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ORIGINAL ARTICLE

Dimensional structure of the demand control support questionnaire: a Brazilian context

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Abstract According to Karasek, job strain results from an interaction between high demands and low decision latitude.

Purpose To reassess the dimensional structure and evaluate the internal consistency of demand control support questionnaire (DCSQ), a shortened version of job content questionnaire that was not sufficiently evaluated in validation studies.

Methods The study investigated 825 workers who completed the DCSQ in Rio de Janeiro, Brazil; to 399 workers, the questionnaire was self-administered at a hospital (2004–2005), and 426 workers were interviewed at nine restaurants (2006–2007). Confirmatory factor analysis using structural equation models was used to test theoretical structure of dimensionality. Internal consistency was evaluated by composite reliability and convergent validity by average variance extracted.

Results Confirmatory factor analysis supported the instrument in three dimensions: demands, skill discretion and decision authority. The best fit model was achieved by removing social support at work and the item *repetitive work* (skill discretion). A cross-loading from *learning new things* on demands and an error measurement correlation between *work fast* and *work intense* were confirmed. Composite reliability was acceptable for all dimensions, except for demands (0.58), which also showed inadequate average variance extracted (0.32). This final model was confirmed in separate analyses according to work setting, but the loadings of demands were lower for restaurant workers.

Conclusion Our results indicated that skill discretion and decision authority formed two distinct dimensions. Additionally, the item *repetitive work* should be removed, as well as one of the items *work fast* or *work intense* (demands). Future research is still required to confirm these findings.

Keywords Validation studies · Factor analysis · Stress, psychological · Demand control model

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Introduction

Proposed by Karasek (1979), the demand control model considers that job strain is a consequence of an interaction between high psychological demands and low job decision latitude, a combined measure of skill discretion and decision authority. The job strain model has been examined in relation to a variety of health outcomes, especially cardiovascular disease (Karasek et al. 1981; Karasek and

Theorell 1990; Schnall et al. 1994; Van der Doef and Maes 1999; Ibrahim et al. 2001). Social support at work was added later to the model as a dimension that may buffer the effect of high strain on stress-related illness (Johnson and Hall 1988).

Several instruments have been used to measure job strain (Landsbergis et al. 2000; Tabanelli et al. 2008), most of which were originally derived from secondary analyses of United States and Swedish surveys. The job content questionnaire (JCQ) consists of 49 items distributed in six dimensions including psychological demands, decision latitude, social support at work and other work characteristics such as physical demands, macro-level decision authority and job insecurity (Landsbergis et al. 2000). Several validation studies have been carried out for JCQ in different subgroups of worker populations in Europe (Escriba-Aguir et al. 2001; Niedhammer 2002; Ferrario et al. 2005; Niedhammer et al. 2006), North America (Brisson and Larocque 2001; Sale and Kerr 2002), Asia (Kawakami and Fujigaki 1996; Cheng et al. 2003; Li et al. 2004; Edimansyah et al. 2006; Eum et al. 2007) and South America (De Araújo and Karasek 2008).

In 1988, Theorell proposed a short version of the JCQ, the Swedish demand control support questionnaire (DCSQ), mostly used in Scandinavian countries, encompassing 17 items distributed in three dimensions: psychological demands, decision latitude and social support at work (Theorell et al. 1988). The four response categories of demands and decision latitude are frequency based (often to never), while in JCQ, they range from totally agree to totally disagree. DCSQ social support items are more oriented toward worksite environment, while in JCQ, they focus on supervisor and coworker support.

Only two validation studies of DCSQ could be identified (Theorell 1996; Sanne et al. 2005), one of them restricted to internal consistency analysis (Theorell 1996). Exploratory factor analysis performed on a representative sample of Norwegian workers confirmed the three-factor solution described earlier (Sanne et al. 2005). However, for men in the four high-status occupations, 0–4 major occupational groups of International Standard Classification of Occupations (ISCO 1988), and for women in the low-status occupations (5–9 ISCO major groups), the decision latitude dimension fitted better when subdivided into skill discretion and decision authority in a four-factor solution.

Since DCSQ originated from JCQ, a common dimensional structure would be expected. Despite differences in length and response categories, Karasek et al. (2007) observed a moderate agreement between those instruments for the high job strain group, when different definitions were applied.

Despite changes occurring worldwide in work environment and employment relationships, the demand control model and related questionnaires (Ferrie et al. 2008), continue to show explanatory power in relation to job stress-related outcomes (Choi et al. 2008; Johnson 2008). However, this theoretical model has not been widely investigated in emerging countries that still have a high percentage of low educated workers, which could interfere with the functional equivalence of instruments across different cultures (Herdman et al. 1998) and may represent a source of heterogeneity in association studies (Stewart and Napoles-Springer 2000; Reichenheim and Moraes 2007).

A cross-cultural adaptation process of the DCSQ for use in the Brazilian context evaluated conceptual, semantic, item, operational, as well as measurement equivalences (Herdman et al. 1998; Alves et al. 2004). However, the latter assessment was restricted to test–retest reliability and internal consistency analysis, indicating that further scrutiny would still be needed (Alves et al. 2004).

Confirmatory factor analysis (CFA) was not performed in none of these DCSQ validation studies. CFA is a multivariate statistical technique that has been increasingly used to test specific prior hypotheses about the structure underlying constructs and compare alternative measurement models with respect to explanatory power. It is also useful to investigate relationships not previously identified in the model, such as error measurement correlation and cross-loadings between items and dimensions. If ignored, they could affect not only the acceptability of a model based on goodness-of-fit indices, but also the parameter estimates (Hair et al. 2006).

The purpose of this study is to examine DCSQ additional psychometric properties, namely, the instrument's dimensional structure concerning number and indicator (item) patterns; potential cross-loadings and error measurement correlations; CFA-based convergent validity and internal consistency. The role of social support at work is evaluated in tandem. Separate analyses according to work setting, hospital and restaurants, are also employed to investigate differences in dimensional structure and in loading patterns.

Methods

Data for the analysis came from two cross-sectional studies in Rio de Janeiro, Brazil. The first one was a multidimensional study of hospital workers. Of 497 estimated active workers in 2004, 424 agreed to participate, and 399 (80.3%) completed the self-administered DCSQ in 2004–2005. Workers aged less than 18 years and those who were inactive or had been working for less than 2 years were excluded.

The second study was performed to evaluate work impairment at nine state-run restaurants. Of 463 active workers in 2006, 426 (92%) completed the DCSQ, which was applied by trained interviewers. Interview assessment was used to avoid missing answers since this occupational group has low reading skills. Workers who were absent, inactive or did not work directly in meal processing were excluded. Both studies were approved by Research Ethics Committee (protocol no 0024.0.009.000-03 and 15/2006, respectively), and all participants signed a free informed consent form.

Both studies employed the Brazilian version of 17-DCSQ (Alves et al. 2004). Each item of psychological demands and decision latitude is scored on a Likert-type ranging from 1 (never/almost never) to 4 (often). For two items, *enough time* and *repetitive work*, scores are reversed. Social support at work consists of six items, and its response categories are also graded on a Likert-type ranging from 1 (strongly disagree) to 4 (strongly agree). Dimension scores are the sum of items and ranged from 5 to 20 for demands and from 6 to 24 for decision latitude and social support at work (see "Appendix").

Univariate analysis was used to describe the sample by age, gender and years of schooling. Job titles were evaluated according to the major ISCO groups (ISCO 1988).

Exploratory factor analysis (EFA) using WLSMV estimator and *geomin* oblique rotation was implemented for each group separately to investigate common dimensional structure, as well as plausible alternatives to the original theoretical model, to be carried out in CFA-type analysis for the whole sample (Muthén 1998–2004).

Confirmatory factor analysis (CFA) using structural equation models was carried out for the whole sample to test the original factor structure proposed by Karasek's theory and outlined in other validation studies (Niedhammer 2002; Sanne et al. 2005).

Three models were tested. Model 1 predicted the existence of three correlated latent factors representing psychological demands, decision latitude and social support at work. Model 2 consisted of four correlated latent factors (due to the subdivision of decision latitude in two factors): psychological demands, skill discretion, decision authority and social support at work. Finally, the third model was based on the original job strain concept (Karasek and Theorell 1990), which does not include social support at work—a potential effect modifier of the association between job strain and health outcomes.

The analyses were carried out in Mplus software, version 5.1 (Muthén and Muthén 1998–2007), which provided the robust weighted least square parameter estimates using a diagonal weight matrix with standard errors and mean- and variance-adjusted chi-square statistics that use a full weight matrix (WLSMV) (Muthén 1993; Muthén and Muthén 1998–2007). Polychoric correlation matrices were appropriately used for categorical variables (Finney and Di Stefano 2006). Estimates of CFA model parameters included standardized factor loadings with the respective 95% confidence interval (95% CI), factor correlations and indicator error variances.

Goodness of fit was evaluated through four fit indices (Brown 2006). To evaluate absolute fit, weighted root mean square residual (WRMR) has been recommended when using WLSMV (Muthén 1998–2004). Values below 0.90 represent adequate fit. Root mean square error of approximation (RMSEA) incorporates a penalty function for poor model parsimony, expressed by model degrees of freedom. Values under 0.06 are recommended, whereas values above 0.10 indicate poor fit and that the model should be rejected (Muthén and Muthén 1998–2007). Comparative fit index (CFI) and Tucker–Lewis index (TLI) represent incremental fit indices (Brown 2006) comparing the hypothesized model to a more restricted nested baseline model, the "null model". CFI and TLI range from zero to one, and values above 0.90 are indicative of adequate fit (Brown 2006). Nested models were compared using the *difftest* option of Mplus, which is adequate to evaluate chi-square difference test when using the WLSMV estimator (Muthén and Muthén 1998–2007).

Modification indices (MI) were also explored in order to identify parameter misfit. The modification index reflects how much the overall model chi-square would decrease if a constrained parameter was freely estimated. Possible correlations between indicator measurement errors or cross-loadings between indicators and factors not previously specified in the model under inspection involving values of modification index equal or more than 10 would be further examined, as well as the magnitude of the corresponding expected parameter changes for freely estimated parameters (Brown 2006).

Factor-based convergent validity (Kline 2005) was assessed through factor loadings, internal consistency and average variance extracted (AVE). It was beyond the scope of this article to investigate construct validity through the assessment of associations/correlations between mean scores of DCSQ dimensions and socio-demographic characteristics or health outcomes.

Composite reliability (CR) was used to estimate the overall scale internal consistency, since the traditional Cronbach's alpha coefficient tends to underestimate the reliability unless under restricted assumptions of equal common factor loadings and uncorrelated measurement errors. For each factor, $\rho_{cr} = \left(\sum_{i=1}^{k}\lambda_i\right)^2 \Big/ \left(\sum_{i=1}^{k}\lambda_i\right)^2 + \sum_{i=1}^{k}\delta_i$, where λ_i represents the standardized factor loadings of i indicators, and δ_i is the related measurement error (uniqueness) that refers to the portion of an indicator not explained by the latent factor. CR ranges from 0 to 1, and values $\geq$0.70 indicate acceptable internal consistency (Hair et al. 2006).

The AVE assesses the amount of variance captured by a common factor in relation to the amount of variance due to random measurement error (Hair et al. 2006). It is calculated as $\rho_{ve} = \sum_{i=1}^{k} \lambda_i^2 \Big/ \left(\sum_{i=1}^{k} \lambda_i^2 + \sum_{i=1}^{k} \delta_i^2 \right)$, where λ_i and δ_i are defined as above. It also vary from zero to one, and a factor shows convergent validity if $\rho_{ve} \geq 0.50$, which is indicative that at least 50% of the variance in a measure is due to the hypothesized latent variable. If AVE < 0.50, then the variance due to measurement error is greater than the variance due to the construct and the convergent validity is questionable (Hair et al. 2006). Factor-based discriminant validity (Kline 2005) was examined using the AVE as benchmark (Hair et al. 2006). In multidimensional models, discriminant validity is said to hold if, for a given factor, $\sqrt{\rho_{ve}}$ is above the absolute value of its correlation with any other factor in the model (Fornell and Larcker 1981). Ninety-five percent confidence intervals (95% CI) for AVE and CR were obtained via bootstrap method with 1,000 replications (Efron and Tibshirani 1993). If correlated measurement errors were considered in the CFA model, the sum of nonzero error covariances multiplied by 2 was added to the denominator (Brown 2006).

The same scrutiny applied for the whole sample was repeated in separate analyses according to work setting, which in our study represents high- and low-level occupational and educational groups (hospital and restaurant workers, respectively), to investigate similarities/disparities in dimensional structure as well as in loading patterns.

Results

Of the 825 workers included in the study, 52.3% were females, mean age was 35.1 years (SD: 10.8), 62.7% had less than 12 years of schooling, and 48.0% were in the highest skill level occupational groups, which represent major ISCO groups 1–4 (ISCO 1988). At the hospital (*H*), there were more females (68.3%) and 30.2% had studied less than 12 years, while at restaurants (*R*), there were more males (62.7%) and most of the workers had studied less than 12 years (92.7%).

Mean scores for DCSQ were 13.3 for psychological demands (95% CI: 13.1; 13.5), 16.3 for decision latitude (95% CI: 16.1; 16.5) and 20.4 for social support at work (95% CI: 20.1; 20.6), with a significant difference ($P < 0.01$) between the work settings for decision latitude (*H*: 17.7 and *R*: 15.0).

Exploratory factor analysis supported a common three-factor solution for both groups, explaining 41% (*H*) and 30% (*R*) of the variance. This closely corresponds to the theoretical structure of the DCSQ (demands, decision latitude and social support at work).

The original three-factor solution showed low loadings for decision latitude, particularly for *initiative* and *repetitive work* (0.39 and 0.19, respectively). Measurement errors were high (≥0.70) for all items of decision latitude and for *enough time* and *conflicting demands* of psychological demands dimension. All goodness-of-fit indices were inadequate. (model 1, Table 1).

Average variance extracted was not acceptable for psychological demands (0.44) and decision latitude (0.31), indicating that convergent validity was questionable for the first model. However, composite reliability was adequate for all three dimensions. Factor correlation between psychological demands and decision latitude was not statistically significant. (model 1, Table 2).

Since the first model showed misfit, a full EFA was implemented. Four empirical factors with the following eigenvalues were disclosed: 4.21 (factor 1), 2.57 (factor 2), 2.36 (factor 3) and 1.10 (factor 4). The ensuing CFA model thus included four factors: demands, skill discretion, decision authority and social support at work. As in model 1, *repetitive work* resulted in a low loading, and in the current model, it was not statistically significant, since the 95% confidence interval included the null value. Measurement errors decreased for *skill level* and *initiative* of skill discretion (model 2, Table 1). Goodness-of-fit indices were still inadequate, in spite of the improvements regarding the previous model (*difftest*: 53.30, $P < 0.01$).

The average variance extracted was low for skill discretion, which means that the variance due to measurement errors was greater than the one due to the latent factor. Composite reliability was moderate for skill discretion. The correlation between skill discretion and social support at work was low and nonsignificant (model 2, Table 2).

Modification indices (MI) showed that freely estimating the loading of *learning new things* on the psychological demands dimension would decrease the model's chi-square by 83.3, with an expected parameter change (EPC) of −0.43. MI also showed that an error measurement correlation between *work fast* and *work intense* would decrease the model's chi-square by 43.4 and yield an EPC of 0.40. Thus, an alternative four-factor model (data not shown) was tested without the item *repetitive work*, in which the cross-loading from *learning new things* on the psychological demands dimension (−0.44; 95% CI: −0.57, −0.31) and the error measurement correlation between *work fast* and *work intense* were confirmed (0.50; 95% CI: 0.42, 0.57).

Since the original job strain concept does not include social support at work, this dimension was removed from the analysis in the third model, wherein the same cross-loading and error measurement correlation described before and detected by MI inspection were considered. The results showed acceptable loadings for all items,

Table 1 Confirmatory factor analyses of the demand control support questionnaire: standardized factor loadings (λ_i) with 95% confidence interval (95% CI), standardized measurement errors (δ_i) and goodness-of-fit indices

Dimension/items	Model 1[a]		Model 2[b]		Model 3[c]	
	λ_i (95% CI)	δ_i	λ_i (95% CI)	δ_i	λ_i (95% CI)	δ_i
Psychological demands (PD)						
1 *Work fast*	0.72 (0.67, 0.77)	0.48	0.71 (0.66, 0.76)	0.49	0.60 (0.53, 0.67)	0.64
2 *Work intense*	0.84 (0.80, 0.88)	0.29	0.84 (0.79, 0.88)	0.30	0.75 (0.68, 0.82)	0.44
3 *Work effort*	0.71 (0.66, 0.76)	0.50	0.72 (0.67, 0.77)	0.48	0.81 (0.74, 0.87)	0.35
4 *Enough time*	0.47 (0.38, 0.56)	0.78	0.47 (0.38, 0.56)	0.78	0.41 (0.32, 0.50)	0.83
5 *Conflicting demands*	0.52 (0.45, 0.58)	0.73	0.51 (0.44, 0.58)	0.74	0.47 (0.40, 0.54)	0.78
Decision latitude (DL)						
Skill discretion (SD)						
6 *Learning new things*	0.50 (0.43, 0.57)	0.75	0.51 (0.42, 0.59)	0.74	0.73 (0.62, 0.84)	0.58
7 *Skill level*	0.44 (0.37, 0.52)	0.81	0.69 (0.61, 0.77)	0.52	0.66 (0.59, 0.74)	0.56
8 *Initiative*	0.39 (0.31, 0.46)	0.85	0.62 (0.54, 0.70)	0.61	0.62 (0.54, 0.70)	0.61
9 *Repetitive work*	0.19 (0.10, 0.28)	0.96	0.08 (−0.02, 0.18)	0.99	.	.
Decision authority (DA)						
10 *How to do the work*	0.75 (0.68, 0.82)	0.44	0.78 (0.70, 0.87)	0.39	0.77 (0.68, 0.85)	0.41
11 *What to do at work*	0.83 (0.76, 0.90)	0.32	0.85 (0.76, 0.93)	0.28	0.87 (0.78, 0.95)	0.25
Social support at work (SSW)						
12 *Pleasant atmosphere*	0.63 (0.58, 0.68)	0.60	0.63 (0.58, 0.68)	0.60	.	.
13 *Relationship at work*	0.81 (0.77, 0.84)	0.35	0.81 (0.77, 0.84)	0.35	.	.
14 *Coworkers support*	0.76 (0.71, 0.80)	0.43	0.76 (0.71, 0.80)	0.43	.	.
15 *Coworkers helpful*	0.71 (0.67, 0.76)	0.49	0.71 (0.67, 0.76)	0.49	.	.
16 *Relationship with supervisors*	0.64 (0.58, 0.70)	0.59	0.64 (0.58, 0.70)	0.59	.	.
17 *Friendly coworkers*	0.81 (0.76, 0.86)	0.34	0.81 (0.76, 0.86)	0.34		
Cross-loading of item 6 on PD					–0.41 (−0.53, −0.29)	
Error measurement correlation (item 1⇔ item 2)					0.43 (0.33, 0.52)	
Goodness-of-fit indices[d]						
WRMR	2.108		1.869		1.073	
RMSEA	0.099		0.090		0.062	
CFI	0.865		0.889		0.968	
TLI	0.881		0.902		0.961	

[a] Model 1: PD ⇔ DL ⇔ SSW

[b] Model 2: PD ⇔ SD ⇔ DA ⇔ SSW

[c] Model 3: PD ⇔ SD ⇔ DA without the item 9, with a cross-loading of item 6 on PD and a correlated measurement error between items 1 and 2

[d] *WRMR* weighted rood mean square residual, *RMSEA* root mean square error of approximation, *CFI* comparative fit index and *TLI* Tucker–Lewis index

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particularly for skill discretion (model 3, Table 1). The cross-loading from *learning new things* on the psychological demands dimension (−0.41; 95% CI: −0.53, −0.29) and the correlation between *work fast* and *work intense* measurement errors (0.43; 95% CI: 0.33, 0.52) were sustained. Goodness-of-fit indices improved vis-à-vis all previous models (model 3, Table 1). The average variance extracted and the composite reliability improved for skill discretion but decreased for psychological demands (model 3, Table 2).

In separate analyses for hospital and restaurant workers, the final model for both groups was the same obtained by the whole sample. The cross-loading remained meaningful (*H*: −0.43; 95% CI: −0.63, −0.23 and *R*: −0.40; 95% CI: −0.60, −0.20) as well as the error measurement correlation between *work fast* and *work intense* (*H*: 0.38; 95% CI: 0.22, 0.54 and *R*: 0.41; 95% CI: 0.29, 0.53). However, the loadings of four items of psychological demands (*work fast*, *work intense*, *enough time* and *conflicting demands*) and of one item of skill discretion (*initiative*)

Table 2 Confirmatory factor analysis of the demand control support questionnaire: average variance extracted (AVE), composite reliability (CR) and factor correlations, with the respective 95% confidence intervals

Dimensions	Model 1[a]		Model 2[b]		Model 3[c]	
Psychological demands (PD)						
AVE	0.44	(0.41, 0.48)	0.44	(0.41, 0.48)	0.32	(0.28, 0.36)
CR	0.79	(0.76, 0.82)	0.79	(0.76, 0.82)	0.58	(0.54, 0.63)
Decision latitude (DL)						
AVE	0.31	(0.28, 0.34)	.	.	.	.
CR	0.70	(0.66, 0.74)	.	.	.	.
Skill discretion (SD)						
AVE	.	.	0.28	(0.24, 0.33)	0.44	(0.37, 0.50)
CR	.	.	0.56	(0.49, 0.62)	0.70	(0.64, 0.75)
Decision authority (DA)						
AVE	.	.	0.66	(0.61, 0.72)	0.67	(0.61, 0.72)
CR	.	.	0.80	(0.76, 0.84)	0.80	(0.76, 0.84)
Social support at work (SSW)						
AVE	0.53	(0.50, 0.57)	0.53	(0.50, 0.57)	.	.
CR	0.87	(0.85, 0.89)	0.87	(0.85, 0.89)	.	.
Factor correlations						
PD ⇔ DL	−0.01	(−0.10, 0.08)	.	.	.	.
PD ⇔ SSW	−0.32	(−0.40,−0.24)	−0.32	(−0.40,−0.24)	.	.
DL ⇔ SSW	0.10	(0.02, 0.19)	.	.	.	.
PD ⇔ SD	.	.	0.28	(0.18, 0.37)	0.46	(0.36, 0.56)
PD ⇔ DA	.	.	−0.12	(−0.21,−0.03)	−0.12	(−0.22,−0.03)
SD ⇔ DA	.	.	0.53	(0.44, 0.62)	0.46	(0.36, 0.55)
SD ⇔ SSW			0.03	(−0.07, 0.13)	.	.
DA ⇔ SSW	.	.	0.11	(0.02, 0.20)	.	.

[a] Model 1: PD ⇔ DL ⇔ SSW

[b] Model 2: PD ⇔ SD ⇔ DA ⇔ SSW

[c] Model 3: PD ⇔ SD ⇔ DA without the item repetitive work, with a cross-loading from learning new things on PD and a correlated measurement error between the items work fast and work intense

"." category not applicable

were lower for restaurant workers (Table 3). Similar patterns were found when inspecting average variance extracted and composite reliability, particularly for demands. The factor correlation between demands and decision authority was nonsignificant for restaurant workers (Table 4).

Discussion

The factor pattern of this study corroborates the demand control model, in which a clear two-factor solution for decision latitude is demonstrated and also detected by Sanne et al. (2005) but restricted to men in the high-status and women in low-status occupation groups. Validation studies with JCQ corroborated the same dimensional structure (Kawakami and Fujigaki 1996; Smith et al. 1997; Karasek et al. 1998; Schreurs and Taris 1998; Niedhammer 2002).

Consistent with several JCQ studies, *repetitive work* showed a nonsignificant loading (Karasek 1979; Smith et al. 1997; Karasek et al. 1998; Escriba-Aguir et al. 2001; Niedhammer 2002; Li et al. 2004; Niedhammer et al. 2006; Eum et al. 2007). The low loadings of *enough time* and *conflicting demands* on the psychological demands dimension were also observed (Karasek 1979; Karasek et al. 1998; Escriba-Aguir et al. 2001; Pelfrene et al. 2001; Niedhammer 2002; Sale and Kerr 2002; Li et al. 2004; Niedhammer et al. 2006).

When JCQ and DCSQ were applied simultaneously to the same Swedish sample, Karasek et al. (2007) found different frequency distributions for the equivalent items *enough time* and *repetitive work*. It might be possible that differences in item wording and/or response categories could lead to different patterns of item–factor relationships, and the findings related to these items would not be comparable.

Lack of correlation between the psychological demands and decision latitude dimensions in model 1 was also observed for the Korean version of JCQ (Li et al. 2004) and in a cross-national comparative study including six samples from USA, Canada, Netherlands and Japan (Karasek et al. 1998). This finding was not corroborated in larges samples of Belgian (Pelfrene et al. 2001) and French workers (Niedhammer 2002), and also when DCSQ was filled out by Norwegian workers (Sanne et al. 2005).

A moderate correlation between skill discretion and decision authority was also detected in four JCQ studies (Karasek 1979; Pelfrene et al. 2001, Niedhammer 2002; Li et al. 2004). Karasek et al. (2007) observed a lower

Table 3 Confirmatory factor analysis of the final model stratified by work setting: standardized factor loadings (λ_i) with 95% confidence interval (95% CI), standardized measurement errors (δ_i) and goodness-of-fit indices

Dimension/items	Hospital		Restaurants	
	λ_i (95% CI)	δ_i	λ_i (95% CI)	δ_i
Psychological demands (PD)				
1 *Work fast*	0.73 (0.65, 0.81)	0.46	0.48 (0.36, 0.60)	0.77
2 *Work intense*	0.84 (0.79, 0.90)	0.29	0.64 (0.52, 0.76)	0.59
3 *Work effort*	0.82 (0.75, 0.88)	0.33	0.84 (0.71, 0.97)	0.30
4 *Enough time*	0.66 (0.58, 0.75)	0.56	0.16 (0.01, 0.32)	0.97
5 *Conflicting demands*	0.56 (0.47, 0.64)	0.69	0.36 (0.25, 0.48)	0.87
Skill discretion (SD)				
6 *Learning new things*	0.81 (0.63, 0.98)	0.54	0.62 (0.42, 0.82)	0.70
7 *Skill level*	0.59 (0.48, 0.69)	0.66	0.74 (0.59, 0.88)	0.46
8 *Initiative*	0.81 (0.70, 0.91)	0.35	0.48 (0.35, 0.60)	0.77
Decision authority (DA)				
10 *How to do the work*	0.78 (0.67, 0.90)	0.39	0.69 (0.51, 0.87)	0.53
11 *What to do at work*	0.81 (0.70, 0.92)	0.34	0.88 (0.67, 0.99)	0.22
Cross-loading of item 6 on PD	−0.43 (−0.63, −0.23)		−0.40 (−0.60, −0.20)	
Error measurement correlation item 1⇔ item 2	0.38 (0.22, 0.54)		0.41 (0.29, 0.53)	
Goodness-of-fit indices[a]				
WRMR	0.689		0.683	
RMSEA	0.048		0.036	
CFI	0.988		0.980	
TLI	0.987		0.974	

Final model: PD ⇔ SD ⇔ DA without the item *repetitive work* (item 9 of SD), with a cross-loading of item 6 on PD and a correlated measurement error between items 1 and 2

[a] *WRMR* weighted rood mean square residual, *RMSEA* root mean square error of approximation, *CFI* comparative fit index and *TLI* Tucker–Lewis index

correlation between skill discretion and decision authority for DCSQ than for JCQ, probably attributed to the different number of items by dimension.

We attempted to test a higher-order factor for decision latitude, represented by skill discretion and decision authority sub-dimensions, correlated to the single-order psychological demands dimension, but the model was under-specified. Although this model has not been tested before for JCQ, it is possible that the few number of items for decision authority in DCSQ and the problems encountered in skill discretion dimension (low loading of *repetitive work* and cross-loading of *learning new things*) contributed for model misspecification. A higher-order CFA model was performed only for decision latitude with Canadian teaching hospital workers, and the best fit model was the one without *repetitive work* and with correlated measurement errors within skill discretion and between skill discretion and decision authority items (Sale and Kerr 2002).

The role of social support at work is not well defined in validation studies of JCQ or DCSQ, although the concept and the interaction effect are well established (Johnson and Hall 1988). The DCSQ validation study (Sanne et al. 2005) and most of the JCQ studies included this dimension in factor analysis (Kawakami and Fujigaki 1996; Escriba-Aguir et al. 2001; Pelfrene et al. 2001; Niedhammer 2002; Li et al. 2004; Edimansyah et al. 2006; Niedhammer et al. 2006; Eum et al. 2007), while others excluded (Brisson et al. 1998; Schreurs and Taris 1998; Sale and Kerr 2002). In our study (models 1 and 2), this dimension presented high item loadings and high internal consistency, which was consistent with Sanne et al. (2005) results.

This is the first published study about psychometric properties of the Demand Control Questionnaire using composite reliability as a measure of internal consistency, since the assumptions for Cronbach's alpha, common factor loadings and uncorrelated measurement errors, were violated.

The final model showed acceptable levels of internal consistency measured by composite reliability, except for the psychological demands dimension in which a correlated measurement error between *work fast* and *work intense* was observed. This correlation was also detected with hospital workers in Canada (Sale and Kerr 2002). One study with a wide range of Belgian workers (Pelfrene et al. 2001) extracted six factors in EFA, where the sixth was represented by *work fast* and *work intense*. Since EFA approach cannot explore error measurement correlation between items, this finding might represent a method

Table 4 Confirmatory factor analysis of the final model stratified by work setting: average variance extracted (AVE), composite reliability (CR) and factor correlations, with the respective 95% confidence intervals

Dimensions	Hospital		Restaurants	
Psychological demands (PD)				
AVE	0.43	(0.37, 0.50)	0.25	(0.20, 0.30)
CR	0.72	(0.66, 0.77)	0.45	(0.36, 0.54)
Skill discretion (SD)				
AVE	0.51	(0.42, 0.61)	0.38	(0.29, 0.46)
CR	0.76	(0.68, 0.83)	0.64	(0.55, 0.72)
Decision authority (DA)				
AVE	0.64	(0.54, 0.73)	0.63	(0.37, 0.88)
CR	0.78	(0.71, 0.85)	0.77	(0.60, 0.93)
Factor correlations				
PD ⇔ SD	0.55	(0.43, 0.66)	0.49	(0.33, 0.65)
PD ⇔ DA	−0.14	(−0.26, −0.02)	−0.01	(−0.16, 0.13)
SD ⇔ DA	0.37	(0.24, 0.50)	0.38	(0.23, 0.52)

Final model: PD ⇔ SD ⇔ DA without the item repetitive work, with a cross-loading from learning new things on PD and a correlated measurement error between the items work fast and the work intense

effect, which means that this sixth factor is not substantively meaningful (Brown 2006).

The fact that *learning new things* loaded not only on the skill discretion, but also on the psychological demands dimension was also observed with Korean health care workers and with Canadian white collar workers, who both filled out the JCQ (Brisson et al. 1998; Eum et al. 2007). This finding could indicate that *learning new things* might represent psychological demands in modern working life (Kawakami et al. 1995).

Limitations of this study include restricted sample composition to two work settings, although there were workers from five and six of the nine major ISCO groups, at restaurants and at the hospital, respectively. The study sample size was in accordance with what is recommended in simulation research for CFA-based analyses (Brown 2006; Gagné and Hancock 2006; Hair et al. 2006) and was similar to other validation studies (Kawakami et al. 1995; Kawakami and Fujigaki 1996; Li et al. 2004; Eum et al. 2007). It has been shown that general rules of thumb are of limited utility, neither for an absolute minimum sample size nor for a critical ratio of sample size to number of indicators or sample size to number of free parameters (Hair et al. 2006). In addition, simulation research has shown that WLSMV estimator performs well with samples as small as 200 for non-normal variables (Brown 2006).

In summary, our results indicate that the psychometric properties of the three DCSQ dimensions (psychological demands, skill discretion and decision authority) were satisfactory in a Brazilian sample population. The different loading patterns found for psychological demands across work settings could indicate that most of these items are sensitive to occupational or educational status, which means that item wording and the whole dimension may have a different meaning across high- and low-level occupational/educational groups. Thus, cross-cultural differences in demands mean scores may not reflect true group variations, but instead socio-occupational bias (Stewart and Napoles-Springer 2000). Additionally, the results suggest that social support at work should not be considered together with demand and control scales in factorial analyses; nevertheless, it could be considered in construct validity studies, by evaluating its association with the demand control scales, as well its role as an effect modifier in association studies of job strain and stress-related health outcomes.

Consistently with other studies, our findings underline the fact that *repetitive work* should be removed from the skill discretion dimension. Similar conclusion could be drawn for the item *learning new things*, since it seems to be nonspecific of skill discretion dimension. In this study, *work fast* and *work intense* are duplicated, which means that one of them should be removed from the psychological demands dimension, but this finding needs further confirmation.

Since DCSQ dimensional structure has been scarcely investigated, it would be desirable to confirm these findings in a Swedish sample, where the questionnaire was originally developed.

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Conflict of interest None.

Appendix

Appendix Demand control support questionnaire

Psychological demands

1. Do you have to work very fast? (*work fast*)
2. Do you have to work very intensively? (*work intense*)
3. Does your work demand too much effort? (*work effort*)
4. Do you have enough time to do everything? (*enough time*)
5. Does your work often involve conflicting demands? (*conflicting demands*)

Decision latitude

Skill discretion

6. Do you have the possibility of learning new things through your work? (*learning new things*)
7. Does your work demand a high level of skill or expertise? (*skill level*)
8. Does your job require you to take the initiative? (*requires initiative*)
9. Do you have to do the same thing over and over again? (*repetitive work*)

Decision authority

10. Do you have a choice in deciding HOW you do your work? (*how to do the work*)
11. Do you have a choice in deciding WHAT you do at work? (*what to do at work*)

Social support at work

12. There is a calm and pleasant atmosphere where I work (*pleasant atmosphere*)
13. We get on well with each other where I work (*relationship at work*)
14. My coworkers support me (*coworkers support*)
15. The others understand if I have a bad day (*coworkers helpful*)
16. I get on well with my supervisors (*relationship with supervisors*)
17. I enjoy working with my coworkers (*friendly coworkers*)

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