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The relationship between job-hopping motives and congruence

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Authors:
Amy A. Hall[1] ⓘ
Brandon Morgan[1] ⓘ
Kleinjan Redelinghuys[1] ⓘ

Affiliations:
[1]Department of Industrial Psychology and People Management, College of Business and Economics, University of Johannesburg, Johannesburg, South Africa

Corresponding author:
Brandon Morgan,
bmorgan@uj.ac.za

Orientation: Many organisational turnover and vocational counselling studies have investigated the outcomes associated with person–environment fit. However, few studies have investigated the potential individual difference variables that might lead to person–environment fit.

Research purpose: This study set out to investigate the relationship between job-hopping motives and interest–environment fit (i.e. congruence).

Motivation for the study: Job-hopping motives reflect individual difference motives that are thought to lead to voluntary turnover behaviour over and above environmental factors. We investigated the hypothesis that these motives might lead to people entering jobs that are congruent with their interests.

Research approach/design and method: A cross-sectional survey research design was used. The Job-Hopping Motives Scale was administered to 197 adults. Multiple linear regression and ridge regression were used to investigate the proposed relationships.

Main findings: Job-hopping motives were related to congruence, with the escape motive showing a negative partial linear relationship with congruence and the advance motive showing no partial linear relationship with congruence.

Practical/managerial implications: Our results suggest that people who have a propensity to frequently change jobs might tend to enter a job that does not necessarily match their interests. This in turn might lead to some of the negative outcomes associated with working in so-called incongruent environments, or alternatively, potentially lead to a cycle of entering and leaving jobs throughout a person's career.

Contribution/value-add: We provide evidence that job-hopping motives might be an organisational-relevant individual difference variable that might lead to working in a congruent or incongruent environment.

Keywords: job-hopping motives; congruence; person–environment fit; vocational interests; turnover.

Introduction

Job-hopping, defined as 'the practice of making frequent voluntary job changes' (Lake, Highhouse, & Shrift, 2018, p. 531), has become an increasingly popular topic of discussion in mainstream media and on employment websites in recent years. A Google search[1] of job-hopping returns posts such as 'Job-hopping is on the rise' (Chatzky, 2018), 'Are you job hopping too much?' (Belli, 2019), and 'The truth about job-hopping: It's not as bad as you think' (Smith & Omoth, 2020). Interest in job-hopping does not seem to be limited to popular press. A Google Scholar and ProQuest[2] search of job-hopping returned 68 200 and 3251 results, respectively, with 15 800 (23%) results on Google Scholar and 679 (20%) results on ProQuest since 2017.

Orientation

Reported statistics on job-hopping range widely and often include arbitrary ratios such as up to 75% of employees plan on leaving the organisation within 5 years (Campbell, 2019) and 64% of employees consider changing work roles to be beneficial (Robert Half, 2018). These statistics, although arbitrary, suggest that employees and employers are concerned about the

1. Conducted on the 1st of July 2021.

2. Conducted on the 1st of July 2021. The ProQuest search included 36 databases and the search was limited to peer-reviewed publications.

Note: Additional supporting information may be found in the online version of this article as Online Appendix 1.

consequences and impact of job-hopping behaviour. A quick perusal of popular press and academic articles supports this notion, with these articles often focusing on *who* job-hops, *how many* people job-hop and what the *outcomes* of job-hopping behaviours are for employees and employers. In this study, we consider psychological processes that might increase job-hopping behaviour and investigate the potential consequences of these psychological processes on person–environment fit.

Various causes of job-hopping behaviours have been investigated in the literature. Some of these causes include, for example, an internal orientation, tendency or impulse toward job-hopping, social influences, turnover cultures within organisations, motivational forces, self-regulatory processes, career advancement goals and/or organisational-relevant variables such as poor working environments (Dougherty, Dreher, & Whitely, 1993; Kanfer, Wanberg, & Kantrowitz, 2001; Khatri, Budwhar, & Fern, 1999; Zimmerman, Swider, & Boswell, 2019). Lake et al. (2018) recently proposed job-hopping motives, the focus of our study, as a causal mechanism in job-hopping. Motives have a long tradition in psychology. They are thought to energise, direct and sustain behaviours (Steers & Porter, 1979), and together with other relevant variables such as personality and social context have often been used to explain behaviour (Kanfer, 1990; Kanfer et al., 2001). With respect to job-hopping, Lake et al. (2018) write that job-hopping motives are 'an overarching withdrawal-related motive that would presumably be present regardless of a person's current work situation and may predict multiple instances of quitting' (p. 544). In other words, these motives could potentially explain why people job-hop over and above, or even regardless of, other external job-related variables.

Research purpose and objectives

Frequently changing jobs can have positive and negative outcomes for employees (or people in general). Positive outcomes include, for example, increase in income and better career progression. Negative outcomes include, for example, negative perceptions from employers when submitting job applications (Nizami & Prasad, 2017; Privalko, 2019; Schmelzer, 2012), reduced organisational commitment, and for employers, the potential loss of talented staff (Cruz-Castro & Sanz-Menendez, 2010; Murrell, Frieze, & Olson, 1996). In this study we investigate if having stronger job-hopping motives can reduce person–environment fit with respect to the match between a person's vocational interests and interests required from an environment. This match is usually referred to as congruence rather than person–environment fit in the career literature (Holland, 1997).

Investigating the potential outcomes of job-hopping motives on congruence is important for several reasons. It has long been recognised that interests motivate behaviour and can be mechanisms through which people achieve work-related goals (Silvia, 2009; Strong, 1943; Wiernik, 2016). Holland (1997) believed that working in congruent environments can

have positive outcomes and that by extension, working in incongruent environments – environments that require different interests than those that a person has – can lead to negative outcomes (also see Strong, 1943). These ideas are generally supported with research showing that congruence and incongruence are related to job satisfaction, job performance, general well-being, engagement and, of interest to this study, turnover intentions, amongst others (Meir, 1989; Van Iddekinge, Roth, Putka, & Lanivich, 2011; Wiernik, 2016).

It has long been thought that people who work in incongruent environments can become more congruent over time as their interests change or develop to fit the environment (e.g. Renninger, 2000; Strong, 1943). Recent research, however, has shown that this is not the case, and that interests (or at least trait interests) do not change much in response to the environment (Etzel & Nagy, 2021; Hanna, Briley, Einarsdóttir, Hoff, & Rounds, 2021; Wilk et al. 1995). This suggests that, whilst holding other relevant variables constant, people in incongruent environments are unlikely to become more interested in their job over time in response to the interests inherent in their job, and they may therefore be more susceptible to the negative outcomes associated with incongruence.

We hypothesise that a strong internal motivation to frequently leave a job can lead to a negative cycle of leaving and starting jobs, where frequently changing jobs can increase the chances of working in incongruent environments, which in turn can interact with these motives and other environmentally relevant variables to fuel further job changes. This hypothesis, however, is based on the premise that job-hopping motives are related to congruence. We therefore set out to determine if there is support for this premise as the first step toward investigating this proposed cycle. In the sections that follow, we discuss job-hopping motives in more detail and then expand on John Holland's model of vocational interests and how these interests fit into person–environment fit broadly and congruence specifically.

Literature review
Job-hopping motives

Shrift (2016) and Lake et al. (2018) used an inductive approach to identify two job-hopping motives. The first motive, called the escape motive, is based on organisational turnover literature, and the second motive, called the advance motive, is based on career psychology.

The escape motive

The escape motive is defined by Lake et al. (2018) as 'a job-hopper's desire to frequently change jobs to immediately escape disliked work environments' (p. 532) and shares similarities with Ghiselli's (1974) so-called hobo syndrome. Ghiselli (1974) noted that some people have an illogical and/or disorganised tendency to move from job to job, independent of the type of job, and that this tendency might be related to frequent job change. Judge and Watanabe (1995) found

support for the hobo syndrome and suggested that it might be related to negative affectivity (e.g. neuroticism), personality characteristics and/or environmental variables. Negative affectivity is likely closely tied to the escape motive, because research has shown that those who score higher on neuroticism tend to view themselves and the world more negatively (e.g. Thomson, 2016) and tend to more frequently generate negative thoughts about events (Perkins, Arnone, Smallwood, & Mobbs, 2015). These negative perceptions may in turn increase the probability that a person leaves their job (Woo, 2011; Zimmerman, 2008). That said, it must be kept in mind that motives are generally context-bound, time-bound and role-bound (i.e. state-like) and therefore not identical to personality (McAdams, 1995). Rather, motives and personality together play 'an important but *different* role in regulating behavior and life outcomes' (Winter, John, Stewart, Klohnen, & Duncan, 1998, p. 246, italics in original).

Lake et al. (2018) found that the escape motive was positively related to impulsivity and negative affect and negatively related to persistence, job involvement, time taken to make a turnover decision and career self-efficacy. Woo (2011), in contrast, found that the hobo syndrome (and by extension the escape motive)[3] was not related to impulsivity directly. Rather, it appears that those who are classified as hobos tend to score lower on Maertz and Griffeths's (2004) affective turnover motive or force and higher on the alternative, behavioural, contractual and constituent turnover motives. These aforementioned motives appear to align with the impulsive quitter and conditional quitter profiles derived by Maertz and Campion (2004), which are marked by more negative affect toward the organisation, and that can lead to leaving a job without an alternative offer available (Maertz & Kmitta, 2012).

Hypothesis 1: People who score higher on the escape motive are less likely to work in environments that match their interests.

The advance motive

The advance motive is defined by Lake et al. (2018) as the 'desire to frequently change jobs as a means of career advancement' (p. 532). This type of motive appears to be used most often in popular press when describing the beneficial outcomes of frequently changing jobs. Pelta (n.d.), for example, writes that people who job-hop tend to do so to advance their career, to increase engagement after mastering current work tasks and/or to learn new skills. It has long been assumed that turnover is a product of tangible individual and/or organisational rewards (March & Simon, 1958). Bigliardi, Petroni and Dormio (2005), for example, found that the employees who perceived greater career opportunities in their organisation tended to have fewer turnover intentions and Zhao and Zhou (2008) found that incremental salary increases and job status were negatively related to turnover.

For Lake et al. (2018), the advance motive reflects 'personal drive, initiative and ambition' (p. 532), which appears to suggest that people leave their jobs for the sake of advancement (see Hall, 2002). These ideas can be found in the work of London (1983), who argued that desire for upward mobility, as one component of career motivation, consists of 'need for advancement, recognition, dominance … money … [and the] ability to delay gratification' (p. 621). In their study, Lake et al. (2018) found that the advance motive is positively related to persistence, self-directed career management, protean career values, the boundaryless mindset and a kaleidoscope career; it is negatively related to impulsivity and normative organisational commitment (Lake et al., 2018). From our reading, it appears that this motive is likely related to Maertz and Campion's (2004) comparison and satisficing quitters, who tend to leave their jobs to advance their career and/or to receive better payment, and who usually have another job available before deciding to leave their current job. Thus, we suspect, as was found by Nguyen & Le (2022), that those who score high on the advance motive, in comparison to the escape motive, do not necessarily engage in actual job-hopping behaviours. Rather, it seems that it is the interaction of this motive with relevant need-fulfilling environmental variables that drives turnover behaviour (e.g. Holland, 1997).

Hypothesis 2: People who score higher on the advance motive are more likely to work in environments that match their interests.

Holland's model of vocational interests

Holland (1959) set out to develop a model of vocational behaviour that was comprehensive, able to integrate existing career models and able to produce observable or testable hypotheses. His theory explicitly took into consideration individual (e.g. hereditary) characteristics and environmental (e.g. cultural) forces associated with vocational choices. Holland believed that these forces lead people to prefer certain activities over others, which in turn lead to interest, competencies and a preferred personal disposition for environments where these activities occur (Holland, 1959, 1997). In this regard, Holland (1959) argued that people search for jobs that help them meet their adjustive orientations. In his 1959 article, Holland classified the U.S. occupational environment into six broad categories. These were called motoric, intellectual, esthetic, supportive, persuasive and conforming, but later changed to the more familiar realistic, investigative, artistic, social, enterprising and conventional (RIASEC) interests (Holland, 1962). People are also classified into these six interest types (Holland, 1997).

Holland noticed that the RIASEC interests approximated a circular pattern of relationship in intercorrelation matrices, which led to the development of the RIASEC hexagon used today (Holland, Whitney, Cole, & Richards, 1969). The circular pattern of relationships is more consistent with Guttman's (1954) circumplex model, though, because the hexagon was only an approximation of the theoretical relationships (Holland et al., 1969). The circumplex model

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3. Maertz and Kmitta (2012) write that the hobo syndrome is not necessarily characterised by negative affect and is more akin to impulsive drifting through occupations. In this study we view the hobo syndrome as similar, but not identical, to the escape motive.

has become central to measuring and explaining congruence in career counselling settings (Holland, 1997; Holland et al., 1969). For example, interest profiles are matched to environmental profiles directly or through some mathematical transformation of scores, such as Euclidean distances (Holland, 1997).

Congruence and person–environment fit

Holland believed that people search for environments that match their interests (Holland, 1962; Holland & Nichols, 1964) and that people in these environments generally have better educational and occupational outcomes (Holland, 1997). This matching is referred to as congruence, which Holland (1997) defined as an environment 'in which a person's preferred activities and special competencies are required and his or her personal disposition and its associated characteristics [i.e. adjustive orientations] … are reinforced' (p. 11). Congruence falls within the general person–environment fit model. Person–environment fit is an umbrella term that captures how well individuals match with their environments based on a certain set of characteristics (Kristof-Brown, Zimmerman, & Johnson., 2005). There are many different types of person–environment fit. These include, for example, person–organisation fit, person–leader fit, person–group fit, person–team fit, person–job fit and person–vocation fit, amongst others (e.g. Ghetta, Hirschi, Wang, Rossier, & Herrmann, 2020; Jansen & Kristof-Brown, 2006).

Muchinsky and Monahan (1987) differentiated between two broad types of person–environment fit, namely, supplementary fit and complementary fit. They defined supplementary fit as the fit of individual characteristics with other individuals in the environment and write that '[p]eople wish to join such organizations because they perceive themselves as having similar values, tastes, and interests as existing members in the organization' (p. 270). Holland (1997) believed that people in environments tend to share similar interests (an assumption that has existed since the early 1900s). From this perspective, congruence can be viewed as the fit between a person and others in the organisation (Muchinsky & Monahan, 1987). Complementary fit is similar to puzzle pieces, where people and environments match in terms of their characteristics and the characteristics (e.g. demands or needs) required from the environment (Muchinsky & Monahan, 1987). We believe that interests can also be classified as complementary fit because interests lead people to develop specialised competencies, needs and reinforcers, and they allow people to leverage their strengths and weaknesses to achieve work goals (e.g. Holland, 1997; Wiernik, 2016).

There is much evidence to support the importance of person–environment fit (e.g. Edwards & Parry, 1993; Pervin, 1968). A recent qualitative study on person–environment fit found that participants who did not fit with their environment experienced reduced motivation, self-confidence, mental well-being, work and personal relationships and work performance, and they experienced increased counterproductive work behaviours (Williamson & Perumal, 2021). There is also evidence that person–environment fit increases turnover attentions and actual turnover behaviours (Tak, 2011). As previously noted, the outcomes of person–environment fit with respect to interest–environment fit is well-documented. These outcomes include, for example, better academic performance (Nye, Butt, Bradburn, & Prasad, 2018), increased educational or job performance and persistence (Nye, Su, Rounds, & Drasgow, 2012) and improved job satisfaction (Wiernik, 2016). These results suggest that working in environments that match people's interests is important.

Development and psychometric properties of the Job–Hopping Motives Scale

Lake et al. (2018) developed the Job-Hopping Motives Scale (JHMS) to measure the escape and advance motives. This is a new measure and has received little research attention. We therefore set out to first investigate the psychometric properties of the measure to ensure that the escape and advance motive scores were reliable and valid in this study. Lake et al. (2018) used Hinkin's (1998) inductive approach, which is typically used when there is little theory available to construct a measure (Hinkins, 1995), to develop the scale. Items were written to reflect job-hopping behaviours to escape work and to advance at work. These items were reviewed by two career counsellors and five PhD students and then administered to a sample of 221 undergraduate university students. All the student participants had had at least one job and approximately 141 of them were currently employed at the time of data collection. Lake et al. (2018) then used exploratory factor analysis on the responses from this sample and removed four items that had 'weak factor loadings or cross-loading[s]' (p. 534).

Factor loadings of the eight remaining items ranged from 0.59 to 0.80 on the escape motive factor ($\overline{\lambda_{11-14}} = 0.68$) and 0.40–0.65 on the advance motive factor ($\overline{\lambda_{15-18}} = 0.58$). It seems that Lake et al. (2018) used a factor loading of |0.35| as a meaningful or salient loading. No items showed cross-loadings based on this rule-of-thumb criterion. However, if |0.30|[4] was used, four items showed cross-loadings and one item showed a potential cross-loading. The off-target loadings ranged from 0.04 to 0.32 for the escape motive factor ($\overline{\lambda_{11-14}} = 0.18$) and 0.09 to 0.31 for the advance motive factor ($\overline{\lambda_{15-18}} = 0.25$). Confidence intervals (or *p* values) were not reported for these factor loadings, making it difficult to determine if these cross-loadings were statistically significant. The cross-loadings, when using |0.30| as a rule-of-thumb, together with the interfactor correlation coefficient of 0.40, suggested that there might be one rather than two distinct factors. Lake et al. (2018) investigated this possibility by modelling one-factor and two-factor models across four other sample groups using confirmatory factor analysis. The differences in fit for all sample groups were statistically

4. We are not suggesting that |0.30| should be used. It is mentioned here as a point of comparison because |0.30| is often used to indicate a meaningful or salient factor loading.

significant, suggesting that the two-factor model had better overall fit. The interfactor correlation coefficients of the two-factor model across the four sample groups ranged from 0.32 to 0.56 with a sample-size-weighted average of 0.40.

Cronbach alpha reliability coefficients ranged from 0.74 to 0.78 (sample-size-weighted $\overline{\alpha}$ = 0.76) for the escape motive scale and from 0.75 to 0.80 (sample-size-weighted $\overline{\alpha}$ = 0.78) for the advance motive scale. Nguyen & Le (2022) found alpha reliability coefficients of 0.72 and 0.77, respectively, although it appears that only three items from each scale were used. These aforementioned alpha coefficients are rather low but expected, given that each scale consists of only four items. Shorter scales are less burdensome for participants but come at the expense of lower reliability and reduced construct coverage (e.g. Widaman, Little, Preacher, & Sawalani, 2011). We used Feldt, Woodruff and Salih's (1987) formula to calculate if the reported alpha coefficients for the four sample groups from Lake et al. (2018) had a statistically significant difference between them and found that there were no statistically significant differences in the Cronbach alpha coefficients for the two scale scores across the four sample groups.

Summary

We have argued that job-hopping motives, or at least the escape motive, might lead to a cycle of leaving and starting jobs, where frequently changing jobs can increase the chance of working in incongruent environments, which in turn can interact with the motives to fuel further job changes. As a precursor to investigating this possibility, we first set out to investigate the hypothesis that people who score high on job-hopping motives might end up in jobs that do not match their interests.

Method

Research approach

We used a cross-sectional quantitative research design.

Research participants

Participants were selected from various organisations using nonprobability sampling throughout 2019. They were required to be working adults who were working at the time of data collection. We originally had 197 responses to the questionnaires. However, nine participants were removed because we were uncertain about the correct RIASEC environmental code for their respective occupations (this coding is explained later). Two participants were removed that appeared to have no work experience, and a further three participants were removed who appeared to have unusual response patterns based on the number of Guttman errors in their responses (Emons, 2008). This left 183 participants in the dataset.

The mean age of the participants was 35.68 (median = 32, minimum = 21, maximum = 68) and consisted of 78 men (42.62%) and 105 women (57.38%). Most of the participants identified as white (n = 85, 46.70%) followed by black or African (n = 51, 28.02%), Indian and/or Asian (n = 33, 18.13%) and mixed race (n = 13, 7.14%). The most spoken home language in the sample was English (n = 89, 48.90%), followed by Afrikaans (n = 42, 23.08%) and isiZulu (n = 10, 5.49%). The participants generally held tertiary education qualifications with 107 (59.12%) indicating that they had a degree, 44 (24.31%) indicating that they had a diploma or certificate and 28 (15.47%) indicating that they had a Grade 12 certificate. The participants had on average worked for six complete years in their current occupation (median = 3, minimum = 0, maximum = 30) and had on average three (median = 3, minimum = 0, maximum = 10) different jobs throughout their careers. For the most part, the participants had full-time employment (n = 153, 85%), with 24 (13.33%) indicating that they had part-time work. Participants worked in various different industries, including but not limited to education, human resources, engineering, finances, information technology and social services.

Measuring instruments

Participants completed the JHMS, the Maslach Burnout Inventory – General Survey (which was required for a different study) and the South African Career Interest Inventory – Short (SACII-SR). The JHMS scale was previously discussed and is therefore not presented here again. The SACII-SR (Morgan & De Bruin, 2019) is a 30-item measure of Holland's six vocational personality types. It has five items per scale and participants respond using a five-point Likert-type scale ranging from strongly disagree to strongly agree. Morgan and De Bruin (2019) found reliability coefficients ranging from 0.80 to 0.89 in their first sample and 0.72 to 0.83 in their second sample for the SACII-SR scale scores. Satisfactory fit to circumplex structure and the correct RIASEC ordering was also found in their study.

Research procedure

Questionnaires for this study and another study were made available in hard-copy format and online using Google Forms. Working adults in the social network of the researchers were approached and invited to complete either the hard-copy or online versions of the questionnaire. LinkedIn and Facebook were used as one approach to invite working adults in their social network to participate. Human Resource managers and/or company owners were also contacted to obtain permission to invite employees in their respective companies to participate. Employees in these companies were sent an invitation to participate via email by the respective manager or owner. Employees who agreed to participate completed the hard-copy questionnaires or the online version of the questionnaires, depending on their preference.

Analysis

Congruence was calculated in two ways: (1) profile correlation coefficients (see Xu & Li, 2020) and (2) the summed distance

between the highest and lowest participant RIASEC score with the highest and lowest environment RIASEC score (the hexagon index).[5] Two approaches were used because there is no agreed-upon measure of congruence and because convergence of results across congruence measures helps reduce potential spurious results because of the type of congruence score calculated. The hexagon index depends on the RIASEC circumplex model holding in this sample. We used the randomisation test of hypothesised order relations (Hubert & Arabie, 1987) and multidimensional scaling to investigate this structure and found the correct RIASEC ordering and satisfactory fit to the model predicted order relations (correspondence index = 0.83, p = 0.015). More details on these two analyses can be obtained from the Online Supplement (https://osf.io/254tz/?view_only=760c5d4866 1a406d9d3db89b8a471dc4).

Environment scores were obtained using O*NET (https://www.onetonline.org/). O*NET contains RIASEC scores for a range of different jobs and is often used to obtain environment scores in congruence research (e.g. Nye et al., 2018). To help us determine each participant's occupation, we asked participants to write down their job title and the main work tasks that they are most often involved in. The first and second author, as well as another master's student, coded the occupations separately. We then met to discuss our coding and attend to discrepancies in occupational codes until agreement was found.

The structural validity of the JHMS was investigated using exploratory factor analysis with ordinary least square extraction and geomin rotation. The *EFAutilities* package version 2.11 (Zhang Jiang, Hattori, & Trichtinger, 2020) in R version 4.0.5 (R Core Team, 2021) was used for this purpose. Jackknifed standard errors were calculated to obtain 95% confidence intervals for the rotated factor loadings and to calculate their statistical significance (Zhang, 2014). The relationship between the job-hopping motives and congruence was investigated using linear regression and ridge regression. The *ridge* package version 2.9 (Cule, Moritz, & Frankowski, 2021) was used for the ridge regression. In Online Appendix 1, we indicate the names of other packages that were used during the analysis but not cited here.

Ethical considerations

Ethical clearance to conduct this study was obtained from the University of Johannesburg, Department of Industrial Psychology and People Management Ethics Committee (reference number: IPPM–2019–360M). Participants received a participant information form and consent form. Participation was voluntary and the data are kept confidential.

Results

Reliability coefficients and structural validity of the South African Career Interest Inventory scale scores

Coefficient omega total reliability coefficients with 95% normal theory confidence intervals for the RIASEC scale scores were as follows: realistic = 0.94 (0.92, 0.96), investigative = 0.89 (0.86, 0.92), artistic = 0.87 (0.84, 0.90), social = 0.85 (0.81, 0.89), enterprising = 0.81 (0.76, 0.86) and conventional = 0.85 (0.81, 0.89). The escape (ES) and advance (AD) motives had coefficient omega total reliability coefficients of 0.76 (0.70, 0.82) and 0.80 (0.75, 0.85), respectively.

Structural validity of the Job-Hopping Motives Scale

The Kaiser-Meyer-Olkin Measure of Sampling Adequacy for the eight JHMS items was 0.86 and ranged from 0.82 (AD7)[6] to 0.90 (ES4). We used parallel analysis of components, sequential χ^2 model testing and the *xgboost* model (Goretzko & Bühner, 2020) to determine the number of major dimensions in the data. Parallel analysis suggested there is one component in the data, whereas sequential χ^2 models and the *xgboost* model suggested that there are two factors in the data, with the *xgboost* model showing a probability of approximately 94% for two factors (compared to approximately 4% for one factor). We proceeded by fitting a one-factor and two-factor model[7] and calculating the number of items with a mean absolute residual ≥ 0.05 (excluding the matrix diagonal). Six items in the one-factor model[8] had mean absolute residuals ≥ 0.05. This suggests that there is some unmodelled variance when one factor is extracted. No items had mean absolute residuals ≥ 0.05 in the two-factor model.[9] However, the inter-factor correlation coefficient was 0.58 (0.40, 0.72), suggesting that the two factors were spaced relatively closely together in two-dimensional space (Figure 1).

In Table 1, we provide the geomin-rotated pattern matrix. Inspection of this table showed that items ES1 through to ES4 had statistically significant factor loadings on the second factor and that items AD5 through to AD8 had statistically significant factor loadings on the first factor. Items ES2 and AD7 had a statistically significant factor loading on both factors (i.e. salient cross-loadings). However, the cross-loadings were only marginally statistically (non-) significant (p = 0.03; p = 0.06). No items showed statistically significant cross-loadings (standard errors for the factor loadings are presented in Online Appendix 1) after applying a Holm-Bonferroni correction.

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5. Wiernik (2016) found that the strongest and weakest interests in a RIASEC profile were the most important predictors of outcomes. We therefore calculated a congruence index using these two interests in comparison to the highest and lowest interest scores in the occupational RIASEC profile. The hexagon interpoint distances were used to calculate this hexagon index where match = 3, adjacent = 2, alternate = 1 and opposite = 0.

6. We use ES to refer to the escape items and AD to refer to the advance items.

7. Geomin rotation results can sometimes reflect a local maximum/minimum rather than the global maximum/minimum (Hattori et al., 2017). We therefore ran 1000 two-factor models with random starting values for each model. No better alternative estimates were obtained.

8. χ^2(20) = 44.57, p < 0.001, RMSEA = 0.08 (0.05, 0.11), ECVI = 0.42 (0.29, 0.50), mean absolute residual = 0.07.

9. χ^2(13) = 12.67, p < 0.001, RMSEA = 0.00 (0.00, 0.07), ECVI = 0.32 (0.28, 0.35), mean absolute residual = 0.02.

The on-target factor loadings ranged from 0.49 to 0.71 ($\overline{\lambda}_{11-14} = 0.63$) for the escape motive and 0.42–0.90 for the advance motive ($\overline{\lambda}_{15-18} = 0.64$). The off-target factor loadings ranged from −0.13 to 0.24 ($\overline{\lambda}_{11-14} = 0.13$) for the escape motive and 0.07 to 0.17 ($\overline{\lambda}_{15-18} = 0.13$) for the advance motive. In Figure 1, we graphically represent the factor loadings in Table 1 with unit length (i.e. communality coefficients set to unity) in two-dimensional space. The grey polygons represent the lowest and highest confidence interval of the four items in each factor.
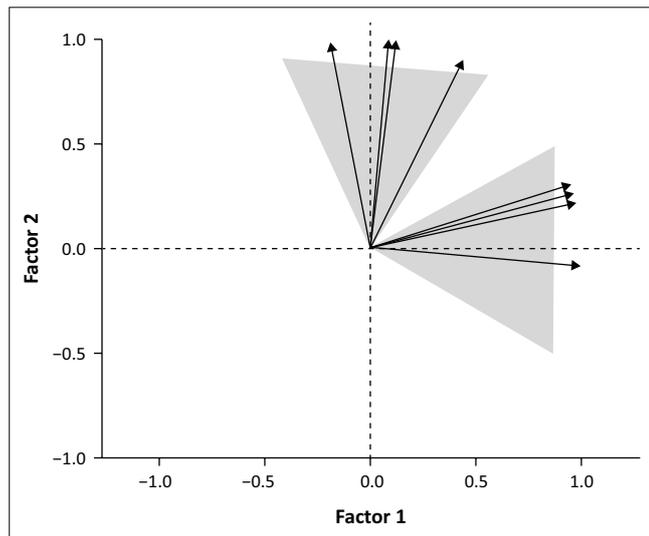
Note: Item labels from the top left quadrant to the bottom right quadrant are Escape (ES)2, ES1, ES3, ES4, Advanced (AD)8, AD6, AD5, AD7. Communality coefficients were set to unity.

FIGURE 1: Two-dimensional representation of factor loadings for the Job-Hopping Motives Scale.

TABLE 1: Geomin-rotated pattern matrix for the two-factor Job-Hopping Motives Scale model.

Items	Factor 1		Factor 2		h^2
	95%	Jackknifed confidence interval	95%	Jackknifed confidence interval	
ES1	0.06	−0.13, 0.25	**0.67**	**0.46, 0.87***	0.49
ES2	**−0.13**	**−0.25, −0.01**	**0.71**	**0.54, 0.89***	0.42
ES3	0.08	−0.17, 00.33	**0.65**	**0.43, 0.87***	0.49
ES4	0.24	−0.03, 0.50	**0.49**	**0.23, 0.75***	0.43
AD5	**0.63**	**0.32, 0.93***	0.16	−0.19, 0.50	0.53
AD6	**0.62**	**0.36, 0.87***	0.17	−0.14, 0.49	0.53
AD7	**0.90**	**0.081, 1.00***	**−0.07**	**−0.15, 0.00**	0.74
AD8	**0.42**	**0.19, 0.65***	0.13	−0.08, 0.35	0.26

ES, escape; AD, advance; h^2, communality coefficient.
Factor loadings in bold are statistically significant at $p < 0.05$.
*, Factor loadings statistically significant after applying a Holm-Bonferroni correction.

Descriptive statistics and correlation coefficients

In Table 2, we present descriptive statistics for the standardised age-weighted number of jobs participants had worked in (i.e., the standardised score of number of jobs divided by age), the escape and advance motives and the profile correlation and hexagon congruence coefficients. Profile correlation coefficients were transformed to Z scores using Fisher's r to Z formula, as this transformation produces scores that better approach a normal distribution. The participants in our sample generally scored higher on the advance motive than on the escape motive.

In Table 3, we present the Pearson correlation matrix for the variables presented above. We first inspected if the relationships between the variables were linear using Ramsey's RESET (Ramsey, 1969) and bivariate scatterplots. No serious concerns were noted. Correlation coefficients in bold in Table 3 are statistically significant at $p < 0.05$ and correlation coefficients with an asterisk are statistically significant after applying a Holm-Bonferroni correction. Inspection of the table shows that participants who scored higher on the escape ($r = 0.25$, $r_{sp} = 0.10$) and advance ($r = 0.28$, $r_{sp} = 0.17$) motive tended to have more jobs throughout their working career and to be less congruent with respect to the profile correlation congruence coefficient (r escape = −0.34, r_{sp} escape = −0.22, r advance = −0.27, r_{sp} advance = −0.09) and hexagon congruence coefficient (r escape = −0.27, r_{sp} escape = −0.16, r advance = −0.24, r_{sp} advance = −0.11). Those participants who had more jobs also tended to be less congruent (r profile correlation = −0.19, r hexagon congruence = −0.12).

Multiple regression

The escape and advance motive are relatively highly correlated ($r = 0.57$), suggesting that multicollinearity might be a concern in the regression model. We calculated the condition index of the scaled design matrix and the variance inflation factors of the two motives to determine if multicollinearity had a large effect on the partial regression coefficient standard errors. We then applied ridge regression to determine if the substantive conclusions of the linear

TABLE 2: Descriptive statistics for weighted jobs, job-hopping motives and congruence coefficients.

Variable	Mean	Median	Standard deviation	Minimum	Maximum	Skewness		Excess kurtosis		Standard error of the mean
						95%	Bias corrected confidence interval	95%	Accelerated confidence interval	
Jobs	0.00	−0.22	1.00	−1.55	3.43	0.70	0.41, 1.06	0.25	−0.42, 1.55	0.07
Escape	2.68	2.50	0.91	1.00	4.75	0.23	0.02, 0.44	−0.54	−0.84, −0.18	0.07
Advance	3.55	3.75	0.92	1.00	5.00	−0.61	−0.88, −0.36	−0.04	−0.55, 0.65	0.07
Escape_R	0.00	−0.06	0.74	−2.11	1.71	−0.13	−0.48, 0.17	0.10	−0.39, 0.79	0.05
Advance_R	0.00	0.06	0.75	−1.76	2.03	−0.09	−0.39, 0.23	−0.02	−0.45, 0.53	0.06
PC	0.41	0.37	0.67	−1.15	2.62	0.32	−0.00, 0.73	0.15	−0.45, 1.28	0.05
HC	5.38	5.00	1.53	2.00	8.00	−0.19	−0.41, 0.03	−0.83	−1.05, −0.51	0.11

Jobs, weighted jobs; Escape_R, residualised escape motive; Advance_R, residualised advance motive; PC, profile correlation; HC, hexagon congruence.
Note: 5000 resamples used.

TABLE 3: Pearson correlation coefficient matrix for job-hopping motives, weighted number of jobs and congruence coefficients.

Variable	Escape	Advance	Weighted jobs	Profile correlation	Hexagon
Escape	-	-	-	-	-
Advance	**0.57*** [0.47, 0.66]	-	-	-	-
Weighted jobs	**0.25*** [0.11, 0.38]	**0.28*** [0.14, 0.41]	-	-	-
Profile correlation	**−0.34*** [−0.46, −0.20]	**−0.27*** [−0.40, −0.13]	**−0.19** [−0.33, −0.05]	-	-
Hexagon	**−0.27*** [−0.40, −0.13]	**−0.24** [−0.37, −0.10]	−0.12 [−0.26, 0.03]	**0.65*** [0.56, 0.73]	-
Residual escape	-	-	0.10 [−0.04, 0.25]	**−0.22*** [−0.36, −0.08]	**−.16** [−.30, −.02]
Residual advance	-	-	**0.17** [0.03, 0.31]	−0.09 [−0.24, 0.05]	−.11 [−.25, 0.04]

Residual escape and residual advance are the residual scores after removing the respective variance attributed to the other motive. 95% confidence intervals based on normal theory. Statistically significant correlation coefficients at $p < 0.05$ in bold.

*, Statistically significant correlation coefficients a*re applying a Holm-Bonferroni correction.*

TABLE 4: Regression of profile correlation on escape and advance motives.

Variable	B	SE	t	p	β	B_{Ridge}	SE_{Ridge}	t_{Ridge}	P_{Ridge}	Rel.Import.
Con.	1.24 [0.86, 1.62]	0.19	6.49	< 0.001	-	-	-	-	-	-
ESC	−0.20 [−0.32, −0.08]	0.06	−3.17	0.002	−0.27 [−0.43, −0.11]	−2.17	0.63	3.44	< 0.001	0.08 [0.03, 0.16]
ADV	−0.08 [−0.21, 004]	0.06	−1.35	0.180	−0.11 [−0.29, 0.06]	−1.06	0.63	1.69	0.092	0.04 [0.01, 0.11]

B, unstandardised beta coefficient; SE, standard error; β, standardised beta coefficient; B_{Ridge}, scaled beta ridge coefficient; Rel.Import, average contribution to R^2 across orderings of regressors. 95% confidence intervals based on normal theory for partial regression coefficients in parentheses. 95% percentile bootstrapped confidence intervals for R^2, adjusted R^2 and relative importance in parentheses. Ridge parameter = 0.113. $R^2 = 0.12$ [0.05, 0.24], adjusted $R^2 = 0.11$ [0.03, 0.23] and $F(2, 180) = 12.46$.
$p < 0.001$.

TABLE 5: Regression of hexagon congruence coefficient on escape and advance motives.

Variable	B	SE	t	p	β	B_{Ridge}	SE_{Ridge}	t_{Ridge}	P_{Ridge}	Rel.Import.
Con.	7.03 [6.15, 7.91]	0.45	15.77	< 0.001	-	-	-	-	-	-
ESC	−0.33 [−0.62, −0.04]	0.15	−2.26	0.025	−0.20 [−0.37, −0.02]	−3.74	1.51	2.47	0.014	0.05 [0.09, 0.12]
ADV	−0.21 [−0.50, 0.07]	0.14	−1.48	0.140	−0.13 [−0.13, 0.03]	−2.59	1.51	1.71	0.087	0.03 [0.01, 0.09]

B, unstandardised beta coefficient; SE, standard error; β, standardised beta coefficient B_{Ridge}, scaled beta ridge coefficient; Rel.Import, average contribution to R^2 across orderings of regressors. 95% confidence intervals based on normal theory for partial regression coefficients in parentheses. 95% percentile bootstrapped confidence intervals for R^2, adjusted R^2 and relative importance in parentheses. Ridge parameter = 0.095. $R^2 = 0.08$ [0.03, 0.18], Adjusted $R^2 = 0.07$ [0.02, 0.17] $F(2, 180) = 8.29$.
$p < 0.001$.

model remained the same. The condition index was 10.22 for the two motives with eigenvalues of the column unit length matrix of 2.92, 0.05, 0.03. The variance inflation factors were both 1.49. Together, these results suggest that multicollinearity might not be too problematic but should also not be ignored. For both regression models, the residuals were approximately normally distributed and had constant variance (Breusch-Pagan $\chi^2[1] = 0.2.13$, $p = 0.14$, Breusch-Pagan $\chi^2[1] = 0.03$, $p = 0.86$), and the partial-residual plots supported linear relationships between the variables.

In Table 4, the regression of the profile correlation congruence coefficient on the escape and advance motive is presented. The two motives explained approximately 12% of the profile correlation coefficient congruence index variance. The escape motive had a statistically significant negative partial relationship with the profile correlation coefficient congruence index ($B = −0.20$, $p = 0.002$), suggesting that those who score higher on the escape motive are more likely to work in environments that do not fit their interests. The same relationship was found with the hexagon correlation coefficient ($B = −0.33$, $p = 0.025$). The advance motive did not show a statistically significant partial regression coefficient with the profile correlation congruence index ($B = −0.08$, $p = 0.180$) or with the hexagon congruence coefficient ($B = −0.21$, $p = 0.140$). Ridge regression produced the same substantive conclusions ($B_{Ridge} = −1.06$, $p = 0.092$; $B_{Ridge} = −2.59$, $p = 0.087$), suggesting that the nonstatistically significant results could not be attributed solely to potential multicollinearity.

Discussion

This study set out to investigate if job-hopping motives were related to congruence as one form of person–environment fit. We first investigated the psychometric properties of the JHMS. Thereafter we examined two hypotheses: (H1) people who score higher on the escape motive are less likely to work in environments that match their interests, and (H2) people who score higher on the advance motive are more likely to work in environments that match their interests.

Psychometric properties

The psychometric properties of the JHMS have not received much attention. Lake et al. (2018) conceptualised the scale as consisting of two dimensions. Their initial study found some evidence that the two scales are measuring different job-hopping motives. However, the interfactor correlation coefficients were relatively large, and there were cross-loadings at the item level. Our results support the two-factor structure of the JHMS, with the items having salient loadings on their intended factors. The interfactor correlation coefficient in our sample was also quite large and was similar to the interfactor correlation coefficient reported by Nguyen & Le (2022). The item level cross-loadings in our study were somewhat smaller than those reported by Lake et al. (2018), with an absolute mean average of 0.13 for both the escape and advance motive, compared to 0.18 and 0.25 found by Lake et al. (2018).

The reason for the relatively large interfactor correlation coefficient is not clear. It is possible that this correlation coefficient is because of a theoretical overlap of the two motives and/or that it is an instrument-specific effect. With respect to the former, theoretical overlap could imply that these two motives are facets of a broader general job-hopping motive. The results presented by Lake et al. (2018) seem to counter this view, because it appears that the two motives do have distinct relationships with criterion variables. With respect to the latter, it is possible that the existing items share too much conceptual similarity. Item ES4, for example, showed a fairly large cross-loading in both our results and Lake et al.'s results (2018). Inspection of the item content also shows that the wording can easily apply to both motives. Regardless, we had evidence for reliability and validity of the job-hopping motive scale scores with the sample group in our study.

Relationship between job-hopping motives and congruence

We first investigated if the two job-hopping motives were related to the number of jobs that participants held. Pearson correlation coefficients showed that the two motives were positively related to the weighted number of jobs held by each participant. Lake et al. (2018), in contrast, found that only the advance motive was related to the log of lifetime number of jobs. Our calculation of number of jobs was different to that of Lake et al. (2018), making direct comparison of results difficult. The semipartial correlation coefficients, however, showed that only the advance motive was related to the weighted number of jobs, supporting Lake et al.'s (2018) findings. We do not want to overinterpret these results because of the limitations in how the variable was calculated. Specifically, we did not ask how many changes in jobs were voluntary or not and/or because of contract or permanent working relationships. It therefore makes it difficult to make any definite conclusions.

The Pearson correlation coefficients showed that the two motives were negatively related to both the profile correlation congruence index and the hexagon congruence index. The semipartial correlation coefficients, however, showed that only the escape motive had statistically significant relationships with these two indices. It is problematic to rely solely on correlation coefficients when there is substantial overlap with other variables (Wiernik, Wilmot, & Kostal, 2015). We therefore used regression to investigate partial relationships. The results showed that the escape motive had a negative statistically significant partial relationship with congruence but that the partial relationship between the advance motive and congruence was not statistically significant. Hypothesis 1 is therefore supported and hypothesis 2 not supported in our study. It is difficult to make a definite conclusion on these two hypotheses from one study on one small sample group. Inspection of the 95% confidence intervals for the advance motive showed that the unstandardised beta coefficients were only marginally overlapping with zero. It is therefore entirely possible that

future studies find that the advance motive is also negatively related to congruence.

We can only postulate on the reason for these relationships in this study. It is possible that those who score higher on the escape motive are more inclined to leave perceived and/or actual undesirable working conditions and that this motive may lead to entering alternative employment without consideration of their interests and the interests required by the environment. Follmer, Talbot, Kristof-Brown, Astrove and Billsberry (2018) found that some people do quickly leave their jobs when they perceive negative fit, although leaving was generally more likely when alternative options were available. On the other hand, people prone to job-hopping under the advance motive may tend to take a more careful approach when leaving their job (Lake et al., 2018). This means that they might take into account whether or not they are interested in the new job before deciding to make the change, especially when advance is meant to further goals. This is somewhat in line with the argument made by Steers and Porter (1979) that people who are less interested in their work will respond more readily to external stimuli when considering alternative employment.

Implications for theory and practice

Our study tentatively supports the psychometric properties of the JHMS, albeit with some caveats. Firstly, the reason for the large interfactor correlation coefficient should be further investigated. That is, clarity should be sought on whether it is a theoretical overlap, a limitation of the instrument itself or a combination of both. One way to address this problem would be to reconsider item wordings and/or add additional items to each scale. That said, the use of the JHMS for research purposes at this stage does seem to be somewhat justified, although more evidence is needed before making any definite claims in this regard.

With respect to the investigated relationships, the results suggest that job-hopping motives might be related to entering jobs that do not necessarily match a person's interests, at least within the boundary conditions of our sample group. This might hold implications for career counsellors, particularly when clients have shown a pattern of frequent job changes (see Hall, 2002). With respect to the escape motive, it might help to initiate a discussion of clients' interests and how it matches, or does not match, with their current job. Those who score high on the advance motive in turn might benefit from a discussion about their career goals and objectives and how these fit into potential job changes and their interests. Future research should investigate if these suggestions are viable.

Limitations and recommendations

There are many limitations in this study that should lead to caution in the interpretation of the results. The first limitation is that the sample size is small and that the sample is not representative of the population. Small sample sizes have statistical implications, such as imprecision in parameter

estimates and greater chances of capitalising on sample-specific peculiarities. The sample composition makes it difficult to determine if the results will generalise to other working adults. For example, our sample size had a fairly large number of highly educated participants. It is possible that those with higher education are more easily able, or more willing, to leave their jobs, regardless of their standing on the two-job hopping motives. The sample group was also mostly clustered around business-type jobs. This makes it difficult to determine if the results are more representative of motives for people who work in enterprising and conventional-type environments or if it would apply to all environments.

A second limitation is that we relied on imperfect conceptualisations of congruence. It is known that reducing complex multivariate relationships to a single univariate fit score can reduce the validity of these measures (e.g. Edwards, Caplan, & Van Harrison, 1998; Wiernik, 2016). Polynomial regression and response surface methodology can overcome these limitations, but in our sample, we did not have commensurate person and environment scores, making application of this approach problematic (Edwards et al., 1998). In addition to this, the small sample size would make statistical estimation of polynomial terms problematic. The environment scores that we used are also based on the O*Net classification. Although, as we previously mentioned, this is common practice, it does introduce some problems. For example, some jobs do not completely match with the available O*Net occupations, meaning that a best alternative had to be chosen. Secondly, the O*Net classification was done by researchers in the United States of America for jobs as they apply to the U.S. job market. It is unclear if the same classifications would apply to jobs in the South African context.

A third limitation is that job complexity was not factored into the analysis. Not all jobs require the same level of interest from employees (Holland, 1997). This means that not all congruence scores are equal. For some occupations, congruence might matter a lot, whereas for other occupations it might matter little. It is possible that complexity could moderate these relationships. We have also not controlled for other potentially confounding variables in the regression model, such as education level, age, gender and so on. Each of these variables might also moderate relationships investigated in this study. We unfortunately did not have sufficient data to consider these moderation effects. We also did not want to conduct post-hoc moderation analyses because this can easily lead to capitalisation on chance relationships.

Conclusion

This study set out to investigate the psychometric properties of the JHMS and to determine if job-hopping motives are related to congruence. The results tentatively support the psychometric properties of the JHMS and suggest that job-hopping motives are negatively related to congruence; however, given the previously reported limitations, we encourage readers to not overinterpret and/or overgeneralise these findings until more comprehensive studies are done.

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