Decomposing Externally Cued Task Switching Costs

Christina Wasylyshyn (christina.wasylyshyn@nrl.navy.mil)
Naval Research Laboratory, 4555 Overlook Ave., S.W.
Washington, DC 20375 USA

Abstract

The double-cued task switching procedure has recently been introduced as a new way to measure externally cued switch costs. In the present individual differences study, two hundred fifty young adults completed measures of task switching, inhibition, and long-term memory. A latent variable approach was taken to examine the relationships among these cognitive measures. Decomposing the externally cued task switching costs into a cue switch component and a task switch component indicated that individual differences in these costs could be explained by benefits of repeated cues rather than by changes in tasks. Individual differences in the cue switch component were predicted by long-term memory scores.

Keywords: cued task switching; switch costs; individual differences; long-term memory; inhibition

Introduction

Recent interest in understanding how people shift their mental sets in response to external cues has led to the development of a new method of measuring task switching costs: the double-cued procedure. In traditional, single-cued procedures, cues and tasks are mapped one-to-one, leading to the possibility that the components of cue switching and task switching are confounded. When there is a change in cue, there must be a change in task; when there is no change in cue, it follows that the task will repeat from the previous trial. Both cue switching and task switching then contribute to the overall switch cost in an undifferentiated manner.

One way to distinguish between cue switching and task switching is to use a double-cued procedure, that is, to have two cues to indicate each task. This leads to three types of trials: cue repeat, cue switch, and task switch. In cue repeat trials, both the cue and the task repeat; this is a traditional nonswitch trial. In cue switch trials, the cue changes but the task remains the same; this is also classified as a nonswitch trial. In task switch trials, both the cue and the task change. The latency differences between cue switch and cue repeat trials are thought to indicate encoding benefits from repeated cues. The latency differences between task switch and cue switch trials are thought to reflect the act of task switching.

In task switching paradigms, responses to the current stimulus trial are slower (and typically less accurate) if the task differs from that completed on the previous trial. Switch costs are thought to indicate how flexibly one can change his/her cognitive configurations, or task sets, to accommodate newly relevant task demands. In order to establish a task set, one must activate relevant task rules (e.g., Mayr & Kliegl, 2003; Rubinstein, Meyer, & Evans, 2001) and minimize interference from competing task sets, possibly through inhibition processes (e.g., see Mayr & Keele, 2000). It has therefore been suggested that one's task set reflects the interaction of task set inertia from previous trial(s), exogenous task set activation from the stimulus itself, and endogenous control input needed to overcome the other two biases and to reconfigure the cognitive system for a change in task (Aron, Monsell, Sahakian, & Robbins, 2004).

Mayr and Kliegl (2003) suggest that performance in double-cued procedures can be explained by two processes: cue-based retrieval of task rules from memory and the application of task rules to the target. The retrieval of task rules produces the cue encoding benefit, while the time involved in applying the mapping rules produces the actual switch cost. Retrieval and loading of task rules from long-term memory is necessary for both nonswitch and switch trials. A repetition of the immediately preceding cue leads to a reactivation of the most recent retrieval process (i.e., positive priming); a cue change, however, requires activation of a new (or less recently activated) retrieval process.

Logan and Bundesen (2003) offer a similar explanation, but one that assumes there is no endogenous control component. The explicit task cue provides enough information to uniquely indicate a response on each trial. There is no task set reconfiguration process between the cue and the target stimulus. Instead, any switch costs that remain beyond the act of cue switching are the result of encoding benefits on nonswitch trials, or priming from related cues, not task switching (see Logan & Bundesen, 2004, for explanation of the process of mediator repetition, and see Schneider & Logan, 2005, for formal model).

Using double-cued procedures, several studies have shown evidence for cue encoding benefits (i.e., responses were faster for cue repeat trials than for cue switch trials). However, after cue encoding effects have been accounted for, the remaining task switch costs have been negligible in some studies (e.g., Logan & Bundesen, 2003, 2004; Monsell & Mizon, 2006 Experiment 1) and substantial in others (e.g., Mayr & Kliegl, 2003; Arrington & Logan, 2005 Experiment 3; Monsell & Mizon, 2006 Experiments 2-6). There are several procedural differences between these studies that may explain some of the differences in results, including type of task cue and frequencies for task switches. Evidence that switch costs result from the processing of the task cue rather than from the switching of tasks has been shown in studies that use salient verbal or pictorial task cues and/or 50% task switch frequency. Evidence for substantial task switch costs over and above any effect of a cue change has been shown in studies that employ arbitrary task cues and/or 33% or less task switch frequency.

Most of the research in task switching and executive control functioning has been experimental in nature. The present study, however, takes a novel individual differences approach to decompose switch costs to determine whether a cue switch component can be differentiated from a task switch component. The overall goal is to establish a representation of the structure of individual differences in the double-cued procedure to determine whether switch costs are more likely to reflect processes involved in the interpretation of instructional cues (i.e., trial to trial change in retrieval path) or the switching of task sets (i.e., trial to trial change in the task itself). In this way, it will be possible to test, at the latent level, if individual differences in the costs incurred reflect a benefit for cue repetition instead of, or as well as, a cost for task switching.

Method

Participants

Two hundred fifty Syracuse University students (169 females, 81 males, mean age = 18.92, SD = 1.21) participated. All students were native English speakers and non-colorblind.

Tasks

There were three task categories: task switching, long-term memory, and inhibition.

Task Switching Switch costs were measured in three task domains: digits, shapes, and verbal. For the digits task, magnitude and parity judgments were made on a series of digits (1, 2, 3, 4, 6, 7, 8, 9). In the shapes task, participants determined either the form or color of an image. Stimuli consisted of combinations of two shapes and two colors. For the verbal task, participants determined if a word shown could be classified as an animal or a non-animal or if it could be classified as something that was smaller or larger than a basketball. Stimuli consisted of 64 high frequency and high imagery nouns obtained from the Toronto Word Pool.

Each task consisted of 120 trials. Cues and targets were combined randomly, with the constraint that cue-repeat, cue-switch, and task-switch trials each occurred on one-third of the trials. In each task, there were four cues. Two meaningful and salient cues distinguished each sub-task, so that one cue noted category membership and the other noted response mapping. In the digits task, participants were presented with one of four cues on each trial: Magnitude, High-Low, Parity, or Odd-Even. Depending on the cue shown, participants pressed the ‘z’ key if the target digit was higher than five or odd, and the ‘/’ key if the target digit was lower than five or even. In the shapes task, one of four cues was presented on each trial: Shape, Triangle-Circle, Color, or Red-Green. Participants pressed the ‘z’ key if the target image was a triangle or was colored red, and the ‘/’ key if the target image was a circle or was colored green. In the verbal task, either the cue Creature, Animal-Nonanimal, Size, or Smaller-Larger was presented on each trial. Participants pressed the ‘z’ key if the target word was an animal or was smaller than a basketball, and the ‘/’ key if the target word was a non-animal or larger than a basketball. On cue repeat trials, both the cue and the task repeated from trial *n* to trial *n+1* (e.g., in the digits task, magnitude followed by magnitude). On cue switch trials, the cue changed but the task repeated (e.g., magnitude followed by high-low). On task switch trials, both the cue and the task changed (e.g., magnitude followed by parity). Participants were given 150 ms for preparation during the cue-stimulus interval and 300 ms for passive dissipation during the response-cue interval.

Inhibition In the flanker task (Eriksen & Eriksen, 1974), participants were asked to identify a centrally presented target letter (either an ‘S,’ ‘C,’ ‘H,’ or ‘K’). This target letter was either presented alone or with three noise letters flanking it on each side. Participants pressed the ‘z’ key when the target letter was S or C, and the ‘/’ key when the target letter was H or K, as quickly as possible. There were four stimulus conditions: (1) no noise (e.g., S), (2) noise same as target (e.g., SSSSSSS), (3) noise response compatible (e.g., SSSCSSS), and (4) noise response incompatible (e.g., SSSKSSS). After completing 32 practice trials, participants completed 160 trials (4 blocks of 40 trials). Trials began with a 500 millisecond fixation cross presented in the center of the screen, followed by a blank screen for 50 milliseconds. The stimulus was then presented until a response was made. The latency difference between the noise response incompatible condition and the no noise condition served as the dependent measure.

Long-Term Memory Participants were asked to learn a list of 30 words. Words were presented one at a time in the center of the screen at a rate of 3 seconds each. After approximately 15 minutes, participants were given a recognition test of all 30 words randomly mixed with 30 foil words. Participants pressed the ‘z’ key if the word was part of the original study list and the ‘/’ key if the word was not presented in the original list. Stimuli were obtained from the Toronto Word Pool and consisted of highly familiar 1 syllable words that were 3, 4, or 5 letters in length. Scores were obtained using a nonparametric form of the discriminability index (i.e., a′), and participants with an a′ score below .5 were excluded from analysis.

Results

Descriptive Statistics

Descriptive statistics are presented in Table 1. All of the measures meet the criteria for univariate normality (Kline, 1998); skews are all less than 3 and kurtosis values are all less than 4. All measures therefore displayed adequate distributional properties for being subjected to latent variable analysis.

Table 1: Descriptive statistics

Variable	M	n	SD	Skew	Kurtosis
Digits					
CR	808.72	241	118.72	0.16	-0.41
CS	1204.26	241	251.99	0.48	0.05
TS	1299.25	241	290.04	0.51	0.49
Shapes					
CR	797.77	245	138.43	0.39	-0.42
CS	1055.56	245	190.16	0.41	0.23
TS	1135.71	245	211.97	0.44	-0.13
Verbal					
CR	934.63	245	149.29	0.44	-0.08
CS	1286.99	245	242.03	0.38	-0.38
TS	1406.19	245	268.89	0.28	-0.16
Delayed Word Recognition	0.83	248	0.08	-0.91	1.29
Flanker	85.39	250	59.92	0.23	0.79

Note: CR = cue repeat trial, CS = cue switch trial, TS = task switch trial

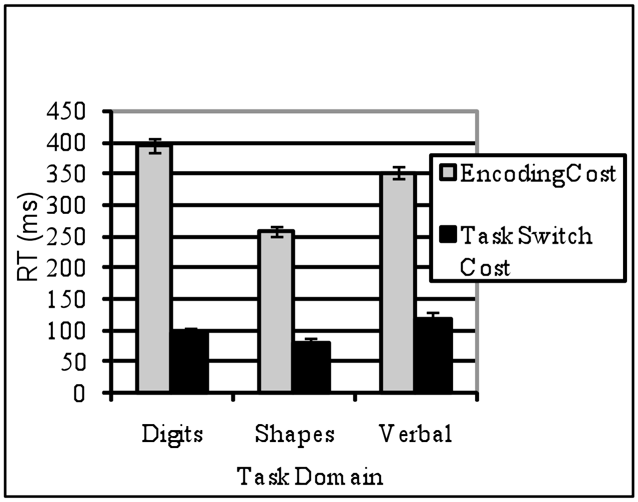

Figure 1: Mean encoding costs and task switching costs (± SE) by task domain (digits, shapes, verbal)

Cue Switching vs. Task Switching: Accounting for Individual Differences in Switch Costs

Figure 1 presents the mean encoding costs and task switching costs (errors bars indicate SE) by task domain (digits, shapes, verbal). Encoding costs were calculated as latency differences between cue switch and cue repeat trials, and task switch costs were computed as latency differences between task switch and cue switch trials. Across task domains (please refer to Table 1), cue repeat trials were the fastest (M = 847 ms), followed by cue switch trials (M = 1182 ms), and finally task switch trials (M = 1280 ms). Cue switch trials were more like task switch trials than cue repeat trials, suggesting that cue repetition effects account for most of the switch cost. Indeed, the majority (77%) of the switch cost is accounted for by the cost of encoding the cue (335 ms cost). However, there is an overall mean residual task switching cost of 98 ms. Planned contrasts indicated that task switch trials were significantly slower than cue switch trials, t = 8.04, p < 0.0001. Therefore, from the means analysis, we can conclude that there is an effect of task switching. Partial correlations between cue repeat and cue switch trials, controlling for performance on task switch trials, were also computed. Controlling for task switch trials did not significantly attenuate the relationship between cue repeat and cue switch trials in any of the task domains; this suggested that the residual effect of task switching might not be useful as an individual differences variable. The next goal was to use Structural Equation Modeling to test for reliable individual differences.

Two models were contrasted to test whether, after accounting for cue encoding, the process of cue-switching is sufficient to explain individual differences in switch costs, or whether an additional task switching process is needed to fully account for these costs. The former will be called the 2-factor model, and the later will be referred to as the 3-factor model. In both models, the manifest measures are the RTs from the cue repeat, cue switch, and task switch trials from the digits, shapes, and verbal tasks. Residual variances of trial type measures employing the same task domain were correlated.

Model fit was assessed using the chi-square test for goodness of fit, the comparative fit index (CFI; Bentler, 1990), the Tucker-Lewis index (TLI; Tucker & Lewis, 1973), and the root mean square error of approximation (RMSEA; Browne & Cudek, 1992). Acceptable model fit is reflected by a nonsignificant chi-square test for goodness of fit, relative fit indices (i.e., CFI and TLI) above .90, and a RMSEA value below .08 (Bentler, 1990; Bentler & Bonnet, 1980). A RMSEA value below .05 indicates excellent fit. Analyses were conducted with AMOS 5 software (Arbuckle, 2003) using maximum likelihood estimation. For all of the SEM models, the factor loadings and interfactor correlations were allowed to vary (Anderson & Gerbing, 1988).

In order to capture what was common, all nine measures were free to load on the first factor, that is, an encoding baseline, because all types of trials (cue repeat, cue switch, and task switch) involve encoding the external cue and thus should have systematic differences in the encoding process. A 1-factor encoding model, however, did not sufficiently account for individual differences in switch costs, $\chi^2(18)$ = 108.88, p = .000, TLI = .881, CFI = .952, RMSEA = .142. The second factor, cue switch, was then introduced to reflect the systematic individual differences associated only with switching a cue; it represents the common variance of the cue switch and task switch trials, once baseline encoding has been accounted for. The 2-factor model (please see

Figure 2; note that the correlations among the residual variances are not included in the figure for ease of interpretation) provided an acceptable fit to the data, $\chi^2(12) = 22.76$, $p = .030$, TLI = .979, CFI = .994, RMSEA = .060. In Figure 2, the larger circles represent the latent variables, and the rectangles represent the scores on the individual indicator tasks that were used to measure each of the latent variables. The smaller circles represent the residual variances of the indictor tasks.

A third factor, task switch, was then introduced to determine if a task switching process, in addition to encoding baseline and cue switching, could better explain individual differences in switch costs. This third factor was equal to the residual common variance of task switch trials only. The 3-factor model (please see Figure 3) provided an excellent fit to the data, $\chi^2(9) = 12.62$, $p = .181$, TLI = .991, CFI = .998, RMSEA = .040. At first glance, this suggests that there is an effect of task switching, over and beyond the processes of cue encoding and cue switching.

Because the 2-factor and 3-factor models are not nested, the Akaike Information Criterion (AIC) and the Expected Cross-Validation Index (ECVI) were used to compare overall model fit (Note: unlike the chi-square difference test, these indices do not provide a statistical comparison of competing models). In general, a model that has the lowest AIC and ECVI values is judged to fit the data better than the alternative model(s) tested (Brown, 2006). For the 2-factor model, AIC = 106.76 and ECVI = 0.429; for the 3-factor model, AIC = 102.62 and ECVI = 0.412. In terms of the AIC and ECVI indices, the 3-factor model best fits the data, however, the 3-factor model was deemed unacceptable due to non-interpretable and ill-fitting parameter estimates of the task switch factor. The variance of the task switch factor was not significant, $p = 0.272$. Only the path loading for the task switch trials in the Shapes task to the task switch factor was significant ($\beta = 0.18$, $p = 0.015$); the path loadings for the Digits and Verbal tasks were not significant.

Therefore, the 2-factor model was accepted. The lower (i.e., better fitting) AIC and ECVI values for the 3-factor model seem to simply reflect adjustment for model complexity compared to the 2-factor model, so that the 3-factor model does not account for any additional systematic differences in switch costs. The act of task switching does not provide reliable information that was not already available from cue switching.

Using Cognitive Measures to Explain Individual Differences in Task Switching

Further support for this claim comes from additional modeling of possible cognitive predictors (inhibition and long-term memory) that might serve to explain some of the systematic differences in each of the latent factors. Please see Table 2 for standardized effects and model fits. It should be noted that because only single indicator predictors are being employed, the following effects are small, but significant. Individual differences in inhibition significantly predict individual differences in encoding, but not cue switching or task switching. That is, individuals who are quicker at encoding the present, most relevant information are also faster at inhibiting previous and/or distracting information. Individual differences in long-term memory

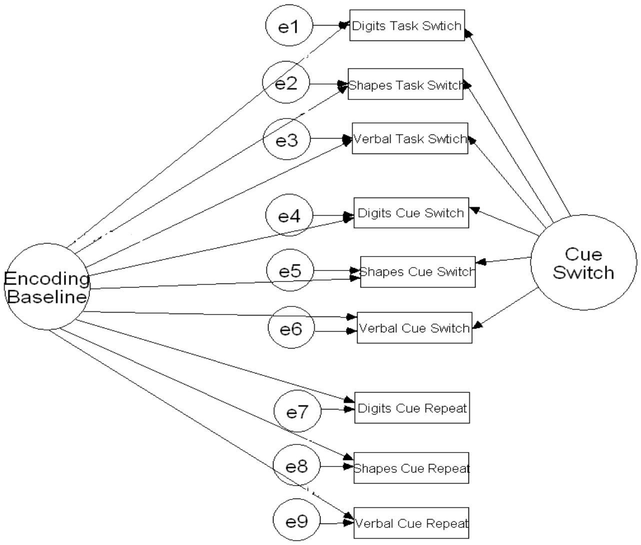

Figure 2: 2-Factor latent variable model to account for switch costs

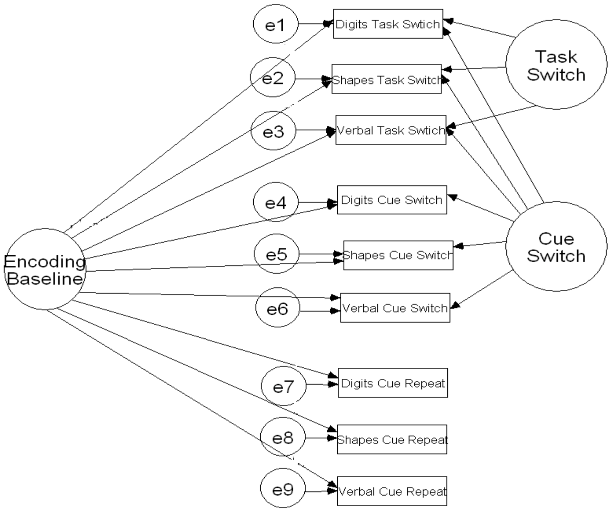

Figure 3: 3-Factor latent variable model to account for switch costs

significantly predict individual differences in cue switching, but not encoding or task switching. That is, individuals who are able to efficiently switch between trials that require a change in information, or change in instructional cue, have higher long-term memory scores.

Table 2: Standardized effects of inhibition and long-term memory on the task switching factor components and model fit statistics

	Inhibition		Long-Term Memory	
	β	p	β	p
Encoding Baseline	**.203***	**.005***	-.039	.586
Cue Switch	-.099	.274	**-.187***	**.036***
Task Switch	.109	.375	-.132	.412
χ^2	22.339		17.37	
df	15		15	
p	.099		.297	
CFI	.996		.999	
TLI	.986		.995	
RMSEA	.044		.025	

Discussion

The overall goal of the present study was to establish a representation of the structure of individual differences in the double-cued procedure. There were two specific aims. The first aim was to examine if individual differences in performance on cue switch and task switch trials could be distinguished at the level of latent variables to determine whether switch costs reflect processes involved in interpreting instructional cues rather than, or in addition to, switching task sets. The second aim was to examine the relationships of the decomposed costs with measures of long-term memory and inhibition to determine the underlying mechanisms or processes that might explain some of the variance in each of the components.

In the aggregate trial-type RT analysis, cue switch trials (M = 1182 ms) appeared more like task switch trials (M = 1280 ms), than like cue repeat trials (M = 847 ms). Although 77% of the switch cost was accounted for by encoding the cue (335 ms), the mean overall residual task switching cost of 98 ms was significant. That is, average performance on task switch trials was significantly slower than average performance on cue switch trials. Using partial correlation, it was found that the first-order correlation between cue repeat and cue switch trials in the digits task was reduced from .76 to .40 after controlling for performance on task switch trials. The correlation between cue repeat and cue switch trials in the shapes task reduced from .80 to .52, and the correlation for the verbal task cue repeat and cue switch trials reduced from .76 to .35. This implied that 63%, 72%, and 59% of the variance shared between the cue repeat and cue switch trials in the digits, shapes, and verbal tasks, respectively, was associated with performance on task switch trials. However, controlling for task switch trials did not significantly attenuate the relationship between cue repeat and cue switch trials in any of the task domains, suggesting that the residual effect of task switching might not be useful as an individual differences variable.

The results of the present study showed that task switching did not serve as a reliable individual differences variable; task switch trials did not provide any additional information that was not already accounted for by the cue switch and cue repeat trials. The residual common variance for the task switch factor was not significant, lending support to the claim made by Logan and Bundesen (2003) that efficient performance does not require an actual act of task switching. It should be noted that this claim can only be made for externally cued paradigms that employ short preparation intervals, as this study only used one interval. Yehene and Meiran (2007) suggest that this may not be the case at longer preparation intervals. However, it should also be noted that in an individual differences study, Friedman and Miyake (2004) could not distinguish switch costs incurred at short preparation intervals from those incurred at longer preparation intervals at the level of latent variables.

That there were no reliable individual differences to account for the act of task switching cannot be attributed to participants' preparatory strategies in response to a high probability of a task switch trial. Recent studies (e.g., Schneider & Logan, 2006; Monsell & Mizon, 2006) have indicated that the frequency of switch trials is related to the magnitude of switch costs, so that the higher the probability of the occurrence of a task switch trial, the smaller the overall switch cost. In the present study, the overall probability of a task switch, *p*(task switch), was 0.33, and the probability of a task switch given a cue switch, *p*(task switch|cue switch) was 0.5. In the Logan and Bundesen (2003) studies, *p*(task switch) = 0.5, and *p*(task switch|cue switch) = 0.67. Unlike the Logan and Bundesen experiments, the present study can rule out the possibility that participants might have strategically controlled their task-set readiness as a function of expectation for a task switch trial, thereby reducing their overall switch costs. Moreover, other procedural precautions were taken in the design of the current study, as suggested by Monsell and Mizon (2006), to capture an endogenous control process, or actual act of task switching, if there was one. For example, the response-stimulus interval was kept constant to avoid confounding active preparation with passive decay, and highly salient cues were used. Finally, the present study can make the claim that increasing the number of target stimuli from 4 in the shapes task, to 8 in the digits task, to 64 in the verbal task did not lead to a task switching effect; participants did not resort to switching task sets in response to the cue as the mapping combinations between cues, targets, and responses got larger.

Conclusions

In summary, individual differences in switch costs were attributed to changes in cue initiated retrieval; switch costs were a consequence of cue priming effects, not a consequence of task changes. The further modeling of cognitive measures to predict individual differences in the component factors indicated that the single inhibition measure was associated with individual differences in the encoding baseline factor, and the single long-term memory measure was related to individual differences in the cue switching factor. It should be noted that although these effects were small, they were theoretically grounded. These results lend support to the claim that the loading of task rules from long-term memory was necessary even on nonswitch trials (Mayr & Kliegl, 2003). Because only one preparation interval was included, the reduction in the switch cost effect across increasing preparation intervals cannot be measured. Therefore, it is not possible to completely rule out an endogenous task set reconfiguration process. The present study can, however, conclude that at short preparation intervals, reliable variance in switch costs could be explained by a cue repetition benefit; an additional task switching process was not needed to fully account for individual differences.

References

Anderson, J. C., & Gerbing, D. W. (1988). Structural equation modeling in practice: A review and recommended two-step approach. *Psychological Bulletin, 103,* 411-423.

Arbuckle, J. L. (2003). Amos (Version 5.0) [Computer software]. Spring House, PA: Amos Development Corporation.

Aron, A. R., Monsell, S., Sahakian, B. J., & Robbins, T. W. (2004). A componential analysis of task-switching deficits associated with lesions of left and right frontal cortex. *Brain: A Journal of Neurology, 127(7),* 1561-1573.

Arrington, C. M., & Logan, G. D. (2005). Voluntary task switching: Chasing the elusive homunculus. *Journal of Experimental Psychology: Learning, Memory, & Cognition, 31(4),* 683-702.

Bentler, P. M. (1990). Comparative fit indices in structural models. *Psychological Bulletin, 23,* 611-626.

Bentler, P. M., & Bonnet, D. G. (1980). Significance tests and goodness of fit in the analysis of covariance structures. *Psychological Bulletin, 88,* 588-606.

Brown, T. A. (2006). *Confirmatory factor analysis for applied research*. New York: Guilford Press.

Browne, M. W., & Cudeck, R. (1992). Alternative ways of assessing model fit. *Sociological Methods and Research, 21,* 230-258.

Eriksen, B. A., & Eriksen, C. W. (1974). Effects of noise letters upon the identification of a target letter in a nonsearch task. *Perception & Psychophysics, 16,* 143-149.

Friedman, N. P., & Miyake, A. (2004). The relations among inhibition and interference control functions: A latent-variable analysis. *Journal of Experimental Psychology: General, 133(1),* 101-135.

Klein, R. B. (1998). *Principles and practice of structural equation modeling*. New York: Guilford Press.

Logan, G. D., & Bundesen, C. (2003). Clever homunculus: Is there an endogenous act of control in the explicit task-cuing procedure? *Journal of Experimental Psychology: Human Perception & Performance, 29(3),* 575-599.

Logan, G. D., & Bundesen, C. (2004). Very clever homunculus: Compound stimulus strategies for the explicit task-cuing procedure. *Psychonomic Bulletin & Review, 11(5),* 832-840.

Mayr, U., & Keele, S. W. (2000). Changing internal constraints on action: The role of backward inhibition. *Journal of Experimental Psychology: General, 129(1),* 4-26.

Mayr, U., & Kliegl, R. (2003). Differential effects of cue changes and task changes on task-set selection costs. *Journal of Experimental Psychology: Learning, Memory, & Cognition, 29(3),* 362-372.

Monsell, S., & Mizon, G. A. (2006). Can the task-cuing paradigm measure an endogenous task-set reconfiguration process? *Journal of Experimental Psychology: Human Perception and Performance, 32(3),* 493-516.

Rubinstein, J. S., Meyer, D. E., & Evans, J. E. (2001). Executive control of cognitive processes in task switching. *Journal of Experimental Psychology: Human Perception and Performance, 27,* 763-797.

Schneider, D. W., & Logan, G. D. (2005). Modeling task switching without switching tasks: A short-term priming account of explicitly cued performance. *Journal of Experimental Psychology: General, 134(3),* 343-367.

Schneider, D. W., & Logan, G. D. (2006). Priming cue encoding by manipulating transition frequency in explicitly cued task switching. *Psychonomic Bulletin & Review, 13(1),* 145-151.

Tucker, L. R., & Lewis, C. (1973). A reliability coefficient for maximum likelihood factor analysis. *Psychometrika, 38(1),* 1-10.

Yehene, E., & Meiran, N. (2007). Is there a general task switching ability? *Acta Psychologica, 126(3),* 169-195.