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Challenging Conventional Wisdom About Who Quits: Revelations From Corporate America

Peter W. Hom
Arizona State University

Loriann Roberson
Columbia University

Aimee D. Ellis
Arizona State University

Findings from 20 corporations from the Attrition and Retention Consortium, which collects quit statistics about 475,458 professionals and managers, extended and disputed established findings about who quits. Multilevel analyses revealed that company tenure is curvilinearly related to turnover and that a job's past attrition rate strengthens the (negative) performance–exit relationship. Further, women quit more than men, while African Americans, Hispanic Americans, and Asian Americans quit more than White Americans, though racial differences disappeared after confounds were controlled for. African American, Hispanic American, and Asian American women quit more than men of the same ethnicities and White Americans, but statistical controls nullified evidence for dual discrimination toward minority women. Greater corporate flight among women and minorities during early employment nonetheless hampers progress toward a more diversified workforce in corporate America.

Keywords: turnover, performance, gender, race, double jeopardy

Effective retention strategies hinge on reliable, timely intelligence about prospective leavers (Judge, 1993; Lyness & Judiesch, 2001; Trevor, Gerhart, & Boudreau, 1997). Thus, whether high or low performers are most liable to quit determines if firms should enact costly steps to curb attrition (Dalton, Todor, & Krackhardt, 1982). Moreover, resolving whether women or people of color are exit prone can alert employers to workplace discrimination and progress toward a more diversified workforce and executive ranks (Brett & Stroh, 1994; Helfat, Harris, & Wolfson, 2006; McCracken, 2000). After all, "the key to lasting diversity is retaining employees from all racial backgrounds" (Zatzick, Elvira, & Cohen, 2003, p. 493). Further, emerging labor shortages in the years ahead—especially among professionals in knowledge-driven industries (Albright & Cluff, 2005; Stevens, 2006) due to impending baby-boomer retirements (Frank, Finnegan, & Taylor, 2004; "Turning Boomers," 2006)—make pinpointing characteristics of potential or current incumbents who quit evermore urgent (Barrick & Zimmerman, 2005; Chatman, 1991).

Peter W. Hom and Aimee D. Ellis, Department of Management, Arizona State University; Loriann Roberson, Teachers' College, Columbia University.

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Correspondence concerning this article should be addressed to Peter W. Hom, Department of Management, Arizona State University, P.O. Box 874006, Tempe, AZ 85287. E-mail: peter.hom@asu.edu

To gather this information, researchers have long explored the personal attributes of those who sever employment ties. Over the years, certain individual correlates of turnover have been repeatedly found, such as higher quits among short-tenure and marginal employees (Griffeth, Hom, & Gaertner, 2000; Meitzen, 1986). Recent studies nonetheless challenged such established findings, revealing that how tenure and performance relate to terminations may depart from the typical pattern and assume different strengths and forms under certain circumstances (Allen & Griffeth, 2001; Griffeth et al., 2000; Salamin & Hom, 2005). Empirical findings for relationships of diversity dimensions to turnover are more variable. Practitioners and scholars have long debated whether women in managerial and professional jobs historically dominated by men are abandoning corporate America (Stroh, Brett, & Reilly, 1996). Unfortunately, claims about women's corporate flight have been grounded more in anecdotal than empirical data (Schwartz, 1989). Even more equivocal is the scarcer evidence for the race–turnover association in well-paid, high-status professions where people of color are underrepresented (Roberson, 2004). Moreover, the joint effects of race and sex on job separations among managers and professionals in corporations remain virtually untested. Past work scrutinized their effects separately (Griffeth et al., 2000), assessing gender effects with mainly Whites (underrepresenting minorities) and assessing race effects with mostly men (undersampling women; Bretz, Boudreau, & Judge, 1994; Payne & Huffman, 2005). Such approaches have left a conspicuous gap in knowledge about exits among professionals who are both female and minority (Berdahl & Moore, 2006; King, 1988), despite persistent stories about their exodus from corporate life (Giscombe & Mattis, 2002).

Because such voids in knowledge stem from inadequate data, we report findings from the Attrition and Retention Consortium

Table 1
Numbers of Women and Men by Racial Group

Ethnicity	Women		Men		Total	
	n	Percentage of row	*n*	Percentage of row	*N*	Percentage of total
White American	84,187	26.10%	238,316	73.90%	322,503	79.82%
African American	12,808	45.63%	15,264	54.37%	28,072	6.95%
Asian American	10,831	31.53%	23,521	68.47%	34,352	8.50%
Hispanic American	5,610	32.96%	11,415	67.04%	17,025	4.21%
Native American	726	34.58%	1,374	65.42%	2,100	0.52%
Total	114,162	28.25%	289,890	71.75%	404,052	

Note. Demographic data provided by 16 of the 20 Attrition and Retention consortium members.

(ARC; www.retentionconsortium.org), composed primarily of *Fortune* 500 corporations that benchmark quit statistics on nearly a half-million employees in a broad spectrum of exempt jobs. ARC emerged because timely quit data for business professions, where departures are most costly and disruptive (Griffeth & Hom, 2001), were nonexistent. In this article, we leverage ARC's unique data set to illuminate current disputes in turnover research. Specifically, this study explicitly tests curvilinear tenure–turnover relationships neglected in organizational psychology (Griffeth et al., 2000). We also consider a new moderator of performance–turnover relationship: the past attrition rate in a job (Steel, 1996). Further, we assess relationships of gender and race to turnover, comparing turnover among five major racioethnic groups and examining the "double jeopardy" hypothesis for professional women of color. All told, our project addresses several controversial and unanswered questions in the turnover literature. In what follows, we review the existing body of evidence about key leaver attributes and discuss how this research contributes to theory and practice.

Curvilinear Relationship Between Company Tenure and Turnover

Among the most robust personal correlates of turnover (Cotton & Tuttle, 1986; Griffeth et al., 2000; Hom & Griffeth, 1995), firm tenure's negative relationship to quits has become a stylized fact (Farber, 1994; Jovanovic, 1979). Theoretical perspectives on attraction–selection–attrition (Schneider, Smith, Taylor, & Fleenor, 1998) and job matching (Jovanovic, 1979) explain why quits decline as job duration grows. During entry, new hires learn if they fit with corporate culture and values, can perform required tasks, and derive satisfaction from available rewards as they experience the job (Meitzen, 1986). Over time, "misfits" exit (leaving most during early employment when person–job mismatches are readily detected; Caldwell & O'Reilly, 1990; Sicherman, 1996), while those who fit remain (exiting less often as seniority accrues). Alternatively, job embeddedness theory posits that job incumbents accumulate more benefits (e.g., pension) and interpersonal ties as they log more time in the firm (T. R. Mitchell & Lee, 2001). Therefore, long seniority deters resignations because veteran employees relinquish ample tenure-driven benefits and disrupt relationships if they switch jobs.

Yet other theories and work have suggested that this standard portrayal of tenure–quit linearity is oversimplified. Tuma (1976) first speculated that the rate of moving out of a job diminishes with job duration at a decelerating rate, presuming that mastering a new job (and amassing rewards) proceeds at a falling rate. In a job attraction model of realistic job previews, Meglino, Ravlin, and DeNisi (2000) later conceptualized quits as a downward sloping convex function of tenure for newcomers (not given previews). Beginning employees hold inflated job expectations, but "their initial encounters with specific job aspects should produce a gradual decline in attraction over time" (p. 410). Job-matching theorists also have posited that "during the initial stages of tenure, the worker discovers the quality of the job match, and if bad information is discovered, he will quit to find a better match" (Meitzen, 1986, p. 164). By extension, learning about match quality by employer and employee should largely end after this period, and quits due to mismatches should fall.

Further, human capital theories foreshadow tenure–quit curvilinearity (Lazear, 1998). Companies invest in new workers by training them in firm-specific skills, while trainees accept subpar wages in exchange. As firm-specific human capital accumulates over time, workers' wages increase to compensate for growing productivity and eventually exceed wages available elsewhere because firm-specific skills are worth less to other employers. Gradually, workers' wage profiles flatten (i.e., upward sloping concave curve; Topel, 1991) as they learn the most important skills first and devote less time to training (Lazear, 1998). Turnover thus becomes increasingly costly for experienced workers who would forfeit rising pay (corresponding to greater job-specific skills) if they exit and fail to find comparable pay in the labor market (Park & Sandefur, 2003; Topel, 1991). Consequently, the quit risk lessens with increasing tenure until this risk reaches some minimum at a time when current wages outstrip market wages.

Cursory graphic depictions have suggested that quits monotonically decline as tenure progresses until the tenure–turnover curve flattens (Burt, 2005; Dickter, Roznowski, & Harrison, 1996; Harrison, Virick, & William, 1996; Lyness & Judiesch, 2001; Meglino, DeNisi, & Ravlin, 1993; Sacco & Schmitt, 2005). Yet empirical corroboration for this curvilinearity is rare, despite its apparent ubiquity across different occupations (Morita, Lee, & Mowday, 1989; Murnane, Singer, & Willett, 1988), sexes (Sicherman, 1996), educational levels (Royalty, 1998), racial minorities (Park & Sandefur, 2003), types of leavers (Sicherman, 1996), and nationalities (Farber, 1994; Light & Ureta, 1992; Weller, 2006). A few labor economic tests have statistically affirmed this trend for broad national or company-wide samples (Ippolito, 1991; Muna-

Table 2
Descriptive Statistics and Variable Intercorrelations

Variable	Mean	Standard Deviation	1	2	3	4	5	6	7	8	9	10	11
1. Quit	.04[a]	.193[a]	—										
2. Firm tenure	7.673[a]	4.3[a]	−.130[*a]	—									
3. Firm tenure^2	77.35[a]	60.145[a]	−.127[a]	.979[*a]	—								
4. Job performance	.123[b]	.330[b]	−.015[*b]	.079[*b]	.063[*b]	—							
5. Past quit rate per job by firm	.035[a]	.027[a]	.075[*a]	−.079[*a]	−.081[*a]	.064[*b]	—						
6. Sex	.28[c]	.45[c]	.040[*c]	−.050[*c]	−.056[*c]	—	.104[*c]	—					
7. African American	.07[c]	.25[c]	.019[*c]	−.057[*c]	−.058[*c]	—	.030[*c]	.105[*c]	—				
8. Asian American	.09[c]	.29[c]	.000[c]	−.125[*c]	−.130[*c]	—	−.025[*c]	.022[*c]	−.083[*c]	—			
9. Hispanic American	−.04[c]	.20[c]	.009[*c]	−.039[*c]	−.041[*c]	—	.009[*c]	.022[*c]	−.057[*c]	−.064[*c]	—		
10. Native American	.01[c]	.07[c]	.001[c]	−.006[*c]	−.008[*c]	—	.010[*c]	.010[*c]	−.020[*c]	−.022[*c]	−.015[*c]	—	
11. Percent women per job by firm	28.25[c]	15.83[c]	.056[*c]	−.040[*c]	−.054[*c]	—	.442[*c]	.352[*c]	.082[*c]	−.007[*c]	.045[*c]	.006[*c]	—
12. Percent minority per job by firm	20.18[c]	8.51[c]	.013[*c]	−.045[*c]	−.078[*c]	—	.139[*c]	.118[*c]	.042[*c]	.225[*c]	.059[*c]	0.00[c]	.335[*c]

Note. Quit is 1 for voluntary quits, 0 for stayers. Sex is 1 for women, 0 for men. African American is 1 for Blacks, 0 otherwise. Asian American is 1 for Asian Americans, 0 otherwise. Hispanic American is 1 for Hispanics, 0 otherwise. Native American is 1 for Native Americans, 0 otherwise. Dashes indicate that correlations between demographic variables and performance are not available due to the method by which participating companies submit data.
[a] $N = 475{,}458$. [b] $N = 362{,}063$. [c] $N = 404{,}052$.
[*] $p < .05$.

singhe & Sigman, 2004; Royalty, 1998; Sicherman, 1996). Following their lead, our investigation explicitly checks and generalizes tenure–quit curvilinearity for corporate professionals whose critical skills are costly and difficult to replace (Griffeth & Hom, 2001; Stevens, 2006).

Such scrutiny is warranted on practical and theoretical grounds. Current assumptions about tenure–quit linearity misleadingly imply that departures steadily decrease throughout employment and that employers should exclusively focus retention efforts on preventing premature exits (presuming nonexistent exits among long-term incumbents). Curvilinearity however suggests that job matching, disillusionment, and human capital investment diminish over time until these processes fade away but allows for nonzero quits among seasoned employees (Lazear, 1998; Meglino et al., 2000; Meitzen, 1986). The aforementioned theories and findings thus imply the following:

> *Hypothesis 1:* Tenure is curvilinearly related to voluntary turnover.

Though less common, some studies disclose nonmonotonic relationships in which turnover is low during initial employment and peaks after some transition period (e.g., probation), subsequently falling (Meglino et al., 1993; Morita et al., 1989). First calling attention to this "cubic" relationship in the National Longitudinal Survey of Youth (1979–1988), Farber (1994) argued that

> the hazard of a job ending will first be quite low as workers and firms learn about the match quality [which takes time and because exiting a job *early* is costly], then will rise as matches revealed to be bad end, and finally will fall as the continuing matches are disproportionately of high quality. (p. 591)

Viewed as a "liability of adolescence," Burt (2005) later termed this relationship form a "kinked decay function" that represents a temporal course for assorted partnerships, such as marriages, joint ventures, and auditor–client relationships. Similarly, socialization scholars identified a "honeymoon" phase during the early phase of a new job that binds newcomers until they later encounter reality shock and leave (Holloran, Mishkin, & Hanson, 1980; Kramer, 1974). All the same, this cubic form merits additional verification as these conclusions have been based on casual inspections of empirical hazard functions rather than on statistical tests (Booth, Fracesconi, & Garcia-Serrano, 1999; Sicherman, 1996). The few existing statistical tests however included national samples of American youth (i.e., National Longitudinal Survey of Youth: 1978–1988, Farber, 1994; National Longitudinal Survey of Youth: 1979–1994, Park & Sandefur, 2003) or German workers (Weller, 2006). Given informal and limited evidence for this curve in one (banking) profession (Burt, 2005), we determine whether this bell-shaped pattern holds more broadly for other professions, testing the following:

> *Hypothesis 2:* A cubic relationship exists between tenure and voluntary turnover.

Moderator of the Performance–Turnover Relationship

Because turnover functionality depends on who leaves (Dalton et al., 1982), many researchers have probed performance's relationship to exits. Meta-analyses have invariably uncovered negative performance–termination correlations (Bycio, Hackett, & Alvares, 1990; Griffeth et al., 2000; McEvoy & Cascio, 1987; Williams & Livingstone, 1994). Though discerning curvilinearity, two of three recent applications of more powerful statistical techniques also find monotonic inverse relationships (Iverson & Deery, 1999; Salamin & Hom, 2005; Trevor et al., 1997). Pooling results from these tests still yields a –.18 correlation. In contrast to extensive inquiries into the nature and strength of this relationship,

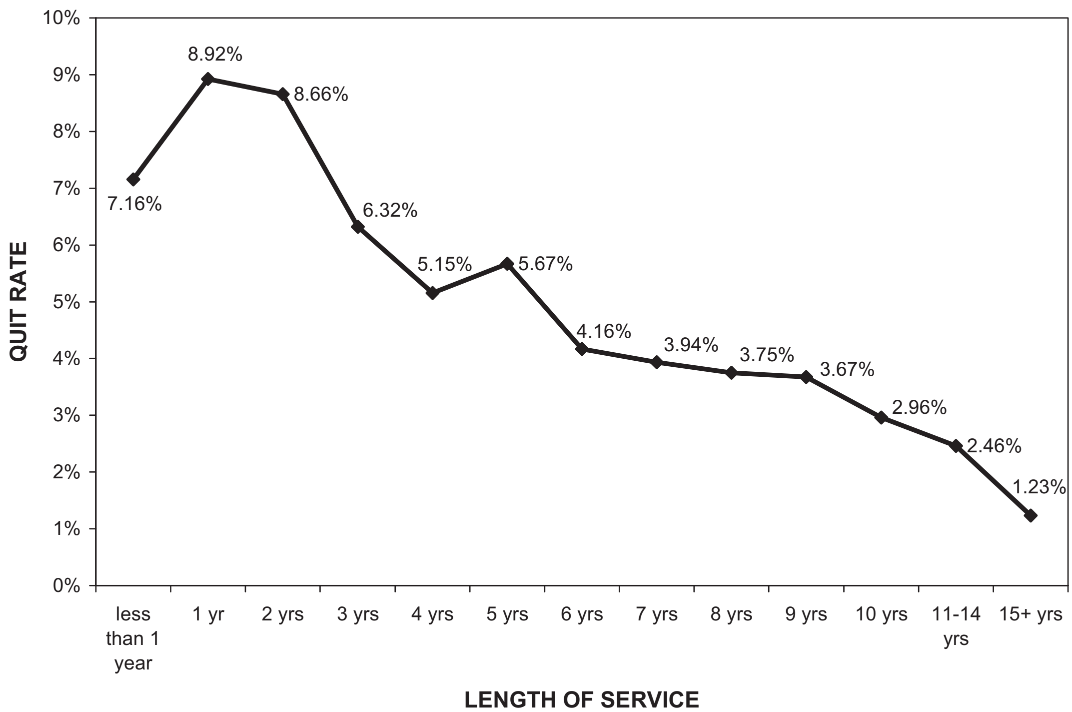


Figure 1. Turnover rates for different years of seniority.

explorations of moderators of this relationship have been infrequent and have inspected a few moderators, such as merit-based rewards, performance visibility, and unemployment rates (Allen & Griffeth, 2001; Griffeth et al., 2000; Salamin & Hom, 2005; Williams & Livingstone, 1994).

A moderator that warrants further scrutiny is job opportunities. Early on, McEvoy and Cascio (1987) proposed that job availability affects performance–turnover associations, such that low unemployment strengthens this relationship because marginal performers can readily change jobs. Although nonsignificant with five studies, their meta-analytic test unexpectedly estimated a stronger association in poor job markets. A more comprehensive meta-analysis also failed to validate unemployment moderation, concluding that "the scarcity or abundance of alternative jobs had little effect on the link between performance and voluntary turnover" (Williams & Livingstone, 1994, p. 294). Even so, they acknowledged that objective joblessness measures may lack construct validity, undermining support for moderation.

We revisit McEvoy and Cascio's (1987) moderator hypothesis with an alternative measure of job availability pioneered by Steel (1996). Compared with conventional labor-market indices such as unemployment rates, a job's *historical retention rate* as demonstrated by Steel (1996)—the retention rate for 1 year prior to the study for each Air Force occupational specialty—was the best objective labor-market proxy accounting for unique variance in military reenlistments and one of two (the other being Department of Labor income projections for a career field) correlating with perceived occupational demand. The historical retention rate—or its obverse, historical attrition rate—in a job thus better captures employment opportunities in an occupation than do customary joblessness measures. Following McEvoy and Cascio's (1987) logic, we envision that the historical rate of job exits conditions how performance relates to departures and accordingly evaluate the following hypothesis:

> *Hypothesis 3:* A job's historical quit rate moderates the performance–turnover relationship, such that this relationship is stronger when the historical quit rate is high.

Corporate Exodus of Women Professionals and Managers

Schwartz (1989) first called attention to women managers' exodus from corporate America, proclaiming that "the rate of turnover in management positions is 2 1/2 times higher among top-performing women than it is among men" (p. 65) on the basis of a company report. Such corporate migration may arise from women's greater discomfort in environments predominated by men (Riordan, Schaffer, & Stewart, 2005; Tsui & Gutek, 1999). Diversity and demography scholars have contended that women in male-majority work settings lack organizational attachment and endure social isolation because they do not self-categorize themselves as ingroup members and do not bond with demographically dissimilar men (Riordan et al., 2005). In support, Elvira and Cohen (2001) showed that women employed in business units where they are numerically rare and most different demographically more often quit. Moreover, domestic responsibilities (since firms may poorly accommodate such obligations or sanction those for using existing accommodations; Hill, Märtinson, Ferris, & Baker, 2004; Johnson, Lowe, & Reckers, in press) pull women managers and professionals away from a job more than do those of their male counterparts (Dalton, Hill, & Ramsay, 1997; Hewlett & Luce, 2005), while inferior work outcomes (e.g., lower pay), sparser

Table 3
Hierarchical Linear Modeling Models for Testing Effects of Job Tenure

Model	Parameter estimate γ_{00}	γ_{01}	γ_{02}	γ_{10}	γ_{20}	γ_{30}	τ_{00}	τ_{11}	τ_{22}
Model 1: Tenure + tenure^2 predicts quits									
L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{tenure}_{ij}) + \beta_{2j}(\text{tenure}_{ij}^2)$									
L2: $\beta_{0j} = \gamma_{00} + U_{0j}$	-3.44^*			-0.01	-0.01^*		0.31^*	0.04^*	0.0002^*
L2: $\beta_{1j} = \gamma_{10} + U_{1j}$									
L2: $\beta_{2j} = \gamma_{20} + U_{2j}$									
Model 2: Tenure + tenure^2 + tenure^3 predict quits									
L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{tenure}_{ij}) + \beta_{2j}(\text{tenure}_{ij}^2) + \beta_{3j}(\text{tenure}_{ij}^3)$									
L2: $\beta_{0j} = \gamma_{00} + U_{0j}$	-3.44^*			-0.03	-0.01	-0.0002	0.31^*	0.04^*	0.0002^*
L2: $\beta_{1j} = \gamma_{10} + U_{1j}$									
L2: $\beta_{2j} = \gamma_{20} + U_{2j}$									
L2: $\beta_{3j} = \gamma_{30}$[a]									
Model 3: Tenure + tenure^2 + control variables predict quits									
L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{tenure}_{ij}) + \beta_{2j}(\text{tenure}_{ij}^2) + \beta_{3j}(\text{jobquits}_{jk})$									
L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_j) + \gamma_{02}(\text{workforce maturity}_j) + U_{0j}$									
L2: $\beta_{1j} = \gamma_{10} + U_{1j}$	-3.48^*	0.00	0.26	-0.01	-0.01^*	6.94^*	0.23^*	0.04^*	0.0002^*
L2: $\beta_{2j} = \gamma_{20} + U_{2j}$									
L2: $\beta_{3j} = \gamma_{30}$[1]									

Note. $N = 475{,}458$. L1 = Level 1; L2 = Level 2; γ_{00} = intercept of L2 regression predicting β_{0j}; γ_{01} = workforce size slope in L2 regression predicting β_{0j}; γ_{02} = workforce maturity slope in L2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}; jobquits_{jk} = past quit rate in job k in firm j; workforce size_j = size of exempt population in firm j; $\text{workforce maturity}_j$ = percent of employees with 10 or more years of seniority in firm j.
[a] For model convergence, effects of past job quit rates are fixed (i.e., $\tau_{33} = 0$).
$^* p \leq .05$.

developmental opportunities (e.g., fewer international assignments), more career obstacles (e.g., exclusion from information networks), and gender and sexual harassment push them away (Brett & Stroh, 1999; Cleveland, Vescio, & Barnes-Farrell, 2005; Eagly & Karau, 2002; Griffeth & Hom, 2001; Laband & Lentz, 1998; Lyness & Heilman, 2006; Lyness & Thompson, 1997; Morris, 2005; Valian, 1999).

These factors for leaving are most pronounced in traditional male-dominated professions in corporate America, where gender inequity is worse (Valian, 1999) and “there has been the most explicit concern about inequitable treatment and glass-ceiling effects” (Petersen & Saporta, 2004, p. 896). Yet empirical research has not clearly validated Schwartz’s (1989) supposition about women’s withdrawal from corporate life (Brett & Stroh, 1994). Higher female turnover has been found among lawyers, auditors, and managers (Arnold & Feldman, 1982; A. Cohen, 1999; Greenhaus, Collins, Singh, & Parasuraman, 1997; Noonan & Corcoran, 2004; Rosin & Korabik, 1995; Sheridan, 1992; Stroh et al., 1996), while a Harris poll reported that women professionals experience more career interruptions than do men (Hewlett & Luce, 2005). By contrast, other researchers did not find higher turnover for women managers (Bretz et al., 1994; Dreher & Cox, 2000; Lewis, 1992; Lyness & Judiesch, 2001; Lyness & Thompson, 1997; McKay et al., 2007), auditors (Dalton et al., 1997), army officers (Payne & Huffman, 2005), professionals in a regulated firm (Petersen & Saporta, 2004), or a nationally representative sample of well-educated women (Royalty, 1998).

Methodological differences across studies may explain conflicting results for female corporate flight. For example, rather than voluntary job quits, many investigators assessed occupational (including switching professional specialties) or labor-market exits (Greenhaus et al., 1997; Hewlett & Luce, 2005; Noonan & Corcoran, 2004), quit intentions (McKay et al., 2007; Rosin & Korabik, 1995), or reasons for leaving among former employees (Dalton et al., 1997; Rosin & Korabik, 1990). Other researchers did not differentiate voluntary from involuntary terminations (Dreher & Cox, 2000; Lewis, 1992; Stroh et al., 1996). Surprisingly, few studies of voluntary turnover behaviors have been conducted in business corporations, the prime target of glass-ceiling complaints (Brett & Stroh, 1999; Rosin & Korabik, 1995). Instead, gender-disparity research has sampled law or accounting firms (Arnold & Feldman, 1982; A. Cohen, 1999; Dalton et al., 1997; Greenhaus et al., 1997; Noonan & Corcoran, 2004; Sheridan, 1992), contexts that exacerbate female exits because their rigid “up-or-out” promotion systems increase work–family conflict during prime child-bearing years (Malos & Campion, 1995; “Why Law Firms Cannot Afford,” 1999).

Alternatively, gender comparisons have relied on broad samples of the well-educated (Hewlett & Luce, 2005; Royalty, 1998), where inclusion of lower paying female-dominated professions (e.g., teachers) inflates female quits (Black & Juhn, 2000; Brett & Stroh, 1999; Murnane et al., 1988). Still other scholars studied public-sector organizations (Lewis, 1992; Payne & Huffman, 2005), where stronger enforcement of equal employment regulations may lessen women’s leaving (Eagly & Karau, 2002). Thus, though important, the bulk of academic studies have not directly addressed corporate desertion among women professionals and managers (Hewlett & Luce, 2005; Schwartz, 1989). Finally, two of

Table 4
Hierarchical Linear Modeling Models for Testing Interaction Between Job Performance and Historical Job Quit Rate

Model	Parameter estimate													
	γ_{00}	γ_{01}	γ_{02}	γ_{10}	γ_{20}	γ_{30}	γ_{40}	γ_{50}	τ_{00}	τ_{11}	τ_{22}	τ_{33}	τ_{44}	τ_{55}
Model 4: Performance + past job quit + Performance × Jobquits predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{performance}_{ij}) + \beta_{2j}(\text{jobquits}_{jk}) + \beta_{3j}(\text{jobperform}_{jk})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	−3.17*			−0.76*	16.06*	−11.62*			0.14*	1.02*	112.45*	51.05*		
Model 5: Performance + past job quit + Performance × Jobquits + control variables predict quits[a] L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{performance}_{ij}) + \beta_{2j}(\text{jobquits}_{jk}) + \beta_{3j}(\text{jobperform}_{jk}) + \beta_{4j}(\text{tenure}_{ij}) + \beta_{5j}(\text{tenure}_{ij}^2)$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_j) + \gamma_{02}(\text{workforce maturity}_j) + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$ L2: $\beta_{4j} = \gamma_{40}$ L2: $\beta_{5j} = \gamma_{50}$	−3.42*	0.00	2.74*	−0.71*	12.49*	−9.43*	−0.07*	−0.01*	0.12*	1.07*	109.91*	102.71*	0.00	0.00

Note. $N = 362{,}063$. L1 = Level 1; L2 = Level 2; γ_{00} = intercept of L2 regression predicting β_{0j}; γ_{01} = workforce size slope in L2 regression predicting β_{0j}; γ_{02} = workforce maturity slope in L2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; γ_{40} = intercept of L2 regression predicting β_{4j}; γ_{50} = intercept of L2 regression predicting β_{5j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2i}; τ_{33} = variance in L2 residual for models predicting β_{3j}; τ_{44} = variance in L2 residual for models predicting β_{4j}; τ_{55} = variance in L2 residual for models predicting β_{5j}; jobquits_{jk} = past quit rate in job k in firm j; performance = 1 for high performers, 0 for nonhigh performers; and jobperform_{jk} = performance × Past Quit Rate in job k in firm j.
[a] For model convergence, this model fixed tenure effects, such that $\tau_{66} = \tau_{77} = 0$.
* $p \leq .05$.

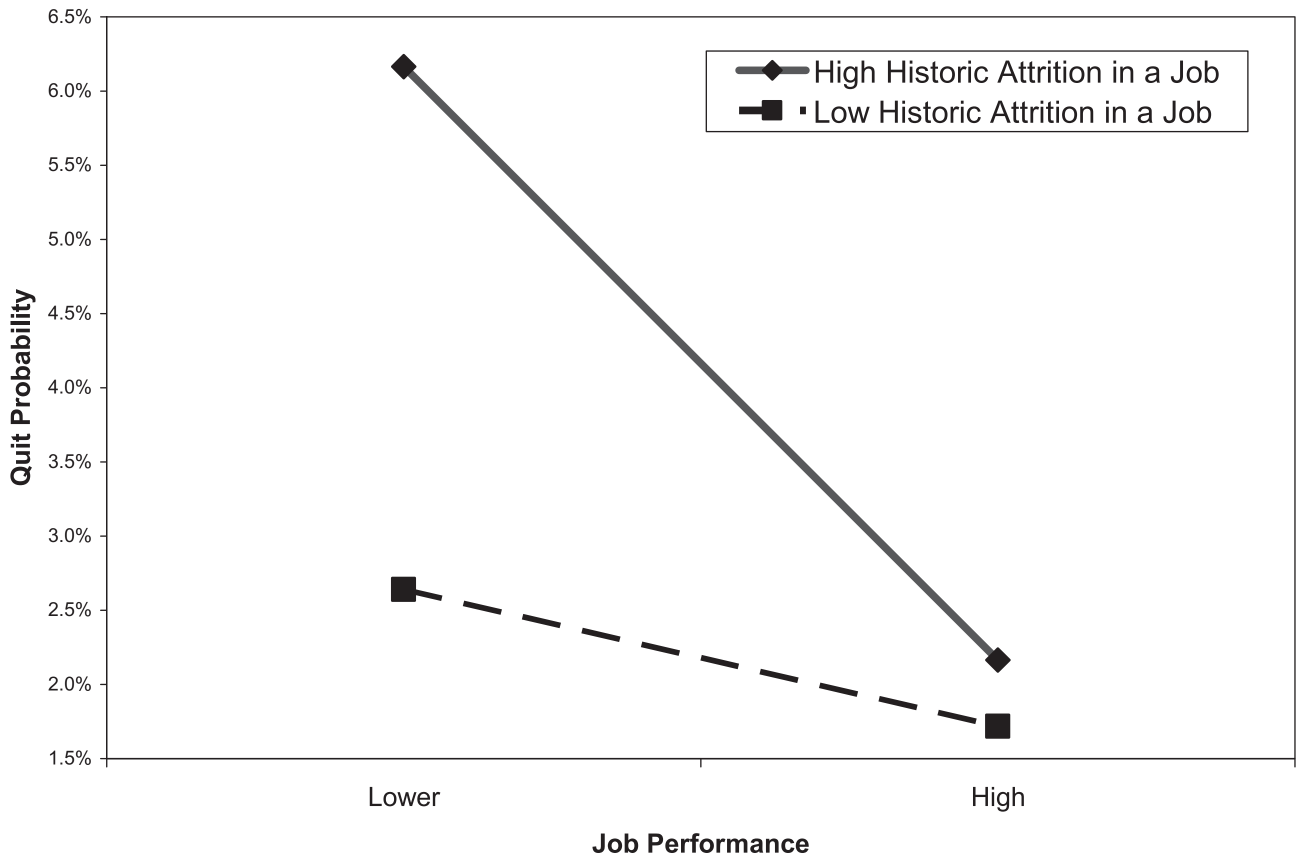


Figure 2. Interaction between job performance and historical job quit rate.

three tests in the private sector failed to detect higher voluntary exits for women, though generalizability of their null findings to other corporate environments remains an open question as each study was performed in a single firm (Bretz et al., 1994; Lyness & Judiesch, 2001; Petersen & Saporta, 2004).

Because accumulated evidence is inconclusive, our investigation revisits longstanding uncertainty over women's corporate exodus with a sample of over 100,000 women from a wide array of professional jobs and 16 companies. We thus offer a broader sample for examining this phenomenon over the single business establishments previously studied (Lyness & Judiesch, 2001; Petersen & Saporta, 2004) as well as more timely statistics (from the year, 2003) about gender disparities (from 1993: Bretz et al., 1994; from 1992–1995: Lyness & Judiesch, 2001; from 1978–1986; Petersen & Saporta, 2004). Reviewing recent data is crucial because the growing female share in traditional male fields renders past estimates, when women were underrepresented, obsolete (Black & Juhn, 2000; Elvira & Cohen, 2001; Shellenbarger, 2003). Although sex differences may shrink with additional controls such as tenure (Booth et al., 1999; Valian, 1999), we advance the following, given prevailing theories about how gender discrimination and/or domestic obligations stimulate female exits (Cleveland et al., 2005; Dalton et al., 1997; Rosin & Korabik, 1995; Stroh et al., 1996):

> *Hypothesis 4:* Women professionals and managers quit more than men.

Relational demography theory and research further have suggested that the male–female gap in terminations is smaller in jobs marked by higher female incumbency, especially staff functions such as human resources and corporate affairs (Elvira & Cohen, 2001; Lyness & Heilman, 2006). Women in jobs where they are tokens or numerical minorities face more prejudice, sexual harassment, social isolation, overachievement pressures, and stereotype threat (Dovidio & Hebl, 2005; Eagly & Karau, 2002; Riordan et al., 2005; Sims, Drasgow, & Fitzgerald, 2005; Wasti, Bergman, Glomb, & Drasgow, 2000). Reinforcing this proposition, Sacco and Schmitt (2005) documented fewer female quits in quick-service restaurants employing more women. Because jobs vary in female representation, we project the following:

> *Hypothesis 5:* A higher proportion of women incumbents in a job reduces gender differences in job separations.

Minority Flight From the Corporate World

The Federal Glass Ceiling Commission (1995) as well as other authors named turnover costs as a major ramification of a glass ceiling for minorities (Cox, 1994; Cox & Blake, 1991; Daniels, 2004; Jones, 1986). Yet reliable evidence for disparate rates of attrition across races in high-status, highly paid professions (Glover, Mynatt, & Schroeder, 2000; Huffman & Cohen, 2004) is scant (Roberson, 2004). Some recent large-scale tests of racial differences have surfaced (Leonard & Levine, 2006; Sacco &

Table 5
Hierarchical Linear Modeling Models for Testing Effects of Sex and Minority Status

Models	Parameter Estimates																	
	γ_{00}	γ_{01}	γ_{02}	γ_{10}	γ_{20}	γ_{30}	γ_{40}	γ_{50}	γ_{60}	γ_{70}	τ_{00}	τ_{11}	τ_{22}	τ_{33}	τ_{44}	τ_{55}	τ_{66}	τ_{77}
Model 6: Sex predicts quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$	−3.31*			0.31*							0.22*	0.03*						
Model 7: Minority status predicts quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{minority}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$	3.25*			0.20*							0.21*	0.03*						
Model 8: Sex + minority status + control variables predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{minority}_{ij}) + \beta_{3j}(\text{tenure}_{ij}) + \beta_{4j}(\text{tenure}_{ij}^{2}) + \beta_{5j}(\text{jobsex}_{jk}) + \beta_{6j}(\text{jobminority}_{jk}) + \beta_{7j}(\text{jobquits}_{jk})$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_j) + \gamma_{02}(\text{workforce maturity}_j) + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$ L2: $\beta_{4j} = \gamma_{40}$[a] L2: $\beta_{5j} = \gamma_{50} + U_{5j}$ L2: $\beta_{6j} = \gamma_{60} + U_{6j}$ L2: $\beta_{7j} = \gamma_{70} + U_{7j}$	−3.70*	0.001*	0.37	0.19*	−0.01	−0.01	−0.01*	0.002	−0.02	10.66*	0.34*	0.04*	0.03	0.01*		0.0001*	0.002*	135.66*

Note. N = 404,052. L1 = Level 1; L2 = Level 2; γ_{00} = intercept of L2 regression predicting β_{0j}; γ_{01} = workforce size slope of L2 regression predicting β_{0j}; γ_{02} = workforce maturity slope of L2 regression predicting β_{0j}; γ_{03} = female percent slope of L2 regression predicting β_{0j}; γ_{04} = minority percent slope of L2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; γ_{40} = intercept of L2 regression predicting β_{4j}; γ_{50} = intercept of L2 regression predicting β_{5j}; γ_{60} = intercept of L2 regression predicting β_{6j}; γ_{70} = intercept of L2 regression predicting β_{7j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}; τ_{33} = variance in L2 residual for models predicting β_{3j}; τ_{44} = variance in L2 residual for models predicting β_{4j}; τ_{55} = variance in L2 residual for models predicting β_{5j}; τ_{66} = variance in L2 residual for models predicting β_{6j}; sex is coded as 1 for women and 0 for men; minority is coded 1 for minorities and 0 for whites; jobsex_{jk} = percent of women incumbents in job k in firm j; jobminority_{jk} = percent of minority incumbents in job k in firm j.
[a] For model convergence, effect of tenure squared is fixed (i.e., $\tau_{44} = 0$).
* $p \leq .05$.

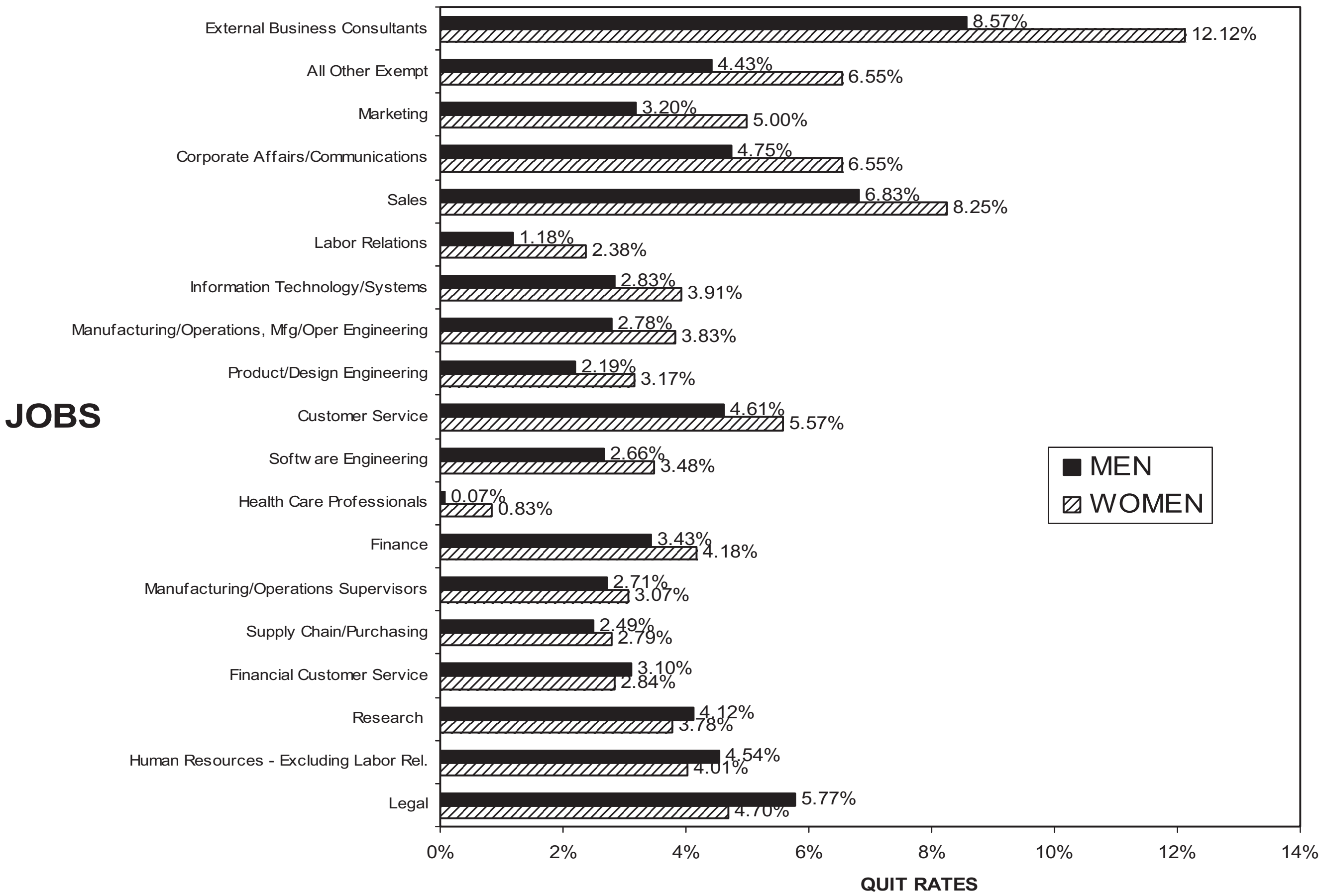


Figure 3. Gender differences for different jobs. Mfg = manufacturing; oper = operations; rel. = relations.

Schmitt, 2005; Zatzick et al., 2003), but in general, "hard numbers are difficult to come by" (Allers, 2005, p. 101) when estimating minority quit rates in exempt occupations. Four studies revealed that minorities exit more than Whites in professions, such as Air Force officers (Baldwin, 2000), managers (Bretz et al., 1994; McKay et al., 2007), nurses (Somers, 1996), and lawyers (Noonan & Corcoran, 2004). In contrast, Ghiselli, La Lopa, and Bai (2001) found more turnover among White than non-White food service managers, although other scholars observed no racial gap in attrition among MBA graduates (Dreher & Cox, 2000), executives (Lyness & Thompson, 1997), and Army officers (Payne & Huffman, 2005). Therefore, empirical research has failed to clearly substantiate accounts that many managers and professionals of color are opting out of corporate life (Cox & Blake, 1991; Daniels, 2004).

Contributing to ambiguity, turnover researchers have typically combined different minority groups together to compare with Whites (Elvira & Cohen, 2001; Griffeth et al., 2000; Payne & Huffman, 2005; Somers, 1996). Though investigators may lack adequate numbers to treat each racial minority as a separate group, this common practice may produce erroneous conclusions should these groups differ in turnover (Roberson, 2004). To illustrate, a meta-analysis by Griffeth and his colleagues (2000) estimated a negligible –.01 correlation between race and turnover based on studies combining minorities into one group. Yet high quits in one racioethnic group may offset low quits in another, obscuring minority–majority differences (McKay et al., 2007; Park & Sandefur, 2003). Indeed, treating different minorities alike perpetuates the thinking that all people of color are an undifferentiated "other" and overlooks each group's unique sociohistorical experiences (Nkomo, 1992).

Our study examines turnover for five racioethnic groups: Asian Americans, African Americans, Hispanic Americans, Native Americans, and White Americans. Their different sociohistorical and contextual experiences may generate dissimilar mobility patterns (Roberson, 2004). Many studies have recounted stronger societal antipathy toward African Americans and Hispanic Americans than toward other minority groups (Greenwald & Banaji, 1995; Hosoda, Stone, & Stone-Romero, 2003; Taylor, 1998; Ziegert & Hanges, 2005). Most notably, stereotype content research has shown that White Americans feel more negative affect toward African and Hispanic Americans (viewing them as threats to their safety, property, or livelihood) than toward Asian and Native Americans (viewing them as threats to group values and ideals; Cottrell & Neuberg, 2005; Fiske, Cuddy, Glick, & Xu, 2002). For example, Cottrell and Neuberg (2005) reported that African and Hispanic Americans aroused fear and anger for White Americans.

Table 6
Hierarchical Linear Modeling Models for Testing Interaction Between Sex and Percent Women Incumbents

Model	Parameter							
	γ_{00}	γ_{10}	γ_{20}	γ_{30}	τ_{00}	τ_{11}	τ_{22}	τ_{33}
Model 9: Sex + percent female incumbents + Sex × Female Incumbents interaction predict quits L1: Logit($quit_{ij}$) = $\beta_{0j} + \beta_{1j}(sex_{ij}) + \beta_{2j}(jobsex_{jk}) + \beta_{3j}(interaction_{jk})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	−3.38*	0.27*	0.01*	−0.01*	0.26*	0.02*	0.0003*	0.0001*
Model 10: Sex + percent female incumbents + Sex × Female Incumbents interaction + control variables predict quits L1: Logit ($quit_{ij}$) = $\beta_{0j} + \beta_{1j}$ (sex) + $\beta_{2j}(jobsex_{jk}) + \beta_{3j}(interaction_{jk}) + \beta_{4j}(minority_{ij}) + \beta_{5j}(tenure_{ij}) + \beta_{6j}(tenure^2_{ij}) + \beta_{7j}(jobminority_{jk}) + \beta_{8j}(jobquits_{jk})$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}$ (workforce $size_j$) + γ_{02}(workforce $maturity_j$) + U_{0j} L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	−3.72*	0.23*	0.01	−0.01*	0.32*	0.03*	0.0002*	0.0001*

Note. N = 404,052. Sex is coded as 1 for women and 0 for men; jobsex is percent of women incumbents in a job k in firm j; interaction is interaction between sex and percent women incumbents; minority is coded as 1 for minorities, 0 for Whites. γ_{00} = intercept of Level (L) 2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}; τ_{33} = variance in L2 residual for models predicting β_{3j}. Model 10 estimated L2 parameters (e.g., γ_{40}, τ_{44}) for control variables, which are omitted but available on request.
* $p \leq .05$.

By contrast, White Americans felt pity toward Native Americans and no specific threat-related emotions toward Asian Americans.

Organizational- and economic-disparity research have sustained this racial pattern. Leonard and Levine (2006) noticed that White Americans abandon workplaces dominated by Hispanic Americans or African Americans. Heightened prejudice toward these groups conceivably underlies their inferior earnings, occupational status, and job security relative to that of White Americans (Brief, Butz, & Deitch, 2005; Frontzek & Johnson, 2003; James, 2000; McKinnon, 2003; Park & Sandefur, 2003; Ramirez & de la Cruz, 2003; Reeves & Bennett, 2003). By contrast, Asian Americans are often deemed a “model minority” given their economic success (Cheng, 1997; Chong & Kim, 2006; Dreher & Cox, 1996; Tang, 1997). Indeed, Leonard and Levine (2006) found that White American exits exhibit no sensitivity to the Asian American workforce share. Because they face more societal prejudice than Asian and Native Americans (Brief, Umphress, et al., 2005; Cottrell & Neuberg, 2005), we reason that African Americans and Hispanic Americans are most prone to quit White-majority corporations. In line with this argument, Leonard and Levine (2006) found that African Americans resign more when White employment expands, while a Ewing Marion Kauffman Foundation survey revealed that African American (26%) and Hispanic American (20%) men contemplate leaving to start their own businesses more so than do White American men (11%; Allers, 2005). Similarly, McKay and associates (2007) documented that African American managers

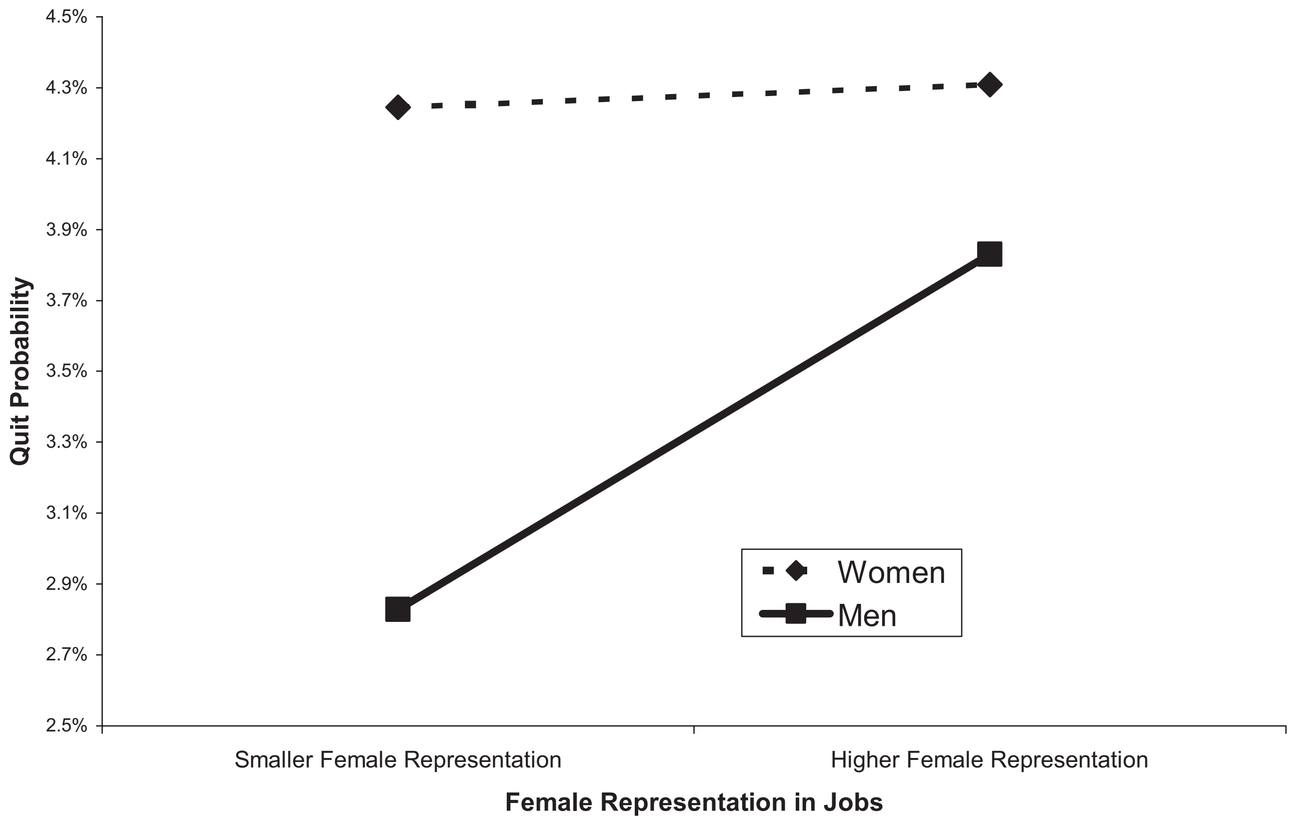


Figure 4. Interaction between sex and percent women job incumbents.

express stronger quit intentions than do White American managers. Given the preponderance of evidence for more pervasive prejudice toward African and Hispanic Americans than Asian and Native Americans, we put forth the following:

> *Hypothesis 6a:* African and Hispanic Americans quit more than White Americans.

> *Hypothesis 6b:* African and Hispanic Americans quit more than Asian and Native Americans.

Facing less hostility than other minorities, Asian and Native Americans are still outgroups to White Americans (Cottrell & Neuberg, 2005; Fiske et al., 2002). Asian Americans are regarded as successful minorities (Golden, 2006) who arouse envy (seen as "competent but cold" according to the "stereotype content model," Cuddy, Fiske, & Glick, 2004; Lin, Kwan, Cheung, & Fiske, 2005). By comparison, Native Americans are seen as "incompetent but warm" according to the stereotype content model (Cottrell & Neuberg, 2005; McKown & Weinstein, 2003), evoking pity and sympathy given past oppression by European settlers (Brown, 1970; Mann, 2005). These emotional reactions can also translate into discriminatory actions. Cuddy and associates (2004) noted that envious stereotypes can potentially erupt into resentment and hostility toward high-status outgroups, while paternalistic stereotypes can lead to social exclusion and patronizing treatment of low-status outgroups. In support, Lin and associates (2005) found that undergraduates who perceive Asians as competent but not socially warm avoid rooming with them, seldom socialize with them, and shun Asian American Studies courses.

Other findings disclose that Asians and Native Americans are disadvantaged relative to White Americans. Asian Americans earn lower returns on their educational investments, are underrepresented in management (but overrepresented in technical professions), must attain higher admission standards to enroll in prestigious colleges, and feel performance pressures to live up to their positive racial stereotypes (M. Bell, Harrison, & McLaughlin, 1997; Cheryan & Bodenhausen, 2000; Golden, 2006). Overshadowed by the model minority myth, most Southeast Asians and Pacific Islanders are poorly educated, underemployed, and perform low-paying menial jobs (Wong, Lai, Nagasawa, & Lin, 1998). Native Americans participate in the labor force at lower rates and have greater unemployment than do other racial groups (Clark & Weismantle, 2003; U.S. Census Bureau, n.d.). They also earn among the lowest median incomes in 2005 ($33,520 for men and $27,977 for women), with only Hispanic Americans having lower incomes (Webster & Bishaw, 2006). Those living near reservations fare worse with even higher joblessness and poorer wages (Gitter & Reagan, 2002; "Survey Shows," 2005). Given this body of evidence, we thus anticipate the following:

> *Hypothesis 6c:* Asian and Native Americans quit more than White Americans.

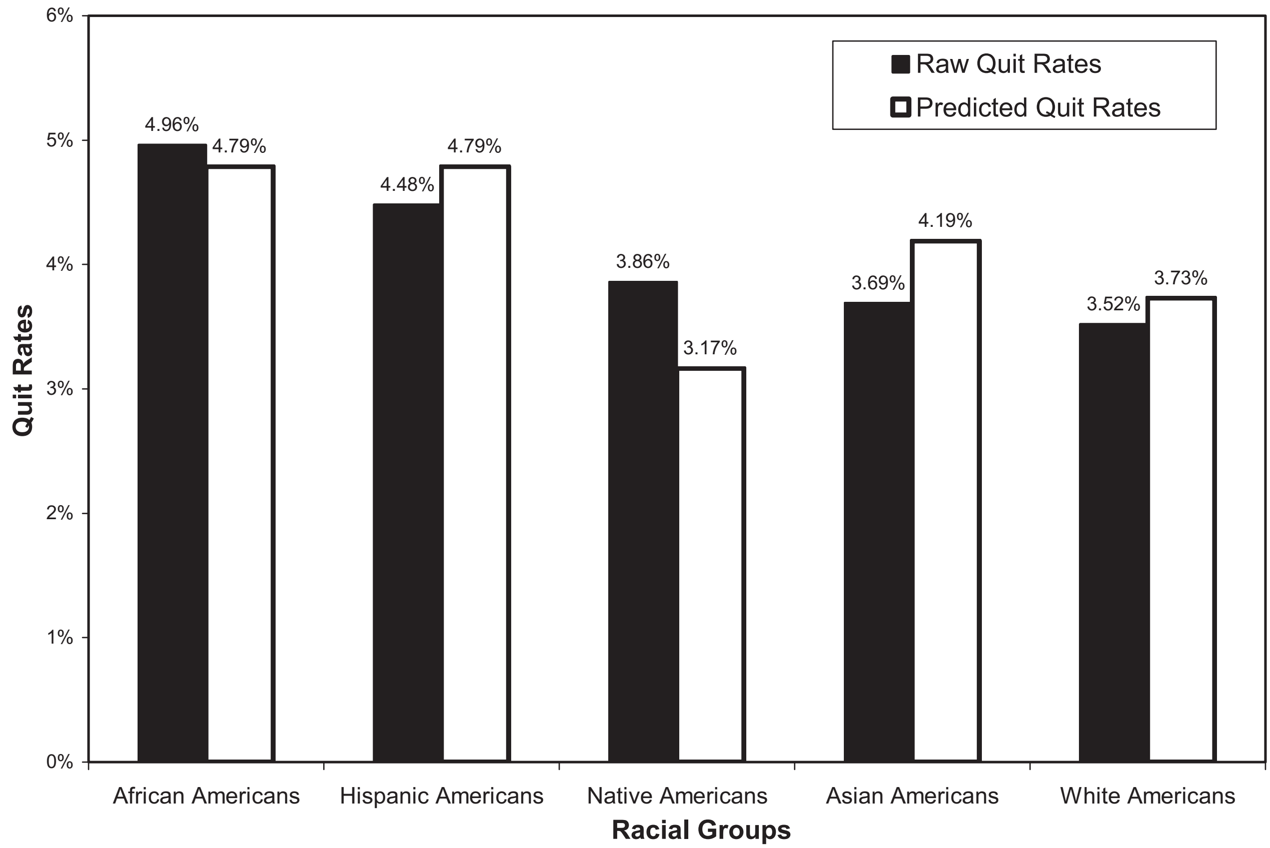


Figure 5. Racial differences in quit rates.

Double Jeopardy: Joint Race and Sex Effects on Turnover

Another unresolved question in the turnover literature is the joint effect of race and sex on quits in managerial and professional fields where women of color have historically been excluded (Dreher & Cox, 2000). Few scholars have addressed how both attributes simultaneously influence corporate withdrawal among exempt women of color, perhaps because they did not have enough representatives from this population (Berdahl & Moore, 2006). Decrying such neglect, hooks (1981) remarked that

> No other group in America has so had their identity socialized out of existence as have black women. We are rarely recognized as a group separate and distinct from black men, or a present part of the larger group "women" in this culture . . . When black people are talked about the focus tends to be on black men; and when women are talked about the focus tends to be on white women. (p. 7)

Similarly, Berdahl and Moore (2006) recounted how longstanding debates about who encounters the worse discrimination—women or minorities—overlook minority women's unique plight.

Our rationale for corporate flight among women of color rests on the "double jeopardy" hypothesis, which holds that "individuals who occupy the lowest position on two or more social categories will experience the most disadvantage of any group" (Browne & Misra, 2003, p. 493). Minority women should face more negative work outcomes than do their male counterparts and Whites of both sexes. Research on various indicators of outgroup disadvantage has sustained this proposition. To illustrate, a comparison between White American and African American men and women found that African American women are least represented in supervisory positions (Petrie & Roman, 2004). Other studies have shown that African and Hispanic American women, compared with other groups, earn the lowest wages (Barnum, Liden, & Ditomaso, 1995; Browne, 1999; King, 1988; Sanchez-Hucles, 1997) and occupy the most marginal jobs (Anderson & Shapiro, 1996; Berdahl & Moore, 2006; Kulis & Miller, 1988). What is more, Berdahl and Moore (2006) reported that women of color, relative to men of color and White American women, encounter the most gender and ethnic harassment at work. Finally, E. L. Bell and Nkomo (2001) observed that African American female professionals perceive greater barriers to advancement, tokenism, and perceive less work group acceptance than do White American female professionals.

Most double jeopardy studies have sampled African American and Hispanic American women, but women of other ethnicities may also face the twofold hardships of sexism and racism. Census data have indicated that, similar to African and Hispanic American women, Asian American women are the most occupationally segregated of all groups and the least upwardly mobile (Giscombe & Mattis, 2002). A Catalyst survey in 30 *Fortune* 1000 corporations finds that female Asian Americans (like African and Hispanic Americans) endure impediments to career progress, such as fewer high visibility assignments and role models of their race (Giscombe & Mattis, 2002). Moreover, Native American women are

Table 7
Hierarchical Linear Modeling Models for Testing Race Differences With Dummy Coding

Model	Parameter									
	γ_{00}	γ_{10}	γ_{20}	γ_{30}	γ_{40}	τ_{00}	τ_{11}	τ_{22}	τ_{33}	τ_{44}
Model 11: Race predicts quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{black}_{ij}) + \beta_{2j}(\text{asian}_{ij}) + \beta_{3j}(\text{hispanic}_{ij}) + \beta_{4j}(\text{native}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$ L2: $\beta_{4j} = \gamma_{40} + U_{4j}$	−3.25*	0.26*	0.12*	0.26*	−0.17	0.21*	0.04*	0.02*	0.06*	0.34*
Model 12: Race + control variables predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{black}_{ij}) + \beta_{2j}(\text{asian}_{ij}) + \beta_{3j}(\text{hispanic}_{ij}) + \beta_{4j}(\text{native}_{ij}) + \beta_{5j}(\text{sex}_{ij}) + \beta_{6j}(\text{tenure}_{ij}) + \beta_{7j}(\text{tenure}_{ij}^{2}) + \beta_{8j}(\text{jobminority}_{jk}) + \beta_{9j}(\text{jobquits}_{jk}) + \beta_{10j}(\text{jobsex}_{jk})$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_{j}) + \gamma_{02}(\text{workforce maturity}_{j}) + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$ L2: $\beta_{4j} = \gamma_{40} + U_{4j}$	−3.69*	0.01	−0.08	0.06	−0.33	0.38*	0.08	0.05	0.05*	0.59*

Note. $N = 404{,}052$. Sex is coded as 1 for women and 0 for men; black is coded as 1 for African Americans, 0 for others; asian is coded as 1 for Asian Americans 0 for others; hispanic is coded as 1 for Hispanic Americans, 0 for others; native is coded as 1 for Native Americans and 0 for others (Whites = reference group). γ_{00} = intercept of Level (L) 2 regression predicting β_{0j}; γ_{01} = workforce size slope of L2 regression predicting β_{0j}; γ_{02} = workforce maturity slope of L2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; γ_{40} = intercept of L2 regression predicting β_{4j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}; τ_{33} = variance in L2 residual for models predicting β_{3j}; τ_{44} = variance in L2 residual for models predicting β_{4j}. Model 12 estimated L2 parameters for control variables, which are omitted but available on request.
* $p \leq .05$.

less likely to graduate from college or find positions as professionals or managers, and, when they do, they usually earn the bottom of the pay scale (Reynolds, 2004). Native American female professionals and managers also undergo strain when balancing commitment to tribal culture with acculturation to corporate life (Muller, 1998).

Because women of color often suffer greater discriminatory experiences (e.g., sexual and ethnic harassment, tokenism, short career ladders, low pay) that induce turnover (Hom & Griffeth, 1995), they should exit more than men of color and White Americans of both sexes. All the same, minority women are rarely the most exit prone according to turnover research, though informal evidence has indicated otherwise (Giscombe & Mattis, 2002). Testing dual marginalization however has not been the prime focus for the bulk of these studies, which have rarely treated both sex and race as main effects in statistical analyses (Booth et al., 1999; Farber, 1994; Ippolito, 1991; Kulis & Miller, 1988; Light & Ureta, 1992; O. Mitchell, 1983; Noonan & Corcoran, 2004; Payne & Huffman, 2005; Sicherman, 1996; Valentine, 2001).

Our research improves on available evidence of corporate attrition among non-White women in high-status, well-paying jobs. Four studies directly sampled professional occupations: midcareer managers and professionals (Dreher & Cox, 2000), academic sociologists (Kulis & Miller, 1988), private-practice lawyers (Noonan & Corcoran, 2004), and military officers (Payne & Huffman, 2005). Excepting Dreher and Cox's (2000) sample, these occupations mainly reside outside corporate settings. Instead, the bulk of studies comprised broad national samples (Booth et al., 1999; Farber, 1994; Light & Ureta, 1992; Valentine, 2001) or corporate-wide workforces (Ippolito, 1991; Sicherman, 1996), combining professional with nonprofessional workers whose attritions differ (Royalty, 1998). More than this, extant tests may lack adequate samples of women of color, weakening findings for twin jeopardy. For example, minority women constituted 1.9% of participants in Kulis and Miller's (1988) study, though most authors failed to report their sample size (Light & Ureta, 1992; Payne & Huffman, 2005; Sicherman, 1996). Further, the sole demonstration that professional women of color depart (sociology faculty) positions at greater rates than do other groups combined voluntary with involuntary separations (Kulis & Miller, 1988). Dreher and Cox (2000) also failed to distinguish voluntary from involuntary exits, but their findings disputed double jeopardy for midcareer managers and professionals. To offset these shortcomings, this research samples over 30,000 minority women in a broad array of business professions and explicitly tests how sex and race jointly affects voluntary terminations. We also extend past research on women of color by assessing how dual discrimination generalizes across different races. Accordingly, we postulate the following:

> *Hypothesis 7:* Minority women quit most, relative to minority men and Whites of both sexes.

Beyond additive effects, some researchers have contended that racism and sexism together produce multiplicative effects

Table 8
Hierarchical Linear Modeling Models for Testing Racial Differences With Contrast Coding

Model	Parameter									
	γ_{00}	γ_{10}	γ_{20}	γ_{30}	γ_{40}	τ_{00}	τ_{11}	τ_{22}	τ_{33}	τ_{44}
Model 13: Race predicts quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{compare1}_{ij}) + \beta_{2j}(\text{compare2}_{ij}) + \beta_{3j}(\text{compare3}_{ij}) + \beta_{4j}(\text{compare4}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$ L2: $\beta_{4j} = \gamma_{40} + U_{4j}$	-3.15^{*}	0.28^{*}	0.01	0.29	0.12^{*}	0.22^{*}	0.13^{*}	0.14^{*}	0.39^{*}	0.01
Model 14: Race + control variables predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{compare1}_{ij}) + \beta_{2j}(\text{compare2}_{ij}) + \beta_{3j}(\text{compare3}_{ij}) + \beta_{4j}(\text{compare4}_{ij}) + \beta_{5j}(\text{sex}_{ij}) + \beta_{6j}(\text{tenure}_{ij}) + \beta_{7j}(\text{tenure}_{ij}{}^{2}) + \beta_{8j}(\text{jobminority}_{jk}) + \beta_{9j}(\text{jobquits}_{jk}) + \beta_{10j}(\text{jobsex}_{jk})$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_{j}) + \gamma_{02}(\text{workforce maturity}_{j}) + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$ L2: $\beta_{4j} = \gamma_{40} + U_{4j}$	-3.69^{*}	0.01	-0.08	0.06	-0.33	0.38^{*}	0.08	0.05	0.05^{*}	0.59^{*}

Note. $N = 404{,}052$. Sex is coded as 1 for women and 0 for men; compare1 is coded as 1/2 for African and Hispanic Americans, −1/2 for Asian and Native Americans, and 0 for White Americans; compare2 is coded as 1/2 for African Americans, −1/2 for Hispanic Americans, and 0 for Asian, Native, and White Americans; compare3 is coded as 1/2 for Asian Americans, −1/2 for Native Americans, and 0 for all other racial groups; compare4 is coded as 1/5 for each racial minority and −4/5 for Whites. γ_{00} = intercept of Level (L) 2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; γ_{40} = intercept of L2 regression predicting β_{4j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}; τ_{33} = variance in L2 residual for models predicting β_{3j}; τ_{44} = variance in L2 residual for models predicting β_{4j}. Model 14 estimated L2 parameters for control variables, which are omitted but available on request.
[a] For model convergence, effect of tenure = squared is fixed (i.e., $\tau_{77} = 0$).
[*] $p \leq .05$.

and denote the extra disadvantage of possessing both devalued attributes simultaneously as "aggravated" or "multiple" jeopardy (King, 1988; Kulis & Miller, 1988). As Berdahl and Moore (2006) explained, "the disadvantages of race and sex compound or multiply each other, making the detrimental effect of both belonging to an ethnic minority and being a woman greater than the additive hypothesis would suggest" (p. 428). In support, Kulis and Miller (1988) observed that minority women sociologists are severely underrepresented in small academic departments given proportions of women and minorities in the academic sociology workforce. By contrast, Barnum and associates (1995) and Berdahl and Moore (2006) documented no interaction between sex and ethnicity on workplace harassment or pay. Such uneven, scant findings demand more clarification. We build on this research stream by exploring the interactive effect of sex and ethnicity on corporate attrition and its racioethnic generalizability (Berdahl & Moore, 2006). Consequently, we test the proposition below:

> *Hypothesis 8:* Gender and race have multiplicative effects on voluntary terminations, such that women of color quit more than do men of color and Whites of both sexes.

Method

Participants

Inaugurated during the height of the "war for talent" (Bodden, Glucksman, & Lasky, 2000), 25 *Fortune* 500 corporations initially formed the ARC consortium to annually share information on attrition rates among (domestic and Canadian) professional and managerial employees by using standard definitions and agreed-on protocols. ARC provides benchmarking statistics for participating firms and hosts an annual conference to review trend data and share effective retention practices. Firms belonging to ARC or submitting data vary from year to year. ARC membership is restricted to companies that have 5,000 or more full-time, permanent employees in North America, excluding human resources consulting firms.

During 2004, 20 member corporations contributed 2003 turnover statistics on 19 occupational fields by firm seniority (see Figure 3). Sixteen of these contributors belonged to the 2003 *Fortune* 500 list. These organizations represent various industries including aerospace and defense, automotive, chemicals, computer software and services, construction, electronics, financial services, insurance, manufacturing, pharmaceuticals, telecommunications,

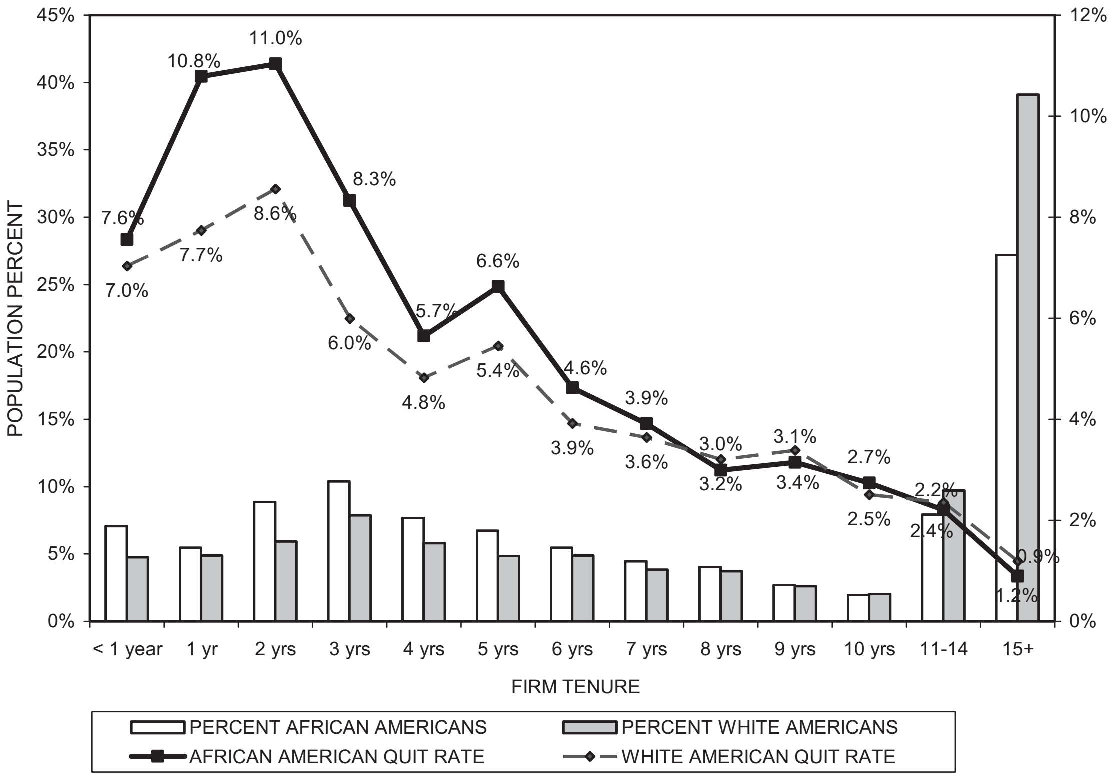


Figure 6. African American and White American differences in quit rates as a function of firm seniority.

and transportation services. The 2003 firm participants employed an average of 23,772.9 professionals and managers, though they also employed abundant nonexempt workers. Our analyses were limited to the combined professional and managerial workforce of participating firms, which totaled 475,458. Of these firms, 12 contributed data on quit rates of their "best performers" (defined as the highest rating category in a firm's appraisal system; total N = 362,063; highest performers = 44,656 or 12.33%), while 16 firms provided data about quit rates for various gender and racial subgroups (N = 404,052). Ten companies provided both demographic and high performer data. Two firms supplied high performer data but no demographic data, while 4 others shared demographic data but no high performer data.

Over 48% of workforce from the 20 firms had been employed by their company for 10 or more years. Table 1 shows demographic composition of the 16 firms submitting quit statistics for demographic subgroups. Of their work force, 79.82% was White American; African Americans made up another 6.95% of this population, Asian Americans another 8.50%, Hispanic Americans 4.21%, and Native Americans 0.52%. Men represented 71.7% of the professional workforce, while women represented 28.3%.

Measurement of Quit Rates

On an Excel spreadsheet with rows defined by jobs and columns by seniority (ranging from less than a year to 15 or more years), ARC companies report the base population for various job and seniority combinations, though the two most senior categories represent broader ranges (11–14 and 15-plus years). ARC defines the base population as all employees who were regular (full- or part-time), active, or on leave at any time during the calendar (2003) year plus all employees who terminated employment during 2003 for any reason (e.g., death, discharge, retirement, layoff, or quits). In a second, adjoining Job × Tenure matrix, firms record the number of quits for each job–tenure cell. Quits represent any voluntary, employee-initiated termination of employment but exclude involuntary attrition (e.g., deaths, discharges, layoffs, and retirements). All 20 participating firms completed and submitted this spreadsheet for their entire workforce. At their discretion, ARC members provide additional spreadsheets for subpopulations. Twelve firms supplied a spreadsheet for their best performers. (For performance comparisons, we computed quit rates for lower—actually "nonhigh"—performers by computing their base Ns and leaver Ns by subtracting a firm's top performer Ns from Ns for its whole population.) Sixteen firms furnished a spreadsheet for each sex–race subgroup (e.g., Asian American men; 10 subgroups total). Performance data and demographic data are not matched, precluding analyses into demographic differences in performance. Tabular data were later converted to individual-level data for statistical analyses.

Statistical Analyses

Because ARC data are nested within firms, we used logistic regression in Raudenbush and Bryk's (2002) hierarchical linear modeling (HLM) program given a binary outcome variable. (Note Level 1 residual variances—or σ^2—are not estimated by HLM logistic regression; Raudenbush, Bryk, Cheong, Congdon, & du

Table 9
Hierarchical Linear Modeling Models for Exploring Explanations for Racial Differences Among Short-Tenure Employees

Level 1 predictors of turnover	Level 2 intercepts for Level 1 predictors			
	Model without control variables	Model with tenure predictors	Model with tenure and percent Asian American concentration in jobs	Model with tenure and percent African and Hispanic Americans in jobs
Black γ_{10}	.18*	.12*	.13*	.09
Asian γ_{20}	−.22*	−.26*	−.09	−.16*
Hispanic γ_{30}	.04	.04	.03	−.001
Native γ_{40}	−.26	−.28	−.45	−.46
Tenure γ_{50}		−.03	−.01	−.01
Tenure^2 γ_{60}		−.01	−.02	−.02
Percent Asian Americans in job γ_{70}			−.04*	
Percent African and Hispanic Americans in job γ_{80}				.01
Historic attrition rate in job γ_{90}			8.30*	9.80*

Note. $N = 170{,}724$. Black is coded as 1 for African Americans, 0 for others; Asian is coded as 1 for Asians, 0 for others; Hispanic is coded as 1 for Hispanics, 0 for others; Native is coded as 1 for Native Americans and 0 for others (Whites = reference group). γ_{10} = intercept of Level (L) 2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; γ_{40} = intercept of L2 regression predicting β_{4j}; γ_{50} = intercept of L2 regression predicting β_{5j}; γ_{60} = intercept of L2 regression predicting β_{6j}; γ_{70} = intercept of L2 regression predicting β_{7j}; γ_{80} = intercept of L2 regressions predicting β_{8j}; γ_{90} = intercept of L2 regressions predicting β_{9j}. No L2 predictors were included.
* $p \leq .05$.

Toit, 2004.) Though primarily interested in effects of lower rather than higher level predictors, modeling nonindependent data offsets the reduction in statistical power of conventional tests of individual-level relationships in grouped data (Bliese & Hanges, 2004). Consequently, HLM results should be accorded more credence than should other supplementary statistical tests that ignore nonindependence (see below). For each hypothesis, we estimated an unconditional random-coefficients regression model that includes theorized predictors in a Level 1 equation predicting quits but does not include Level 2 predictors. Representing the mean quit–predictor slope across firms, the Level 2 intercept (e.g., γ_{10}) in the Level 2 equation predicting this Level 1 slope (e.g., β_{1j}) tests the hypothesis. No control variables were initially included because we seek results that confirm (or disconfirm) previous meta-analytical and industry reports based on "simple" bivariate relationships. For additional comparability with prior bivariate statistics that ignore nested data structures, we also computed Cramer's V (to assess tenure–quit relationship), phi coefficient (to assess performance–quit correlation), and Z tests of group differences in proportions (to assess demographic differences in turnover).

We next repeated HLM tests with Level 1 and Level 2 control variables (estimating random intercept-only models; Raudenbush & Bryk, 2002) because predictor effects may shrink or disappear with statistical controls (Booth et al., 1999; Sicherman, 1996). Given their influence on quits (Haveman & Cohen, 1994; Mobley, 1982; Shaw, Delery, Jenkins, & Gupta, 1998), all HLM tests controlled firm workforce size and "workforce maturity" (percent of employees with 10 or more years of service) when predicting firm-level quit rates (or Level 1 intercepts) in Level 2 regressions. Apart from these Level 2 control variables, we computed mean quit rates for jobs in the previous year (2002) for each firm to control differential job opportunities across occupations in Level 1 equations (cf. Steel, 1996). The HLM test of the interaction between performance and past job attrition also included tenure as a Level 1 control variable given its association with quits (Griffeth et al., 2000).

HLM tests of sex and race effects further controlled job differences in minority and female representation (which vary across firms; besides historic job attrition and tenure) in Level 1 equations given their impact on minority, female, and White male quits (Lyness & Judiesch, 2001; Sacco & Schmitt, 2005; Tsui, Egan, & O'Reilly, 1992; Wasti et al., 2000). For racial comparisons, we used dummy coding to contrast each minority with Whites and contrast coding to compare African and Hispanic Americans with Asian and Native Americans (J. Cohen, Cohen, West, & Aiken, 2003). Additionally, we controlled minority status when assessing sex effects (as minority status might increase female quits more than male quits) and controlled sex when assessing race effects (as women's greater exit propensity may magnify racial differences in quits; Booth et al., 1999). We used grand-mean centering for all continuous predictors excepting nominal variables (Hofmann, Griffin, & Gavin, 2000; Hofmann & Gavin, 1998).

Results

Table 2 reports intercorrelations among variables. Turnover is negatively related to tenure and job performance, although women quit more. Moreover, historical attrition rate in a job, percent of women in a job, and percent of minorities in a job are positively associated with quits.

Curvilinear Tenure–Turnover Relationship

According to Cramer's V, we found a significant bivariate relationship between tenure and turnover ($V = .13$, $p < .05$). Figure 1 further reveals that quits steadily decline after the 1st year of employment. To test Hypothesis 1, we also carried out a logistic regression predicting turnover from firm tenure and a quadratic tenure term. Model 1 in Table 3 corroborates this curvilinear trend

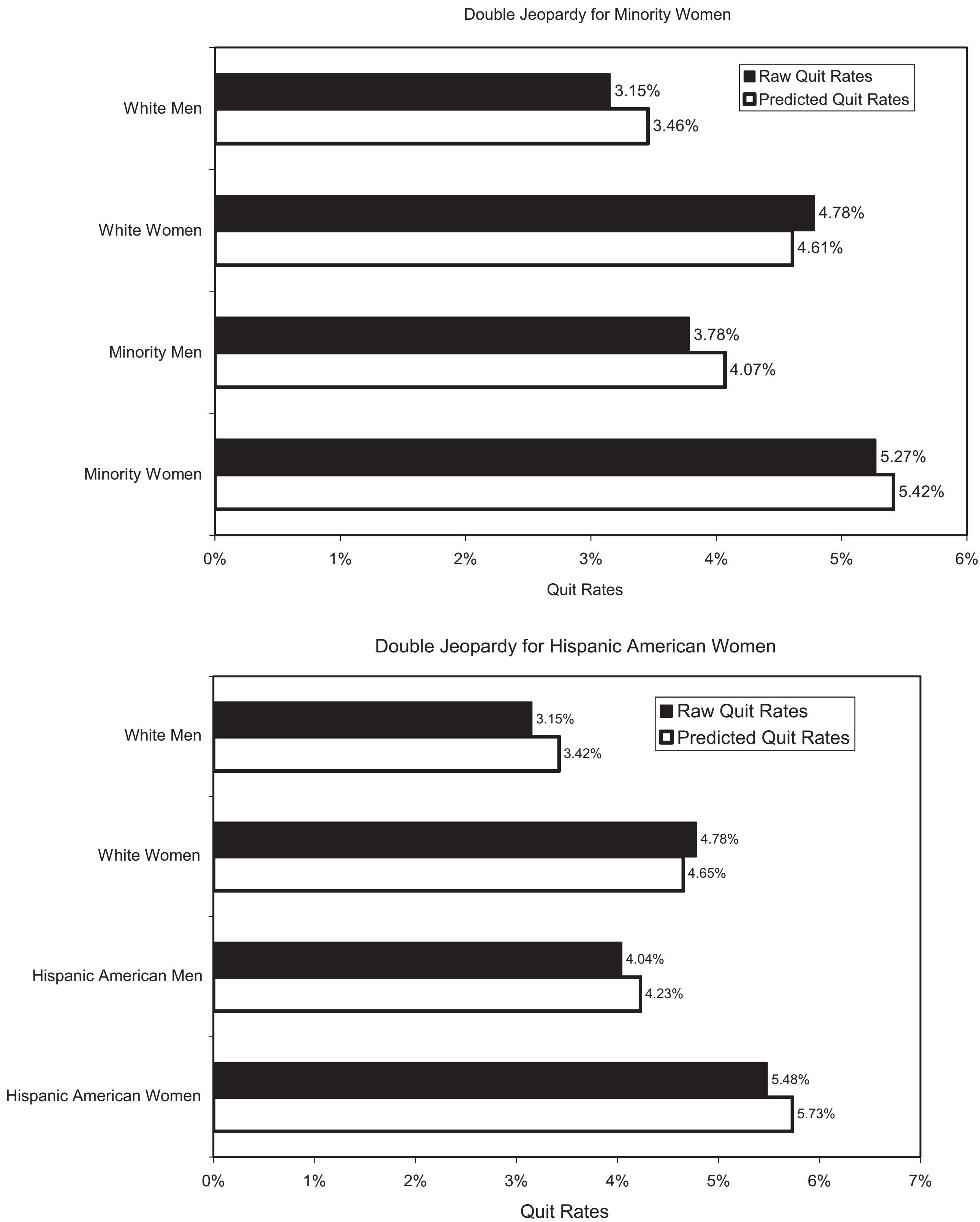
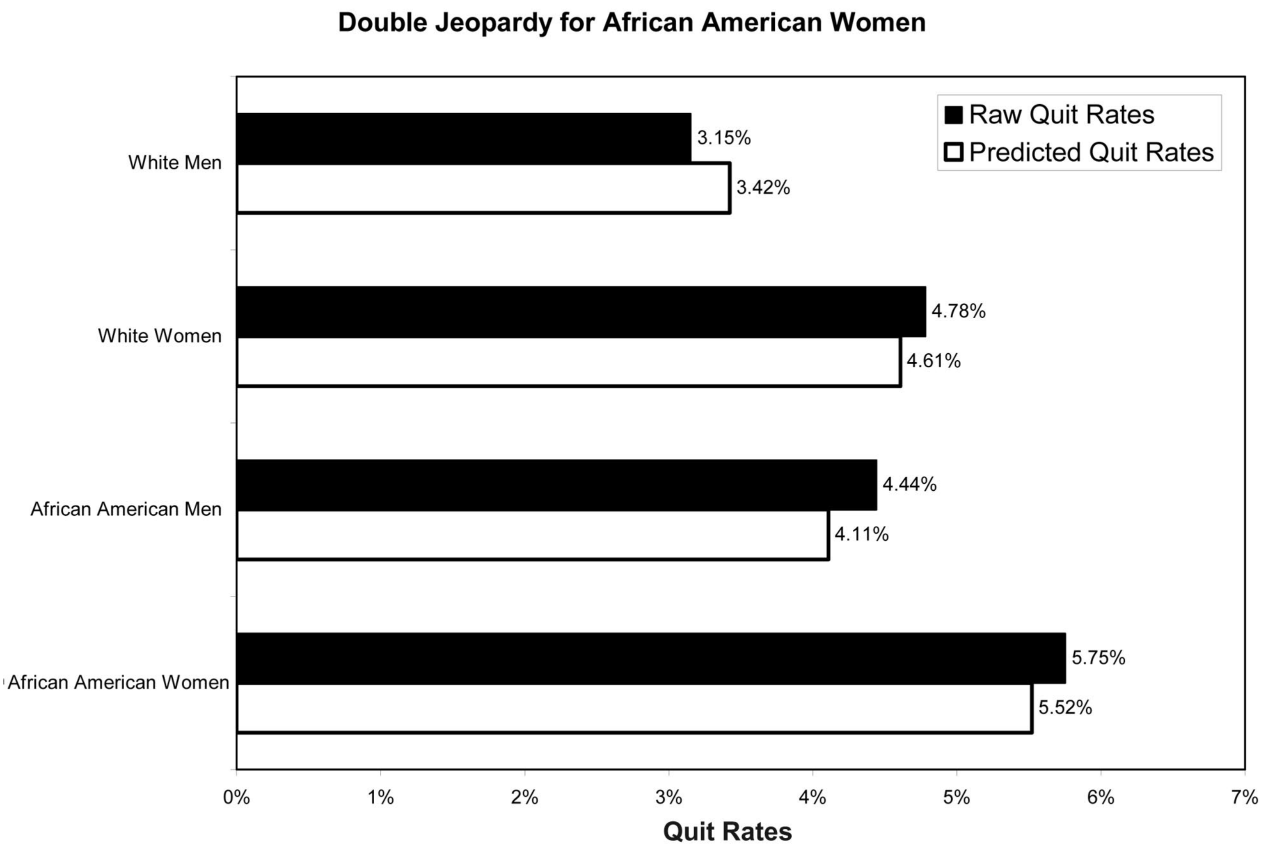


Figure 7. Quit rates between minority women, minority men, White men, and White women.



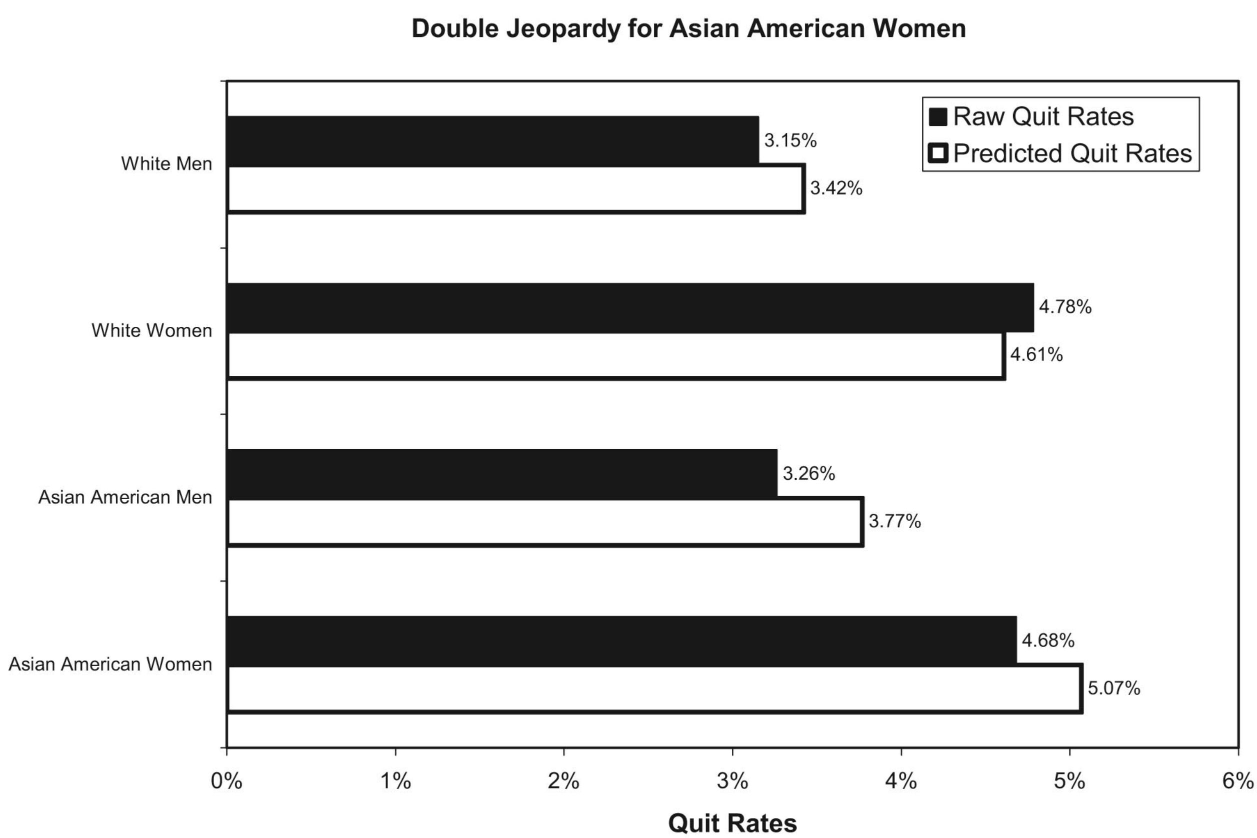


Figure 7. (*Continued*)

Table 10
Hierarchical Linear Modeling Models for Testing Double Jeopardy for Minority and African American Women

Model	Parameter Estimate γ_{00}	γ_{10}	γ_{20}	τ_{00}	τ_{11}	τ_{22}
Model 15: Sex + minority status predicts quits[a] L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{minority}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$	−3.33*	0.30*	0.17*	0.23*	0.03*	0.02*
Model 16: Sex + African American predict quits[b] L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{black}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$	−3.34*	0.31*	0.19*	0.23*	0.03*	0.06*
Model 17: Sex + African American + control variables predict quits[b] L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{black}_{ij}) + \beta_{3j}(\text{tenure}_{ij}) + \beta_{4j}(\text{tenure}_{ij}^2) + \beta_{5j}(\text{jobsex}_{jk}) + \beta_{6j}(\text{jobblack}_{jk}) + \beta_{7j}(\text{jobquits}_{jk})$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_j) + \gamma_{02}(\text{workforce maturity}_j) + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$	−3.71*	0.21*	0.004	0.19*	0.04	0.06

Note. Sex is coded as 1 for women and 0 for men; minority is coded as 1 for minorities and 0 for whites; black is coded as 1 for African Americans and 0 for White Americans; jobblack_{jk} is the percentage of African American incumbents in job k in firm j. γ_{00} = intercept of Level (L) 2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}. Model 17 estimated L2 parameters for control variables, which are omitted but available on request.
[a] $N = 404{,}052$. [b] $N = 350{,}575$.
* $p \leq .05$.

(tenure^2 $\gamma_{20} = -0.01$, $p < .05$). However, we did not verify the cubic relationship predicted by Hypothesis 2 (tenure^3 $\gamma_{30} = -0.0002$, $p > .05$; Model 2 in Table 3). When control variables were included (Model 3 in Table 3), the quadratic tenure term remained significant ($\gamma_{20} = -0.01$, $p < .05$). In line with Steel's (1996) finding, the prior year's rate of attrition in a job predicted quits in the current year ($\gamma_{30} = 6.94$, $p < .05$). In short, these findings uphold a quadratic rather than cubic relationship between tenure and quits, sustaining Meglino et al.'s (2000) job attraction and Lazear's (1998) human capital formulations.

Moderators of Job Performance–Turnover Association

Reinforcing previous meta-analyses (Griffeth et al., 2000), we found that performance was negatively related to quits ($\phi = -.01$, $p < .05$). Testing Hypothesis 3, an HLM test found a significant interaction between performance and historical quit rates in jobs ($\gamma_{30} = -11.62$, $p < .05$; Model 4 in Table 4). After adding control variables, the Performance × Past Job Attrition interaction remained significant ($\gamma_{50} = -9.43$, $p < .05$; see Model 5 in Table 4). In line with past work (Williams & Livingstone, 1994), being among the best performers in a company lowers the odds of departures by 50.84% ($\gamma_{10} = -0.71$, $p < .05$). To plot the interaction, we used moderator scores +1 *SD* above and –1 *SD* below the mean historic job attrition rate for high and nonhigh performers (J. Cohen et al., 2003; Raudenbush & Bryk, 2002). Sustaining McEvoy and Cascio's (1987) prediction, Figure 2 shows a stronger performance–turnover relationship when jobs have high historical quit rates, implying that plentiful employment opportunities allow low performers to exit. Unlike ambiguous findings based on unemployment statistics (McEvoy & Cascio, 1987; Williams & Livingstone, 1994), using a proxy based on historic attrition rates in jobs verified McEvoy and Cascio's (1987) longstanding thesis about unemployment moderation.

Sex Differences in Quit Rates

According to a *Z* test of group differences in proportions, women (4.88%) quit more than men (3.20%) at a 52.2% higher rate ($Z = 25.44$, $p < .05$). Controlling nonindependency among observations from the same firm (Bliese & Hanges, 2004), an HLM test showed that women quit at a 36.34% higher rate than did men ($\gamma_{10} = 0.31$, $p < .05$; Model 6 in Table 5). From this HLM equation (Raudenbush & Bryk, 2002), we computed "predicted" quit rates for women (4.74%) and men (3.52%), which yield a more accurate index of sex differences by controlling for nonindependent data. An HLM test with individual- and firm-level control variables reaffirmed this gender gap ($\gamma_{10} = 0.19$, $p < .05$; Model 8 in Table 5). To better capture occupational differences, we performed logistic regression outside HLM to include dummy-coded variables to represent occupational differences. HLM cannot handle the same dummy-coded job variables because some job categories were nonexistent in companies, and their absence would yield dummy variables without variance (i.e., all study participants

Table 11
Hierarchical Linear Modeling Models for Testing Double Jeopardy for Hispanic American Women

Model	Parameter Estimate					
	γ_{00}	γ_{10}	γ_{20}	τ_{00}	τ_{11}	τ_{22}
Model 18: Sex + Hispanic American predict quits L1: Logit (quit$_{ij}$) = β_{0j} + β_{1j}(sex$_{ij}$) + β_{2j}(hispanic$_{ij}$) L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$	−3.34*	0.32*	0.22*	0.24*	0.04*	0.06*
Model 19: Sex + Hispanic American + control variables predict quits L1: Logit(quit$_{ij}$) = β_{0j} + β_{1j}(sex$_{ij}$) + β_{2j}(hispanic$_{ij}$) β_{3j}(tenure$_{ij}$) + β_{4j}(tenure$_{ij}{}^2$) + β_{5j}(jobsex$_{kj}$) + β_{6j}(jobhispanic$_{kj}$) + β_{7j}(jobquits$_{kj}$) L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}$(workforce size$_j$) + γ_{02}(workforce maturity$_j$) + U_{0j} L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$	−3.67*	0.21*	0.01	0.19*	0.01*	0.03*

Note. N = 339,528. Sex is coded as 1 for women and 0 for men; Hispanic is coded as 1 for Hispanic Americans and 0 for White Americans; jobhispanic$_{jk}$ is the percent of Hispanic American incumbents in a job k in firm j. γ_{00} = intercept of Level (L) 2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}. Model 19 estimated L2 parameters for control variables, which are omitted but available on request.
* $p \leq .05$.

would have zero scores). To overset this limitation, we ran logistic regression separately for each firm. In each firm, we computed a (firm-specific) set of dummy job variables and regressed quits onto this set and other individual-level control variables and gender. These tests showed that the sex disparity in quits persisted even when job differences were more fully controlled. Collectively, these tests sustain Hypothesis 4, demonstrating higher corporate attrition among women.

Revealing occupational moderation, Figure 3 shows uneven gender differences in turnover for various jobs, while Z tests revealed that women quit more than men in 9 of the 19 jobs (i.e., all other exempt jobs, customer service, finance, information technology, marketing, sales, manufacturing/operations, product engineering, software engineering). Sex differences were nonsignificant in the other 10 positions. For greater clarification, we next tested the interaction between sex and extent of female incumbency in jobs with HLM (cf. Zatzick et al., 2003). (We note that ARC jobs do vary in female concentration: SD = 15.84% and M = 28.25%.) Model 9 in Table 6 discloses a significant interaction ($\gamma_{30} = -0.01$, $p < .05$). Generating predicted quit probabilities from HLM regression, we plotted this interaction for men and women who work in jobs with more (one SD above the mean female proportion in jobs) or less (one SD below the mean) women incumbents (see Figure 4). This plot partly sustained Hypothesis 5. In line with Tsui et al.'s (1992) and Chatman and O'Reilly's (2004) tests, men more readily leave when women increasingly populate jobs. Opposing earlier findings (Elvira & Cohen, 2001; Reskin, McBrier, & Kmec, 1999; Sacco & Schmitt, 2005), women quits did not fall when more members of their sex enter an occupation. This interaction ($\gamma_{30} = -0.01$, $p < .05$; Model 10 in Table 6) stayed significant when HLM controlled other extraneous influences.

Racial Differences in Turnover

A statistical test of group differences in proportions concluded that minorities (4.30%) quit more than Whites (3.52%; Z = 10.51, $p < .05$). Similarly, HLM indicated that minorities are more exit prone ($\gamma_{10} = 0.20$, $p < .05$; Model 7 in Table 5); minority status boosts turnover odds by 22.14%. Comparing different minorities, Z tests indicated that African (4.96% quit rate) and Hispanic (4.48%) Americans exit more than White Americans (see "raw" quit rates in Figure 5). Another HLM test (Model 11 in Table 7) found that all racial minorities exhibit greater turnover (African American $\gamma_{10} = 0.26$, $p < .05$, 29.69% higher quit rate than White Americans; Asian American $\gamma_{20} = 0.12$, $p < .05$, 12.75% higher; Hispanic American $\gamma_{30} = 0.26$, $p < .05$, 29.69% higher) excepting Native Americans ($\gamma_{40} = -0.17$, $p > .05$). Figure 5 reports their "predicted" quit rates with the HLM equation that takes into account the nested data structure (Bliese & Hanges, 2004). Importantly, higher minority quit rates relative to those of Whites vanished once the HLM test included control variables (see Model 12 in Table 7), repudiating Hypotheses 6a and 6c.

To evaluate Hypothesis 6b, we compared racial minorities with one another with contrast coding. HLM tests showed that African and Hispanic Americans quit more than Asian and Native Americans ($\gamma_{10} = 0.28$, $p < .05$; Model 13 in Table 8). Meanwhile, African and Hispanic Americans did not differ in turnover (γ_{20} = 0.01, $p > .05$), nor did Asian and Native Americans ($\gamma_{30} = 0.29$, $p > .05$). Disputing Hypothesis 6b, quit differences among minor-

Table 12
Hierarchical Linear Modeling Models for Testing Double Jeopardy for Asian American Women

Model	Parameter Estimate γ_{00}	γ_{10}	γ_{20}	τ_{00}	τ_{11}	τ_{22}
Model 20: Sex + Asian American predict quits L1: Logit $(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{asian}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$	-3.34^*	0.31^*	0.10^*	0.24^*	0.03^*	0.01^*
Model 21: Sex + Asian American + control variables predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{asian}_{ij}) + \beta_{3j}(\text{tenure}_{ij}) + \beta_{4j}(\text{tenure}_{ij}^{2}) + \beta_{5j}(\text{jobsex}_{jk}) + \beta_{6j}(\text{jobasian}_{jk}) + \beta_{7j}(\text{jobquits}_{jk})$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_j) + \gamma_{02}(\text{workforce maturity}_j) + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$	-3.90^*	0.20^*	-0.07	0.50^*	0.03	0.05

Note. $N = 356{,}855$. Sex is coded as 1 for women and 0 for men; asian is coded as 1 for Asian Americans and 0 for White Americans; jobasian_{jk} is the percent of Asian American incumbents in a job k in firm j. γ_{00} = intercept of L2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}. Model 21 estimated L2 parameters for control variables, which are omitted but available on request.
$^* p \leq .05$.

ities disappeared when control variables were included in the HLM regression (Model 14 in Table 8), refuting the expectation that African and Hispanic Americans exit more than Asian and Native Americans.

Additional HLM tests ascertained that company tenure primarily accounted for the apparently higher turnover among African, Hispanic, and Asian Americans relative to White Americans. These analyses showed that these minorities were concentrated in "higher risk" short-tenure categories, where most attrition occurs (Griffeth et al., 2000). Statistically controlling only tenure and tenure^2 in HLM equations thus "erased" racial gaps in attrition. To illustrate this "confound," Figure 6 presents quit rates for African and White Americans by firm tenure as well as proportion of each race in each tenure category (see bar charts). As is evident, African Americans leave more often than White Americans, especially during early employment. Yet 51.7% of African Americans belong to short-tenure categories (0–6 years) compared with 39% of White Americans. As a result, holding constant such group differences in tenure abolished the turnover gap between these populations. All told, these post hoc tests conclude that minorities are leaving at elevated rates primarily because their ranks include disproportionately more newcomers who are coping with assimilation pressures and learning if the job is the right match for them (Meitzen, 1986).

All the same, Figure 6 suggests more pronounced quit disparities between short-tenure African Americans and White Americans that are not fully explained by tenure differences. We thus carried out additional post hoc HLM tests to explore factors behind racial differences in quits for employees with 6 or less years of service. Table 9 shows racial dissimilarities for short-tenure incumbents: African Americans leave more than White Americans, but Asian Americans exit less. These racial differences remained even when controlling tenure in HLM. To further pinpoint why these two minorities vary in quits, we investigated proportional representation of each minority in jobs as a potential mediator of the race–turnover relationship.

Following Baron and Kenny's (1986) mediation preconditions, we regressed minority concentration in jobs onto racial membership in HLM Level 1 equations. HLM regressions established that beginning Asian Americans tend to hold jobs that include more members of their race (e.g., 39.2% were product and software engineers), while short-tenure African Americans more often occupy jobs employing more African and Hispanic Americans (e.g., 55.5% of African Americans work in sales, customer service, information technology, and "all other exempt" jobs though White Americans numerically dominate all professions). Next, HLM tests demonstrated differential quit rates for these jobs, showing less turnover for all incumbents (both non-Whites and Whites) in jobs employing more Asian Americans but showing more turnover in jobs having more African and Hispanic Americans. Finally, we added these mediators to Level 1 HLM equations predicting quits, including historical job attrition as a control variable given its potential relationship to racial composition in jobs (though past attrition rate did not vary by race). Table 9 reveals that adding Asian American incumbency in jobs eliminated the effect of Asian American ethnicity on quits, while adding African and Hispanic American representation in jobs nullified the impact of African American ethnicity.

Testing Joint Effects of Race and Sex on Turnover

The quit rate for minority women is higher than that for White men (69.9% or [5.25% – 3.09%]/[3.09%]) and minority men (40.2% higher) as well as White women (10.6% higher) according to *Z* tests (performed on "raw quit rates" in Figure 7). We first tested the double jeopardy thesis with all minorities

Table 13
Hierarchical Linear Modeling Models for Testing Double Jeopardy for Native American Women

Model	Parameter Estimate					
	γ_{00}	γ_{10}	γ_{20}	τ_{00}	τ_{11}	τ_{22}
Model 22: Sex + Native American predict quits L1: Logit (quit_{ij}) = $\beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{native}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$	−3.35*	0.33*	−0.08	0.25*	0.05*	0.22*
Model 23: Sex + Native American + control variables predict quits L1: Logit(quit_{ij}) = $\beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{native}_{ij}) + \beta_{3j}(\text{tenure}_{ij}) + \beta_{4j}(\text{tenure}_{ij}^{2}) + \beta_{5j}(\text{jobsex}_{jk}) + \beta_{6j}(\text{jobnative}_{jk}) + \beta_{7j}(\text{jobquits}_{jk})$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_j) + \gamma_{02}(\text{workforce maturity}_j) + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$	−3.74*	0.21*	−0.34	0.26*	0.04	0.40

Note. N = 324,603. Sex is coded as 1 for women and 0 for men; native is coded as 1 for Native Americans and 0 for White Americans; jobnative_{jk} is the percent of Native American incumbents in job k in firm j. γ_{00} = intercept of Level (L) 2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}. Model 23 estimated L2 parameters for control variables, which are omitted but available on request.
* $p \leq .05$.

combined. An HLM test found significant effects for gender and minority status (sex $\gamma_{10} = 0.30$, $p < .05$; minority status $\gamma_{20} = 0.17$, $p < .05$; see Model 15 in Table 10). On the basis of this logistic regression, we found that minority women leave at a 56.75% higher rate than do White men. The HLM logistic regression that included control variables however found minorities in general were not more apt to leave than were Whites ($\gamma_{20} = -0.01$, $p > .05$ in Model 8 in Table 5).

Because racial minorities have dissimilar quit propensities, we repeated these tests with each minority group. *Z* tests (on raw quit rates in Figure 7) showed that the African American female turnover rate was 32.8% higher than the African American male quit rate, 20.6% higher than the White American female quit rate, and 85.3% higher than the White American male quit rate. HLM logistic regression further demonstrated that being a woman raises the quit odds by 36.34%, while being African American elevates quit odds by 20.92% (Model 16 in Table 10). In combination, these demographic effects indicated that African American women quit at a 61.29% greater rate than that of White American men. All the same, the HLM logistic regression with control variables rejected double jeopardy for African American women (see Model 17 in Table 10).

Similarly, we found Hispanic American women quit more than Hispanic American men (36.2% higher), White American women (14.9% higher), and White American men (76.6% higher) based on *Z* tests (see Figure 7). HLM logistic regression also showed that Hispanic American women exit at a 67.50% higher rate than that of White American men (sex $\gamma_{10} = 0.32$, $p < .05$; Hispanic status $\gamma_{20} = 0.22$, $p < .05$ in Model 18 in Table 11). Once again, an HLM regression controlling for other turnover influences disconfirmed the twin jeopardy proposition for Hispanic American women (Model 19 in Table 11).

Z tests found that Asian American women quit more than White American men (50.7% higher) and Asian American men (43.3% higher) but quit less than White American women (1.9% less; see Figure 7). More definitively, HLM model (No. 20) in Table 12 estimated significant sex ($\gamma_{10} = 0.31$, $p < .05$) and race ($\gamma_{20} = 0.10$, $p < .05$) effects. This equation projects that Asian women exit at a 48.11% higher rate than that of White men. All the same, adding control variables to the HLM analysis (Model 21 in Table 12: sex $\gamma_{10} = 0.20$, $p < .05$; race $\gamma_{20} = -0.07$, $p > .05$) disproved dual discrimination for female Asian Americans.

Finally, the Native American women turnover rate (4.13%) did not differ from those of White Americans or Native American men (3.71%) according to *Z* tests (see Figure 7). Rejecting Hypothesis 8, HLM logistic regression detected no ethnicity effect for Native Americans ($\gamma_{20} = -0.08$, $p > .05$), finding a sex effect ($\gamma_{10} = 0.33$, $p < .05$; Model 22 in Table 13). Similarly, HLM regression with control variables detected no joint sex and race effects (Model 23 in Table 13).

We further explored which control variables undergirded dual marginalization evidence for African, Hispanic, and Asian American women. Post hoc HLM tests determined that minority women have less seniority than do men of their own race, White women, and men and that controlling such tenure differences was enough to undermine support for this thesis. To illuminate this finding, we show in Figure 8 that short-tenure African American women exit more than African American men and White Americans of both sexes and that they overpopulate (as a proportion of their population) short-tenure categories relative to the other three groups.

To ascertain whether this conclusion holds for short-tenure women of color, we carried out more HLM tests with employees with 6 or less years of tenure. Table 14 suggests evidence for twin jeopardy for short-tenure African American women. Yet this finding was nullified when tenure was again controlled because these women are still overrepresented in the 2nd through 8th year of employment compared with African American men and White Americans (see Figure 8). By contrast, another statistical test

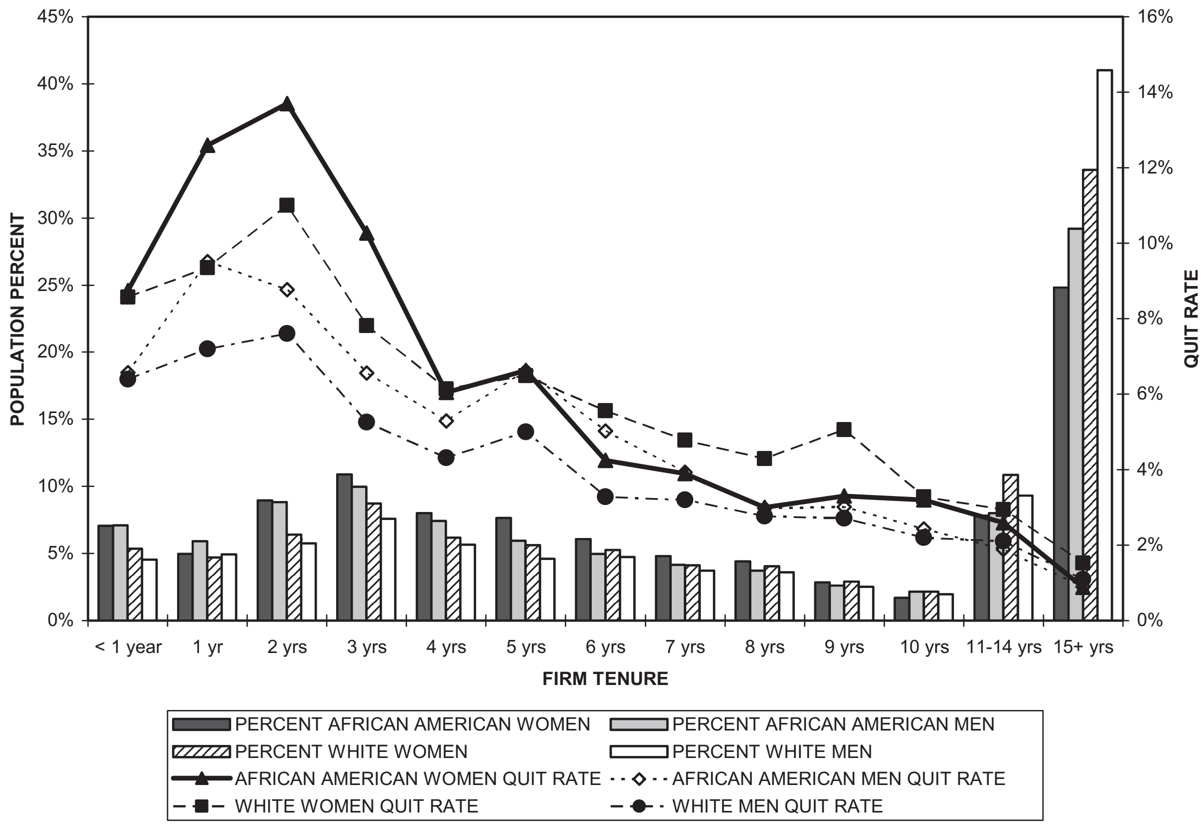


Figure 8. African American and White American quit rates as a function of firm tenure.

revealed that Asian American women are no different from White American men in turnover because sex and ethnicity have opposing effects on attrition for these women (see Table 14). Finally, controlling Asian American representation in jobs suppressed the effect of Asian American ethnicity, implying that short-tenure Asian Americans are more stable because they work in (more desirable) fields where members of their ethnicity are more numerous, which deters exits (see Table 14).

Testing Aggravated Jeopardy for Minority Women

Tables 15 to 19 report HLM regression tests, which disconfirmed Hypothesis 8. Whether including control variables or not, HLM logistic regression disclosed no significant Race × Gender interactions when all minorities were combined or when each race was considered separately.

Discussion

This consortium inquiry broadens as well as challenges prevailing academic and industry wisdom about the types of people who sever employment ties. Qualifying previous meta-analyses (Cotton & Tuttle, 1986; Griffeth et al., 2000; Hom & Griffeth, 1995), ARC findings indicate that quits monotonically decline with increasing firm tenure at a decelerating rate. Though recognized by labor economists and sociologists (Munasinghe & Sigman, 2004; Tuma, 1976), organizational psychologists have largely overlooked tenure–turnover curvilinearity. While conventional linearity is roughly accurate for the early period of employment, this form becomes increasingly imprecise when portraying quit risks for more senior incumbents. These findings corroborate predictions implied by theoretical formulations by Meglino and associates (2000) and Lazear (1998) as well as generalize this relationship form for a wide range of business professions in corporate America (Ippolito, 1991; Royalty, 1998; Sicherman, 1996).

Moving beyond prior work on moderators of performance–turnover associations (Allen & Griffeth, 2001), we demonstrated that employment opportunities (high historical attrition rate in a job) strengthen this relationship, such that marginal performers exit more than high performers when work alternatives are plentiful. Indeed, Steel's (1996) job availability proxy better upheld McEvoy–Cascio's (1987) unemployment moderation thesis than did previous tests using objective joblessness indices (Williams & Livingstone, 1994). Conceivably, Steel's (1996) measure of past attrition in a job may reflect job quality rather than job availability. Jobs with historically high quit rates might be least desirable and satisfying.[1] Contrary to this interpretation, Steel (1996) however showed that his measure correlated with objective and subjective measures of job opportunities. Trevor's (2001) research further challenged this alternative interpretation, as job performance may represent a form of *movement capital:* human capital traits that permit greater ease of movement. He observed that job dissatis-

[1] We thank a reviewer for this suggestion.

Table 14
Hierarchical Linear Modeling Models for Testing Explanations for Double Jeopardy for Short-Tenure Employees

	Level-2 intercepts for Level-1 predictors		
Level-1 predictors of turnover	Model without control variables	Model with tenure predictors	Model with tenure and percent Asian American concentration in jobs
Sex γ_{10}	.26*	.24*	.20*
Black γ_{20}	.13*	.08	.10
Asian γ_{30}	−.21*	−.24*	−.09
Hispanic γ_{40}	.03	.04	.01
Native γ_{50}	−.24	−.32	−.54*
Tenure γ_{60}		−.03	−.02
Tenure^2 γ_{70}		−.01	−.01
Percent Asian Americans in job γ_{80}			−.03*
Historical attrition rate in job γ_{90}			7.79*

Note. $N = 170{,}724$. Sex is coded 1 for women and 0 for men. Black is coded as 1 for African Americans, 0 for others; Asian is coded as 1 for Asians, 0 for others; Hispanic is coded as 1 for Hispanics, 0 for others; Native is coded as 1 for Native Americans and 0 for others (White Americans = reference group). γ_{10} = intercept of Level (L) 2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; γ_{40} = intercept of L2 regression predicting β_{4j}; γ_{50} = intercept of L2 regression predicting β_{5j}; γ_{60} = intercept of L2 regression predicting β_{6j}; γ_{70} = intercept of L2 regression predicting β_{7j}; γ_{80} = intercept of L2 regression predicting β_{8j}; γ_{90} = intercept of L2 regression predicting β_{9j}. No L2 predictors were included.
* $p \leq .05$.

faction most induces departures among incumbents with high rather than low movement capital. If historical quit rates reflect poor job quality instead of job availability, then high performers would have been more prone to exit when historical attrition rates are high, which was not the case (see Figure 2). Rather, our Performance × Historical Job Attrition moderation paralleled Trevor's (2001) other finding: Employees with low movement capital are more responsive to the external employment market.

Our ARC project furnishes the strongest evidence to date that the turnover of female professionals and managers is higher than that of their male counterparts. Rather than oppose Schwartz's (1989) claim, our findings buttress widespread anecdotal reports in the business and popular press about women's corporate flight (Shellenbarger, 2006). These findings are fairly robust for statistical analyses including other turnover influences, such as tenure and percent female incumbency, and the gender gap in attrition remained despite additional statistical controls for these turnover influences. This demonstration is invaluable because validating the gender–turnover relationship establishes the groundwork for further scrutiny of the theoretical mechanisms mediating this relationship (Baron & Kenny, 1986). Prevailing theoretical accounts suggest that sex discrimination in compensation, job level, promotions, and/or work–family interference may explain why women flee corporate America (Lyness & Judiesch, 2001; Stroh et al., 1996). These factors may explain why our finding diverges from two prior corporate studies. Lyness and Judiesch (2001) found that women managers quit less than men managers— both with and without controls for promotion, salary, and job level. However, near gender parity in management (42% women) in the financial services firm they studied may have lessened women's exits. The greater number of promotions and pay increases awarded to women relative to men in the highly regulated federal contractor studied by Petersen and Saporta (2004) may have similarly attenuated sex differences in turnover. Such varying conclusions across studies suggest the importance of examining multifirm data in future study of mediators and moderators of the gender–turnover relationship.

We considered one moderator of gender differences in turnover—proportion of female incumbents in a job—and expected that female turnover would decrease with rising female representation. Instead, we found that the behavior of men, but not women, was related to female concentration in jobs. While partially upholding (for men) the similarity–attraction thesis that members of a particular demographic subgroup take flight when demographically dissimilar members increasingly populate the workplace, this pattern could be due to a different process. Chatman and O'Reilly (2004), using workgroups as the unit of analysis, found that rising female representation was associated with higher turnover of both men and women. They argued that due to women's lower status in society, rising representation of women in a workgroup lowers its perceived status and attractiveness for men and women alike. Chatman and O'Reilly studied workgroups where gender heterogeneity is likely more salient than at the job level at which we examined the data. Yet gender differences in status may also underlie our results. Research in communication suggests that men are more attuned to status cues than are women (Tannen, 1995), and this may explain why male quits were sensitive to female representation even at the job level.

Admittedly, the gender gap in turnover seems small: 3.20% male quit rate versus 4.88% women quit rate. Yet a computer simulation by Martell, Lane, and Emrich (1996) projected that sex differences explaining merely 1% of variance in initial performance ratings substantially lowered promotion rates for women, resulting in 35% female incumbency in senior management (though equal numbers of women and men staffed the bottom of an eight-level job hierarchy). Martell (1999, cited in Eagly & Karau, 2002) later showed that adding a very small gender bias at each yearly review generated the minuscule proportion of female senior executives typical of corporations (Helfat et al., 2006). From these demonstrations, Eagly and Karau (2002) concluded that "slight prejudice that is consistently acted on greatly reduces women's chances of rising to high-level positions in organizations" (p. 589). By extension, the "slightly" higher women's quit rate in ARC firms (in tandem with smaller female representation at each job

Table 15
Hierarchical Linear Modeling Models for Testing Aggravated Jeopardy for Minority Women

Model	Parameter Estimate							
	γ_{00}	γ_{10}	γ_{20}	γ_{30}	τ_{00}	τ_{11}	τ_{22}	τ_{33}
Model 24: Sex + minority status + Sex × Minority interaction predicts quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{minority}_{ij}) + \beta_{3j}(\text{interaction}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	−3.35*	0.32*	0.21*	−0.08	0.24*	0.05*	0.05*	0.05*
Model 25: Sex + minority status + Sex × Minority interaction + control variables predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{minority}_{ij}) + \beta_{3j}(\text{interaction}_{ij}) + \beta_{4j}(\text{tenure}_{ij}) + \beta_{5j}(\text{tenure}_{ij}^2) + \beta_{6j}(\text{jobsex}_{jk}) + \beta_{7j}(\text{jobminority}_{jk}) + \beta_{8j}(\text{jobquits}_{jk})$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_j) + \gamma_{02}(\text{workforce maturity}_j) + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	−3.70*	0.21*	0.003	−0.07	0.35*	0.04	0.03	0.03

Note. $N = 404{,}052$. Sex is coded as 1 for women and 0 for men; minority is coded as 1 for minorities and 0 for Whites; interaction is a product between sex and minority status. γ_{00} = intercept of Level (L) 2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}; τ_{33} = variance in L2 residual for models predicting β_{3j}. Model 25 estimated L2 parameters for control variables, which are omitted but are available on request.
* $p \leq .05$.

level) may drastically shrink the female talent pool for advancement into executive posts (Petersen & Saporta, 2004).

While gender disparity in quits persisted after control variables were accounted for, statistical controls eliminated racioethnic differences in the total sample. Secondary tests specifically identified racial differences in tenure as being responsible for higher turnover among minorities and minority women. These findings suggest that socialization challenges faced by all beginning employees on entry into corporate life explain why people of color, whose numbers include more newcomers, are most liable to quit (Griffeth & Hom, 2001). It would seem that ARC establishments are more successful at recruiting minorities (giving rise to highly skewed frequency distributions for minorities toward short-tenure categories noted in Figures 6 and 8) but are less successful at forestalling their early attrition. Perhaps this pattern is not surprising. Studies have found that organizational entry is no longer the major barrier to corporate advancement for people of color (Kilian, Hukai, & McCarty, 2005). Indeed, surveys of diversity practices have reported that recruiting is one of the most common strategies used in organizations (Kulik & Roberson, in press). McKay and Avery (2005) noted that diversity recruitment practices, while effective at bringing people into the organization, may ironically contribute to high early turnover if they raise expectations for a positive diversity climate that is not fulfilled. As accumulated experience is typically needed to reach senior levels in the organization, heightened early departure rates among minorities and minority women contract the "pipeline" for future leadership (Eagly & Karau, 2002). In a vicious circle, entering minorities may interpret numerical rarity of executives of color as an indicator of barriers to advancement, motivating them to seek opportunities elsewhere (Elvira & Cohen, 2001). In short, the finding that racial gaps in turnover vanish with statistical controls for tenure should not be construed as a hallmark of corporate progress toward a diversified workforce. Our data underscore the importance of increased attention to retention and advancement as well as recruitment in diversity initiatives.

Post hoc analyses with only those at 6 or less years of tenure revealed that tenure differences insufficiently account for racial differences for African Americans and Asian Americans. For these groups, holding tenure constant did not eliminate racial differences. The pattern partly supports Hypotheses 6a and 6b that African American employees quit more than White or Asian American employees. This finding suggests that beyond socialization difficulties there are additional early work experiences that distinguish African Americans from other groups. For example, Thomas and Gabarro (1999) compared the career histories of White and minority male executives (who were largely African American). The successful minority male executives experienced slower career progress than did White male executives during their early years in a corporation, such that the two groups seemed to follow different career trajectories (see also James, 2000; Thomas, 2001). Such discouraging experiences may prompt beginning African Americans to believe there are limited advancement opportunities for their group (Rosin & Korabik, 1995), inducing early exits (Roberson, 2004).

Post hoc tests further indicated that the proportion of African and Hispanic Americans in jobs lies behind the corporate flight of beginning African Americans. That is, we determined that African Americans occupy jobs with greater numbers of African and Hispanic Americans (though White Americans constitute the over-

Table 16
Hierarchical Linear Modeling Models for Testing Aggravated Jeopardy for African American Women

Model	Parameter Estimate							
	γ_{00}	γ_{10}	γ_{20}	γ_{30}	τ_{00}	τ_{11}	τ_{22}	τ_{33}
Model 26: Sex + African American + Sex × African American interaction predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{black}_{ij}) + \beta_{3j}(\text{interaction}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	−3.35*	0.32*	0.21	−0.05	0.26*	0.05*	0.13*	0.10*
Model 27: Sex + African American + Sex × African American interaction + control variables predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{black}_{ij}) + \beta_{3j}(\text{interaction}_{ij}) + \beta_{4j}(\text{tenure}_{ij}) + \beta_{5j}(\text{tenure}_{ij}{}^{2}) + \beta_{6j}(\text{jobsex}_{jk}) + \beta_{7j}(\text{jobblack}_{jk}) + \beta_{8j}(\text{jobquits}_{jk})$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_{j}) + \gamma_{02}(\text{workforce maturity}_{j}) + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	−3.65*	0.19*	−0.01	−0.05	0.17*	0.05	0.08	0.04

Note. $N = 350{,}575$. Sex is coded as 1 for women and 0 for men; black is coded as 1 for African Americans and 0 for White Americans; interaction is a product between sex and African American status; jobblack_{jk} is the percent of African American incumbents in job k in firm j. γ_{00} = intercept of Level (L) 2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}; τ_{33} = variance in L2 residual for models predicting β_{3j}. Model 27 estimated L2 parameters for control variables, which are omitted but available on request.
* $p \leq .05$.

whelming majority in all jobs) and that such jobs stimulate turnover for incumbents of all races. Several possible reasons exist for this ethnic clustering. Unequal distributions of racioethnic groups across occupations may reflect different career and educational choices. For example, African Americans are underrepresented in science and engineering professions compared with Asian Americans (Tang, 1997). Yet overrepresentation of Hispanic and African Americans in certain jobs also suggests some racial segregation. This accords with past findings that although racial segregation is most evident in skilled manual jobs it also occurs, albeit less strongly, in jobs requiring college degrees (Tomaskovik-Devey, 1994). Some approaches to diversity management may also exacerbate racial clustering. Organizations may recruit African and Hispanic Americans for sales and customer service, specifically to target corresponding markets (Sacco & Schmitt, 2005). This strategy may be most common in organizations holding an “access and legitimacy” perspective on diversity, where diversity is valued for its ability to increase access to different markets (Ely & Thomas, 2001). Ely and Thomas (2001) argued that this strategy can enhance workforce diversity but also can confine such diversity to those parts of the organization serving niche markets. They further cautioned that this practice can result in perceptions that diverse employees are desired only in particular functions and that opportunities for advancement in the larger organization are limited. These perceptions may be true as prior research has shown that job quality improves as the proportion of White male incumbents increases (Tomaskovik-Devey, 1994). Thus, jobs with higher African and Hispanic American concentrations may have less status, truncated career paths, lower pay, and less influence (Chatman & O’Reilly, 2004; Hom & Griffeth, 1995; Tsui et al., 1992), prompting turnover by incumbents of all ethnicities.

Post hoc tests also revealed that tenure differences did not account for racial differences in turnover for short-tenure Asian Americans (see Table 9). However, in contrast to Hypothesis 6c, low-tenure Asian Americans quit significantly less than White Americans. Asian Americans also tended to be clustered in jobs with more members of their race. However, for Asian Americans, ethnic clustering in jobs reduced turnover. Several explanations may clarify why Asian American job concentration mediates the (negative) relationship between Asian American ethnicity and quits. For one, Asian Americans may hold higher status or more attractive jobs (e.g., product engineers) than do other minorities (e.g., customer service). From a similarity–attraction perspective, Asian American employees may be more likely to stay because of greater opportunities for interactions with similar others whose ethnic assembly does not symbolize lower status (Chatman & O’Reilly, 2004). Alternatively, a stronger collectivist orientation for this group may predispose Asian Americans to remain (Palich, Hom, & Griffeth, 1995). Indeed, collectivist people are attracted to ethnically similar work units, valuing ingroups with shared demographic traits (Triandis, 1995). Comprising more technical fields, this ethnic job cluster may also include more nonresident aliens who must stay with their sponsoring organization for the (3-year) duration of their visa (awarded for scarce technical skills; “Out of Visas,” 2006; Watts, 2001; Zavodny, 2003). Premature departures

Table 17
Hierarchical Linear Modeling Models for Testing Aggravated Jeopardy for Asian American Women

Model	Parameter Estimate							
	γ_{00}	γ_{10}	γ_{20}	γ_{30}	τ_{00}	τ_{11}	τ_{22}	τ_{33}
Model 28: Sex + Asian American + Sex × Asian American interaction predict quits L1: Logit($quit_{ij}$) = β_{0j} + β_{1j}(sex_{ij}) + β_{2j}($asian_{ij}$) + β_{3j}($interaction_{ij}$) L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	−3.35*	0.33*	0.14*	−0.09	0.26*	0.05*	0.04*	0.08*
Model 29: Sex + Asian American + Sex × Asian American interaction + control variables predict quits L1: Logit($quit_{ij}$) = β_{0j} + β_{1j}(sex_{ij}) + β_{2j}($asian_{ij}$) + β_{3j}($interaction_{ij}$) + β_{4j}($tenure_{ij}$) + β_{5j}($tenure_{ij}^2$) + β_{6j}($jobsex_{jk}$) + β_{7j}($jobasian_{jk}$) + β_{8j}($jobquits_{jk}$) L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}$(workforce $size_j$) + γ_{02}(workforce $maturity_j$) + U_{0j} L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	−3.90*	0.21*	−0.05	−0.05	0.50*	0.04	0.05	0.03*

Note. $N = 356{,}855$. Sex is coded as 1 for women and 0 for men; Asian is coded as 1 for Asian Americans and 0 for White Americans; interaction is a product between sex and Asian American status; and $jobasian_{jk}$ is the percent of Asian American incumbents in job k in firm j. γ_{00} = intercept of Level (L) 2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}; τ_{33} = variance in L2 residual for models predicting β_{3j}. Model 29 estimated L2 parameters for control variables, which are omitted and available on request.
$^* p \leq .05$.

would risk deportation, especially if another sponsor cannot be found. In sum, these arguments imply stronger work connections and more sacrifice from leaving, suggesting that greater job embeddedness enhances the retention of Asian Americans (T. R. Mitchell & Lee, 2001).

Going beyond past work on dual discrimination (Berdahl & Moore, 2006; Kulis & Miller, 1988; Sanchez-Hucles, 1997), we substantiated the joint effects of race and sex on voluntary terminations among exempt workers for certain non-White Americans. African American and Hispanic American women quit more often than their male counterparts or White Americans. With few exceptions (Hosoda et al., 2003; Powell & Butterfield, 1997), diversity research primarily concludes that African and Hispanic American women face worse working conditions than do African American and Hispanic American men or White Americans (Barnum et al., 1995; Berdahl & Moore, 2006; Giscombe & Mattis, 2002). Extending these findings, an HLM test also revealed that Asian American women quit more than men of their own race and White Americans of either sex. Beyond this, we uncovered little evidence that dual marginalization plagues Native American women. Nevertheless, we caution that double jeopardy evidence evaporates when we statistically controlled other turnover influences—notably, tenure. Rather than double marginalization, these latter tests imply that occupancy of women of color in low-tenure positions—where they risk greater socialization difficulties (e.g., poor mentoring, receive less meaningful assignments)—impels them to quit.

Our research further disputes the multiplicative version of the double jeopardy thesis for all racial minorities, which is consonant with the few available tests of this proposition (Berdahl & Moore, 2006; Hosoda et al., 2003; Nkomo & Cox, 1989). Moreover, our work contradicts a "parallel process" model that presumes no additive effects for race and sex (Hosoda et al., 2003). According to this thinking, women and minorities encounter similar forms of discrimination because they are subjected to similar forms of negative stereotyping (Hosoda et al., 2003). By implication, non-Whites and White women would have similar turnover rates that exceed that of White men. Yet such subgroup differences predicted by the parallel process model were not borne out by ARC data.

Study Limitations

Though immense, the ARC data are limited. Specifically, the present cross-sectional evidence for tenure–turnover curvilinearity should be replicated with longitudinal data and survival analyses. While consistent with survival analyses (Royalty, 1998; Sicherman, 1996), our cross-sectional findings may reflect cohort (or generational) differences rather than simply tenure effects. Thus, our bell-shaped tenure–quit curve for knowledge workers should be replicated with panel data on more fine-grained time units (e.g., weeks, Farber, 1994) and analyzed with a log-logistic model (Park & Sandefur, 2003), which may more readily detect Burt's (2005) "kinked decay function." Moreover, our truncated measure for company tenure may have underestimated tenure–turnover curvilinearity. Because quit rates are low beyond the 10th year of employment (nearing 1% at the 15th year) and remained low throughout lengthy terms of service, ARC firms reported quit statistics for broad tenure categories for long seniority. Senior-

Table 18
Hierarchical Linear Modeling Models for Testing Aggravated Jeopardy for Hispanic American Women

Model	Parameter Estimate							
	γ_{00}	γ_{10}	γ_{20}	γ_{30}	τ_{00}	τ_{11}	τ_{22}	τ_{33}
Model 30: Sex + Hispanic American + Sex × Hispanic American interaction predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{hispanic}_{ij}) + \beta_{3j}(\text{interaction}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	−3.35*	0.33*	0.27*	−0.14	0.24*	0.04*	0.06*	0.01
Model 31: Sex + Hispanic American + Sex × Hispanic American interaction + control variables predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{hispanic}_{ij}) + \beta_{3j}(\text{interaction}_{ij}) + \beta_{4j}(\text{tenure}_{ij}) + \beta_{5j}(\text{tenure}_{ij}^2) + \beta_{6j}(\text{jobsex}_{jk}) + \beta_{7j}(\text{jobhispanic}_{jk}) + \beta_{8j}(\text{jobquits}_{jk})$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_j) + \gamma_{02}(\text{workforce maturity}_j) + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	−3.66*	0.21*	0.04	−0.08	0.19*	0.03	0.05*	0.03

Note. $N = 339{,}528$. Sex is coded as 1 for women and 0 for men; hispanic is coded as 1 for Hispanic Americans and 0 for White Americans; interaction is a product between sex and Hispanic American status; jobhispanic_{jk} is the percent of Hispanic American incumbents in job k in firm j. γ_{00} = intercept of Level (L) 2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}; τ_{33} = variance in L2 residual for models predicting β_{3j}. Model 31 estimated L2 parameters for control variables, which are omitted and available on request.
* $p \leq .05$.

tenure categories may thus have weakened statistical tests of the quadratic tenure term in logistic regression and understated a potentially more pronounced curvilinear relationship. Further, the consortium's dichotomous performance measure may have attenuated moderating effects by underestimating the performance–turnover relationship. Another limitation of our research is the relatively small number of companies reporting performance and demographic data. Additionally, ARC findings, where the average quit rate was 3.89% in 2003, may not generalize to other companies, especially in high turnover industries. Still, the ARC quit rate is comparable with national quit statistics for professional workers, as the U.S. Department of Labor (2004–2005) finds that the mean rate of turnover for professional and business services was 3.27% in 2003.

The corporate consortium currently does not collect measures on other factors (e.g., education, hierarchical level, pay) that might underpin group differences in leaving (Lyness & Judiesch, 2001). Such omitted factors may be behind the gender disparity that remained despite inclusion of a full array of control variables into statistical analyses. Though our sample included managers, ARC does not collect data about incumbents' job level, precluding gender and ethnic comparisons among managers and executives (cf. Lyness & Judiesch, 2001). On the basis of Equal Employment Opportunity Commission categories, we found that ARC racial groupings were also broad, preventing inquiry into termination differences among different ethnic subgroups. For example, Asian Americans have all too often been viewed as a homogenous group without regard to nationality, nativity status, or Westernization (Leong, 2001; Tang, 1997). Such within-group differences are material. Tang (2000) observed unequal career outcomes for native- and foreign-born Asian engineers, whereas Leong (2001) noticed that highly acculturated Asians report greater job satisfaction and earn better performance reviews than do less acculturated Asians.

Implications

Quite obviously the ARC data yield major implications for both research and practice. ARC was formed to provide organizations with a representative, accurate picture of turnover so that the firms could better understand their own patterns of attrition and compare them with those of other firms. Thus, our findings should be similarly useful for other comparable organizations. We can only speculate on the reasons underlying turnover decisions, but our results identify the characteristics of those at greater risk for turnover, identifying areas where management might want to pinpoint retention strategies. For scholars, our large sample provides a more accurate picture of trends in turnover and the relationship of turnover to employee characteristics, both challenging and supporting current theories. For example, our findings clearly establish corporate attrition among professional and managerial women. Both scholarly and applied research must focus on discovering why. Theories are available to guide such inquiry, and propositions regarding work–family conflict and continued bias in promotions and compensation have been advanced. On an encouraging note, being a numerical minority in male-dominated jobs did

Table 19
Hierarchical Linear Modeling Models for Testing Aggravated Jeopardy for Native American Women

Model	Parameter Estimate							
	γ_{00}	γ_{10}	γ_{20}	γ_{30}	τ_{00}	τ_{11}	τ_{22}	τ_{33}
Model 32: Sex + Native American + Sex × Native American interaction predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{native}_{ij}) + \beta_{3j}(\text{interaction}_{ij})$ L2: $\beta_{0j} = \gamma_{00} + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	-3.35^{*}	0.33^{*}	-0.01	-0.30	0.24^{*}	0.04^{*}	0.32^{*}	0.28
Model 33: Sex + Native American + Sex × Native American interaction + control variables predict quits L1: $\text{Logit}(\text{quit}_{ij}) = \beta_{0j} + \beta_{1j}(\text{sex}_{ij}) + \beta_{2j}(\text{native}_{ij}) + \beta_{3j}(\text{interaction}_{ij}) + \beta_{4j}(\text{tenure}_{ij}) + \beta_{5j}(\text{tenure}_{ij}^{2}) + \beta_{6j}(\text{jobsex}_{jk}) + \beta_{7j}(\text{jobnative}_{jk}) + \beta_{8j}(\text{jobquits}_{jk})$ L2: $\beta_{0j} = \gamma_{00} + \gamma_{01}(\text{workforce size}_{j}) + \gamma_{02}(\text{workforce maturity}_{j}) + U_{0j}$ L2: $\beta_{1j} = \gamma_{10} + U_{1j}$ L2: $\beta_{2j} = \gamma_{20} + U_{2j}$ L2: $\beta_{3j} = \gamma_{30} + U_{3j}$	-3.74^{*}	0.21^{*}	-0.22	-0.42	0.26^{*}	0.04	0.38	0.45

Note. $N = 324{,}603$. Sex is coded as 1 for women and 0 for men; native is coded as 1 for Native Americans and 0 for White Americans; interaction is a product between sex and Native American status; jobnative_{jk} is the percent of Native American incumbents in job k in firm j. γ_{00} = intercept of Level (L) 2 regression predicting β_{0j}; γ_{10} = intercept of L2 regression predicting β_{1j}; γ_{20} = intercept of L2 regression predicting β_{2j}; γ_{30} = intercept of L2 regression predicting β_{3j}; τ_{00} = variance in L2 residual for models predicting β_{0j}; τ_{11} = variance in L2 residual for models predicting β_{1j}; τ_{22} = variance in L2 residual for models predicting β_{2j}; τ_{33} = variance in L2 residual for models predicting β_{3j}. Model 33 estimated L2 parameters for control variables, which are omitted and available on request.
* $p \leq .05$.

not impel more female quits, yet reasons why men leave jobs with increasing numbers of women merit further scrutiny.

While turnover is most likely for all employees during initial employment, African Americans are especially at risk during this time. Our findings suggest that they may face different—if not more difficult—socialization hurdles than do White or Asian Americans (James, 2000; Thomas, 2001; Thomas & Gabarro, 1999). By choice or administrative fiat (Lefkowitz, 1994), they often occupy jobs having more African and Hispanic American incumbents—work environments that motivate them to flee. Thus, companies must monitor the presence of such ethnic enclaves, and scholars should study why and how they arise. To illustrate, Lefkowitz (1994) reported an ethnic drift toward supervisor–subordinate race congruency over time. It is possible that management may encourage ethnic clustering with good intentions, believing that newcomers of color can receive more psychosocial support from same-race others. Our results suggest that such practices can also have negative outcomes. The differential relationship of Asian versus African American/Hispanic American representation in jobs to turnover was not expected but is consistent with the reasoning underlying Hypothesis 6a and 6b, that White Americans hold greater negative affect toward African and Hispanic Americans. Unfortunately, our results suggest the continued salience of race as a status indicator and organizational reproduction of societal patterns, even at the exempt levels studied here. Future research should examine organizational variation in racial and gender clustering in jobs and their relationships to turnover. Theoretically, we might expect these patterns to vary with the organizational climate for diversity (Cox, 1994) or organizational paradigm for diversity (Ely & Thomas, 2001).

Finally, corporations should systematically monitor the retention of women of color who may suffer disadvantages exacerbated by being double minorities (Kulis & Miller, 1988). Indeed, one *Fortune* 500 firm specifies that turnover rates of women of color in each functional area should not exceed comparable rates for the rest of its workforce (Giscombe & Mattis, 2002). Moreover, in-depth corporate research should pinpoint their unique disadvantages, such as inaccessibility to high visibility assignments and powerful mentors during initial employment when they are most susceptible to leaving (Giscombe & Mattis, 2002). To illustrate, Giscombe and Mattis (2002) found that African American women professionals and managers perceive more stereotyping about their race and more discomfort by other employees around members of their racial group than do Asian American women.

In summary, our project extends knowledge of turnover research by establishing that quitting is curvilinearly related to tenure and that historical quit rates in jobs can shape how performance relates to leaving. In addition, we provide invaluable evidence on sex– and race–turnover relationships, substantiating industry and journalistic accounts that women and racial minorities (excepting Native Americans) vacate exempt occupations at higher rates than do men and White Americans, respectively (Daniels, 2004; Mandel & Farrell, 1992; Schwartz, 1989; though statistical controls implicate different causes for different races). Our demonstrations animate further explorations into theoretical mechanisms behind demographic disparities in corporate attrition rather than presume their nonexistence (given heretofore ambiguous evidence). Advancing the turnover literature, we further show that African American, Asian American, and Hispanic American women flee corporate America more than do their male counterparts and White Americans of both sexes. Though the prevalence of "rookies" among minorities and minority women may underlie this corporate exodus, their premature attrition constricts the pipeline of people of color available for future leadership (Helfat et al., 2006; Thomas, 2001). In sum, this study confirms encouraging reports that organizational entry for historically excluded women and minorities is a declining problem, but it also raises serious concerns about these groups' full participation in corporate America given their premature attrition.

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