Personal Relationships, **15** (2008), 533–550. Printed in the United States of America.

Women's weekly relationship functioning and depressive symptoms

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Abstract
In an exploration of the links between relationship quality and depression, the extent to which women's weekly reports of depressive symptoms vary as a function of same-week relationship functioning was tested. A sample of 161 married or cohabiting U.S. women completed measures of relationship functioning, mood, and depressive symptoms weekly for 12 weeks. In a series of hierarchical linear models, results of within-subject analyses indicated that depressive symptoms were negatively associated with same-week relationship functioning. Weekly fluctuations in mood did not account for these associations. Results of between-subjects analyses suggested that women low in stereotypical masculinity and in relationships of shorter duration are particularly likely to show increased depressive symptoms during weeks when they experience poorer relationship functioning than usual.

There is a well-established cross-sectional association between marital distress and depression (Proulx, Helms, & Buehler, 2007; Whisman, 2001) and growing evidence that relationship discord prospectively predicts poorer emotional well-being (Proulx et al., 2007) and onset of major depression (Whisman & Bruce, 1999). Women's psychological well-being, in particular, appears to be strongly linked to the current functioning of their intimate partnerships (Proulx et al., 2007). Building on these findings, researchers have suggested that changes in relationship functioning within an individual may be associated with corresponding changes in that individual's depressive symptoms. For example, stress-generation models of depression (Davila, Bradbury, Cohan, & Tochluk, 1997) posit that increased marital stress leads to increased depressive symptoms, which lead to further declines in relationship functioning, and so on in a negative cycle. Similarly, the marital discord model of depression (Beach, Sandeen, & O'Leary, 1990) proposes that marital distress heightens risk for depression, while alleviation of marital discord may promote recovery from depressive episodes. In fact, based on the notion that improvements in relationship quality may predict corresponding improvements in depressive symptoms, some researchers have proposed couples therapy as a treatment for depression in maritally distressed individuals, most often wives (e.g., Beach & Jones, 2002).

Despite the growing body of theory, the association between changes in relationship quality and depression within individuals, on which these theories and interventions are

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This research was supported in part by grants from the National Institute of Mental Health to the first author (F32MH077434) and by grants from the National Institute of Child Health and Human Development to the second (R01HD047564) and third authors (R01HD053314). We gratefully acknowledge Galena Rhoades and Dave Atkins, who provided valuable advice and assistance with statistical analyses, as well as Ben Karney, for his generous counsel on the design and analysis of multiwave data sets in early stages of this project.

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based, has rarely and only recently been directly tested (cf. Davila, Karney, Hall, & Bradbury, 2003). The vast majority of studies supporting a link between relationship functioning and depression are cross-sectional (see Whisman, 2001) and, therefore, do not address changes in either variable. Moreover, longitudinal studies have primarily examined between-subject associations, which demonstrate that adults with better initial relationship functioning tend to have lower depressive symptom levels at a later time point than do those with poor initial relationship functioning. These findings do not tell us whether the increases in relationship quality individuals experience (i.e., within-person changes) are associated with corresponding within-person decreases in depressive symptoms. Without direct evidence of a within-individual association between relationship satisfaction and depressive symptoms, we lack a strong research-based rationale for offering couples interventions to treat depression. The field clearly needs more information regarding the extent to which changes in relationship quality are reliably associated with corresponding changes in depressive symptoms.

Noting the deficits in the literature, researchers have recently begun to use methodological and analytic strategies that allow for the direct estimation of the association between within-person changes in relationship quality and depressive symptoms (e.g., hierarchical linear modeling; Bryk & Raudenbush, 1987). In this method, each variable is measured at multiple time points for each study participant. Hierarchical linear models are used to estimate change in relationship quality and change in depressive symptoms over time for each individual and to calculate the within-subject association between the changes in the two variables. Findings from three studies using this methodology have supported the hypothesized associations, showing that decreases in marital quality tend to co-occur with increases in severity of depressive symptoms (Davila et al., 2003; Karney, 2001; Kurdek, 1998). These studies have significantly advanced our knowledge, providing the first direct evidence of a within-subject association between relationship functioning and depressive symptoms.

Nonetheless, all three studies collected data at the relatively long intervals of 6 months to 1 year, limiting conclusions about the extent to which more rapid or short-term changes in relationship functioning and depressive symptoms are associated. Although there is evidence that perceptions of marital satisfaction can fluctuate within individuals over intervals as short as 1 week (McNulty & Karney, 2001), we do not know whether fairly short-term fluctuations in couple functioning are associated with corresponding fluctuations in depressive symptoms. This gap in our knowledge clearly limits applicability of findings to clinical interventions. Couples therapies to treat co-occurring depression and relationship distress tend to include 15–20 sessions that occur on a weekly basis (e.g., Jacobson, Dobson, Fruzzetti, Schmaling, & Salusky, 1991). Therefore, studies relying on assessments spaced 6 months apart lack the sensitivity to capture either the fairly rapid changes in relationship satisfaction likely to occur during couples therapy or their associations with changes in depressive symptoms. In addition, lengthy intervals between assessments can render it difficult to control for intervening events and often require individuals to report on events and perceptions that occurred some time ago, introducing problems associated with retrospective recall. Together, these limitations suggest the potential utility of measuring relationship functioning and depressive symptoms more frequently, which would allow us to capture links between short-term fluctuations in the key variables (thereby increasing the clinical applicability of findings), limit the influence of intervening events, and allow participants to report on relationship events and symptoms of depression closer to the time they occur.

Accordingly, the first purpose of the present study was to test for the presence of within-subject associations between short-term fluctuations in relationship functioning and depressive symptoms. In a sample of women in cohabiting or married relationships, we assessed relationship functioning (operationalized as levels of perceived satisfaction with the relationship and amount of couple conflict) and depressive symptoms each week for

12 weeks and estimated the within-subject associations between weekly reports of relationship functioning and weekly reports of depressive symptoms. Although relationship functioning may also influence men's mental health, we chose to study these phenomena in women because of the more central role that relationships appear to play in women's psychological functioning. Women tend to be at greater risk for depressive reactions to interpersonal difficulties than men (Kendler, Thornton, & Prescott, 2001; Proulx et al., 2007), which may contribute to their rates of major depression and subdiagnostic depressive symptoms being approximately twice as high as men's (Kessler, 2003; Kessler, Zhao, Blazer, & Swartz, 1997). It is particularly important, then, to understand in more detail the links between relationship functioning and depression among women and to identify protective factors that may buffer women from depressive reactions to relationship stress.

Although we are aware of no previous studies that examine the short-term within-subject associations between relationship quality and depression, several previous findings are relevant. First, there is evidence that other psychosocial variables predict daily fluctuations in depressive symptoms (e.g., negative cognitive responses to daily stressors; Hankin, Fraley, & Abela, 2005). Further, Thompson and Bolger (1999) demonstrated that individuals' daily depressive symptoms and their partner's daily feelings about the relationship covaried over the course of 1 month; however, these authors did not test within-person associations. Together, these studies suggest that fairly rapid fluctuations in depressive symptoms may covary with fluctuations in psychosocial factors such as relationship functioning. Building upon these findings, as well as the studies documenting within-subject associations between relationship quality and depression when measured at 6-month to 1-year intervals (Davila et al., 2003; Karney, 2001; Kurdek, 1998), we hypothesized that women's depressive symptoms, measured each week, would vary as a function of same-week relationship functioning.

Although the frequent, repeated measurement of key variables was advantageous in many ways, it did introduce the possibility that variability in individual's current mood state from week to week might account for the week-to-week variation in individuals' self-reported relationship functioning and depressive symptoms, resulting in a spurious association between the two variables. To address this possibility, we also measured participants' mood each week and reran the within-subject analyses including weekly mood as a control variable. Although we expected self-reports of both relationship functioning and depressive symptoms to be associated with participants' current mood state, we predicted that fluctuations in participants' mood would not account for the week-to-week associations between the relationship functioning variables and the depressive symptoms.

We designed this study to estimate the simultaneous covariation between relationship functioning and depressive symptoms; we did not intend to assess directionality of effects or causal relations between the variables. Nonetheless, because our understanding of how couple relationships and depression interrelate would be greatly augmented if we were able to shed light on whether changes in relationship functioning generally precede changes in depressive symptoms or if changes in depressive symptoms tend to precede changes in relationship functioning, we also explored temporal relations between variables. We conducted exploratory time-lagged associations, in which the previous week's relationship functioning was used to predict depressive symptoms and in which the previous week's depressive symptoms were used to predict the relationship functioning variables.

As we seek to understand the links between women's intimate partnerships and depression, it is crucial to identify moderators of the within-subject associations between relationship functioning and depressive symptoms. Individual and relationship characteristics that are associated with stronger within-subject associations may represent risk factors because they predict greater vulnerability to depression at times when relationship quality declines and greater risk of relationship distress when depressive symptoms become elevated. Conversely, characteristics associated with a weaker within-subject association may buffer individuals from depressive reactions to relationship dissatisfaction and conflict. Unfortunately, research on moderators of the links between relationship quality and depression is extremely limited (cf. Whisman, 2001), leaving us with little knowledge regarding the characteristics that may serve as risk or protective factors. The few studies to date assessing within-subject relations have only assessed the moderating effect of one personality variable, neuroticism, demonstrating that newlyweds with higher neuroticism tend to show greater increases in depressive symptom levels at times when their marital satisfaction declines than do those low in neuroticism (Davila et al., 2003; Karney, 2001). Therefore, the second purpose of this study was to investigate the potential moderating effects of two theoretically indicated personality traits (feminine and masculine gender roles) and relationship characteristics (marital status and relationship length) on the week-to-week associations between women's relationship functioning and depressive symptoms.

Gender roles, defined as individuals' levels of psychological femininity and masculinity, represent particularly plausible moderators of the links between close relationship functioning and depression. The high rates of depression in women (Kessler, 2003) may result from the more central role of interpersonal relationships in women's than men's lives, identities, and self-concepts (e.g., Bilsker, Schiedel, & Marcia, 1988; Gilligan, 1982), which leaves them more vulnerable to depression in response to relationship distress (e.g., Jack, 1991). The extent to which women value close relationships and base their identities in their relationships, however, varies widely across individuals and may be better captured by their degree of psychological femininity than their biological sex. Psychological femininity is a cluster of traits characterized by a communal orientation, interpersonal warmth, and concern with the maintenance of interpersonal relations (e.g., Bem, 1974). Women show large individual differences in femininity (Helmreich, Spence, & Wilhelm, 1981) and those with higher levels show a greater tendency to prioritize their romantic relationships over personal goals (Hammersla & Frease-McMahan, 1990). Because highly feminine women tend to be relationship focused and to derive self-worth from their relationships (Helmreich et al., 1981), they may be at risk for dysphoria or depression in times of relationship stress or dissatisfaction (Waelde, Silvern, & Hodges, 1994). Consistent with this notion, cognitive and psychodynamic theorists have long proposed that individuals who base their self-worth on the quality of their relationships are vulnerable to depression in the face of interpersonal loss, conflict, or rejection (Beck, Rush, Shaw, & Emery, 1979; Blatt, D'Affitti, & Quinlan, 1976). Thus, we hypothesized that women who report high psychological femininity would show stronger week-to-week associations between relationship functioning and depressive symptoms than would those low in femininity.

Women's levels of stereotypically masculine characteristics may also influence the extent to which their weekly depressive symptoms vary as a function of their same-week relationship functioning. Having an agentic orientation, engaging in instrumental behaviors, high self-efficacy, and greater assertiveness characterize psychological masculinity, an essentially orthogonal construct from femininity (Bem, 1974; Helmreich et al., 1981). We propose that masculinity may be a protective factor against depressive reactivity to relationship distress. Masculinity predicts less depression and dysphoria in the face of various types of life stress (Brazelton, Greene, & Gynther, 1996; Nezu, Nezu, & Peterson, 1986), perhaps because highly masculine individuals tend to engage in more active and problem-focused (vs. emotion-focused and ruminative) coping strategies in response to stressful situations (Conway, Giannopoulos, & Stiefenhofer, 1990; Lengua & Stormshak, 2000). These characteristic responses are likely to decrease risk for depression according to behavioral, cognitive, and gender-based theories of depression (e.g., Abramson, Seligman, & Teasdale, 1978; Beck et al., 1979; Nolen-Hoeksema, 1987). Because women with higher levels of stereotypically masculine characteristics are likely to feel more efficacious and to use active coping strategies when faced with reduced relationship quality, they may be less likely to show depressive reactions to that stress. Thus, we hypothesized that higher masculinity would be associated with a weaker within-subject association between weekly relationship functioning and depressive symptoms.

We also explored whether the relationship characteristics of marital status and relationship length moderate the week-to-week associations between couple functioning and depressive symptoms. To date, studies document within-subject associations only in newlywed couples followed through the first years of marriage (Davila et al., 2003; Karney, 2001; Kurdek, 1998), limiting generalizability of findings to couples in longer term marriages and in other types of relationships. Determining whether relationship functioning and depressive symptoms covary within individuals in nonmarried cohabiting relationships is particularly important given the large proportion of women who cohabit (Bumpass & Lu, 2000) and evidence that cohabiters are at higher risk than married couples for relationship distress (Stanley, Whitton, & Markman, 2004). Therefore, we tested hypotheses in a sample of both cohabiting and married women with a broad range of relationship lengths and assessed whether the week-to-week association between relationship functioning and depressive symptoms differed by marital status or relationship length. Due to lack of existing data, we made no directional hypothesis about the moderating effect of marital status. Based on meta-analytic findings that the cross-sectional association between marital quality and emotional well-being was stronger in marriages of 8 years or less than in longer marriages (Proulx et al., 2007), we tentatively hypothesized that longer relationship duration would be linked with a weaker within-subject association.

Method

Participants and procedure

Participants were women currently in a cohabiting or married relationship who responded to U.S.-based advertisements for participants to take part in a 12-week Internet-based study on relationships and individual well-being. We recruited them primarily through newspaper and electronic newsletter advertisements (46%) and an advertisement posted on the Web site for a relationship-enhancement program (14%). The remainder of participants heard about the study through forwarded e-mails from friends or colleagues (20%) or searching the Internet (6%). Fourteen percent did not provide recruitment information. Recruitment source was not associated with differences in any key variables or associations. Participants agreed to provide information about their relationships, behaviors, and feelings each week for 12 weeks using a password-protected Web site. Each week, they received e-mail reminders to complete the questionnaires out of sight of their partner. Data were time and date stamped; we excluded any data provided more than 2 days from the expected date. At completion of Time 12, we e-mailed participants a brief study description and instructions for obtaining information regarding study results. Compensation for participation was entry into a lottery for US$250; participants received one entry for each completed time point and 10 additional entries for completing all 12 weeks.

One hundred eighty-two women participated in the study.[1] We excluded 9 women in same-sex relationships from present analyses due to the small cell size, along with 8 women with insufficient data and 4 multivariate outliers (detected by large mahalanobis distances in multivariate analyses). This created a final sample of 161 women in heterosexual cohabiting or married relationships who provided 1,396 weeks of data. Each participant provided a median number of 10 (range = 2–12) data points; 21% of the sample ($n = 34$) completed all 12 data points. Analyses (described below) used all available data from each of the 161 women.

The average age of participants was 32.60 years (range = 20–59 years). Participants were

1. Thirty-one men also completed measures; however, because this number was low and we were interested in examining hypothesized relations in women, we did not include men in analyses.

primarily Caucasian (85% Caucasian, 3% Hispanic, 3% African American, 2% Asian American, 4% multicultural, and 2% Native American). Thirty-seven percent of the women were in a dating, cohabiting relationship, and 63% were married. Mean relationship length was 7.8 years but varied widely across participants (range = 3 months to 37 years; 50% of women's relationships were longer than 5 years and 25% were longer than 10 years). Married participants had been in their relationships significantly longer (M = 10.24 years) than had cohabiting participants (M = 3.53 years), $t(159) = 6.44$, $p < .001$. Median income was in the range of US$30,000–39,000. The sample was highly educated; all participants had completed high school, mean years of education was 17 (1 year of graduate school), 70% graduated college, and 40% had a masters level or higher degree. Given 2002 U.S. Census estimates that only about one fourth of the population has a bachelor's degree or higher, our methodology appears to have selected for highly educated participants. These demographics are consistent with evidence that Web-based studies generally obtain diverse but not representative samples (e.g., Best, Krueger, Hubbard, & Smith, 2001) and, in this way, are similar to most convenience samples.

Measures

Participants completed a battery of questionnaires at Time 1 assessing constructs assumed to be relatively stable (i.e., person-level measures), including the proposed moderators. We assessed variables presumed to change from week to week, including relationship satisfaction, couple conflict, several relationship behaviors, and depressive symptoms, weekly for 12 weeks. We present only those measures relevant to the present hypotheses.

Person-level measures

Marital status and relationship length. A basic demographic information form assessed marital status (dummy coded; 0 = *cohabiting*, 1 = *married*) and relationship length. Participants reported no breakups during the study.

Psychological femininity and masculinity. Using the short form of the Personal Attributes Questionnaire (Spence, Helmreich, & Stapp, 1975), participants described themselves on a 5-point bipolar adjective scale for each of 24 characteristics (e.g., 1 = *not at all competitive*, 5 = *very competitive*). The Femininity scale contains characteristics women stereotypically possess to a greater degree than men (e.g., emotional expressiveness, relationship focus). The Masculinity scale contains characteristics males stereotypically possess more than females (e.g., instrumentality, assertiveness). Both scales are fairly stable over time, internally consistent (femininity α = .85 and masculinity α = .82; Spence et al., 1975; α is a measure of a scale's reliability) and have demonstrated construct validity, correlating with other measures of sex-role characteristics (e.g., Holmbeck & Bale, 1988). In this sample, internal consistency was somewhat lower (αs = .71 for femininity and .67 for masculinity).

See Table 1 for descriptive statistics for each of the person-level characteristics and

Table 1. *Descriptive statistics and correlations of person-level variables*

Variable	*M*	*SD*	%	1	2	3
1. Married[a]			64			
2. Relationship length (years)	7.83	7.11		.46**		
3. Femininity	24.90	3.11		.02	.01	
4. Masculinity	20.54	3.59		.07	.07	.11

[a]Dummy coded (0 = *cohabiting*, 1 = *married*).
**$p < .01$.

the correlations among them. As expected, marital status correlated moderately with relationship length, but no other correlations between person-level variables were significant. Consistent with previous research (e.g., Spence et al., 1975), masculinity and femininity demonstrated a small positive correlation that was not statistically significant.

Weekly measures

Depressive symptomatology. We assessed depressive symptoms using two versions of the Center for Epidemiological Studies-Depression Scale (CESD; Radloff, 1977). At Time 1, participants completed the original 20-item measure, rating how often they experienced a variety of depressive symptoms during the past week (e.g., "I felt sad," "I felt bothered by things that don't usually bother me," "I could not get going," "I did not feel like eating") on a 4-point scale (0 = *rarely or none of the time*, 3 = *most or all of the time*). This well-validated measure has high levels of internal consistency (e.g., Eaton & Kessler, 1981). Although the full 20-item measure was too long to be used for the frequent repeated measurement, we included it at Time 1 to assess validity of a brief version of the CESD and assess depression level in this sample. Internal consistency was excellent ($\alpha = .90$). Average Time 1 CESD score, calculated as the sum of the 20 items, was 13.47 ($SD = 9.61$). Using an overall score of 16 as a cutoff (e.g., Derogatis, Lynn, & Maruish, 1999), 47 women (29%) reported symptom levels indicative of clinically significant depression.

To reduce the burden of a 20-item measure when given at multiple weekly intervals, we used a 10-item brief version of the CESD at Times 2–12. We chose the 10 items (item numbers 1, 3, 6, 8, 12, 14, 16, 17, 18, and 20 from the original measure) because, in a longitudinal sample of married couples (Stanley et al., 2001), they maximized similarity of factor loadings of items across sex, maximized internal consistency for both sexes (αs = .84 for women and .81 for men), and yielded scores that correlated extremely highly with scores from the original 20-item scale ($r = .92$). In the current sample, the 10-item scale demonstrated high internal consistency (average $\alpha = .88$, range = .86–.91) and correlated strongly with the full CESD (Time 1 $r = .93$). Analyses used the 10-item brief scale for all time points (i.e., we excluded the other 10 items at Time 1) for consistency of measurement.

Relationship functioning. We chose relationship satisfaction and conflict as indicators of relationship functioning because both variables discriminate between couples who are currently distressed versus nondistressed (Crane, Allgood, Larson, & Griffin, 1990; Weiss & Heyman, 1997), prospectively predict long-term marital outcomes (e.g., Clements, Stanley, & Markman, 2004), and relate to the development of depression (e.g., Beach et al., 1990; Coyne, 1976). We used a 12-item measure of *relationship satisfaction* designed for a daily diary study (McNulty & Karney, 2001). Participants rated the degree to which they were satisfied with nine specific aspects of their relationship (e.g., sex life) and three general aspects (e.g., "How satisfied are you with your partner?") during the past week on a 7-point scale (1 = *not at all satisfied*, 7 = *very satisfied*). This instrument demonstrates within-subject variance across time periods as short as 1 day (McNulty & Karney, 2001). Internal consistency was excellent (average $\alpha = .94$, range across waves = .92–.95). We measured *relationship conflict* using the 8-item Negative Conflict scale of the Communication Skills Test (Jenkins & Saiz, 1995). Participants rated the frequency of negative relationship conflict events, including withdrawal, negative conflict, escalation, and invalidation, during the past week on a 7-point scale (1 = *never*, 7 = *most of the time*). Mean response across items represented participants' relationship conflict scores. The subscale has previously shown internal consistency and evidence of validity (Stanley et al., 2001; Whitton et al., 2007). In the current sample, alphas ranged from .90 to .95 across data points (average $\alpha = .92$). Because relationship satisfaction and relationship conflict were strongly associated at each time point (average $r = .70$, range = .62–.79), we created a composite variable representing the average of the participant's

Table 2. *Descriptive statistics for weekly measures*

Variable	Mean across weeks	Between-person variance	Within-person variance
Relationship functioning	5.52	0.93	0.43
Negative mood	1.91	0.37	0.23
Positive mood	3.38	0.40	0.48
Depressive symptoms	6.55	19.47	15.20

satisfaction and (reverse-scored) conflict ratings, which we labeled *relationship functioning*.[2] Scores reflect participants' mean response (1–7) across items; higher scores represent better functioning.

Negative and positive mood. We assessed weekly mood using the Positive and Negative Affect Scales (PANAS; Watson, Clark, & Tellegen, 1988). Participants rated the extent to which they felt each of 10 positive and 10 negative mood states during the past week (1 = *very slightly or not at all*, 5 = *extremely*). The PANAS scales have demonstrated utility in assessing mood fluctuations over the course of days and weeks (Watson, 1988; Watson et al., 1988). In this sample, the Positive Affect Scale and the Negative Affect Scale each had good internal consistency across time points (αs ranged from .89 to .94 and from .89 to .92, respectively).

Table 2 displays descriptive statistics for the weekly measures. Preliminary analyses indicated that each of the weekly measures demonstrated significant between-person variance, as well as significant variance over time within individuals. The within-person variance in depressive symptoms across the 12 weeks appears to have been clinically meaningful; using eight as a clinical cutoff on the brief CESD, the majority (n = 92, 57%) of the sample fluctuated from nonclinical to clinical levels of depressive symptoms, while 22 women (14%) were always in the clinical range and 47 (29%) were always in the subclinical range.

2. Because the two measures used the same 1–7 scale and had similar standard deviations, we did not standardize them prior to creating the composite. Retaining the original scale also aids with interpretation. Analyses run on the satisfaction and conflict measures separately yielded virtually identical results; we do not present these.

Analyses

Due to the multilevel structure of the data, in which repeated weekly measures were nested within individuals, we conducted analyses using multilevel models (also called hierarchical linear models or multilevel random coefficient models). We used the hierarchical linear modeling (HLM) program, Version 6.02 (Raudenbush, Bryk, Cheong, & Congdon, 2000) to examine a two-level model of relationship functioning and depressive symptoms. In Level 1 (within-person) analyses, coefficients were estimated representing the within-person, week-to-week associations between relationship functioning and depressive symptoms. Level 2 (between-person) analyses then tested whether these within-person associations varied as a function of individual differences on person-level characteristics. That is, Level 2 analyses examined whether individual differences on variables such as marital status and gender roles moderated the within-person links between relationship functioning and depressive symptoms.

Masculinity and femininity were standardized before entry into the multilevel analyses in order to facilitate interpretation of their respective moderating effects on the associations between weekly relationship functioning and depressive symptoms. In addition, standardizing continuous person-level variables removes the influence of differences among the variances of the measures on the parameter estimates (Nezlek, 2001).

Results

Tests of week-to-week within-subject associations

Weekly relationship functioning and depressive symptoms

To test for within-subject, week-to-week associations between relationship functioning and depressive symptoms, we specified the following within-person (Level 1) model:[3]

$$Y_{ij} = \pi_{0i} + \pi_{1i}(\text{Relationship functioning})_{ij} + \varepsilon_{ij}$$

where Y_{ij} represents the depression score of person j at week i. In this equation, π_{0i} represents the individual's average level of depressive symptoms across time and π_{1i} represents the individual's within-subject association between relationship functioning and depressive symptoms. Relationship functioning was group mean centered.[4] In this data structure, group is defined as the individual participant; therefore, the coefficient for relationship functioning is an estimate of the association between each woman's weekly deviations from her own average score on relationship functioning and her weekly deviations from her average depressive symptoms score. To test whether the within-person association between weekly relationship functioning and depressive symptoms was significant in this sample, we analyzed the Level 1 coefficients at Level 2 (the person level) using the following model:

$$\pi_{0i} = \beta_{00} + r_{0i}$$
$$\pi_{1i} = \beta_{10} + r_{1i}$$

In the equation predicting π_{0i} (individuals' average depressive symptom level across time), β_{00} represents the average depression score across all subjects, and r_{1i} is a random effect (varies across individuals) that represents individual differences in average depression. In the equation predicting π_{1i} (individuals' within-subject association between relationship functioning and depression), β_{10} is a fixed-effect coefficient that represents the average within-subject association between relationship functioning and depression across subjects, and r_{1i} is a random effect representing individual differences in the within-subject association. The coefficient that represents the mean within-subject association (β_{10}) is shown in Table 3 (see Model 1).

Consistent with Hypothesis 1, there was a significant within-person association between relationship functioning and depressive symptoms. The negative coefficient (-2.56, $p < .001$) indicates that in weeks that women reported lower than usual relationship functioning, they reported more depressive symptoms than usual. The effect size, calculated using the formula $r = \sqrt{t^2/(t^2 + df)}$, was in the large range ($r = -.67$), indicating that 43% of the variance in depressive symptoms from week to week was accounted for by individuals' weekly deviations from their own average relationship functioning score. Within individuals, higher levels of weekly depressive symptoms were significantly associated with lower levels of same-week relationship functioning, such that a weekly 1-point increase in relationship functioning (on a 1–7 scale) was associated with a 2.56-point reduction in that individual's depressive symptom score (on a scale from 0 to 30).

3. Previous research using 6-month intervals has shown that, within individuals, depressive symptoms do not change in a linear fashion but vary at each assessment around the individual's mean (Davila et al., 2003; Karney, 2001). Nevertheless, we also estimated a Level 1 equation including a variable representing number of days since the first assessment (i.e., testing for a linear slope in symptoms over time). There was no evidence of linear change in depressive symptoms (mean slope $= -.006$; $t = -1.20$, $p > .05$) and no other variable predicted variance across subjects in linear change. Therefore, we did not include it in analyses.
4. There is disagreement in the field regarding whether it is necessary to include the mean score of predictor variables across time in the Level 2 model when using group mean centering of those predictors. Some researchers argue that if the mean scores are not included in Level 2, the model will be underspecified because the variance in the mean of the predictor is not included (e.g., Snijders & Bosker, 1999). To address this issue, we reran all analyses with the means of any predictor variables that were included in Level 1 entered as terms in the Level 2 model. Results were virtually identical to those presented here and therefore we do not present them. For similar reasoning and procedure, see Nezlek and Allen (2006).

Table 3. *Within-person associations of weekly relationship functioning with weekly depressive symptoms, with and without controlling for weekly mood states*

Model	Predictor variables	Coefficient	t	Effect size r
1	Relationship functioning	−2.56***	−11.49	−.67
2	Positive mood	−2.62***	−13.17	−.72
	Negative mood	4.00***	16.94	.79
	Relationship functioning	−0.80***	−5.28	−.39

Note. $N = 161$. Coefficients are from models predicting depressive symptoms.
***$p < .001$.

Models controlling for weekly mood

Next, to test whether the observed association between weekly relationship functioning and weekly depressive symptoms would persist when controlling for weekly mood, we estimated a second Level 1 model including positive and negative mood as additional predictors:

$$Y_{ij} = \pi_{0i} + \pi_{1i}(\text{Positive mood})_{ij} + \pi_{2i}(\text{Negative mood})_{ij} + \pi_{3i}(\text{Relationship functioning})_{ij} + \varepsilon_{ij}$$

in which Y_{ij} again represents the depression score of person j at week i. In this function, the coefficient π_{3i} estimates the week-to-week association between the relationship functioning variable and the depressive symptoms controlling for the association between mood and depressive symptoms. See Table 3 (Model 2) for fixed-effect coefficients representing the average value of π_{3i} in this sample from the corresponding person-level model.

>As expected, the within-subject associations between weekly mood and depressive symptoms were large (coefficients $= -2.62$ and 4.00, $p < .001$, for positive and negative mood, respectively). This indicates that weekly self-reports of positive and negative mood and of depressive symptoms do, in fact, strongly covary within individuals. After accounting for the effects of mood, however, relationship functioning continued to contribute substantial unique variance to the prediction of depressive symptoms (coefficient $= -0.80$, $p < .001$). In fact, when controlling for mood, the effect size for relationship functioning remained in the medium range (−.39) and accounted for 15% of the variance in weekly depressive symptoms.

Lagged-week analyses

To explore directionality, we conducted lagged-week analyses to examine the temporal sequence across weeks. In HLM equations parallel to those presented above, depressive symptoms during a given week were predicted by the previous week's relationship functioning while controlling for the previous week's depressive symptoms. In parallel equations, relationship functioning was predicted by the previous week's depressive symptoms, controlling for the previous week's functioning. These analyses revealed no statistically significant lagged associations from one week's relationship functioning to the following week's depression and no significant lagged associations from one week's depressive symptoms to the next week's relationship functioning. The reliability estimates, which in HLM represent the proportion of each parameter's variance that can be treated as meaningful or true variance, for all three variables from the previous week were very low (.06–.22), however, likely due to the relatively low number of time points; time-lagged analyses were designed for data sets with large numbers of assessments (e.g., 150 time points are not uncommon). Consequently, the null results should be interpreted cautiously.

Models controlling for autocorrelation

Finally, an assumption of regression-based analytic procedures is that the residuals (error terms) are independent. Repeated-measurement data with many time points, especially at short intervals, may violate this assumption if there is a correlation between residuals from one time point to the next (i.e., *autocorrelation*). Although HLM naturally controls for some correlation of residuals, we tested for the presence of autocorrelation and whether it impacted our results using the hierarchical multivariate linear models program of HLM. Specifically, we refit our hypothesized models including an autoregressive, Lag 1 correlation model, which specifies that time points closer together correlate more strongly than do time points farther apart. This model did indicate a significant effect of autocorrelation and inclusion of the autoregressive term improved the fit of the model ($p < .001$). None of the substantive findings were altered, however, when autocorrelation was taken into account. Therefore, for simplicity and space, we report results without adjusting for autocorrelation.

Tests of potential moderators of within-subject associations between relationship functioning and depressive symptoms

Results thus far demonstrated that, within individuals, weekly reports of relationship functioning were associated in expected directions with same-week depressive symptoms and that weekly fluctuations in mood did not account for this association. The second purpose of the study was to examine whether this within-person association was equal across participants or moderated by certain relationship or individual characteristics. We tested these hypotheses using Level 2 analyses, in which we estimated the coefficients from Level 1 analyses by adding the moderators to the person-level model:

$$\pi_{0i} = \beta_{00} + \beta_{01}(\text{Femininity}) + \beta_{02}(\text{Masculinity}) + \beta_{03}(\text{Marital status}) + \beta_{04}(\text{Relationship length}) + r_{0i}$$

$$\pi_{1i} = \beta_{10} + \beta_{11}(\text{Femininity}) + \beta_{12}(\text{Masculinity}) + \beta_{13}(\text{Marital status}) + \beta_{14}(\text{Relationship length}) + r_{1i}$$

All moderators were grand mean centered. We were primarily interested in β_{11}–β_{14}, fixed-effect coefficients representing the association between the given moderator and the within-subject association between relationship functioning and depressive symptoms. Nonetheless, we also entered each moderator into the equation predicting average depression across time (π_{0i}) so that hypothesized moderating effects would be estimated controlling for associations between the moderator and the average depressive symptom level. Results showed that higher masculinity and longer relationship length were each associated with lower average depression (see top half of Table 4). We tested hypotheses by examining β_{11}–β_{14} (displayed in the bottom half of Table 4); significant values for each coefficient indicate that the given moderator is related to the within-subject association between relationship functioning and depressive symptoms, controlling for the influence of the other moderators and their associations with average depression over time.

Contrary to hypotheses, femininity did not moderate the within-person associations between relationship functioning and depressive symptoms ($\beta_{11} = -.01$, *ns*). In contrast, masculinity demonstrated the hypothesized moderating effect ($\beta_{12} = .53$, $p < .01$). Recall that we standardized the person-level (Level 2) variables of masculinity and femininity prior to entry in analyses, which makes interpretation of the coefficients via simple slope decompositions straightforward. When predicting weekly depressive symptoms, an increase of 1 *SD* in the masculinity score (equal to a 1-point increase because the measure was standardized) was associated with a 0.53 increase in the coefficient for weekly relationship functioning. The mean coefficient for relationship functioning was −2.56. Therefore, for a woman 1 *SD* above the mean on masculinity, the predicted coefficient for relationship satisfaction as a predictor of depressive symptoms was −2.03 (−2.56 + 0.53), $t = -4.89$, $p < .01$. (We conducted tests of the significance of simple slopes derived from interactions in HLM using newly available free software; Preacher, Curran, & Bauer, 2006). For a woman 1 *SD* below the mean in masculinity, the predicted coefficient was −3.09 (−2.56, − 0.53), $t = -7.73$, $p < .01$. The larger negative coefficient associated with lower masculinity indicates that the negative

Table 4. *Summary of moderation analyses*

Predictor variables	Coefficient	t
Predicting average depressive symptoms across time points (π_{0i})		
Intercept	6.59**	18.81
Femininity	−0.43	−1.06
Masculinity	−1.13**	−3.04
Married	1.31	0.91
Relationship length (years)	−1.19**	−3.66
Predicting the within-subject association between relationship functioning and depressive symptoms (π_{1i})		
Intercept	−2.56**	−11.68
Femininity	−0.01	0.22
Masculinity	0.53**	2.91
Married	−0.63	−1.24
Relationship length (years)	0.44*	2.00

Note. $N = 161$.
$^{*}p < .05$. $^{**}p < .01$.

within-person association between relationship satisfaction and depressive symptoms was stronger for women with low psychological masculinity, although the association is significant at all observed levels of masculinity. Figure 1 illustrates the interaction. Masculinity differentiated women's experience of depressive symptoms on weeks when women experienced relationship functioning that was poorer than their own average, but not on weeks when their relationship functioning was higher than usual.

Marital status showed no moderating effect. Relationship length, however, did moderate the week-to-week association between relationship functioning and depressive symptoms ($\beta_{14} = .44$, $p < .05$). Because 1 point on the relationship length scale indicated 1 year, the coefficient of .44 indicates that an increase of 1 year in relationship length was associated with an increase of .44 in the coefficient for relationship functioning, resulting in a weaker negative association. To illustrate these effects in a manner that allows straightforward comparison with the effects of other moderators, we calculated the coefficient (i.e., slope) for relationship functioning predicting depressive symptoms for women in relationships with lengths 1 *SD* below (0 years 8 months) and above (14 years 10 months) the sample mean. For women in relatively new relationships, the coefficient = −3.04, $t = 6.79$, $p < .01$, and for women in long-term relationships, the coefficient = −2.15, $t = 5.17$, $p < .01$. This indicates that weekly depressive symptoms covaried with relationship functioning for all women but did so more strongly for women in relationships of shorter duration. As Figure 1 illustrates, relationship length differentiated women's depression levels at low, but not high, levels of within-person relationship functioning.

Discussion

In the present study, we examined within-subject associations between women's weekly reports of depressive symptoms and their weekly reports of relationship functioning. Consistent with stress-generation models of depression (Davila et al., 1997) and the marital discord model of depression (Beach et al., 1990), depressive symptoms were negatively associated with same-week relationship functioning levels. That is, during weeks when women reported levels of relationship functioning that were lower than their average levels, they also tended to report higher than average depressive symptoms.

The results of this study further our understanding of the associations between intimate couple relationships and individual well-being

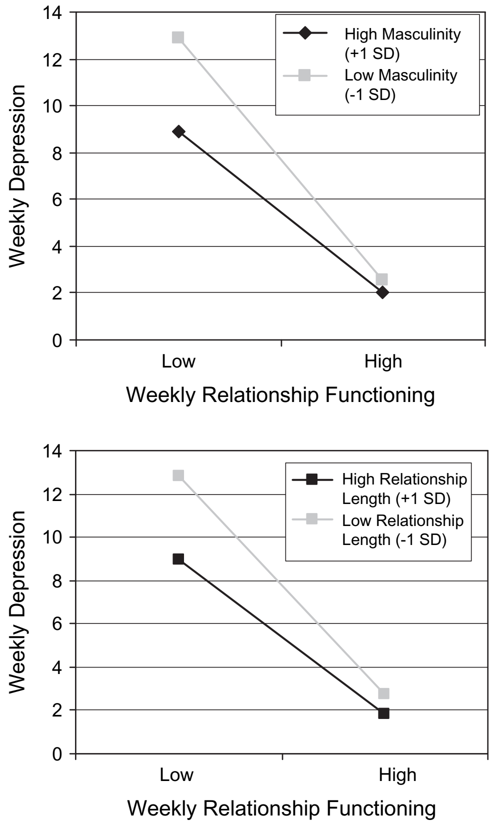

Figure 1. The moderating effects of masculinity and relationship length on the within-subject association between weekly relationship functioning and depressive symptoms.
Note. High and low levels of masculinity and relationship length were defined as ± 1 *SD* from the grand mean.

by demonstrating that small, short-term fluctuations in relationship functioning, which are likely normal in most relationships, tend to co-occur with changes in individuals' well-being. Major drops or boosts in couple functioning are not necessary for the change to be linked with a simultaneous change in depressive symptoms. Rather, the effect size of the week-to-week association between relationship functioning and depressive symptoms ($-.67$) was comparable to effect sizes observed in studies using 6-month intervals between assessments ($-.60$ to $-.74$; Karney, 2001). Thus, the week-to-week flux in individuals' relationship functioning and depressive symptoms appear to covary at a magnitude similar to the covariance between larger, long-term shifts in marital quality and depression. Given that over the course of the study, the majority of women fluctuated between nonclinical and clinical levels of depressive symptoms, it appears that the types of changes in relationship functioning observed across 12 weeks are linked with variation in depressive symptoms that is clinically meaningful. Although we need replication of these findings before drawing confident conclusions, the link between changes in relationship quality and depressive symptoms on a week-to-week basis preliminarily suggests that couple-based interventions, typically offered at weekly intervals for around 3–6 months, have the potential to benefit partners' depressive symptoms. Similarly, individual treatments that decrease an individual's depressive symptoms may promote co-occuring positive changes in the client's intimate relationships.

Fluctuations in individuals' mood from week to week did not account for the within-person associations between relationship functioning and depressive symptoms, suggesting that the observed associations are not attributable to common self-report method variance. Moreover, this finding mitigates concerns that in this nonclinical sample, the within-person variability in the CESD simply reflects changes in feeling states rather than changes in depressive symptoms. By showing that even when controlling for variance in current mood states, women's weekly reports of depressive symptoms covaried with their same-week reports of relationship satisfaction and conflict, the present findings highlight the potential role of relationship events in predicting the course of women's clinically meaningful depressive symptomatology.

Guided by the marital discord model of depression (Beach et al., 1990), we conceptualized the present findings as reflecting an impact of weekly relationship functioning levels on same-week depressive symptoms. The implicit assumption that relationship events impact individuals' well-being more so than vice versa is common in studies of within-person associations between variables assessed on a weekly or daily basis (see Nezlek & Gable, 2001). The main findings, however, were derived from correlational data, prohibiting statements about directionality of effects. Importantly, time-lagged analyses did not reveal any effects of relationship functioning from the previous week on a given week's depressive symptoms or any effects of depression from the previous week on a given week's relationship quality. This finding echoes those of Karney (2001) and Kurdek (1998), suggesting a simple covariation between relationship satisfaction and depressive symptoms rather than a directional relationship. As such, the present data support Kurdek's observation that, although it may be tempting to speculate about causal relations between the two variables, the evidence currently points toward a model of reciprocal change, in which decreases in relationship satisfaction and increases in depressive symptoms accompany one another rather than one preceding the other.

This study also makes an important contribution to the literature by identifying psychological masculinity as a moderator of the within-person link between weekly depressive symptoms and relationship functioning. Women with high levels of masculinity who experienced a weekly decrease in their relationship quality were less likely to show elevated depressive symptoms that week than were those low in masculinity. This finding is consistent with the hypothesis that higher levels of psychological masculinity may serve as a protective factor for women, buffering them from depressive reactions to dips in relationship well-being and from adverse impacts of increased dysphoria on their relationship quality. In light of the strong theoretical base suggesting that passive coping styles, hopelessness, and helplessness confer risk for depression (Abramson et al., 1978; Beck et al., 1979; Nolen-Hoeksema, 1987), it is likely that the active, agentic coping strategies stereotypically masculine individuals use (e.g., Lengua & Stormshak, 2000) may buffer individuals from potential depressive effects of stressors. Applied to these findings, highly masculine women's emotional well-being may be less dependent on the week-to-week fluctuations in relationship functioning (and vice versa) because they are able to cope with decreases in relationship functioning in a less depressogenic manner and to manage increased depressive symptoms without allowing them to taint perceptions of relationship quality. Clinically, this suggests that augmenting individual's stereotypically masculine, active coping strategies may help break cycles of worsening relationship distress and increasing dysphoria. Given that cognitive therapy for depression often includes efforts to improve self-efficacy and teach effective coping strategies (e.g., Beck et al., 1979), it may be that treatment of comorbid depression and marital distress would benefit from including these cognitive strategies. We did not, however, assess coping strategies in this study; future research is needed to directly test whether coping styles account for effects of masculinity on links between couple functioning and depression.

In contrast to masculinity, femininity did not moderate the within-subject association between relationship satisfaction and depressive symptoms. This was surprising, given the rich theory behind proposed links between femininity and sensitivity to relationship stress (Jack, 1991; Waelde et al., 1994). While the nonsignificant moderating effect of femininity might be an anomaly of this data set, perhaps due to the high education level of women in this sample or the somewhat low reliability of the femininity measure, there is evidence to suggest otherwise. The femininity scale showed higher reliability than did the masculinity scale, and the median values of masculinity and femininity were similar to those found in other samples. Thus, it is possible that this pattern of findings reflects a greater importance of instrumental, masculine characteristics than expressive, relationship-focused feminine characteristics in determining the strength of covariation between depression and marital distress. Future research should explore whether this pattern of findings is specific to depression or if femininity may impact within-subject associations between relationship quality and other psychological problems. For example, evidence that femininity is associated with lower rates of substance use and antisocial behavior (Lengua & Stormshak, 2000) suggests that high femininity may buffer women against substance use and acting out in the face of relationship distress.

The current results also extend previous research, which documented the within-subject association between satisfaction and depression in early marriage, by demonstrating that relationship functioning and depressive symptoms also covary within individuals in nonmarried cohabiting relationships and longer term marriages. Marital status did not moderate the within-person associations, suggesting that nonmarried cohabiting women may be just as much at risk for increased depressive symptoms when encountering relationship distress—and for decreased relationship functioning when experiencing dysphoria—as are married women. This is an important finding to note in light of the high prevalence of cohabiting unions (Bumpass & Lu, 2000), which are at heightened risk for relationship problems (Stanley et al., 2004). A large number of women who experience distress in their cohabiting unions are likely to be at risk for elevated depressive symptoms, although they may fall under the radar in the United States because their relationships are often not legally recognized. Thus, an important clinical implication of the present results is that therapists might be well advised to pay attention to current relationship changes occurring for single but cohabiting women presenting with mild depression and may be justified in using couples-based interventions to treat these clients.

Results of moderation analyses also indicated that although the within-person association between weekly relationship functioning and depressive symptoms was present across relationships of different lengths, it was significantly stronger for women in relatively new relationships than for women in long-term relationships. This finding, together with previous evidence that marital quality and personal well-being are more strongly correlated in marriages of less than 8 years than in longer marriages (Proulx et al., 2007), may reflect a self-selection bias, in which women in marriages of longer duration are a select group who have made it past the years during which risk for divorce is highest. It is plausible that women whose personal well-being is most closely tied to concurrent relationship functioning may be more likely to dissolve their relationships early on, while women who remain in their relationships long term are (and always were) less affected by short-term fluctuations in functioning levels. Alternately, it may be that women in relatively new relationships are particularly attentive to any problems in their relationship during a given week. As relationships are maintained over longer time periods, partners' commitment and perceptions of relationship stability tend to increase, which may serve to limit their emotional reactivity to short-term fluctuations in conflict and satisfaction levels (e.g., Stanley, Blumberg, & Markman, 1999). That is, although early in relationships a hostile fight might make a woman feel heightened dysphoria associated with perceptions that her relationship is in trouble, after she and her partner repeatedly weather conflicts successfully and are able to repair the relationship afterward, such conflicts may be less distressing.

This study has several limitations. As mentioned above, because we derived the present findings from correlational data and because lagged analyses were nonsignificant, we can make no conclusions about the direction of effects. To examine the temporal relations between relationship functioning and depressive symptoms more sensitively, future research is needed that captures many more data points from each participant and would therefore generate reliable estimates of the lagged effects of each variable on the other. We draw conclusions with recognition of the nonrepresentative nature of our convenience sample, which may affect general applicability of findings. Participants were highly educated and, because the study required use of the Internet, women of low socioeconomic status were unlikely to participate. Because we recruited this sample in the United States, it is also possible that findings would be different in a different cultural context, for example, in a culture where the masculine characteristics of instrumentality and independence are less highly valued. Nevertheless, the sample, composed of community (nonstudent and nonclinic) women, was far broader than undergraduate samples often used in relationship studies. Further, Web-based samples, despite not being representative of the overall U.S. population, tend to yield similar results as telephone or laboratory studies

(Best et al., 2001; Gosling, Vazire, Srivastava, & John, 2004). We included no men in analyses; future research should explore whether the associations are also present among men. It would be particularly interesting to, in a large representative sample of men and women, simultaneously assess sex and gender roles as potential moderators of the within-person association between relationship functioning and depressive symptoms, to test whether differences in gender roles may account for any sex differences in the strength of the link between relationship functioning and depressive symptoms. Finally, because we assessed depressive symptoms rather than diagnostic depression, further study is needed before drawing conclusions regarding influences between relationship functioning and major depression. Nonetheless, around one third of our sample did report initial depressive symptom levels indicative of clinically significant depression, and over half of the sample fluctuated between clinical and nonclinical levels of symptoms, suggesting that our findings may be relevant to women with clinically significant depression.

Despite the recognized limitations, this study adds to our understanding of the links between relationship functioning and mental health by demonstrating that women's weekly depressive symptoms vary as a function of weekly relationship functioning. We observed these associations among women in both cohabiting relationships and long-term marriages, extending previous findings from newlywed-only samples. In addition, the current findings contribute to knowledge about individual and relationship characteristics that moderate associations between relationship functioning and depression in women, by demonstrating that high levels of stereotypical masculinity reduce the extent to which weekly fluctuations in relationship satisfaction and conflict are associated with corresponding changes in depressive symptoms. Women in relationships of longer duration also showed a smaller week-to-week association between relationship functioning and depressive symptoms than did women in relationships of shorter duration. Together, these findings suggest that women in newer relationships and women who are low in stereotypically masculine characteristics may be at heightened risk for dysphoria when experiencing relationship distress and for compromised relationship functioning during periods of increased depressive symptoms.

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