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# An examination of the psychometric properties of the Connor-Davidson Resilience Scale-10 (CD-RISC10) among accounting and business students

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ABSTRACT

Using a sample of 546 undergraduate accounting and business students from four US universities, one on the East Coast, two in the South, and one on the West Coast, this study examined the efficacy of the Connor-Davidson Resilience Scale-10 (Campbell-Sills & Stein, 2007; CD-RISC 10) for use with accounting and other designated business majors. The analyses included an examination of possible demographic differences in overall score, the scale's factor structure, the invariance of its factor structure across major and gender, the scale's reliability, and its convergent and divergent validity. The results indicate significant inter-major and gender differences in scores. Most troubling, female accounting majors report the lowest resilience levels, significantly below those recorded for male accounting majors, and male and female non-accounting majors. However, the scale has a common factor structure. We further find that a two-factor solution provides a superior fit to the data compared to the single factor structure used in most prior research. Spearman-Brown reliability coefficients, item-total correlations, and coefficient alphas each support the reliability of the items loading on the scale for the full sample, as well as for each of the above-referenced demographic subsamples.



## 1. Introduction

This study assesses the psychometric properties of the Connor-Davidson Resilience Scale 10 (Campbell-Sills & Stein, 2007; CD-RISC 10) among a sample of accounting and other designated business majors. The results provide evidence of the scale's utility as a measure of one's ability to persevere in the face of adversities thereby increasing the likelihood that one can better adapt to the demands of the academic environment. Validating this measure with these cohorts may be particularly beneficial given the scale's potential as a tool for assessing individual resilience levels as part of a broader interventional strategy to address students' concerns about inadequate resilience training prior to entering the workforce.

Although CD-RISC 10 has proven to be an efficacious resilience measure with numerous subgroups (Aloba, Olabisi, & Aloba, 2016; Coates, Phares, & Dedrick, 2013; Gucciardi, Jackson, Coulter, & Mallett, 2011; Smith, Emerson, & Schuldt,

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2018), including college students in general (Campbell-Sills & Stein, 2007), its suitability for measuring resilience in accounting and business students is not a foregone conclusion because business students have been found to differ from general population college students in substantive ways. For example, business students have been found to diverge from the generic student body across each of the Big Five personality dimensions (Lounsbury, Smith, Levy, Leong, & Gibson, 2009). Furthermore, while there is extensive research that differentiates the characteristics of business school students from the rest of the student body, there is also considerable evidence supporting the contention that the personality attributes of accounting students can be distinguished from those completing other business administration majors. For example, accounting majors tend to prefer structured activities, while management majors prefer more ambiguous tasks (Chacko, 1991). Noel, Michaels, and Levas (2003) found accounting majors to be more reserved, restrained, and practical and that they preferred working with ideas rather than people when compared to either marketing or information systems majors. Accounting majors were also found to be significantly more honest than marketing majors, but simultaneously lower in open-mindedness and flexibility (Kochunny, Rogers, & Ogbuehi, 1992), and Pringle, DuBose, and Yankey (2010) show that accounting students score higher in conformity and lower in extraversion relative to other business majors. Moreover, Fortin and Amernic (1994) found that accounting majors were lower in self-esteem than other business majors. This finding in particular has important implications because an individual's resiliency is directly associated with their self-esteem and self-efficacy (Byrne, Flood, & Griffin, 2014).

Accordingly, this study evaluates the efficacy of the CD-RISC 10 for accounting majors as a distinct population, in addition to a constituent of the overall business student population. We also conduct inter-group comparisons of mean CD-RISC 10 scores between designated accounting and other business majors. Such comparisons may shed light on possible connections between measured resilience levels of students in college, and entry-level staff turnover levels in accounting and other business fields. We also evaluate potential dissimilarities based on location, gender, and university with our sample of students from four geographically and mission-diverse universities in the United States. These results should provide an interesting view of students from different parts of the country who attend universities of different size with dissimilar missions, as well as providing for an assessment of the stability of the scale's psychometric properties.

This study's context differs from previous studies that examined the psychometric properties of the CD-RISC 10. Despite the scale's popularity, it has not been specifically validated for use with accounting and other business students. This, and the factor structure and potential gender invariance issues discussed below, motivate this effort to evaluate the CD-RISC 10's factorial validity among these student cohorts as well as the equivalence of the CD-RISC 10 measurement structure across gender lines. In addition, this study extends the literature by looking at sample differences across business majors and university types. A rigorous assessment of the CD-RISC 10 should provide evidence regarding its appropriateness for use with these distinct student populations, as well as its viability as a tool for human resource personnel at college business programs who wish to measure student resilience levels as part of their student counseling efforts.

Our analyses show that the scale reflects a two-factor structure with the elements loading on two unique dimensions: motivation and toughness, which each load on the single higher-order factor of resilience. We demonstrate that the scale has good construct validity across several dimensions including gender, major, and geographic location. We also show that although the relative levels of resilience vary across major and gender, the scale is invariant with regard for factor structure and factor loadings for each group.

The remainder of this study is structured as follows. First, we review the literature on resilience in college students and motivate why the examination of resilience in accounting and business students is a worthwhile endeavor. Second, we provide more detail on the development and nature of the CD-RISC 10, as well as issues that must be addressed before considering utilization of the scale with these student cohorts. Third, we outline the methods employed to assess the reliability and validity of the CD-RISC 10, followed by a presentation of the results. We next discuss the significance of the results in terms of the scale's utility in assessing resilience in both accounting and business students. Lastly, we conclude with an evaluation of the study's inherent limitations and provide suggestions for future research.

## 2. Literature review

Many students find entry into academia to be a stressful, demanding and emotional experience as they are thrust into a world that requires a high degree of initiative, autonomy, and self-regulation (Byrne et al., 2014). Students who lack the capacity to function efficiently in this stressful environment are more likely to prematurely withdraw from the university and/or suffer from poor academic performance while still in attendance (Byrne et al., 2014; Murtaugh, Burns, & Schuster, 1999). Not surprisingly, stress has become a prevalent topic of academic research which has linked stress to higher levels of academic burnout (Law, 2010), lower academic performance (Gabre & Kumar, 2012), and poor health (Lim, Tam, & Lee, 2013).

Of particular concern are the sources and potential remedies for stress as it relates to business students in general, and accounting students in particular. In an investigation that validated an instrument intended to capture feelings of stress in a population of business and accounting students, Smith, Rosenberg, and Haight (2014) recounted that business students are exposed to stress from a wide variety of sources, including heavy workloads, increases in tuition and attendant debt, the requirement to manage competing demands related to coursework, athletic and other campus activities, pressures to sustain a high level of performance to secure lucrative employment opportunities, and dealing with the often conflicting requirements of different courses. These difficulties are exacerbated by a host of outside influences such as the demands of outside

employment, familial obligations and/or participation in non-academic activities. In addition to the above-referenced health and performance consequences, business students are reported to have significant communication apprehension (Simons, Higgins, & Lowe, 1995) and generally exhibit lower well-being and higher tendencies toward substance use (Vansteenkiste, Duriez, Simons, & Soenens, 2006).

The potential for excessive stress in school to follow students to the workplace is particularly concerning. Law (2010) noted that "students enrolled in business programs at universities who experience high and persistent levels of stress in school may find themselves ill prepared for the additional stressors found in the workforce" (p. 195). In the public accounting arena for example, stress can arise from serving multiple stakeholders with competing or conflicting agendas (Bagley & Reed, 2012) as well as from time and budget pressures, work overload, and increasing commercialism (Buchheit, Dalton, Harp, & Hollingsworth, 2016). Excessive stress in public accounting has been linked to high levels of turnover and/or the intent to do so (e.g., Fogarty, Singh, Rhoads, & Moore, 2000; Hermanson, Houston, Stefaniak, & Wilkins, 2016) and job burnout has been shown to contribute to absenteeism (Herda & Lavelle, 2012) and both reduced job satisfaction and performance (e.g., Jones, Norman, & Wier, 2010; Smith, Davy, & Everly, 2007). Smith and Emerson (2017) used a sample of professional auditors to investigate the role of resilience in the association between stress and reduced audit quality practices and found that it exerted a negative influence on stress arousal, auditor burnout and attenuated the prevalence of dysfunctional audit behaviors.

The distinction between environmental stressors and cognitive-affective appraisal of those factors is critical to understanding the link between stress and resilience. The above-referenced discussion outlines a variety of challenges that students face during their studies and later in the workplace. These challenges (e.g., heavy course loads, outside work, time budget pressure, etc.) represent environmental stressors that all students face to varying degrees. However, stress arousal as defined by Girdano and Everly (1986), is the "fairly predictable arousal of psycho-physiological (mind-body) systems which, if prolonged, can fatigue or damage the system to the point of malfunction or disease" (p. 5). As Smith and Emerson (2017) note, individuals differ in their susceptibility to stressors, thus factors that would trigger excessive stress arousal in one person would have little or no impact on another. It is the cognitive-affective appraisal of a stressor as threatening that triggers stress arousal.

Because stressors are pervasive in the business school environment and the workplace, understanding potential remedies and ameliorating factors can provide valuable insight. Resilience may be such a factor, as it has been defined as one's capacity to adapt and endure when faced with stressors and adversities, and one's resilience level appears to influence one's threshold for reacting to stressors (Ong, Bergeman, Bisconti, & Wallace, 2006; Coutu, 2002). In an academic context, resilience may be defined as the "the heightened likelihood of success in school and other life accomplishments despite environmental adversities brought about by early traits, conditions, and experiences" (Wang, Haertal, & Walberg, 1994, p. 46). Martin and Marsh (2006) note that "academic resilience is relevant to all students because at some point all students may experience some level of poor performance, adversity, challenge, or pressure" (p.267).

Wagnild and Young (1993) characterize resilience as a generalized personality characteristic that mitigates the adverse effects associated with various stressors. This definition views resilience as a *trait* as opposed to a *state*, which can be thought of as a highly specific, one-time behavior (Block & Kremen, 1996). While resilience may be dispositional, it is undeniable that resilience exhibits state-like qualities in that it is responsive and adaptive to environmental factors and is capable of being developed and managed (Cadogan-McClean, 2009; Luthans, 2002; Norman et al., 2005; Rutter, 1985). In fact, Jacelon (1997) conceptualizes resilience as a combination of traits and learning. Ginsburg (2015) analogizes resilience to coping, and states that it is effectively a preventative action in which one engages to effectively handle stressors. Thus, resilience can be construed as a set of actions, thoughts and behaviors that an individual can learn over time.

Resilience can also be conceptualized as stress resistance, i.e., the ability to effectively cope with stressors without negative consequences. Resilient individuals are able to anticipate positive outcomes under adverse conditions, visualize work demands as challenges to be overcome, with confidence in their ability to succeed (Britt & Jex, 2015). Personality facets, coping mechanisms, and social support structures all influence one's level of resilience.

There is a growing body of empirical evidence that supports a significant negative relationship between resilience and stress in various academic contexts. For example, Ahern and Norris (2011) found resilience to have a negative effect on stress among a sample of 166 community college students in the Southeast U.S. Fishman (2012) found that the resilience factor of equanimity significantly mediated the relationship between perceived stress and both physiological and psychological stress reactions among a sample of 115 community college students in upstate New York.[1] Hartley (2012) conducted a multi-group analysis self-reported scores of resilience, mental health, and social support between 605 general population undergraduate students with those of 121 students in campus mental health offices that offered college counseling, psychiatric-support, and disability-support services. The author found that participants seeking assistance from campus mental health offices reported significantly lower levels of resilience, significantly higher levels of psychological stress, and significantly lower levels of satisfaction with social supports.

Fishman (2012), citing Bartone, Roland, Picano, and Williams (2008), states that, "Researchers are currently focusing on the existence of resilient behaviors in individuals who appear to have a psychological hardiness, which results in their ability to effectively deal with stressful situations and events that otherwise cause individuals to become psychologically and physically

[1] Fishman (2012), referencing Wagnild and Young (1990) define equanimity as, "the ability to create a balanced perspective in one's responses, avoiding extremes and being accepting of what life brings" (p. 1011).

distressed" (p. 5). This is reflective of Zunz (1998) who notes that individuals will call on their available resources to deal with situations they perceive as stressful including social support systems, problem solving skills, self-efficacy, and resilience. Smith and Emerson (2017) summarize this relationship by stating that, "In effect, resilience serves as a coping mechanism that individuals employ to attenuate the effects of stress by increasing the threshold at which stress arousal is activated" (p. 4).

Prior research into resilience among college students documents that resilient individuals exhibit improved symptomology in the face of stress (Campbell-Sills, Cohan, & Stein, 2006; Fredrickson, Tugade, Waugh, & Larkin, 2003) and demonstrate positive academic performance and persistence (Hartley, 2011). Resilience has also been associated with a number of attributes, including active coping (Campbell-Sills et al., 2006; Li, 2008; Perez, Espinoza, Ramos, Coronado, & Cortes, 2009; Steinhardt & Dolbier, 2008), family and social support networks (Miller, 2002; Perez et al., 2009), and an internal locus of control (Banyard & Cantor, 2004; Yeager & Dweck, 2012).

The benefits of resilience in college students have been documented in a variety of settings and contexts. In terms of navigating the many challenges of the college experience, researchers have shown that resilience is associated with easier transition and adjustment for new students (Banyard & Cantor, 2004; Yeager & Dweck, 2012), greater enrollment persistence through the academic year (Shields, 2001), and higher academic performance (Hartley, 2011; Perez et al., 2009; Yeager & Dweck, 2012). Resilience has also been connected to diminished levels of depression, anxiety and other negative psychiatric symptoms (Campbell-Sills et al., 2006; Fredrickson et al., 2003; Steinhardt & Dolbier, 2008).

Moreover, research on college resilience intervention programs also suggests that resilience is not an immutable trait but can be inculcated in students as a way to mitigate the deleterious effects of academic stressors (Byrne et al., 2014; Clauss-Ehlers & Wibrowski, 2007; Martin & Marsh, 2006; Steinhardt & Dolbier, 2008). Indeed, Yeager and Dweck (2012) further note that individual-level resilience can be developed in the classroom and that the subsequent enhanced resilience levels are associated with better academic performance.

Although resilience among accounting and business student populations has received little attention in prior research, there are indications that these students possess characteristics that may uniquely equip them to manage the stressors of rigorous business programs. For example, Lounsbury et al. (2009) find that business students scored higher for several resilience related personality traits, including extraversion, conscientiousness, tough-mindedness, assertiveness, and emotional stability. Both extraversion and conscientiousness were found to be associated with resilience among general population college students (Campbell-Sills et al., 2006), while emotional stability, assertiveness, and tough-mindedness align with other resilience-related factors such as active coping, social support, and internal locus of control.

Other studies using samples of business school students show that greater psychological capital (e.g., hope, optimism, motivation) is positively correlated with academic performance (Luthans, Luthans, & Jensen, 2012).

Nevertheless, there are also indications that the rigor of accounting coursework and other business major curricula may elevate students' stress, which may lead to maladaptive coping behaviors (Law, 2010). Vansteenkiste et al. (2006) find that business students' "lower well-being and higher substance" (p. 2892) use relates to their extrinsic (i.e., materialistic) value orientations, which suggests that they may have innate predispositions that exacerbate the harmful effects of environmental stressors. Moreover, business students report higher tendencies and more tolerance of cheating (Klein, Levenburg, McKendall, & Mothersell, 2007), which, similar to substance use, can be viewed as a maladaptive coping mechanism.

In sum, the extant research suggests that studying resilience in accounting and business students is a worthy endeavor. Improving our understanding of resilience in these student populations will likely serve an informative purpose for prospective employers, such as public accounting firms, whose success hinges critically on the work ethic and integrity of their employees. These propositions appear particularly cogent with respect to accounting majors, and despite the fact that resilience is a highly valued attribute that employers expect of new employees, many recent accounting graduates believe that their program did not adequately exploit the available opportunities for them to develop this skill (Webb & Chaffer, 2016).

In addition, evidence of the efficacy of collegiate resilience intervention programs offers the possibility of adopting similar intervention programs to alleviate excessive stress before the onset of negative personal and professional consequences. Thus, to help facilitate these ends, we seek to validate a resilience measurement instrument for use in these student populations by assessing its construct reliability, discriminate validity, and its psychometric properties.

## 3. The Connor-Davidson Resilience Scale

Connor and Davidson (2003) developed the original Connor-Davidson Resilience Scale to measure the "personal qualities that enable one to thrive in the face of adversity" (p. 76). The original self-report instrument consisted of 25 items. However, our analyses evaluate the abbreviated 10-item version described below. Campbell-Sills and Stein (2007) developed the CD-RISC 10 as a viable alternative to the original scale after a review of its efficacy revealed an unstable factor structure. The revised scale has been reported to possess excellent psychometric properties and has generally been shown to be unidimensional in nature. It has been utilized in varied settings with diverse subject groups,[2] and "captures the core features of resilience," i.e., "the ability to tolerate experiences such as change, personal problems, illness, pressure, failure, and painful feelings" (Campbell-Sills & Stein, 2007, pp. 1026–1027). The instrument uses a five point Likert scale to evaluate a person's ability to cope with stressors, deal with adversity, adapt to change, see problems in a humorous light, etc. (see Table 4 for a complete list of scale items).

[2] For a review, see http://www.cd-risc.com/user-guide.php.

The CD-RISC 10 is measured as the summated total for all of the individual items. The higher one's score, the higher one's level of perceived resilience.

The CD-RISC 10 has widespread popularity as evidenced by its use in at least 150 studies and by the 964 citations, (as of January 2019), of Campbell-Sills and Stein (2007). Despite the scale's popularity, there are issues that must be addressed before endorsing its viability for measuring resilience among accounting and business students. For example, while CD-RISC 10 has proven efficacious in measuring resilience in numerous subgroups (Coates et al., 2013; Gucciardi et al., 2011), including college students (Campbell-Sills & Stein, 2007), our literature review suggests business students differ from the population of college students at large across a variety of key personality dimensions (Lounsbury et al., 2009; Vansteenkiste et al., 2006).

In addition, Aloba et al. (2016) recently reported a two-factor solution for the CD-RISC 10 with a sample of student nurses in Southwestern Nigeria. Specifically, they found six items that loaded on a factor that they named Toughness, and four items that loaded on a factor they named Motivation (see Table 1 below for a complete list of scale items). Admittedly, this was with a disparate sample. However, this finding was replicated by Smith et al. (2018) using a sample of auditors employed at U.S. public accounting firms. Moreover, Green et al. (2014) also identified two unique factors when evaluating post-9/11 veterans, but these factors were derived from the original 25-item scale and are substantively different from the items used in the CD-RISC 10. Collectively, these results raise concerns about the factor structure of the scale when employed in a collegiate context.

Gender differences in CD-RISC 10 scores are also a concern that needs to be addressed before endorsing the scale's utilization with accounting and business students due to reports that the scale may not be gender invariant. For example, Campbell-Sills, Forde, and Stein (2009) found that in a general community population ($n$ = 764), women's scores were significantly lower than men on the CD-RISC 10, a finding replicated by Rahimi, Baetz, Bowen and Balbuena (2014) among a sample of Canadian medical students. Moreover, a recent meta-analysis found gender to have a significant moderating effect on the association between resilience and mental health (Hu, Zhang, & Wang, 2015). Smith et al. (2018) found that men had significantly higher scores on the CD-RISC 10 compared to their female counterparts, but also found the scale to be valid and the factor structure to be stable across each of these groups. These findings call into question the invariance of the CD-RISC 10's factor structure with regard to gender.

It should be noted that it is not our contention that accounting and business students are inherently different in the manner in which they process the items that comprise the CD-RISC 10. Thus, we do not have a specific *a priori* expectation for their levels of resilience as compared to other college students enrolled in high-stress majors (e.g., engineering). However, having a validated instrument available to quantify resilience for this population will be advantageous. Further, the inconsistencies noted above motivate us to examine the scale's psychometric properties to ensure that it is a valid and reliable

**Table 1**
Mean resilience score comparisons between accounting and non-accounting majors.[a]

Panel A: Comparisons of mean resilience scores by major and gender

| | Accounting majors | | | Non-accounting majors | | | | |
|---|---|---|---|---|---|---|---|---|
| | $n$ | μ | $\sigma$ | $n$ | μ | $\sigma$ | F-Value | p-value |
| Full sample: | | | | | | | | |
| ($n$ = 546) | 201 | *27.52* | *6.11* | 345 | *29.04* | 5.66 | *8.68* | *.003* |
| Gender: | | | | | | | | |
| Female ($n$ = 244) | 107 | *26.39* | 6.18 | 137 | *28.91* | 5.61 | *11.02* | *.001* |
| Male ($n$ = 295) | 94 | 28.80 | 5.79 | 201 | 29.26 | 5.66 | 0.428 | .514 |

Panel B: Two-way analysis of variance test results (dependent variable = resilience)

| Source | Type III SS | *df* | Mean squares | F-Ratio | p-value |
|---|---|---|---|---|---|
| Major | 274.992 | 1 | 274.992 | 8.241 | *0.004* |
| Gender | 236.814 | 1 | 236.814 | 7.097 | *0.008* |
| Major * Gender | 129.866 | 1 | 129.866 | 3.892 | *0.049* |
| Error | 17,851.465 | 535 | 33.367 | | |

Panel C: Bonferroni-adjusted test of inter-group differences

| | | | | 95% Confidence interval | |
|---|---|---|---|---|---|
| Gender*Major$_1$ | Gender*Major$_2$ | Difference | $p$-Value | Lower | Upper |
| ***Female Accounting*** | ***Male Accounting*** | ***−2.405*** | ***0.020*** | −4.568 | −0.243 |
| ***Female Accounting*** | ***Female Non-Accounting*** | ***−2.513*** | ***0.005*** | −4.486 | −0.539 |
| ***Female Accounting*** | ***Male Non-Accounting*** | ***−2.871*** | ***0.000*** | −4.702 | −1.041 |
| Male Accounting | Female Non-Accounting | −0.107 | 1.000 | −2.156 | 1.941 |
| Male Accounting | Male Non-Accounting | −0.466 | 1.000 | −2.377 | 1.446 |
| Female Non-Accounting | Male Non-Accounting | −0.359 | 1.000 | −2.053 | 1.336 |

The inter-group difference for each of the bold and italicized comparisons is statistically significant (at p < .05 or lower).

[a] While the full sample consisted of 546 respondents, only 539 designated gender. The analyses in Panel B and Panel C are based on the 539 respondents who designated gender.

instrument that can be utilized in future studies involving both accounting and other business majors, and as a tool to measure resilience by business school administrators.

## 4. Methods

### 4.1. Subjects

The full sample consists of 546 students matriculating in a variety business courses at three public universities, one on the East Coast, two in the South, and private university located on the West Coast. All hold AACSB business accreditation, and three of the four schools additionally hold separate AACSB accounting accreditation. Each university's human subjects committee reviewed and approved the instrument package, which was subsequently administered in class at approximately the same time during the academic term at each institution. The instructors were not in attendance during administration, and assurances of anonymity were given to the student participants. The demographics of this convenience sample ($n$ = 546) are as follows: 293 (54%) of the respondents attended the universities in the south, 169 (31%) came from the university in the east, and 84 (15%) attended the West Coast university. No significant inter-school differences were measured in mean CD-RISC 10 scores, nor on any of the scales used to assess its convergent or divergent validity, thus prompting us to use the full sample data when analyzing the psychometric properties of the CD-RISC 10.

### 4.2. Analyses

The first set of analyses were designed to assess the significance of measured CD-RISC 10 mean score differences by major and gender. These differences were evaluated via a series of analyses of variance. We then examined the intra-scale correlations among individual CD-RISC 10 items.

We evaluated factor structure via a series of confirmatory factor analyses (CFA) using maximum likelihood (ML) estimation procedures that incorporated Satorra and Bentler (2001) scaling corrections. These procedures calculated the Satorra-Bentler chi-square value (SB$\chi^2$) and additional robust statistics (described below). Byrne (2006) argues that evaluating measurement instruments by means of CFA analysis is appropriate for scales "that have been fully developed and their factor structures validated"[3] (p. 118). The Satorra-Bentler rescaled estimate was appropriate with our sample due to the high Mardia's normalized estimate values generated in preliminary ML analyses indicating that the data were not normally distributed.[4] Bentler and Wu (2002) state that the Satorra-Bentler scaled $\chi^2$ "is the most widely studied and generally accepted best alternative test statistic for model evaluation under nonnormality" (p. 250). We began by evaluating the one-factor model with the full sample. We then assessed the two-factor structure uncovered by Aloba et al. (2016) given: (1) the intuitive appeal of the two factors; (2) the statistical robustness of their results; (3) that their sample of student nurses represented a professional, albeit culturally distinct, population; and, (4) the above-referenced confirmatory evidence reported by Smith et al. (2018). We next evaluated a second-order factor model that incorporated the above-referenced two factors, but related both to a second-order factor to explain the predicted significant correlation between the two first-order components, given that the majority of prior studies uncovered a univariate factor structure for the scale.[5,6]

Our demographic finding (reported below) of significant inter-major and inter-gender differences in mean CD-RISC 10 scores motivated us to assess possible inter-group scale factor structure differences. We also considered the output of the Wald model tests to determine if there were statistically non-significant parameters in the initial models.[7] Similarly, we evaluated the output of the Lagrange Multiplier (LM) test to assess whether model fit could be enhanced by adding additional model parameters.[8]

[3] This study's justification for the CFA analysis of the aforementioned CD-RISC 10 univariate factor structure model is predicated upon: (1) its popularity as noted above; (2) its favorable psychometric properties as noted in a methodological review of resilience scales by Windle, Bennett, and Noyes (2011); and (3) generally consistent results obtained from previous testing of its factor structure (with the exceptions noted above). Interestingly, in their comprehensive assessment of the psychometric qualities of 15 popular resilience scales, Windle et al. (2011) concluded that each of those scales were in need of additional validation.

[4] Our initial CFA output using the maximum likelihood (ML) method indicated significant multivariate kurtosis with a Mardia's Coefficient of 24.66 and a normalized estimate of 18.74. Bentler (2006) states that evidence of nontrivial positive kurtosis exists when normalized estimate values exceed 3.00, but further notes that modeling statistics may not be affected until the values exceed 5–6 or higher.

[5] To preempt any concerns with respect to our selection of Aloba et al. (2016) two-facture structure for our CFA analyses, we conducted an exploratory factor analysis of the CD-RISC 10's factor structure using a holdout sample of 560 randomly selected accounting and business majors from the target institutions that provided our primary study data. This principal components analysis with Varimax rotation produced two factors (discussed below) with similar factor loadings as those reported by Aloba et al. (2016).

[6] Gorsuch (1983) discusses the means by which factoring of correlations among primary factors can cause second order factors. Factoring a set of redundant variables may result in the emergence of multiple factors of narrow scope. A broader scope higher order factor may emanate from the correlation of these narrow factors. This premise motivated our testing of a second-order factor model.

[7] The Wald test is a post-hoc (i.e., not theory driven) procedure that is sample-specific. Therefore, to determine if this study's uncovered relations are applicable to other populations, replications with other samples are warranted.

[8] Byrne (2006) describes how the LM test is used on CFM tests to evaluate how selected indicator variables load on specific factors. In this study, we incorporated this test to uncover misfitting parameters, i.e., variables that should be added to a factor to enhance model fit. As in the case of the Wald test, confirmation with another sample is needed to determine whether the relations uncovered herein hold.

To assess overall fit of the above-referenced models, we evaluated the: (1) SB$\chi^2$ statistic; (2) Wheaton, Muthen, Alwin, and Summers (1977) relative/normed chi-square ($\chi^2/df$);[9] (3) the robust normed and nonnormed fit indices (NFI and NNFI); (4) the robust comparative fit index (CFI); (5) the Average Off-Diagonal Absolute Standardized Residual (AOASR); and, (6) the adjusted root mean squared error of approximation (RMSEA) for nonnormal conditions.[10] Bentler (1990) originally prescribed fit index values of 0.90 or higher as indicative of good model fit. However, Hu and Bentler (1999) revised the fit index cutoff value to close to 0.95, as well as a revised RMSEA cutoff of 0.06 or lower for relatively good fit. Jöreskog and Sörbom (1988) proposed that AOASR values of under 0.05 indicate good model fit.

We conducted additional assessments of the CD-RISC 10's reliability and validity. We evaluated the CD-RISC 10's internal consistency by calculating Cronbach's alpha and composite reliability coefficients for the full scale as well as the Motivation and Toughness subscales. We assessed convergent validity by conducting a correlation analysis of CD-RISC 10 scores with scores on the General Health Questionnaire-12 (GHQ-12: Goldberg & Williams, 1988) and the Academic Hardiness Scale (AHS: Benishek & Lopez, 2001) as revised by Creed, Conlon, and Dhaliwal (2013). The GHQ-12, developed to be a screening instrument for measuring general psychological health in community and non-psychiatric clinical settings (Molina, Rodrigo, Losilla, & Vives, 2014), has been widely used in both clinical and research settings.[11] (Creed et al., 2013) note that the Academic Hardiness Scale is based on Kobasa's (1979a, 1979b) conception of hardiness, and is intended to measure both overall academic hardiness as well as its the individual components of challenge, commitment, and control. As both psychological health and hardiness are associated with resilience (Creed et al., 2013) we expect to see a strong, positive and significant association between these two measures and the CD-RISC 10.[12]

We assessed divergent validity by two means. First, we correlated responses on the CD-RISC 10 with scores on the Stress Arousal Scale 4 (Smith, Everly, & Haight, 2012; SAS4) and the Maslach Burnout Inventory-General Survey Students (MBI-GS (S): Maslach, Jackson, & Schwab, 2016). The SAS4 is a four-item scale designed to capture cognitive-affective precursors to the physiological stress response. The MBI-GS (S) is a 16-item scale intended to capture symptoms of burnout in adult college and university students: it consists of three sub-scales: (1) Professional Efficacy measures satisfaction with past and present accomplishments and explicitly assesses one's expectations of continued effectiveness at school; (2) Cynicism captures an indifference or a distant attitude towards one's studies; and, (3) Exhaustion seeks to assess feelings of being overextended and exhausted by one's studies. Given that resilience as measured by the CD-RISC 10 is designed to measure one's ability to succeed under adverse circumstances (Connor & Davidson, 2003), we anticipate an attenuated or negative association between responses on the CD-RISC 10 and scores on the SAS4 and MBI-GS (S). We calculated both Pearson and Spearman rank correlations for both the convergent and discriminant validity assessments.[13] Because we conducted multiple significance tests simultaneously, we also evaluated the Pearson correlations' significance by computing Bonferroni adjusted probabilities (Wilkinson, 1999).

## 5. Results

### 5.1. Analysis of demographic differences and descriptive statistics

Of the 539 students reporting gender, 244 were women and 295 were men. Fifty percent ($n = 274$) of the 545 reporting age were 21–22 years old, 186 (34 percent) were 22–24 years old, 43 (8 percent) were under 20 years old, and 42 (8 percent) were over 24 years old. The majority of the 546 students reporting academic level were upper-classmen or graduate students. Specifically, 250 (46 percent) were seniors, 185 (34 percent) were juniors and 14 (3 percent) were graduate students. The remainder were sophomores ($n = 97$; 18 percent). Nearly 37 percent ($n = 201$) indicated that they were accounting majors, with the others reporting their primary major as management and marketing ($n = 149$; 27 percent), finance ($n = 79$; 14 percent), other (e.g., economics, entrepreneurship, general business, international business; $n = 80$; 15 percent), and information systems ($n = 37$; 7 percent).

Table 1 reports the results of a series of demographic comparisons of CD-RISC 10 scores between accounting and non-accounting majors.[14] As indicated in Panel A, the overall mean score for accounting majors (27.52) is significantly lower than that for non-accounting majors (29.04). In addition, while there is no significant inter-major mean score difference between males, the mean score for female accounting majors (26.39) is significantly lower than that for their non-accounting major counterparts (28.91). To further assess the significance of the major and gender differences, we conducted a two-way analysis of variance with Bonferroni-adjusted calculations of the significance of inter-group differences. Panel B indicates that there are

[9] Hooper, Coughlin, and Mullen (2008, 54) propose that, "the statistic that minimises the impact of sample size on the Model Chi-Square is Wheaton et al. (1977) relative/normed chi-square ($\chi^2$/df)". Although there is no consensus regarding an acceptable ratio for this statistic, recommendations range from as high as 5.0 (Wheaton et al., 1977) to as low as 2.0 (Tabachnick & Fidell, 2007).

[10] Fogarty et al. (2000) argue for the utilization of multiple measures to assess model fit as there is no one definitive measure.

[11] Gnambs and Staufenbiel (2018) concluded after two meta-analyses of the GHQ-12 with 76,473 and 410,640 participants, respectively, that the scale is essentially unidimensional in nature. Their findings prompted this study's incorporation of the GHQ-12 as a unidimensional measure of general health.

[12] The GHQ-12 items in this study are positively scored, i.e., the higher the total score, the more positive one's level of psychological health.

[13] Spearman rank correlations, being non-parametric in nature, are useful when the data are not normally distributed and provide statistically more conservative correlation coefficients.

[14] Though not reported in tabular format, each of the CD-RISC 10 inter-item correlations were positive and statistically significant with the full sample ($n = 546$) data.

significant inter-group differences by major, gender, and the interaction of major and gender. Panel C indicates that the source of these differences is the fact that the mean score for female accounting majors is significantly lower than that of any other group. In addition, though not reported in Table 1, the mean CD-RISC 10 score for the full sample ($\bar{x}$ = 28.48; s = 5.91; $n$ = 546) was significantly lower ($t$ = 10.51; $df$ = 1,31; $p$ < .001) than the score reported ($\bar{x}$ = 31.78; s = 5.41; $n$ = 764) for a "community sample" by Campbell-Sills et al. (2009). The mean CD-RISC 10 score for the sample excluding the female accountants ($\bar{x}$ = 29.05; s = 5.66; $n$ = 432) was also significantly lower ($t$ = 8.24; $df$ = 1,194; $p$ < .001) than that for the community sample.

### 5.2. Factorial validity analyses

Table 2 presents an evaluation of the goodness-of-fit statistics for the one, two and second-order factor models. Panel A of Table 2 provides the goodness-of-fit for the full sample. The NFI, NNFI, and CFI for the one-factor model all fall short of Hu and Bentler (1999) prescribed minimum of 0.95, while the SB$\chi^2$/$df$ ratio exceeds its recommended upper threshold of <2. However, the AOASR (0.031) and RMSEA (0.055) fall below their respective upper thresholds of 0.05 and 0.06, respectively.

The two-factor model test exhibited better fit than the one-factor model; however, post hoc (i.e., the WALD and Lagrange Multiplier) test output indicated that item 5 (i.e., can bounce back from sickness and hardship) should be an indicator of the Toughness factor, as opposed to the Motivation factor as previously specified in the Aloba et al. (2016) model. All of the goodness-of-fit indices for the revised two-factor model met their most rigorous benchmarks for well-fitting models. Moreover, this model exhibited better fit than the one-factor model as indicated by the results of the SB chi-square difference test ($\Delta$SB$\chi^2$ = 25.100, $df$ = 1, $p$ = <.001). Given its mathematical equivalence to the two-factor model, the second-order factor model test (Model 3) generated identical fit statistics.

### 5.3. Supplemental inter-group analyses

As reported in Panel C of Table 1, the mean CD-RISC 10 score for female accounting majors was significantly lower than the corresponding scores for male accounting majors, male non-accounting majors, and female non-accounting majors

**Table 2**
Goodness-of-fit indices CD-RISC 10 confirmatory factor analyses.

| MODEL | SB$\chi^2$ | df | p | SB$\chi^2$/df | NFI | NNFI | CFI | AOASR | RMSEA | RMSEA 90% Conf. Interval |
|---|---|---|---|---|---|---|---|---|---|---|
| PANEL A: Full sample ($n$ = 546) | | | | | | | | | | |
| 1. One factor model | 93 | 35 | <0.001 | 2.657 | 0.919 | 0.932 | 0.947 | 0.031 | 0.055 | 0.042–0.069 |
| 2. Two factor model[a,b] | 58 | 34 | 0.006 | 1.706 | 0.949 | 0.971 | 0.978 | 0.027 | 0.036 | 0.019–0.052 |
| 3. Second-order factor model | 58 | 34 | 0.006 | 1.706 | 0.949 | 0.971 | 0.978 | 0.027 | 0.036 | 0.019–0.052 |
| PANEL B: Inter-major analysis | | | | | | | | | | |
| *Female Accounting Majors* ($n$ = 107) | | | | | | | | | | |
| 1. One factor model | 50 | 35 | 0.043 | 1.429 | 0.834 | 0.923 | 0.940 | 0.053 | 0.065 | 0.012–0.101 |
| 2. Two factor model[c] | 46 | 34 | 0.088 | 1.353 | 0.850 | 0.941 | 0.955 | 0.050 | 0.057 | 0.000–0.096 |
| 3. Second-order factor model | 46 | 34 | 0.088 | 1.353 | 0.850 | 0.941 | 0.955 | 0.050 | 0.057 | 0.000–0.096 |
| *Male Accounting & Other Business Majors* ($n$ = 432) | | | | | | | | | | |
| 1. One factor model | 73 | 35 | <0.001 | 2.086 | 0.912 | 0.938 | 0.952 | 0.032 | 0.060 | 0.045–0.075 |
| 2. Two factor model[d] | 41 | 34 | 0.199 | 1.206 | 0.942 | 0.951 | 0.989 | 0.022 | 0.021 | 0.000–0.043 |
| 3. Second-order factor model | 41 | 34 | 0.199 | 1.206 | 0.942 | 0.951 | 0.989 | 0.022 | 0.021 | 0.000–0.043 |
| PANEL C: Multi-sample inter-group analysis | | | | | | | | | | |
| 1. Configural second-order model | 86 | 68 | 0.068 | 1.265 | 0.927 | 0.978 | 0.983 | 0.050, .022[g] | 0.031 | 0.000–0.050 |
| 2. Model with designated first-order factor loadings constrained to be equal between samples[e] | 92 | 75 | 0.094 | 1.223 | 0.923 | 0.982 | 0.985 | 0.056, .024[g] | 0.029 | 0.000–0.047 |
| 3. Model with designated first and second-order factor loadings constrained to be equal between samples[f] | 91 | 76 | 0.116 | 1.197 | 0.923 | 0.984 | 0.986 | 0.066, .026[g] | 0.027 | 0.000–0.046 |
| Standard for acceptance | NA | NA | >0.05 | <2.0–3.0 | >0.95 | >0.95 | >0.95 | <0.05 | <0.06–0.08 | NA |

NFI = Normed Fit Index; NNFI = Non-normed Fit Index; CFI = Comparative Fit Index; AOASR = Average Off-Diagonal Absolute Standardized Residual; RMSEA = Root Mean Square Error of Approximation.

[a] The Satorra-Bentler (SB)$\chi^2$ difference test does not capture the true mathematical difference between the two values, but rather a calculated scaled difference that takes into account the effect of non-normal data distribution.

[b] SB$\chi^2$ difference (in relation to the one-factor model) = 25.100, $df$ = 1, $p$ = <.001.

[c] SB$\chi^2$ difference (in relation to the one-factor model): female accounting majors = 8.040, $df$ = 1, $p$ = .005.

[d] SB$\chi^2$ difference (in relation to the one-factor model): male accounting & other business majors = 19.033, $df$ = 1, $p$ = <.001.

[e] SB$\chi^2$ difference (in relation to the configural model) = 4.780, $df$ = 7, $p$ = .687.

[f] SB$\chi^2$ difference (in relation to Model 2) = 0.134, d$f$ = 1, $p$ = .714.

[g] AOASR values are for female accounting majors, and male accounting and other business majors, respectively.

(there were no other significant inter-group mean score differences). To facilitate meaningful inter-group comparisons, we bisected the sample into female accounting majors ($n = 107$) and all other students that reported their gender ($n = 432$). Bisecting the sample in this fashion allows us to determine if there is a significant inter-group difference in scale factor structure analogous to the significant mean score difference between female accounting majors and the other aforementioned groups.

Panel B of Table 2 includes a summary of goodness-of-fit statistics for the one, two, and second-order factor models for the female accounting majors and other majors. The one-factor solution generated acceptable CFI (0.952) and RMSEA (0.060) values for the other majors, less so for the female accounting majors with only the SB$\chi^2$/*df* statistic meeting its rigorous standard of acceptance (2.0). Chi-square difference tests indicated that the two factor and second-order factor models were better fitting for both female accounting ($\Delta\chi^2 = 8.040$, $df = 1$, $p = .005$) as well as the other majors ($\Delta\chi^2 = 19.003$, $df = 1$, $p = <.001$).

We cross-validated the second-order factor model by testing for invariance between female accounting and other majors by incorporating the measurement models for each in a multi-sample analysis. First, using the procedure specified by Byrne (2006), we tested a multi-group model that specified the same factor loadings for each group but imposed no equality constraints on the parameters. Byrne (2006) states that this, "allows for the invariance tests to be conducted across the two groups simultaneously" and, "the fit of this configural model provides the baseline value against which all subsequently specified invariance models are compared" (p. 234). Panel C of Table 2 indicates good fit for this multi-group configural model. There was no significant loss of fit compared to the configural model ($\Delta\chi^2 = 13.737$, $df = 8$, $p = .088$) in the test of the invariance of the first-order factor loadings.[15] However, the Lagrange Multiplier test results indicated that the constrained factor loading for item 2 ("I can deal with whatever comes my way") should be released, indicating that it is not equivalent between female accounting majors and the other majors. Releasing this constraint also resulted in no significant loss of fit compared to the configural model ($\Delta\chi^2 = 4.780$, $df = 7$, $p = .687$). Similarly, there was no significant loss of fit ($\Delta\chi^2 = 4.344$, $df = 2$, $p = .114$) in the test for the invariance of both first and second-order factor loadings, i.e. the fit of Model 3 relative to Model 2.[16] However, the Lagrange Multiplier test results indicated that the constrained factor loading for item 3 ("I try to see the humorous side of things when I am faced with problems") should be released. Releasing this constraint also resulted in no significant loss of fit relative to Model 2 ($\Delta\chi^2 = 0.134$, $df = 1$, $p = .714$). Fig. 1 illustrates the standardized path coefficients for the second-order factor model.

### 5.4. Reliability analyses

Table 3 presents CD-RISC 10 mean scores, standard deviations, and reliability statistics for the full sample as well as the Motivation and Toughness subscales. Factor loading coefficients ranged from 0.510 for CD-RISC3 to 0.731 for CD-RISC9 on the full scale, 0.757 for CD-RISC3 to 0.777 for CD-RISC2 on the Motivation subscale, and 0.657 for CD-RISC4 to 0.752 for CD-RISC9 on the Toughness subscale. The Cronbach's alpha of 0.855 for the full scale and 0.831 for the Toughness subscale exceeded Nunnally (1978) prescribed threshold of 0.70, thus demonstrating adequate internal consistency, as did their respective Spearman-Brown reliability coefficients. While the alpha coefficient is not comparably supportive for the Motivation subscale, it should be noted that this value is directly impacted by the number of scale items. As Tavakol and Dennick (2011), citing Nunnally and Bernstein (1994) and Streiner (2003) note, "If the test length is too short, the value of alpha is reduced... Thus, to increase alpha, more related items testing the same concept should be added to the test" (p. 53). Thus, the lower reliability value measured for the Motivation subscale should be interpreted accordingly.

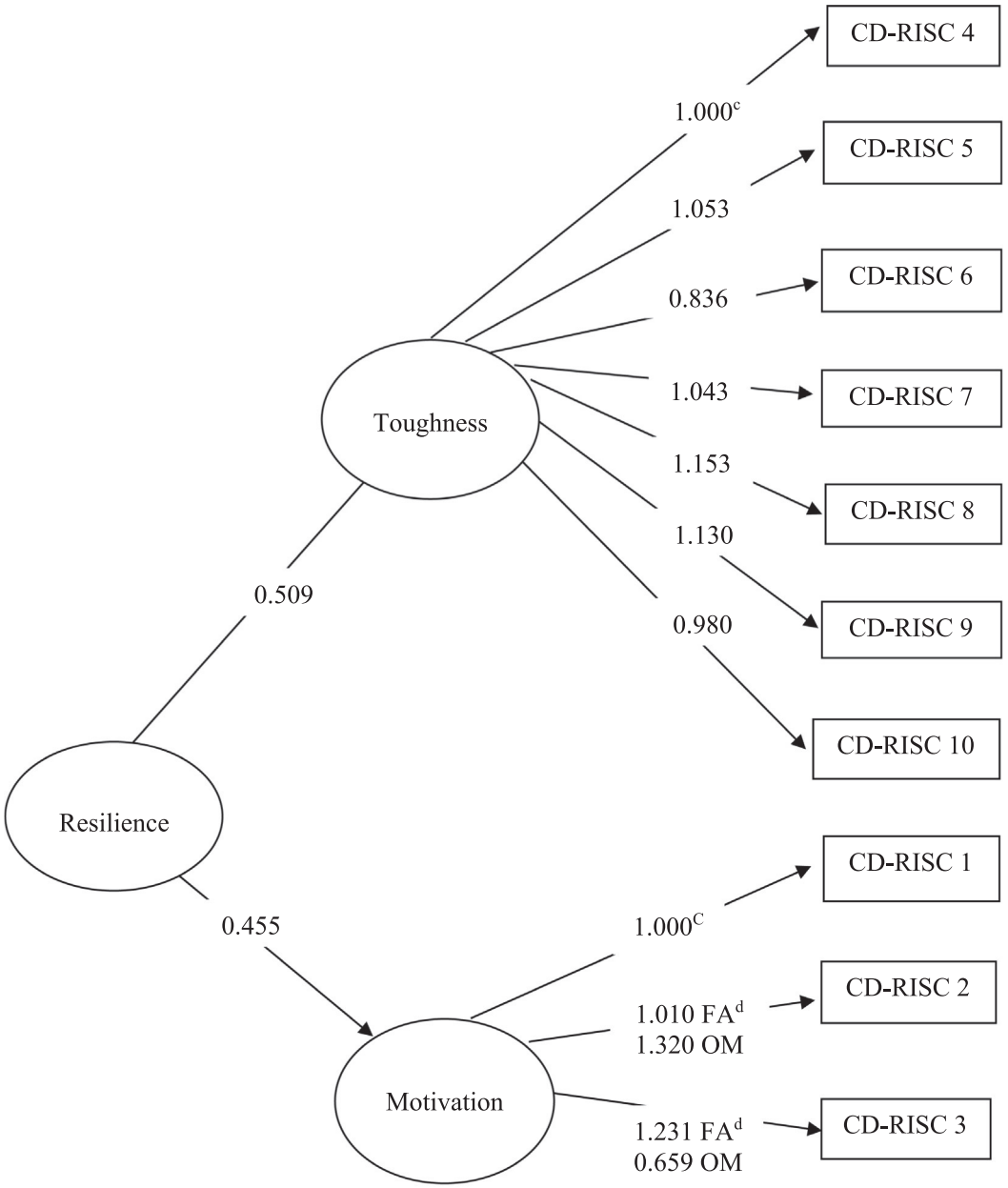
### 5.5. Convergent and divergent validity analyses

Table 4 presents the results of the correlations between the CD-RISC 10 scores and the measures designed to assess convergent and divergent validity. We measured significant positive correlations between the scores on the CD-RISC 10 and both those on the GHQ-12 and AHS scales, and with each of the AHS subscales. Conversely, there was a significant negative correlation between CD-RISC 10 scores and those on the SAS4 and MBI-GS (S) scales, and with each of the MBI-GS (S) subscales. The same pattern of correlations exists between the Motivation and Toughness dimensions of each of the other measures with one exception; the correlation (−0.112) between Motivation and the Exhaustion dimension of the MBI-GS (S) is non-significant.[17]

[15] Byrne (2006) states that, "invariance holds if goodness-of fit related to this (multi-group) model is deemed adequate and there is minimal difference in fit from that of the configural model" (p. 239). Byrne also notes that that when testing for invariance of factor loadings in second-order factor structures, it is customary to assess whether first-order factor loadings operate in the same manner between groups prior to assessing invariance of first and second-order factor loadings.

[16] Cheung and Rensvold (2002) define minimal difference in fit as a change in CFI between models of <0.01.

[17] Ginsburg (2008) argue that in situations where large sample sizes result in small correlations being statistically significant, that "the statistical significance is almost meaningless and should probably be ignored"(p. 233). Thus, we posit that the negative correlation between the CD-RISC 10 scores and those on the SAS4 and MBI-GS supports the divergent validity between the CD-RISC 10 and the latter two scales.




[a] All estimable coefficients between individual CD-RISC 10 items and first order factors, and between each first order factor and the second-order factor, are significant at $p < .01$.
[b] For diagramming ease and reader interpretability, error and disturbance terms are omitted.
[c] One measure of each construct must be fixed to 1.000 to establish the scale of that latent construct.
[d] FA = Female accounting majors, OM = Other majors.

**Fig. 1.** Multi-sample analysis path coefficients for the second-order factor model.[a,b]

### *5.6. Analysis of the relationship between resilience and stress arousal*

As noted in our discussion of the results in Table 4, there appears to be a significant negative relationship between resilience (as measured by the CD-RISC 10) and stress arousal (as measured by the SAS4). This finding is congruent with that of Smith et al. (2014, p. 52) who compared the scores on a two-item version of the CD-RISC to those on the Perceived Stress Scale −10 (Cohen & Williamson, 1988), a widely used measure of perceived stress. Moreover, Smith and Emerson (2017) found (via structural equations model analysis) a significant negative association between resilience and stress arousal

**Table 3**
CD-RISC 10 item means, standard deviations, item-total correlations, Cronbach Alphas, and Spearman-Brown Reliability Coefficients ($n$ = 546).

| Scale items | $\bar{x}$ | s | Item-total correlations | | |
|---|---|---|---|---|---|
| | | | Full Scale | Motivation | Toughness |
| CD-RISC 1. Able to adapt when change occurs (M) | 4.031 | 0.763 | 0.642 | 0.762 | |
| CD-RISC 2. Can deal with whatever comes my way (M) | 3.924 | 0.789 | 0.722 | 0.777 | |
| CD-RISC 3. Try to see the humorous side of problems (M) | 3.726 | 1.043 | 0.510 | 0.757 | |
| CD-RISC 4. Coping with stress can make me stronger (T) | 3.625 | 0.971 | 0.622 | | 0.657 |
| CD-RISC 5. Can bounce back from sickness and hardship (T) | 3.986 | 0.899 | 0.688 | | 0.724 |
| CD-RISC 6. Can overcome obstacles to achieve goals (T) | 4.305 | 0.750 | 0.653 | | 0.667 |
| CD-RISC 7. Under pressure can stay focused (T) | 3.601 | 0.928 | 0.681 | | 0.715 |
| CD-RISC 8. Not easily discouraged by failure (T) | 3.439 | 0.986 | 0.714 | | 0.741 |
| CD-RISC 9. Is a strong person when dealing with difficulties (T) | 4.025 | 0.870 | 0.731 | | 0.752 |
| CD-RISC 10. Can handle unpleasant or painful feelings (T) | 3.774 | 0.912 | 0.685 | | 0.687 |
| Cronbach alpha | | | 0.855 | 0.627 | 0.831 |
| Split-half Correlation | | | 0.806 | 0.515 | 0.765 |
| Spearman-Brown reliability coefficient | | | 0.893 | 0.680 | 0.867 |

**Table 4**
Pearson and Spearman Correlations of CD-RISC 10 with GHQ-12, AHS, SAS4, and MBI-GS(S)[a,b,c].

| Scale | CD-RISC 10 (Full scale) | Motivation (F1) | Toughness (F2) |
|---|---|---|---|
| General Health Questionnaire -12 | 0.450 (0.439) | 0.326 (0.311) | 0.449 (0.439) |
| Academic Hardiness Scale: | 0.421 (0.422) | 0.275 (0.284) | 0.434 (0.430) |
| Commitment | 0.225 (0.217) | 0.106 (0.108) | 0.250 (0.238) |
| Challenge | 0.233 (0.250) | 0.198 (0.235) | 0.220 (0.219) |
| Control | 0.465 (0.472) | 0.322 (0.327) | 0.457 (0.474) |
| Stress Arousal Scale 4 | −0.308 (−0.327) | −0.228 (−0.237) | −0.284 (−0.320) |
| Maslach Burnout Inventory – General Survey Students: | −0.376 (−0.381) | −0.242 (−0.229) | −0.389 (−0.397) |
| Exhaustion | −0.211 (−0.201) | *−0.112 (−0.100)* | −0.229 (−0.222) |
| Academic Efficacy | −0.359 (−0.380) | −0.242 (−0.257) | −0.367 (−0.383) |
| Cynicism | −0.286 (−0.303) | −0.198 (−0.199) | −0.289 (−0.310) |

[a] Spearman correlation coefficients appear in parentheses.
[b] All correlations except between Motivation and Exhaustion are significant at $p < .001$. Probability values were calculated for the Pearson's correlations using the Bonferroni adjustment which provides protection for simultaneous tests of multiple correlations.
[c] Correlations between CD-RISC10 scores and those for the GHQ-12 and AHS are intended to assess convergent validity, whereas those between the CD-RISC10 and the SAS4 and the MBI-GS(S) are assessments of divergent validity.

among a sample of auditors in public accounting.[18] These findings motivated us to conduct a more in-depth analysis of this relationship. Specifically, we analyzed a general linear model in which in which stress arousal was set as the dependent variable, and resilience, gender, and major as the predictors. The results presented in Table 5 indicate that resilience is the sole statistically significant predictor of stress arousal. It is also noteworthy that after controlling for the effects of gender and major, the relationship between resilience and stress arousal is still significant at $p < .0001$.

## 6. Discussion

The objectives for this study were two-fold: (1) we sought to validate the CD-RISC 10 for use with accounting and business majors; and, (2) to assess the scale for factor loading invariance across major, academic level, gender and age in a sample from four geographically and mission-diverse universities. The data support the internal consistency and reliability of the items loading on the scale, as well the scale's convergent and divergent validity. Specifically, there were significant positive correlations with the scales predicted to be similar (i.e., the GHQ-12 and AHS scales), and negative associations with those predicted to be dissimilar (i.e., the SAS4 and MBI-GS scales).

This study's factor structure analyses found significantly better fit for a two-factor model similar (though not identical) to that of Aloba et al. (2016) and Smith et al. (2018) and a second-order (i.e., hierarchical) factor model in which the two first-order factors, Motivation and Toughness, are similarly associated with a single higher-order factor, i.e., Resilience. While accounting majors reported significantly lower CD-RISC-10 scores than their other business major counterparts, the second-order factor model was also major invariant with respect to factor structure and factor loadings. In addition, while

[18] The authors used the same instruments to measure resilience and stress arousal that are used in the present study. As did Smith and Emerson (2017), we acknowledge that the relationship between resilience and stress arousal might be non-recursive to some degree. However, by definition, highly resilient individuals should be less influenced by stress than should their less resilient counterparts.

**Table 5**
Multi-way analysis of the influence of resilience, gender, and major on stress arousal.

| Source | Type III SS[1] | df | Mean Squares | F-Ratio | *p*-Value |
|---|---|---|---|---|---|
| Resilience | 384.484 | 1 | 384.484 | 39.884 | 0.000 |
| Gender | 4.074 | 1 | 4.074 | 0.423 | 0.516 |
| Accounting Major | 1.689 | 1 | 1.689 | 0.175 | 0.676 |
| Accounting Major*Resilience | 2.318 | 1 | 2.318 | 0.240 | 0.624 |
| Gender*Resilience | 11.386 | 1 | 11.386 | 1.181 | 0.278 |
| Gender*Accounting Major | 31.037 | 1 | 31.037 | 3.220 | 0.073 |
| Gender*Accounting Major*Resilience | 27.534 | 1 | 27.534 | 2.856 | 0.092 |
| Error | 5118.837 | 531 | 9.640 | | |

[1] Our data were unbalanced, i.e., they had an unequal number of observations in each group (i.e., gender and academic major). Kuznetsova, Brockhoff, and Christensen (2017) state that "In terms of hypothesis tests this type (Type III ANOVA) is the easiest one to interpret both in unbalanced and balanced cases" (p. 17). This statement aligns with Shaw and Mitchell-Olds (1993) who, in citing Speed, Hocking, and Hackney (1978) and Milliken and Johnson (1984), state "When all treatment combinations are observed, but the number of observations per cell varies, the Type 111 method provides the most readily interpretable tests of the null hypotheses of no main effect of Factor A and B" (p. 1643).

women reported significantly lower CD-RISC 10 scores than men, the second-order model was totally gender-invariant (i.e., configural and metric) with respect to factor structure.

It is noteworthy that accounting majors report the lowest levels of resilience compared to other business students. However, as indicated by the two-way analysis of variance reported in Table 2, female accounting majors were the outlier group in the sample with a mean CD-RISC 10 score (26.39) significantly lower than that for the other three groups. This finding motivated the effort reported in Panel C of Table 4 to determine if this mean score difference manifested in non-invariant factor loadings between groups. While the two-factor and second-order factor models generated acceptable and better fit statistics than the one-factor model for both groups as noted in Panel B of Table 4, the factor loadings for items 2 and 3 were non-invariant between groups. Interestingly, both items are designed to measure the Motivation factor. That is, items 2 and 3 influenced female accounting majors at intensities that varied with those reported by male accounting and both male and female non-accounting majors, though the loadings were statistically significant for both groups. Confirmation of this phenomenon with other student samples appears warranted.

A number of potential factors could contribute to our finding of lower reported resilience among female accounting majors. One potential factor is that females with low resilience may self-select into the accounting major. For example, Simons et al. (1995) found that female accounting majors reported the highest levels of oral communication apprehension among business majors in their sample. The authors state that this finding "might be the result of self-selection into a major perceived to have less oral communication demands" (p. 169). Because communication apprehension and resilience are negatively correlated (Glaser, Butler, & Pryor, 1998), it is possible that female business students with low resilience are naturally more inclined to choose accounting as an academic major.

Other potential factors pertain to female perceptions of mathematical self-efficacy. Prior research finds that female college students report lower mathematical self-efficacy than male college students (Betz & Hackett, 1983) and that implicit gender-math stereotypes negatively impact math performance among female college students (Kiefer & Sekaquaptewa, 2007). Given the mathematical demands of the accounting major, female accounting students may perceive a weak internal locus of control with respect to their major course work, which can contribute to lower reported resilience as reported above. Alternatively, low resilience scores may reflect social desirability biases that make individuals more likely to provide responses that conform to social norms of behavior. Given prior research that finds these biases tend to be stronger among women (Chung & Monroe, 2003), it is possible that female accounting majors report lower resilience because gender-math stereotypes are more likely to shape their perceptions on how they handle math-related coursework. These potential explanations notwithstanding, further research as to possible causes for this phenomenon appears warranted.

Another interesting finding is the homogeneity of resilience scores among the student populations of the four universities. The four universities have different missions: a land-grant university in the south, a relatively new comprehensive university in the South, a traditional liberal arts university on the East Coast, and a private institution on the West Coast. It was expected that the student populations would differ somewhat; the finding of consistency was unexpected. The AACSB Accreditation to which all four schools are subject might explain, in part, the score homogeneity.

## 7. Limitations and conclusions

There are notable caveats to this study despite its finding of empirical support for the psychometric properties of the CD-RISC 10 among a population of business school students. As Campbell-Sills et al. (2009) note, the CD-RISC 10, being a self-report measure, "may be influenced by reporting biases (e.g. social desirability) and current emotional state (e.g. individuals in a positive emotional state may overestimate their resilience; the reverse may be true for those in a negative emotional state)" (p. 1011). In the prior section, we discussed the potential for social desirability bias to affect resilience scores, particularly for female accounting majors. Such a bias may hinder the validity of the CD-RISC 10 for affected respondents. In addition, the scale measures a person's perception of their ability to bounce back from adversity, not necessarily that individual's actual ability to do so. Furthermore, respondent answers are potentially vulnerable to exogenous influences. Campbell-Sills

and Stein (2007), in summarizing their evaluation of the scale, note that the CD-RISC 10 captures a characteristic that is able to differentiate individuals who are able to function well after adversity versus those who are not so capable, yet they concede "...other causal relationships could explain the observed pattern of results" (p. 1027). For example, the responses of an individual who is experiencing anxiety symptoms might be negatively influenced by the presence of those symptoms.

Our convergent and divergent (i.e., discriminant) validity assessments were also limited to the heterotrait-homomethod approach. That is, the other measures utilized were also self-report measures. Thus, corroborative evidence to support the validity of this study's findings awaits future efforts that incorporate additional data-gathering methods and different means of assessing resilience. We did not comprehensively evaluate the CD-RISC 10's test-retest reliability due to data availability issues.[19] However, previous studies have reported favorable test-retest reliability, albeit with diverse international subject groups (e.g., see Notario-Pacheco et al., 2011; Wang, Shi, Zhang, & Zhang, 2010).

The above-referenced limitations notwithstanding, the CD-RISC 10 appears to be a practical measure of resilience, as it can be completed and scored in just a few minutes, thus confirming its efficacy in both clinical and research settings. Clinically, student support services professionals at colleges and universities can use it as an initial screening measure to identify students who might benefit from resilience training. For research purposes, the scale's brevity facilitates inclusion of additional measures into studies that might otherwise be excluded due instrument length constraints. For these reasons, we argue for additional consideration of the CD-RISC 10 for inclusion in future studies aimed at examining issues related to stress, performance, retention, and well-being of accounting and business students.

This study's findings also have potential ramifications for women entering the accounting profession. Future research aimed at verifying the reported mean resilience score and factor loading differentials reported by female accounting majors would serve as the first step toward efforts to understand the underlying influences on this phenomenon. In turn, knowledge of these influences could lead to interventions designed to enhance resilience levels among female accounting majors, which have the potential to benefit them in their studies as well as in their careers. Considering that women comprised 52% of Bachelor's and Masters' of accounting enrollees in U.S. colleges and universities in 2015–2016, and 46% of all professional staff of US CPA firms in 2016 (AICPA, 2017), the value of further research in this area should be evident.

This study's finding of a significant negative relationship between resilience and stress arousal suggests that stress management training could be a critical component of resilience enhancement interventions targeted for female accounting majors and other vulnerable student populations. Smith et al. (2014) uncovered a similar relation between resilience and stress arousal, prompting them to suggest that a stress arousal assessment can be used "...as an initial screening measure for distressed students, as well as a means of assessing student progress during counseling" (p. 55). They further document the potential consequences of excessive stress in terms of diminished academic performance among students and the resulting deleterious impact on faculty, school administrators, and future employers. The link between resilience and stress clearly supports an integrated approach to enhancing student well-being.

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## Appendix A. Supplementary material

Supplementary data to this article can be found online at https://doi.org/10.1016/j.jaccedu.2019.01.002.

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[19] Faculty volunteers at each institution agreed to allow students to be administered the survey package during class time. Given the required sacrifice of subject matter coverage due to this protocol, the authors thought it prudent not to request a second in-class solicitation for additional data. However, one of the authors at the East Coast school conducted a two-week test–retest of CD-RISC 10 scores with a sample of 53 accounting majors. The test and retest mean scores were 30.70 and 30.40, respectively, and the correlation coefficient between scores was 0.62. Terwee et al. (2012) indicate that a sample size of 53 falls into the "good" category for assessing test–retest reliability, whereas Tabachnick and Fidell (2007) argue for a total sample size of 10 times the number of scale items for psychometric evaluation of a measure. In the latter case, a minimum sample size of 100 would be required to adequately assess the test–retest reliability of the CD-RISC 10.

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