Classical and modern prejudice: Attitudes toward people with intellectual disabilities

Nazar Akrami a,*, Bo Ekehammar a, Malin Claesson b, Karin Sonnander b

a Department of Psychology, Uppsala University, Box 1225, SE-751 42 Uppsala, Sweden
b Department of Neuroscience, Psychiatry Ulleråker, Uppsala University, Sweden

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Abstract

In two studies, Study 1 and Study 2, we examine whether attitudes toward people with intellectual disabilities, like sexism and racism, consist of two forms—a classical and a modern, where the classical is overt and blatant and the modern is more subtle and covert. Self-report scales tapping these two forms were developed in Study 1. Based on confirmatory factor analyses, the results in Study 1 supported our hypothesis and revealed that the modern and classical forms are correlated but distinguishable. This outcome was replicated in Study 2. Construct and discriminatory validations of the scales provided further support for the distinction. The theoretical and practical importance of the results is discussed in relation to previous research on attitudes toward people with intellectual disabilities and other social outgroups.

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1. Introduction

Employing recent methods and theorizing in the domain of social attitudes, the present paper is an attempt to shed light on attitudes toward people with intellectual disabilities and their relation to other social attitudes (e.g., toward ethnic groups, women, lesbians and gay men). Historically, people with intellectual disabilities were considered as a low-power group in most, if not all, societies and they were repeatedly victims of discrimination. The common finding in
research on attitudes toward people with intellectual disabilities (e.g., Handler, Bhardwaj, & Jackson, 1994; Leyser, Kapperman, & Keller, 1994; Pittock & Potts, 1988; Rimmerman, 1998) is that people, to various degrees, harbor negative attitudes toward persons with disabilities. Negative attitudes toward people with disabilities hinder them from bringing about their life goals (Antonak & Livneh, 2000) and narrow their possibilities to fully exercise the rights conferred on them by community law, for example, the right to free movement and the right to vote.

In search for an explanation of these negative attitudes, and for moderating factors, researchers have examined variables, like gender, age, education, socioeconomic status, and belief in a just world, with a mixed pattern of results. Whereas some researchers have found these variables to be related to attitudes toward people with disabilities (e.g., Furnham, 1995; Leyser et al., 1994), others have not (e.g., Hudson-Allez & Barrett, 1996; Krajewski & Flaherty, 2000).

The research referred to above has, almost entirely, been based on self-report measures and the reported changes across time have been found to be in the direction of more tolerant attitudes toward people with intellectual disabilities (e.g., Antonak & Livneh, 2000). One can question whether this reflects a real attitude change or whether the change is due to various response biases in the instruments used. This is an important issue, because social psychologists have questioned the ability of traditional self-reports to accurately reflect prejudicial or negative attitudes that individuals may harbor against social groups (Akrami, Ekehammar, & Araya, 2000; Dovidio & Gaertner, 1986; McConahay, Hardee, & Batts, 1981). Examining whether there is any change in attitudes toward people with intellectual disabilities, Haar, de Vries, and van Maren (2000) found, that after years of campaigning, attitudes toward these people were more or less the same. However, and more interesting for the present study, Haar et al. (2000) found that there is a change in terminology used to describe people with intellectual disabilities. We argue that this change reflects new ways of expressing prejudice beliefs.

The susceptibility of traditional self-report questionnaires to various biases is further supported by people’s tendency to present themselves as socially or politically “correct” which may prevent an open expression of negative attitudes and prejudice (e.g., Crosby, Bromley, & Saxe, 1980). Consequently, researchers have recently begun to make a distinction between classical (old-fashioned, blatant, overt) and modern (subtle, covert) forms of prejudice (e.g., Akrami et al., 2000; Glick & Fiske, 1996; McConahay et al., 1981; Swim, Aikin, Hall, & Hunter, 1995), where the classical form taps direct or open prejudice and the modern covert or subtle. Sears (1988) characterized modern prejudice by three components: denial of continued discrimination, antagonism toward minority group demands, and resentment about special favors for minority groups. These classifications have been found to be identifiable in various cultural contexts and for different types of discrimination (e.g., Akrami et al., 2000; Swim et al., 1995). Thus, we suggest that the distinction between modern and classical types of prejudiced attitudes is identifiable also for attitudes toward people with mental disabilities.

In Study 1, we develop and validate a modern and a classical scale measuring attitudes toward people with intellectual disabilities, and we examine whether these two forms can be distinguished. Following previous research on modern and classical attitudes in other domains, we anticipate that the responses will be characterized by a correlated two-factor structure, representing a modern and a classical factor, respectively. In Study 2, using another sample, we examine the replicability of the results obtained in Study 1. In addition, we examine the relation of prejudice toward people with intellectual disabilities with sexism, racism, and attitudes toward lesbians and gay men.
2. Study 1

A pool of items was generated reflecting the ideas underlying classical and modern attitudes toward people with intellectual disabilities. The modern items were generated to reflect three components: denial of continued discrimination, antagonism toward minority group demands, and resentment about special favors for minority groups (Sears, 1988). Further, the classical items were generated to reflect classical (old-fashioned) attitudes toward people with intellectual disabilities in accord with previous research in other domains (e.g., McConahay et al., 1981). In line with previous research (e.g., Akrami et al., 2000; Pettigrew & Meertens, 1995; Swim et al., 1995; but see Coenders, Scheepers, Sniderman, & Verberk, 2001), we expect the modern and the classical attitudes to be correlated but characterized by a two-factor structure, representing a modern and a classical factor, respectively.

To examine the discriminant validity of the modern and the classical scales, we first expected participants to score higher on the modern than the classical scale. This expectation was based on the contention that the overt nature of the classical scale, as compared with the modern, restrains participants’ responses. Second, we included a set of variables assessing participants’ education concerning, experience of, and various types of contact with, people with intellectual disabilities. We expected these variables to show a differential pattern of relations to the modern and classical scales. Specifically, we expected participants with education about, and experience and contact with, people with intellectual disabilities to score lower on the modern scale. On this point, we base our argument on previous research where contact and education have been shown to be a key factor in the development of more positive or tolerant attitudes toward people with disabilities (e.g., Hudson-Allez & Barrett, 1996; Pittock & Potts, 1988; Rimmerman, 1998). Because of its overt nature, the classical scale is not expected to differentiate between no- and yes-responders on the mentioned variables.

A further examination of the discriminant validity of the scales is made on the basis of the participants’ gender, where men are expected to express more prejudicial beliefs than women. This expectation was based on previous research revealing that men score higher on, for example, ethnic prejudice (Akrami et al., 2000), sexism (Ekehammar, Akrami, & Araya, 2000; Swim et al., 1995), and constructs related to prejudicial beliefs, such as social dominance orientation (Sidanius & Pratto, 1999). However, because of the overt nature of classical prejudicial attitudes, we expect that the magnitude of the differences between men and women is higher for the modern as compared with the classical scale.

As to construct validation, we expect the two scales to be positively correlated with social dominance orientation (SDO), an individual differences variable measuring “one’s degree of preference for inequality among social groups” (Pratto, Sidanius, Stallworth, & Malle, 1994, p. 741). Measures of social dominance orientation have been found to be highly correlated with negative attitudes toward mentally retarded people (Claesson, Sonnander, & Ekehammar, 2000; Ekehammar, Akrami, Gylje, & Zakrisson, 2004), racial prejudice (e.g., Ekehammar et al., 2004; Sidanius & Pratto, 1999), sexism (e.g., Ekehammar et al., 2000; Sidanius, 1993), and negative attitudes toward homosexuals (e.g., Ekehammar et al., 2004; Whitley, 1999).

2.1. Method

2.1.1. Participants

There were 235 participants, 146 women and 89 men in the age range of 18–43 years ($M = 23.9$ years). The participants, students at Uppsala University, represented various academic
disciplines: natural sciences and technology \((n = 51)\), philosophy \((n = 53)\), economics \((n = 12)\), physiotherapy \((n = 14)\), behavioral science \((n = 38)\), medicine \((n = 41)\), and education \((n = 46)\). Participation in the study was voluntary and participants did not receive any payment for their participation.

2.1.2. Procedure and questionnaire

Responses to all items were made on Likert-type 4-point scales ranging from strongly disagree (1) to strongly agree (4). In connection with their lectures at the different departments, participants were given envelopes containing a questionnaire of three parts: (1) a preliminary set of items for the Modern and Classical attitudes scales toward people with intellectual disabilities, (2) items for assessing participants’ education about, experience of, and various types of contact with people with intellectual disabilities, (3) a Swedish modification of the Social Dominance Orientation Scale (for reliability and validity information, see Pratto et al., 1994), which included items such as Increased social equality and Inferior groups should stay in their place. Items were randomly mixed and appropriate items were reversed. The internal consistency of the Social Dominance Orientation Scale was found to be quite satisfactory in the present sample \((\alpha = 0.87)\).

2.1.3. Scale construction

Based on Sears’ (1988) classification system (denial of continued discrimination; antagonism toward out-groups’ demands; lack of support for policies designed to help out-groups), 25 potential items were constructed to reflect beliefs and ideas underlying classical and modern prejudiced attitudes toward people with intellectual disabilities. An exploratory principal components factor analysis of these items yielded a two-factor solution, using a scree test for determining the number of factors. Most modern items loaded on the first factor and the classical items on the second. Items that loaded equally on both factors or on the “wrong” factor were eliminated from subsequent analyses. The 19 final items, 8 classical and 11 modern, are presented in Appendix A.

2.1.4. LISREL analyses

Confirmatory factor analyses were carried out on the item covariance matrix for the combined sample using LISREL 8.14 (Jöreskog & Sörbom, 1993). As in previous research (e.g., Akrami et al., 2000), three separate models were tested: (1) The one-factor model examining if the correlation matrix of the modern and classical items is best represented by one latent factor. (2) The uncorrelated two-factor model examining if the correlation matrix can be reduced to two separate and uncorrelated latent factors (a modern and a classical). (3) The correlated two-factor model examining if the correlation matrix can be reduced to two separate and correlated factors.

The best factor solution was determined by using \(\chi^2\) tests for the fit between model and data. However, because \(\chi^2\) values are influenced by sample size, five additional indices were used to compare the goodness-of-fit of the models (Bollen, 1989). The root mean square error of approximation (RMSEA) measures the discrepancy per degree of freedom. Values close to or lower than 0.05 indicate a satisfactory fit (Browne & Cudeck, 1993). The goodness-of-fit Index (GFI) measures the relative amount of variance and covariance jointly accounted for by the model. The more variance accounted for by the model, the better the fit. The GFI can range from 0 to 1, with higher values indicating better fit (Bollen, 1989; Jöreskog & Sörbom, 1993). The additional three fit indices (NFI, NNFI, CFI) measure how much better a model fits as
compared to other models (Bollen, 1989). These indices too, can range from 0 to 1, with higher values indicating better fit.

### 2.2. Results and comments

#### 2.2.1. Exploratory factor analysis

Based on a scree test, an exploratory principal components analysis of the final set of 19 modern and classical items revealed a two-factor structure, where the classical items loaded on one factor and the modern items on the other. Exploratory factor loadings, item means, standard deviations, and corrected item-total correlations are presented in Table 1. Forming a classical and a modern scale by adding item scores in accord with the factor analytic outcome showed that the internal consistency reliability was found to be satisfactory for the modern ($\alpha = 0.71$) but low ($\alpha = 0.63$) for the classical scale.

#### 2.2.2. Confirmatory factor analyses

The fit indices showed that the correlated two-factor model fits the data relatively well (see Table 2). Furthermore, the $\chi^2$ tests indicated that the correlated two-factor model gave a significantly better fit than the uncorrelated two-factor model, $\Delta \chi^2 (1, N = 235) = 37, p < 0.000$. The correlated two-factor model also yielded a better fit than the one-factor model, $\Delta \chi^2 (1,$

<table>
<thead>
<tr>
<th>Item</th>
<th>$M$</th>
<th>S.D.</th>
<th>$r$</th>
<th>Exploratory</th>
<th>Confirmatory</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td>Modern</td>
<td>Classical</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td>Modern</td>
<td>Classical</td>
</tr>
<tr>
<td>Classical</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1</td>
<td>1.23</td>
<td>0.55</td>
<td>0.41</td>
<td>0.13</td>
<td>0.32</td>
</tr>
<tr>
<td>2</td>
<td>1.22</td>
<td>0.48</td>
<td>0.45</td>
<td>0.27</td>
<td>0.35</td>
</tr>
<tr>
<td>3</td>
<td>1.70</td>
<td>0.67</td>
<td>0.61</td>
<td>0.01</td>
<td>0.65</td>
</tr>
<tr>
<td>4</td>
<td>1.46</td>
<td>0.76</td>
<td>0.60</td>
<td>0.00</td>
<td>0.55</td>
</tr>
<tr>
<td>5</td>
<td>1.37</td>
<td>0.67</td>
<td>0.57</td>
<td>0.06</td>
<td>0.57</td>
</tr>
<tr>
<td>6</td>
<td>1.23</td>
<td>0.54</td>
<td>0.48</td>
<td>0.08</td>
<td>0.45</td>
</tr>
<tr>
<td>7</td>
<td>1.31</td>
<td>0.58</td>
<td>0.60</td>
<td>0.26</td>
<td>0.56</td>
</tr>
<tr>
<td>8</td>
<td>1.08</td>
<td>0.32</td>
<td>0.51</td>
<td>0.05</td>
<td>0.63</td>
</tr>
<tr>
<td>Modern</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1</td>
<td>1.11</td>
<td>0.34</td>
<td>0.35</td>
<td>0.34</td>
<td>0.18</td>
</tr>
<tr>
<td>2</td>
<td>1.11</td>
<td>0.32</td>
<td>0.33</td>
<td>0.30</td>
<td>0.29</td>
</tr>
<tr>
<td>3</td>
<td>1.89</td>
<td>0.82</td>
<td>0.64</td>
<td>0.61</td>
<td>0.14</td>
</tr>
<tr>
<td>4</td>
<td>1.55</td>
<td>0.63</td>
<td>0.46</td>
<td>0.38</td>
<td>0.17</td>
</tr>
<tr>
<td>5</td>
<td>1.17</td>
<td>0.43</td>
<td>0.52</td>
<td>0.62</td>
<td>0.00</td>
</tr>
<tr>
<td>6</td>
<td>1.76</td>
<td>0.60</td>
<td>0.52</td>
<td>0.44</td>
<td>0.31</td>
</tr>
<tr>
<td>7</td>
<td>2.42</td>
<td>0.96</td>
<td>0.62</td>
<td>0.60</td>
<td>0.04</td>
</tr>
<tr>
<td>8</td>
<td>1.44</td>
<td>0.62</td>
<td>0.46</td>
<td>0.43</td>
<td>0.00</td>
</tr>
<tr>
<td>9</td>
<td>1.37</td>
<td>0.68</td>
<td>0.53</td>
<td>0.50</td>
<td>0.16</td>
</tr>
<tr>
<td>10</td>
<td>1.26</td>
<td>0.57</td>
<td>0.51</td>
<td>0.52</td>
<td>0.09</td>
</tr>
<tr>
<td>11</td>
<td>1.35</td>
<td>0.54</td>
<td>0.59</td>
<td>0.63</td>
<td>0.10</td>
</tr>
</tbody>
</table>

Note: Item wordings are to be found in Appendix A. The response scale ranges from 1 to 4, with higher scores indicating or recoded to indicate more prejudiced responses. All correlations are significant at $p < 0.001$. The highest factor loading for each item in the exploratory analysis is italicized.
All factor loadings for the correlated two-factor solution (see Table 1, right-hand part) were significantly (ts varying between 3.4 and 8.2, ps varying between 0.000 and 0.001) greater than 0. Further, the correlation between the modern and the classical factor in the correlated two-factor model was found to be high, \( r = 0.55, p < 0.000 \). Although correlated, this pattern of results shows that a modern and a classical attitude factor concerning people with intellectual disabilities could be distinguished.

### 2.2.3. Discriminant validity

The discriminant validity of the modern and classical scales was examined by conducting \( t \)-tests on the means for participants responding \textit{yes} and \textit{no} to questions regarding whether they had been educated about, had experience of, and had any type of contact with people with intellectual disabilities. The analyses were conducted for the modern (Table 3, upper part) and the classical (Table 3, lower part) scales separately. As can be seen in Table 3, the modern scale, with significant differences between yes- and no-responders on all six items, shows a better discriminant validity than the classical scale, where differences between yes- and no-responders were found on only two of the six questions.

However, to arrive at a more proper psychometric assessment of discriminant validity, we conducted a factor analysis on the validation variables in Table 3. The analysis revealed a one-factor solution, tentatively denoted Knowledge and Experience of People with Intellectual Disabilities, indicating a highly shared variance among the variables. Based on the factor analysis, we calculated factor scores for each participant which were then correlated with participants’ total scores on the modern and classical prejudice scales. The analyses revealed a significant correlation of the factor scores with modern prejudice, \( r = -0.25 (p < 0.00) \), but not with classical prejudice, \( r = -0.10 (p < 0.13) \). A further analysis revealed that the correlation between the modern scale and the discriminatory factor was significantly \( [t(232) = 2.49, p < 0.01] \) higher than the corresponding figure for the classical scale.

An analysis of the mean scores showed that the mean modern scale score (\( M = 1.49, \text{ S.D.} = 0.31 \)) was significantly (\( t = 7.47, p < 0.000 \)) higher than the mean classical scale score (\( M = 1.32, \text{ S.D.} = 0.31 \)). This difference in means, we argue, is a further evidence for the discriminant validity of the modern and classical scales. More important, the difference in means supports the non-reactivity of the modern scale on the one hand and the overt nature of the classical scale on the other.

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### Table 2

Goodness-of-fit indices for three confirmatory factor models in Study 1 and Study 2

<table>
<thead>
<tr>
<th>Model</th>
<th>( \chi^2 )</th>
<th>d.f.</th>
<th>( \chi^2/d.f. )</th>
<th>GFI</th>
<th>RMESA</th>
<th>NFI</th>
<th>NNFI</th>
<th>CFI</th>
</tr>
</thead>
<tbody>
<tr>
<td>Study 1</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>One-factor</td>
<td>307</td>
<td>152</td>
<td>2.01</td>
<td>0.875</td>
<td>0.068</td>
<td>0.559</td>
<td>0.667</td>
<td>0.705</td>
</tr>
<tr>
<td>Uncorrelated two-factor</td>
<td>282</td>
<td>152</td>
<td>1.86</td>
<td>0.895</td>
<td>0.055</td>
<td>0.594</td>
<td>0.720</td>
<td>0.752</td>
</tr>
<tr>
<td>Correlated two-factor</td>
<td>245</td>
<td>151</td>
<td>1.62</td>
<td>0.906</td>
<td>0.048</td>
<td>0.647</td>
<td>0.795</td>
<td>0.820</td>
</tr>
<tr>
<td>Study 2</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>One-factor</td>
<td>309</td>
<td>152</td>
<td>2.03</td>
<td>0.817</td>
<td>0.087</td>
<td>0.599</td>
<td>0.705</td>
<td>0.739</td>
</tr>
<tr>
<td>Uncorrelated two-factor</td>
<td>295</td>
<td>152</td>
<td>1.94</td>
<td>0.842</td>
<td>0.073</td>
<td>0.617</td>
<td>0.730</td>
<td>0.761</td>
</tr>
<tr>
<td>Correlated two-factor</td>
<td>256</td>
<td>151</td>
<td>1.70</td>
<td>0.854</td>
<td>0.066</td>
<td>0.667</td>
<td>0.799</td>
<td>0.824</td>
</tr>
</tbody>
</table>

\( N = 235 \) = 62, \( p < 0.000 \). All factor loadings for the correlated two-factor solution (see Table 1, right-hand part) were significantly (ts varying between 3.4 and 8.2, ps varying between 0.000 and 0.001) greater than 0. Further, the correlation between the modern and the classical factor in the correlated two-factor model was found to be high, \( r = 0.55, p < 0.000 \). Although correlated, this pattern of results shows that a modern and a classical attitude factor concerning people with intellectual disabilities could be distinguished.

**Note.** GFI, goodness-of-fit Index; RMESA, root mean square error of approximation; NFI, Bentler–Bonett’s normed fit index; NNFI, Bentler–Bonett’s non normed fit index; CFI, Bentler’s comparative fit index. Values in italics indicate the best fit within the study.
Examining the gender differences, we conducted two regression analyses with the participants’ gender as the independent variable and the total mean scores on the modern and classical scales as dependent variables. The analysis revealed a significant relation between gender and the modern as well as the classical scores \((r = 0.21, p < 0.001, \text{in both cases})\).

Specifically, and as expected, men scored higher than women on both the modern \((M = 1.44, \text{S.D.} = 0.32 \text{ for women and } M = 1.58, \text{S.D.} = 0.28 \text{ for men})\) and the classical \((M = 1.27, \text{S.D.} = 0.26 \text{ for women and } M = 1.41, \text{S.D.} = 0.36 \text{ for men})\) scales.

However, testing the regression slopes revealed no significant difference \([t(308) = 0.02, p = 0.49, \text{one-tailed}]\) between the gender–modern scale relation as compared to the gender–classical scale relation. This indicates that the magnitude of the difference between men and women was the same for the modern and the classical scales. This lack of difference limits the support for the discriminant validity of the scales.

### Table 3
Mean modern and classical prejudice toward people with intellectual disabilities for each discriminatory variable in Study 1

<table>
<thead>
<tr>
<th>Variable</th>
<th>Yes</th>
<th>No</th>
<th>t(233)</th>
<th>p</th>
</tr>
</thead>
<tbody>
<tr>
<td>M</td>
<td>S.D.</td>
<td>N</td>
<td>M</td>
<td>S.D.</td>
</tr>
<tr>
<td>Modern Education about people with intellectual disabilities</td>
<td>1.45 0.28 89</td>
<td>1.52 0.33 146</td>
<td>1.76 0.04</td>
<td></td>
</tr>
<tr>
<td>Experience of people with intellectual disabilities</td>
<td>1.40 0.28 40</td>
<td>1.51 0.32 195</td>
<td>2.22 0.01</td>
<td></td>
</tr>
<tr>
<td>Know people with intellectual disabilities</td>
<td>1.36 0.28 35</td>
<td>1.52 0.31 200</td>
<td>2.89 0.00</td>
<td></td>
</tr>
<tr>
<td>Meet people with intellectual disabilities</td>
<td>1.43 0.31 52</td>
<td>1.51 0.31 183</td>
<td>1.68 0.05</td>
<td></td>
</tr>
<tr>
<td>Care about people with intellectual disabilities</td>
<td>1.43 0.29 139</td>
<td>1.59 0.32 96</td>
<td>4.01 0.00</td>
<td></td>
</tr>
<tr>
<td>Work with people with intellectual disabilities</td>
<td>1.40 0.31 50</td>
<td>1.52 0.30 185</td>
<td>2.48 0.01</td>
<td></td>
</tr>
</tbody>
</table>

Classical Education about people with intellectual disabilities | 1.27 0.26 89 | 1.36 0.33 146 | 1.98 0.02 |
| Experience of people with intellectual disabilities | 1.32 0.29 40 | 1.33 0.31 195 | 0.13 0.45 |
| Know people with intellectual disabilities | 1.29 0.21 35 | 1.33 0.32 200 | 0.81 0.21 |
| Meet people with intellectual disabilities | 1.32 0.34 52 | 1.33 0.30 183 | 0.06 0.48 |
| Care about people with intellectual disabilities | 1.27 0.28 139 | 1.40 0.34 96 | 3.11 0.00 |
| Work with people with intellectual disabilities | 1.29 0.25 50 | 1.34 0.32 185 | 1.02 0.15 |

*Note. All t-tests are one-tailed. Higher M indicates more negative attitudes. The response scale ranges from 1 to 4 for all scales.*

Examining the gender differences, we conducted two regression analyses with the participants’ gender as the independent variable and the total mean scores on the modern and classical scales as dependent variables. The analysis revealed a significant relation between gender and the modern as well as the classical scores \((r = 0.21, p < 0.001, \text{in both cases})\). Specifically, and as expected, men scored higher than women on both the modern \((M = 1.44, \text{S.D.} = 0.32 \text{ for women and } M = 1.58, \text{S.D.} = 0.28 \text{ for men})\) and the classical \((M = 1.27, \text{S.D.} = 0.26 \text{ for women and } M = 1.41, \text{S.D.} = 0.36 \text{ for men})\) scales.

However, testing the regression slopes revealed no significant difference \([t(308) = 0.02, p = 0.49, \text{one-tailed}]\) between the gender–modern scale relation as compared to the gender–classical scale relation. This indicates that the magnitude of the difference between men and women was the same for the modern and the classical scales. This lack of difference limits the support for the discriminant validity of the scales.

#### 2.2.4. Construct validity

For validation purposes, we computed product–moment correlation coefficients \((r)\) of total scores on the modern and classical scales with total scores on the SDO scale. The correlation analyses revealed that the modern \((r = 0.39)\) and the classical \((r = 0.34)\) scale scores were positively and significantly \((p < 0.000)\) correlated with the SDO scores, thus the higher people’s social dominance orientation the higher their prejudice toward people with intellectual disabilities. Assuming that social dominance is positively correlated with negative attitudes toward people with mental retardation (see Claesson et al., 2000; Ekehammar et al., 2004) and minority groups, the positive correlations lend support for the modern and classical scales’ construct validity.

Although the present results give support for the distinction between modern and classical attitudes toward people with intellectual disabilities, the outcome is to be replicated before
drawing a firm conclusion. This caution is further motivated by the fact that the internal consistency reliability of the classical scale was found to be low in Study 1.

3. Study 2

The main purpose of Study 2 was to replicate the confirmatory factor analyses conducted in Study 1, employing the same 19 items on another sample. Further, in addition to the Swedish version of the SDO scale (Pratto et al., 1994), we included scales measuring modern and classical racial prejudice (Akrami et al., 2000) and sexism (Ekehammar et al., 2000) as well as a scale measuring homophobic attitudes (Ekehammar & Akrami, in preparation). These scales were included for further assessment of the construct validity of the modern and classical scales measuring attitudes toward people with intellectual disabilities. In line with previous research (Ekehammar et al., 2004) and in accord with Allport’s (1954) theorizing, we expect the SDO, racism, sexism, and homophobia scales to correlate positively with modern and classical attitudes toward people with intellectual disabilities. Also, a further discriminant validation was conducted by examining the total mean for the modern and the classical scale. In accord with the original model of modern and classical attitudes, we expect higher scores on the modern as compared to the classical scale.

In addition, as in Study 1, we make a further examination of the discriminant validity of the scales on the basis of the participants’ gender, where we expect men to express more prejudicial beliefs than women. Further, we expect the gender differences to be larger for the modern scale as compared with the classical scale.

3.1. Method

3.1.1. Participants

There were 156 participants, 77 women and 79 men, aged 18–57 years ($M = 23.8$ years), students at Uppsala University and at the local authority-administered adult education. The former were students from various academic disciplines, such as social science, behavioral science, medicine, economics, technology, and dentistry. They received cinema vouchers for their participation.

3.1.2. Procedure and questionnaire

All participants were tested individually on a computer. All instructions and item wordings were presented on the computer screen. The participant’s responses were automatically stored in a file. All items were randomly mixed within each scale. Appropriate items were reversed when coded. The items of all scales were answered on 5-step Likert-type scales ranging from strongly disagree (1) to strongly agree (5). For all scales the average score across items was computed for all participants.

The computerized questionnaire consisted of the following scales: The Modern Attitude Scale toward People with Intellectual Disabilities, the Classical Attitude Scale toward People with Intellectual Disabilities, the Modern Racial (example: Discrimination against immigrants is no longer a problem in Sweden) and the Classical Racial Prejudice Scale (example: Immigrants do not keep their homes tidy; for reliability and validity information, see Akrami et al., 2000), the Modern Sexism (example: Discrimination of women is no longer a problem in Sweden) and the Classical Sexism Scale (example: Women are generally not very talented; for reliability and validity information, see Ekehammar et al., 2000), the Attitude to Homosexuality Scale.
3.2. Results and comments

3.2.1. Confirmatory factor analyses

As in Study 1, we conducted confirmatory factor analyses on the item intercorrelation matrix for the combined sample using LISREL 8.14. The fit indices for the tested models are presented in Table 2 (lower part). The results show that the correlated two-factor model fits the data well. Further, the correlated two-factor model showed a better fit than the uncorrelated two-factor model, $\Delta \chi^2 (1, N = 156) = 39, p < 0.000$, and the one-factor model, $\Delta \chi^2 (1, N = 156) = 55, p < 0.000$, whereas the two-factor uncorrelated model showed a better fit than the one-factor model $\Delta \chi^2 (1, N = 156) = 14, p < 0.000$. Further, the factor loadings for the correlated two-factor solution were significantly greater than 0 ($t$ varying between 3.5 and 8.3, $p$s varying between 0.000 and 0.001). These results replicate those from Study 1 and lend support for the distinction between modern and classical prejudice found in other attitude domains, and shows that the distinction is valid, also, for attitudes toward people with intellectual disabilities.

3.2.2. Discriminant validity

Analysis of the total mean scores of the modern and classical scales revealed that participants’ scores on the modern scale were significantly [$t(155) = 3.14, p < 0.002$] higher than their scores
on the classical scale (see Table 4). Again, the results replicate those from Study 1 and indicate the differential covert/overt nature of the modern and classical scales.

Two separate regression analyses with the participants’ gender as the independent variable and the total scale scores as dependent variables revealed a significant relation between gender and the modern ($r = 0.27$, $p < 0.001$) but not the classical ($r = 0.10$, $p < 0.21$) scale scores. Specifically, men ($M = 1.86$, S.D. = 0.48) scored higher than women ($M = 1.59$, S.D. = 0.48) on the modern scale, whereas there was no significant difference between men ($M = 1.64$, S.D. = 0.53) and women ($M = 1.54$, S.D. = 0.48) on the classical scale. Regression slope testing revealed a marginally significant difference [$t(308) = 1.52$, $p = 0.065$, one-tailed)] between the gender–modern scale relation and the gender–classical scale relation. Although marginally significant, the difference in regression slopes indicates that the magnitude of the difference between men and women is higher for the modern as compared with the classical scale. This is in line with our predictions and lends support to the discriminant validity of the scales.

3.2.3. Construct validity

Product–moment correlation coefficients ($r$) between the total scores of the modern and classical prejudice scales toward people with intellectual disabilities and the other prejudice scales were calculated. The results, presented in Table 4, show that these correlations are positive and relatively high, and further that the modern and classical scales were highly positively correlated. This pattern of correlations supports the construct validity of the scales and the contention that various types of prejudice are related.

4. General discussion

We hypothesized that attitudes toward people with intellectual disabilities, like those toward other minority groups and women, are expressed in a modern (covert) and a classical (overt) way. We suggested that these two forms of prejudiced attitudes are distinguishable, but correlated. We also predicted that attitudes toward people with intellectual disabilities are closely related to other types of prejudicial beliefs.

In Study 1, and in accord with our hypothesis, the results supported the correlated two-factor model. The results on the discriminatory aspects of the scales indicated the superiority of the modern scale and, at the same time, revealed the reactivity of the classical scale. It was also found that attitudes toward people with intellectual disabilities and social dominance orientation were correlated. Study 2 replicated the findings from Study 1 and gives further support to the distinction between a modern and a classical form of prejudice. Further, we found that the mean modern scale scores were significantly higher than those of the classical, which replicated the corresponding results of Study 1 and indicates that the responses to the classical scale are restrained due to its overt nature. The replicated gender differences in prejudicial attitudes lend support to the discriminant validity of the scales as well. This support was, however, slightly limited in Study 1, where the differences between men and women were equal for both scales. In addition, it was found that social dominance orientation and prejudicial beliefs, such as sexism, racism, homophobia, and negative attitudes toward people with intellectual disabilities share common variance (see also e.g., Claesson et al., 2000; Ekehammar & Akrami, 2003; Ekehammar & Sidanius, 1982).

We argue that the conclusions from the findings of the present paper are multifarious. From a theoretical point of view, replicating findings on other types of prejudice and from other cultural contexts, the present study supports the original model of covert and overt attitudes (for a review,
see Duckitt, 1992). Further, the results disclose the common structure or the common face of prejudicial beliefs (Allport, 1954). In addition, the findings integrate the research on attitudes toward people with intellectual disabilities with the main body of attitude research, specifically, with recent models of attitude change and expression of prejudicial beliefs. Finally, the findings offer practical support for the research on prejudice toward individuals with intellectual disabilities by signaling caution about the conventional scales’ sensitivity to, for example, social desirability concerns (Antonak & Livneh, 1994).

However, in recent years, the model of covert and overt attitudes has been subjected to criticism. Some researchers, questioning the distinction, have argued that, for example, the modern racial prejudice scale, like the classical scale, is reactive and susceptible to motivational bias (e.g., Coenders et al., 2001; Fazio, Jackson, Dunton, & Williams, 1995). Another line of criticism is based on the high correlation often found, like in the present study, between the modern and classical scales, arguing for that the two forms are not to be distinguished (e.g., Coenders et al., 2001). In addition, there have been signs of poor psychometric qualities for the classical scale, which can motivate a cautious position as to whether there is full psychometric support for the distinction. In the present study, for example, the reliability data for the classical scale can be considered low, at least in Study 1. Finally, the samples in the two present studies comprised university students, although from various academic disciplines, which might limit the external validity of our findings. The points above motivate some caution, but should, however, not overthrow the theoretical importance of the findings. Besides the theoretical importance of our findings, there is also an applied advantage of using the modern scale as a reliable instrument to assess negative attitudes toward people with intellectual disabilities more accurately as compared to the classical scale.

Finally, taking the question of attitudes toward people with intellectual disabilities to the main arena of attitude research, the present paper could be a natural step toward examining these attitudes’ more implicit or automatic function using an experimental design analyzing implicit rather than explicit attitudes (cf. Greenwald & Banaji, 1995). For example, what impact can the activated social category people with intellectual disabilities have on our implicit attitudes, and further, on our behavior and daily life social interaction?

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Appendix A

Item contents of the classical and modern scales of attitudes toward people with intellectual disabilities

Classical

1. The basic reasons for many of the social and economic problems that people with intellectual disabilities suffer from are due to their own mental weaknesses.
2. Even though there are some exceptions, it seems that most people with intellectual disabilities simply lack those qualities that community members should have.
3. People with intellectual disabilities should live in protected places because of the dangers in society.
4. It would be unwise for a person without any intellectual disability to marry a person with intellectual disabilities.
Appendix A (Continued)

5. People with intellectual disabilities do not have the character strength that people without intellectual disability have.
6. It seems that people with intellectual disabilities do not take the opportunities offered by the society.
7. Like all other people, people with intellectual disabilities have goals and meaning in their lives. (R)
8. People with intellectual disabilities often commit crimes.

Modern

Denial of continuing discrimination
4. Most people with intellectual disabilities are no longer victims of discrimination in Sweden.
8. People with intellectual disabilities are in general treated in the same way as people without intellectual disability in society.
9. Negative attitudes in society make the lives of people with intellectual disabilities difficult. (R)
10. It is easy to understand that people with intellectual disabilities and their relatives still struggle against the injustice they suffer in society. (R)

Antagonism toward demands
2. People with intellectual disabilities are getting too demanding in their push for equal rights.
3. People with intellectual disabilities have more to offer society than they have been given the opportunity to. (R)
6. The situation for people with intellectual disabilities is good as it is.
7. People with intellectual disabilities get too little attention in the media. (R)
11. There have been enough societal efforts in favor of people with intellectual disabilities.

Resentment about special favors
1. Society takes more care of people with intellectual disabilities than is fair to other groups.
5. It is right that people with intellectual disabilities sometimes get special support from society to find appropriate jobs. (R)

Note: Items with (R) have reversed coding.

References

Ekehammar, B., & Akrami, N. Attitudes to homosexuality among Swedish women and men, Uppsala University, Sweden, in preparation.


