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THE JOB DEMAND-CONTROL MODEL OF JOB STRAIN ACROSS GENDER

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ABSTRACT

This study investigated (a) the structural and measurement equivalence of measures of the job demand-control (JDC) model of job strain for men and women (as operationalised by the Sources of Work Stress Inventory), (b) whether a common or separate regression equations should be used for men and women in the JDC model, and (c) the strain and buffer hypotheses associated with the JDC model. These objectives were pursued by means of factor analysis, item response theory analysis, and moderated hierarchical multiple regression analysis. The results show that the use of a common regression equation might give biased results. The results also provided stronger support for the strain hypothesis than for the buffer hypothesis.

OPSOMMING

Hierdie studie het ondersoek ingestel na (a) die strukturele en metingsekwivalensie van metings van die werkseise-kontrole (WEK) model van werkspanning vir mans en vroue (soos geoperasionaliseer deur die Bronne van Werkstresinventaris), (b) die wenslikheid van 'n gemeenskaplike teenoor afsonderlike regressievergelykings vir mans en vroue, en (c) die spanning en buffer hipoteses wat met die WEK model geassossiëer word. Hierdie doelstellings is ondersoek aan die hand van faktorontleding, item responsteorie-ontleding en gemodereerde hiërargiese meervoudige regressie-ontleding. Die resultate wys dat die gebruik van 'n gemeenskaplike regressievergelyking sydige resultate mag oplewer. Die resultate toon verder sterker ondersteuning vir die spanning hipotese as vir die buffer hipotese.

Different kinds of jobs introduce different levels of work stress, due to, amongst others, the following environmental factors: the amount of work needed to be done, the amount of decision-making authority an individual has, and the extent to which an individual can choose to employ his or her skills (Theorell & Karasek, 1996). Theorell and Karasek call these three environmental elements respectively: job demands, decision authority, and skill discretion. The latter two elements jointly constitute job control (which Theorell and Karasek also refer to as decision latitude). Karasek (1979) introduced a model of job strain that accounts for the relationship between job demands, job control, and negative health and psychological outcomes. This model is most often referred to as the job demand-control model (hereafter referred to as the JDC model).

The demands component of the model is most often conceptualised as time pressure due to a heavy workload (Fernet, Guay & Senécal, 2004; Karasek & Theorell, 1990), but it may be broadened to also include role ambiguity and role conflict. The job control dimension is often conceptualised as the sum of two components, namely skill discretion and decision authority. The decision authority component refers to the opportunity to make independent decisions and to have a say in what happens in the workplace. Fernet et al. (2004) pointed out that the decision authority component concerns "opportunities for control and decision and therefore ... job control per se" (p. 45). In agreement with this view Wall, Jackson, Mullarkey and Parker (1996) recommended that measures of job control should focus on the decision authority component. The skill discretion component mostly addresses task variety and appears only obliquely related to job control.

The JDC model has been expanded to include a support dimension and is referred to as the job demand-control-support model (JDCS model). The expanded model makes the additional prediction that low levels of social support from supervisors and peers will contribute to job strain. In the present study the focus falls on the original JDC model rather than on the expanded model.

The basic premise of Karasek's (1979) JDC model is that job demands and job control interact in such a way that they create different psychosocial work experiences for the individual, depending on the respective magnitudes of job demands and job control. Karasek (1979) classified these work experiences into four types of jobs, namely high-strain jobs (high demands and low control), active jobs (high demands and high control), low-strain jobs (low demands and high control), and passive jobs (low demands and low control).

The strain and buffer hypotheses of the JDC model

Van der Doef and Maes (1999) drew a distinction between two hypotheses associated with the JDC model. The first hypothesis, labelled the strain hypothesis, predicts that job demands and job control combine additively to produce negative psychological and health outcomes in environments characterised by high job demands and low job control. The strain hypothesis holds the practical implication that job demands *and* job control need to be addressed to reduce job strain. The second hypothesis, labelled the buffer hypothesis, predicts that job demands and job control combine multiplicatively and that job control moderates the negative effect of job demands on health and well-being. Specifically, high job control is predicted to ameliorate the negative effects of high job demands (Karasek, 1979). The buffer hypothesis holds the practical implication that improved health and psychological well-being in employees may be obtained by increasing job control without reducing job demands. In this regard Wall et al. (1996) stated that "... increased control reduces the effects of stressors by allowing individuals to face demands when they are best able to do so and in ways they find most acceptable" (p. 155).

Research has provided relatively consistent and convincing support for the strain hypothesis, but less so for the buffer hypothesis (Van der Doef & Maes, 1999). Subsequent to Van der Doef and Maes' narrative meta-analysis several studies have tested the strain and buffer hypotheses concurrently. The general picture arising from these more recent studies is that the data supports the general strain hypothesis, but not the buffer hypothesis (e.g. Pelfrene, Vlerick, Kittel, Mak, Kornitzer & de Backer, 2002; Rafferty, Friend & Landsbergis, 2001; Van der Doef et al., 2000; Verhoeven, Maes, Kraaij & Joekes, 2003).

The role of gender in the JDC model

Although researchers typically include gender as a covariate in studies of the JDC model, the role of gender has not yet been fully explored. Van der Doef and Maes (1999) reported that studies using women as participants more often failed to support the strain hypothesis and concluded that men and women may react differently to the effects of high-strain work, with men appearing more vulnerable to the negative effects of high job demands and low job control. In this regard, Vermeulen and Mustard (2000) concluded that characteristics of the work place, such as job demands and job control, may have a greater impact on the psychological well-being of men compared with women. This has prompted some researchers to refer to the JDC model as a "male model" (cf. Johnson & Hall, 1988).

A common analytic strategy employed by job demand-control researchers is to include gender along with other demographic variables, such as age, in the first step of a hierarchical multiple regression analysis. The effect of this strategy is to partial out from the dependent and independent variables variance that is shared with gender (and the other demographic variables included in the first step of the hierarchical analysis). The regression weights associated with gender in the first step are typically small, indicating little differences between men and women in regard to their means on dependent variables such as burnout, job dissatisfaction and work stress (cf. Rafferty et al., 2001; Rodríguez Bravo, Peiró & Schaufeli, 2001; Van der Doef et al., 2000). A shortcoming of these analyses, however, is that the assumption is made that a common regression equation, with equal slopes and intercepts, applies for men and women. This, however, is an empirical issue that has to be tested directly rather than assumed (cf. Pedhazur, 1997).

Verhoeven et al. (2003) used separate regression equations for men and women and found that job demands, job control, and social support accounted for the same amount of variance in outcome variables for men and women. However, a potentially more appropriate and explicit test of the matter of a common versus separate regression equations would entail testing one of two (or both) of the following conditions. A first test would be to investigate whether the inclusion of the multiplicative effects of gender × job control, and gender × job demands account for a statistically significant amount of variance in the outcome variables, over and above that accounted for by job demands, job control, the multiplicative effect of job demands × job control, and gender as a main effect. A second test would be to determine whether the inclusion of gender as a main effect accounts for a statistically significant amount of variance in the outcome variables, over and above that accounted for by job demands and job control and their multiplicative effect. The first test is a test of equal regression slopes, whereas the second test is a test of equal intercepts. Only when the hypothesis of equal regression slopes is not rejected is it appropriate to test whether the intercepts are equal (Pedhazur, 1997). Barnett and Brennan (1997) did explicitly test for a multiplicative effect of gender and job conditions and found that, after controlling for the gender-related covariates of full-time employment, occupational prestige and household income, the effects of job demands and job control with psychological distress did not depend on gender.

One possible reason why the question of a common versus separate regression equations for men and women has received little research attention may be that researchers automatically assume that a common regression equation will hold. This assumption implies that job demands and job control have the same effect on men and women and is possibly not warranted. A second possible reason is that gender is a very complex independent variable. Gender is correlated with many other variables, each of which may be correlated with or influence the variables in the JDC model. Obvious potentially confounding variables include job or occupational status (with women probably holding lower status jobs in general), differences in remuneration, full-time versus part-time employment, and parenthood, but there may be many other such variables. Controlling for all confounding variables associated with gender becomes an impossible task and attempting to do so may give misleading results. However, we believe that it is important to empirically ascertain whether the relations of job strain with job demands and job control are the same for important demographic groups, such as men and women. Failure to do so may lead to bias in the estimates of the regression coefficients. On the basis of the meta-analysis of Van der Doef and Maes (1999) one might expect the effects of job demands and job control to be upwardly biased for women, and downwardly biased for men.

Construct and measurement equivalence for men and women

The construct and measurement equivalence of the concepts in the JDC model for men and women is another aspect that has received little research attention. All studies that combine the data of men and women or compare the measures of men and women proceed on the assumption that the scales are perceived in the same way and have the same meaning for the two groups. This, however, is an empirical matter that has to be explicitly investigated. To the extent that measures of the JDC model differ for men and women the observed relations between the concepts may be biased.

Santavirta (2003) demonstrated that the factor structure of the Finnish version of the Job Content Questionnaire was similar for men and women, suggesting that qualitatively similar constructs were measured for the two groups. A confirmatory factor analysis of the Job Content Questionnaire in Belgium showed that a model with equal factor loadings for men and women, which is a minimum requirement for measurement equivalence, fit the data satisfactorily (Pelfrene, Vlerick, Mak, De Smet, Kornitzer & De Backer, 2001).

A more stringent test of measurement equivalence, however, would require that the item discrimination parameters (which are conceptually similar to factor loadings) and item location parameters (which are conceptually equivalent to the "difficulty" or "endorsability" of items) in an item response theory analysis be equal for men and women (Embretson & Reise, 2000). Items for which the parameters differ across groups are said to demonstrate *differential item functioning* (DIF) and the inclusion of such items in a scale may lead to biased measures for one or more groups. Psychometricians distinguish between uniform DIF, which is present when the item location parameter of an item differs across groups, and non-uniform DIF, which is present when the item discrimination parameter of an item differs across groups. In the Rasch model (Rasch, 1960), which is the item response theory model employed in the present study, all items are required to have equal discrimination parameters and only the item location parameters are allowed to vary. Hence, from a Rasch perspective it is usually only uniform DIF that is investigated. Items that display non-uniform DIF are identified with item fit statistics (R.M. Smith, 2004a). In the context of Likert-type items, as is the case in the present study, an item displays DIF if individuals from different groups (such as men and women), but with equal standings on the trait of interest, have different expected raw scores for the item. To the best of our knowledge no item response theory analyses in general, and Rasch analyses in particular, of measures of the JDC model have been published.

Aims and hypotheses of the present study

We set out to test the construct and measurement equivalence for men and women of the measures of the JDC model used in this study. In this regard we aimed to show (a) that the factor structure of the measures are qualitatively similar for men and women and (b) that the items function equivalently for men and women by indicating that the Rasch item location parameters are equivalent for the two groups. Following the demonstration of construct and measurement equivalence we aimed to test the

strain and buffer hypotheses of the JDC model with regard to general work stress and to investigate whether a common or separate regression equations should be used for men and women. We tested the following substantive hypotheses:
Hypothesis 1 (Gender hypothesis). Gender moderates the relations of job demands, job control, and their product-term with general work stress.
Hypothesis 2 (Buffer hypothesis). Job control moderates the effect of job demands on general work stress.
Hypothesis 3 (Strain hypothesis). Job demands and job control combine additively to explain general work stress.

METHOD

Participants and procedure
The participants were 481 employees (205 men, 276 women) from two post-school academic institutions in the Gauteng area and an MBA programme. The participants represented many different academic faculties, as well as different organisational levels within the respective institutions. All participants held full-time positions in their respective organisations. The average age of the men was 36.71 years (SD = 12.26) and that of the women was 37.99 years (SD = 11.23). All participants were volunteers and the information was used for research only.

Instruments
General work stress. The 9-item General Work Stress Scale of the Sources of Work Stress Inventory (de Bruin & Taylor, 2003) asks questions about the respondents' overall level of work-related stress and was used to operationalise job strain. Sample items are "Do you get so stressed at work that you want to quit?" and "Do you find it difficult to sleep at night because you worry about your work?" Respondents answer on a five-point Likert-type scale where the ordered response options are: Never (1), Rarely (2), Sometimes (3), Often (4), and Always (5). Reliable scores were obtained with the General Stress scale in the present study (Cronbach's α = 0,91).

Job demands. The Workload scale of the Sources of Work Stress Inventory was used to operationalise job demands. The scale consists of seven items where an individual is asked to indicate the degree to which his or her workload is a source of work stress for him or her. Sample items are "Having to take work home at night" and "Receiving work at a faster pace than I can handle". This scale measures the workload and time pressure components of the job demands dimension of the JDC model. The response format is a five-point Likert-type scale, with responses ranging from 1 to 5 in the following order: None at All, Very Little, Some, Quite a Lot, and Very Much. Reliable scores were obtained with the Workload scale in the present study (Cronbach's α = 0,89).

Job control. The Lack of Autonomy scale of the Sources of Work Stress Inventory was used to operationalise job control. The scale consists of seven items that provide an indication of the degree to which an individual feels that he or she is limited in their ability to function autonomously, whether it be due to constraints imposed on them by the actual work environment, or by the nature of their job. The scale measures the decision authority component, which Wall et al. (1996) have described as the central aspect of job control. Sample items are: "Not being consulted on changes at work that affect me" and "Having to do my work according to a rigid set of rules". The response format is identical to that of the Workload scale. Reliable scores were obtained with the Lack of Autonomy scale in the present study (Cronbach's α = 0,90).

Analyses
Factor analysis.
As a first step we examined the convergent and discriminant validities of the items of the three scales by subjecting them to separate maximum likelihood factor analyses for men and women. The number of factors was decided on the basis of theoretical expectations, the scree test and the eigenvalues-greater-than-one criterion. The factors were obliquely rotated to the Direct Oblimin criterion. The similarity of the factor structures for men and women were evaluated by means of coefficients of congruence. A general rule of thumb is that coefficients > 0,90 indicate factor similarity. The fit of the factor model to the empirical data was judged by means of the root mean squared residual (RMSR), for which values below 0,06 are thought to indicate satisfactory fit.

Rasch rating scale analysis.
Standard parametric statistical techniques, such as multiple regression and ANOVA, proceed on the assumption that continuous variables are measured on at least a linear interval scale. However, raw total scores for psychological scales are often non-linear and at best ordinal (Pedhazur, 1997; Wright, 1999). The Rasch model (Rasch, 1960), which is one of a family of item response theory models, can be used to transform ordinal scores to interval measures (Andrich, 1988; Wright & Mok, 2004). In addition, the Rasch model sheds important light on the psychometric integrity of the items that constitute a scale and on the scale itself.

The Rasch rating scale model (Andrich, 1978; Linacre; 2004, Wright & Masters, 1982) is appropriate for the analysis of Likert-type items that share the same set of ordered response categories. The rating scale model specifies that the probability of endorsing a particular category of an item is a function of (a) the person's standing on the latent trait that is measured by the scale, (b) the overall severity or affective intensity of the item, and (c) the category thresholds, which reflect the difficulty in choosing a particular response category rather than the one directly preceding it (Bond & Fox, 2001). For an item set with k ordered response categories, k – 1 category thresholds are estimated. The person and item parameters were estimated with the Winsteps Version 3.56 software package (Linacre, 2005), which uses the joint maximum likelihood estimation method. According to the rating scale model the probability of person n endorsing category j on item i is estimated by the following formula:

$$Pnij\ (x_{nij} = j \mid \theta_n, \beta_i, \delta_i) = \frac{e^{(\theta_n - \beta_i - \delta_i)}}{1 + e^{((\theta_n - \beta_i - \delta_i)}}$$

where Pnij ($x_{nij} = j \mid \theta_n, \beta_i, \delta_{ij}$) is the probability of person n on item i endorsing category j (x = j), given person ability (θ_n), item location (β_i) and the category threshold (δ_i), and e is the natural log function.

On the basis of the estimated person and item parameters expected item scores can be computed for each individual. These expected scores are then compared with the observed scores. If the data fits the model the discrepancy between the model expected scores and the observed scores should be small (R.M. Smith, 2004a; Wright & Masters, 1982). The data-model discrepancies are summarised in the infit mean square statistic. The infit mean square has an expected value of 1, if the data fits the model. Following the guidelines presented by Adams and Khoo (1996), infit mean squares ranging between 0,75 and 1,33 will be taken as indicating satisfactory fit. Items with poor fit may be measuring constructs other than the one of interest and may detract from the measurement quality of the scale. Good fit between the data and the model indicates that a single unidimensional construct underlies responses to the items.

The reliability of the estimated Rasch person measures is expressed as a person separation reliability index, which is similar in meaning and interpretation to Cronbach's alpha coefficient. However in the Rasch person separation reliability index the total variance is based on the estimated person measures rather than the observed raw scores, and the error variance is defined as the average of the squares of the standard errors of measurement for each person (E.V. Smith, 2004).

The Rasch model specifies that the item location parameters should be invariant over different ability or demographic groups. From this perspective differential item functioning (DIF) is present when one or more items have different location parameters (R.M. Smith, 2004b) In this study we employed the DIF analysis routine of the Winsteps Version 3.56 (Linacre, 2005) software package, which compares the location parameters of an item over groups, while holding constant (a) the item locations and thresholds of all other items, and (b) the trait estimates of all individuals. We set the criterion for statistical significance at $p < 0,01$.

Hierarchical multiple regression analyses.
The Rasch person measures were used in all further analyses with regard to the JDC model. The hypotheses were tested with hierarchical multiple regression analysis, following the approach recommended by Pedhazur (1997). In the first step job demands and job control were entered as a block. In the second step the multiplicative term of job demands × job control was entered. In the third step gender (dummy coded: men = 0, women = 1) was entered. In the fourth step the multiplicative terms of gender × job demands, and gender × job control were entered as a block. Finally, in the fifth step the multiplicative term of gender × job demands × job control were entered.

To test the gender hypothesis the increment in the proportion of explained variance in general work stress in steps five, four and three were tested for statistical significance. A statistically significant result for step five would indicate that gender moderates the multiplicative effect of job demands × job control. A statistically significant result for step four would indicate that at least one of the slopes of job demands and job control differs for men and women and that separate regression equations should be used for them. A statistically significant result at step three, which is only tested if a non-significant results is obtained in step four, would indicate that at least one of the intercepts of job demands and job control differ for men and women and that separate regression equations should be used for them.

The buffer hypothesis was tested by testing the increment in the proportion of variance accounted for in general work stress after adding the multiplicative term of job demands × job control to the regression equation (the second step in the hierarchical analysis).

Finally, the strain hypothesis was tested by testing the proportion of variance accounted for in general work stress by the additive effects of job demands and job control (the first step in the hierarchical analysis).

RESULTS

Factor analysis

To supervise the convergent and discriminant validity of the questionnaire items used in the study, and to assess the equivalence of the measured constructs across gender, we subjected the items of the Workload, Lack of Autonomy, and General Work Stress scales to maximum likelihood factor analyses for men and women separately. On the basis of the scree-test, number of eigenvalues greater than unity, *and* theoretical expectations, we extracted three factors in both groups, which were obliquely rotated to the Direct Oblimin criterion. For men and women the root mean squared residual was 0,05, indicating satisfactory fit between the data and the three-factor model. The coefficients of congruence of the corresponding factors for men and women were as follows: General work stress = 0,97, job control = 0,98, and job demands = 0,96. These results indicated very similar patterns of high and low factor pattern coefficients across gender and it appeared safe to conclude that the three factors represented qualitatively similar constructs for men and women.

On the basis of these results the analysis was repeated for the combined group of men and women. The oblique target rotated factor pattern matrix of the combined group is presented in Table 1[1]. Inspection of the factor loadings (factor pattern coefficients) shows that each factor was well defined. The items of the three scales grouped in accord with theoretical expectations: the first factor loaded saliently (> 0,40) on the items of the General Work Stress scale and was labelled *general work stress*, the second factor loaded saliently on the items of the Lack of Autonomy scale and was labelled *job control*, and the third factor loaded saliently on the items of the Workload scale and was labelled *job demands*. Each variable had a salient loading on the factor it was expected to define, but no salient loadings on any other factor. The correlations between the three factors were as follows: General work stress and job control, r = -0,50 for men and r = -0,37 for women; General work stress and job demands, r = 0,53 for men and r = 0,48 for women, and job demands and job control, r = -0,34 for men and women. These correlations suggest that the relation between general work stress and job control differs for men and women, a point that we return to later in the article.

TABLE 1
OBLIQUE ROTATED FACTOR PATTERN MATRIX OF THE WORKLOAD, AURONOMY AND GENERAL WORK STRESS ITEMS

Item		Factor		
		Workload	Autonomy	General Work stress
WL1	No time for hobbies	**0,584**	0,033	0,123
WL2	Work quickly	**0,616**	0,039	0,135
WL3	Take work home	**0,823**	-0,093	0,005
WL4	Work over weekends	**0,627**	0,050	0,000
WL5	Cut back on social life	**0,665**	0,058	0,019
WL6	Too few hours in day	**0,821**	-0,092	0,112
WL7	Receive work at fast pace	**0,755**	0,052	0,089
LA1	Changes happen too slow	0,163	**0,613**	-0,047
LA2	Rigid rules	0,016	**0,745**	-0,032
LA3	Policies and procedures prevent proper work	0,140	**0,730**	-0,051
LA4	Unable to be creative	0,064	**0,746**	0,025
LA5	Others make decisions about me	0,099	**0,747**	-0,034
LA6	Not consulted on changes that affect me	0,121	**0,671**	0,035
LA7	Ask permission before doing anything	0,123	**0,740**	-0,089
GS1	Wish for different job	-0,106	0,269	**0,759**
GS2	Want to quit	-0,074	0,286	**0,769**
GS3	Worry about waking up and going to work	0,007	0,231	**0,706**
GS4	Difficult to sleep at night	0,170	0,084	**0,567**
GS5	So stressed forget to do important tasks	0,207	0,064	**0,536**
GS6	So stressed difficult to concentrate on tasks	0,254	0,058	**0,578**
GS7	Spend a lot of time worrying about work	0,247	-0,024	**0,464**
GS8	Feel cannot cope with work anymore	0,366	-0,003	**0,544**
GS9	So stressed that you lose your temper	-0,024	0,220	**0,528**

Note. Factor pattern coefficients > 0,40 are printed in boldface.
WL = Workload
LA = Lack of autonomy
GS = General work stress

Overall the findings of the factor analysis supported the convergent and discriminant validity of the items of the three scales and empirically indicated that job demands, job control and general work stress are separate, but correlated constructs. This is important because excessive overlap between job demands and measures of job strain have been held responsible for failures to find support for the buffer hypothesis (Van der Doef & Maes, 1999; Wall et al., 1996). Importantly, the results also show that qualitatively similar constructs were measured for men and women.

Item response theory analysis

To further investigate the psychometric quality of the three scales and to establish measurement equivalence for men and women, we subjected the items of each scale to item response theory analyses. Specifically, the fit between the items and the requirements of the Rasch rating scale model were examined. For the General Work Stress scale the mean of the infit mean squares was 0,99 (SD = 0,15), which is close to the expected value of 1,00 (see Table 2). The infit mean squares for the individual items ranged from 0,78 for item GS6 to 1,22 for item GS7, showing that all the items demonstrated satisfactory fit. Moreover, all the items had strong item-measure correlations, ranging from 0,66 for item GS9 to 0,80 for item GS2. The person separation reliability of the General Work Stress scale was 0,89, which might be described as satisfactory.

Table 2
Item location parameters and infit statistics for the general work stress, workload and lack of autonomy scales

Item label	Item location parameter	Standard error	Infit mean square	Infit *t*	Item-measure correlation
		General work stress			
GS7	-0,55	0,06	1,22	3,3	0,69
GS9	0,05	0,07	1,17	2,6	0,66
GS4	0,03	0,07	1,12	1.8	0,74
GS8	0,11	0,07	1,05	0,8	0,74
GS3	0,05	0,07	0,98	-0,3	0,76
GS5	0,46	0,07	0,93	-1,1	0,70
GS1	-0,60	0,06	0,89	-1,8	0,78
GS2	0,11	0,07	0,79	-3,5	0,80
GS6	0,34	0,07	0,78	-3,7	0,76
Mean	0,00	0,06	0,99	-0,2	
SD	0,36	0,00	0,15	2,4	
		Workload (Job demands)			
WL1	0,02	0,05	1,23	3,3	0,71
WL4	-0,16	0,05	1,20	3,0	0,72
WL2	-0,33	0,05	1,08	1,3	0,72
WL5	0,18	0,05	0,98	-0,3	0,75
WL3	0,47	0,05	0,94	-0,8	0,76
WL7	0,07	0,05	0,83	-2.6	0,77
WL6	-0,25	0,05	0,76	-3,9	0,79
Mean	0,00	0,05	0,99	0,0	
SD	0,25	0,00	0,16	2,5	
		Lack of Autonomy (Job control)			
LA1	-0,10	0,06	1.12	1.7	0,73
LA7	-0,22	0,06	0,98	-0,2	0,75
LA6	-0,50	0,06	1.07	1.0	0,76
LA2	0,27	0,06	0,99	-0,1	0,75
LA4	0,30	0,06	0,95	-0,7	0,77
LA5	-0,09	0,06	0,91	-1,3	0,78
LA3	0,33	0,06	0,94	-0,9	0,77
Mean	0,00	0,06	0,00	-0,1	
SD	0,29	0,00	0,07	1,0	

The mean of the infit mean squares for the Workload scale (job demands) was 0,99 (SD = 0,16), which is close to the expected value of 1.00, indicating satisfactory overall fit (see Table 2). The infit mean squares of the individual items ranged from 0,76 for item WL6 to 1,23 for item WL1, showing that all the items demonstrated satisfactory fit. All the item-measure correlations were strong and ranged from 0,71 for item WL1 to 0,79 for item WL8. The person separation reliability of the Workload scale was 0,84, which might be described as satisfactory.

The mean of the infit mean squares for the Lack of Autonomy scale (job control) was 0,99 (SD = 0,07), which is very close to the expected value of 1,00, The infit mean squares of the items ranged from 0,94 for item LA3 to 1.12 for item LA1, showing that all the items demonstrated satisfactory fit (see Table 2). All the item-measure correlations were strong and ranged from 0,73 for item LA1 to 0,78 for item LA5. The person separation reliability was 0,86, which might be described as satisfactory.

Differential item functioning

We investigated DIF by comparing the item location parameters of the three scales for men and women. Two items from the General Work Stress scale showed statistically significant DIF, namely items GEN1 [DIF contrast = -0,41, t(462) = -3,19, p = 0,0015] and GEN8 [DIF contrast = 0,40, t(458) = 3,00, p = 0,0029]. Men found item GEN1 relatively easier to agree with, whereas women found item GEN8 relatively easier to agree with. Ordinal logistic regression showed that conditional on trait level, uniform DIF with regard to gender explained approximately 0,95% of the variance in item GEN1 and 1.07% of the variance in item GEN8.

Two items from the Workload scale (job demands) showed statistically significant DIF, namely items WL4 [DIF contrast = -0,50, t(438) = -4.75, p < 0,0001] and WL5 [DIF contrast = -0,40, t(438) = -3,76, p = 0,0002]. Men found both items relatively easier to agree with. Ordinal logistic regression showed that conditional on trait level, uniform DIF with regard to gender explained approximately 1.20% of the variance in item WL47 and 0,67% in item WL48. Only one item from the Lack of Autonomy scale (job control), namely LA1, showed statistically significant DIF [DIF contrast = 0,33, t(447) = 2,83, p = 0,0049]. Item LA1 was relatively easier to agree with for women than for men. Ordinal logistic regression showed that conditional on trait level, uniform DIF with regard to gender explained approximately 0,76% of the variance in item LA1.

Overall, the results of the DIF analyses showed that across the three scales five items (approximately 22%) showed statistically significant, but practically unsubstantial DIF. On the basis of these results and that of the factor analysis it seemed safe to assume that each of the three scales could be considered sufficiently unidimensional and internally consistent to justify the computation of a total score or person measure for each participant. The results also showed that the scales measured the same traits and functioned equivalently for men and women. The person separation reliabilities indicated that each scale reliably separated individuals with different trait levels. Linear person measures (in logits) were obtained for all three scales and these person measures were used in all further analyses (see Table 3).

The means of the men and women with regard to the linear combination of general work stress, job demands and job control were compared by means of a discriminant analysis, which showed no statistically significant multivariate differences [Wilk's Λ = 0,995, $\chi^2(3)$ = 2.510, p = 0,473]. Univariate analysis of variance also showed no statistically significant mean differences between men and women with regard to the three variables. The zero-order correlations for the men and women combined of the general work stress, job demands and job control person measures were as follows: job demands and job control, r = -0,40 (p < 0,001), general work stress and job demands, r = 0,53 (p < 0,001), and general work stress and job control, r = -0,46 (p < 0,001).

Table 3
Means and standard deviations of the Rasch person measures for men and women

Scale	Men		Women	
	Mean	SD	Mean	SD
General work stress	-1,382	1,850	-1,220	1,591
Workload (Job demands)	-0,557	1,578	-0,327	1,624
Lack of Autonomy (Job control)	-0,541	1,687	-0,497	1,683

Note. The means and standard deviations are expressed in logits.

Moderated hierarchical multiple regression analysis
The results of the moderated hierarchical regression analysis are summarised in Table 4. The test of hypothesis one, the gender hypothesis, showed that the addition of the multiplicative term of gender × job demands × job control in step five of the hierarchical analysis was statistically non-significant and that this term accounted for no additional variance in general work stress [ΔR^2 = 0,00, F(1, 467) = 0,025, p = 0,874]. However, in step four the addition of gender × job demands, and gender × job control produced a small, but statistically significant increment in the proportion of explained variance in general work stress [ΔR^2 = 0,01, F(2, 468) = 3.879, p = 0,021]. Inspection of the standardised partial regression coefficients showed that the coefficient associated with gender × job control was statistically significant [β = -0,163, t = -2.599, p = 0,010], but not the coefficient associated with gender × job demands [β = 0,008, t = 0,126, p = 0,900]. The semi-partial correlation for gender × job control suggested that it uniquely accounted for a small amount of variance in general work stress (r_{sp} = 0,095), whereas the semi-partial correlation for gender × job demands showed that it uniquely accounted for virtually no variance in general work stress (r_{sp} = 0,005). Nonetheless, on the basis of the statistically significant finding, hypotheses two and three were tested separately for men and women. The results of these analyses are summarised in Table 4.

Table 4 shows that the results did not support hypothesis two (the buffer hypothesis). For men the addition of job demands × job control to the regression equation in step two led to a statistically non-significant and very small increment in the proportion of explained variance in general work stress [ΔR^2 = 0,00, F(1, 198) = 0,468, p = 0,495]. Similar results were obtained for the women [ΔR^2 = 0,00, F(1, 269) = 0,839, p = 0,361].

In contrast, there was strong support for hypothesis three (the strain hypothesis). For men the additive effects of job demands and job control accounted for approximately 40% of the variance in general work stress [R^2 = 0,403, F(2, 199) = 67.162, p < 0,001], whereas for women the additive effects accounted for approximately 34% of the variance [R^2 = 0,337, F(2, 270) = 68.677, p < 0,001]. These results suggest that the JDC model did a slightly better job in explaining the variance in general work stress for men than for women. Inspection of the zero-order and semi-partial correlation coefficients in step one shows that job control had a stronger relationship with general work stress for men [r = -0,538, r_{sp} = -0,357], than for women [r = -0,413, r_{sp} = -0,215]. To shed further light on this issue, 90% confidence intervals were constructed around the R^2 point-estimates of the men and women. For men the lower limit was 0,307 and the higher limit was 0,484, whereas for women the lower limit was 0,256 and the higher limit was 0,410, These results show that the 90% confidence intervals of the two groups show considerable, but not complete overlap. It is noteworthy that the confidence interval of the women included the R^2 point-estimate of the men, and vice versa.

Table 4
Hierarchical multiple regression analysis for general work stress for the total group

Variables entered	Steps				
	1	2	3	4	5
	Combined group				
1. Job demands	0,375*				
Job control	-0,280*				
2. Job demands × job control		-0,029			
3. Gender			0,000		
4. Gender × job demands				0,005	
Gender × job control				0,095*	
5. Gender × job demands × job control					-0,006
R^2	0,36*	0,36*	0,36*	0,37*	0,37*
ΔR^2	0,36*	0,00	0,00	0,01*	0,00
	Women				
1. Job demands	0,408*				
Job control	-0,215*				
2. Job demands × job control	-0,045				
R^2	0,34*	0,34*			
ΔR^2	0,34*	0,00			
	Men				
1. Job demands	0,337*				
Job control	-0,357*				
2. Job demands × job control	-0,038				
R^2	0,40*	0,40*			
ΔR^2	0,40*	0,00			

Note. The coefficients associated with each variable are semi-partial correlations at that step of the hierarchical analysis.
ΔR^2 = change in the squared multiple correlation.
* *p* < 0,05

DISCUSSION

This study had two broad aims. The first was to test the strain and buffer hypotheses associated with Karasek's (1979) JDC model. Both hypotheses predict that the highest job strain is experienced in environments characterised by high job demands and low job control. However, they differ in that the strain hypothesis predicts that job demands and job control have additive effects, whereas the buffer hypothesis predicts that job demands and job control have a multiplicative effect and that high job control can ameliorate the negative effects of high job demands. In testing these two hypotheses researchers routinely control for gender by including it as a covariate in the first step of hierarchical regression analyses. This step, however, proceeds on the assumption that the measures used in the study are perceived in the same way by men and women and that a common regression equation can be used for men and women. We believe, however, that measurement equivalence and the use of common versus separate regression equations are empirical matters. Therefore, the second aim of this study was to investigate the questions of measurement equivalence and the use of a common versus separate regression equations for men and women. We start the discussion with reference to the second research question.

The results of the factor and item response theory analyses showed that the psychometric properties of the general work stress, job demands and job control measures were satisfactory overall and appeared equivalent for the men and women. The factor analysis indicated that the three scales measure qualitatively similar constructs for men and women and that for both groups each item served as an indicator only of the factor to which it is assigned by the scoring key. The factor loadings associated with all items were strong, suggesting that each item was saturated with the corresponding constructs. The results also supported the construct validity of the three scales by showing that each item demonstrated convergent and discriminant validity. On the basis of these results it was concluded that the three factors represented theoretically *and* empirically distinctive, but correlated constructs. This is important against the background of criticisms that overlap in content and shared affective judgement in responding to items of job demands and job strain, leads to spurious correlations between these two variables (Wall et al., 1996).

The results of the item response theory analyses confirmed that for each of the three scales a single line of enquiry runs through the items and that it is appropriate to combine the items to obtain a single score or measure. The DIF analysis showed that five of the 23 items were slightly biased, but from a practical measurement perspective the impact of the bias was minimal. The reliabilities of the person measures obtained with the scales were shown to be highly satisfactory. On the basis of these results we concluded that men and women perceived the general work stress, job demands and job control measures in the same way and that comparable measures were obtained for the two groups.

Although measurement equivalence was obtained, the results of this study showed that from a statistical perspective the assumption of a common regression equation for men and women did not hold. This result runs counter to the common practice of controlling for gender differences in the first step of a hierarchical multiple regression analysis. Specifically, our results showed that job control had statistically significant different slopes for men and women. Comparison of the zero-order correlations, semi-partial correlations, and regression weights showed that job control is more strongly related to general work stress for men than for women. However, although statistically significant, the inclusion of gender as a moderator led to a very modest increment in the proportion of variance accounted for (approximately 1%). These results are consistent with that of Vermeulen and Mustard (2004) and Griffin, Fuhrer, Stansfeld and Marmot (2002) who reported that men respond more adversely to low job control than women. A possible reason for this is that men's identity and self-worth is more closely tied to the work place. In this regard men are more likely to value their "organisational roles as a central life interest" (Dodd-McCue & Wright, 1996, p. 1086) and that men may be more likely to interpret a lack of control in the work place as a threat to their general self-worth.

With regard to our first research aim our results provided no support for the buffer hypothesis for men or women. The results showed that the multiplicative model did not fit the data any better than the additive model. The product term of job demands and job control was not statistically significant and accounted for less than 1% of additional variance in general work stress, after the additive effects of job demands and job control had been partialled out. Hence, the evidence obtained in this study is consistent with a large body of research that did not provide support for the idea that job control moderates the effect of job demands on psychological well-being.

However, our results did provide strong support for the strain hypothesis. The additive effects of job demands and job control accounted for slightly more variance in general work stress for men (approximately 40%) than for women (approximately 34%). This is probably due to the stronger relationship of job control with psychological well-being for men as explained in an earlier paragraph. One potentially important implication of this finding is that the use of a common regression equation may slightly underestimate the strength of the additive effects of job demands and job control for men, and slightly overestimate the effects for women.

In summary, this study provides strong support for the strain hypothesis of the JDC model for men and women, but no support for the buffer hypothesis. In this respect the results of this study are consistent with the majority of studies that examined the same hypotheses. The study contributes in that it is one of very few studies that explicitly addressed the measurement equivalence of the concepts of the JDC model for men and women. The study also contributed in that it showed that, contrary to common practice, separate regression equations had to be used for men and women. This may not be true in other investigations into the strain and buffer hypotheses, but it remains an empirical question that has to be explicitly tested.

This study has limitations, of which three are highlighted. In the first place the participants were volunteers from only two higher education institutions and it is not clear that the results will generalise to other individuals in these organisations or to individuals from other organisations. However, the overall similarity of the results obtained in this study with those of many other studies suggests that the findings may be of general importance. Secondly, this study made use of self-reports with regard to job demands, job control and general work stress. Previous researchers have pointed out that self-reports may lead to inflated correlations between job demands, job control and psychological well-being, which may lead to an overestimation of the main effects of job control and job demands at the expense of the likelihood of observing a significant multiplicative effect (Van der Doef, Maes & Diekstra, 2000; Wall et al. 1996). The factor analysis of the three scales used in this study empirically showed that they represented separate constructs, but the problem of shared method variance remains.

Thirdly, the use of gender as an independent variable is problematic due to its correlations with a vast range of other variables that may influence or be correlated with general work stress, job control or job demands. Hence, no causal statements with regard to the role of gender can be made. For instance, it can be safely assumed that gender is related to occupational status, with women typically enjoying lower status than men (although there increasingly are women who enjoy higher occupational status than men, one might generally still expect that men hold jobs with higher status). It may well be that the observed differences between men and women in this study partially or fully reflect differences in occupational status, which in turn might be influenced by or be correlated with a vast range of other confounding variables. Like other researchers (cf. Barnett & Brennan, 1997), we might have controlled statistically for confounding variables such as occupational status and income. However, this would have informed us about a potential world where men and women have equal occupational status – a world that does not yet exist. Moreover, it is impossible to control for all confounding variables. In this regard we concur with methodologists such as Lord (1969), Pedhazur (1997), Reichardt (1979), and Wolins (1982) who hold the opinion that controlling for confounding variables in the absence of random assignment to groups is an almost impossible task and that attempts to do so might provide misleading results. In short, our results showed that different regression equations were necessary for men and women in this study. This may not be true in other studies.

In closing we restate our belief that when different demographic groups, such as men and women or different cultures, are included in studies of the JDC model, the psychometric equivalence of the measures for the groups should be empirically tested. Only when it is shown that the psychometric properties are equivalent for the groups, can one be sure that the scales, inventories or questionnaires are perceived in the same way, and only then are comparisons or combinations of the data of the groups meaningful. In addition to psychometric equivalence, it is also prudent to investigate whether common versus separate regression equations should be used for the relevant groups. Failure to do so may lead to overestimation of the effects of job demands and job control on psychological and health outcomes for some groups, and underestimation for others. We hope that in future studies such tests of psychometric equivalence and the equivalence of regression equations will become routine, not only with regard to gender, but also with other potentially important demographic variables such as ethnicity or culture.

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