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The Financial Review 51 (2016) 527–557

The Financial Review

The Effects of Executive, Firm, and Board Characteristics on Executive Exit

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Abstract

We estimate a hazard model of the probability of top corporate executives exiting their firms over the period 1996–2010. Our main findings are that: (1) female executives have greater likelihoods of exit than males, (2) the likelihood of exit increases with the independence of the board and decreases with the fraction of the board that is female and the average age of board members, and (3) a higher percentage of independent directors on the board lowers the probability of exit more for females than for males. Further, controlling for exit risk reduces the well-documented compensation differential between men and women.

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The authors would like to thank seminar participants at the Western Economic Association annual meeting (Seattle, 2013) and Loyola Marymount University, as well as two anonymous referees and the editor for valuable insights. All remaining errors are our own.

Keywords: corporate executives, job turnover, board of directors, firm performance, hazard models

JEL Classifications: G30, G34, J16, J44

1. Introduction

Significant change in the composition of the executive suite has occurred in recent decades, particularly in the length of time executives remain employed in their firms as well as in the gender composition of the profession. Cappelli and Hamori (2005) find that among Fortune 100 firms, the average tenure of executives dropped by more than five years from 1980 to 2001. Also, in 1980 there were no female executives in these firms, but over the subsequent 20-year period, women not only entered the executive ranks, but they also reached the highest ranks faster than men did. These trends suggest that the workforce has become increasingly dynamic and that corporate stakeholders recognize value in diversity. However, the relative scarcity of women in the top ranks attracts media attention when one leaves her position unexpectedly. The forced departure of Carly Fiorina, Hewlett-Packard's chief executive, in 2005 made national headlines. More recently, the sudden departure of Jill Abramson as executive editor of the *New York Times* was discussed in a *Wall Street Journal* article, as she was the second top female to leave the firm suddenly in fewer than three years (Trachtenberg and Marr, 2014).

In this paper, we use a data set compiled from multiple sources to examine the factors that determine the tenure of top executive officers with their firms as well as how those factors differ by gender. While previous studies have compared the reasons for exit between male and female executives (Becker-Blease, Elkinawy and Stater, 2010), we focus on how long executives remain employed. Executive tenure is an issue for corporate governance due to the potential relation between tenure and shareholder wealth, and for public policy as the longevity of executives is related to wealth acquisition and influence. While long tenure can promote stability and allow a corporation to strengthen its market position as the executive gains experience, excessive tenure can also lead to entrenchment and to overly risk-averse behavior that is detrimental to firm survival and performance. Indeed, McClelland, Barker and Oh (2012) find that chief executive officer (CEO) paradigms become increasingly obsolete as tenure increases, albeit only in dynamic industries where the operating environment undergoes frequent change. Gender differences in executive tenure may also reflect imperfections in the labor market.

While the primary motivation for our study is to understand tenure, approaching the problem with a survival analysis model allows us to readily address censoring problems that are posed by the data. Nonetheless, as there is a tight theoretical connection between the risk of exit from a state or condition in any given time period and the total amount of time spent there, our approach of studying the hazard of executive exit from large corporations amounts to modeling the distribution of executive tenures within those firms.

Our results suggest that female executives have higher probabilities of exit from their firms than do male executives, all else equal. We also find that age has a positive effect and that compensation has a positive but diminishing effect on the hazard of exit. CEOs have lower probabilities of exit than do lower-ranked officers, and directors have higher probabilities of exit than do nondirectors. The size of the board and the average age of board members reduce the likelihood of an executive's exit, as does the fraction of the board that is female, whereas a higher fraction of independent directors increases the likelihood of exit. Firm financial performance, accounting performance, and firm value also affect the hazard of exit.

We further explore gender differences in exit risk and find that the relation is affected by certain board and firm attributes. Holding a director position lowers the risk of exit more for women than for men, as do increases in board independence, industry-adjusted return on assets, and industry-adjusted stock returns. Finally, we examine the well-documented compensation gap between men and women, and find it is substantially reduced, if not eliminated, by controlling for the risk that the executive will exit the firm. Overall, these results are consistent with the notion that the hazard of executive exit not only responds to personal attributes, the opportunity costs of leaving top managerial positions, and aspects of corporate governance and firm performance, but that it can also have a role in shaping employment outcomes for top executives.

2. Literature review

A number of studies focus on drawing implications of corporate executive job tenure for various firm outcomes. Musteen, Barker and Baeten (2006) find that the CEOs of nonprofit organizations become more conservative as their tenure increases. Walters, Kroll and Wright (2007) report a nonlinear effect of CEO tenure on shareholder returns surrounding acquisition announcements, suggesting fewer agency problems in the early years of CEO tenure followed by increased entrenchment in later years. Masters-Stout, Costigan and Lovata (2008) examine the relation between CEO tenure and goodwill impairment decisions, which give the CEO an opportunity to potentially manage earnings. They find that new CEOs impair more goodwill than do their senior counterparts, suggesting that differences in tenure have different corporate governance effects associated with U.S. financial accounting rules. Graham, Harvey and Puri (2015) survey more than 1,000 CEOs and chief financial officers (CFOs) and find that the delegation of decision-making authority decreases with CEO tenure. Their findings suggest that as CEOs become more knowledgeable, they are less likely to share the capital structure decision with others. Gong (2011) examines CEOs who were in office between 1992 and 2007, and finds that both CEO compensation and shareholder value aggregate naturally over CEO tenure.

While executive tenure is a common control variable in corporate governance research, a review of the literature reveals relatively few papers focusing on the determinants of executive tenure, particularly in the economics and finance literature. One of the earliest studies on this issue is Salancik and Pfeffer (1980), who examine the determinants of CEO tenure in 84 U.S. corporations in 1972. Not surprisingly, they find that firm performance positively affects tenure. They also find that the concentration of stock ownership is a significant determinant of tenure. However, they do not find an association between the proportion of inside directors and the years spent as chief executive. More recently, Wang, Davidson and Wang (2010) find that CEO tenure was not significantly shortened with the passage of the Sarbanes-Oxley Act in 2002, a finding they attribute to increased risk aversion by CEOs due to the increased monitoring level associated with the new legislation. Brookman and Thistle (2009) use survival analysis to examine CEO tenure from 1993 through 2001 and find that the risk of termination increases until the CEO has been in office for 13 years and then falls. Current and prior year stock returns negatively affect the hazard of termination but these effects diminish over time. They also find that greater CEO compensation, stock ownership, and age decrease the hazard while greater firm size, greater use of leverage, and more frequent board meetings increase the hazard. Finally, Paul and Sahni (2010) examine the types of firms that females choose to work for from 1994 to 2002 and find that female non-CEO executives have lower tenure in office than all other executives. However, their tenure model does not include certain governance characteristics, such as board composition, that may influence the relation.

Most of the aforementioned studies focus on the tenure or risk of termination of the chief executive. However, important decision making also occurs at the lower ranks such as CFO and vice presidents, and thus an investigation of the factors that affect the tenure of non-CEO executives would enrich our understanding of corporate governance. In addition, none of these studies thoroughly explores gender differences in the effects of executive and firm attributes on tenure or the risk of exit from an executive position.

3. Data and descriptive statistics

3.1. Data sources and variables

We obtain our data from *ExecuComp*, *Riskmetrics* (formerly the *Investor Responsibility Resource Center* or *IRRC*), COMPUSTAT, and CRSP. *ExecuComp* provides data on (roughly) the top five executives in Standard & Poor's 1500 firms for the years 1992–2010 although due to limitations on the board data available from *Riskmetrics* we use only data from 1996 and beyond. *ExecuComp* reports each executive's salary, total direct compensation (TDC; which includes salary, bonuses, the total value of restricted stock granted, the total value of stock options granted, long-term incentive payouts, and all other total annual compensation), gender, job title, and tenure with the firm, which is defined as the number of years credited toward retirement.[1]

Riskmetrics provides data on the board of directors, including board size, the numbers of male and female directors, and the number of independent board members. COMPUSTAT provides accounting and financial data, and CRSP provides stock price information at a variety of frequencies. The *Riskmetrics*, COMPUSTAT, and CRSP data are then merged with the *ExecuComp* data by firm and year.

Underlying any survival process is a random variable that measures the duration of the spell. The underlying duration variable in our analysis is the tenure of executives with their firms. We use two main variables in *ExecuComp* to determine the elapsed length of service of an executive as of any given sample year. First, the "Joined_Co" variable records the date the executive joined the company. Second, the variable "Ret_Yrs" records the number of years credited toward retirement according to the company's retirement plan. Of course, executives could be awarded years of service in the pension plan as a perk without having actually served those years, and some pension plans allow a maximum of 30 credited years. Unfortunately, the data are not complete, and information is missing for some executives. In general, when measuring elapsed tenure with the firm, we place more weight on the date the executive joined the company as opposed to the number of credited years toward retirement (in cases where the two are in conflict). If no other information is available, we simply use credited retirement years. However, in cases of notable conflict between the two variables, clear errors in the data, or questionable observations, we use electronic filings and news searches to determine the length of service. The final data set includes a panel of 8,707 executives and 46,046 executive-year pairs.

We use the following guidelines to determine tenure with the firm, which is intended to measure the number of full years an executive worked at the company prior to the year of observation as recorded in *ExecuComp*. Service that began before September 1 of a given year is counted as a full year in order to make the observations we fill in reasonably consistent with those already appearing in the database where, in many cases, executives appear to have received full years for partial years of service. We treat subsidiaries of a parent company as the same firm as the parent, so if the executive transferred from the subsidiary to the parent or vice versa, that is counted as continued service with the same company.[2] We count executives of companies whose

[1] Executives could well have begun their careers with their firms in lower-level positions than we observe. If they never ascend to a top-level executive position, we never observe them in our data set. Thus, our paper does not address the determinants of job tenure for corporate employees in general. Our analysis instead focuses on the probability that an executive remains with the firm an additional year, given that they have survived with the firm long enough or been successful enough to obtain a position sufficiently high to be recorded in *ExecuComp*.

[2] Since *ExecuComp* lists executives with two positions at distinct companies (e.g., at a parent and at a subsidiary) twice, we treat executives with two listed positions in *ExecuComp* as having two distinct positions, except that the tenure of the executive is treated as continuous over the two firms.

firm are bought or acquired by some other company as starting a new position with the acquirer on the transaction completion date. If *ExecuComp* has no information about the retirement years of executives or the date they joined the company, we omit those executives from the sample.

The covariates we use to explain the probability that an executive exits the firm in a given year (in the sense that they are reported by *ExecuComp* as being an officer of the firm in one year but not the next)[3] consist of executive-level controls, board controls, firm financial and stock return controls, industry controls, and year controls. Our executive-level controls obtained from *ExecuComp* include dummies for female, CEO, and director; age and age squared; and TDC (which includes all forms of compensation including salary, bonuses, options, deferred compensation, etc.) and its square. In some (unreported) specifications, we replace TDC with salary and discuss the few differences that arise. *ExecuComp* also includes industry identifiers that we use to group firms into the 48 Fama-French (1997) industries. We also construct a dummy for whether the firm currently has a female CEO (other than the executive herself) based on the information available in *ExecuComp*.

The board of director controls obtained from *Riskmetrics* include the number of directors on the board other than the executive; the percentage of other directors on the board who are female; the percentage of other directors who are independent; and the average age of the other board members.[4] We also include controls for firm financial and accounting performance that are available from COMPUSTAT and CRSP. From COMPUSTAT, we obtain the firm's return on assets (net income divided by total assets) in each sample year, which we then industry-adjust using the 48 Fama-French categories;[5] the firm's book-to-market ratio (book value per share divided by the stock's closing price per share at the end of the fiscal year); and the firm's total assets in each year. We measure firm value as Tobin's q, calculated as (market cap –

[3] We rely on *ExecuComp* to record accurately when executives exit the firm, which we interpret as when *ExecuComp* no longer lists the executive with the firm. Executives who disappear and then reappear in the data set in the same position are treated as missing for the year when they are not observed, but their spell continues unbroken for observations after the missing year. After stepping down as an executive, some executives might remain with the firm for a brief period in a nonexecutive capacity. While the disappearance from *ExecuComp* might not represent the correct year an executive actually left the firm, in the overwhelming majority of the cases we looked at more closely, any discrepancy is at most one year. In any case, we can interpret the hazard as the probability of leaving an executive position with the firm if not necessarily the firm itself.

[4] We exclude the executive in question from calculations of female CEO, total number of board members, percentage of board members who are female, percentage who are independent, and average age of board members in order to avoid potential endogeneity concerns. If an executive is the only member of the board, the number of other board members and the percent of other directors who are female or independent are set to 0, and the average age of the other board members is taken to be the age of the sole executive board member.

[5] The industry-adjusted value of a variable (return on assets, stock returns, or Tobin's q) is the residual from an ordinary least squares (OLS) regression of the variable on 47 of the 48 Fama-French industry categories and 14 of the 15 year dummies.

stockholders' equity + book value of assets)/book value of assets, where for "book value of assets" we simply use "total assets." We also construct an alternative measure of Tobin's q as (market cap + short-term and long-term debt)/book value of assets, with very similar results. We industry-adjust both Tobin's q measures. From CRSP we obtain the firm's annual stock return (the percentage change in the common stock's closing price from the end of the previous to the end of the current fiscal year), which we also industry-adjust. We control for time by including a set of year dummies for the years 1996–2010. (We actually have to exclude the dummies for both 2009 and 2010 due to multicollinearity.) Table 1 provides a description of all of the variables used in the analysis.

3.2. Censoring problems

The construction of the data leads to two issues with regard to censoring. First, since we begin the sample in 1996 due to data limitations, we never observe those who started and ended their tenures prior to 1996. It follows that we oversample executives who have survived a relatively long time in their firms. Moreover, those who are observed in the sample in the initial year of 1996 are effectively entering the sample at a later date than when they started working with their firms. This is known as "stock sampling," "left truncation," or "delayed entry" in the survival analysis literature. Second, since our data set ends in 2010, some executives who are observed in 2010 will have employment periods that would continue beyond the end of the data set if such data were available. In other words, their employment spells are unfinished, and their true employment durations are unknown. This is known as a "right censoring" problem. Fortunately, survival analysis methods are well suited to addressing both left truncation and right censoring. This is why, even though the primary motivation for our paper is understanding the factors affecting job tenure, we examine this process in a survival analysis or hazard model framework.

3.3. Descriptive statistics

Descriptive statistics for the variables in our sample are provided in Table 2. The means indicate that 5.8% of the executives in our sample are females. We also see that 23.5% are CEOs, so the remaining 76.5% are other "chief" officers or vice presidents of various types. The average age of the executives in our sample is 52.0 and these executives earn an average base salary of $464,000. However, the average TDC is $3.1 million, reflecting the large role of nonsalary compensation for top executives. The average number of directors on the board (excluding the executive in question) is 9.3; about 10.2% of the other directors are female, and about 50% are independent. The average return on assets among the firms in the sample is 4.0%, the average annual stock return is about 7.2%, and the average value of Tobin's q is 1.97.

Table 1

Variable descriptions

This table defines the variables in the data.

Variable	Definition	Source
Female	Dummy variable for females.	E
CEO	Dummy variable for CEOs.	E
Director	Dummy variable for directors.	E
Age	Executive's age in years.	E, AI
Salary	Executive's base salary in thousands of dollars.	E
Total direct compensation	Executives total compensation in thousands of dollars (includes salary, bonuses, total value of restricted stock granted, net value of stock options exercised, long-term incentive payouts, and the total of all other forms of compensation).	E
CEO is female	Dummy variable for whether the CEO of the firm is female (other than the executive herself; so this equals 0 for female CEOs as well as for officers of firms without female CEOs).	E
Number of other directors	Number of directors on the board (other than the executive herself), so this value is the number of directors minus 1 for executives who are also directors; for those who are not directors it is just the total number of directors.	R
Pct. of other directors female	Percentage of directors on the board (other than the executive herself) who are female, so for female directors this value is the number of other females on the board divided by number of directors minus 1; for male directors it is the number of females on the board divided by the number of directors minus 1; for nondirectors it is the percent of the full board that is female.	R
Pct. of other directors independent	Percentage of directors on the board (other than the executive herself) who are independent, where independence is defined as not being a current or former employee of the firm, not being a relative of a current or former employee, and not reporting a financial association with or interest in the firm. This variable equals the number of independent directors divided by the number of directors of the firm other than the executive.	R
Avg. age of other directors	Average age of the directors on the board (other than the executive herself). For nondirectors this value is the average age of the full board. If the executive is the only member of the board, this value is set equal to the age of the executive alone.	R
Industry-adjusted return on assets	Return on assets (ROA) = net income/total assets. Industry-adjusted value is the residual from an OLS regression of ROA on a full set of Fama-French industry dummies and a full set of year dummies.	CO
Book-to-market ratio	Book value per share/closing price per share at end of fiscal year.	CO

(*Continued*)

Table 1 (*Continued*)

Variable descriptions

Variable	Definition	Source
Total assets	Total assets of the firm in millions of dollars.	CO
Industry-adjusted annual stock return	Annual stock return = percentage change in closing price from end of previous fiscal year to end of current fiscal year. Industry-adjusted value is the residual from an OLS regression of annual stock return on a full set of Fama-French industry dummies and a full set of year dummies.	CR
Industry-adjusted Tobin's q	Tobin's q = (market capitalization – stockholders' equity + total assets)/total assets. Stockholders' equity represents book value of equity and total assets represents book value of assets. Industry-adjusted value is the residual from an OLS regression of Tobin's q on a full set of Fama-French industry dummies and a full set of year dummies.	CO
Industry-adjusted alternative Tobin's q	Alternative Tobin's q = (market capitalization + short-term debt + long-term debt)/total assets. Total assets are used to represent the book value of assets. Industry-adjusted value is the residual from an OLS regression of the alternative Tobin's q on a full set of Fama-French industry dummies and a full set of year dummies.	CO
Fama-French industry categories	Dummy for Fama-French (1997) industry codes 1–48.	F
Year 1996–Year 2010	Dummies for years 1996 through 2010, respectively.	E

E, *ExecuComp*; R, *Riskmetrics*; CO, COMPUSTAT; CR, CRSP; AI, authors' Internet searches; F, Fama and French (1997).

4. Empirical methods and hypotheses

4.1. Survival analysis methods

A natural approach to modeling variables that measure duration is survival analysis, a method in which one predicts the conditional probability that the time spent in some state of interest such as good health, employment, etc. (known as a "spell") ends, given that it has lasted to the present time, rather than predicting the length of the spell directly. In our data, a spell consists of an executive's tenure with a particular firm, and the last year in which the executive is reported to be with that firm in *ExecuComp* is interpreted as the end of the employment spell.

The conditional probability of a spell ending in the current time period, given that it has lasted until the current period, is referred to as the "hazard" of ending the

Table 2

Summary statistics

This table presents summary statistics about the data.

Variable	Mean	SD	Min	Max
Female	0.058	0.234	0	1
CEO	0.235	0.424	0	1
Director	0.395	0.489	0	1
Age	52.0	7.7	20	94
Salary ($ thousands)	464	314	0	8,100
Total direct compensation ($ thousands)	3,108	12,233	0	1,589,839
CEO is female (other than the exec.)	0.013	0.112	0	1
Number of other directors	9.3	2.9	0	34
Proportion of other directors female	0.102	0.097	0	1
Proportion of other directors independent	0.502	0.267	0	1
Average age of other directors	60.2	4.2	33	86
Return on assets (proportion)	0.040	0.142	−5.78	0.872
Book-to-market ratio	0.495	0.616	−44.1	12.1
Total Assets ($ millions)	16,285	74,998	16.77	2,223,299
Annual stock return (proportion)	0.072	0.723	−0.999	46.73
Tobin's q	1.971	1.708	0.369	78.56
Alternative Tobin's q (only 45,920 observations)	1.622	1.743	0.043	78.42
Year 1996	0.044	0.205	0	1
Year 1997	0.054	0.225	0	1
Year 1998	0.060	0.238	0	1
Year 1999	0.063	0.244	0	1
Year 2000	0.068	0.252	0	1
Year 2001	0.076	0.264	0	1
Year 2002	0.075	0.264	0	1
Year 2003	0.080	0.272	0	1
Year 2004	0.079	0.270	0	1
Year 2005	0.073	0.261	0	1
Year 2006	0.071	0.257	0	1
Year 2007	0.074	0.261	0	1
Year 2008	0.071	0.257	0	1
Year 2009	0.063	0.243	0	1
Year 2010	0.049	0.215	0	1
Number of observations			46,046	

Variable definitions provided in Table 1.

spell. We estimate the hazard using a Cox proportional hazards (PHs) model, which is specified as follows:[6]

$$\lambda(t) = \lambda_0(t) \exp(\boldsymbol{x}(t)\boldsymbol{\beta}), \tag{1}$$

[6] The terminology "proportional hazards" comes from the fact that, if the covariates are fixed over time (i.e., $x(t) = x$ for all t), then different values of x act multiplicatively to shift the baseline hazard function, so at any given time the values of the hazard function at two values of x are proportional to each other. As a result, at any two values of x the paths of the log hazard function over time are parallel.

where $\lambda(t)$ is the hazard rate at time t, $\lambda_0(t)$ is the baseline hazard function (which governs how the hazard evolves over time when all covariate values are set to some baseline value), $\boldsymbol{x}(t)$ is a vector of (possibly time-varying) covariates, and $\boldsymbol{\beta}$ is a vector of coefficients to estimate.[7]

Taking the logs of both sides of Equation (1) allows us to interpret the coefficients $\boldsymbol{\beta}$:

$$\log \lambda\,(t) = \log \lambda_0\,(t) + \boldsymbol{x}(t)\,\boldsymbol{\beta}. \tag{2}$$

Equation (2) shows that the log hazard rate is linear in the parameters, so that the coefficient β_k represents the proportional change in the hazard rate $\lambda(t)$ due to a one-unit change in x_k. In cases where we interact the female dummy D with various xs, the coefficient on the interaction term Dx_k represents the difference in the effect of x_k on the log hazard for women versus men. If the coefficient is positive (negative), it means that an increase in x_k increases the log hazard, and hence the hazard, more (less) for women than for men. Thus the signs of the coefficients on the interaction terms bear a similar interpretation as they would in a linear regression model.

Assuming a random sample of completed spells of length $t_1 < t_2 < \ldots < t_n$ taken from the population, with the possibility of censoring ignored for the moment, the log-likelihood for the Cox PHs model is:

$$\ln L = \sum_{i=1}^{n} \boldsymbol{x}_i(t_i)\,\boldsymbol{\beta} - \sum_{i=1}^{n} \ln\left\{\sum_{j\in R(t_i)} \exp\left(\boldsymbol{x}_j(t_i)\,\boldsymbol{\beta}\right)\right\}, \tag{3}$$

where $\boldsymbol{x}_i(t_i)$ is the vector of covariate values for person i (the person who fails at time t_i) measured at time t_i, $R(t_i)$ is the "risk set" at time t_i, which is the set of individuals who have not yet failed as of time t_i, and $\boldsymbol{x}_j(t_i)$ is the vector of covariate values for person j in the risk set at time t_i. One then estimates the parameter vector $\boldsymbol{\beta}$ by maximizing ln L.[8] Since the baseline hazard function $\lambda_0(t)$ does not appear in the above expression for $\ln L$, it is allowed to take on an arbitrary time path in this approach. This feature is one of the advantages of the PHs model and why it is sometimes referred to as a "semi-parametric" approach.

Hazard models easily handle both of the censoring problems encountered in our data. In the case of right censoring, the log likelihood is of the same form as above,

[7] Time-varying covariates can be accommodated in hazard models by splitting up the data on each individual into intervals over which the covariates are fixed. This accommodation requires effectively treating each time interval for each individual as a separate observation. Thus with time-varying covariates the standard errors are clustered by person in order to account for the fact that realizations of variables in different time intervals for the same person are not independent.

[8] Although we assume a sample of distinct failure times in this description, the methods can be readily extended to incorporate tied failure times by adjusting the likelihood for the fact that multiple individuals can fail from the risk set at a given failure time. Two commonly used techniques for dealing with tied failure times are the Breslow and Efron approximations. We use the Efron approximation because it performs better in cases with large numbers of ties, as often happens when duration data are grouped into discrete intervals.

except the risk set at time t_i now includes those who have neither failed nor been censored as of t_i. Under left truncation (or delayed entry), the log-likelihood for the Cox model again has the above form, except that subjects who are left truncated are not included in risk sets prior to the time at which they enter the sample. (If left truncation were ignored, then individuals who enter late would be included in all risk sets prior to the time at which they fail or are censored. Including subjects in all risk sets is not possible in our data set because data on the subjects simply do not exist prior to when they first come under observation, which means they cannot be included in those risk sets.)[9]

It is important to emphasize that these methods yield parameter estimators that are consistent and asymptotically normal, with a known variance-covariance matrix, only if the censoring and truncation processes are independent. Independent censoring means that those who are censored have similar hazard rates, given their covariate values, as those who are not censored. Independent censoring seems plausible in this case since the reason that executives are censored is simply because our data set ends in 2010. Independent truncation means that those who are late entrants into the sample have similar hazard rates, given their covariates, as those who have been in the sample since they were first at risk of failure. For executives who are first observed relatively few years after they were initially hired as top executives, independent truncation also seems plausible, but the assumption may be more problematic for executives with employment durations that are already very long as of 1996 since there will be no corresponding nontruncated members of the sample with similar durations. Thus, our model may not accurately capture the factors contributing to extremely long employment durations.

The above model assumes that the duration variable T is in continuous time. In our data, however, the only possible duration values are discrete since we only have what amounts to one observation per year for each executive, and they are observed to either exit or not in that year. In the technical jargon of survival analysis, our data are called "grouped" duration data as the discreteness of the data is not intrinsic but due to the aggregation of the data to a yearly frequency.[10]

4.2. Hypotheses

All else equal, we expect female executives to have higher hazards of exit from top executive positions than do male executives due to the historic underrepresentation of females in the corporate sector, possible imperfections in the executive labor

[9] Further, we treat the data as single spell format although we allow for clustering of standard errors by executive.

[10] Discrete time hazard models can also be estimated as logit or conditional logit models, while also accounting for right censoring and truncation, but the results turn out very similar whether we estimate the model in continuous or discrete time. Thus, for ease of exposition, we estimate our models as if the data were in continuous time.

market, and voluntary departures due to family and/or other household responsibilities. CEOs should have lower hazards of exit than lower-ranked officers due to higher opportunity costs of exit, the fact that they must demonstrate a high level of aptitude or knowledge of the firm or industry in order to secure such positions, and the greater power they can exert within the firm to repel attempts to remove them. For similar reasons, directors should have lower hazards of exit than nondirectors.

We expect age to have a nonlinear effect on the hazard: for low ages, the probability of exit should decrease as executives become more settled in their careers, but this effect should diminish over time and eventually become positive as the executive enters the later stages of the career cycle toward retirement. Salary should likewise have a nonlinear effect: for low salaries, an increase in salary should lower the probability that the executive leaves the position due to a higher opportunity cost of exit, but as the salary becomes large, the risk of exit should start to increase because the executive can afford to retire, has a high enough profile to seek other lucrative opportunities, or is more likely to face termination by a cost-conscious board.

The risk of exit is expected to be lower in large firms because size is typically associated with greater visibility, and large companies generally offer more opportunities for advancement and the development of managerial skills. Industry-adjusted return on assets, stock returns, and firm value (measured by Tobin's q) are expected to reduce the hazard of exit because better-performing firms should be more attractive places to work and may also encourage the board to retain current executives. However, better-performing firms could also make an executive more attractive to other firms, resulting in higher exit probabilities. Studies that examine the determinants of executive turnover generally find that executives are more likely to be terminated following poor firm performance, particularly when the board is more independent or more gender diverse (e.g., Weisbach, 1988; Huson, Malatesta and Parrino, 2004; Adams and Ferreira, 2009; although for the importance of industry performance, see Jenter and Kanaan 2015). All else equal, then, the risk of executive exit should be negatively related to firm performance and value. Book-to-market ratio is expected to have an ambiguous effect on the hazard since firms with high book value in relation to market value are generally more mature firms likely to attract executives who value longevity. On the other hand, these firms also tend to have fewer growth opportunities, which could lower the opportunity cost of leaving.

We expect that a greater number of board members will reduce the hazard of exit because coordination problems on larger boards may make it more difficult to remove executives. However, more independent boards should find it easier to remove executives, so that the hazard of exit should be higher when the board consists of a greater proportion of outsiders. A higher average age of board members should reduce the hazard because older boards may have more established relationships with the firm's executives and be more reluctant to remove them.

This study is especially interested in identifying firm and board characteristics that could lead to differences in the risk of exit between men and women. In general, we hypothesize that the factors that increase female representation in executive positions are the same factors likely to reduce the hazard of women executives leaving their positions. Carter, Simkins and Simpson (2003) find that the proportion of women and minorities on the board decreases as the number of insiders increases. In addition, Skaggs, Stainback and Duncan (2012) find that the proportion of female directors is positively associated with the percentage of females at the management level. To the extent that the presence of women at the top of the management hierarchy, greater gender diversity on the board, and more independence on the board are perceived as signals of more welcoming and equitable work environments by female executives, we expect that women are more likely to remain employed in firms (and hence have lower exit rates) where either the CEO is female, the board has better female representation, or the board is more independent.

The incentives of the board to more closely monitor the firm may also imply different effects by gender for various board characteristics. Larger boards with their weaker ability to coordinate their decisions might more often fail to see or reward the potential in talented female executives, resulting in a higher hazard for women relative to men as the board becomes larger. We also expect that the exit probability for females will be relatively greater with an aging board because older boards consist mostly of individuals from a generation where women were less represented in the executive ranks.

We also investigate whether directors and CEOs have different hazard rates by gender. If female directors or CEOs are disproportionately targeted for termination by male-dominated boards or shareholder groups, then we would expect female directors and CEOs to have higher hazards of exit than their male counterparts. Bertrand and Hallock (2001) show that smaller firms tend to have more female representation in the top management positions. However, Skaggs, Stainback and Duncan (2012) find greater female managerial representation in larger establishments. In addition, larger firms have larger boards, and Carter, Simkins and Simpson (2003) find that board diversity increases with board size. Due to these mixed results, it is unclear whether the hazard of exit for females is different from that of male executives as the firm grows in size. Better-performing firms, on the other hand, may be so because they are better able to attract and retain talented executives regardless of gender. If so, then firms with higher industry-adjusted return on assets, stock returns, and firm value might be expected to produce lower hazard rates for female relative to male executives. A similar remark could apply to book-to-market although higher book-to-market may indicate older, better-established firms with more entrenched "old boy" networks. Table 3 records the hypotheses associated with each variable.

5. Empirical results

5.1. Proportional hazards diagnostics

We begin by examining the validity of the PH assumption with the help of several diagnostics. Although technically speaking, hazards are no longer proportional with

Table 3

Hypotheses

This table lists the expected sign of the coefficient of each variable when predicting the hazard of exit.

Variable	Hypothesized effect on hazard
Female	+
CEO	–
Director	–
Age	Nonlinear: − and then +
Salary/compensation	Nonlinear: − and then +
CEO is female	?
Number of other directors	–
Pct. of other directors female	?
Pct. of other directors independent	+
Average age of other directors	–
Return on assets	–
Book-to-market ratio	?
Total assets	–
Stock return	–
Tobin's q	–
Female*CEO	+
Female*Director	+
Female*Age	?
Female*Salary/compensation	?
Female*CEO is female	–
Female*Number of other directors	+
Female*Pct. of other directors female	–
Female*Pct. of other directors independent	–
Female*Average age of other directors	+
Female*Return on assets	–
Female*Book-to-market ratio	?
Female*Total assets	?
Female*Stock return	–
Female*Tobin's q	–

time-varying covariates in the model, which means the model is more flexible in this case than the PH moniker suggests, we can check whether the assumption holds for the time-invariant covariates, of which there are two: gender and industry. Thus, in Figure 1 we plot the log Kaplan-Meier (KM) estimate of the survivor function against the log analysis time (i.e., the time spent at risk of the failure event) by gender and industry.[11] Under the PH assumption, the curves for men and women should

[11] Let T be the duration variable that governs survival times. Then the KM estimator is a nonparametric estimator of the survivor function $S(t) = \Pr[T \geq t]$ defined by $\hat{S}(t) = \prod(n_j - d_j)/n_j$, where the product is taken over all failure times t_j that precede t, n_j is the number of individuals at risk at time t_j, and d_j is the number of failures at time t_j.

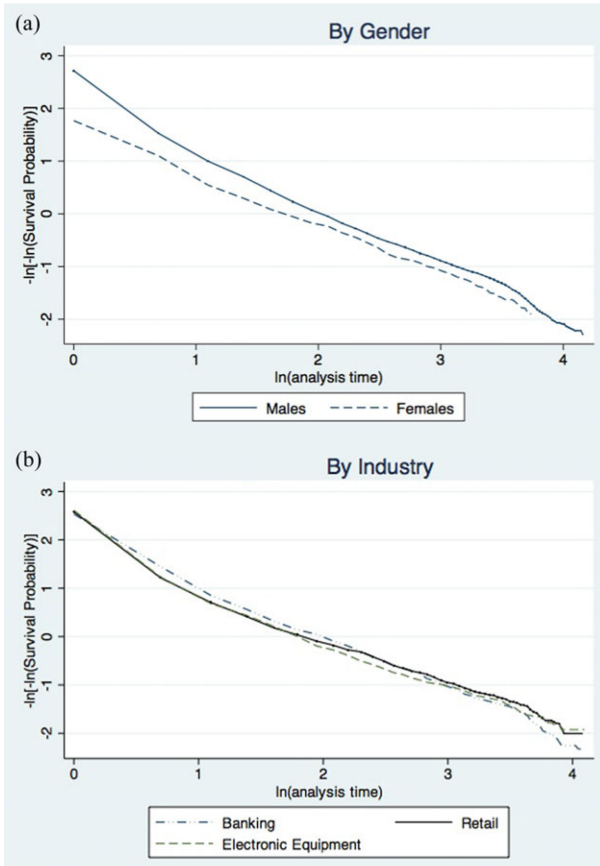

Figure 1

Proportional hazards plots

This figure graphs the log-log of survival probabilities (nonparametric Kaplan-Meier estimates) against the log of analysis time. Panel (a) compares survival probabilities across genders while Panel (b) compares survival probabilities across three of the more popular industries in the sample.

be parallel, as should the curves for any pair of industry groups. We see that the graphs for males and females are roughly parallel except possibly at early analysis times (Figure 1a), suggesting the PH assumption is reasonable with respect to gender. Because of the large number of industry groups, the graph for industries (Figure 1b) shows only three of the top industries in our sample (banking, electronic equipment, and retail) but even these show clear deviations from parallel. Thus it appears that the PH assumption is not valid for industry.

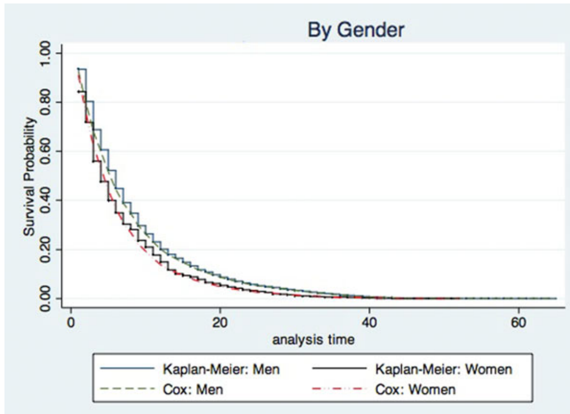

Figure 2

Kaplan-Meier and Cox survivor function estimates

This figure compares Kaplan-Meier and Cox survivor function estimates for both men and women. Cox survivor function estimates are obtained from a Cox proportional hazards model with gender as the only explanatory variable.

In cases in which the PH assumption is not met, an alternative method is to estimate a stratified model, which means the baseline hazard function is allowed to differ across the strata of the variable(s) for which the assumption is violated. In Figure 2, we graph the KM estimate and the survivor function estimate predicted from the Cox model against analysis time for men versus women. If the PH assumption is met, the Cox estimates should be close to the KM estimates for each group since the KM estimator does not impose the PH assumption. For both men and women this appears to be the case through all analysis times.

The final PH diagnostics we consider are the Schoenfeld residuals. For the variable x_k and the observation j failing at time t_j, the Schoenfeld residual is essentially the difference between the value of covariate x_k for person j and the weighted average of the covariate values for all other individuals in the risk set R_j, where the weights are the estimated relative hazards from the Cox model. If, contrary to the PH assumption, the coefficient β_k varies over time, then the Schoenfeld residual r_{jk} can be written in the form $r_{jk} = \alpha_k h(t) + \varepsilon$ where ε is a zero mean error term. Thus, a test of the PH assumption can be conducted by estimating a linear regression of the Schoenfeld residuals on $h(t)$, which is often just t, and then testing the hypothesis that the coefficient is zero. When we conduct this test for gender, the estimated coefficient on time has a p-value of 0.26, so we again fail to reject the PH assumption. In contrast, when we test the Schoenfeld residuals for industries, eight of the 48 Fama-French industry dummies have coefficients that are significant at the 0.10 level or better. In view of these diagnostic results, our modeling approach will be to impose the PH assumption for gender but to stratify the model by industry.

5.2. Regression results: Baseline specifications

We now discuss the estimates of the Cox PH model, which are presented in Table 4. Model 1 provides the results with a simple set of executive controls, including gender, rank, quadratics in age and TDC, and the director dummy. The estimates imply that females have a hazard of exit that is 29.7% greater than that of males, controlling for other personal attributes. Of course, this result does not explain why women have a higher exit risk, and several possible explanations exist, including labor market imperfections on one hand, and voluntary career interruptions, delays, or truncations due to family and household responsibilities on the other. Unfortunately, our data do not give enough details on reasons for exit to distinguish between these possibilities.

The remaining estimates in Model 1 indicate that, not surprisingly, CEOs have much lower probabilities of exit in a given year than other "chief" officers or vice presidents.[12] The coefficients on the age terms suggest only a linear increase in the hazard of exit as the executive ages, as the quadratic term is statistically insignificant. Executives with more life experience have a greater risk of separation from the firm, and presumably much of this effect is related to voluntary succession at retirement.

TDC also has a nonlinear effect on the hazard. The coefficients imply that the effect of TDC increases up to a maximum at about $55 million and decreases thereafter. Since $55 million is above the 99th percentile value of total compensation, practically speaking we find that the hazard of exit increases at a decreasing rate over the sample range of compensation values. That the hazard generally increases with TDC suggests that increases in total compensation reflect the expansion of an executive's outside opportunities in the market or that firms are more motivated to remove executives with higher total compensation packages.

Perhaps surprisingly, directors are found to have a 10.8% higher hazard of exit than nondirectors, which could be because executives who are also directors are more inclined to leave the firm when their terms as directors expire. Alternatively, their accountability to shareholders may make directors more common targets for removal than nondirectors.

In Model 2 of Table 4, we report the results obtained on adding the board controls. Executives in firms with a female CEO have a 25.0% higher hazard of exit than those in firms with male CEOs. This result could indicate that other executives (who are primarily male) have less desire to work for female CEOs. Alternately, difficulties

[12] Of course, a greater than 100% reduction in the hazard rate is not literally possible, but it would be rare in practice to observe a CEO with the same amount of compensation, for example, as a vice president. CEOs are also generally older and much more likely to be directors than lower-ranked officers, so the thought experiment of holding all other variables constant in estimating the effect is almost completely hypothetical.

Table 4

Hazard estimates for executive exit

This table presents estimates of the coefficients of the Cox proportional hazards model. Standard errors are clustered by executive. Total direct compensation and total assets are deflated by the consumer price index for each year. Industry dummies are the Fama-French (1997) industry categories. Year controls consist of dummies for years 1996–2010, with 2009 and 2010 both excluded due to multicollinearity. Industry-adjusted values are residuals of OLS regressions of return on assets, stock returns, or Tobin's q on 47 of the 48 Fama-French (1997) industry dummies and 14 of the 15 year dummies.

	(1)	(2)	(3)	(4)
Variable	Base model coefficient [std. error]	+ Board coefficient [std. error]	+ Firm coefficient [std. error]	+ Yr & Indus coefficient [std. error]
Female	0.297***	0.297***	0.304***	0.298***
	[0.055]	[0.054]	[0.055]	[0.056]
CEO	−1.018***	−1.048***	−1.068***	−1.030***
	[0.052]	[0.053]	[0.053]	[0.052]
Age	0.071***	0.091***	0.087***	0.044**
	[0.019]	[0.021]	[0.021]	[0.020]
Age squared/10^2	−0.017	−0.032	−0.029	0.013
	[0.020]	[0.020]	[0.020]	[0.018]
Total direct comp./10^6	0.035***	0.039***	0.047***	0.040***
	[0.006]	[0.006]	[0.006]	[0.007]
Total direct comp. squared/10^{14}	−0.032***	−0.037***	−0.046***	−0.042***
	[0.010]	[0.011]	[0.012]	[0.012]
Director	0.108***	0.096**	0.095**	0.084**
	[0.037]	[0.038]	[0.038]	[0.038]
CEO is female (other than exec.)		0.250**	0.224**	0.129
		[0.112]	[0.113]	[0.114]
Number of other board members		−0.019***	−0.022***	−0.014**
		[0.006]	[0.005]	[0.006]
Proportion of other directors female		−0.602***	−0.502***	−0.665***
		[0.158]	[0.158]	[0.174]
Proportion of other directors independent		0.454***	0.459***	0.289***
		[0.060]	[0.060]	[0.089]
Average age of other directors		−0.044***	−0.046***	−0.044***
		[0.004]	[0.004]	[0.004]
Ind.-adjusted return on assets (proportion)			−0.334***	−0.405***
			[0.060]	[0.061]
Book-to-market ratio			−0.035	−0.021
			[0.027]	[0.029]
Total assets/10^{11}			0.022	0.016
			[0.020]	[0.019]
Industry-adjusted annual stock return (%)			0.027***	0.086***
			[0.005]	[0.017]
Industry-adjusted Tobin's q			−0.094***	−0.074***
			[0.015]	[0.014]
Number of observations	46,046	46,046	46,046	46,046
Number of subjects	8,707	8,707	8,707	8,707
Number of exits	5,557	5,557	5,557	5,557
Log pseudo-likelihood	−38,156	−38,041	−37,971	−19,623
Stratified by industry?	NO	NO	NO	YES
Year controls?	NO	NO	NO	YES

***, **, * indicate statistical significance at the 0.01, 0.05 and 0.10 level, respectively.

with entering the executive labor market may relegate female CEOs to weaker firms, yielding greater executive turnover. The size of the board has a negative effect on the likelihood of leaving: each additional member of the board (other than the executive) reduces the hazard rate by 1.9%. This result is consistent with the hypothesis that larger boards experience coordination failures that make it more difficult to remove executives from their positions.

The proportion of directors who are female has a negative effect on the chances of exit whereas the proportion who are independent has a positive effect. If the fraction of the board that is female (independent) increases by 10 percentage points, the probability of exit for an executive falls by 6.0% (rises by 4.5%). It could be that female boards are more lenient or supportive to executives while more independent boards are stricter and more demanding. Alternatively, better-functioning firms where executives are more apt to stay might be more inclined to appoint female directors while firms with high executive turnover might be more inclined to appoint independent directors as an attempt to stabilize the firm. The average age of the members of the board also has a negative effect on the hazard: 4.4% for each one-year increase in average age. This may be because older boards are more lenient toward executives or that they make better hiring decisions, which ultimately results in fewer firings.

Firm controls are added to Model 3 of Table 4. An increase in the firm's industry-adjusted return on assets lowers the hazard of exit as does an increase in industry-adjusted Tobin's q, whereas an increase in industry-adjusted stock returns raises the hazard. This result could reflect that stock returns are a more visible measure of firm performance, resulting in more lucrative external opportunities for executives, whereas accounting measures such as return on assets and firm value may be used more for internal evaluation or promotion decisions. The size of the firm, as measured by total assets, has a statistically insignificant effect on the chances of exit, as does the firm's book-to-market ratio.

When we add year controls to the model and stratify by industry (Model 4), most of the results remain the same in terms of sign and significance and similar in magnitude. The main exception is that the dummy for having a female CEO is now statistically insignificant.[13]

The shapes of the survivor functions for various groups, based on the estimates of Model 3 in Table 4 with year controls added, are presented in Figure 3 below, while Figure 4 displays the estimated instantaneous hazard functions for these same groups. According to Figure 4, the hazard of executive exit follows a complex nonmonotonic time path. Figures 3 and 4 show that at any given time, women are more likely to

[13] In unreported specifications, we re-estimate the models with salary used as the compensation measure in place of TDC. While most of the results remain similar, there are some differences worth noting. Salary has the opposite pattern of coefficients from TDC (a negative linear term a positive quadratic term), the size of the board has an insignificant instead of a negative effect, and total assets have a positive instead of an insignificant effect.

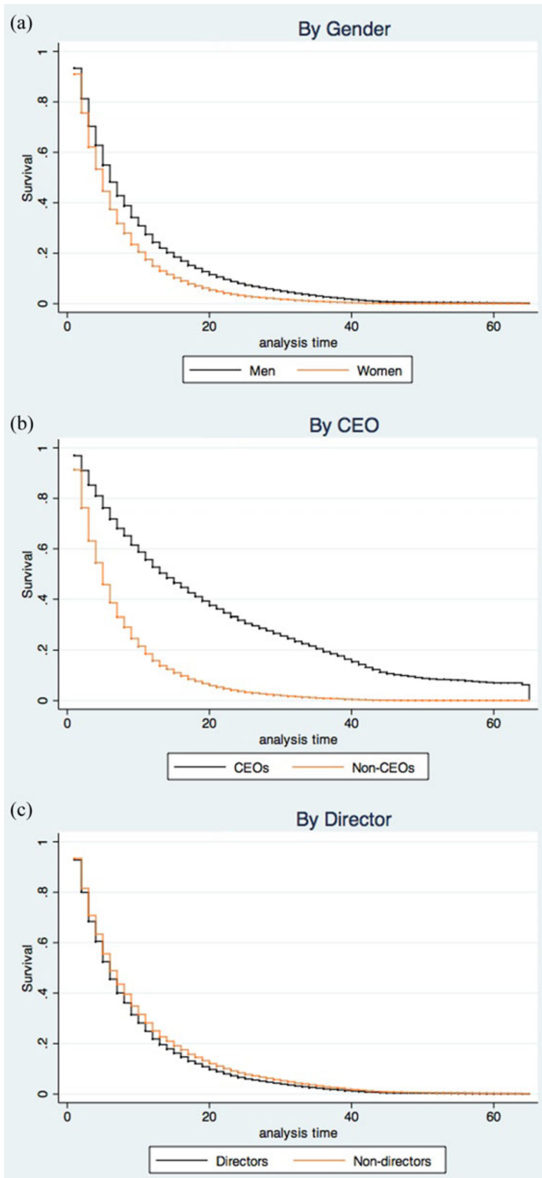

Figure 3

Estimated survivor functions from Cox model

This figure presents estimates of the survivor function, which is the probability that an executive remains in service at a firm for longer than a certain number of years. Survival probabilities are calculated from a Cox proportional hazards model that includes gender (a), CEO (b), director (c), quadratics in age and total direct compensation, board controls, firm financial controls, and year controls.

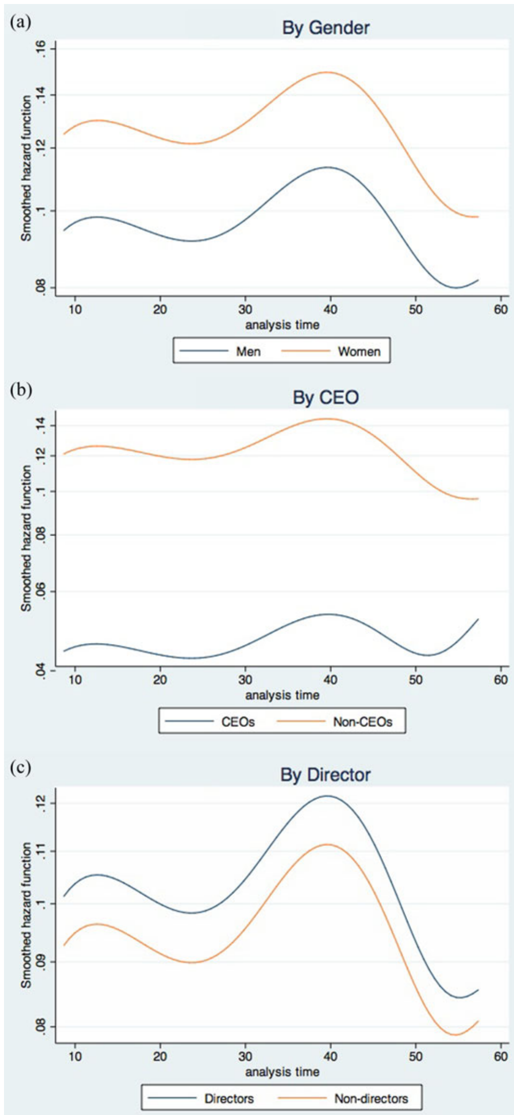

Figure 4

Estimated instantaneous hazard functions

This figure presents estimates of the hazard function, or the instantaneous probability that an executive leaves an executive capacity conditional on reaching a given length of service. Hazard functions are calculated from a Cox proportional hazards model that includes gender (a), CEO (b), director (c), quadratics in age and total direct compensation, board controls, firm financial controls, and year controls.

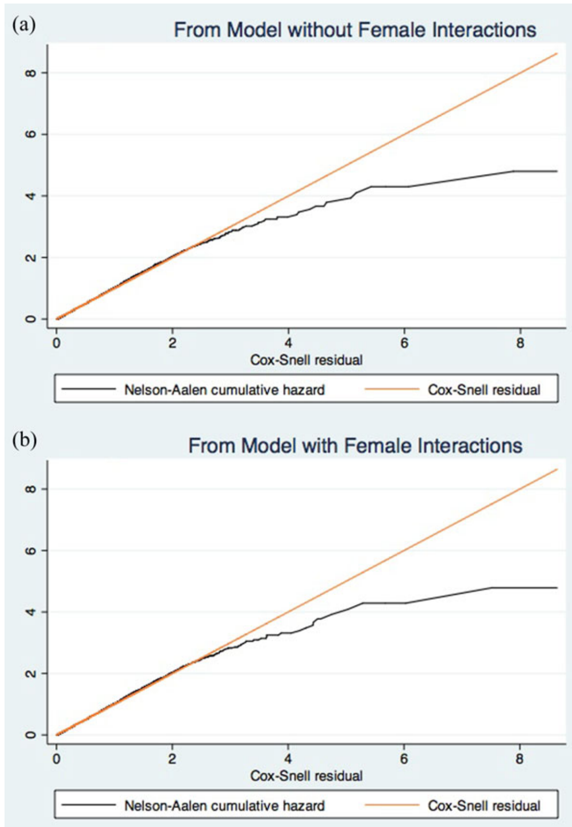

Figure 5

Cumulative hazard function of Cox-Snell residuals

This figure compares the Cox-Snell residuals to the Nelson–Aalen estimated residuals. Panel (a) calculates the Cox-Snell residuals from a Cox proportional hazards model that includes gender, CEO, quadratics in age and total direct compensation, director, board controls, firm financial controls, and year controls. Panel (b) calculates the Cox-Snell residuals from a Cox proportional hazards model that includes all explanatory variables in Panel (a) as well as interactions of the female variable with all of these variables. The Nelson–Aalen estimator is a nonparametric estimator of the cumulative hazard function obtained from raw failure time data without controls, where the original failure time variable has been replaced by the Cox-Snell residual.

exit than men, CEOs are less likely to exit than other officers, and directors are more likely to exit than nondirectors.

To assess the fit of these models, Figure 5 presents a graph of the cumulative hazard function for the Cox-Snell residuals from Model 3 in Table 4 with year controls added. The Cox-Snell residual for the jth observation (failing at time t_j) is essentially just the predicted cumulative hazard function at time t_j. Under the PH specification, the Cox-Snell residuals have an exponential distribution with hazard rate equal to one, which implies that their cumulative hazard function is a 45-degree line. Thus, we plot the Nelson–Aalen estimator of the cumulative hazard for the Cox-Snell residuals with a 45-degree line included in the graph for reference.[14] We see that at relatively low values of the Cox-Snell residuals, the cumulative hazard fits the 45-degree line almost perfectly, but at higher values the deviations from the 45-degree line become large. Although a reduction in fit is to some extent expected due to the reduction in the effective number of observations caused by prior failures and censoring, the graph does suggest that for large failure times the predictive power of the model is relatively poor. The low predictive power at large failure times could be a result of the delayed entry problem mentioned earlier: executives with long tenures began their careers well before the start of our sample period, and so there do not exist comparable executives within the sample on which to base predictions about their longevity.

5.3. Regression results: Specifications with female interactions

In this section, we interact the female variable with the executive, firm, and board-level variables to determine if these variables have different effects by gender. The estimates are presented in Table 5. The results in the first column are the uninteracted coefficients, which show the effects of each variable for men. Since men comprise the vast majority of the sample, these effects are very similar to what we see in the models that assume equal effects for women and men (i.e., Model 4 of Table 4).

The coefficients on the interaction terms show the difference in the effect of each variable for women versus men. We first note that the coefficients for females are jointly significantly different from the coefficients for males ($p < 0.01$). The effect of being a director is significantly more negative for women than for men, which may indicate that women face a higher opportunity cost of leaving firms where they are directors, perhaps because directorships are harder to obtain for women than men. Or, female directors may be highly coveted in order to demonstrate a firm's commitment to diversity and equality, so firms work harder to retain them.

The percentage of directors on the board who are independent also has a significantly more negative effect for women than for men. Thus, a higher proportion of independent directors increases the hazard rate significantly less for women than it does for men. This result is consistent with a scenario in which female executives are

[14] The Nelson–Aalen estimator is a nonparametric estimator of the cumulative hazard function $\Lambda(t)$ defined by $\hat{\Lambda}(t) = \sum_{t_j \leq t} d_j / n_j$, where d_j is the number of failures at time t_j and n_j is the number at risk of failure at time t_j. In this calculation, we replace the actual failure time t_j with the Cox-Snell residual.

Table 5

Hazard estimates for executive exit (results with female interactions)

This table presents estimates of the coefficients of the Cox proportional hazards model with female interaction terms. Standard errors are clustered by executive. Total direct compensation and total assets are deflated by the consumer price index for each year. Industry dummies are the Fama-French (1997) industry categories. Year controls consist of dummies for years 1996–2010, with 2009 and 2010 both excluded due to multicollinearity. Industry-adjusted values are residuals of OLS regressions of return on assets, stock returns, or Tobin's q on 47 of the 48 Fama-French (1997) industry dummies and 14 of the 15 year dummies.

Variable	Coefficient for men [std. error]	Interaction with female [std. error]
Female	–	−0.673
	–	[2.570]
CEO	−1.051***	0.542
	[0.053]	[0.374]
Age	0.045**	0.025
	[0.020]	[0.099]
Age squared/10^2	0.012	−0.031
	[0.019]	[0.097]
Total direct comp./10^6	0.039***	0.152***
	[0.007]	[0.045]
Total direct comp. squared/10^{14}	−0.041***	−0.751***
	[0.012]	[0.226]
Director	0.109***	−0.735***
	[0.039]	[0.247]
CEO is female (other than exec.)	0.206	−0.397
	[0.126]	[0.312]
Number of other board members	−0.014**	−0.006
	[0.006]	[0.023]
Proportion of other directors female	−0.591***	−0.733
	[0.174]	[0.569]
Proportion of other directors independent	0.329***	−0.658***
	[0.090]	[0.244]
Average age of other directors	−0.044***	0.015
	[0.004]	[0.014]
Ind.-adjusted return on assets (proportion)	−0.395***	−1.258**
	[0.064]	[0.504]
Book-to-market ratio	−0.023	0.055
	[0.028]	[0.131]
Total assets/10^{11}	0.005	0.028
	[0.020]	[0.037]
Industry-adjusted annual stock return (%)	0.093***	−0.030*
	[0.020]	[0.016]
Industry-adjusted Tobin's q	−0.074***	0.014
	[0.014]	[0.047]
Number of observations	46,046	
Number of subjects	8,707	
Number of exits	5,557	
Log pseudo-likelihood	−19,603	
Stratified by industry?	YES	
Year controls?	YES	
Joint significance of interactions	$\chi^2 = 45.71$***	

***, **, * indicate statistical significance at the 0.01, 0.05 and 0.10 level, respectively.

on average more effective than their male colleagues and in which more independent boards are more focused on evaluating executives based on observed talent and performance than boards that are more stacked with insiders. Alternatively, women may feel more comfortable working in firms that have more independent boards, perhaps because such boards are less male-dominated or less susceptible to the formation of old-boy networks even when they have a preponderance of males.

Industry-adjusted return on assets and industry-adjusted stock returns also have a significantly more negative effect on the probability of leaving for women than for men, albeit only at the 10% significance level in the case of stock returns. These results are consistent with a scenario in which women must do more than men to prove their worth with tangible financial results in order to be retained. Thus, when the performance of the firm improves, women experience a greater increase than men in the likelihood of being retained, or a greater decrease in the probability of being dismissed. (It may also suggest that the outside opportunities of women expand at a slower rate as firm performance improves.) Alternatively, it may be that better-performing firms are the ones that tend to hire and retain women, or that firms that retain women longer tend to be the ones that perform better. These possibilities could again signify that women in a male-dominated profession tend to be more able on average than their male peers.

TDC appears to have an effect that differs between women and men, but in fact it does not differ qualitatively. The values of the coefficients for men imply that the hazard increases as compensation increases up to a value of about $47.5 million, approximately the 99th percentile of TDC for men, so practically the hazard increases with TDC for the entire sample. However, the values of the coefficients for females imply that the hazard increases as compensation increases only up to the value of about $10 million. But $10 million is about a 97th percentile value of TDC for women, so over the vast majority of the sample the effect of TDC on the hazard of exit increases at a decreasing rate for women, just as for men.[15]

6. Application to the gender compensation gap

We now use the hazard analysis of the risk of executive exit to explain the gender compensation gap that is well documented in the literature (e.g., Bell, 2005; Elkinawy and Stater, 2011). Part of this difference could be due to a simple form of discrimination of lower pay for substantially identical work. But more subtle forms of discrimination could cause female executives to obtain lower levels of compensation

[15] When we use salary in place of TDC in the model with female interactions, we obtain mostly similar results. The exceptions are that the salary terms have the mirror image pattern from the compensation terms (negative linear term, positive quadratic term), the size of the board and the proportion of female members have insignificant rather than negative effects, total assets have a positive rather than insignificant effect, and the female-salary interactions have a different pattern than the female-TDC interactions (insignificant linear term, positive quadratic term rather than positive linear term and negative quadratic term).

when relegated to less remunerative posts. For instance, discrimination might force female executives to lead firms in less desirable industries (Paul and Sahni, 2010). Likewise, female executives could join firms with weaker future prospects and hence a higher probability of involuntary turnover.

Another possibility, which has not yet been formally investigated in the literature, is that women are penalized in the labor market for having a higher risk of turnover. We examine this question by including the predicted value of the hazard in a compensation regression. We therefore estimate the following model of compensation for an arbitrary executive i in year t:

$$\ln(TDC_{it}) = \gamma_0 + \gamma_1 Female_i + \gamma_2 \lambda_{it} + \boldsymbol{\delta} \boldsymbol{w}_{it} + \varepsilon_{it}, \tag{4}$$

where γ_1 and γ_2 are the coefficients of interest, *Female* is an indicator variable for female executives, λ_{it} is the hazard rate for executive i at time t, $\boldsymbol{\delta}$ is a vector of coefficients to estimate, $\boldsymbol{w}_{it}$ is a vector of control variables (potentially including executive characteristics other than gender, board information, firm financials, year dummies, and industry dummies), and ε_{it} is an error term assumed to have mean zero conditional on the female variable, λ, and $\boldsymbol{w}$. We estimate the model with pooled OLS (POLS).[16] To obtain λ_{it}, we estimate Cox models following the corresponding columns of Table 4 and obtain the predicted hazard for each executive in each year.[17]

Table 6 presents estimates of the compensation regressions with and without the hazard variable. In Model 1, which only controls for executive characteristics other than gender in $\boldsymbol{w}_{it}$ and excludes the hazard variable, we find that females earn 7.9% less TDC than males. But once the hazard is included, there is no remaining gender gap because γ_1 is insignificantly different from zero. The coefficient on the hazard, however, is negative and significant, indicating that executives with a higher risk of exiting earn less. The same pattern of results persists in Model 5, which includes executive characteristics other than gender, board information, firm financials, and year and industry dummies in $\boldsymbol{w}_{it}$. In Model 5 without the hazard variable, female executives obtain 13.6% less compensation. After including the hazard, female compensation is again insignificantly different from male compensation while the hazard again has a negative and significant effect on total compensation.

Thus, the results indicate that, once we account for the fact that executives with a higher risk of exit earn lower compensation, the difference in compensation by gender is reduced or even eliminated. That is, lower female compensation relative to males can potentially be explained by females having a greater risk of exit. The other

[16] Alternatively, we could estimate the model with panel data techniques. However, fixed effects cannot be used because the gender effect we are interested in cannot be identified from the fixed effect. Random effects could be used, but since that method achieves efficiency gains over POLS only if the variance-covariance matrix of the errors follows a specific form, random effects may not be superior to POLS. Moreover, POLS is consistent under weaker assumptions than random effects (e.g., strict exogeneity is not needed for consistency of POLS).

[17] To avoid simultaneity, we replace TDC with salary in the Cox model that determines the hazard.

Table 6

Total direct compensation estimates based on hazard of exit

This table presents regressions of total direct compensation on gender and various combinations of controls. Some specifications include the predicted hazard of exit as a regressor. Standard errors are clustered by executive. Executive controls include female, CEO, age, age squared, and director. Board controls include female CEO, number of board members, percent of female directors, percent of independent directors, and average age of board members. Firm controls include industry-adjusted return on assets, book-to-market ratio, total assets, industry-adjusted stock returns, and industry-adjusted Tobin's q. Year controls include dummies for 1996–2009, and industry controls include dummies for 47 of the 48 Fama-French (1997) industry categories. "Hazard" is the predicted hazard ratio from an industry-stratified Cox proportional hazards model with executive, board, firm, and year controls (salary and salary squared are added to the executive controls).

	(1)		(2)		(3)		(4)		(5)	
Vars.	Coeff [SE]	Coeff [SE]	Coeff [SE]	Coeff [SE]	Coeff [SE]	Coeff [SE]	Coeff [SE]	Coeff [SE]	Coeff [SE]	Coeff [SE]
Female	−0.079***	−0.009	−0.136***	−0.071***	−0.143***	−0.079***	−0.155***	−0.074***	−0.136***	−0.033
	[0.023]	[0.023]	[0.022]	[0.022]	[0.021]	[0.021]	[0.020]	[0.020]	[0.020]	[0.020]
Hazard		−0.265***		−0.251***		−0.234***		−0.295***		−0.382***
		[0.009]		[0.009]		[0.009]		[0.010]		[0.010]
N	46,030	46,030	46,030	46,030	46,030	46,030	46,030	46,030	46,030	46,030
R^2	0.157	0.172	0.245	0.257	0.306	0.316	0.331	0.344	0.364	0.382
Exec?	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y
Board?	N	N	Y	Y	Y	Y	Y	Y	Y	Y
Firm?	N	N	N	N	Y	Y	Y	Y	Y	Y
Year?	N	N	N	N	N	N	Y	Y	Y	Y
Indus?	N	N	N	N	N	N	N	N	Y	Y

***, **, * indicate statistical significance at the 0.01, 0.05 and 0.10 level, respectively.

intermediate specifications such as Model 2 through Model 4 of Table 6 continue to permit the interpretation of outright gender-based simple discrimination, as the coefficient on female remains significantly less than zero even after the hazard is included in the regression. However, the coefficient on female is invariably reduced by about half when we include the hazard. Therefore, the results suggest that at least a substantial portion of the gender compensation gap is due to the risk of exit and not necessarily the simple discrimination of lower pay for equivalent work.[18]

7. Conclusion

This paper contributes to the corporate governance literature by investigating a particular facet of governance, the tenure of executives, which unlike executive turnover has received comparatively little attention in financial economics research. Our study allows us to look at the tenure of CEOs as well as at the tenure of executives at the lower ranks by studying the factors affecting their exit probabilities in a given year. Of special interest to us is whether there are gender differences in the effects of executive, firm, and board characteristics on the probability of exit from the firm. The focus on gender is becoming more prevalent in the finance literature as females increase their representation in the executive ranks. Thus, understanding the determinants of executive exit deepens our knowledge of professional labor market dynamics.

Our findings suggest that issues related to gender are important factors in determining the exit rates of executives. Women are consistently found to have a higher hazard rate than men, all else equal, and the proportion of directors on the board who are female reduces hazard rates for executives in general. There is also evidence that boards that are more independent reduce hazard rates more for women than men. These results may indicate that women have difficulty gaining or maintaining access to entrenched networks of insiders within the firm. Firms with better accounting and stock performance appear to retain females more readily than males, suggesting that women may have to demonstrate superior financial results in order to be retained at similar rates as men. Being a director reduces hazards more for women than for men, perhaps because female directors are highly valued in a still male-dominated profession. We also find that the compensation gap between male and female executives may to some extent be driven by the relatively higher exit probabilities of women. In

[18] Even though we use salary instead of TDC to predict the hazard, some spurious correlation between the predicted hazard and TDC may still be present. Thus, as an alternative, we also estimate (unreported) specifications in which we predict the hazard without any compensation or salary measures at all (i.e., we use all controls in Model 4 of Table 4 except for TDC and TDC squared). Using this approach the reduction in the gender compensation gap (in the Model 5 of Table 6 with the full set of controls) when the hazard is included is smaller, on the order of about 12% instead of 45% or more, and females are still predicted to earn significantly less than males. The effect of the hazard remains negative and significant in this alternative approach, but is only about 20% as large in magnitude.

other words, if women executives were not observed to have a higher risk of exit, the compensation differential might be smaller.

An implication of our results is that female executives seeking longer careers with their firms may wish to seek out firms with more female directors, more independent directors, and strong accounting and stock performance (after adjusting for industry norms). These implications must be taken with caution, however, as the precise mechanism behind the gender differences that we find cannot be definitively ascertained with present data. For instance, it could be that better-run firms with lower executive turnover tend to be the ones that hire and retain females or that female executives who ascend to the top of the profession are more highly able on average than their male counterparts (though this may indicate the presence of barriers to entry into the profession for females). In either of these cases, lower hazard rates for females would be associated with markers of strong governance and financial performance without these factors being the primary cause of female retention. Further research will be required in order to disentangle these competing explanations and assess the relative merits of each. Gender equity in exit rates and job tenure is important in that developing and retaining the human capital of talented executives, regardless of gender, is conducive to the financial health of the world's largest and most influential companies.

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