cue and the target stimulus increased. The reduction in switch cost with increasing cue-target interval is a standard finding, often attributed to endogenously controlled preparation processes (e.g., Mayr & Kliegl, 2000; Rogers & Monsell, 1995). Logan and Bundesen, however, ascribed the reduction in cue-switch cost over the cue-target interval to a fading benefit of repetition priming of cue encoding; as time for processing the task cue increases, even the encoding of a changed cue may be completed before the target arrives, leaving no room for a repetition benefit.

With the cost of switching cues assessed separately from the cost of switching tasks, Logan and Bundesen (2003) discovered that there was no *task-switch* cost in their experiments; the entire cost of switching was due to a change in task cue. Logan and Bundesen (2003, 2004) captured these results in a mathematical model that successfully fitted their data by assuming that cue encoding was more efficient under cue repetition, but without assuming any cost due to a change in the task. Further, the benefit of repetition in cue encoding could be attributed to an exogenous process of repetition priming. Taking together their modelling and empirical results, Logan and Bundesen concluded that there was no evidence for endogenously controlled processes in adapting to explicitly cued task switches.

Using a very similar method involving a 2:1 mapping of cues to tasks, Mayr and Kliegl (2003) also found a substantial cue-switch cost that faded as the cue-target interval increased. But unlike Logan and Bundesen (2003, 2004), they observed

a significant task-switch cost that could not be attributed to cue switching. This task-switch cost was dissociated from the cue-switch cost by virtue of being constant across increasing cue-target intervals. Mayr and Kliegl accounted for these results by proposing two serial processing stages: (a) cue-based retrieval of task rules, possibly based on intentional preparation, and (b) application of the task rules to the target stimulus, which is affected by exogenous processes such as response priming when a task is repeated. The second stage is more efficient when the task is repeated but it cannot be initiated until the target stimulus arrives.

The discrepancy between the Mayr and Kliegl (2003) and the Logan and Bundesen (2003) studies with respect to the presence versus absence of a task-switch cost is of particular interest because task-switch cost is potentially associated with endogenous cognitive control operations. An important difference between the two studies lies in the nature of the task cues that were used. Mayr and Kliegl used arbitrary cues (single letters) that had no relationship to the task stimuli (coloured shapes), whereas Logan and Bundesen used related word phrases to cue each task (e.g., *parity* and *odd-even* were used to cue a parity-judgement task applied to digits). A comparison of the size of switch cost produced when cues are meaningful words versus arbitrary symbols has shown that switch costs are smaller with meaningful cues (Logan & Schneider, 2006; Schneider & Logan, 2011).

A clue as to why the nature of the task cues might modulate task-switch cost was provided by Arrington, Logan, and Schneider (2007). Although they, like Logan and Bundesen (2003), used meaningful cues in a 2:1 mapping of cues to tasks, they required subjects to make an overt response to each task cue before the target stimulus was presented. This requirement led to the emergence of a task-switch cost in addition to a cue-switch cost. Moreover, the task-switch cost was particularly strong when the two cues for a task required the same overt response. This arrangement was assumed by Arrington et al. to induce subjects to generate a relatively abstract task

representation common to either cue for a task. We suggest that a similar kind of task representation is invoked when arbitrary task cues are used, as in the Mayr and Kliegl (2003) study, and that this task representation is involved in the observed task-switch cost. Arrington et al. speculated that this cost may reflect any of a number of possible processes, including endogenous or exogenous task reconfiguration and negative priming of a task set induced by the target stimulus.

Additional clues as to the nature of the task-switch cost that can be separated from cue-switch cost may be found in studies of voluntary task switching. In these experiments, subjects are not explicitly cued to perform a particular task on each trial, but instead are free to choose which task to perform (with constraints such as performing each task about equally often overall). Ostensibly, there are no cue encoding operations in this procedure, but when the task chosen by the subject entails a task switch, endogenous control operations similar to those implicated in cued task switching contexts may be required. Arrington and Logan (2004b) showed that subjects engaged in a voluntary-switch paradigm took longer to respond after switching tasks and that this switch cost was reduced as the interval between the response on trial N and presentation of the target on trial $N + 1$ was increased (see also Liefoghe, Demanet, & Vandierendonck, 2009). This reduction in switch cost might reflect the greater opportunity to prepare for a switch with increasing response-stimulus interval (RSI).

In a direct comparison of task-switch cost under explicitly cued and voluntary task-switching instructions, Arrington and Logan (2005) used a 2:1 mapping of cues to tasks for the explicit group. The cues were arbitrary colours, and, consistent with the observation above regarding the use of arbitrary cues, they obtained evidence for a task-switch cost in addition to a cue-switch cost. The task-switch cost under explicit cues, however, did not vary across cue-target intervals (consistent with Mayr & Kliegl, 2003), whereas the voluntary group, for whom a random colour cue was presented as a warning signal at the start of each trial, showed decreasing task-switch costs as the

cue-target interval increased. Arrington and Logan suggested that the task-switch cost for the voluntary group, which operated with little environmental support (warning cues were not correlated with the task performed), reflected endogenous control. For the explicitly cued task-switch cost, where the response was fully specified by the combination of task cue and task-specific stimulus-response mapping rules, Arrington and Logan argued that there is little need for endogenous control and ascribed this task-switch cost to priming of exogenous response-retrieval processes on task-repetition trials.

There are reasons, however, to suggest that the task-switch costs seen with explicitly cued task switching and with certain versions of voluntary switching arise from a common endogenous control mechanism. First, under voluntary task switching, the reduction in task-switch cost with increasing RSI appears to be due in large part to the inclusion of task-selection processes in the measure of response time. A much weaker reduction in task-switch cost with voluntary switching was found by Arrington and Logan (2005, Experiment 6) when an overt task-selection response was made at the start of each trial, so that time to respond to the target stimulus did not include task-selection operations. A weak or absent reduction in task-switch cost with increasing RSI is consistent with task-switch costs found by Arrington and Logan using the explicit-cueing paradigm (see also Mayr & Kliegl, 2003). Second, Yeung (2010) demonstrated that both voluntary-switch and explicitly cued subjects showed *switch-cost asymmetry* when the two cued tasks differed in difficulty (i.e., switching from the more difficult task to the less difficult one required more time than switching in the other direction). Although Liefoghe, Demanet, and Vandierendonck (2010) did not find a switch-cost asymmetry among subjects operating under voluntary task-switch instructions, it turned out that their two tasks (word reading and colour naming) did not differ in difficulty, so it is not surprising that no asymmetry in switching was obtained.

We propose, then, that a common set of mechanisms underlies the task-switch cost that can be

measured independently of cue-switch cost for both endogenously controlled (voluntary) and explicitly cued task switches. To test this proposal, we conducted an experiment that compared voluntary and explicitly cued task switching while using identical cue and target presentations for both groups. For the voluntary-switch group, subjects made an overt response to indicate their task selection for a particular trial, which caused a visual cue associated with the selected task immediately to appear. This procedure ensured that task-choice operations would not contaminate measurement of the time required to respond to the target stimulus (Arrington & Logan, 2005). A variable cue-target interval ensued, and the target stimulus was then presented. Each subject in the explicitly cued group was yoked to a subject in the voluntary-switch group and received the corresponding cue-target sequences. The time interval between a voluntary-switch subject's response to the target on one trial and that subject's task-selection response on the next trial varied, and that variation was replicated in the intertrial intervals experienced by the yoked, explicitly cued subject.

The introduction of cues in the voluntary-switch group was intended to create identical cueing experiences for voluntary-switch and explicitly cued subjects. We were then able to use the formal cue-encoding model developed by Logan and colleagues (e.g., Arrington & Logan, 2004a; Logan & Bundesen, 2003, 2004) to assess separately the contribution of cue switches and task switches to switch costs for both groups. This approach allowed us to compare voluntary and explicitly cued task switching using a formal model to decompose switch costs into component processes. Three versions of this model were considered, Models 2, 2 + 1, and 3. In Model 2, switch costs are assumed to result from repetition priming of the cue-encoding process on trials where the cue from the previous trial is repeated. This effect is captured in the model by a cue-encoding time parameter (designated by μ in the equations below) reflecting the mean time required to encode a task cue. This parameter takes on a smaller value for cue repetitions than for cue switches due to repetition priming from the

previous trial. Time available for cue encoding is manipulated by varying the stimulus-onset asynchrony (SOA) between the cue and the target. Response time in the cue-repetition case is then modelled as

$$RT = RT_{\text{Base}} + \mu_r \times \exp(-SOA/\mu_r), \quad (1)$$

where RT_{Base} is the time required to process and respond to the target stimulus, and the function \exp refers to the constant e raised to the indicated power. With longer SOA, less time is required for cue encoding operations beyond the SOA interval, producing a smaller contribution to measured response time. For the cue-switch condition, regardless of whether a task switch results, response time is defined as

$$RT = RT_{\text{Base}} + \mu_s \times \exp(-SOA/\mu_s), \quad (2)$$

where μ_s is constrained to be greater than μ_r . In this model, the only source of switch cost is cue-encoding time, and no task-switch cost is assumed.

In Model 2 + 1, in addition to repetition priming of cue encoding, an endogenous task-switching operation is included. When the task cue is repeated from the previous trial, no task-switch process is invoked, and response time is determined by Equation 1. When the task is repeated from the previous trial but the cue changes, response time is defined by Equation 2. When the task and cue both change, cue-encoding and task-switch operations are captured as follows:

$$\begin{aligned} RT = & RT_{\text{Base}} + \exp[-SOA/\mu_s] \times (\mu_s + \mu_t) \\ & + \frac{1/\mu_s}{1/\mu_s - 1/\mu_t} \\ & \times (\exp[-SOA/\mu_t] - \exp[-SOA/\mu_s]) \\ & \times \mu_t, \end{aligned} \quad (3)$$

where μ_t is the mean time required for the task-switching operation to complete. Note that in Equation 3, the task-switching process occurs during the SOA so that with a sufficiently long SOA, the operation may be completed, and no

task-switch cost will be obtained—that is, there will be no *residual switch cost* attributable to task-switching operations (cf. Meiran, 1996; Rogers & Monsell, 1995).

Finally, in Model 3 there is a benefit for repetition of the cue (as in Model 2) but also a benefit for target processing when the task repeats. Because this benefit applies to operations performed on the target, it is not dependent on the duration of the SOA but instead is captured by the RT_{Base} component of the model. This parameter takes on a larger value for task switches than for task repetitions, so the three conditions, cue repetition, cue switch, and task switch, respectively, are modelled as follows:

$$RT = RT_{\text{BaseRep}} + \mu_r \times \exp(-SOA/\mu_r), \quad (4)$$

$$RT = RT_{\text{BaseRep}} + \mu_s \times \exp(-SOA/\mu_s), \quad (5)$$

$$RT = RT_{\text{BaseSw}} + \mu_s \times \exp(-SOA/\mu_s). \quad (6)$$

Target processing time is assumed to take longer with a task switch, so we have the constraint that $RT_{\text{BaseSw}} > RT_{\text{BaseRep}}$ (where Sw denotes switch, and Rep denotes repetition).

We considered a conceptual variant of Model 3, in which a possibly endogenous processing operation sensitive to a task switch may be carried out during the SOA and so is not directly tied to the processing of the target stimulus. Unlike Model 2 + 1, however, the completion of this process is not guaranteed, even with a long SOA, yielding a residual switch cost. This proposal was motivated by De Jong's (2000) *failure to engage hypothesis*, whereby it is assumed that endogenous reconfiguration of a task set is needed to accommodate a task switch, but is sometimes postponed until the target stimulus occurs. Engagement in preparation may fail for several reasons, among them the failure to maintain the intended task set either because of competition from the previous task set (Bub, Masson, & Lalonde, 2006; De Jong, Berendsen,

& Cools, 1999) or because of decay (Altmann & Gray, 2002), and interference with formation of top-down intentions caused by competition between tasks (Poljac & Yeung, 2014).

This proposal leads to a clear prediction regarding response-time distributions. When reconfiguration processes are engaged, and a relatively long SOA is provided, subjects should be equally prepared to respond in task-repeat and task-switch conditions, leading to no task-switch cost. Completion of preparation prior to arrival of the target stimulus also means that response time will be relatively short. Therefore, if there is a situation in which little or no task-switch cost occurs, it should be when response latency is particularly short. If the subject does not engage a reconfiguration process on a switch trial, then the system is unprepared when the stimulus arrives, leading to a relatively long response time and a clear advantage for the task-repeat condition. Thus, with a long SOA, there should be little or no task-switch cost among the shortest response latencies, but a cost should emerge for longer latencies (where task reconfiguration presumably was not engaged). In addition to fitting formal models to response-time means, then, we also examined the response-time distributions for evidence of failure to engage in an endogenous task-set reconfiguration operation.

In our experiment, subjects were tested under voluntary-switch or explicitly cued instructions. Two different colour cues were assigned to each of two tasks to enable separate assessment of the cost of switching cues (without switching tasks) and the cost of switching tasks, which typically requires a cue switch (Logan & Bundesen, 2003; Mayr & Kliegl, 2003). Arbitrary colour cues were used so that not only cue-encoding efficiency but also a potentially endogenous component of task preparation might contribute to performance (Arrington et al., 2007; Logan & Bundesen, 2004; Mayr & Kliegl, 2003). Given that task cues were crucial for the explicitly cued group but redundant for the voluntary-switch group, there was every reason to expect cue-switch costs to arise only in the former group. Our primary interest lies with the question of whether task-switch

costs would be found for one or both groups and whether these costs would be qualitatively similar, with respect to both modulation by cue–target interval and the pattern of costs across the response time distributions.

No previous work has reported the use of overt, redundant cues under voluntary-switch instructions, so it was not entirely clear how a redundant cue would influence performance in the voluntary-switch group. Therefore, we included an additional voluntary-switch group in which subjects were not presented redundant task cues, although they were required to make an overt response on each trial to indicate which task had been selected for that trial.

EXPERIMENTAL STUDY

Method

Subjects

A sample of 104 students received extra credit in an undergraduate psychology course for participating in the experiment. Thirty-six subjects were randomly assigned to the *voluntary-switch* (with task cues) group and 36 to the *explicitly cued* group based on their order of appearance in the laboratory. The first subject was assigned to the voluntary-switch group, and the next subject was yoked to the first subject and tested under explicitly cued instructions. The remaining subjects were similarly designated for participation in the two groups. The remaining 32 subjects, tested in a later academic term, participated under voluntary-switch instructions without task cues (the *voluntary/no-cue* group). A sample size of approximately 30 subjects per group in this paradigm provides substantial sensitivity to distinguish between null effect and effect-present models using estimated Bayesian posterior probabilities as described in the Results section.

Procedure

Subjects were tested individually in a quiet room. Stimulus display was controlled by a Macintosh computer. Subjects made parity and magnitude

judgements to the eight digits ranging from 1 to 9, omitting 5. They were first given two practice blocks of 16 trials in which each task was performed in isolation. For subjects in the two cued groups (voluntary switch and explicitly cued), each trial began with a fixation cross at the centre of the computer monitor, which was replaced after 500 ms by a coloured rectangular frame (blue, green, red, or yellow). Two colours were used for each task, and this assignment was counterbalanced across subjects. After a delay of 150 or 1200 ms, the target digit appeared in the centre of the frame. These two intervals were selected to cover the range typically examined in task-switching experiments. Eight consecutive trials of each block used one of these two delays, and the other eight trials used the other delay. Target digits were presented in a random sequence across the trials within a block. Responses were made on a button box using the index and middle fingers of the left hand for one task and the index and middle fingers of the right hand for the other task. Assignment of tasks to hands was counterbalanced across subjects.

Each trial started automatically 600 ms after a response was made on the preceding trial. We did not vary the intertrial interval to equate total time between response completion and target onset for the next trial as others have done (e.g., Arrington & Logan, 2004b) because we could not precisely control the interval between response on one trial and the onset of the task cue for the next trial. This interval could not be controlled because subjects in the voluntary-switch group freely determined how long they would take to decide which task to perform on a given trial, and variation in those choice times would work against precise control over response–cue intervals. Moreover, Mayr and Kliegl (2003) showed that in an explicit cueing paradigm using arbitrary task cues, variations in response–target interval did not affect switch costs when cue–target interval was held constant. A display at the bottom of the computer monitor was continuously available, which indicated which hand and finger were to be used to respond for each task and stimulus category, respectively, but no colours were included in that display.

Following single-task training, subjects in the voluntary-switch group were told that they would be performing the two different tasks on a series of stimuli and that they were to choose which task to perform at the beginning of each trial. They were instructed to select each task about equally often and in a random sequence. Each trial was run as in the training phase, except that the fixation cross remained in view until the subject pressed the left-most or right-most response button to indicate which task he or she selected for that trial. On this button press, the fixation cross was replaced by a coloured rectangle, randomly selected from the two possible colours for that task, and the trial was completed as in the training phase. Thus, when a subject elected to perform the same task as that on the previous trial, the current trial was equally likely to be a cue-repeat or a cue-switch trial. Subjects began a block of 16 practice trials with the same cue-target interval as that used for the first training block. Then the first critical block of 40 trials with that same cue-target interval was presented, followed by seven additional blocks of 40 trials alternating between the two cue-target interval values (150 and 1200 ms), for a total of 320 critical trials. Cue-target interval was manipulated in blocks rather than as a random sequence so that subjects would be encouraged to engage in endogenous task preparation when they could be sure that sufficient time would be available (Rogers & Monsell, 1995). A break was provided after every 80 critical trials.

The procedure for subjects in the explicitly cued group was the same as that for voluntary-switch subjects, except that instead of selecting the task for each trial, these subjects were presented a task sequence and colour-cue sequence yoked to one of the subjects in the voluntary-switch group. The fixation cross on each trial remained in view for the same amount of time as that required by the corresponding voluntary-switch subject to press a button to indicate the task selection, thereby yoking subjects with respect to response-to-cue interval.

Subjects in the voluntary/no-cue group were treated exactly the same as subjects in the voluntary-switch group, except that no colour cues were presented.

Results

The inferential tests we report were carried out by computing an estimate of the Bayesian posterior probability that the observed data more strongly support a model that assumes that an effect is present relative to a model that assumes no effect (Wagenmakers, 2007; see also, Masson, 2011). This procedure uses the Bayesian information criterion (BIC) to estimate the posterior odds that the observed data favour one model over the other. In applying this method one assumes, as is done in the standard application of analysis of variance, that errors of measurement are normally distributed. We also make the reasonable assumption that the prior odds equal 1, indicating that the null and alternative models are deemed equally likely before the data are taken into account. The estimated posterior odds can be converted to posterior conditional probabilities, which we designate as p_{BIC} , representing the probability that a particular model is preferred, given the observed data. Because there are only two competing models, their posterior probabilities sum to 1.0. The values of these probabilities can be placed into broad categories that characterize the strength of evidence, as suggested by Raftery (1995): .50-.75 = weak; .75-.95 = positive; .95-.99 = strong; >.99 = very strong. To guide the reader who is more comfortable with null-hypothesis significance testing conventions, we note that any effect that we present as being more strongly supported than the null model with at least positive evidence would be significant at least at the .05 level and more often than not beyond the .01 level.

Task choice

Task selection characteristics among the voluntary-switch subjects were similar to what has been reported in previous work on voluntary task switching (e.g., Arrington & Logan, 2004b, 2005; Vandierendonck, Demanet, Liefvooghe, & Verbruggen, 2012). Namely, subjects showed a slight tendency to select the parity task more frequently than the magnitude task (mean for parity task = .513, 95% confidence interval = $\pm .009$), and subjects were somewhat more likely to repeat

a task than to switch (mean switch probability = .403, 95% confidence interval = $\pm .044$). Electing to switch tasks was not modulated by the cue–target interval in effect for the current block of trials (mean switch probability = .405 and .401 for the short and long interval conditions, respectively), as indicated by a Bayesian analysis that showed positive support for a null difference model, $p_{\text{BIC}} = .849$. Subjects in the voluntary/no-cue group were about equally likely to choose either task (mean for parity task = .495, 95% confidence interval = $\pm .020$), and more often opted to repeat a task than to switch (mean switch probability = .364, 95% confidence interval = $\pm .061$). There was weak evidence that switching tasks was more likely to occur when the long cue–target interval was in force rather than the short interval (.374 vs. .354, respectively), $p_{\text{BIC}} = .644$.

Task performance

For each of the three groups of subjects, the magnitude judgement task was performed more quickly than the parity judgement task (799 ms vs. 888 ms, overall), $p_{\text{BIC}} > .998$ in all cases. Our primary analyses were directed toward a comparison between subjects in the voluntary-switch and explicitly cued groups, averaging across judgement tasks. We consider the voluntary/no-cue group later. In the analysis of task performance, we assessed response time as a function of whether the current trial was a task repetition or a task switch. The first trial after a break was excluded from analyses because of the relatively long period between it and the preceding trial. To analyse response times, trials on which an error (wrong response hand—a task error; or wrong response finger—a classification error) occurred were excluded, as were trials with response times below 200 ms or above 2800 ms. The upper cut-off was established so that we removed no more than 0.5% of eligible trials (Ulrich & Miller, 1994). The use of two cues per task allowed us to assess task performance in each of three conditions: cue repeat, cue switch, and task switch. Cue–target interval and instructional group were the other factors in our analyses.

Response time means are shown in Figure 1. Cue-switch cost was defined as the difference

between performance in the cue-repeat and cue-switch conditions. This effect was modulated by both instructional group and cue–target interval, as indicated by a Bayesian analysis showing very strong evidence for a three-way interaction between cue condition (cue repeat vs. cue switch), instructional group, and cue–target interval, $p_{\text{BIC}} > .999$. For voluntary-switch subjects, there was weak evidence favouring a null cue-switch cost, $p_{\text{BIC}} = .574$, with no indication of a change across cue–target interval, $p_{\text{BIC}} = .853$, favouring a null effect. This outcome is sensible given that the cues were redundant for these subjects. For explicitly cued subjects, there was a large cue-switch cost that decreased substantially with increased cue–target interval (from 175 ms to 20 ms), $p_{\text{BIC}} > .999$ for the interaction between switch condition and cue–target interval.

In contrast to the dissociation between voluntary-switch and explicitly cued subjects with respect to cue-switch effects, the two groups showed remarkably similar results for task-switch cost, defined as the difference between cue-switch and task-switch conditions. There was a substantial overall task-switch cost of 103 ms, $p_{\text{BIC}} > .999$, which did not vary as a function of instructional group, $p_{\text{BIC}} = .861$ favouring the null interaction model. Response time decreased across cue–target interval for the explicitly cued group, $p_{\text{BIC}} > .999$, but it

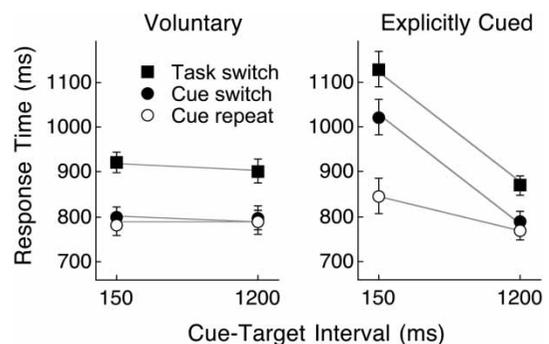


Figure 1. Mean response time for the voluntary-switch and explicitly cued groups as a function of switch condition and cue–target interval. Grey lines indicate the predicted values from Model 3. Error bars are 95% within-subject confidence intervals, appropriate for comparing means across switch/repeat conditions (Loftus & Masson, 1994; Masson & Loftus, 2003).

did so equally for both the task-switch and cue-switch conditions, $p_{BIC} = .779$, favouring a null interaction model, yielding an overall task-switch cost that was comparable to that seen for voluntary-switch subjects. This reduction in response time with an increase in the cue–target interval for the cue-switch and task-switch conditions was apparently related to processing of a cue switch that was required in both of those conditions.

Task errors (executing the wrong task) were quite rare, although they were more frequent for explicitly cued subjects than for voluntary-switch subjects (1.5% vs. 0.3%), $p_{BIC} > .999$. This advantage for voluntary-switch subjects may have been due to those subjects using the same hand to indicate their task selection and to respond to the target on each trial. Making the task-selection response with a particular hand could have prepared that hand for use in response to the upcoming target stimulus. Explicitly cued subjects made no task-selection response and so did not experience this preparation. Task errors showed a cue-switch cost for explicitly cued subjects, whereas no such effect was found for voluntary-switch subjects (mean cost = 1.2% vs. -0.2%), $p_{BIC} = .978$ for the interaction effect. There was no task-switch cost in this error measure, $p_{BIC} = .894$ for the null effect model. Response errors (incorrect classification of a stimulus) were slightly more common for voluntary-switch subjects than for explicitly cued subjects (3.1% vs. 2.0%), but a Bayesian analysis provided only weak evidence for this effect, $p_{BIC} = .633$. There was weak evidence for a cue-switch cost on this measure that was greater for the explicitly cued subjects than for the voluntary-switch subjects (mean cost = 1.2% vs. -0.1%), $p_{BIC} = .729$. No task-switch cost was present for response errors, $p_{BIC} = .888$ for the null model.

Condition means for response time in the voluntary/no-cue group is shown in the top panel of Figure 2. There was a clear switch cost for these subjects, $p_{BIC} > .999$, but the evidence is ambivalent with respect to whether this cost was affected by the response-stimulus interval, $p_{BIC} = .526$. Task errors were very rare for these subjects (mean < 0.1%) so no analyses are reported. For response errors, more occurred with a short

cue–target interval than with a long interval (4.6% vs. 3.3%, respectively), $p_{BIC} = .869$, although there was weak evidence that this effect held only following a task switch, $p_{BIC} = .629$. Data from the cue-repeat and task-switch conditions for the voluntary-switch group are also shown in Figure 2 for comparison. The response time patterns were very similar for the two groups of subjects, although subjects in the voluntary-switch group were generally slower in their responses.

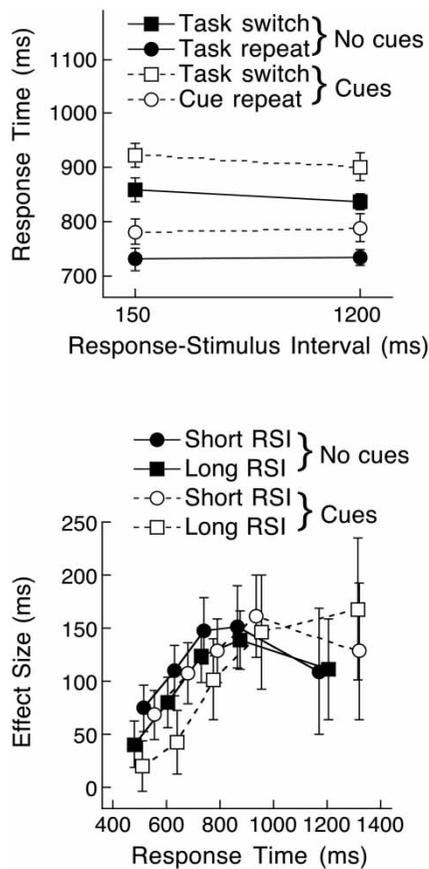


Figure 2. Mean response time as a function of switch condition and response-stimulus interval (upper panel), and mean task-switch effect as a function of response-stimulus interval (RSI) and response-time quantile (lower panel) for the voluntary/no-cue group. Corresponding data for the voluntary-switch group tested with cues are presented for comparison. Error bars are 95% confidence intervals for each effect.

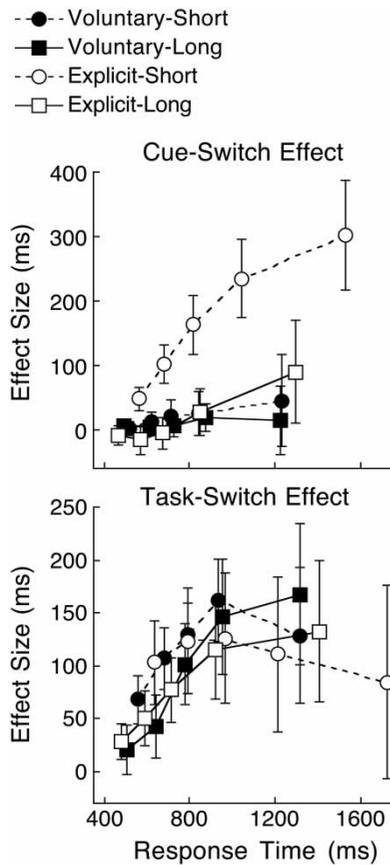


Figure 3. Mean cue-switch effect (upper panel) and task-switch effect (lower panel) as a function of instructional group, cue-target interval, and response-time quantile. Data points are plotted on the horizontal axis according to the mean response time for that quantile, averaged across the two conditions contributing to the effect. Error bars are 95% confidence intervals for each effect.

Modelling of response-time data

Model fitting attempts were restricted to the two groups of subjects who were presented with task cues. We first considered cue-encoding Models 2 and 2 + 1 from Logan and colleagues (Arrington & Logan, 2004a; Logan & Bundesen, 2003, 2004; Schneider & Logan, 2005). These versions of the model assume that the same base response time value holds for all three conditions (task switch, cue switch, and cue repeat). Under these versions of the model, there should be hardly any task-switch effect in the long cue-target interval

condition because with 1200 ms of processing time available, cue encoding and any set-switching process are highly likely to be complete by the time the target is presented. Figure 1 clearly shows, however, that there was a substantial task-switch effect for both cue-target intervals. Consequently, these models did not fit very well the data of either the voluntary-switch or the explicitly cued groups (root mean squared deviation, $RMSD > 30$ ms in all cases when fitting data aggregated over subjects within a group; fits obtained using the Solver function in Microsoft Excel to minimize RMSD between observed and predicted response times).

Model 3 includes the possibility that a process following cue encoding may differ between task-switch and task-repetition conditions. Arrington et al. (2007) considered this process to be a target-encoding operation, and it is incorporated into the cue-encoding model as an additional residual processing-time parameter for the task-switch condition. This four-parameter model fitted the aggregated data very well for both groups of subjects ($RMSD = 5$ ms in both cases), and the resulting parameter values are shown in Table 1. Predicted values generated by the model are indicated by grey lines in Figure 1. This model was also fitted to individual subjects, and the average values for each parameter, computed across the subjects within each condition, are shown in Table 1. Using the individual subject fits, we tested for influences of instructional condition and switch condition on the parameter values. The cue-encoding parameters and the RT-base parameters were examined separately. For the cue-encoding parameters, μ_r and μ_s , their values were substantially larger for the explicitly cued group than for the voluntary-switch group, $p_{BIC} > .999$. The parameter for the cue/task-switch conditions (μ_x) took on a larger value than the parameter for the cue-repeat condition (μ_r), $p_{BIC} > .999$, although this effect was present only for the explicitly cued subjects, $p_{BIC} > .999$ for the group by cue-condition interaction. For the RT-base parameters, the instructional groups did not differ, $p_{BIC} = .879$ for the null difference model, but the parameter for the task-switch

Table 1. Values of best fitting parameters and measure of goodness of fit for Model 3 based on aggregate data and the mean of best fitting parameters for individual subjects

Measure	Voluntary-switch subjects					Explicitly cued subjects				
	μ_r	μ_s	$RT_{BaseRep}$	RT_{BaseSw}	RMSD	μ_r	μ_s	$RT_{BaseRep}$	RT_{BaseSw}	RMSD
Aggregate	9	86	788	905	5	179	384	768	860	5
Mean	88	105	770	888	39	170	379	760	851	38
SEM	17	18	27	37	5	21	30	25	34	4

Note: RT = response time; RT_{Base} is the time required to process and respond to the target stimulus; Rep = repetition; Sw = switching; RMSD = root mean squared deviation; SEM = standard error of the mean; all values in ms.

condition had a larger value than the parameter for the cue-repeat and cue-switch conditions, $p_{BIC} > .999$.

Response-time distribution analysis

We examined the response-time distributions for task performance to evaluate De Jong’s (2000) failure-to-engage hypothesis. The critical question was whether the task-switch effect was robust throughout the response-time distribution for both short and long cue–target intervals, or was weak for the shortest response times under a long cue–target interval. If subjects sometimes fail to engage preparatory processes even when sufficient time is provided (as when a long cue–target interval is presented), then task-switch costs should be particularly evident when response times are longer. With short response times, preparatory processes probably would have been completed given a long cue–target interval, even for a task switch, rendering a reduced or null task-switch cost. This effect should not appear for a short cue–target interval because there would not be adequate time for endogenous preparatory processes to be completed.

We partitioned each subject’s response-time data for each condition into five quantiles. The first quantile for a particular condition contained the shortest 20% of the correct response times in that condition, the second quantile contained the next shortest 20%, and so on. The mean response time within each quantile was computed for each subject, and cue-switch and task-switch effects were computed for each quantile. Each effect was plotted as a function of response time quantile,

producing what is referred to as *delta plots* (Ridderinkhof, 2002).

The delta plot for cue-switching, averaged across subjects, is shown in the top panel of Figure 3 for the voluntary-switch and the explicitly cued groups. Cue-switch effects were minimal for the voluntary-switch subjects across all quantiles. For explicitly cued subjects, the cue-switch effect with the long cue–target interval became apparent only at the longest response-time quantile. This finding suggests that subjects generally were able to adjust to the change in task cue when given adequate processing time, but occasional failures led to long response times. With a short cue–target interval, however, cue-switch costs were clearly present even at the shortest response-time quantile, indicating that explicitly cued subjects consistently required more time for processing a cue change than was available with the short preparation interval.

The pattern of task-switch effects across response-time quantiles was similar for both instructional groups, with little or no task-switch effect in the shortest quantile when the cue–target interval was long (Figure 3, bottom panel). Indeed, considering only the first quantile and averaging over the two instructional groups, the task-switch effect was considerably larger with a short cue–target interval than with a long interval (86 ms vs. 24 ms), $p_{BIC} > .999$. Larger task-switch effects were apparent at later quantiles, even with a long cue–target interval, consistent with De Jong’s (2000) proposal that subjects often fail to engage preparatory operations even when ample

time is provided. Successful engagement presumably leads to short response times and little or no task-switch cost. With a short cue–target interval, robust task-switch costs are present even in the shortest response quantile.

The delta plot for switch cost in the voluntary/no-cue group is shown in the lower panel of Figure 2 and reveals a pattern similar to that seen in the delta plot of task-switch cost for the other two groups (delta plots for task-switch cost in the voluntary-switch group is shown for comparison). Specifically, only a small switch cost (41 ms) is seen in the first response-time quantile when a long cue–target interval was in effect, and this cost was clearly smaller than when a short cue–target interval was in effect (75 ms), $p_{\text{BIC}} = .813$. The switch-cost effect converges for short and long cue–target intervals at later response-time quantiles, as with the other two instructional groups and as predicted by De Jong's (2000) failure to engage hypothesis.

Discussion

We have shown a clear dissociation between explicitly cued and voluntary task-switching operations with respect to cue-switch cost, but strong similarities between them regarding task-switch cost. Cue encoding is crucial when subjects rely on cues to signal the upcoming task, and a version of the cue-encoding model proposed by Logan and colleagues (Arrington & Logan, 2004a; Logan & Bundesen, 2003, 2004) provided a good account of our data: Cue repetition speeded responding but only for explicitly cued subjects, not for voluntary-switch subjects. One might suggest that the lack of a cue-switch cost for the voluntary-switch group was the product of the requirement that voluntary-switch subjects make an overt task-choice response before receiving each cue. This possibility seems unlikely because Arrington et al. (2007) obtained robust cue-switch costs when explicitly cued subjects were required to make an overt response to each task cue before the target stimulus was presented.

The appearance of a task-switch cost, particularly for the explicitly cued group, required the

cue-encoding model to include an effect of task switch on the base response-time parameter. This outcome is consistent with the Arrington et al. (2007) finding of a task-switch cost when explicitly cued subjects overtly responded to each task cue before presentation of the target stimulus. That condition is somewhat similar to our voluntary-switch condition. Importantly, however, even our explicitly cued subjects showed this effect without having made any response to the task cues. Moreover, both groups of subjects showed no overall influence of cue–target interval on task-switch cost, suggesting a similar cause for task-switch costs in the two groups.

This outcome is not entirely consistent with the pattern of task-switch costs reported by Arrington and Logan (2005, Experiment 3). Whereas we obtained no reduction in average task-switch cost with increasing cue–target interval or RSI for any of our three groups, Arrington and Logan found a reliable reduction in switch cost for their voluntary-switch subjects with increasing RSI. They used a 2:1 mapping of cues to tasks for their explicitly cued subjects and obtained uniform task-switch costs across cue–target intervals, as we did. The difference in the results for our respective voluntary groups probably is due to the fact that their subjects did not make a response to indicate their task selection prior to presentation of the target. Rather they made a response only to the target, so switch costs measured in that experiment included possible influences of task-choice processes. Furthermore, when Arrington and Logan (2005, Experiment 6) had voluntary-switch subjects make a response to indicate task selection prior to presentation of the target stimulus, there was only a small reduction in switch cost with increasing response–target interval (the interval between making a task-choice response and appearance of the target). Finally, Arrington and Logan randomly varied this interval across trials, whereas in our experiments that factor was blocked. Mixing long and short intervals may have placed subjects at an exaggerated disadvantage when a short-interval task-switch trial occurred (see Altmann, 2004; Rogers & Monsell, 1995, for effects of blocking vs. mixing interval durations). Nevertheless, it might be informative to replicate

the present experiment using randomized rather than blocked presentation of different cue–target intervals.

To account fully for our results, a version of the cue-encoding model was required that included an additional process to capture task-switch cost. Arrington et al. (2007; see also Arrington & Logan, 2004a) considered a number of candidates for this additional process. One possibility is an endogenous form of task-set reconfiguration, although the fact that task-switch cost was not reduced when a long cue–target interval was used seems to contradict this option. The other two suggestions considered by Arrington et al. were (a) an exogenous form of task-set reconfiguration and (b) negative priming of the cued task set. Similarly, Mayr and Kliegl (2003) proposed that the task-switch cost they observed was due to exogenous response priming when the task was repeated, rather than to an endogenous control operation.

We propose that the additional process (beyond cue encoding) responsible for task-switch cost is a form of endogenous task preparation (e.g., task-set reconfiguration) and, consistent with De Jong (2000), that this process is susceptible to failures to engage. Our examination of response-time distributions for short and long cue–target intervals or RSI indicated that although the overall magnitude of task-switch cost did not substantially vary with interval length, qualitative differences did emerge. With a longer interval, task-switch cost was weak or absent in the first response-time quantile relative to what was found with a short interval. Clear task-switch costs were apparent in later quantiles, and the size of these costs converged for short and long intervals. This pattern of results is consistent with the proposal that a time-consuming reconfiguration process is optionally engaged during the cue–target interval or RSI on task-switch trials. If successfully completed, a rapid response to the target is possible, and little or no task-switch cost is obtained. An exogenous process that is independent of the cue–target interval and is triggered by onset of the target stimulus is not consistent with this pattern of results.

Lien, Ruthruff, Remington, and Johnston (2005) proposed an alternative to De Jong's

(2000) failure to engage hypothesis, in which preparation was claimed to include only specific stimulus–response pairs associated with an upcoming task. Thus, rather than preparing a task set in an all-or-none manner on some trials, subjects instead consistently prepare only a subset of stimulus–response pairs in advance of the actual target stimulus. Lien et al. provided evidence for this idea using tasks in which three stimulus values were mapped to unique fingers for responding. They expected that these mappings would be prepared in sequence from left to right (index, middle, and ring fingers, respectively, on the right hand), so that the first stimulus–finger pair (including the index finger) consistently would be prepared first. Indeed, they found no switch cost, at any point in the response-time distribution, when the target stimulus belonged to that first stimulus–response pair.

There are two reasons to suspect that the Lien et al. (2005) proposal will not provide a full account of task-switch cost. First, it is possible that the stimulus–response encoding operations engaged by their subjects (i.e., the sequential coding of stimulus–response pairs) were peculiar to the details of the tasks they used. In typical task-switching experiments that include two tasks, classes of stimuli rather than individual items are associated with a particular response (e.g., odd digits are assigned to the index finger), and often there are only two classes of stimuli. A more generic or abstract coding of stimulus–response rules may be adopted under these circumstances (e.g., Schneider & Logan, 2007).

Second, the Lien et al. (2005) account assumes that there are inherent limits on the ability of the cognitive system to prepare a task in advance. It is not possible, on this view, fully to prepare for an upcoming task, so that any preparation that is carried out must be partial. Contrary to this assumption, Verbruggen, Liefvooghe, Vandierendonck, and Demanet (2007) showed that by using brief cue durations (<150 ms) coupled with a reasonably long cue–target interval in an explicit task-cueing paradigm, switch costs could be drastically reduced and even eliminated. Thus, when circumstances force subjects to use

task cues efficiently, preparation for a task switch may consistently be completed in full prior to the arrival of the target stimulus. This finding raises questions about whether there is a need to postulate a process exogenously triggered by the arrival of the target stimulus when explaining task-switch cost (cf. Rogers & Monsell, 1995; Rubenstein, Meyer, & Evans, 2001).

We suggest that the component of switch cost that we demonstrated to be due specifically to task switching (rather than to cue switching) reflects inconsistent success at completing an endogenous form of task-set reconfiguration during the cue–target interval. In addition, we have identified important commonalities in the nature of this task-switch cost across explicit task cueing and voluntary task switching, while at the same time demonstrating a dissociation between these two task-switching paradigms with respect to cue-encoding operations. The pattern of task-switch effects, both at the level of condition means and across the response-time distributions, was the same for explicitly cued and voluntary task switching, providing evidence for a common mechanism underlying the preparation for a task switch in these two situations.

Original manuscript received 7 January 2013
Accepted revision received 29 November 2013
First published online 5 March 2014

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