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Multinational Corporations and Stock Price Crash Risk

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Abstract

A nascent literature in finance and accounting on tail risk in individual stock returns concludes that bad news hoarding by corporate managers engenders sudden, extreme crashes in a firm's stock price when the bad news is eventually made public. This literature finds that firm-specific crash risk is higher among firms with more severe asymmetric information and agency problems. A hitherto disjointed literature spanning the fields of international business, finance, and accounting suggests that geographic dispersion in a firm's operations, and especially dispersion across different countries, gives rise to organizational complexities and greater costs of monitoring that can exacerbate asymmetric information and agency problems. Motivated by the confluence of arguments and findings from these two strands of literature, this paper examines whether stock price crash risk is higher among multinational firms than domestic firms. Using a large sample of U.S. headquartered firms during 1987-2011, we find robust evidence that multinational firms are significantly more likely to crash than domestic firms. Moreover, we show that the difference in crash risk between multinational and domestic firms is most acute among firms with weaker corporate governance mechanisms, including weaker shareholder rights, less independent boards, and less stable institutional ownership. Our analysis indicates that stronger monitoring from each of these three governance mechanisms significantly attenuates the positive relation between crash risk and multinationality. Our findings are robust to the use of alternative measures of crash risk and to controlling for known determinants of crash risk identified in prior studies. Our study offers new insights that should hold value for scholars and market participants interested in understanding the implications of heighted agency problems that multinational firms are likely to encounter and scholars and market participants interested in developing models that more accurately predict tail risk in the equity returns of individual firms...

Key Words: *Multinational, Crash Risk, Tail Risk, Corporate Governance*

JEL classification: *F32, G11, G17, G14, G30, G34*

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Introduction

Scholars in international business fields have long been interested in understanding the benefits and costs of internationalization. Numerous studies suggest that there are potential benefits to establishing operations outside of a firm's home country, including scale and scope economies (Tallman and Li, 1996, Lu and Beamish, 2004), arbitrage opportunities (Rugman and Verbeke, 2004), operational flexibility (Kogut, 1983), lower tax liabilities (Hines and Rice, 1994, Rego, 2003), diversification when firms can diversify internationally at a lower cost than individual shareholders (Errunza and Senbet, 1981, 1984, Kogut and Kulatilaka 1994), and the exploitation of firm-specific, intangible assets (Caves 1971, Morck and Yeung, 1991). On the other hand, there is also a considerable literature viewing internationalization through the lens of agency theory. According to this literature, expanding a firm's operations across international boundaries induces greater complexity in a firm's operational structure and increases the difficulty and cost to investors of monitoring the firm, which in turn gives rise to greater agency costs via more severe asymmetric information and moral hazard problems (Lee and Kwok, 1988, Geringer et al., 1989, Mitchell et al., 1992, Nohria and Ghoshal, 1994, Burgman, 1996, Hitt et al., 1997, Sanders and Carpenter, 1998, Riahi-Belkaoui and Picur, 2001, Denis et al., 2002, Doukas and Pantzalis, 2003, Tihanyi et al., 2003, Bushman et al., 2004, Lee et al., 2008, Black et al., 2014, Tsao et al., 2016). In this paper, we add new insights to this literature by empirically investigating whether stock price crash risk is higher among multinational firms than domestic firms.

The recent financial crisis of 2008-2009 has renewed interest among scholars, policy makers, and market participants in understanding and modeling extreme negative outcomes (i.e., tail risk) in financial markets. Recently, interest among scholars in this topic has migrated to the study of severe price crashes of individual firms. A prominent line of thought that has emerged from this literature is that bad news hoarding by corporate managers leads to sudden, extreme price drops when the accumulated bad news is eventually made known to the market. Kothari et al. (2009) and Graham et al. (2005) discuss numerous factors that can incent managers to delay the release of negative news. In the theory of Jin and Myers (2006), managers withhold and accumulate bad news for extended periods until the cost or difficulty of concealing their negative private information becomes too high, at which point the accumulated bad news tends to come out all at once or very quickly. This causes a crash in the stock price, which can be empirically identified as an extreme left-tail outlier in the distribution of weekly or daily firm-specific (idiosyncratic) returns. The Jin and Myers (2006) model of bad news hoarding and stock price crashes has birthed a new and growing empirical literature focused on identifying corporate activities and/or firm characteristics that incent or facilitate bad news hoarding and which, therefore, predict stock price crashes. In general, this literature finds that crash risk is positively (negatively) related to factors that exacerbate (attenuate) information asymmetry and agency conflicts between managers and investors (Hutton et al., 2009, Kim et al., 2011a, 2011b, An and Zhang, 2013, Callen and Fang, 2013, 2015a, Kim et al., 2014, Xu et al., 2014, Yuan et al. 2016).

Based on the rationale that multinational firms face heightened asymmetric information and agency problems, we conjecture that managers of multinational firms have greater opportunities to engage in bad news hoarding, as well as greater opportunities to engage in self-serving activities (e.g., empire building or excess perk consumption) that incentivize bad news hoarding. Motivated by this line of thought, coupled with the findings of recent literature concluding that bad news hoarding engenders stock price crashes, we propose and empirically test the hypothesis that crash risk is higher among multinational firms than among their domestic counterparts. As in prior studies of crash risk, our empirical analysis uses firm-specific information available at the close of a given fiscal year to predict the probability of a firm-specific crash during the following fiscal year, where a crash is defined as an extreme, negative outlier in the distribution of the firm's weekly idiosyncratic returns.[1] Using a large panel of U.S. headquartered firms spanning the period 1987-2011, we find that one-year-ahead crash risk is significantly higher among multinational firms than domestic firms, both in simple univariate comparisons and in multivariate regressions that control for

[1] *Specifically, we follow prior literature (Hutton et al., 2009, Kim et al., 2011a, 2011b) and define a crash as a weekly firm-specific log-return that is 3.1 or more standard deviations below the firm's average during the year.*

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known predictors of crash risk identified by prior studies. After establishing a positive relation between multinationality and crash risk, we further examine whether this relation varies with the strength of the firm's corporate governance mechanisms. This additional analysis is motivated by studies suggesting that effective governance mechanisms are necessary to mitigate the risk of elevated agency problems in multinational firms or complex firms in general (Bushman et al., 2004, Luo, 2005, Jiraporn et al., 2006, Gande et al., 2009, Tsao et al, 2016). We find robust evidence that the difference in crash risk between multinational and domestic firms is most acute among firms with weaker corporate governance mechanisms, including weaker shareholder rights, less independent boards, and less stable institutional ownership. Our analysis indicates that stronger monitoring along each of these three dimensions of governance significantly attenuates the positive relation between crash risk and multinationality. All of our findings are robust to the use of alternative measures of firm-specific crash risk employed by prior studies, including the negative skewness of firm-specific weekly returns (Chen et al., 2001; Kim et al., 2011a, 2011b).

Our study makes several important contributions. First, we add to the literature on agency costs in multinational firms. Numerous studies have examined the relation between multinationality and measures of operating performance or firm value, including Grant (1987), Gomes and Ramaswamy (1999), Qian (2002), Denis et al. (2002), Contractor et al. (2003), Doukas and Lang (2003), Doukas and Kan (2006), and Gande et al. (2009). However, as discussed by Tsao et al. (2016) and Wiersema and Bowen (2011), this line of empirical research has produced mixed evidence and conflicting conclusions. Our study brings a new yet relevant perspective to this literature. To our knowledge, our paper is the first to elucidate the presence of heightened tail risk in the equity returns of multinational firms relative to domestic firms, and especially among firms in which agency costs are likely to be higher due to weaker governance. We therefore offer significant new insights that should hold value for scholars interested in understanding the implications of heightened agency problems in multinational firms. Second, our study adds to the growing literature on tail risk in equity markets. Recent studies of firm-specific crash risk have successfully identified several firm characteristics and corporate activities that predict future crash risk. Our study identifies an observable firm characteristics, i.e., whether a firm is multinational or domestic, with a robust empirical relation to crash risk and also provides evidence on the extent to which certain corporate governance mechanisms affect this relation, which should help to further inform future empirical and theoretical research on bad news hoarding and crash risk. Finally, the findings of our study should have value for various market participants and practitioners. For example, the empirical evidence of Mitton and Vorkink (2007) suggests that skewness in individual security returns, which is determined by the probability of extreme returns, influences investors' portfolio decisions, while recent studies by Boyer et al. (2010) and Conrad et al. (2013) suggest that idiosyncratic skewness is a priced risk factor in the cross-section of expected stock returns. In addition, Xiong et al. (2016) conclude that investors concerned with the mean, variance, and skewness of portfolio returns can benefit from more accurate forecasts of skewness of individual assets in a portfolio. We show that multinationality is a significant predictor of future idiosyncratic return skewness, thus our findings should be particularly useful to investors and practitioners interested in developing valuation or portfolio decision models that explicitly incorporate ex ante forecasts of skewness. Another practical arena where our work is relevant is the options market for individual stocks, since option values depend on the likelihood of extreme returns (Bakshi et al., 1997; Pan, 2002). Our findings should therefore be of interest to market participants and practitioners interested in developing practically useful option pricing models that take account of variation in expected crash risk across individual stocks.

The remainder of our paper proceeds as follows. The second (next) section discusses related literature and presents our empirical hypotheses. The third section discusses our data and methodology. The fourth section discusses our empirical results and the last section concludes.

Literature Review

As discussed by Kothari et al. (2009), managers typically possess superior private information relative to outside investors and can exercise significant discretion over the flow of information to capital markets, and in particular over how quickly or slowly certain types of firm-specific information are communicated to

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investors. Managers' tendency to withhold bad news stems from a standard agency problem where managerial disclosure preferences are not aligned with those of shareholders, who are presumed to prefer timely disclosures. Kothari et al. (2009) discuss various factors that can incent managers to delay the release of bad news for prolonged periods, including concerns about current-period performance-based compensation, prospects for promotion, future employment, post-retirement benefits such as directorships, and the potential for termination. In addition, non-monetary incentives, such as empire building, perquisite consumption, and the desire to maintain the esteem of peers, could also motivate managers to conceal and withhold negative private information (Ball, 2009). Survey evidence from Graham et al. (2005) also suggests managers delay the release of bad news in hopes that subsequent (positive) events will allow them to "bury" the bad news.

Our first hypothesis is based on the confluence of arguments and findings from two separate strands of literature. The first is the recent literature on firm-specific crash risk, which has its origins in the theory of Jin and Myers (2006). In Jin and Myers (2006), managers withhold and accumulate bad news up to some threshold level, at which point it becomes too costly or difficult for the manager to continue. When this threshold is reached, the accumulated bad news is revealed to the market all at once, resulting in a large crash in the stock price. Based on the Jin and Myers (2006) rationale that bad news hoarding engenders stock price crashes, recent empirical studies attempting to identify factors that predict firm-specific crash risk have focused on corporate activities or firm characteristics that exacerbate (or mitigate) information asymmetry and agency conflicts between managers and investors, including earnings management (Hutton et al., 2009), aggressive tax avoidance strategies (Kim et al., 2011a), stock option compensation (Kim et al., 2011b), stable institutional ownership (Callen and Fang, 2013, An and Zhang, 2013), corporate social responsibility (Kim et al., 2014), and the religiosity of the area where a firm is headquartered (Callen and Fang, 2015a).[2]

The second strand of literature we draw from in order to formulate our main empirical hypothesis spans the fields of international business, accounting, and finance. This literature suggests that internationalization of a firm's operations gives rise to greater information asymmetry between managers and investors. This can occur through multiple channels. First, multinational firms are more organizationally complex and typically have multiple subsidiaries or business segments located in foreign countries. In the U.S., firms are required to disclose, on a regular basis, large amounts of detailed, *aggregated* information regarding the firm's financial condition and performance (e.g., income statements, balance sheets, cash flow statements, etc.). However, the rules governing reporting of individual geographic segments require far less disclosure of segment-specific information. Thus, while investors can observe the aggregate cash flows, liabilities, operating costs, current and long-term assets, etc., of the firm as a whole, they typically cannot observe the same set of detailed information for the individual foreign (and domestic) segments of a multinational firm. This information aggregation problem can result in substantial information asymmetry between managers and outside shareholders (Gilson et al., 2001, Krishnaswami and Subramaniam, 1999, Bushman et al., 2004, Liu and Lai, 2012). Second, investors typically have more knowledge about their home country than about foreign countries, thus it stands to reason that investors will tend to be less informed about a firm's foreign operations than its domestic operations (Ashbaugh and Pincus, 2001). In addition, differences in cultures, customs, languages, competitors, regulations, and political systems across different countries make it more difficult and costly for investors to become knowledgeable about foreign business environments, which in turn diminishes investors' incentives and ability to monitor a firm's foreign operations, thus widening the asymmetric information gap between investors and managers. Empirical evidence supporting the notion that information asymmetry is higher in multinational firms is provided by Duru and Reeb (2002), who find that analysts' earnings forecasts are systematically less accurate for multinational firms than domestic firms. In addition, evidence presented by Thomas (1999), Callen et al. (2005), and Khurana et al. (2003) indicates that investors and analysts systematically underestimate the

[2] *Using samples of public firms in China, Xu et al. (2014) find that crash risk is associated with excess managerial perk consumption, while Yuan et al. (2016) finds that crash risk is higher when directors and officers are protected from financial liability via liability insurance.*

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persistence of foreign earnings. Thomas (1999) and Callen et al. (2005) conclude that investors do so because of a lack of understanding of firms' foreign operations due in part to poor disclosure.

Numerous studies suggest that greater information asymmetry in multinational firms gives rise to greater agency costs via greater conflicts of interest between owners and managers, including Lee and Kwok (1988), Geringer et al. (1989), Mitchell et al. (1992), Nohria and Ghoshal (1994), Burgman (1996), Hitt et al. (1997), Sanders and Carpenter (1998), Riahi-Belkaoui and Picur (2001), Denis et al. (2002), Tihanyi et al. (2003), Bushman et al. (2004), Lee et al. (2008), Black et al. (2014), and Tsao et al. (2016). When investors are less informed about the firm and the activities of its managers, they are less equipped to effectively monitor managers' actions and assess managers' performance. This gives rise to greater opportunities for managers to engage in self-serving actions that are not in the best interest of shareholders. As observed by Kothari et al. (2009), bad new hoarding is, in and of itself, a self-serving action that conflicts with the interests of shareholders, since shareholders prefer timely disclosures. The preceding discussion suggests that more severe information asymmetry in multinational firms should facilitate managers' ability to hoard bad news for extended periods. Moreover, Black et al. (2014) point out that the greater complexity and information asymmetry in multinational firms facilitates self-serving resource diversion by managers, such as excess perquisite consumption or empire building. As discussed by Kim et al. (2011a), Xu et al. (2016), and Ball (2009), managers engaged in substantial resource diverting activities for prolonged periods have incentives to conceal their actions and the resulting negative consequences, which is a form of bad news hoarding. Thus, the preceding discussion suggests that managers of multinational firms have greater opportunities to engage in bad news hoarding, as well as greater opportunities to engage in self-serving activities that incentivize bad news hoarding. This leads to our first formal hypothesis.

***H1**: Crash risk is higher among multinational firms than domestic firms, all else equal.*

Bushman et al. (2004), Luo (2005), Jiraporn et al. (2006), Gande et al. (2009), and Tsao et al. (2016) argue that heightened agency problems in multinational firms can be mitigated by stronger corporate governance mechanisms. Mangers of multinational firms with stronger corporate governance should be subject to greater levels of monitoring and discipline, which should help to attenuate agency problems, while managers of multinationals with weak corporate governance mechanisms should have greater opportunities to engage in bad news hoarding and/or the resource diverting activities that incentivize bad news hoarding. We therefore expect the impact of multinationality on crash risk to be attenuated for firms with stronger governance. This leads to our second formal hypothesis.

***H2**: The (positive) difference in crash risk between multinational and domestic firms is attenuated by stronger monitoring from the firm's corporate governance mechanisms, all else equal.*

Research and Methods

Data and Sample

Our data sources include the Center for Research in Security Prices (CRSP) for stock prices, returns, shares outstanding, and trading volumes, Compustat for financial accounting and geographic segment data, the Institutional Shareholder Service (ISS) Governance and Directors databases[3] for data on governance provisions and boards of directors, and Thomson Reuters' Institutional (13f) Holdings database for data on shareholdings of institutional investors. Our baseline sample consists of firms that are covered by CRSP and Compustat during fiscal years 1987-2011[4] with publicly traded common stock (CRSP share codes of 10, 11, or 12) and headquarters located in the United States (Compustat Foreign Incorporate Code = USA). We exclude firms-years with non-positive book assets and firm-years with fewer than thirty non-missing weekly returns in CRSP. We also require that firms have the necessary information in CRSP

[3] The ISS databases were formerly known as RiskMetrics and IRRC.

[4] We start our sample period in 1987 because the data items required to construct several of the variables in our analysis are unavailable before this, including discretionary accruals (Hutton et al., 2009) and long-run effective tax rates (Kim et al., 2011a). In addition, ISS data on governance provisions and boards of directors are available starting in 1990 and 1997, respectively.

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and Compustat required for the construction of our baseline control variables, which are described in detail in a later subsection entitled "Baseline Control Variables." After applying these data screens we are left with a baseline sample of 104,929 firm-years.[5]

Estimation of Firm-Specific Returns and Crash Risk Measures

We estimate firm-specific (idiosyncratic) weekly returns for each firm-year in the sample using a five-factor model that includes the three factors of Fama and French (1993), the momentum factor of Carhart (1997), and an industry factor:

$$R_{i,\tau} = \alpha_i + \beta_i R_{m,\tau} + s_i SMB_\tau + h_i HML_\tau + u_i UMD_\tau + \gamma_i IND_{i,\tau} + \varepsilon_{i,\tau} \qquad (1)$$

where, for firm *i* in week τ, R_i is the weekly stock return of firm *i*, R_m is the weekly return on the CRSP value-weighted index, *SMB*, *HML*, and *UMD* are the weekly returns on the Small-Minus-Big, High-Minus-Low, and Up-Minus-Down portfolios that capture size, book-to-market, and return momentum effects, respectively, and IND_i is the weekly return on a value-weighted index that includes all firms in firm *i*'s Fama-French industry.[6] The residual from Equation (1) is defined as the weekly firm-specific return in week τ.

Following Hutton et al. (2009) and Kim et al. (2011a, 2011b), we estimate Equation (1) separately for each firm-year and define firm-specific log-returns as the natural logarithm of one plus the firm-specific return. Firm *i* is defined as experiencing a stock price crash in a given week if the firm-specific log-return is 3.09 or more standard deviations below firm *i*'s mean weekly firms-specific log-return during that fiscal year. The cutoff of 3.09 standard deviations is chosen to generate a frequency of 0.1% in the normal distribution.[7] As in Hutton et al. (2009) and Kim et al. (2011a, 2011b), our primary response variable of interest, denoted as *CRASH*, is a binary variable that equals one if a firm experiences a stock price crash in the given fiscal year and zero otherwise. As in prior studies of crash risk, our empirical analysis uses information available in year *t* to predict crash risk in year *t*+1. Thus, measurement of our explanatory variables begins in the first year of our sample period, 1987, while the measurement of our crash risk variables begins in 1988.

Table 1 reports the frequency of firm-specific crashes for each year, starting in 1988 and ending in 2011. As reported at the bottom of Table 1, 18.5% of firm-years in the sample contain a crash. As noted by Kim et al. (2011a, 2011b), the frequency of crashes is generally larger in the 2000's than in the 1990's and peaks in 2008, which may be a by-product of the financial crisis of 2008-2009. Table 1 also reports the mean firm-specific and raw stock returns during crash weeks by year and for the total sample. On average, stock prices fall by 23.5% on an adjusted (firm-specific) basis and 24.5% on an unadjusted (raw) basis during the week of a crash.

We check the robustness of our conclusions using two alternative measures of crash risk derived by Chen et al. (2001) and employed by Kim et al. (2011a, 2011b), Callen and Fang (2013, 2015a, 2015b), and An and Zhang (2013). We briefly describe these variables below. Definitions of these variables as well as all other variables used in our empirical analyses can also be found in the appendix. The first alternative crash risk measure that we use is *NCSKEW,* which is defined as negative one multiplied by the skewness of the firm's weekly firm-specific log-returns during the given fiscal year. Skewness is computed as the sample third central moment divided by the sample standard deviation cubed. This measure captures the magnitude of leftward skewness in the distribution, with larger (or less negative) values of *NCSKEW*

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[5] *Throughout our paper, any references to "years" denote fiscal years (not calendar years) unless otherwise stated.*

[6] *We use the updated industry classification scheme of Fama and French (1997) that groups firms into 49 industries. Kenneth French has generously provided the algorithm for mapping SIC codes to the 49 Fama-French industries on his website: http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html.*

[7] *Under the assumption that firm-specific weekly log-returns are independently normally distributed, this definition of a crash implies roughly a 5% probability of observing a crash for a given firm in a given year. As shown in Table 1, the incidence of crashes in our sample is considerably higher than this, which is not surprising given that weekly log-returns are not normally distributed (although they do provide a closer approximation to normality than non-logged returns). Nonetheless, we use the 0.1% normal distribution cutoff as in Hutton et al. (2009) and Kim et al. (2011a, 2011b) as a convenient and reasonable means of identifying extreme return observations.*

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Table 1: Frequency of Stock Price Crashes and Average Crash Returns

This table reports the frequency of stock price crashes by fiscal year for firms in our sample. The table also reports average firm-specific and raw stock returns during weeks that contain a crash.

Fiscal Year	Number of Firms	Number of Firms with a Stock Price Crash	Percentage of Firms with a Stock Price Crash	Average Firm-Specific Return during Crash Weeks	Average Raw Return during Crash Weeks
1988	3,872	531	13.71%	-25.4%	-25.6%
1989	4,142	672	16.22%	-25.9%	-26.5%
1990	4,116	853	20.72%	-22.8%	-24.0%
1991	4,071	691	16.97%	-22.2%	-23.4%
1992	4,021	685	17.04%	-21.1%	-22.1%
1993	4,037	596	14.76%	-23.3%	-24.5%
1994	4,732	680	14.37%	-23.7%	-25.2%
1995	4,903	628	12.81%	-25.5%	-26.8%
1996	5,227	726	13.89%	-26.8%	-27.5%
1997	5,326	777	14.59%	-29.2%	-29.4%
1998	5,414	958	17.69%	-27.0%	-27.1%
1999	5,243	827	15.77%	-26.7%	-28.5%
2000	5,096	997	19.56%	-20.0%	-20.6%
2001	4,783	966	20.20%	-18.4%	-19.6%
2002	4,713	1,071	22.72%	-18.3%	-19.5%
2003	4,505	856	19.00%	-17.7%	-19.0%
2004	4,290	911	21.24%	-19.6%	-22.1%
2005	4,089	925	22.62%	-28.5%	-30.0%
2006	3,965	924	23.30%	-26.5%	-24.9%
2007	3,874	911	23.52%	-18.6%	-19.5%
2008	3,793	1,080	28.47%	-18.6%	-20.6%
2009	3,706	770	20.78%	-23.3%	-24.3%
2010	3,634	699	19.24%	-18.9%	-19.7%
2011	3,377	724	21.44%	-18.7%	-20.7%
Total	104,929	19,458	18.50%	-23.5%	-24.5%

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indicating greater crash risk. The second alternative measure of crash risk that we use is "down-to-up" volatility, *DUVOL,* which captures asymmetric volatilities between below-mean and above-mean returns. For a given firm-year, *DUVOL* equals the standard deviation of the firm's weekly firm-specific log-returns that are below the firm's mean in that year divided by the standard deviation of the firm's weekly firm-specific log-returns that are above the firm's mean in that year. As with *NCSKEW*, larger values of *DUVOL* are interpreted as indicating greater crash risk.

Table 2 reports sample descriptive statistics (mean, standard deviation, median, 25th percentile, and 75th percentile) for $NCSKEW_{t+1}$ and $DUVOL_{t+1}$, as well as for all other variables used in our analysis. Since our empirical analysis uses information from year *t* to predict crash risk in year *t*+1, throughout we subscript our crash risk measures with the *t*+1 subscript. As reported in Table 2, the sample mean and median of $NCSKEW_{t+1}$ is -0.174, indicating that weekly firm-specific log-returns are positively skewed for the typical firm-year in our sample. This is consistent with prior studies (Kim et al., 2011a, 2011b, Callen and Fang, 2013). The mean and median of $DUVOL_{t+1}$ are both below one, which indicates a tendency toward greater above-mean volatility than below-mean volatility for the typical firm-year, which is also consistent with prior studies.

Table 2: Summary Statistics for Sample Firms

This table reports summary statistics of variables used in our empirical analysis. The unit of observation is a firm-year (fiscal). All variables are defined in the appendix.

Variable	Mean	Standard Deviation	25th Percentile	Median	75th Percentile	N
Crash Risk Measures						
$CRASH_{t+1}$	0.185	0.389	0.000	0.000	0.000	104,929
$NCSKEW_{t+1}$	-0.174	0.882	-0.600	-0.174	0.234	104,929
$DUVOL_{t+1}$	0.983	0.495	0.711	0.900	1.142	104,929
Multinational Variable						
MN_t	0.328	0.470	0.000	0.000	1.000	104,929
Corporate Governance Variables						
SH_RIGHTS_t	3.528	1.420	3.000	3.000	5.000	29,828
$BOARD\_IND_t$	0.669	0.210	0.571	0.714	0.818	21,003
$IIO\_STABILITY_t$	0.0112	0.0026	0.0102	0.0117	0.0130	79,511
Baseline Control						
$SIZE_t$	5.402	2.128	3.838	5.278	6.879	104,929
BM_t	0.683	0.618	0.309	0.552	0.882	104,929
$LEVERAGE_t$	0.220	0.206	0.039	0.178	0.343	104,929
ROA_t	-0.034	0.224	-0.022	0.019	0.062	104,929
$DTURN_t$	0.003	0.075	-0.018	0.000	0.019	104,929
$NCSKEW_t$	-0.156	0.758	-0.578	-0.169	0.226	104,929
$SIGMA_t$	0.063	0.039	0.035	0.053	0.081	104,929
$ALPHA_t$	0.001	0.011	-0.004	0.001	0.006	104,929
Additional Control Variables						
$DIS\_ACCRUALS_t$	0.243	0.250	0.087	0.164	0.302	87,448
$LRETR_t$	0.348	0.302	0.135	0.302	0.416	90,078

Identifying Multinational Firms

Beginning in 1977, the U.S. Securities and Exchange Commission (SEC) rules governing geographic segment reporting require U.S. firms to report limited audited financial information for individual foreign segments (those located outside the U.S.) that account for 10% or more of the firm's total consolidated sales, profits, or assets. This information is reported in the Compustat Geographic Segment files and includes segment-specific sales and identifiable assets as well as other items that may be voluntarily

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reported by firms, such as segment-specific earnings and capital expenditures.[8] Following Denis et al. (2002), Gande et al. (2009), Tsao et al. (2016), Liu and Lao (2012), and Jiraporn et al. (2006), we define a firm as multinational in fiscal year *t* if it reports sales by one or more foreign segments that year. For the purposes of testing our hypotheses regarding differences in crash risk across multinational and domestic firms, we construct an indicator variable, MN_t, which equals one if the firm is multinational in year *t* and zero otherwise.

Table 2 reports sample descriptive statistics for MN_t. Since it is an indicator variable, the sample mean of MN_t (0.328) indicates that 32.8% of firms-years in the sample correspond to multinational firms. As would be expected, a year-by-year tabulation (unreported for brevity) reveals that the proportion of multinational firms has risen significantly over time, from 24% of sample firms in 1987 to 45% of sample firms in 2010. This trend is consistent with what prior studies have documented.

Corporate Governance Variables

We consider three important aspects of corporate governance that have been extensively studied in the finance and accounting literature: shareholder rights, the extent to which a firm's board is independent of top management, and institutional ownership.

We use the ISS Governance database to construct our measure of shareholder rights, which is based on the widely used "entrenchment index" developed by Bebchuk, Cohen, and Ferrell (2009). In their seminal study, Gompers, Ishii, and Metrick (2003) develop an index of shareholder rights comprised of twenty-four unique governance provisions included in the ISS database that affect the balance of power between a firm's managers and shareholders. These include provisions that managers can use to block unwanted takeovers, even when they are desired by a majority of shareholders, and, more generally, provisions that prevent a majority of shareholders from imposing their will on management. Gompers et al. (2003) show that firms with higher shareholder rights, as measured by lower values of their governance index, have significantly higher valuations and long-run stock returns. Bebchuk et al. (2009) extend the study of Gompers et al. (2003) by examining individually the twenty-four unique provisions that comprise the Gompers et al. (2003) governance index. They find that six of the twenty-four provisions are strongly associated with firm value and stock return performance, while the remaining 18 provisions in the Gompers et al. (2003) governance index are not. These six provisions include four constitutional provisions that limit the ability of shareholders to impose their will on management (staggered boards, limits to shareholder bylaw amendments, supermajority requirements for mergers, and supermajority requirements for charter amendments), and two explicit takeover prevention provisions (poison pills and golden parachutes). Based on these findings, Bebchuk et al. (2009) construct their "entrenchment index" by adding one point for each of the six provisions a firm has in place. The entrenchment index therefore ranges from zero to six, with higher values corresponding to lower shareholder rights.

We construct a measure that is increasing in shareholder rights, denoted as SH_RIGHTS_t, which equals six minus the firm's entrenchment index in year *t*.[9] SH_RIGHTS_t therefore ranges from zero to six, with higher values corresponding to higher shareholder rights.[10] Managers and directors of firms with lower

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Prior to 1998, firms were required to report segment-specific sales, identifiable assets, and earnings for each foreign segment. From 1998 onward, firms are only required to report sales and identifiable assets for each foreign segment. A peculiarity of the Compustat Geographic Segment files is that segment-specific identifiable assets are consistently missing for most firms, which could be due to a lack of standardization in how and where this information is reported in annual reports across different firms. Segment-specific sales, however, are consistently non-missing in the Compustat Geographic Segment files, which is why researchers using this database typically rely on sales by foreign segments to identify multinational firms.

[9] *We base our measure of shareholder rights on the entrenchment index of Bebchuck et al. (2009) rather than the governance index of Gompers et al. (2003) because Bebchuck et al. find that the six provisions comprising the entrenchment index are the most consequential with respect to their impact on shareholder value and because some of the twenty-four provisions comprising the Gompers et al. governance index are not reported in the ISS database after 2006, while all six provisions comprising the Bebchuk et al. entrenchment index are.*

[10] *We test H2 by regressing measures of crash risk on interactions between our governance measures and our indicator variable denoting multinationality. The coefficients on interactions in a regression are most easily interpreted when the variables have been*

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shareholder rights are viewed as being more insulated from the discipline of a hostile takeover or the activist efforts of shareholders, thus higher levels of shareholder rights correspond to better governance. In addition to having lower stock returns and lower valuations (Bebchuk et al., 2009, Chi and Lee, 2010, Cremers and Ferrell, 2014), empirical evidence suggests that firms with lower shareholder rights make more value destroying acquisitions (Harford et al., 2012, Masulis et al., 2007). Furthemore, Jiraporn et al. (2006) find that weaker shareholder rights is associated with substantially lower firm values in highly diversified firms, while Dittmar and Marhrt-Smith (2007), in their analysis of the market's valuation of firms' cash holdings, find that a dollar of cash held by a firm with strong shareholder rights has roughly twice the market value of a dollar of cash held by a firm with weak shareholder rights. Dittmar and Mahrt-Smith (2007) conclude that weaker shareholder rights are associated with lower cash holdings valuations because managers of such firms are more likely to waste cash. The ISS Governance database begins in 1990 and covers roughly 30% of the firms in our sample during that year, with coverage generally increasing over time to about 40% of our sample in 2010.[11] Thus, our regressions that include SH_RIGHTS_t will have a lower number of reported observations than that of the full baseline sample.

The next corporate governance variable that we consider is the fraction of independent outsiders on the board of directors. We use the ISS Directors database to construct our measure of board independence, $BOARD\_IND_t$, which equals the number of independent directors (as defined by ISS) divided by the total number of directors on the firm's board in year *t*. ISS classifies a director as independent if he/she is not employed by the firm in any other capacity and has no material connections to the firm other than holding a board seat. Directors who are captured by management, especially those employed directly under the CEO, face significant conflicts of interest that reduce their incentives to effectively monitor the firm's managers. Independent directors are presumed not to be hampered by such conflicts of interest, and are thus viewed as being more effective monitors. Empirical evidence indicates that firms with more independent boards make more voluntary disclosures (Lim et al., 2007), are less likely to commit corporate fraud (Beasley, 1996, Uzun et al., 2004), and less likely to manipulate reported earnings (Klein, 2002, Osma, 2008). Furthermore, Weisbach (1988) finds that CEO turnover is more sensitive to firm performance when the board is dominated by independent outsiders, while other works suggest that firms with more independent boards make better decisions with respect to tender offer bids (Byrd and Hickman, 1992) and the adoption of poison pills (Brickley et al., 1994). Our measure of board independence, $BOARD\_INT_t$, ranges from zero to one, with higher values corresponding to greater board independence and more effective monitoring. The ISS Directors database begins in 1996 and covers roughly 23% of the firms in our baseline sample during that year, with coverage generally increasing over time to 39% of our sample in 2010.

Our final governance variable, which measures the stability of a firm's institutional ownership, is based on Callen and Fang (2013) and is computed using data from Thomson Reuters' Institutional (13f) Holdings database. Callen and Fang (2013) hypothesize that stable (or long-term) institutional investors are more likely to be concerned with long-term value maximization, rather than short-term accounting profits, and thus more likely to have strong monitoring incentives, which should lessen bad news hoarding and crash risk. On the other hand, investors focused on short-term trading profits (transient investors) should have little or no incentives to monitor management. Callen and Fang (2013) further argue that transient institutional investors may even exacerbate bad news hoarding and crash risk due to greater pressure on managers to maximize short-term earnings at the expense of long-term value. Using a measure originally developed by Elyasiani et al. (2010), which measures the volatility, or instability, of a firm's institutional ownership, Callen and Fang (2013) document a significant positive relationship between the instability of a firm's institutional ownership and one-year-ahead crash risk, leading them to conclude that more stable institutional ownership mitigates crash risk, while less stable institutional ownership exacerbates crash risk.

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constructed as intuitively as possible. We therefore construct all our governance variables so that they have minimum values of zero and are increasing with better governance.

[11] *Between 1990 and 1997, the ISS Governance database covers firms in the S&P 1500. After 1997, coverage is expanded to include some smaller and mid-size firms.*

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For the purposes of our study, we begin by constructing the variable used by Callen and Fang (2013), denoted as *StdI* in their study,[12] which is defined as the average standard deviation of quarterly institutional shareholding proportions across all institutional investors in the firm over a 5-year period (year *t*-4 to year *t*). Specifically, for a given firm *i* in fiscal year *t*, *StdI* is computed using data on institutional shareholdings over the twenty quarters comprising years *t*-4 to *t* as follows:

$$StdI_i = \sum_{j=1}^{J_i} Std(p_{i,q}^j)/J_i \qquad (2)$$

where $p_{i,q}^j$ is the proportion of firm *i* held by institutional investor *j* at quarter *q* (*q* = 1, 2, …, 20), $Std(p_{i,q}^j)$ is the time series standard deviation of institutional investor *j*'s quarterly shareholding proportions in firm *i* during the twenty quarters comprising fiscal years *t*-4 to *t*, and J_i is the number of institutional investors in firm *i.* The rationale for this measure is based on the idea that institutional investors focused on short-term trading profits should exhibit more volatile shareholding proportions, while investors focused on long-term value maximization should exhibit less volatile shareholding proportions. *StdI* is therefore viewed as an inverse measure of a firm's institutional ownership stability. To obtain a measure that is increasing in institutional ownership stability, we first winsorize *StdI* at the 1st and 99th sample percentiles as in Callen and Fang (2013) and then perform the following transformation to arrive at our measure of institutional ownership stability for each firm in the sample:

$$IIO\_STABILITY_{i,t} = \max(StdI) - StdI_{i,t} \qquad (3)$$

where max(*StdI*) is the maximum value of *StdI* across all firm-years in the sample. As with our other corporate governance measures, *IIO_STABILITY* has a sample minimum value of zero. As previously discussed, Callen and Fang (2013) find that firm's with more stable institutional ownership have lower one-year-ahead crash risk. In addition, McCahery et al. (2016), in their survey of institutional investors, provide evidence that long-term institutional investors actively monitor and discipline managers through "behind the scenes" interventions, which lends further support to the notion that managers of firms with more stable institutional ownership are subject to better monitoring.

Baseline Control Variables

We derive our set of control variables from Chen et al. (2001), Hutton et al. (2009), Kim et al. (2011a, 2011b), Callen and Fang (2013, 2015a, 2015b), and An and Zhang (2013). We briefly define the control variables below and report standard summary statistics for these variables in Table 2. All monetary variables are measured in millions of 2012 U.S. dollars using the consumer price index to adjust for inflation. To mitigate the potential influence of extreme outliers in our regressions, we follow Callen and Fang (2013, 2015a, 2015b) and winsorize all control variables at the 1st and 99th percentiles, although we have verified in untabulated results that our conclusions hold without winsorizing the control variables. Detailed definitions of each control variable can also be found in the appendix.

Our baseline set of controls includes the following variables: $SIZE_t$, BM_t, $LEVERAGE_t$, ROA_t, $DTURN_t$, $NCSKEW_t$, $SIGMA_t$, and $ALPHA_t$. $SIZE_t$ is defined as the natural logarithm of the market value of the firm's common equity at the close of fiscal year *t*. BM_t is the book value of the firm's common equity divided by the market value of common equity and $LEVERAGE_t$ is the firm's total debt (long-term debt plus short-term debt) divided by the book value of total assets, all measured at the close of year *t*. ROA_t is income before extraordinary items scaled by the book value of total assets at the close of year *t*, which measures the firm's accounting profitability. Prior studies have generally found that crash risk is positively related to firm size and negatively related to the book-to-market ratio. Findings regarding measures of accounting profitability are mixed, with some studies finding that more profitable firms have higher future crash risk (Callen and Fang, 2013, 2015a, 2015b) and others finding that more profitable firms have lower future crash risk (Hutton et al., 2009; Kim et al., 2011a, 2011b). Similarly, findings regarding the relation between

[12] *See Equation (1) in Callen and Fang (2013).*

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leverage and crash risk are also mixed in prior studies.[13] $DTURN_t$ is the detrended average monthly share turnover in year t, which equals the average monthly share turnover during year t minus the average monthly share turnover during year t-1, where monthly share turnover is calculated as monthly trading volume (number of shares) divided by the total number of shares outstanding during the month. This variable is viewed as a proxy for differences of opinion among investors (Chen et al., 2001). The variable $NCSKEW_t$ equals negative one multiplied by the skewness coefficient of the firm's weekly firm-specific log-returns in year t. The variable $SIGMA_t$ is the standard deviation of the firm's weekly firm-specific log-returns in year t. $DTURN_t$ and $NCSKEW_t$ generally exhibit a positive relationship with one-year-ahead crash risk in prior studies (Chen et al., 2001; Kim et al., 2011a, 2011b; Callen and Fang, 2013, 2015a, 2015b). Prior studies have documented mixed results regarding the relation between crash risk and $SIGMA_t$.[14] $ALPHA_t$ is defined as the natural logarithm of one plus the estimated intercept from the firm's five-factor model estimated during year t (α from Equation (1)). This variable measures the firm's idiosyncratic stock return performance during year t. Chen et al. (2001) find that crash risk is higher among firms with higher past stock return performance.

Additional Control Variables

In some of our regressions, we also include variables that capture the firm's earnings management and tax avoidance activities as additional controls. Hutton et al. (2009) argue that reported earnings that include large (in absolute value) discretionary accruals, which managers can use to manipulate earnings, are less transparent and therefore more difficult for investors to assess. They therefore hypothesize that crash risk should be greater among firms with larger (in absolute value) discretionary accruals. Their findings indicated that crash risk increases at a decreasing rate as the absolute value of past discretionary accruals increases, i.e., crash risk is positively related to the absolute value of discretionary accruals and negatively related to the square of the absolute value of discretionary accruals. We therefore include in some of our regressions the variable $DIS\_ACCRUALS_t$, which is defined as the sum of the absolute value of annual discretionary accruals over the last three fiscal years (years t, t-1, and t-2). This variable is constructed exactly as in Hutton et al. (2009) and Kim et al. (2011a) using the modified Jones model of discretionary accruals originally developed by Dechow et al. (1995).[15] Since Hutton et al. (2009) find that the positive relation between crash risk and $DIS\_ACCRUALS_t$ declines at higher levels of $DIS\_ACCRUALS_t$, we also include the square of this variable in our regressions. Kim et al. (2011a) argue that aggressive tax avoidance activities, like complex tax shelters, can create opportunities for managers to conceal negative information and mislead investors. They find empirical support for their hypothesis that tax avoidance is positively associated with crash risk. Following Kim et al. (2011a), in some of our regressions specifications, we include the variable $LRETR_t$, which is a measure of the firm's long-run effective tax rate. $LRETR_t$ is defined as in Dyreng et al. (2008) and Kim et al. (2011a) and is computed as total taxes paid during the last five fiscal years (years t-4 to t) divided by total pre-tax income net of special items during the same period. The Compustat data items used to construct the variables $DIS\_ACCRUALS_t$ and $LRETR_t$ are missing for a nontrivial portion of firm-years in our sample, hence the regression specifications that include these variables have fewer observations than our baseline specification.

Regression Models

We test H1 in a multivariate setting using the following regression model:

$$CRASH_{i,t+1} = \alpha_0 + \alpha_1 MN_{i,t} + \sum_{q=2}^{m} \alpha_q (q^{th}\ Control_{i,t}) + \mu_{i,t} \quad (4)$$

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[13] *Hutton et al. (2009) and Callen and Fang (2013) find that leverage is significantly and negatively related to crash risk. In other studies, however, the coefficient on leverage is either significantly positive or insignificant in many cases (Kim et al., 2011a, 2011b, An and Zhang, 2013).*

[14] *Chen et al. (2001) find that future crash risk, as measured by NCSKEW, is negatively related to past return volatility, as do An and Zhang (2013) using multiple measures of crash risk. On the other hand, Kim et al. (2011a, 2011b) and Callen and Fang (2015b) find that crash risk (using multiple measures) is positively related to past return volatility, while Callen and Fang (2013, 2015a) report mixed results across different measures of crash risk.*

[15] *This variable is denoted as OPAQUE in Hutton et al. (2009) and ACCM in Kim et al. (2011a).*

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where for firm *i* in fiscal year *t*, $CRASH_{i,t+1}$ is the previously defined binary variable denoting whether the firm crashes in year *t+1*, $MN_{i,t}$ is the previously defined indicator variable denoting whether the firm is multinational or domestic in year *t*, and the control variables are previously defined. As in Hutton et al. (2009) and Kim et al. (2011a, 2011b), we estimate this model using a logistic regression that includes year dummies and industry dummies corresponding to the 49 Fama-French industries based on the updated industry classification scheme of Fama and French (1997). A positive and statistically significant estimate of the coefficient α_1 indicates that stock price crashes are more likely to occur among multinational firms than domestic firms, thus we interpret a positive and significant estimate of α_1 in equation (4) as evidence in favor of H1.

We test H2 by augmenting the right hand side of Equation (4) with the previously defined corporate governance variables, *SH_RIGHTS*, *BOARD_IND*, and *IIO_STABILITY*, and their interactions with *MN*, which produces the following regression specification:

$$CRASH_{i,t+1} = \alpha_0 + \alpha_1 MN_{i,t} + \alpha_2 MN_{i,t} \times SH\_RIGHTS_{i,t} + \alpha_3 MN_{i,t} \times BOARD\_IND_{i,t} + \alpha_4 MN_{i,t} \times IIO\_STABILITY_{i,t} + \alpha_5 SH\_RIGHTS_{i,t} + \alpha_6 BOARD\_IND_{i,t} + \alpha_7 IIO\_STABILITY_{i,t} + \sum_{q=8}^{m} \alpha_q (q^{th}\ Control_{i,t}) + \mu_{i,t} \quad (5)$$

This specification allows the relation between crash risk and multinationality to vary across different levels of the corporate governance variables, *SH_RIGHTS*, *BOARD_IND*, and *IIO_STABILITY*. H2 predicts that the positive relation between crash risk and multinationality is attenuated by stronger governance, thus H2 implies significantly negative estimates of α_2, α_3, and α_4.

As previously discussed, we check the robustness of our conclusions using two alternative measures of one-year-ahead crash risk, $NCSKEW_{t+1}$ and $DUVOL_{t+1}$.When these crash risk measures are used as the dependent variable, we estimate the regression models with OLS that include industry and year dummies, as in Kim et al. (2011a, 2011b), Callen and Fang (2013, 2015) and An and Zhang (2013).

Results and Discussion

Univariate Comparisons

We begin our empirical analysis by performing univariate comparisons of crash risk across multinational and domestic firms. In Table 3, we first compare the proportion of firms with a crash in year *t*+1 ($CRASH_{t+1}$ = 1) across multinational (MN_t = 1) and domestic firms (MN_t = 0). As reported in Table 3, domestic firms exhibit an unconditional one-year-ahead crash probability of 0.175, indicating that 17.5% of the firm-years classified as domestic experience a crash in the following year. In contrast, multinational firms exhibit an unconditional one-year-ahead crash probability of 0.207 (or 20.7%), yielding a difference between the two subsamples of 3.2 percentage points. This difference is highly significant (*z*-statistic = 12.63) and corresponds to an 18% (0.032/0.175) rise in the probability of crashing. Table 3 also reports mean and median values of our alternative crash risk measures, $NCSKEW_{t+1}$ and $DUVOL_{t+1}$, across the two subsamples. The mean (median) values of $NCSKEW_{t+1}$ are -0.209 (-0.119) for domestic firms and -0.103 (-0.124) for multinational firms. The difference in means (and medians) is highly significant, indicating that multinational firms have significantly higher crash risk as measured by $NCSKEW_{t+1}$. Likewise, multinational firms have significantly higher mean and median values of $DUVOL_{t+1}$. Overall, the results of these univariate tests indicate that, on an unconditional basis, one-year-ahead crash risk is indeed higher among multinational firms.

Table 3: Multinationality and Crash Risk: Univariate Comparisons

This table compares measures of one-year-ahead crash risk across multinational firms (MNt = 1) and domestic firms (MNt = 0). All variables are defined in the appendix. We use a two-proportion z-test to test for statistical significance of the difference in the proportion of firms with a crash. We use a two-sample t

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test (Wilcoxon test) to test for statistical significance of the difference in mean (median) values of NCSKEWt+1, and DUVOLt+1. *, **, and *** indicate statistical significance of the difference at the 10%, 5%, and 1% levels respectively in a two-tailed test.

	Multinational Firms (MN_t = 1)	Domestic Firms (MN_t = 0)	Test statistic (difference)
No. of Firm-Years	34,455	70,474	
Proportion of firm-years with a crash in year t+1 ($CRASH_{t+1}$ = 1)	0.207	0.175	12.63***
Mean of $NCSKEW_{t+1}$	-0.103	-0.209	18.59***
Median of $NCSKEW_{t+1}$	-0.124	-0.199	19.83***
Mean of $DUVOL_{t+1}$	1.009	0.971	12.33***
Median of $DUVOL_{t+1}$	0.927	0.887	18.69***

Regression Analysis

In the first two models reported in Table 4, we perform formal tests of H1 with logistic regressions that model the probability of a stock price crash as a function of whether the firm is multinational or domestic and the previously described set of control variables. The table reports coefficients along with test statistics in parentheses, which are based on standard errors adjusted for heteroscedasticity and clustering by firm.

After controlling for known predictors of crash risk, we find robust evidence that multinational firms are indeed more likely to crash than domestic firms. Model (1) in Table 4 reports the results of a logistic regression of $CRASH_{t+1}$ on MN_t and the baseline control variables. The coefficient on MN_t is positive and statistically significant at the 5% level, indicating that one-year-ahead crash risk is higher among multinational firms than among domestic firms. In model (2) of Table 4, which includes the additional control variables, $DIS\_ACCRUALS_t$ and $LRETR_t$, we again observe a positive coefficient on MN_t that is now significant at the 1% level. The results from models (1) and (2) in Table 4 therefore provide evidence in support of H1.

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Table 4: Multinationality and Crash Risk: Regression Analysis

This table reports coefficient estimates from regressions where the dependent variable is CRASHt+1, NCSKEWt+1 or DUVOLt+1. All variables are defined in the appendix. The unit of observation is a firm-year (fiscal). When the dependent variable is CRASHt+1, we use a logistic regression. When the dependent variable is NCSKEWt+1 or DUVOLt+1, we use OLS. All regressions include fiscal year dummies and industry dummies (coefficients unreported) that correspond to the updated 49 Fama-French industries (Fama and French, 1997). Z-statistics (t-statistics) are reported in parentheses below the logistic (OLS) coefficients and are based on heteroskedastic-consistent standard errors adjusted for clustering by firm. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels respectively in two-tailed tests.

Dependent Variable	$CRASH_{t+1}$				$NCSKEW_{t+1}$		$DUVOL_{t+1}$	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
MN_t	0.044*	0.063*	1.377	1.503	0.414*	0.455*	0.227*	0.256**
	(2.03)	(2.66)	(3.55)	(3.73)	(3.05)	(3.11)	(3.26)	(3.45)
MN_t x SH_RIGHTS_t			-	-	-	-	-	-0.010*
			(-	(-	(-2.00)	(-1.94)	(-1.80)	(-1.92)
MN_t x $BOARD\_IND_t$			-	-	-	-	-	-
			(-	(-	(-2.82)	(-2.79)	(-2.59)	(-2.66)
MN_t x			-	-	-	-	-	-
			(-	(-	(-1.74)	(-1.78)	(-2.09)	(-2.16)
SH_RIGHTS_t			0.019	0.012	0.011	0.010	0.001	0.002
			(0.85)	(0.50)	(1.38)	(1.19)	(0.33)	(0.45)
$BOARD\_IND_t$			0.396	0.390	0.076	0.076	0.043*	0.045
			(2.58)	(2.25)	(1.59)	(1.40)	(1.67)	(1.55)
$IIIO\_STABILITY_t$			-25.70	-	-14.20	-12.29	-7.544	-6.514
			(-	(-	(-1.50)	(-1.14)	(-1.59)	(-1.19)
$SIZE_t$	0.014*	0.027*	-0.005	-	0.047*	0.043*	0.020*	0.017**
	(2.17)	(3.63)	(-	(-	(6.05)	(5.03)	(4.95)	(3.76)
BM_t	-	-	-0.089	-	-	-	-	-
	(-4.86)	(-3.64)	(-	(-	(-2.88)	(-2.91)	(-3.11)	(-3.13)
$LEVERAGE_t$	0.087*	0.070	0.138	0.064	0.004	-0.027	0.017	0.002
	(1.91)	(1.38)	(1.17)	(0.50)	(0.09)	(-0.56)	(0.76)	(0.10)
ROA_t	0.062	0.049	0.807	0.847	0.307*	0.255*	0.181*	0.152**
	(1.33)	(0.97)	(3.51)	(3.54)	(4.18)	(3.21)	(4.92)	(3.84)
$DTURN_t$	0.708*	0.642*	0.379	0.397	0.210*	0.226*	0.101*	0.107*
	(6.24)	(5.24)	(1.44)	(1.43)	(2.21)	(2.25)	(1.92)	(1.93)
$NCSKEW_t$	0.145*	0.141*	0.092	0.084	0.031*	0.031*	0.014*	0.013**
	(12.37)	(10.60)	(3.52)	(2.94)	(3.08)	(2.83)	(2.74)	(2.47)
$SIGMA_t$	-	-	0.687	0.022	0.963*	0.823*	0.419*	0.374
	(-4.91)	(-5.23)	(0.61)	(0.02)	(2.36)	(1.81)	(2.00)	(1.63)
$ALPHA_t$	4.253*	4.508*	3.817	4.509	3.853*	4.153*	1.822*	2.029**
	(4.72)	(4.49)	(1.33)	(1.48)	(3.78)	(3.82)	(3.39)	(3.56)
$DIS\_ACCRUALS_t$		0.480*		0.159		-0.056		-0.040
		(4.30)		(0.52)		(-0.54)		(-0.75)
$DIS\_ACCRUALS_t^2$		-		0.140		0.109		0.064
		(-4.05)		(0.48)		(1.12)		(1.24)
$LRETR_t$		-		-		-		-
		(-3.16)		(-		(-2.23)		(-3.38)
Pseudo R^2 / R^2	0.018	0.019	0.023	0.024	0.027	0.028	0.027	0.028
No. of Observations	104,92	77,191	17,25	14,85	17,25	14,85	17,25	14,854

In Models (3) and (4) of Table 4, we test H2 by repeating the logistic regressions in models (1) and (2) while adding the three corporate governance variables SH_RIGHTS_t, $BOARD\_IND_t$, and $IIO\_STABILITY_t$ and their interactions with MN_t. In both of these models, the coefficient on MN_t is positive and significant at the 1% level, while the coefficients on the interactions of MN_t with SH_RIGHTS_t, $BOARD\_IND_t$, and $IIO\_STABILITY_t$ are all negative and significant at the 5% level. These findings strongly support H2, since

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they indicate that the positive relation between the probability of a crash and multinationality diminishes as the strength of each corporate governance mechanism increases. In models (5) and (6), which include $NCSKEW_{t+1}$ as the dependent variable, the coefficient on MN_t is positive and significant at the 1% level, while the coefficients on the interactions of MN_t with SH_RIGHTS_t, $BOARD\_IND_t$, and $IIO\_STABILITY_t$ are all negative and significant at the 10% level or better. We repeat the same two regressions in specifications (7) and (8) using $DUVOL_{t+1}$ as the dependent variable. We again observe positive and highly significant coefficient on MN_t and negative coefficients on the interactions of MN_t with the three governance variables, which are significant at the 10% level or better. The results from Table 4 therefore provide robust evidence in support of H1 and H2. Summarily, we find that one-year-ahead crash risk is higher among multinational firms than domestic firms, all else equal. This positive relation is attenuated by stronger corporate governance mechanisms, as measured by stronger shareholder rights, greater board independence, and more stable institutional ownership.

In Table 4, the coefficient estimates on the control variables are generally consistent with the findings of prior works. $SIZE_t$, $DTURN_t$, $NCSKEW_t$, and $ALPHA_t$ have positive and significant coefficients in all or a majority of the models, while BM_t, has a negative and significant coefficient in a majority of the models. These results are in line with Chen et al. (2001), Kim et al. (2011a, 2011b), Hutton et al. (2009), Callen and Fang (2013, 2015a, 2015b), and An and Zhang (2013).

The coefficient on ROA_t is positive and significant in most of the models in Table 4, which is consistent with the results of Callen and Fang (2013, 2015a, 2015b). The coefficients on $LEVERAGE_t$ and $SIGMA_t$ exhibit inconsistencies in either their signs or statistical significance across the different models in Table 4. This is unsurprising, given that prior studies have reported mixed results regarding these two variables. $LRETR_t$ has a negative and significant coefficient in three of the four models in which it is included, indicating that firms with lower long-run effective tax rates have greater crash risk, which is consistent with the main findings of Kim et al. (2011a). In addition, in model (2) of Table 4, $DIS\_ACCRUALS_t$ has a significantly positive coefficient while the square of this variable has a significantly negative coefficient. This is consistent with the findings of Hutton et al. (2009). In the remaining three models that include this variable (models (4), (6), and (8)), neither $DIS\_ACCRUALS_t$ nor its square is significant. This is unsurprising given that the samples used to estimate those three models are necessarily limited to the year 1996 and years thereafter, since they include the variable $BOARD\_IND_t$ (which is available starting in 1996). Hutton et al. (2009) find that the significant relation between one-year-ahead crash risk and discretionary accruals (which they denote as *OPAQUE* in their study) disappears completely after the passage of the Sarbanes Oxley (SOX) Act in 2002. More than half the observations used to estimate models (4), (6), and (8) in Table 4 are from the post-SOX period, while less than a third of the observations used to estimate model (2) are from the post-SOX period, which explains the significance of $DIS\_ACCRUALS_t$ in model (2) and the lack of significance in models (4), (6), and (8).

In a logistic regression, the coefficients do not represent marginal effects as they do in a linear regression. The marginal effect of a particular independent variable on the dependent variable in a logistic regression depends on the value of that particular independent variable as well as the values of all other independent variables in the model. To elucidate the economic significance of the impact of multinationality on the probability of a stock price crash, as well as how this impact varies with the strength of the firm's corporate governance mechanisms, in Table 5 we report marginal effects that correspond to the marginal difference in crash probability between multinational and domestic firms at different strengths of the three corporate governance mechanisms. In Panel A of Table 5, we report marginal effects derived from model (4) of Table 4 that represent the marginal difference in crash probability between multinational (MN_t=1) and domestic firms (MN_t=0) at different levels of SH_RIGHTS_t, while holding all other independent variables in the model (including $BOARD\_IND_t$ and $IIO\_STABILITY_t$) at their sample means, which is common convention when estimating marginal effects derived from a logistic regression. The second column in Panel A of Table 5 shows that, among firms with the weakest shareholder rights (SH_RIGHTS_t = 0) the marginal difference in the probability of a crash between multinational and domestic firms is 0.0679 (or 6.79 percentage points). Given that the unconditional crash probability in our sample is 0.185 (or 18.5%, see Table 2), the estimate of 0.0679 is economically meaningful. Among firms with the next lowest value of one for SH_RIGHTS_t, the

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second column in Panel A of Table 5 reports a slightly lower marginal difference in one-year-ahead crash probability between multinational and domestic firms of 0.0542, which is again economically meaningful. As shown in Panel A, the marginal effect of MN_t on one-year-ahead crash probability continues to decline as shareholder rights increase. In fact, among firms that have the two highest values for SH_RIGHTS_t of five and six, the marginal differences in crash probability between multinationals and domestics, 0.0030 and -0.0089 respectively, are small and close to zero.

Table 5: Marginal Effect of Multinationality on Crash Risk at Different Levels of Corporate Governance

This table reports marginal effects derived from regression Models (4), (6), and (8) of Table 4 that correspond to the marginal difference in crash risk between multinational firms (MN_t=1) and domestic firms (MN_t=0) at different levels of the three corporate governance variables.

	Marginal Effect of MN_t on $CRASH_{t+1}$ [from Table 4, Model (4), holding other independent variables at their sample means]	Marginal Effect of MN_t on $NCSKEW_{t+1}$ [from Table 4, Model (6), holding the other two governance variables at their sample means]	Marginal Effect of MN_t on $DUVOL_{t+1}$ [from Table 4, Model (8), holding the other two governance variables at their sample means]
Panel A: Marginal effect of Multinationality on Crash Risk at Different Levels of Shareholder Rights			
SH_RIGHTS_t			
0 (min)	0.0679	0.1028	0.0534
1	0.0542	0.0820	0.0430
2	0.0409	0.0612	0.0326
3	0.0279	0.0403	0.0221
4	0.0153	0.0195	0.0117
5	0.0030	-0.0013	0.0012
6 (max)	-0.0089	-0.0221	-0.0092
Panel B: Marginal effect of Multinationality on Crash Risk at Different Levels of Board Independence			
$BOARD\_IND_t$			
0.00 (min)	0.0803	0.1638	0.0840
0.40 (10th %ile)	0.0454	0.0833	0.0437
0.50 (20th %ile)	0.0364	0.0632	0.0336
0.60 (30th %ile)	0.0274	0.0431	0.0235
0.67 (40th %ile)	0.0213	0.0297	0.0168
0.71 (50th %ile)	0.0170	0.0202	0.0120
0.75 (60th %ile)	0.0137	0.0130	0.0084
0.80 (70th %ile)	0.0091	0.0029	0.0034
0.86 (80th %ile)	0.0038	-0.0085	-0.0024
0.89 (90th %ile)	0.0009	-0.0150	-0.0056
1.00 (max)	-0.0095	-0.0373	-0.0168
Panel C: Marginal effect of Multinationality on Crash Risk at Different Levels of Institutional Ownership Stability			
$IIO\_STABILITY_t$			
0.0000 (min)	0.2145	0.2469	0.1514
0.0081 (10th %ile)	0.0678	0.0903	0.0544
0.0097 (20th %ile)	0.0428	0.0587	0.0348
0.0106 (30th %ile)	0.0300	0.0417	0.0242
0.0112 (40th %ile)	0.0212	0.0294	0.0166
0.0117 (50th %ile)	0.0139	0.0189	0.0102
0.0122 (60th %ile)	0.0073	0.0092	0.0041
0.0127 (70th %ile)	0.0009	-0.0005	-0.0019
0.0133 (80th %ile)	-0.0054	-0.0104	-0.0080
0.0138 (90th %ile)	-0.0120	-0.0209	-0.0145
0.0145 (max)	-0.0202	-0.0348	-0.0232

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The second column in Panel B of Table 5 shows how the marginal difference in crash probability between multinationals and domestics changes at different levels of board independence. The values we display for $BOARD\_IND_t$ correspond to the sample minimum of this variable (zero), followed by the sample 10th , 20th, 30th, 40th, 50th, 60th, 70th, 80th, and 90th percentile values and, finally, the sample maximum value of $BOARD\_IND_t$ (one). As shown in the second column in Panel B of Table 5, the marginal effect of MN_t on one-year-ahead crash probability tends to be large and positive at lower levels of $BOARD\_IND_t$ and declines steadily to trivially small values at relatively high levels of $BOARD\_IND_t$. In the second column of Panel C, Table 5, we repeat the same exercise showing how the marginal effect of MN_t on one-year-ahead crash probability changes at different levels of institutional ownership stability, and we again observe a similar pattern of relatively large and positive marginal effects at low levels of $IIO\_STABILITY_t$ which decline to trivially small values at the highest levels of $IIO\_STABILITY_t$. In the third and fourth columns of Table 5, we repeat the same kind of tabulation for the marginal effect of MN_t on $NCSKEW_{t+1}$ and $DUVOL_{t+1}$ using coefficient estimates from models (6) and (8), respectively, in Table 4. For each governance variable, these marginal effects are computed while holding the other two governance variables at their sample means. Thus, they can be directly inferred from the coefficients reported in Table 4 and the sample means reported in Table 2, since these models are estimated with linear regressions. As Table 5 shows, the marginal effect of MN_t on these crash risk measures is relatively large and positive at low levels of each of the three corporate governance variables and tends to decline to much smaller values as the strength of the given governance variable increases. Overall, the results in Table 5 indicate economically significant differences in crash risk between multinational and domestic firms with weaker shareholder rights, less independent boards, or less stable institutional ownership. In addition, Table 5 also illustrates that each of these three governance mechanisms has an economically significant moderating effect on the impact of multinationality on crash risk.

Additional Tests and Robustness Checks

We perform a number of additional tests and robustness checks of our main results. For brevity, we discuss these results below but do not tabulate them in tables.

We have considered several other corporate governance variables that have been studied in the finance literature. These include board size, defined as the natural logarithm of the number of directors on the board, whether the CEO is also the chairman of the board, and blockholder ownership, defined as the proportion of outstanding shares owned by institutional investors who own 5% or more of the firm's outstanding shares. However, we do not find significant evidence that these variables are related to crash risk or that they affect the relation between crash risk and multinationality in our sample.

We examine whether the results from our tests persist when we use an expanded version of the five-factor model to estimate firm-specific weekly returns. Specifically, we augment Equation (1) with one weekly lag and one weekly lead of each factor, yielding the following model:

$$R_{i,\tau} = \alpha_i + \beta_{1i}R_{m,\tau-1} + \beta_{2i}SMB_{\tau-1} + \beta_{3i}HML_{\tau-1} + \beta_{4i}UMD_{\tau-1} + \beta_{5i}IND_{i,\tau-1} + \beta_{6i}R_{m,\tau} + \beta_{7i}SMB_{\tau} + \beta_{8i}HML_{\tau} + \beta_{9i}UMD_{\tau} + \beta_{10i}IND_{i,\tau} + \beta_{11i}R_{m,\tau+1} + \beta_{12i}SMB_{\tau+1} + \beta_{13i}HML_{\tau+1} + \beta_{14i}UMD_{\tau+1} + \beta_{15i}IND_{i,\tau+1} + \varepsilon_{i,\tau} \quad (6)$$

The leads and lags are included to correct for potential biases in the loading coefficients caused by non-synchronous trading (Dimson, 1979). Most prior studies of crash risk use an expanded version of the market model. However, we view the five-factor model as more robust because it captures systematic size, value, momentum, and industry effects. We estimate the model in Equation (6) for each firm-year in the sample and then use the log-transformed residuals to compute our crash risk measures. We find that our results remain robust.

Finally, we rerun our regressions in Table 4 using alternative controls for size, leverage, profitability, and past stock return performance. For size, we use the natural logarithm of total assets. For leverage, we use market leverage, defined as the book value of debt divided by the market value of assets, where the market value of assets equals book assets minus the book value of common equity plus the market value of common equity. For profitability, we use return-on-equity as in Hutton et al. (2009), defined as income

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before extraordinary items scaled by the book value of common equity. For past stock return performance, we use the firm's annual market adjusted stock return in year *t* as in Chen et al. (2001), defined as the firm's raw stock return minus the return on the CRSP value-weighted index. We find that our regression results are robust to these alternative specifications.

Conclusion

In a large panel of U.S. firms during 1987–2011, we find robust evidence that multinational firms are more likely than domestic firms to experience extreme stock price crashes. Our findings are robust to alternative measures of crash risk as well as controlling for known predictors of one-year-ahead crash risk identified by extant literature, including firm size, leverage, profitability, book-to-market ratio, past stock return performance, return volatility, return skewness, trading volume, discretionary accruals, and tax avoidance. We also show that the effect of multinationality on crash risk is especially strong among firms with weaker corporate governance mechanisms, specifically weaker shareholder rights, less independent boards, and less stable institutional ownership. We find that better governance along each of these three dimensions substantially reduces the impact of multinationality on crash risk. Our findings are consistent with a greater propensity for bad news hoarding in multinational firms relative to domestic firms that is especially strong among firms with weaker governance but significantly attenuated in firms with stronger governance.

Our paper makes several important contributions. First, our research adds to literature in international business, finance, and accounting viewing multinational firms from an agency perspective by documenting an economically significant, and hitherto unknown, consequence of heighted agency problems in multinational firms. Second, our research extends recent academic efforts focused on understanding and modeling tail risk in financial markets and, more specifically, the growing literature aimed at identifying firm characteristics and corporate activities that portend firm-specific stock price crashes. Our research shines the spotlight on an observable firm characteristic with a statistically and economically significant association with future crash risk. Given that several recent studies suggest that investors are concerned with tail risk in the returns of individual stocks (Conrad et al., 2013; Boyer et al., 2010; Yan, 2011; Barberis and Huang, 2008; Brunnermeier et al., 2007; Mitton and Vorkink, 2007), our findings should be especially useful to market participants, practitioners, and scholars interested in developing predictive models of ex ante tail risk in individual stock returns.

References

An, H. and Zhang, T. (2013). Stock price synchronicity, crash risk, and institutional investors. *Journal of Corporate Finance* 21, 1-15. DOI: 10.1016/j.jcorpfin.2013.01.001.

Ashbaugh, H. and Pincus, M. (2001). Domestic Accounting Standards, International Accounting Standards, and the Predictability of Earnings. *Journal of Accounting Research* 39, 417- 434. DOI: 10.1111/1475-679X.00020.

Bakshi G., Cao, C. and Chen Z. (1997). Empirical performance of alternative option pricing models. *Journal of Finance* 52, (2003)–2049. DOI: 10.1111/j.1540-6261.1997.tb02749.x.

Ball, R. (2009). Market and Political/Regulatory Perspectives on the Recent Accounting Scandals. *Journal of Accounting Research* 47, 277–323. DOI: 10.1111/j.1475-679X.2009.00325.x.

Barberis, N. and Huang, M., (2008). Stocks as Lotteries: The Implications of Probability Weighting for Security Prices. *American Economic Review* 98, 2066–2100. DOI: 10.1257/aer.98.5.2066.

Beasley, M. (1996). An Empirical Analysis of the Relation between the Board of Director Composition and Financial Statement Fraud. *Accounting Review* 71, 443–465.

Bebchuk, L.A., Cohen, A., and Ferrell, A. (2009). What Matters in Corporate Governance? *Review of Financial Studies* 22, 783-827. DOI: 10.1093/rfs/hhn099.

Peer-reviewed Academic Journal published by SSBFNET with respect to copyright holders.

www.manaraa.com

Black, D.E., Dikolli, S.S., and Dyreng, S.D. (2014). CEO Pay-for-Complexity and the Risk of Managerial Diversion from Multinational Diversification. *Contemporary Accounting Research* 31, 103-135. DOI: 10.1111/1911-3846.12024.

Boyer, B., Mitton, T., and Vorkink, K. (2010). Expected Idiosyncratic Skewness. *Review of Financial Studies* 23, 169–202. DOI: 10.1093/rfs/hhp041.

Brickley, J., Coles, J., and Terry, R. (1994). Outside directors and the adoption of poison pills. *Journal of Financial Economics* 35, 371–390. DOI: 10.1016/0304-405X(94)90038-8.

Brunnermeier, M., Gollier, C., and Parker, J. (2007). Optimal Beliefs, Asset Prices and the Preference for Skewed Returns. *American Economic Review*, 97,159–65. DOI: 10.1257/aer.97.2.159.

Burgman, T. (1996). An Empirical Examination of Multinational Capital Structure. *Journal of International Business Studies* 27, 553–570. DOI: 10.1057/palgrave.jibs.8490143.

Bushman, R., Chen, Q., Engel, E., and Smith, A. (2004). Financial accounting information, organizational complexity, and corporate governance systems. *Journal of Accounting and Economics* 37, 167-201. DOI: 10.1016/j.jacceco.2003.09.005.

Byrd, J. and Hickman, K. (*1992). Do outside directors monitor managers? Evidence from tender offer bids. Journal of Financial Economics* 32, 195–222. DOI: 10.1016/0304-405X(92)90018-S.

Callen, J.L. and Fang, X. (2015a). Religion and Stock Price Crash Risk. *Journal of Financial and Quantitative Analysis* 50, 69-195. DOI: 10.1017/S0022109015000046.

Callen, J.L. and Fang, X. (2015b). Short interest and stock price crash risk. *Journal of Banking and Finance* 60, 181-194. DOI: 10.1016/j.jbankfin.2015.08.009.

Callen, J.L. and Fang, X. (2013). Institutional investor stability and crash risk: Monitoring vs. short-termism? *Journal of Banking and Finance* 37, 3047–3063. DOI: 10.1016/j.jbankfin.2013.02.018.

Callen, J.L., Hope, O.-K., and Segal, D. (2005). Domestic and Foreign Earnings, Stock Return Variability, and the Impact of Investor Sophistication. *Journal of Accounting Research* 43, 377–412. DOI: 10.1111/j.1475-679x.2005.00175.x.

Carhart, M.M. (1997). On the Persistence in Mutual Fund Performance. *Journal of Finance* 52, 57-82. DOI: 10.1111/j.1540-6261.1997.tb03808.x.

Caves, R. (1971). International corporations: The Industrial Economics of Foreign Investment. *Econometrica* 38, 1-27.

Chen, J., Hong, H., and Stein, J. (2001). Forecasting crashes: trading volume, past returns, and conditional skewness in stock prices. *Journal of Financial Economics* 61, 345–381. DOI: 10.1016/S0304-405X(01)00066-6.

Chi, J.D. and Lee, D.S. (2010). The conditional nature of corporate governance. *Journal of Banking and Finance* 34, 350–361. DOI: 10.1016/j.jbankfin.2009.08.001.

Conrad, J., Dittmar, R.F., and Ghysels, E. (2013). Ex Ante Skewness and Expected Stock Returns. *Journal of Finance* 68, 85-124. DOI: 10.1111/j.1540-6261.2012.01795.x.

Contractor, F.J., Kundu, S.K., and Hsu, C.C. (2003). A Three-Stage Theory of International Expansion: The Link between Multinationality and Performance in the Service Sector. *Journal of International Business Studies* 34, 5–18. DOI: 10.1057/palgrave.jibs.8400003.

Cremers, M. and Ferrell, A. (2014). Thirty Years of Shareholder Rights and Firm Value. *Journal of Finance* 69, 1167–1196. DOI: 10.1111/jofi.12138.

Denis, D.J., Denis, D.K., and Yost, K. (2002). Global Diversification, Industrial Diversification, and Firm Value. *Journal of Finance* 57, 1951-1979. DOI: 10.1111/0022-1082.00485.

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Dimson, E. (1979). Risk measurement when shares are subject to infrequent trading. *Journal of Financial Economics* 7, 197–226. DOI: 10.1016/0304-405X(79)90013-8.

Dittmar, A. and Mahrt-Smith, J. (2007). Corporate governance and the value of cash holdings. *Journal of Financial Economics* 83, 599–634. DOI: 10.1016/j.jfineco.2005.12.006.

Doukas, J.A. and Kan, O.B. (2006). Does Global Diversification Destroy Firm Value? *Journal of International Business Studies* 37, 352–371. DOI: 10.1057/palgrave.jibs.8400203.

Doukas, J.A. and Lang, L.H.P. (2003). Foreign Direct Investment, Diversification and Firm Performance. *Journal of International Business Studies* 34, 153–172. DOI: 10.1057/palgrave.jibs.8400014.

Doukas, J.A. and Pantzalis, C. (2003). Geographic Diversification and Agency Costs of Debt of Multinational Firms. *Journal of Corporate Finance* 9, 59–92. DOI: 10.1016/S0929-1199(01)00056-6.

Duru, A. and Reeb, D. (2002). International Diversification and Analysts' Forecast Accuracy and Bias. *Accounting Review* 77, 415.433. DOI: 10.2308/accr.2002.77.2.415.

Dyreng, S.D., Hanlon, M., and Maydew, E.L. (2008). Long-Run Corporate Tax Avoidance. *Accounting Review* 83, 61–82. DOI: 10.2308/accr.2008.83.1.61.

Elyasiani, E., Jia, J., and Mao, C. (2010). Institutional ownership stability and the cost of debt. *Journal of Financial Markets* 13, 475–500. DOI: 10.1016/j.finmar.2010.05.001.

Errunza, V. and Senbet, L. (1981). The Effects of International Operations on the Market Value of the Firm: Theory and Evidence. *Journal of Finance*, 36, 401–417. DOI: 10.1111/j.1540-6261.1981.tb00455.x.

Errunza, V. and Senbet, L. (1984). International Corporate Diversification, Market Valuation and Size-Adjusted Evidence. *Journal of Finance* 39, 727–745. DOI: 10.1111/j.1540-6261.1984.tb03663.x.

Fama, E.F. and French, K.R. (1993). Common Risk Factors in the Returns on Stocks and Bonds. *Journal of Financial Economics* 33, 3-56. DOI: 10.1016/0304-405X(93)90023-5.

Fama, E.F. and French, K.R. (1997). Industry costs of equity. *Journal of Financial Economics* 43, 153-193. DOI: 10.1016/S0304-405X(96)00896-3.

Gande, A., Schenzler, C., and Senbet, L.W. (2009). Valuation Effects of Global Diversification. *Journal of International Business Studies* 40, 1515–1532. DOI: 10.1057/jibs.2009.59.

Geringer, J.M., Beamish, P.W., and DaCosta, R.C. (1989). Diversification Strategy and Internationalization: Implications for MNE Performance. *Strategic Management Journal* 10, 109–119. DOI: 10.1002/smj.4250100202.

Gilson, S., Healy, P., Noe, C., amd Palepu, K. (2001). Analyst Specialization and Conglomerate Breakups. *Journal of Accounting Research* 39, 565–582. DOI: 10.1111/1475-679X.00028.

Gomes, L.K. and Ramaswamy, K. (1999). An Empirical Examination of the Form of the Relationship between Multinationality and Performance. *Journal of International Business Studies* 30, 173–188. DOI: 10.1057/palgrave.jibs.8490065.

Gompers, P., Ishii J., and Metrick, A. (2003). Corporate Governance and Equity Prices. *Quarterly Journal of Economics* 118, 107-155. DOI: 10.1162/00335530360535162.

Graham, J.R, Harvey, C.R., and Rajgopal, S. (2005). The economic implications of corporate financial reporting. *Journal of Accounting and Economics* 40, 3–73. DOI: 10.1016/j.jacceco.2005.01.002.

Grant, R.M. (1987). Multinationality and Performance among British Manufacturing Companies. *Journal of International Business Studies* 18, 79–89. DOI: 10.1057/palgrave.jibs.8490413.

المنارة للاستشارات
www.manaraa.com

Harford, J., Humphery-Jenner, M., and Powell, R. The sources of value destruction in acquisitions by entrenched managers. *Journal of Financial Economics* 106, 247–261. DOI: 10.1016/j.jfineco.2012.05.016.

Hines J.R. and Rice, E.M. (1994). Fiscal Paradise: Foreign Tax Havens and American Business. *Quarterly Journal of Economics* 109, 149-82. DOI: 10.3386/w3477.

Hitt, M.A., Hoskisson, R.E., and Kim, H. (1997). International Diversification: Effects on Innovation and Firm Performance in Product Diversified Firms. *Academy of Management Journal* 40, 767–798. DOI: 10.2307/256948.

Hutton, A.P., Marcus, A.J., and Tehranian, H. (2009). Opaque financial reports, R^2, and crash risk. *Journal of Financial Economics* 94, 67-86. DOI: 10.1016/j.jfineco.2008.10.003.

Jin, L. and Myers, S.C. (2006). R^2 around the world: New theory and tests. *Journal of Financial Economics* 79, 257-292. DOI: 10.1016/j.jfineco.2004.11.003.

Jiraporn, P., Kim, Y.-S., Davidson, W.N., and Singh, M. (2006). Corporate governance, shareholder rights and firm diversification: An empirical analysis. *Journal of Banking and Finance* 30, 947–963. DOI: 10.1016/j.jbankfin.2005.08.005.

Khurana, I. K., Pereira, R., and Raman, K.K. (2003). Does Analyst Behavior Explain Market Mispricing of Foreign Earnings for U.S. Multinational Firms? *Journal of Accounting, Auditing and Finance* 18, 453–78. DOI: 10.1177/0148558X0301800403.

Kim, J.-B., Li, Y., and Zhang, L. (2011a). Corporate tax avoidance and stock price crash risk: Firm-level analysis. *Journal of Financial Economics* 100, 639–662. DOI: 10.1016/j.jfineco.2010.07.007.

Kim, J.-B., Li, Y., and Zhang, L. (2011b). CFOs vs. CEOs: Equity incentives and crashes. *Journal of Financial Economics* 101, 713–730. DOI: 10.1016/j.jfineco.2011.03.013.

Kim, Y., Li, H., and Li, S. (2014). Corporate social responsibility and stock price crash risk. *Journal of Banking and Finance* 43, 1-13. DOI: 10.1016/j.jbankfin.2014.02.013.

Klein, A. (2002). Audit committee, board of director characteristics, and earnings management. *Journal of Accounting and Economics* 33, 375–400. DOI: 10.1016/S0165-4101(02)00059-9.

Kogut, B. (1983). Foreign Direct Investment as a Sequential Process. *The Multinational Corporation of the 1980s*, edited by C. Kindelberger. Cambridge, MA: MIT Press.

Kogut, B. and Kulatilaka, N. (1994). Operating Flexibility, Global Manufacturing, and the Option Value of a Multinational Network. *Management Science* 40, 123-139. DOI: 10.1287/mnsc.40.1.123.

Kothari, S.P., Shu, S., and Wysocki, P.D. (2009). Do Managers Withhold Bad News? *Journal of Accounting Research* 47, 241–276. DOI: 10.1111/j.1475-679X.2008.00318.x.

Krishnaswami, S. and Subramaniam, V. (1999). Information asymmetry, valuation, and the corporate spin-off decision. *Journal of Financial Economics* 53, 73–112. DOI: 10.1016/S0304-405X(99)00017-3.

Lee, H.-Y., Mande, V., and Son, M. (2008). A Comparison of Reporting Lags of Multinational and Domestic Firms. *Journal of International Financial Management and Accounting* 19, 28-56. DOI: 10.1111/j.1467-646X.2008.01015.x.

Lee, K. and Kwok, C. (1988). Multinational Corporations vs. Domestic Corporations: International Environmental Factors and Determinants of Capital Structure. *Journal of International Business Studies* 19, 195–217. DOI: 10.1057/palgrave.jibs.8490381 .

Lim, S., Matolcsy, Z., and Chow, D. (2007). The Association between Board Composition and Different Types of Unitary Disclosure. *European Accounting Review* 16, 555–583. DOI: 10.1080/09638180701507155.

المنارة للاستشارات
www.manaraa.com

Liu, C.-L. and Lai, S.-M. (2012). Organizational Complexity and Auditor Quality. *Corporate Governance: An International Review* 20, 352–368. DOI: 10.1111/j.1467-8683.2012.00914.x.

Lu, J.W. and Beamish, P.W. (2004). International Diversification and Firm Performance: The S-Curve Hypothesis. *Academy of Management Journal 47*, 598–609. DOI: 10.2307/20159604.

Luo, Y. (2005). Corporate governance and accountability in multinational enterprises: Concepts and agenda. *Journal of International Management* 11, 1-18. DOI: 10.1016/j.intman.2004.11.001.

Masulis, R., Wang, C., and Xie. F. (2007). Corporate Governance and Acquirer Returns*. Journal of Finance* 62, 1851-1889. DOI: 10.1111/j.1540-6261.2007.01259.x.

McCahery, J.A., Sautner, Z., and Starks, L.T. (2016). Behind the Scenes: The Corporate Governance Preferences of Institutional Investors. *Journal of Finance*, forthcoming. DOI: 10.1111/jofi.12393.

Mitchell, W., Shaver, M., and Yeung, B. (1992). Getting There in a Global Industry: Impact on Performance of Changing International Presence. *Strategic Management Journal* 13, 410–432. DOI: 10.1002/smj.4250130603.

Mitton, T. and Vorkink, K. (2007). Equilibrium Underdiversification and the Preference for Skewness. *Review of Financial Studies* 20, 1255–1288. DOI: 10.1093/rfs/hhp041.

Morck, R. and Yeung, B. (1991). Why investors Value Multinationality. *Journal of Business* 64, 165-187. DOI: 10.1086/296532.

Nohria, N. and Ghoshal, S. (1994). Differentiated Fit and Shared Values: Alternatives for Managing Headquarters-Subsidiary Relations. *Strategic Management Journal* 15, 491–502. DOI: 10.1002/smj.4250150606.

Osma, B.G. (2008). Board Independence and Real Earnings Management: The Case of R&D Expenditure. *Corporate Governance: An International Review* 16, 116–131. DOI: 10.1111/j.1467-8683.2008.00672.x.

Pan, J. (2002). The jump-risk premia implicit in options: Evidence from an integrated time-series study. *Journal of Financial Economics* 63, 3–50. DOI: 10.1016/S0304-405X(01)00088-5.

Qian, G. (2002). Multinationality, product diversification, and profitability of emerging us small- and medium-sized enterprises. *Journal of Business Venturing* 17, 611–633. DOI: 10.1016/S0883-9026(01)00080-5.

Rego, S.O. (2003). Tax-Avoidance Activities of U.S. Multinational Corporations. *Contemporary Accounting Research* 20, 805-833. DOI: 10.1506/VANN-B7UB-GMFA-9E6W.

Riahi-Belkaoui, A. and Picur, R.D. (2001). The Impact of Multinationality on the Informativeness of Earnings and Accounting Choices. *Managerial Finance* 27, 82–94. DOI: 10.1108/03074350110767466.

Rugman, A.M. and Verbeke, A. (2004). A Perspective on Regional and Global Strategies of Multinational Enterprises. *Journal of International Business Studies* 35, 3–18. DOI: 10.1057/palgrave.jibs.8400073.

Sanders, W.G. and Carpenter, M.A. (1998). Internationalization and Firm Governance: The Roles of CEO Compensation, Top Team Composition, and Board Structure. *Academy of Management Journal* 42, 158–178. DOI: 10.2307/257100.

Tallman, S. and Li, J. (1996). Effects of International Diversity and Product Diversity on the Performance of Multinational Firms. *Academy of Management Journal* 39, 179–189. DOI: 10.2307/256635.

Thomas, W. (1999). A test of the market's mispricing of domestic and foreign earnings. *Journal of Accounting and Economics* 28, 243-267. DOI: 10.1016/S0165-4101(00)00007-0.

المنارة للاستشارات

www.manaraa.com

Tihanyi, L., Johnson, R.A., Hoskisson, R.E., and Hitt, M.A. (2003). Institutional Ownership Differences and International Diversification: The Effects of Boards of Directors and Technological Opportunity. *Academy of Management Journal* 46, 195–205. DOI: 10.2307/30040614.

Tsao, S.-M., Lu, H.-T., and Keung, E.C. (2016). Internationalization and Auditor Choice. *Journal of International Financial Management and Accounting*, forthcoming. DOI: 10.1111/jifm.12056.

Uzun, H., Szewczyk, S. H., and Varma, R. (2004). Board Composition and Corporate Fraud. *Financial Analysts Journal* 60, 33–43. DOI: 10.2469/faj.v60.n3.2619.

Weisbach, M. (1988). Outside directors and CEO turnover. *Journal of Financial Economics* 20, 421–460. DOI: 10.1016/0304-405X(88)90053-0.

Wiersema, M. and Bowen, H.P. (2011). The Relationship between International Diversification and Firm Performance: Why it remains a puzzle. *Global Strategy Journal* 1, 152-170. DOI: 10.1002/gsj.5.

Xiong, J.X., Idzorek, T.M., and Ibbotson, R.G. (2016). The Economic Value of Forecasting Left-Tail Risk. *Journal of Portfolio Management* 42, 114-123. DOI: 10.3905/jpm.2016.42.3.114.

Xu, N., Li, X., Yuan, Q., and Chan, K.C. (2014). Excess perks and stock price crash risk: Evidence from China. *Journal of Corporate Finance* 25, 419–434. DOI: 10.1016/j.jcorpfin.2014.01.006.

Yan, S. (2011). Jump risk, stock returns, and slope of implied volatility smile. *Journal of Financial Economics* 99: 216–33. DOI: 10.1016/j.jfineco.2010.08.011.

Yuan, R., Sun, J., Cao, F., (2016). Directors' and officers' liability insurance and stock price crash risk. *Journal of Corporate Finance* 37, 173–192. DOI: 10.1016/j.jcorpfin.2015.12.015.

Appendix

Variable	Definition
$CRASH_{t+1}$	Equals one if the firm experienced a stock price crash during fiscal year t+1 and zero otherwise. A stock price crash is defined as a weekly firm-specific log-return that is 3.09 or more standard deviations below the firm's mean weekly firm-specific log-return during the given fiscal year.
$NCSKEW_{t+1}$	Negative one multiplied by the coefficient of skewness of the firm's weekly firm-specific log-returns during fiscal year t+1. The coefficient of skewness equals the sample third central moment divided by the cube of the sample standard deviation.
$DUVOL_{t+1}$	The ratio of the standard deviation of the firm's weekly firm-specific log-returns that are below the firm's mean in fiscal year $t+1$ divided by the standard deviation of the firm's weekly firm-specific log-returns that are above the firm's mean in fiscal year t+1.
MN_t	Equals one if the firm is a multinational firm in fiscal year t and zero otherwise. A firm is defined as multinational if it reports sales by a foreign (non-U.S.) subsidiary.
$SIZE_t$	Natural logarithm of the firm's inflation-adjusted market value of common equity at the close of fiscal year t, measured in millions of 2012 U.S. dollars.
BM_t	Book value of common equity divided by the market value of value of common equity at the close of fiscal year t.
$LEVERAGE_t$	Total debt divided by total assets at the close of fiscal year t.
ROA_t	Income before extraordinary items scaled by the book value of total assets at the close of fiscal year t.
$DTURN_t$	The firm's average monthly share turnover during fiscal year t minus the firm's average monthly share turnover during fiscal year t-1,where monthly share turnover is

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	defined as monthly trading volume (number of shares) divided by the total number of shares outstanding during the month.
$NCSKEW_t$	Negative one multiplied by the coefficient of skewness of the firm's weekly firm-specific log-returns during fiscal year *t*. The coefficient of skewness equals the sample third central moment divided by the cube of the sample standard deviation.
$SIGMA_t$	The standard deviation of the firm's weekly firm-specific log-returns during fiscal year *t*.
$ALPHA_t$	The natural logarithm of one plus the firm's estimated intercept coefficient (alpha) from the five-factor model (see Equation (1)). The five factor model is estimated using all the firm's weekly stock returns during fiscal year *t*.
$LRETR_t$	The firm's long-run effective tax rate, as defined in Dyreng et al. (2008) and Kim, et al. (2011a). Equals total taxes paid during the last five fiscal years (years *t*-4 to *t*) divided by total pre-tax income net of special items during the same period.
$DIS\_ACCRUALS_t$	The sum of the absolute value of discretionary accruals over the last three fiscal years (years *t*, *t*-1, and *t*-2), as defined in Hutton et al. (2009) and Kim et al. (2011a). Discretionary accruals are estimated using the modified Jones model of discretionary accruals of Dechow et al. (1995).
SH_RIGHTS_t	Six minus the firm's entrenchment index as defined by Bebchuk et al. (2009). The entrenchment index is constructed by adding one point for each of the following governance provisions a firm has in place in fiscal year *t*: staggered board, limits to shareholder bylaw amendments, supermajority requirements for mergers, supermajority requirements for charter amendments, poison pill, and golden parachutes.
$BOARD\_IND_t$	Number of independent directors on the board divided by the total number of directors on the board in fiscal year *t*.
$IIO\_STABILITY_t$	For firm *i* in fiscal year *t*, $IIO\_STABILITY_{i,t}$ = max(*StdI*) - $StdI_{i,t}$, where $StdI_{i,t}$ is defined below and max(*StdI*) is the sample maximum of *StdI* across all firms-years in the sample. Following Callen and Fang (2013) $StdI_{i,t}$ is defined as the average standard deviation of quarterly institutional shareholding proportions across all institutional investors in firm *i* over a 5-year period (year *t*-4 to year *t*). Formally, for a given firm *i* in fiscal year *t*, $StdI_i = \sum_{j=1}^{J_i} Std(p_{i,q}^j)/J_i$ where $p_{i,q}^j$ is the proportion of firm *i* held by institutional investor *j* at quarter *q* (*q* = 1, 2, …, 20), $Std(p_{i,q}^j)$ is the time series standard deviation of institutional investor *j*'s quarterly shareholding proportions in firm *i* during the twenty quarters comprising fiscal years *t*-4 to *t*, and J_i is the number of institutional investors in firm *i*. We follow Callen and Fang (2013) and winsorize *StdI* at the sample 1st and 99th percentiles.

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