Does impulsivity Contribute to the Item-Position Effect?

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Abstract:
This paper reports an investigation of whether the personality trait impulsivity contributes to the item-position effect observable in reasoning data. The item-position effect refers to the dependence of item positions and item statistics. Lozano (2015) showed that impulsivity predicts the item-position effect. Nevertheless, an attempt to replicate this finding by Ren, Gong, Chu, and Wang (2017) failed. To further investigate the proposed relationship a sample of 284 undergraduate students completed a set of Advanced Progressive Matrices (APM) as well as the Barratt Impulsiveness Scale (BIS). Confirmatory factor models were used for analyses in our study. Results did not provide evidence for a relationship between impulsivity and the item-position effect. Sample characteristics and cultural differences are discussed as possible reasons for these results.

Keywords:
impulsivity, item-position effect, APM, BIS, impulsivity, confirmatory factor analysis

Introduction

The item-position effect is a method effect that is apparent in the dependency of item statistics on the positions of the items in the sequence of items of a scale. Major hypotheses explaining the item-position effect are, for example, the learning hypothesis (Ren, Wang, Altmeyer, & Schweizer, 2014) and the hypothesis suggesting impulsivity as source of the item-position effect (Lozano, 2015). In this paper an investigation of the hypothesis focusing on impulsivity is reported.

Item-Position Effect

The concept of the item-position effect states that the responses to the items of a scale depend on the positions of the items within the sequence of items. A number of experimental investigations demonstrated such an effect in various personality and ability scales (e.g., Campbell & Mohr, 1950; Hamilton & Schuminsky, 1990; Knowles & Byers, 1996). Learning and memory seem to be sources of this effect since a position effect was frequently observed when investigating memory processes or is ascribed to learning processes (see Sederberg, Howard, & Kahana, 2008). However, there may also be other sources of this effect as, for example, impulsivity (Lozano, 2015).

The item-position effect was also found to increase the amount of systematic variance along with the sequence of items of a scale (Knowles, 1988). Such systematic variation can be expected to find its expression in the covariance pattern of the items. Detection and representation of systematic variation in data is the purpose of factor analysis that operates with matrices of covariances or correlations as input. Investigations show that the item-position effect can be represented as a factor of a confirmatory factor model (Schweizer, 2012; Schweizer, Schreiner, & Gold, 2009). Furthermore, there are demonstrations that the factor assumed to represent the item-position effect does not tap the sources of other method effects as, for example, the difficulty factor (Schweizer & Troche, 2018; Zeller, Reiß, & Schweizer, 2017) or the speed factor (Zeller, Reiß, & Schweizer, 2020).

The separation of the item-position effect from the effect due to the source focused by the scale can be achieved by decomposing the observed variances and covariances. The separated parts are due to two sources: the source that is focused by the scale and the source of the item-position effect. In investigations on the structural validity of a scale only the source that is focused by the scale is taken into consideration. This is achieved by the congeneric model (Jöreskog, 1971). The extended version of this model regarding the jth manifest item random variable Y_j ($j = 1, \ldots, p$) describes this variable as composed of a true part and an error part. The true part is considered as the sum of an item-specific true score random variable ν_j ($j = 1, \ldots, p$) and the person-specific product of factor loading λ_j ($j = 1, \ldots, p$) and latent true score random variable η that reflects the source focused by the scale and is considered as the person attribute of interest. Furthermore, there is the error part that is represented by the error random variable ε_j ($j = 1, \ldots, p$). In the following Equation 1 taken from the extended congeneric measurement model a linear relationship among these parts is assumed such that

$$Y_j = \nu_j + \lambda_j \eta + \varepsilon_j \qquad (1)$$

The congeneric test theory considers the item-specific latent true score random variable as item easiness and the factor loading on the latent true score random variable as item discriminability (Carmines & McIver, 1981; Lucke, 2005; McDonald, 1999).

If variances and covariances provide the input to confirmatory factor analysis, the item-specific true score random variable needs to be eliminated from the model. Variances and covariances are due to individual differences, i.e. the person-specific and error components of the model. The elimination of the item-specific part of the manifest item random variable is made obvious by replacing the upper case letter Y by the corresponding lower case letter y so that

$$y_j = \lambda_j \eta + \varepsilon_j \qquad (2)$$

The confirmatory factor model according to Equation 2 provides the outset for the design of the confirmatory factor model that captures the item-position effect since the item-position effect is usually observed in the items of scales that are thought to represent a construct in the first place. The extension of Equation 2 by an additional factor is necessary for capturing the item-position effect. In order to achieve the separation of the effects and to assure that the position-effect factor covers solely this effect, it is useful to fix the factor loadings according to the expectations regarding the unfolding of the item-position effect along with the sequence of items. Based on Knowles' (1988) results a monotonically increasing size of the effect is expected. A simulation study designed according to the relational pattern of data of the well-known ability measure Advanced Progressive Matrices (APM; Raven, Raven, & Court, 1997) compared different ways of realizing the increase (Zeller, Krampen, Reiß, & Schweizer, 2017). In this study piecewise adaptation did best but did not differ substantially from the quadratic increase. Assuming such an increase, the following quadratically increasing function f_q is to be selected for representing the item-position effect:

$$f_q(j) = (j-1)^2 / (p-1)^2 \quad (3)$$

The extension of the confirmatory factor model for the representation of the item-position effect can be achieved by integrating the product of constraint $\lambda_j^{f_q}$ (j = 1, ..., p) and the factor η_p, so that the following Equation 4 for the jth item is obtained:

$$y_j = \lambda_j \eta_g + \lambda_j^{f_q} \eta_p + \varepsilon_j \quad (4)$$

The subscript g (for *genuine*) is added to the latent true score random variable η in order to distinguish it from η_p. There is also the possibility to constrain the factor loadings on the latent true score random variable η_g in order to assure that this factor does not accommodate other effects:

$$y_j = \lambda_j^{f_g} \eta_g + \lambda_j^{f_q} \eta_p + \varepsilon_j \quad (5)$$

The superscript signifies that the loading is a constraint, which is achieved by means of function *f*. The superscript f_g indicates that this is a constraint for the factor loading of the genuine factor. Furthermore, the fixation of factor loadings must be accompanied by the estimation of the variance of the corresponding latent random variable.

Confirmatory factor analysis of data obtained by binary items of a reasoning scale on the basis of the described confirmatory factor model means analyzing binary data by means of a statistical procedure that expects continuous and normally distributed data. To overcome this discrepancy, we use the probability-based covariance in combination with a variance-stabilizing link transformation (Schweizer, 2013; Schweizer, Ren, & Wang, 2015). A link function for bridging the gap between model and data is suggested by the generalized linear model approach (McCullagh & Nelder, 1985). In the present case the link function relates the variances predicted by the model to the actual variances characterizing the data. The link function is realized as a weight function *w* with the probability to respond correctly π_j (j = 1, ..., p) as argument. It is defined such that

$$w(j) = \sqrt{\pi_j(1-\pi_j)} \quad (6)$$

for j = 1, ..., p. This function provides weights that serve as multipliers to each true component of the measurement models. This weight function serves well in combination with fixed factor loadings.

The weight function in Equation 6 needs to be integrated into Equations 4 and 5 for achieving an equation that represents the source captured by the scale and the source of the item-position effect in addition to considering the distributional differences of data and latent variables:

$$y_j = \lambda_j \eta_g + w(j) \lambda_j^{f_q} \eta_p + \varepsilon_j \quad (7)$$

and

$$y_j = w(j)\lambda_j^{f_g}\eta_g + w(j)\lambda_j^{f_q}\eta_p \quad (8)$$

where each true part receives a weight.

Impulsivity

Impulsivity is a relatively stable personality trait that has been found to play a role in a variety of areas of life such as education, work, social adaptation, and health (Olmstead, 2014). It is considered as a "predisposition toward rapid, unplanned reactions to internal or external stimuli without regard to the negative consequences of these reactions to the impulsive individual or to others" (Moeller, Barratt, Dougherty, Schmitz, & Swann, 2001, p. 1784). Impulsivity is a multidimensional construct that comprises various aspects of thinking, feeling, and action (Cyders, 2012). As a consequence, there are numerous ways of assessing impulsivity. There are tasks focusing on response inhibition and on decision processes (e.g. Logan & Cowan, 1984; Reynolds, 2006), and there are many scales requiring self-descriptions. One of the most frequently used measures is the *Barratt Impulsiveness Scale* (BIS) in its current Version 11 (BIS-11; Patton, Stanford, & Barratt, 1995). The structural model of this scale includes six first-order factors and three second-order factors. The second-order factors (attentional, motor, and non-planning impulsiveness) are achieved by merging two first-order factors each. Attempts to replicate the first-order structure did not yield satisfactory results, whereas the second-order structure appears to show sufficient replicability (Stanford et al., 2009; Vasconcelos, Malloy-Diniz, & Correa, 2012).

A relationship between impulsivity and the item-position effect was proposed and demonstrated by Lozano (2015). Lozano's approach has grown out of the research on the relationship between impulsivity and intelligence. Impulsivity usually negatively affects performance on complex cognitive tasks such as intelligence tests (e.g., Lozano, Gordillo, & Pérez, 2014; Russo, De Pascalis, Varriale, & Barratt, 2008). Furthermore, impulsivity appears to be especially disadvantageous for the executive functions of working memory (Whitney, Jameson, & Hinson, 2004). Moreover, there are demonstrations that the item-position effect is related to attention and learning (Ren, Goldhammer, Moosbrugger, & Schweizer, 2012; Ren et al., 2014). Given these previous results, it is argued that impulsivity can be expected to be more disadvantageous for the item-position effect than for the core of intelligence. Lozano's study included two impulsivity scales, the UPPS-P Impulsive Behavior Scale (Lynam, Smith, Whiteside, & Cyders, 2006) and BIS-11, in their Spanish adaptations. For each one of them a negative relationship with the position-effect factor (UPPS-P: −.53, BIS-11: −.48) is reported in a sample of 232 undergraduate students. An attempt to replicate the results by Ren, Gong, Chu, and Wang (2017) failed.

Objective of Study

The aim of the present study was to provide further evidence for evaluating the proposed relationship of impulsivity and the item-position effect.

Method

Participants

The sample included 284 participants, who were undergraduate students of Goethe University Frankfurt, Germany. The mean age of the sample was 22.8 years (*SD* = 4.2). Two times as many females as males could be recruited. Data were collected using a secure online questionnaire system. Participation was voluntary; as incentive students were rewarded with course credit or a financial reward.

Instruments

Advanced Progressive Matrices (APM)
As matrices problems were repeatedly shown to contain the item-position effect (e.g., Kubinger, Formann, & Farkas, 1991), a shortened version of APM Set II (Raven et al., 1997) that included 18 items was selected for the assessment of intelligence. The APM items required participants to select one out of eight options to complete an incomplete 3 x 3 matrix of geometric forms. In this point, the present study differs from the Lozano (2015) study that used all 36 APM items. However, Mackintosh and Bennett (2005) found that the reduced set of items covered the same range of item difficulties as the complete set of items. Participants were allowed 20 minutes for completing the items. This was half of the time recommended for the full version.

Barratt Impulsiveness Scale (BIS)
The latest revision BIS-11 (Patton et al., 1995) included 30 items, which were answered on a four-point scale (rarely/never, occasionally, often, almost always/always). Three broad dimensions of impulsivity were measured: attentional impulsiveness (e.g., I don't "pay attention"), motor impulsiveness (e.g., I act on the spur of the moment), and non-planning impulsiveness (e.g., I plan tasks carefully[1]). Spinella (2007) presented a shortened English language version of BIS-11 with 15 items (BIS-15), which identified the second-order factors of attentional, motor, and non-planning impulsiveness equally well. A German translation of this short version was developed by the authors (see Krampen, Schweizer, Reiß, & Gold, 2016) and used to assess impulsivity in the present study. A pilot study based on a student sample showed the appropriateness of the translated German version. The three-dimensional structure was supported (Krampen et al. 2016).

Statistical Investigation

In the first step the quality of measurement models was investigated. Two types of measurement models were designed for the APM items. One type was a hybrid model (see Equation 7). This model included one factor with free factor loadings (the genuine factor) and another one with factor loadings fixed according to Equation 3 (the position-effect factor). In the other type all factor loadings were fixed; the factor loadings on the factors representing the construct were of equal size. In both types there were 18 manifest variables according to the number of items. Since the CFI results of one model missed the lower limit for acceptable results, the two easiest items that showed almost no variance were excluded. This reduced the number of manifest variables from 18 to 16.

1 Reverse scored item

The measurement model regarding impulsivity was a one-factor model. It included three manifest variables corresponding to the three subscale scores. Since setting the variance parameter to one in combination with free factor loadings and free error estimates implied zero degrees of freedom, several alternative models were specified by setting factor loadings equal to one and by setting error variances equal to each other. These models had to be compared for identifying the best fitting model with one degree of freedom.

In the second step the measurement models for intelligence and the measurement model for impulsivity were combined to obtain comprehensive models that enabled the estimation of the correlation between the position-effect factor and the impulsivity factor.

The statistical investigation was conducted using the maximum likelihood estimation method. Model fit was evaluated by root mean square error of approximation (RMSEA), standardized root mean square residual (SRMR) and comparative fit index (CFI). Furthermore, information on chi-square (χ2) and Akaike information criterion (AIC) were recorded. Analyses were carried out using LISREL (Jöreskog & Sörbom, 2006).

Results

The results regarding model fit of the APM models are provided in Table 1.

All RMSEA and SRMR results indicated good model fit. In contrast, all CFIs were close to the lower limit of acceptable results, that was .90. Since the CFI for the two-factor model with fixed factor loadings and 18 manifest variables was below .90, the two items showing the smallest variance (the first item and the second item) were eliminated. The elimination of these items led to an acceptable CFI for the two-factor model with fixed factor loadings. The hybrid models led to slightly better fit results than the two-factor models with fixed factor loadings with the exception of the AICs.

The fit results for the BIS models are reported in Table 2.

Table 1 Fit Results Obtained when Investigating APM Items by Two-factor Confirmatory Factor Model (*N* = 284)

Manifest variables	χ^2 (df)	RMSEA	SRMR	CFI	AIC
Two-factor model *with* fixed factor loadings					
1 - 18	239.8 (151)	.05	.07	.89	279.8
3 - 18	193.7 (118)	.05	.07	.90	229.7
Hybrid two-factor model					
1 - 18	210.6 (134)	.05	.06	.91	284.6
3 - 18	170.5 (103)	.05	.06	.91	236.5

The results reported in the first to third rows were obtained by investigating one-factor confirmatory factor models with different factor loadings fixed to one. The last row includes the result for the one-factor confirmatory factor model with the second and third error components fixed to equal sizes. The fit statistics indicated good model fit for the one-factor confirmatory factor model with the first factor loadings fixed to one (first row) and the one-factor confirmatory factor model with the second and third error components fixed to equal sizes (last row). There was no substantial difference between these two good models (CFI difference = 0.0). However, since the one-factor confirmatory factor model with the second and third error components fixed to equal sizes showed the slightly better AIC, this model was selected for the investigation of the relationship between impulsivity and intelligence.

Table 2 Fit Results Obtained when Investigating BIS Scores by One-factor Confirmatory Factor Model (*N* = 284)

Feature	χ^2 (df)	RMSEA	SRMR	CFI	AIC
$\lambda_1 = 1$	0.35 (1)	.00	.02	1.00	10.35
$\lambda_2 = 1$	15.53 (1)	.22	.12	.91	25.53
$\lambda_3 = 1$	26.58 (1)	.30	.14	.84	36.58
$\varepsilon_2 = \varepsilon_3$	0.03 (1)	.00	.00	1.00	10.03

Table 3 Fit Results Obtained when Investigating APM Items by Two-factor Confirmatory Factor Model (*N* = 284)

Manifest variables	χ^2 (df)	RMSEA	SRMR	CFI	AIC	$r_{\text{p-BIS}}$	*t*	*p*
				Two-factor model *with* fixed factor loadings				
1 - 18	290.1 (204)	.04	.07	.91	344.1	–.11	–0.77	ns
3 - 18	193.7 (165)	.04	.06	.91	291.0	–.13	–0.82	ns
				Hybrid two-factor model				
1 - 18	261.3 (187)	.04	.06	.93	349.3	–.12	–1.02	ns
3 - 18	218.2 (150)	.04	.06	.92	298.2	–.12	–1.09	ns

Note: BIS was represented according to the one-factor model with the errors components of the second and third scales set equal.

The results achieved in investigating the comprehensive models and the relationship between impulsivity and intelligence are included in Table 3.

Statistics signify good model fit according to RMSEA and SRMR and acceptable model fit according to CFI. The standardized correlations between impulsivity and intelligence vary between .11 and .13. None of them reached level of significance.

Discussion

Impulsivity is a personality trait with possible consequences for performance. Persons showing a high level of impulsivity are unlikely to concern themselves with the same task for a longer time span and to ignore what can potentially draw their attention away. A high level of impulsivity is not in line with a high level of attention in the sense of concentration and it is not a favorable precondition of learning (as a source of the item-position effect). Rather, it appears to be a high level of attention that was found to relate to the item-position effect (e.g. Ren et al., 2012). Therefore, it is reasonable to expect a negative correlation of impulsivity and the item-position effect.

The results of the present study do not provide evidence in favor of this expectation. They do not replicate the results achieved by Lozano (2015). Instead, they are in line with the results achieved by Ren et al. (2017). On the one hand, this is surprising, because the present study seems to show considerable similarity to the Lozano (2015) study: samples are drawn from the undergraduate students' population and sample statistics are comparable. Furthermore, since APM is a non-verbal test, the difference between the Spanish and German versions should be minor. On the other hand, there are some crucial variations: Lozano (2015) used a paper-pencil version of APM without time limit, while the current study used a computer-based version with a certain time limit. This may affect test performance and its relationship with impulsivity. Another possible reason for the failure to replicate Lozano's results are differences between the BIS versions since the translation of a questionnaire from one language into another language can cause dissimilarity. Furthermore, BIS scores in our sample were quite similar to those observed in other student samples (e.g., Russo et al., 2008, Whitney et al., 2004). BIS scores in Lozano's sample showed a considerably higher range. Cultural differences in impulsive behavior may also play a role (Lee & Kacen, 2008). Further research is necessary to clarify this point.

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