Opinion Leadership Types or Continuous Opinion Leadership Traits?

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Accepted for publication in the *International Journal of Psychology*.

Abstract

Opinion leadership is typically conceptualized as a continuous personality trait. However, many authors adhere to the view of qualitatively different opinion leadership types and apply arbitrary criteria to split continuous trait scores into two groups (i.e., opinion leaders vs. non-leaders). The present study is the first to empirically evaluate this approach. A sample of $N = 3{,}812$ adults (67% women) was administered a validated opinion leadership scale. Finite mixture models examined whether the latent trait distribution can be represented by a set of discrete trait levels that reflected distinct opinion leadership types. The results did not give support to a discrete typology that distinguished leaders from non-leaders. Rather, opinion leadership was best characterized as a continuous trait.

Keywords: personality type, social influence, mixture modeling, non-parametric factor analysis

Opinion Leadership Types or Continuous Opinion Leadership Traits?

Individual differences in social influence determine the degree to which people can shape attitudes, decisions, and overt behaviors of their friends, family members, and coworkers. A trait reflecting the ability to informally influence others and thereby promote the diffusion of new ideas and trends in a social group is opinion leadership (cf. Batinic, Appel, & Gnambs, 2016; Flynn, Goldmith, & Eastman, 1996). Opinion leadership is a central concept in such diverse fields as marketing, political research, or health communication that has attracted worldwide interest (see Weimann, Tustin, van Vuuren, & Joubert, 2007). Recently, a fresh impetus to opinion leadership research around the world has resulted from the increasing popularity of social media and lead to numerous studies highlighting peer influences on discussion boards or social networking sites (e.g., Weeks, Ardèvol-Abreu, & de Zúñiga, 2015). Even international public opinion surveys such as the representative Eurobarometer[1] surveys conducted each month in the European member states routinely include measures to stratify attitudes on current topics by levels of opinion leadership. Although many of these studies administered different instruments that varied with regard to the precise construct definitions and the breadth of the operationalized constructs (Trepte & Scherer, 2010), the scales shared a common focus and described individuals that informally influence their social peer group. Moreover, in line with typical personality traits all these scales conceptualized opinion leadership as a continuous trait.

In contrast, some authors adhered to the idea of qualitatively different opinion leadership types that distinguished opinion leaders from non-leaders (e.g., Chan & Misra, 1990; Goldsmith & Flynn, 1994; Lyons & Henderson, 2005; Vernette, 2004). Thus, they used the instruments constructed for measuring continuous opinion leadership traits and, subsequently, applied some arbitrary criterion to split the sample into two artificial groups

[1] http://ec.europa.eu/COMMFrontOffice/publicopinion/

that supposedly distinguished opinion leaders from non-leaders. Some authors used the sample's mean or the theoretical mean based on the response scale to divide the sample into two groups (Chan & Misra, 1990). Others split their sample in such a way as to place a certain percentage of their sample into the group of opinion leaders (Goldsmith & Flynn, 1994; Lyons & Henderson, 2005; Vernette, 2004). They argued on theoretical accounts that only a minority of the population (e.g., 10%) is expected to act as opinion leader and, thus, should be viewed as such in a given sample. Common to these studies is that they created two artificial groups based on some arbitrary criterion without any psychometric evaluation if, indeed, the sample at hand represented a mixture of two distinct subpopulations.

As highlighted by several methodological studies (e.g., Bissonnette, Ickes, Bernstein, & Knowles 1990; Ruckera, McShanea, & Preacher, 2015) these post-hoc classifications can introduce a non-negligible bias into research findings. Discretizing continuous variables to create artificial groups not only results in a loss of power to identify meaningful effects but more seriously can also increase Type I errors by identifying spurious effects that do not actually exist. Therefore, it is important to empirically evaluate whether distinct subpopulations of respondents can be identified before creating opinion leadership groups. To address this research gap, the present study examined the responses from a large sample of adults on a validated opinion leadership scale. Finite mixture models (Hallquist & Wright, 2014) were estimated to identify potential subgroups of respondents that might reflect the opinion leadership dichotomy of leader versus non-leader. Moreover, age differences between different opinion leadership types were studied for different scoring schemes to highlight the danger of drawing misleading conclusions depending on the way the opinion leadership groups were created.

Method

Participants

The respondents were members of a commercial market research panel from Germany who repeatedly participate at anonymous web-based surveys in exchange for minor incentives (e.g., gift coupons). The present sample included 1,212 men, 2,496 women, and 104 individuals that did not report their gender. They were between 14 and 85 years old (M = 30.97, SD = 11.88). Most respondents had an educational level equivalent to university entrance qualifications (45%) or had already obtained a university degree (24%).

Instruments

Generalized opinion leadership was measured with nine items (e.g., “I usually succeed if I want to convince someone about something.”) on 5-point response scales from 1 “do not agree at all” to 5 “agree completely” (Gnambs & Batinic, 2011a). Previous research demonstrated good psychometric properties of this instrument including a unidimensional factor structure, high test-retest reliability, and good construct validity (e.g., Batinic et al., 2016; Gnambs & Batinic, 2011b, 2012). In the present sample, the scale score (M = 2.93, SD = 0.57) had coefficient alpha and omega hierarchical reliabilities of .84 and .71, respectively.

Statistical Analyses

The administered items were scaled using a graded response model (Samejima, 1969) that estimated a single latent factor representing the opinion leadership trait. Subgroups of respondents that might reflect different opinion leadership types were identified using a non-parametric factor approach (see Hallquist & Wright, 2014). Thus, finite mixture models with two or more latent classes were specified that estimated the probability of belonging to a given latent class for each respondent. Across the different classes strict measurement invariance of the latent factor was enforced. Moreover, the factor variances in each class were fixed to zero. In this way, the latent trait distribution can be represented by a set of discrete levels along the trait continuum with homogenous respondents within each class that reflect distinct opinion leadership types. Subsequently, the assumption of homogeneity within classes was relaxed by estimating different variances within class. This resulted in a semi-parametric factor model that allows for a potential non-normal trait distribution (see Hallquist & Wright, 2014). The number of subgroups (i.e., latent classes) were identified by a stepwise procedure comparing models with different numbers of classes based on the Bayesian Information Criterion (BIC) where lower values indicate models that more closely approximate the empirical data. Moreover, adjusted likelihood ratio tests (Lo, Mendell, & Rubin, 2001) were used to compare models with a given number of classes to one with one class less; statistical significant results would indicate more support for the model with more classes over the model with fewer classes. Finally, the sample size of each class was used to guide decisions on the practical relevance of a given class solution. A two-classes solution with higher opinion leadership scores in the smaller class would support the assumption of qualitatively distinct opinion leadership types (i.e., leaders vs. non-leaders), whereas multiple classes with approximatively linear increasing mean opinion leadership scores would rather fall in line with a continuous trait representation. All mixture models were estimated in M*plus* 7 (Muthén & Muthén, 1998-2012) with a robust maximum likelihood estimator.

Results

The factor score distribution of the opinion leadership trait for the entire sample without specifying any latent classes (see Figure 1) exhibited an approximatively normal shape and revealed no evidence of multiple local maxima that might indicate two or more qualitatively different respondent types. In the next step, various non-parametric factor models were estimated that specified between two and eight latent classes. The respective model fit indices are summarized in Table 1. The model with two classes exhibited an inferior fit as indicated by the significant ($p < .05$) likelihood-ratio test and the large BIC as compared to models with more latent classes. Thus, there was no support for two qualitatively distinct opinion leadership types. To examine whether the respondents might be grouped into three or more discrete levels along the trait continuum, I also examined the three to eight class models. However, the different model fit indices did not converge to a common solution. The likelihood ratio test identified no superior fit of the 4-class model as compared to the 3-class model (p = .29). However, the BIC of the latter was rather large as compared to models with more latent classes and, moreover, indicated a worse fit than the initial model without any latent classes (ΔBIC = 1,192.2). The best fit in terms of the BIC was achieved by a model with seven classes. However, in this model many classes were rather small; only three classes had class proportions exceeding five percent. More importantly, the latent factor means in the seven latent classes exhibited an approximately linear increase of the opinion leadership trait (see Figure 1). Thus, the assumption that opinion leadership would be better represented by qualitatively distinct types rather than a continuous trait yielded no support. Finally, the previous analyses were replicated by relaxing the homogeneity assumption within classes and estimating different variances for each class. Fit indices for one to three class models favored the two class solution with BIC_1 = 75,219.5, BIC_2 = 75,088.1, and BIC_3 = 75,101.6, respectively. However, the second class was quite small (1.4% of the sample) and, in contrast to opinion leadership theory, had a smaller mean (M = -0.33, SD = 5.63) than the larger class (M = 0.00, SD = 1.18). Again, the analyses did not support the assumption of a distinct group characterized by particularly high levels of opinion leadership.

To demonstrate the consequences of using arbitrary cutoffs for the derivation of opinion leadership types, three classification schemes were adopted. In line with prevalent practice (see Vernette, 2004), the top 5%, 10%, or 15% scorers on the opinion leadership scale were classified as opinion leaders, whereas the remaining sample were considered non-leaders. Given that opinion leadership is subject to pronounced age differences (Batinic et al., 2016), the standardized mean difference in the respondents' age was calculated for the three scoring schemes. If only the top 5% of the respondents were considered as opinion leaders, the age difference between the two groups would amount to Cohen's d = .20, 95% CI [.05,

.34]. In contrast, using either the top 10% or 15% scorers as opinion leaders would lead to age differences of *d* = .09, 95% CI [.01, .19] and *d* = -.02, 95% CI [-.06, .10], respectively. Thus, depending on the arbitrarily chosen classification scheme researchers would draw different conclusions regarding age differences between opinion leaders and non-leaders.

Discussion

Contemporary views on opinion leadership consider the concept as a continuous trait of interindividual differences in social influence (cf. Batinic et al., 2016; Flynn et al., 1996; Weimann et al., 2007). In practice, however, many researchers operate with an implicit typology and try to classify their respondents into two distinct groups that distinguish opinion leaders from non-leaders. Thus, they try to identify subgroups of particularly influential individuals within a sample. Despite the intuitive appeal of this approach (e.g., allowing for an easy communication of group comparisons on key variables to the general public), it is unknown whether empirical data actually reflects distinct opinion leadership types. Therefore, the present study applied finite mixture models to identify homogenous subgroups of respondents that fall in line with the hypothesized opinion leadership dichotomy. However, despite being based on a large sample including over 3,000 respondents and using a validated instrument these analyses found no support for discrete opinion leadership types. Rather, opinion leadership was best conceptualized as a continuous trait. These findings also align with several previous taxometric analyses showing that individual differences in personality are typically continuous rather than categorical (e.g., Foster & Campbell, 2007; Marcus, Lilienfeld, Edens, & Poythress, 2006). Therefore, previous attempts deriving two groups reflecting leaders and non-leaders should be viewed with due caution because there is little evidence that these subgroups actually exist. Moreover, given that discretizing continuous scores can introduce substantial biases into statistical results (e.g., Bissonnette et al., 1990;

Ruckera et al., 2015), the practice of deriving post-hoc opinion leadership groups without appropriate psychometric evaluations should be abandoned.

In conclusion, the study failed to provide support for qualitatively distinct opinion leadership types but suggested a continuous opinion leadership trait. Future studies are encouraged to extend this line of research to related instruments measuring other variants of opinion leadership (Flynn et al., 1996; Weimann et al., 2007). However, it is expected that the reported findings will generalize to many of these scales because they exhibit strong convergent validities with the administered instrument (Gnambs & Batinic, 2011b). Importantly, researchers adhering to the concept of different opinion leadership types are well advised to apply appropriate psychometric models (see Hallquist & Wright, 2014) to identify the hypothesized types before applying arbitrary criteria to create artificial groups for which empirical support may not be available. Particularly, international, cross-cultural studies need to demonstrate that comparable opinion leadership types exist in all samples before deriving conclusions on the determinants and consequences of opinion leadership in different countries.

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Table 1.

Fit Indices for Various Non-Parametric Factor Models

Model	*logLik*	Number of parameters	BIC	LRT-*p*	Smallest class proportion
0 classes [a]	-37.524.3	37	75,353.0	-	-
1 class	-41.996.2	44	84,354.6	-	-
2 classes	-39,012.2	46	78,402.9	< .001	.50
3 classes	-38,075.1	48	76,545.2	< .001	.21
4 classes	-37.696.2	50	75,803.8	.29	.04
5 classes	-37.471.9	52	75,371.7	< .001	.02
6 classes	-37,379.0	54	75,202.4	< .001	< .01
7 classes	-37,348.9	56	75,158.6	.05	< .01
8 classes	-37,341.7	58	75,160.7	.01	< .01

Note. *logLik* = Logarithm of the model likelihood; BIC = Bayesian Information Criterion; LRT-*p* = *p*-value associated with the adjusted likelihood ratio test (Lo, Mendel, & Rubin, 2001).

[a] Factor model with no latent classes and estimated latent factor variance.

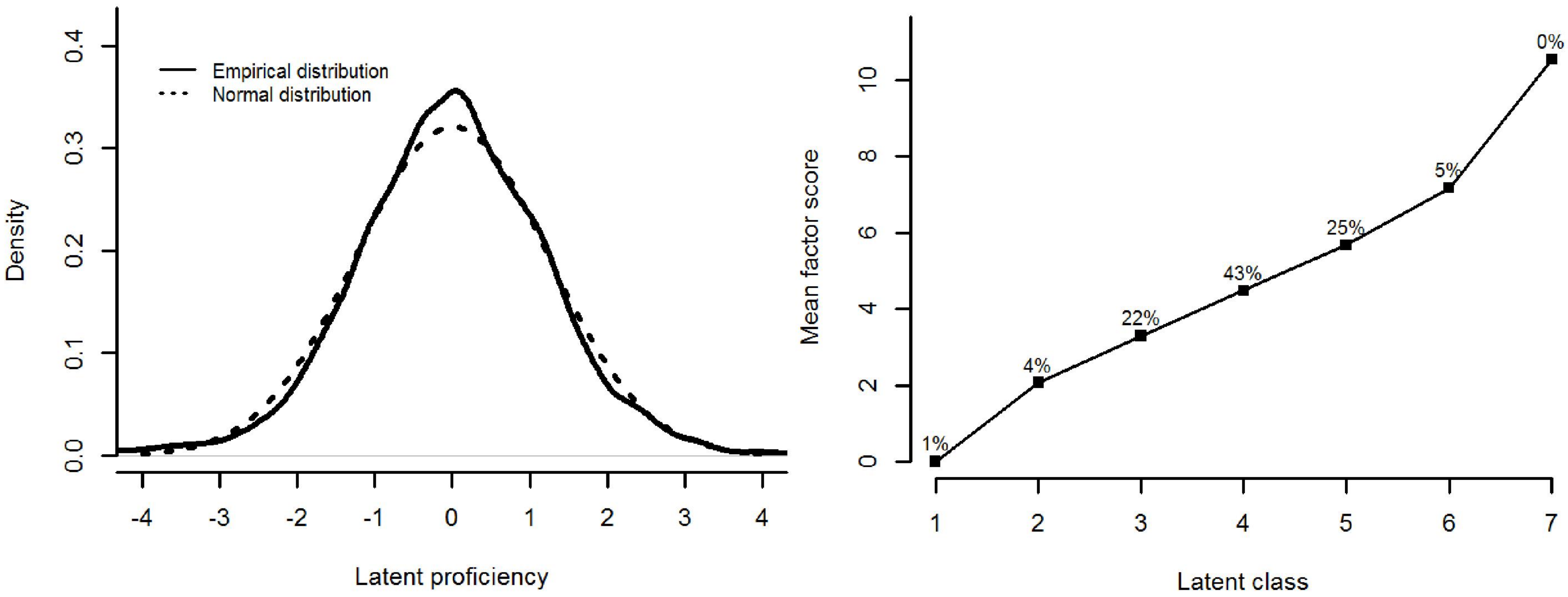

Figure 1. Factor score distribution of opinion leadership (left panel) and mean factor scores with proportions of sample size in latent classes (right panel)