Neuroticism and interpretive bias as risk factors for anxiety and depression

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Vinograd, M., Williams, A., Sun, M., Bobova, L., Wolitzky-Taylor, K. B., Vrshek-Schallhorn, S., Mineka, S., Zinbarg, R. E., & Craske, M. G. (2020). Neuroticism and interpretive bias as risk factors for anxiety and depression. Clinical Psychological Science, 8(4), 641-656. https://doi.org/10.1177/2167702620906145

Abstract:

Neuroticism has been associated with depression and anxiety both cross-sectionally and longitudinally. Interpretive bias has been associated with depression and anxiety, primarily in cross-sectional and bias induction studies. The purpose of the current study was to examine the role of interpretive bias as a prospective risk factor and a mediator of the relation between neuroticism and depressive and anxious symptoms in young adults assessed longitudinally. Neuroticism significantly predicted a broad general-distress dimension but not intermediate fears and anhedonia-apprehension dimensions or a narrow social-fears dimension. Neuroticism also significantly predicted negative interpretive bias for social scenarios. Negative interpretive bias for social scenarios did not significantly predict dimension scores, nor did it mediate the relation between neuroticism and general distress or social fears. These results suggest that although neuroticism relates to negative interpretive bias, its risk for symptoms of depression and anxiety is at most weakly conferred through negative interpretive bias.

Keywords: neuroticism | interpretive bias | anxiety | depression | open data

Article:

Major depressive disorder and anxiety disorders are some of the most prevalent of all psychiatric disorders; lifetime prevalence rates in the United States are 18.3% and 33.7%, respectively (Kessler, Petukhova, Sampson, Zaslavsky, & Wittchen, 2012). These disorders are associated with significant impairments in occupational, social, and physical functioning (McKnight & Kashdan, 2009; McKnight, Monfort, Kashdan, Blalock, & Calton, 2016). Their prevalence as well as their associated personal and societal cost underscore the need to accurately identify individuals at risk so that they may receive preventive care. Despite progress in identifying at-risk individuals, a significant subset is not accurately identified before the onset of symptoms, which typically occurs in late adolescence or young adulthood (Kessler et al., 2012). Elucidation of risk factors also informs specifically targeted prevention strategies and treatment interventions. For these reasons, the goal of our study is to extend existing research by evaluating the relation between two risk factors for depression and anxiety: neuroticism and interpretive bias.

Neuroticism is a trait disposition to experience negative affect that has been reliably measured as young as early childhood to midchildhood (Lamb, Chuang, Wessels, Broberg, & Hwang, 2002). It has been shown to relate to mood and anxiety symptoms in both clinical (Malouff, Thorsteinsson, & Schutte, 2005) and nonclinical samples (Jylhä & Isometsä, 2006; Saklofske, Kelly, & Janzen, 1995; Uliaszek et al., 2009). Furthermore, it has been associated with unipolar mood and anxiety disorders in cross-sectional studies (Kotov, Gamez, Schmidt, & Watson, 2010) and has been predictive of these disorders in longitudinal studies (Clark, Watson, & Mineka, 1994; Kendler, Gatz, Gardner, & Pedersen, 2006; Kendler, Kuhn, & Prescott, 2004; Zinbarg et al., 2016). Evidence to date supports the characterization of neuroticism as an established risk factor for unipolar mood and anxiety disorders (Griffith et al., 2010). Despite this robust association, the mechanisms through which neuroticism confers risk for depression and anxiety are not well understood.

Interpretive bias is a cognitive pattern in which individuals interpret ambiguous stimuli (e.g., words, scenarios) as having either positive or negative emotional qualities. Biased information processing has been posited as a key factor in the development and maintenance of unipolar mood and anxiety disorders (Beck & Haigh, 2014), and interpretive bias specifically is included in models of the development and maintenance of social anxiety disorder, generalized anxiety disorder, panic disorder, and depression (Hirsch, Meeten, Krahé, & Reeder, 2016). In terms of empirical studies, interpretive bias has been shown to relate to mood and anxiety symptoms (Constans, Penn, Ihen, & Hope, 1999; Huppert, Foa, Furr, Filip, & Mathews, 2003; Lawson, MacLeod, & Hammond, 2002; Lee, Mathews, Shergill, & Yiend, 2016). Moreover, it has been cross-sectionally associated with major depression, social anxiety disorder, generalized anxiety disorder, and panic disorder (Amin, Foa, & Coles, 1998; Everaert, Podina, & Koster, 2017; Hirsch et al., 2016; Mathews & MacLeod, 2005; Mogg, Bradbury, & Bradley, 2006; Stopa & Clark, 2000).

Longitudinal studies of interpretive bias are more limited in number. In a study of young adult women, interpretive bias was predictive of new onsets of panic disorder approximately 17 months later, even after adjusting for anxiety sensitivity and fear of bodily sensations (both established risk factors for panic disorder; Woud, Zhang, Becker, McNally, & Margraf, 2014). Negative interpretations of posttraumatic stress disorder (PTSD) symptoms within 4 months of a trauma predicted PTSD severity at 9 months (Dunmore, Clark, & Ehlers, 2001). Negative interpretive bias also predicted depression symptoms 4 to 6 weeks later in a sample of undergraduates (Rude, Wenzlaff, Gibbs, Vane, & Whitney, 2002), although symptoms were minimal in severity. Finally, negative interpretive bias predicted diagnoses of major depression during an 18- to 28-month follow-up period, even when adjusting for baseline depressive symptoms, although only for individuals who completed the task while under cognitive load (Rude, Valdez, Odom, & Ebrahimi, 2003).

Other work in adult samples has primarily relied on induction of an interpretive bias to assess causal relations, albeit limited to anxiety symptoms. In these studies, a bias toward a particular interpretation is trained through feedback, which resulted in increases in state anxiety and anxiety reactivity to a subsequent laboratory stressor following induction of a negative bias (Grey & Mathews, 2000; Mathews & MacLeod, 2002; Mathews & Mackintosh, 2000; Wilson, MacLeod, Mathews, & Rutherford, 2006). These studies suggest that experimental induction of interpretive bias can have a causal effect on vulnerability to anxiety. The possible causal effect of interpretive bias on depressive symptoms has been less well studied. One study found that never-disordered daughters of recurrently depressed mothers interpreted ambiguous words and stories more negatively and ambiguous words less positively than did daughters of never-disordered mothers following the induction of a sad mood (Dearing & Gotlib, 2009). Given that maternal depression is a significant risk factor for depression in offspring, these results suggest that interpretive bias before the onset of depression may be a marker of increased risk. Further prospective and longitudinal research is warranted to evaluate the role of interpretive bias as a risk factor for anxiety and depression. A primary goal of our study was to address this gap in the research.

The association between neuroticism and interpretive bias has been reliably demonstrated. For example, individuals high on neuroticism were more likely to make negative judgments on a number of cognitive tasks following a negative mood induction (Rusting, 1999) and to interpret ambiguous information in a negative way (Salemink & van den Hout, 2010) than individuals scoring low on neuroticism. Moreover, participants who scored low on trait anxiety and neuroticism and high on extraversion were more likely to produce positive interpretations of words in an auditory homophone task (Byrne & Eysenck, 1993). Byrne and Eysenck (1993) found that within the upper tertile of trait-anxiety, neuroticism was significantly correlated with the number of negative interpretations. Finally, studies using cognitive bias modification for interpretation have demonstrated reductions in trait anxiety, closely related to neuroticism, following the induction of a benign interpretive bias (Mathews, Ridgeway, Cook, & Yiend, 2007) or the reduction of a negative bias (Brosan, Hoppitt, Shelfer, Sillence, & Mackintosh, 2011). Although neuroticism and interpretive bias have been shown to relate to one another, no research has examined the relation between these two risk factors in the prospective prediction of psychopathology.

We sought to examine the relations among neuroticism, interpretive bias, and symptoms in a sample of young adults at risk for anxiety and depression. First, we tested whether neuroticism prospectively predicted interpretive bias. We hypothesized that neuroticism would significantly predict negative interpretive bias longitudinally given the existing cross-sectional evidence and the effects of interpretive-bias training on trait anxiety. Second, we tested whether interpretive bias prospectively predicted symptoms of depression and anxiety. In accord with cognitive theories (Beck & Haigh, 2014), we hypothesized that greater levels of negative interpretive bias would predict anxiety and depression symptoms. We used the trilevel model of depression and anxiety to model symptoms (Naragon-Gainey, Prenoveau, Brown, & Zinbarg, 2016; Prenoveau et al., 2010). The trilevel model includes a broad general distress structural factor, two intermediate factors (fears and anhedonia-apprehension1), and several disorder-specific narrow factors, including social fears. General distress accounts for the overlap in anxiety and depressive symptoms, whereas fears and anhedonia-apprehension capture symptoms unique to anxiety and depressive disorders, respectively. The social-fears dimension is most closely linked with social anxiety disorder clinical severity ratings and clinician ratings of social anxiety disorder symptom features. We chose to examine the social-fears dimension because of the social content of some of the interpretive-bias stimuli in our study as well as the robust evidence for interpretive bias for social content among individuals with social anxiety disorder (Hirsch et al., 2016).

We then tested whether interpretive bias partially mediates the relation between neuroticism and symptoms. The construct of neuroticism has been criticized as a “black box” that does not elucidate specific pathways by which negative affect contributes to psychopathology and as “a universal accompaniment of abnormal functioning (both psychological and biological) that by itself has little descriptive or explanatory value” (Claridge & Davis, 2001, p. 383; but see Zinbarg et al., 2016). Conceivably, interpretative bias is one pathway through which neuroticism confers risk. Individuals who are most vulnerable to negative affect may be especially likely to appraise outcomes as negative as a function of mood congruency or being better able to learn and remember material that is congruent in valence with one’s emotional state (Bower, 1981). Likewise, trait congruency, in which individuals process information in ways that are consistent with their underlying personality traits, suggests that individuals high in neuroticism would demonstrate a negative interpretive bias (Rusting, 1998). The evidence cited above regarding associations between neuroticism and negative interpretive bias (e.g., Byrne & Eysenck, 1993; Rusting, 1999; Salemink & van den Hout, 2010) supports the notion of mood and trait congruency, as does the finding that neuroticism positively correlated with affective priming for negative but not for positive words (Robinson, Ode, Moeller, & Goetz, 2007). Repeated interpretation of ambiguous events as negative in combination with other cognitive and biological factors might then lead to depression and anxiety (Mathews & MacLeod, 2005).

Indirect support for the mediational role of interpretive bias derives from evidence for negative relationship-specific interpretations mediating an association between high neuroticism and decreased relationship satisfaction in couples (Finn, Mitte, & Neyer, 2013). Furthermore, rumination and worry, cognitive processes that can be altered by interpretive-bias modification paradigms (Hirsch et al., 2016), have been shown to mediate the effect of neuroticism on depressive and anxiety symptoms in adult samples (Roelofs, Huibers, Peeters, Arntz, & van Os, 2008). Both of these studies, however, measured neuroticism, mediators, and outcomes at the same point in time. Temporal precedence has been identified as an important element of mediation models for establishing causality (MacKinnon, Fairchild, & Fritz, 2007). The longitudinal design of the current study allows us to establish the degree to which neuroticism predicts interpretive bias and the degree to which interpretive bias mediates the relation between neuroticism and subsequent symptoms, which fills an important gap in the literature.

The task used in the present study measured positive and negative interpretive bias toward both social and nonsocial scenarios. The social items broadly consisted of ambiguous interpersonal scenarios, whereas the nonsocial items related to bodily sensations, physical illness and flying, as well as scenarios about physical violence or harm to oneself or others. Studies of interpretive bias in individuals with panic disorder and generalized anxiety disorder have demonstrated disorder-specific biases of fear of bodily sensations and general threat, respectively, whereas interpretive bias regarding social scenarios may be particularly relevant for social anxiety disorder and depression (Hirsch et al., 2016). In the current study, negative interpretive bias for ambiguous social scenarios was hypothesized (a) to predict general distress, fears, anhedonia-apprehension, and social-fears dimensions and (b) to partially mediate the relationship between neuroticism and each of these dimensions, whereas negative interpretive bias for nonsocial scenarios was hypothesized (a) to predict general distress and fears and (b) to partially mediate the relationship between neuroticism and these dimensions only.

Method

Participants

Participants were recruited through two suburban high schools in Los Angeles and Chicago as part of the Youth Emotion Project (YEP). A total of 1,976 students, recruited from 3 consecutive years of 11th graders at each school, completed a 23-item version of the neuroticism scale of the revised Eysenck Personality Questionnaire (EPQ-R-N; H. J. Eysenck & Eysenck, 1975). The EPQ-R-N includes items such as “Does your mood often go up and down?” and “Do you worry about awful things that might happen?” Given the established association between neuroticism and depression and anxiety, students were categorized as either low (endorsement of 7 or fewer items), medium (endorsement of more than 7 but less than 12 items), or high EPQ-R-N scorers (endorsement of 12 or more items). In selecting the final sample of 1,269 students to invite for further study participation, the high-EPQ-R-N group was oversampled to increase the number of participants likely to develop unipolar mood and anxiety disorders in the follow-up period. Efforts were made to maintain equal numbers of male and female students across the EPQ-R-N categories. A total of 668 students agreed to participate in the study following parental assent, of which 627 completed their baseline assessment. Following the baseline assessment, neuroticism, depression, and anxiety symptoms were assessed once every 6 to 12 months via a battery of self-report questionnaires. All procedures were approved by the University of California, Los Angeles, and Northwestern University Institutional Review Boards.

A total of 199 participants completed the interpretive-bias task across both sites. One participant was excluded because he or she completed the interpretive-bias task twice, and 3 participants’ data were excluded because of mislabeling at data extraction. The interpretive-bias task was administered approximately 6 years after the baseline assessment for each cohort (n = 195; M = 6.44 years, SD = 0.92). Given the intervals of time used in existing longitudinal literature, neuroticism and symptom measures that were collected closest in time to the interpretive-bias task were used for the current analyses. Neuroticism and pretask symptoms were measured on average 6.70 months (SD = 11.13) before the interpretive-bias task, and posttask symptoms were measured on average 7.95 months (SD = 4.60) following the task. Studies examining the relation between neuroticism and interpretive bias have typically measured the two constructs at the same time (Byrne & Eysenck, 1993; Rusting, 1999; Salemink & van den Hout, 2010) or 1 month or less apart (Brosan et al., 2011; Mathews et al., 2007). Studies examining the longitudinal relation between interpretive bias and symptoms of psychopathology that did not rely on bias induction methods have had time intervals of weeks (Rude et al., 2002) to months (Dunmore et al., 2001). Therefore, the time intervals used in the current study are equivalent to or much greater than those used in previous studies. Symptoms before the interpretive-bias task were included in analyses to assess the degree to which interpretive bias predicted (or mediated) anxious and depressive symptoms when adjusting for preexisting symptoms. Of the 195 participants with usable interpretive-bias data, 195 had data for the preexisting trilevel model factor scores and neuroticism factor, and 172 had measures for the trilevel model outcome scores.

Measures

Interpretive-bias task The interpretive-bias task was completed on participants’ home computers via the online interface WebEx. The task was modeled from that used by Mathews and Mackintosh

(2000) and included 20 brief scenarios, three sentences in length on average. Ten scenarios were social situations, and 10 scenarios were nonsocial situations (see Table 1 for examples of a social and a nonsocial item). Participants were instructed to read each scenario carefully and imagine themselves in each scenario. Each scenario was presented with a title, and each ended with a target word with missing letters, represented by blank spaces, that participants were asked to complete by typing their response in a box that appeared below each scenario on the screen (e.g., "Your friend asks you to give a speech at her wedding reception. You prepare some remarks and when the time comes, get to your feet. As you speak, you notice some people in the audience start to l - - gh.") This target word did not resolve the ambiguity of the given scenario but was presented to facilitate encoding of the scenarios. On the following screen, participants were given the correct answer to the word and asked to indicate if their response was correct by typing "Y" for yes-correct or "N" for no-incorrect. Participants then answered a simple yes/no comprehension question to further improve encoding (e.g., "Did you stand up to speak?"). After reading and answering comprehension questions for all 20 scenarios, participants completed a filler task lasting between 2 and 5 min (a nonaffective go/no-go task).

Table 1. Example of Social and Nonsocial Items Used in Interpretive-Bias Task

Item elements	Example
Social scenario	Your friend asks you to give a speech at her wedding reception. You prepare some remarks and when the time comes, get to your feet. As you speak, you notice some people in the audience start to l - - gh.
Comprehension question	Did you stand up to speak?
Response options	• As you speak, people in the audience laugh appreciatively (positive) • As you speak, some people in the audience find your efforts laughable (negative) • As you speak, everyone in the audience bursts into applause (positive foil) • As you speak, you notice somebody in the audience start to yawn (negative foil)
Nonsocial scenario	You have been feeling dizzy occasionally, and decide to get a checkup. You make an appointment right away. Your doctor takes your blood pressure and listens to your chest, and then tells you to relax while giving you his o p - n - o n
Comprehension question	Did you delay before going to the doctor?
Response options	• The doctor tells you that there is absolutely nothing to worry about (positive) • The doctor tells you to relax and gives you an opinion on your disease (negative) • The doctor tells you that you have made a complete recovery (positive foil) • The doctor tells you that you will need another course of treatment (negative foil)

Following the filler task, participants were presented with the title of each scenario and four response options and asked to rank the degree to which each option was similar to the corresponding scenario on a 4-point Likert scale (1 = very different in meaning, 4 = very similar in meaning). The four options corresponded to a positive interpretation (As you speak, people in the audience laugh appreciatively), a negative interpretation (As you speak, some people in the audience find your efforts laughable), a positive foil interpretation (As you speak, everyone in the audience bursts into applause), and a negative foil interpretation (As you speak, you notice somebody in the audience start to yawn). The foil options were intended to be positively or negatively valenced but not logically related to the scenario. Participants were not instructed to rank order the four response options, that is, they could rate each option independently of the other three.

Self-report measures To construct the hierarchical neuroticism model (Zinbarg et al., 2016) and the trilevel model of anxiety and depression symptoms (Prenoveau et al., 2010), we used indicators from a number of self-report measures. In the hierarchical neuroticism model, we used subscale averages as indicators (see Zinbarg et al., 2016, supplemental material), and we used items as indicators for the trilevel model. Here, we report consistency coefficients of the neuroticism and cognitive vulnerability measures only. Hierarchical omega (ωh; Zinbarg, Revelle, Yovel, & Li, 2005) is reported in the case of multidimensional models. Alpha is provided as the internal consistency metric for subscales because it is a close approximation of ωh when a measure is unidimensional (McDonald, 1999, p. 93). Full models from the present analyses, including factor loadings from each model, can be found at https://osf.io/jn8hd/.

The neuroticism scale from the International Personality Item Pool-NEO-PI-R (IPIP-NEO-PI-R; Goldberg, 1999) includes 60 items rated on a 5-point Likert scale (1 = very inaccurate, 5 = very accurate). Items include "Become overwhelmed by events" and "Am often down in the dumps." Reliability was very good among the participants who completed the measure at the pretask assessment (α = .96; ωh = .71).

The Behavioral Inhibition Scale is a subscale of the Behavioral Inhibition Scale/Behavioral Activation Scale (BIS/BAS; Carver & White, 1994). The BIS consists of seven items rated on a 4-point Likert scale (1 = quite untrue of you, 4 = quite true of you) and includes items such as "I worry about making mistakes" and "Criticism or scolding hurts me quite a bit." The BIS/BAS has demonstrated acceptable convergent and discriminant validity (Carver & White, 1994). Among the subset of participants who completed the interpretive-bias task, internal consistency of the BIS was good (α = .83).

The neuroticism subscale of the Big Five Mini-Markers Scale (Saucier, 1994) consists of eight items. Respondents are asked to rate how accurately certain adjectives describe them on a 9-point Likert scale (1 = extremely inaccurate, 9 = extremely accurate). The scale includes adjectives such as "fretful" and "moody." Evaluated with the participants who completed the interpretive-bias task, reliability was sufficient (α = .85).

The Cognitive Style Questionnaire (Alloy et al., 2000; Haeffel et al., 2008) is a 24-item questionnaire that assesses inferential style for negative and positive hypothetical events. In the current study, only the 12 negative events were used. For each event, participants provide scores ranging from 1 to 7 on five scales: internality (α = .90), stability (α = .94), globality (α = .90), likelihood of negative consequences (α = .93), and negative implications about the self (α = .94).

The Dysfunctional Attitudes Scale–Form A (DAS; Weissman & Beck, 1978) is a 64-item questionnaire that combines the 40 items from the original version of the DAS with 24 items that are age appropriate from the Cognitive Vulnerability to Depression Project (Alloy et al., 2000). Participants endorsed items on a 7-point scale. The two DAS subscales, need for approval (α = .80) and need for achievement (α = .94), were reliable in our sample.

The Personal Style Inventory (PSI-II; Robins et al., 1994) is a 48-item questionnaire consisting of two scales: sociotropy and autonomy. Each item is endorsed on a 6-point scale. Robins, Bagby, Rector, Lynch, and Kennedy (1997) reported good to excellent reliability of the sociotropy (α = .90) and autonomy (α = .86) subscales in college students. Reliability coefficients were similar in the baseline measurement of our sample for both sociotropy (α = .90) and autonomy (α = .84). However, these scales were more highly correlated at baseline in our study than they were in Robins et al. (1994; r = .18). Reliability coefficients of sociotropy (α = .94) and autonomy (α = .90) were excellent in the present sample.

The Anxiety Sensitivity Index–Expanded Form (ASI-X; Li & Zinbarg, 2007) is a 29-item questionnaire that includes the 16 items from the original ASI (Reiss, Peterson, Gursky, & McNally, 1986) and 13 newer items. On this measure, participants evaluate statements regarding fears and cognitions about symptoms of anxiety in terms of how well each statement applies to them on a 5-point Likert scale (1 = very little, 5 = very much). Items are categorized into three subscales: physical concerns, social concerns, and mental incapacitation concerns. There was evidence of good general factor saturation ($\omega h = .71$), and each subscale alone evidenced excellent internal consistency in our sample (physical concerns: $\alpha = .91$; social concerns: $\alpha = .88$; mental incapacitation concerns: $\alpha = .94$).

Participants were administered 10 items from the 50-item Fear Survey Schedule–II (FSS-II; Geer, 1965). These items were scored to yield three specific fear subscales: animals, heights, and blood and injury. These 10 items were identified by Zinbarg and Barlow (1996) as excellent markers of specific fears. On this measure, participants rate their degree of fear to each object on a 7-point Likert scale (0 = none, 6 = terror).

The Situational Fears Questionnaire (SFQ) was adapted from the Albany Panic and Phobia Questionnaire (Rapee, Craske, & Barlow, 1994). The SFQ is a 22-item measure that assesses fear of activities that produce physical sensations and agoraphobic situations.

Participants completed the self-consciousness subscale of the Social Phobia Scale, or SPS (Mattick & Peters, 1988). Items describe situations that involve being observed or evaluated by others, such as public speaking, and participants rate how typical anxiety-related feelings are for them in those situations on a 5-point Likert scale.

The Inventory to Diagnose Depression (IDD; Zimmerman, Coryell, Corenthal, & Wilson, 1986) is a 21-item questionnaire that measures how respondents have been feeling in the past week. Items roughly correspond to the Diagnostic and Statistical Manual of Mental Disorders criteria for a major depressive episode. The IDD demonstrated adequate reliability and validity in a sample of college students (Goldston, O'Hara, & Schartz, 1990).

The Mood and Anxiety Symptoms Questionnaire (MASQ; Watson et al., 1995) is a 90-item measure that assesses for the severity of mood and anxiety symptoms in the past week. The modified MASQ used in our study excluded the suicidality item because of Institutional Review Board concerns and had a total of 89 items. The MASQ has five subscales: general distress, general anxiety, general depression, anxious arousal, and anhedonic depression. These subscales have shown strong internal consistency in student and adult samples in past research (Watson et al., 1995).

Data-analytic strategy

Data analysis proceeded in several steps. First, we evaluated the temporal measurement invariance of the trilevel model. Next, we extracted general distress, fears, anhedonia-apprehension, and social fears factor scores from configural invariant versions of the trilevel model (Prenoveau et al., 2010) at the YEP assessment waves relevant to the present study. Next, we used the hierarchical neuroticism model to test regressions between pretask and posttask general distress, fears, anhedonia-apprehension, social fears, and the general neuroticism factor (GNF). We then specified a structural equation model (SEM) to test the relations between the eight interpretive-bias variables and the GNF. Here, we sought to parse the distinctions between response bias as evidenced by an effect of the foil variables, as opposed to a more specific interpretive bias, evidenced by a significant effect of the nonfoil variables. Our next step was to examine whether negative interpretive bias for social scenarios, the only interpretive-bias variable that was significantly related to the GNF, predicted posttask general distress, fears, anhedonia-apprehension, or social fears, adjusting for these factors' pretask scores. Finally, we tested for mediation in two models: one in which posttask general distress was regressed on pretask general distress, the GNF, and negative interpretive bias for social scenarios and one in which posttask social fears was regressed on pretask social fears, the GNF, and negative interpretive bias for social scenarios. In each analysis, two indirect paths transmitting the effects of the GNF on the dimension were evaluated: paths from the GNF through negative interpretive bias for social scenarios and paths from the GNF through pretask scores on the dimension. Full results from the regression analyses can be found at https://osf.io/jn8hd/.

The temporal measurement invariance of the trilevel model was evaluated according to established fit standards: comparative fit index (CFI) $\geq .9$, root mean square error of approximation (RMSEA) ≤ 0.06, and standardized root mean square residual (SRMR) $\leq .08$ (Hu & Bentler, 1999). We note that data from the full YEP sample was used in these analyses rather than symptom data from only the subset of participants who completed the interpretive-bias task. To obtain separable indices of anxiety and depressive symptoms that accounted for their overlap, we extracted factor scores from configural invariant forms of the trilevel model at every assessment wave and then used those factor scores that corresponded to the preinterpretive and postinterpretive-bias task assessments. Extracted factor scores represented the general distress, fears, anhedonia-apprehension, and social-fears dimensions of the trilevel model.

We used a modification of the hierarchical model of neuroticism reported by Zinbarg et al. (2016) to test hypotheses regarding their GNF and interpretative bias. Zinbarg et al. reported acceptable fit for the hierarchical measurement model of neuroticism in the full sample of participants recruited into YEP who completed the baseline assessment, $\chi 2(339, N = 607) = 973.99$, $p < .001$; CFI = .92; RMSEA = 0.056; SRMR = .05. The primary modification used in our study was that the EPQ-R-N (H. J. Eysenck & Eysenck, 1975) was not included in the confirmatory factor analysis (CFA) model of neuroticism because the EPQ-R-N was only administered at the screening assessment. When the modified model used in our study was fit to the full sample of participants who completed the baseline assessment, model fit was acceptable, $\chi 2(228, N = 606) = 622.26$, $p < .001$; CFI = .93; RMSEA = 0.053; SRMR = .05.2 In addition, the GNF obtained high factor determinacy (.96) in our analyses, which made use of only a subset of participants from the baseline sample.

In all steps involving the GNF, we fixed the hierarchical neuroticism measurement model loadings to values obtained in Zinbarg et al. (2016). This was done for several reasons. First, the current sample is a subset of the sample used by Zinbarg et al., thus, the estimates obtained by Zinbarg et al. are more precise than estimates that could be obtained from this subsample. Second, fixing the neuroticism measurement model loadings reduced the number of parameters estimated for each SEM and thereby minimized identification and convergence problems (e.g., Bollen, 1996; Yuan & Chan, 2002). Doing so also prevented relations with symptoms and interpretive bias from biasing estimates of the measurement model loadings (e.g., Hoshino & Bentler, 2011).

We used Mplus (Version 8; Muthén & Muthén, 2017) to conduct all analyses. For each analysis, we accommodated missing data with full information maximum likelihood (FIML) estimation. For analyses involving neuroticism and invariance tests of the trilevel model, we used CFA. Such an approach confers several advantages in the context of our study, including the opportunity to connect our previously validated hierarchical model of neuroticism and hierarchical trilevel model of anxiety and depression symptoms to new hypotheses involving interpretative bias.

Results

Descriptive statistics

The final sample (N = 172) was 69.8% female and 30.2% male. This unequal gender distribution occurred because female students were more likely to agree to complete the EPQ-R-N, be invited to participate to fill the high-risk category, and agree to participate in the study if invited at baseline. The sample was 50.0% White, 14.1% Hispanic/Latinx, 13.4% African American/Black, 12.8% more than one ethnicity, 5.2% other, 4.1% Asian, and 0.6% Pacific Islander and were categorized as 16.9% low risk, 27.3% medium risk and 55.8% high risk on the basis of their EPQ-R-N scores. Participants were an average age of 23.28 years old (SD = 0.91) when they completed the task. Descriptive statistics and reliability coefficients (α, ωh) for the interpretive-bias task are presented in Table 2.

Table 2. Descriptive Statistics for Interpretive-Bias Task

Variable	*n*	*M*	*SD*	α	ω_h
Social scenarios					
Negative bias	170	28.60	5.21	.68	.49
Negative foil bias	170	15.02	5.38	.85	.73
Positive bias	170	27.40	5.61	.70	.56
Positive foil bias	170	16.15	5.63	.83	.60
Nonsocial scenarios					
Negative bias	170	25.61	6.07	.75	.63
Negative foil bias	170	17.67	5.88	.82	.69
Positive bias	170	26.86	5.70	.75	.48
Positive foil bias	170	17.89	5.82	.80	.61

Measurement invariance of the trilevel symptom model

We obtained adequate evidence of configural invariance at waves T1 through T13 (see Supplemental Material available online; Williams, Craske, Mineka, & Zinbarg, 2019). Model nonconvergence led us to specify slightly constrained versions of the trilevel model at T15 and T17, in which the first indicator's loadings and thresholds for each factor were constrained to their corresponding T1 values to aid model fit. The same procedure was conducted for Item 61 from the MASQ because this item contributed to model misfit. At the final assessment (T19), the configural invariant version of the trilevel model did not converge, and the aforementioned approach did not result in model convergence. We instead specified an alternative, fully scalar invariant model instead. Although model fit at this wave was poor, factor score correlations revealed no divergence from values obtained using previous waves. For fit statistics from each wave's final model, see the Supplemental Material.

Regressions

Neuroticism and trilevel symptomsOur first SEM contained the adjusted (see Data-Analytic Strategy) hierarchical measurement model of neuroticism (Zinbarg et al., 2016) as well as a regression of posttask general distress, fears, anhedonia-apprehension, and social fears on pretask general distress, fears, anhedonia-apprehension, social fears, and the GNF. Adjusting for pretask general distress, fears, anhedonia-apprehension, and social fears, the GNF significantly predicted posttask general distress, $\beta = 0.27$, $b = 0.22$, 95% CI = [0.07, 0.37], SE = 0.08, $p = .004$). The GNF was not a significant predictor of posttask fears, $\beta = -0.03$, $b = -0.02$, 95% CI = [−0.17, 0.12], SE = 0.07, $p = .734$; anhedonia-apprehension, $\beta = -0.09$, $b = -0.08$, 95% CI = [−0.26, 0.10], SE = 0.09, $p = .385$; or social fears, $\beta = -0.04$, $b = -0.03$, 95% CI = [−0.19, 0.13], SE = 0.08, $p = .733$.

Neuroticism and interpretive bias The GNF was significantly associated with the total score for negative interpretive bias for both social and nonsocial scenarios (see Table 3). Because participants did not generate their own responses to the scenarios and the task included a delay between scenario presentation and the ranking of possible interpretations, our interpretive-bias task is considered an “off-line” measure. Such measures can be confounded by response bias when participants selectively endorse negatively valenced response options regardless of their validity. This effect is more likely to occur when there is a delay between stimulus presentation and response selection (Hirsch et al., 2016). To address this limitation of the task used in the present study, we sought to parse interpretive bias from response bias for both negative social and negative nonsocial scenarios. As described above, the task includes corresponding negative and positive foil responses for each scenario. The foil responses are valenced but not logically related to the scenario; the scores on these items can be used as an index of response bias.

Table 3. Results From SEM Regression of Interpretive-Bias Variables on the GNF

Interpretive-bias variable	β	*b*	95% CI	*SE*	*p*
Positive social	0.01	0.07	[−0.71, 0.85]	0.40	.860
Negative social	0.16	0.80	[0.10, 1.50]	0.36	.025
Positive foil social	0.12	0.65	[−0.12, 1.41]	0.39	.096
Negative foil social	0.20	1.05	[0.34, 1.76]	0.36	.004
Positive nonsocial	0.07	0.38	[−0.43, 1.19]	0.41	.352
Negative nonsocial	0.14	0.85	[0.03, 1.67]	0.42	.044
Positive foil nonsocial	0.15	0.85	[0.06, 1.64]	0.41	.036
Negative foil nonsocial	0.24	1.40	[0.61, 2.18]	0.40	.001

Note: The 95% confidence interval (CI) estimates reflect the bounds of the unstandardized regression coefficients. GNF = general neuroticism factor; SEM = structural equation model.

To assess the effects of interpretive bias versus response bias, we first saved the residuals from a model in which negative bias scores were regressed on negative foil bias scores to obtain an interpretive-bias measure independent of response bias variance. We then specified a SEM regressing the residuals on the GNF. For the social scenarios, the regression coefficient between the residual and the GNF only approached significance, $\beta = 0.12$, $b = 0.62$, 95% CI = [−0.08, 1.31], SE = 0.35, $p = .081$. However, in such an analysis, it is also important to consider what percentage of the regression of negative bias on the GNF remains after residualizing for negative foil bias, which for social scenarios is approximately 75%, $\beta = 0.16$ vs. $\beta = 0.12$. In other words, negative response bias accounts for only a small portion, approximately 25%, of the significant association between the GNF and negative interpretations for social scenarios. This suggests that the association between the GNF and negative interpretations for social scenarios is more reflective of a negative interpretive bias than a response bias.

A similar analysis regressing the residuals in negative bias scores after accounting for negative foil bias scores for nonsocial scenarios, however, suggested that negative response bias accounts for the vast majority of the association between the GNF and negative interpretations for nonsocial scenarios. The regression coefficient between the nonsocial residual and the GNF did not near significance, $\beta = 0.004$, $b = 0.02$, 95% CI = [−0.70, 0.75], SE = 0.37, $p = .953$. Furthermore, the percentage of the regression of negative bias for nonsocial scenes on the GNF that remains after residualizing for negative foil bias for nonsocial scenarios is only approximately 3%, $\beta = 0.14$ vs. $\beta = 0.004$. Thus, the association between the GNF and negative interpretations for nonsocial scenarios is largely reflective of a response bias. As a result, negative bias for nonsocial scenarios was not tested as a predictor of the trilevel model factors or as a mediator of the relation between the GNF and symptoms.

Negative interpretive bias for social scenarios and trilevel symptoms Our next SEM tested the regression of posttask general distress, fears, anhedonia-apprehension, and social fears on pretask general distress, fears, anhedonia-apprehension, social fears, and negative interpretive bias for social scenarios. These analyses revealed significant regression coefficients for the trilevel factors' autoregressive paths. Pretask general distress, fears, anhedonia-apprehension, and social fears were significant predictors of posttask general distress, $\beta = 0.70$, $b = 0.66$, 95% CI = [0.56, 0.76], SE = 0.05, $p < .001$; fears, $\beta = 0.77$, $b = 0.79$, 95% CI = [0.70, 0.89], SE = 0.05, $p < .001$); anhedonia-apprehension, $\beta = 0.53$, $b = 0.51$, 95% CI = [0.39, 0.63], SE = 0.06, $p < .001$); and social fears, $\beta = 0.44$, $b = 0.48$, 95% CI = [0.34, 0.62], SE = 0.07, $p < .001$, respectively, when adjusting for negative interpretive bias for social scenarios. In addition, pretask anhedonia-apprehension was predictive of posttask social fears, $\beta = -0.16$, $b = -0.13$, 95% CI = [−0.23, −0.02], SE = 0.05, $p = .019$.

Adjusting for each of the pretask trilevel symptom factors, negative interpretive bias for social scenarios did not significantly predict posttask general distress, $\beta = 0.08$, $b = 0.01$, 95% CI = [−0.005, 0.03], SE = 0.009, $p = .151$; fears, $\beta = 0.05$, $b = 0.009$, 95% CI = [−0.008, 0.03], SE = 0.01, $p = .279$; anhedonia-apprehension, $\beta = 0.06$, $b = 0.01$, 95% CI = [−0.01, 0.03], SE = 0.01, $p = .339$; or social fears, $\beta = 0.004$, $b = 0.001$, 95% CI = [−0.02, 0.02], SE = 0.01, $p = .957$. Correlations between pretask general distress, fears, anhedonia-apprehension, and social fears with negative interpretive bias for social scenarios were generally small and nonsignificant (see Table 4).

Mediating effects of negative interpretive bias for social scenarios

Mediation of the GNF on posttask-general-distress effect To investigate the possibility of partial mediation of the effect of the GNF on posttask general distress through negative interpretive bias for social scenarios, we regressed general distress on its pretask scores, the GNF, and negative interpretive bias for social scenarios. In this analysis, we examined the strength of three paths through which the GNF may transmit its effects on general distress: through negative interpretive bias for social scenarios, through pretask general distress, or directly. We did so using the bootstrapping approach recommended by Preacher, Zhang, and Zyphur (2016). In total, the model

Table 4. Bivariate Correlations Among General Neuroticism Factor, Interpretive-Bias Task, and Trilevel Model Factor Scores

Variable	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16
1. GNF	—															
2. Positive social	−.01	—														
3. Negative social	.16*	.39**	—													
4. Positive foil social	.15	.23**	−.07	—												
5. Negative foil social	.26*	.07	.15	.72**	—											
6. Positive nonsocial	.06	.73**	.44**	.28**	.22**	—										
7. Negative nonsocial	.16*	.46**	.69**	.13	.33**	.52**	—									
8. Positive foil nonsocial	.16*	.36**	.06	.71**	.69**	.45**	.26**	—								
9. Negative foil nonsocial	.30**	.23**	.29**	.64**	.80**	.33**	.51**	.70**	—							
10. Pretask social fears	.12	.06	.06	−.04	.00	.06	.02	.04	.01	—						
11. Pretask fears	.11	.07	.14	−.20**	−.11	.09	.07	.02	−.05	.03	—					
12. Pretask AA	−.36**	.13	.03	−.20*	−.17*	.05	−.01	−.18*	−.18*	.03	.15*	—				
13. Pretask GD	.71**	−.12	.13	−.07	.06	−.05	.03	−.02	.12	.02	.02	−.09	—			
14. Posttask social fears	.12	−.07	.05	−.11	−.12	−.06	−.06	−.12	−.06	.45**	.11	−.13	.05	—		
15. Posttask fears	.13	.04	.17*	−.17*	−.03	.08	.11	.02	.03	.00	.78**	.15*	.09	.09	—	
16. Posttask AA	−.22**	.12	.08	−.13	−.04	−.01	.09	−.12	−.06	−.05	.08	.54**	−.03	−.14	.11	—
17. Posttask GD	.63**	-.04	.18*	-.04	.14	.00	.12	.05	.17*	.07	.09	-.09	.71**	.02	.04	-.02

Note: Social = interpretive bias for social scenarios; nonsocial = interpretive bias for nonsocial scenarios; AA = anhedonia-apprehension; GD = general distress.
*p < .05 (two-tailed). **p < .01 (two-tailed).

used 1,000 bootstrap draws. The total effect of the three paths combined was significant, $\beta = 0.63$, $b = 0.56$, 95% CI = [0.44, 0.67], SE = 0.06, $p < .001$. We also detected significant direct paths, $\beta = 0.27$, $b = 0.24$, 95% CI = [0.07, 0.47], SE = 0.09, $p = .009$, and indirect paths, $\beta = 0.35$, $b = 0.31$, 95% CI = [0.17, 0.44], SE = 0.07, $p < .001$, from the regressions of posttask general distress on the GNF and posttask general distress on the GNF through pretask general distress, respectively. However, the indirect effect of the GNF on posttask symptoms through negative interpretive bias for social scenarios was not significant, $\beta = 0.01$, $b = 0.009$, 95% CI = [−0.002, 0.04], SE = 0.01, $p = .336$ (see Fig. S1 in the Supplemental Material). Given these effect-size estimates, even if this indirect effect were significant, it is estimated to account for less than 3% of the total effect of the GNF on posttask general distress.

It is important to note that the estimate of the indirect effect of the GNF on posttask general distress through social interpretation bias has a very tight confidence interval (−0.002, 0.04). Thus, even if the population indirect effect size is the largest value in our confidence interval for it (0.04) and the population total effect of the GNF on posttask general distress is the smallest value in our confidence interval for it (0.44), the result shows that the indirect effect is still quite small and accounts for less than 10% of the total effect of the GNF on posttask general distress.

Mediation of the GNF on posttask-social-fears effect Although we did not observe a significant relation between the GNF and posttask social fears, previous research has found that negative interpretive bias in social anxiety disorder may be especially evident in the context of social stimuli, including scenarios (Hirsch et al., 2016). Given the significant effects of the GNF on negative interpretive bias for social scenarios, we constructed another mediation model to examine whether negative interpretive bias for social scenarios mediated the relation between the GNF and posttask social fears. Three paths from the GNF to posttask social fears were examined, in line with the previous analysis: the indirect effect of the GNF on posttask social fears through pretask social fears, the indirect effect of the GNF on posttask social fears through negative interpretive bias for social scenarios, and the direct path from the GNF to posttask social fears. As in the previous analysis, we used the Preacher et al. (2016) bootstrapping approach with 1,000 bootstrap draws to examine these effects. The total effect of the model was not significant, $\beta = 0.14$, $b = 0.11$, 95% CI = [−0.02, 0.24], SE = 0.07, $p = .100$ (see Fig. S2 in the Supplemental Material). The specific indirect path from the GNF to posttask social fears through negative interpretive bias for social scenarios was also not significant, $\beta = 0.001$, $b = 0.001$, 95% CI = [−0.01, 0.02], SE = 0.007, $p = .894$. In contrast, the indirect effect from the GNF to posttask social fears through pretask social fears was significant, $\beta = 0.06$, $b = 0.05$, 95% CI = [0.009, 0.10], SE = 0.02, $p = .035$.

Discussion

The aim of the current study was to examine the relation between two established risk factors, neuroticism and interpretive bias, in the prospective prediction of anxious and depressive symptoms in a sample of young adults. Neuroticism measured approximately 6 months before an interpretive-bias task significantly predicted a broad general distress factor measured approximately 8 months after the interpretive-bias task, when adjusting for preexisting general distress. Neuroticism also significantly predicted negative interpretive bias for social scenarios. Negative interpretive bias for social scenarios did not significantly predict general distress, fears, anhedonia-apprehension, or social-fears-symptom dimensions prospectively, nor did it significantly mediate the relation between neuroticism and general distress or neuroticism and social fears.

The evidence that neuroticism predicts general distress when adjusting for pretask general distress replicates prior findings, including prior analyses investigating the prediction of first onsets of anxiety and unipolar mood disorders in the present sample (Zinbarg et al., 2016). Neuroticism did not significantly predict posttask fears, anhedonia-apprehension, or social fears above pretask levels of these symptom dimensions. Taken together, our findings are consistent with the view that neuroticism is a marker of general rather than specific internalizing symptom risk (Jeronimus, Kotov, Riese, & Ormel, 2016; Kotov et al., 2010; Ormel, Jeronimus, et al., 2013).

Neuroticism significantly predicted negative interpretive bias for social scenarios prospectively. This finding expands on existing evidence for a cross-sectional relation between neuroticism and interpretive bias (e.g., Byrne & Eysenck, 1993) as well as experimental studies that have found effects of cognitive bias modification for interpretation on the reduction of trait anxiety (Brosan et al., 2011; Mathews et al., 2007). Other research on the prospective prediction of interpretive bias from neuroticism (e.g., Rusting, 1999) has been limited to shorter time frames (e.g., 1–2 weeks), whereas there was approximately 6 and a half months between the two measures in our study. Note that this effect was not attributable to negative response bias, which accounted for only a small portion (about 25%) of the relationship between neuroticism and negative bias for social scenarios. Our study analyzed the negative foil response option, meant to be valenced but not logically related to the target scenario, relative to the nonfoil option, which allows us to conclude that the relation between neuroticism and negative bias for social scenarios is indicative of an interpretive bias rather than response bias. This finding is especially important given that the task was an off-line measure and such measures are thought to be more prone to demand effects, selection bias, and response bias (Hirsch et al., 2016).

Although neuroticism also predicted negative bias for nonsocial scenarios, this effect appears to be largely due to a response bias rather than an interpretive bias. The social scenarios were generally related to interpersonal situations, whereas the nonsocial items included scenarios related to bodily sensations, physical illness and flying, as well as scenarios about physical violence or harm to oneself or others. Existing research that established a relation between neuroticism and interpretive bias has used social scenarios (Salemink & van den Hout, 2010) or valenced homophones (Byrne & Eysenck, 1993; Rusting, 1999). It is possible that associations between neuroticism and interpretive bias for nonsocial items were mitigated by the wide variation in nonsocial content. It has also been suggested that the relation between neuroticism and interpretive bias may surface only when participants are under stressful conditions or have undergone a mood induction (Ormel, Bastiaansen, et al., 2013). In accord, as reported earlier, individuals high on neuroticism are more likely to make negative judgments on cognitive tasks after a negative mood induction (Rusting, 1999). Likewise, there is evidence to suggest that trait anxiety, closely related to neuroticism, and state anxiety interact to engender interpretive bias (M. W. Eysenck, 2000). Conceivably, elevated neuroticism predicts negative interpretive bias across nonsocial as well as social items only when state anxiety is also elevated, something that was not measured in our study.

Negative interpretive bias for social scenarios did not significantly predict posttask general distress, fears, anhedonia-apprehension, or social fears symptom factors when adjusting for each of the factors measured before the task. The posttask symptom factors were measured approximately 8 months following the interpretive-bias task, which is at the upper limit of time intervals used in existing prospective research. This relatively long interval may have contributed to the null results found for the role of interpretive bias as a predictor and mediator given that interpretive bias may predict dimensional symptoms only at shorter intervals. Furthermore, a lack of idiosyncrasy has been posed as a specific limitation of tasks in which participants rank disambiguated response options (Hirsch et al., 2016), as was done in the current study, and may have also contributed to our null findings. For example, it has been suggested that the use of stimuli that do not reflect depression-relevant concerns may lead to null findings (Everaert et al., 2017). In a study of panic disorder, panic-related, but not general threat-related, interpretive bias predicted new onsets of panic disorder (Woud et al., 2014). Furthermore, negative interpretations of PTSD symptoms predicted PTSD severity at 9 months (Dunmore et al., 2001), which suggests that disorder-specific stimuli may be pertinent to investigations of the prospective relation of interpretive bias and anxiety symptoms. The social content of our interpretive-bias task pointed to the relevance of the social-fears dimension of the trilevel model, which has been closely associated with clinician ratings of social anxiety disorder symptoms (Naragon-Gainey et al., 2016; Prenoveau et al., 2010). Inclusion of this dimension did not change the overall pattern of the results, however, because we did not detect evidence that negative interpretive bias for social scenarios was a prospective predictor of posttask social fears.

Finally, we hypothesized that negative interpretive bias would partially mediate the relation between neuroticism and dimensions of the trilevel model. Prior data provide indirect support for this hypothesis because rumination and worry, cognitive processes that can be altered by interpretive bias modification tasks (Hirsch et al., 2016), have been shown to mediate the effect of neuroticism on depressive and anxiety symptoms in adult samples (Roelofs et al., 2008). Our study design provides the temporal precedence necessary to test mediation, with neuroticism measured before the purported mediator (interpretive bias), which was measured before symptom outcomes. The current findings do not support the role of interpretive bias as a mediator of the relation between neuroticism and symptoms of anxiety or depression. Even our more specific test of the mediating effect of negative interpretive bias for social scenarios in the prospective relation between pretask and posttask social fears, the most conceptually relevant dimension of the trilevel model, failed to support a mediating role of interpretive bias. Therefore, the mechanism through which neuroticism confers risk for psychiatric symptoms remains unclear. One possibility, supported by the existing literature, is that individuals with greater levels of neuroticism are prone to increased emotional reactivity to stressors and use of maladaptive coping strategies (Gunthert, Cohen, & Armeli, 1999; Suls & Martin, 2005). Heightened emotional reactivity and maladaptive coping may then confer risk for anxious and depressive symptoms.

There are a number of limitations in the present study. First, our sample consisted of only 172 participants, a relatively small sample size given the multivariate analyses that were used. Some of the confidence intervals around our point estimates are large, and in these cases, it is possible that the population values are different from those obtained in the current sample. Thus, a larger replication sample would be valuable. Second, it is possible that interpretive bias is best conceptualized as a facet of neuroticism. Indeed, the model of neuroticism used in the current study includes cognitive facets of neuroticism, including dysfunctional attitudes and negative inferential style. Facets of a construct can have a causal role, however, akin to symptoms of a disorder having a casual role in maintaining a disorder. For example, negative cognitions are a facet of depression (i.e., most effective treatments for depression will lead to a reduction in negative cognitions) but also ascribed a prominent causal role in cognitive therapy of depression (Hollon, DeRubeis, & Evans, 1987). In the current study, we sought to test whether interpretive bias, a possible cognitive facet of neuroticism, has a causal role in the effects of neuroticism (i.e., symptoms). Finally, the current study only includes one assessment of interpretive bias, so we are unable to distinguish between effects of between-person as opposed to within-person associations. The use of repeated measurements of interpretive bias to address this limitation, especially as related to neuroticism, is an important next step for future research.

Although we did not find evidence for the mediating role of interpretive bias between neuroticism and symptoms of anxiety and depression—and suggest that any such mediating role is, at most, quite small—neuroticism significantly predicted general distress and negative interpretive bias for social scenarios. The findings lend support to existing research establishing neuroticism as a risk factor for internalizing symptoms and bolster evidence for neuroticism's role in the prospective prediction of interpretive bias among young adults. The identification of measurable prospective predictors of psychiatric symptoms is important for the improvement of prevention and intervention strategies.

Acknowledgments

M. Sun is now at the Department of Psychological and Brain Sciences, Dartmouth College. L. Bobova is now at the Psychology Department, Adler University, Chicago Campus. K. B. Wolitzky-Taylor is now at the Department of Psychiatry and Biobehavioral Sciences, University of California, Los Angeles. S. Vrshek-Schallhorn is now at the Department of Psychology, University of North Carolina, Greensboro. A version of the project in the current manuscript was presented by M. Vinograd at the 2016 Anxiety and Depression Association of America Annual Conference, San Francisco, CA.

Transparency

Action Editor: Scott O. Lilienfeld
Editor: Scott O. Lilienfeld

Author Contributions

M. Vinograd and M. G. Craske developed the study concept. M. Vinograd, S. Mineka, R. E. Zinbarg, and M. G. Craske contributed to the study design. Data collection and cleaning was overseen by L. Bobova, K. B. Wolitzky-Taylor, and S. Vrshek-Schallhorn. M. Vinograd, A. Williams, and M. Sun performed the data analysis and interpretation under the supervision of R. E. Zinbarg. M. Vinograd and A. Williams drafted the manuscript, and R. E. Zinbarg and M. G. Craske provided critical revisions. All of the authors approved the final version for submission.

Open Practices

All data have been made publicly available via Open Science Framework and can be accessed at https://osf.io/jn8hd/. The complete Open Practices Disclosure for this article can be found at http://journals.sagepub.com/doi/suppl/10.1177/2167702620906145.

This article has received the badge for Open Data. More information about the Open Practices badges can be found at https://www.psychologicalscience.org/publications/badges.

Notes

Declaration of Conflicting Interests

The author(s) declared that there were no conflicts of interest with respect to the authorship or the publication of this article.

Funding

The research reported and the preparation of this article was supported by National Institute of Mental Health Grants R01-MH065651 and R01-MH065652 (to M. G. Craske, S. Mineka, and R. E. Zinbarg) and T32-MH015750 (to M. Vinograd).

1. Anhedonia-apprehension was named anxious-misery in earlier publications using the trilevel model.
2. To compare the meaning of the factors in the original Zinbarg et al. (2016) model with the factors in the modified model used here, we saved factor scores from each model in the full sample of participants who completed the baseline assessment and correlated the two sets of factor scores. The correlations among corresponding factors across the two models had a mean of .956, a median of .990, and a standard deviation of .099. In contrast, the correlations among noncorresponding factors across the two models had a mean of .134, a median of .116, and a standard deviation of .137. Thus, the factors in the model used here show excellent convergent validity and discriminant validity with the factors in the original Zinbarg et al. model and have nearly identical meaning.

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