Regulating Pay in Banking: Does CEO Bonus Compensation Increase Bank Default Risk?*

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29 March 2012

Abstract

We investigate the link between the incentive mechanisms embedded in CEO cash bonuses and the riskiness of banks. For a sample of U.S. and European banks, we employ the Merton distance to default model to show that increases in CEO cash bonuses lower the default risk of a bank. However, we find no evidence of cash bonuses exerting a risk-reducing effect when banks are financially distressed or when banks operate under weak bank regulatory regimes. Our results link bonus compensation in banking to financial stability and caution that attempts to regulate bonus pay need to tailor CEO incentives to the riskiness of banks and to regulatory regimes.

JEL Classification: G21; G33; J33
Keywords: banks; default risk; executive compensation

*We thank Elizabeth Cooper, John Duca, Bob DeYoung, Alex Edmans, Michael Faulkender, Kevin Keasey, Claire Mathews, Kevin McMeeking, Robert Prilmeier, Duane Stock, Aydin Ozkan and Larry Wall for comments and suggestions on previous versions of this paper. We also thank Vassiliki Bamiatzi and Konstantinos Bozos and participants at seminars organized by the Universities of Cagliari, Edinburgh, Exeter, and Hull as well as the participants at the 2011 FMA meetings in New York and Porto and the 11th FDIC Annual Bank Research Conference in Washington DC for helpful comments. Francesco Vallascas gratefully acknowledges financial support from a CAREFIN research grant and Jens Hagendorff gratefully acknowledges financial support from The British Academy (grant reference: SG100316). We are grateful to Christopher Keasey for excellent research assistance. A previous version of this paper was disseminated under the title 'CEO Remuneration and Bank Default Risk: Evidence from the U.S. and Europe'. The usual disclaimer applies.

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Abstract

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1 Introduction

Compensation practices in the banking industry are widely believed to have induced excessive risk-taking and, thus, to have played an important role in causing the recent financial crisis (e.g., Bebchuk and Spamann, 2009; IMF, 2010; Board of Governors et al., 2010). This paper analyzes the effects of CEO cash bonuses on the level of bank default risk for a sample of U.S. and European banking firms.

By way of preview, our analysis shows that increases in CEO cash bonuses lower the default risk of banks. We explain this finding by arguing that, because bonus payments are contingent on bank solvency, they lower the risk preferences of CEOs. We then show that the risk reducing effect of bonuses holds when we control for other types of incentive compensation. Among other things, we control for unobserved CEO heterogeneity, which has recently been identified as a major determinant of the level and structure of managerial pay incentives (Graham et al, 2012; Coles et al. 2011), to illustrate that our findings on bonuses are robust to the effects of other forms of CEO compensation.

We then demonstrate that the risk-reducing effect of cash bonuses disappears as banks move closer to the point of default. At the most risky banks, our results show that bonus payments promote rather than mitigate bank risk-taking. This finding is consistent with the view that financially distressed banks seek to maximize the value of the financial safety net. Finally, we analyze the bonus-risk relationship across bank regulatory regimes of varying power. We demonstrate that powerful regulatory regimes amplify the risk-reducing effect of CEO bonus payments. By contrast, when bank regulatory regimes are weak (and the prospect of regulatory intervention low) bonuses are no longer risk-reducing. We argue that it is likely that the risk-reducing incentives embedded in CEO bonus plans are counterweighed by incentives for banks to shift risk to regulators which are weak.

Our investigation comes against the background of an emerging policy consensus in the U.S. and Europe that compensation practices and bonus payments in particular have promoted excessive risk-taking at financial firms. For instance, the Financial Stability Board (2009) concludes in a report for the Group of Twenty that 'high short-term profits led to generous bonus payments to employees [which] amplified the excessive risk-taking that severely threatened the global financial system' (pg. 3). In response, U.S. compensation guidelines for CEOs and other senior executives at large banks come close to dictating that bonuses be paid in equity based instruments, that the payment be deferred and that vesting be performance-contingent (Board of Governors et al., 2010, pg. 33). Similarly, the Committee of European Banking Supervisors (2010) proposes in its draft guidelines on remuneration policies that at least 50% of variable compensation (including bonus payments) should be non-cash.[1] Crucially, our finding that CEO bonuses are linked to lower default risk challenges the view underlying these regulatory guidelines that cash-based executive compensation is linked to high risk-taking in the banking industry.

Given the attention which CEO compensation has recently attracted, it is important to gain a better understanding of how the structure of CEO compensation contracts relates to the riskiness of banks. However, most of the extant literature focuses on the risk incentives exerted by executives holding stock options (e.g., Hubbard and Palia, 1995; Hagendorff and Vallascas, 2011; DeYoung et al., 2012), while largely overlooking the role of CEO cash bonuses as a risk-taking incentive. Exceptions to this are studies by Harjoto and Mullineaux (2003) and Fahlenbrach and Stulz (2011). However, Harjoto and Mullineaux restrict their analysis to showing that CEO bonuses and stock volatility are positively linked and Fahlenbrach and Stulz study the effects of CEO bonuses on bank performance (rather than bank risk) during the recent financial crisis.

The lack of empirical work on CEO bonus plans on bank risk is unfortunate for two reasons. First, CEO cash bonuses are an important component of executive pay. In our sample of U.S. and European banks, CEO bonus payments make up a third of total CEO compensation. Second, the effect of bonus payments on managerial risk preferences is likely to differ from that exerted by option grants. Figure 1 illustrates the shape of a typical bonus function. CEO cash bonuses generally become payable once an earnings-based target over a one-year period has been met (Murphy, 2000). After exceeding this threshold, the CEO payoffs from bonus plans increase in performance until capped at a maximum payout. Therefore, for earnings performance exceeding the threshold at

[1] Previously, the Financial Services Authority (FSA) introduced a new remuneration code for UK financial firms which effectively limited basic compensation and bonus payments to 50% of CEO compensation (with the rest paid in equity and equity-like instruments). Also, both the UK and France have introduced a 50% tax on bonuses paid during 2009. (see IMF, 2010, *Appendix 2;* 'FSA Unveils Tough New Pay and Bonus code', *Financial Times,* July 30 2010; 'Big Bonuses Face Curbs', *Financial Times,* 17 May 2010).

which bonus payments become payable, CEO bonus plans do not provide convex payoffs (Smith and Stulz, 1985) and should not promote excessive risk-taking (see Noe et. al., 1996; Duru, 2005). This is in sharp contrast to CEO holdings of stock options. Since the payoffs from option holdings are convex functions of the volatility of stock returns, options incentivize CEOs to increase the volatility of share prices by engaging in riskier activities (Guay, 1999; Rajgopal and Shivaram, 2002; Coles et al. 2006; Hagendorff and Vallascas, 2011).

This paper extends the literature on the link between executive compensation in banking and risk-taking by providing the first cross-country evidence on the bonus-risk relationship in the banking industry. Existing studies have limited their focus on a U.S. market context to analyze the link between bonus payments and bank performance (Fahlenbrach and Stulz, 2011), stock volatility (Harjoto and Mullineaux, 2003) and bank default risk relative to non-cash compensation (Balachandran et al., 2010).

Next to the international dimension, the present analysis augments the results of previous studies in three ways. First, we explore how the managerial risk-taking preferences related to cash bonus payments differ at distressed banks. In these banks, shareholder-designed compensation contracts are likely to promote risk-shifting to bondholders and regulators because shareholders gain more from high-risk activities which maximize the value of the financial safety net than from attempts to increase low levels of charter value (Keeley, 1990; Park and Peristiani, 2007). Second, we exploit the cross-country dimension of our sample to test the prediction that regulatory regimes interact with pay incentives in shaping bank risk-taking (see Noe et al., 1996).

Third, our analysis is also able to shed some light on the debate whether bonuses exert undesirable short-term effects on banks. A popular criticism of bonuses in banking is that they fuel a culture of short-termism where excessive risks are taken to meet short-term performance goals (e.g. Financial Stability Board, 2009, pg. 13). Our results show that changes in cash bonuses have measurable long-term effects which are of even higher magnitude than their short-term effects. Therefore, our findings are at odds with the view that the effect of bonuses is short-term and add to a growing literature which shows that bonus compensation is an essential part of the long-term compensation policy of a firm (Gibbons and Murphy, 1992; Indjejikian et al., 2011).

We analyze the bonus-risk relationship by means of a partial adjustment model and by using a dynamic panel GMM estimator. Essentially, our results are based on an instrumental variable approach which uses lagged values of pay variables as instruments (which we demonstrate to be valid) for pay variables. This approach controls for unobservable heterogeneity, simultaneity issues (e.g., when CEO pay incentives are jointly determined with a firm's investment and other financial policies; see Coles et al., 2006) as well as for dynamic endogeneity (e.g., when current risk affects future CEO pay which, in turn, affects future risk-taking).[2] Wintoki et al. (2012) demonstrate the importance of controlling for dynamic endogneity in corporate governance research. The authors find that few of the previously identified relationships between corporate governance and firm performance are confirmed after controlling for a firm's history through a dynamic panel GMM estimator.

The paper proceeds as follows. The next section discusses the theoretical background to our study and formulates expectations which we test in the analysis. Section 3 describes the sample of U.S. and European banks. Section 4 discusses the methodology used to measure distance to default, while Section 5 discusses the empirical design adopted to analyze the relationship between CEO cash bonuses and default risk. We present our empirical results in Section 6. We conclude and offer some policy implications in Section 7.

2 Theoretical Background and Hypotheses Development

2.1 Theoretical Background

Executive compensation policy serves as a mechanism to reduce conflicts between managers and shareholders over the deployment of corporate resources and the riskiness of the firm. Over the past two decades, performance-related pay has increased markedly in and outside the banking industry. Generally, the increasing use of performance-sensitive compensation is consistent with agency theory which suggests that optimal executive compensation needs to align the interests of risk-averse managers with those of risk-neutral shareholders by motivating managers to commit to

[2] It is worth noting that the results we report in this paper do not rely on us using a GMM specification. In unreported tests, we obtain qualitatively similar findings when we employ a simple within-fixed effect estimator that does not include the lag value of the distance to default as an explanatory variable. However, a fixed effect model does not account for endogeneity between pay and risk and it does not allow us to distinguish between the short- and the long-term effects of bonus compensation on distance to default. The results of these additional tests are available from the authors upon request.

risk-increasing but positive net present value (NPV) projects (Jensen and Meckling, 1976; Smith and Stulz, 1985).

In the banking industry, the monitoring of managerial risk-taking is subject to a theoretically well-defined moral hazard problem. Banks are highly-leveraged institutions. Given the call option properties of equity, bank shareholders face limited downside risk but stand to benefit more from high-risk strategies that increase the volatility of bank assets than shareholders in less leveraged firms (Jensen and Meckling, 1976; John and John, 1993; Noe et al., 1996). Shareholder incentives to increase risk are further aggravated by deposit insurance and the operation of implicit bail-out policies which insulate many bank creditors from losses in the event of institutional failure and which reduce their propensity to limit excessive risk-taking by executives (see John et al., 2000, John et al., 2010).[3] As a result, shareholders are likely to use their control over CEO pay to encourage managerial risk-taking and the shifting of risk to regulators and bondholders (Benston et al., 1995; Hubbard and Palia, 1995; Bolton et al., 2010). While the effectiveness of pay-based incentives to take risk will vary across banks (e.g., a bank's future growth prospects have been identified as a mitigant to managerial risk-taking [Keeley, 1990]), bonuses may play an important role in affecting bank risk overall. We explain how CEO bonus plans affect bank risk-taking in the next subsection.

2.2 CEO Bonuses and Bank Risk-taking

Theoretical work posits that CEO cash bonuses can play an important role in mitigating managerial incentives to engage in risk-shifting. In a theoretical model, Smith and Stulz (1985) show that as long as cash bonuses increase linearly with corporate performance, the payoffs linked to a bonus plan are non-convex and, therefore, not inherently risk-rewarding. However, when performance is below the earnings-based threshold at which bonuses become payable, bonus plans resemble a call option on the performance measure. In this case, bonus plan payoffs will be convex and offset the concavity of the CEO's risk-averse utility function. By contrast, when performance is above the threshold at which bonuses become payable (and below the bonus cap), the slope of a bonus plan is

[3] While many European countries either did not have explicit deposit insurance during our sampling period or had insurance schemes which were less generous than the U.S. scheme, the operation of implicit or explicit bail-out policies in many European countries may not cause regulatory moral hazard on a scale comparable to that caused by the blanket guarantees of the Federal Deposit Insurance Corporation in the U.S..

linear with respect to performance and will not incentivize risk-averse CEOs to increase bank risk in order to secure higher bonus payments.[4]

Other work suggests that, rather than having no effect on risk-taking, bonuses lower the risk preferences of the CEO. One such argument goes that, because bonus payments can only be received in a state of solvency, they incentivize CEOs to avoid bankruptcy (Brander and Poitevin, 1992; John and John, 1993). Consistent with this, Duru et al. (2005) find that earnings-based cash bonuses make managers seek stable cash flows to meet contractual debt obligations. The authors show that the costs of debt financing decrease for firms which pay higher CEO cash bonuses and contend that this reflects the lower agency costs of debt and reduced risk-shifting incentives in these firms.

In spite of the arguments above, claims that cash bonuses encourage banks to engage in 'excessive' risk-taking continue to be made (for instance, in Financial Stability Board, 2009). However, the case for bonuses increasing risk tends to rest on two assumptions both of which have recently been challenged by empirical evidence. The first assumption is that cash bonus contracts do not sufficiently expose managers to downside risk and, therefore, reward managers for taking on more risk to achieve the performance goals underlying bonus contracts. In contrast to this, empirical studies have shown that bonus contracts tend to punish underperformance more than they reward strong performance (Leone et al., 2006; Indjejikian et al., 2011).[5]

The second assumption is that, by making bonus payments contingent on annual performance targets, shareholders design bonuses to affect short-term behavior and managers pursue higher risk-taking strategies to achieve these short-term goals. While some evidence exists that managers may

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[4] Kim et al. (2008) confirm the theoretical predictions made by Smith and Stulz (1985) using estimates of the performance thresholds underlying individual CEO bonus plans. However, it is important to bear in mind that, because the conditions under which managerial option plans are granted are seldom disclosed, estimates of the position of a CEO on the bonus payoff function will be imprecise. This is why our analysis is unable to empirically examine this aspect of risk-taking on bonus payments directly.

[5] One explanation advanced for bonus contracts producing asymmetric payoffs is that CEOs bonus payments are partly based on performance gains which have yet to materialize at the time the bonus becomes payable. Should the anticipated performance not materialize, it will be difficult for companies to claw back overpaid funds from the CEO. However, this so-called 'settling up problem' has been questioned recently. For instance, Shaw and Zhang (2010) are only able to replicate Leone et al.'s (2006) finding that cash compensation is more sensitive to downside than to upside risk when underperformance on the basis of stock returns, but not using other performance measures.

ration productive effort in order to maximize bonuses payments (Holthausen et. al., 1995; Bouwens and Kroos, 2011),[6] this does not rule out that bonuses can be designed as long-term compensation tools. In fact, Gibbons and Murphy (1992) and Indjejikian et al. (2011) show that bonus plans provide managers with long-term incentives to exert effort.These studies document that companies consider a trade-off between bonuses and career incentives over time horizons of multiple years when devising managerial compensation packages.

Evidently, cash bonuses are not the only form of incentive compensation that may affect CEO risk-taking. Stock and stock options may cause CEOs to overcome their risk aversion and incentivize them to engage in risk increasing projects (Smith and Stulz, 1985; Coles et al., 2006). Stock options may be particularly effective in this regard, because they have convex slopes with respect to firm value and increase the sensitivity of CEO wealth to stock return volatility (see Guay, 1999). By contrast, pension rights and other forms of unsecured, unfunded debt claims which CEOs hold may constrain managerial risk-taking (Sundaram and Yermack, 2007; Edmans and Liu, 2011).

The narrow empirical evidence available for the U.S. banking industry reaches conflicting conclusions on the role of CEO bonus payments and bank risk. Harjoto and Mullineaux (2003) report a positive association between bonus payments and return volatility (risk). Balachandran et al. (2010) proffer some evidence that the sum of bonus and other cash incentives reduces the probability of bank default. Based on the theoretical arguments above, and in line with the findings of this latter study, we expect to find that higher CEO bonus payments lower the default risk of a bank.

Previous studies have identified two factors that may moderate the bonus-risk relationship. First, cash bonuses may promote managerial risk-shifting at financially distressed banks (Noe et al., 1996; Benston and Evan, 2006). At distressed banks with poor growth prospects, higher risk-taking maximizes the value of the safety net to a bank (e.g., Keeley, 1990; Park and Peristiani, 2007). This is because when the value attributed to the growth prospects of a bank (the so-called

[6] The notion that bonuses impact short-term CEO behavior has been advanced in Weizman's (1980) 'ratcheting' argument. Accordingly, managers faced with earnings-based bonus contracts may intentionally withhold productive effort to avoid future losses which would result if favorable performance at present led to higher performance targets in the immediate future. Holthausen et. al. (1995) and Bouwens and Kroos (2011) have found evidence consistent with ratcheting by management.

charter value) is low, the prospect of losing charter value in the event of institutional failure is not a sufficient deterrent against risk-shifting behavior. Consistent with this, Akerlof and Romer (1993) propose the 'looting hypothesis'. When senior management knows that there is a high likelihood a bank could fail, they are likely to undertake riskier strategies in an effort to boost short term performance. While such behavior may increase the probability of bank default, managers expect to benefit in terms of continuous employment and future pay should their gamble pay off. Similarly, Edmans and Liu (2011) argue that since CEO cash bonuses are only paid in states of solvency, CEOs at distressed firms may engage in risk-shifting to 'gamble for solvency' with little regard for the liquidation value of a firm.[7]

The second factor which may moderate the bonus-risk relationship is the overall strength of the bank regulatory regime. When the prospect of regulatory intervention is low, cash bonuses may lose their risk reducing effect (Noe et al. 1996). Webb (2008) suggests that as regulatory monitoring weakens, bank executives become less risk averse. Consequently, it could be argued that when regulatory environments are less strict and the probability of regulatory intervention is low, CEOs may be become more responsive to incentives to shift risk to regulators.

We sum up our expected findings as follows. We expect to find that CEO bonus payments mitigate risk-shifting incentives. However, when banks are financially distressed or when the regulatory regime is weak, we do not expect to find that CEO bonus payments lower the default risk of a bank.

[7] Similarly, if CEOs believe that previously paid bonuses will not have to be returned in the event of institutional failure in a future year, it could also be argued that bonus payments at distressed banks will insulate CEOs from the longer-term negative effects of high-risk, negative NPV projects and provide incentives to shift earnings over time (Duru et al., 2005). However, there are two exceptions to this—although, admittedly, both of these have been difficult to enforce in the past. In the U.S., the FDIC may sue bank managers for gross negligence if the case can be made that a CEO contributed to the failure of bank. Also, some CEO may have agreed to clawback provisions which stipulate repayment of cash bonuses in the event of subsequent underperformance.

3 Sample and Remuneration Data

3.1 The Sample

Our analysis is based on 117 listed banks (76 U.S. bank and 41 European banks) in the period 2000 – 2008. For the U.S. subsample, we collect compensation data from Standard & Poor's ExecuComp database. Since ExecuComp's coverage is restricted to firms currently or previously included in the S&P 500, S&P MidCap 400, and S&P SmallCap 600, we eliminate many banks at this point. We extract compensation data for firms with SIC codes between 6000 and 6300. We require at least five years of compensation data to implement the dynamic panel model as described in Section 5. We then exclude firms with SIC 6282 (investment advice) and SIC 6111 (federal credit agency). Out of firms with SIC codes 6211 (Security Brokers and Dealers) and 6199 (Finance Services), we retain Citigroup in the sample. Based on these criteria, our search yielded an initial sample of 103 U.S. banks.

Since we wish to use identical data sources for all sample banks, we further stipulate that banks need to have five years of accounting data on Fitch's IBCA's Bankscope database and equity return data on Datastream over the same time period. We lose 5 banks as a result of this criterion. Further, since our default risk calculations require daily market data, we ensure that the calculations are reliable and stipulate that sample banks have at least 70% of non-zero daily returns on Datastream each year. We lose a further 22 banks as a result of this liquidity criterion. Our final sample contains 76 U.S. banks.

To build the subsample of European banks, we start with the universe of listed banks which are listed on Bankscope and chartered in Europe (EU-15 plus Norway and Switzerland). After excluding institutions with SIC codes 6211 and 6199, we identified 200 banks. Next, we stipulate that sample banks have at least five years of accounting data and equity return data subject to the same liquidity criterion we applied to the U.S. sample. Applying these criteria halved the number of banks in our sample.

Finally, we stipulate that sample banks need to have CEO remuneration data available for at least five consecutive years. We hand collect compensation data from annual reports and governance reports. The levels of CEO pay disclosure vary widely across European countries. In a first group of countries (Belgium, France, Ireland, Italy, the Netherlands, and the UK), pay disclosure requirements are very high and comparable to those applicable to listed U.S. firms. In a second group of countries (Austria, Germany, and Spain), detailed disclosure of CEO pay is not required, but a small number of banks still disclose CEO pay data. Finally, a third group of countries (Switzerland, Greece, Portugal, Norway, Finland, Sweden and Luxembourg) does not disclose CEO pay data for most of the sample period and typically reports pay data aggregated at the level of the board. In order to be sampled, European banks have to be based in the first and second group of countries. Since we require five years of CEO pay data for our dynamic panel, we are unable to include banks in the third group that have recently started to publish CEO pay data. Our final sample contains 41 European banks.

Table 1 near here

The distribution of the final sample by year and country is reported in Table 1. European sample banks are mostly located in the UK and in Italy which reflects the concentration of listed banks in both of these countries. Further, in spite of having selected the largest listed U.S. banks, the average size of U.S. banks in our sample is smaller than the European banks in our sample.[8]

3.2 Bonus Data

For U.S. banks, we extract compensation data from ExecuComp. We obtain the variables SALARY, BONUS and OTHCOMP to construct measures of CEO cash compensation.[9] While ExecuComp offers a much broader range of remuneration data than we extract, the range of compensation data we examine in this study is determined by the availability of similar data for European banks. For the latter group of banks, we use annual reports and governance reports to hand collect CEO data on basic salary, cash bonuses, and other forms of cash compensation.

Using the information collected, we construct two remuneration variables. **LGBONUS** is the logarithmic transformation of 1 plus the total cash bonus received by the CEO (in $ thousands; for

[8] We assess the implications of size differences between the U.S. and European banks for our results in the robustness section where we re-estimate our main models after excluding smaller banks. Our main results are not affected if we exclude the smallest banks from the sample.

[9] We do not use the data field TDC1 as a measure of total compensation, because it includes a valuation of new stock option grants and no comparable data on CEO stock options grants are available for most European banks.

European banks, currency conversions are based on IMF exchange rates at the date of the annual report). The relative importance of bonus payments is captured by **RELBONUS** which is the ratio of CEO cash bonuses to total cash compensation.

Table 2 reports descriptive statistics on the bonus variables and demonstrates that cash bonuses are key components of CEO pay in the banking industry. Panel A of Table 2 shows that the average (median) bonus paid over the sample period is $1.234 ($0.461) million. On a bank-year basis, cash bonuses on average contribute one-third to total CEO compensation.[10]

Next, Panel B of Table 2 reports basic descriptive statistics for European and U.S. banks. The analysis of the two sub-samples shows that the value of bonuses and total compensation appears to be larger in Europe. Nonetheless, the ten largest bonus payments are all made to U.S. bank CEOs. For instance, the CEO of Citigroup, Sanford Weill, received a bonus of $29 million in 2003, while the CEO of Countrywide Financial Corporation, Angelo Mozilo, received a bonus of around $20 million in the same year. In Europe, Josef Ackermann, the CEO of Deutsche Bank, received a cash bonus of $12 million in 2007, followed by the CEO of Unicredit, Alessandro Profumo, who received a bonus of $8.8 million during the same year.

Table 2 near here

Since bank size is a major determinant of executive compensation policies, Panel C of Table 2 reports summary statistics for the sample after excluding banks located in the lowest quartile of the distribution of average total assets during the sample period. After excluding these smaller (mainly U.S.) banks, the summary statistics provide few indications that U.S. and European compensation practices are different.[11]

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[10] Interestingly, the share of bonus payments declined to 10% of total CEO pay in 2008. While the declining share of bonus payments denotes a sharp change in the remuneration practices at banks in the aftermath of the recent financial crisis, our results are not be affected by this since our regression models employ lagged explanatory variables (see Section 5). The last year of compensation data employed in our models is 2007.

[11] In untabulated tests, we evaluate whether there are statistically significant differences between the compensation practices of European and U.S. banks. We employ the Wilcoxon rank-sum test to examine if differences between U.S. and European banks in Panel C of Table II are statistically significant. We do not find significant differences between U.S. and European Banks. The results of these tests are available upon request.

4 Measuring Bank Default Risk

We measure bank risk via the Merton distance to default (DD) model. This model boasts a wide range of empirical (e.g., Vassalou and Xing, 2004; Sundaram and Yermack, 2007; Vallascas and Hagendorff, 2011) and commercial applications. Default frequencies estimated using the DD model correlate around 80% with default frequencies estimated using the commercially applied and in many ways similar Moody's KMV proprietary model (Bharath and Shumway, 2008). The DD model has been demonstrated to provide a suitable indicator of bank fragility which outperforms other market-based indicators such as subordinated debt spreads in predicting bank default (Gropp et al., 2006). An additional advantage of DD is that it may be calculated for all listed banks (while other market-based risk indicators such as CDS spreads are only available for a limited number of banks, especially in Europe).

The Merton model measures default risk as the number of standard deviations that the market value of bank assets are above default point (the point where the market value of assets is less than the book value of total liabilities). Formally, DD at the end of year t is expressed as:

$$\mathrm{DD}_t = \frac{\ln\left(V_{A,t}/L_t\right) + \left(r_f - 0.5\sigma_{A,t}^2\right)T}{\sigma_{A,t}\sqrt{T}}, \tag{1}$$

where $V_{A,t}$ is the market value of assets at the end of the fiscal year t, L_t is the book value of total liabilities, r_f is the risk-free rate (proxied by the yield on 1-year U.S. treasury bills for U.S. banks and by the 1-year interbank [EURIBOR] rate for European banks, $\sigma_{A,t}$ is the annualized asset volatility at t (based on volatility of equity returns in a given year), and T is the time to maturity (conventionally set to one year). The computation of DD_t requires estimates of $V_{A,t}$ and $\sigma_{A,t}$ neither of which are directly observable. Following Akhigbe et al. (2007), Vassalou and Xing (2004) and Hillegeist et al. (2004), we infer the values for $V_{A,t}$ and $\sigma_{A,t}$ through an iterative process based on the Black-Scholes-Merton pricing model. Specifically, we express the market value of a firm's equity ($V_{E,t}$) as a function of the asset value by solving the following system of nonlinear equations:

$$V_{E,t} = V_{A,t}N(d_{1,t}) - X_t e^{-r_f T} N(d_{2,t}) \tag{2}$$

$$\sigma_{E,t} = \left(\frac{V_{A,t}}{V_{E,t}}\right) N\left(d_{1,t}\right) \sigma_{A,t} \tag{3}$$

Equation (2) defines $V_{E,t}$ as a call option on the market value of the bidder's total assets, with $d_{1,t} = \frac{\ln\left(V_{A,t}/L_t\right)+\left(r_{f_t}+0.5\sigma^2_{A,t}\right)T}{\sigma_{A,t}\sqrt{T}}$ and $d_{2,t} = d_{1,t} - \sigma_{A,t}\sqrt{T}$. Equation (3) is the optimal hedge equation that relates the standard deviation of a bank's equity value to the standard deviation of a bank's total assets value (both on an annualized basis).

To solve this system of equations, we employ as starting values for $\sigma_{A,t}$ the historical volatility of equity (computed on a yearly basis using daily data) multiplied by the ratio of the market value of equity to the sum of the market value of equity and the book value of total liabilities, i.e. $\sigma_{A,t} = \sigma_{E,t}V_{E,t}/(V_{E,t} + L_t)$. A Newton search algorithm then identifies the daily values for $V_{A,t}$ and $\sigma_{A,t}$ in an iterative process which we then employ to compute DD_t as in Equation (1).[12]

Table 3 near here

In Table 3 we offer some preliminary evidence on the relationship between DD and CEO bonus pay. More precisely, the Table reports the mean and median values of our risk measure for banks that show low and high values of LGBONUS and RELBONUS. The results show that the mean and median values of DD are higher in banks where CEOs receive larger bonus payments. In addition, as reported in Column 3, the differences between the two groups of banks are statistically significant at customary levels. The results of this simple univariate test support theoretical arguments in favor of a risk-reducing effect of cash bonuses.

[12] The distance to default of the banks in our sample assumes particularly low values in 2008—the peak of the recent financial crisis. More precisely, the mean (median) DD in 2008 is equal to 1.27 (1.27) against a value of 4.41 (4.14) for the rest of the sample period. DD, therefore, seems to be particularly effective in tracking the evolution of bank default risk.

5 Empirical Design

5.1 The Model

In this paper, we build on previous work and model the relationship between CEO remuneration and bank risk-taking in a dynamic framework.[13] We assume that banks have an internally optimal, unobservable target distance to default to which they revert over time. We also assume that reversion to this target level of risk will not be instantaneous, but will take place via observable changes in distance to default over time. Therefore, observable adjustments in risk over time ($\mathrm{DD}_{i,t} - \mathrm{DD}_{i,t-1}$) will be a fraction of the differences between the target level of risk ($\mathrm{DD}^*_{i,t}$) and levels of risk existing at t-1.

$$\mathrm{DD}_{i,t} - \mathrm{DD}_{i,t-1} = \lambda \left(\mathrm{DD}^*_{i,t} - \mathrm{DD}_{i,t-1}\right) + \varepsilon_{i,t}, \tag{4}$$

In this framework, λ is the adjustment speed toward the long-term distance to default targeted by a bank. The adjustment speed will lie between the [0, 1] interval. If $\lambda = 0$, no observable risk adjustment between risk in a given time period $t$and the risk observed in the previous period occurs, whereas $\lambda = 1$ indicates that adjustment to the target distance of default is instantaneous.

Since the target level of risk is unobservable, we express $\mathrm{DD}^*_{i,t}$ as a function of bonus payments, BONUS_{t-1}, a vector of bank characteristics $\mathbf{X}_{i,t-1}$, and unobservable bank attributes (B_i):

$$\mathrm{DD}^*_{i,t} = \gamma \mathrm{BONUS}_{i,t-1} + \beta \mathbf{X}_{i,t-1} + B_i \tag{5}$$

Therefore, substituting Equation (5) into Equation (4), distance to default at a given time period t and its determinants will be driven by the following process:

$$\mathrm{DD}_{i,t} = (1-\lambda)\mathrm{DD}_{i,t-1} + \lambda \left(\gamma \mathrm{BONUS}_{i,t-1} + \beta \mathbf{X}_{i,t-1} + B_i\right) + \varepsilon_{i,t} \tag{6}$$

Here, $\lambda \times \beta$measures the short-term effect of a variable on distance to default in a given period, while β captures the long-term effect of a variable by measuring its effect on the unobservable target

[13] Among others, Shrieves and Dahl (1992), Rime (2001) and Jokipii and Milne (2011) use a dynamic framework to model the capital buffers held by banks.

level of default risk. Consequently, the impact of each explanatory variable on the target level of risk $DD^*_{i,t}$ can be derived by scaling the coefficients in Equation (6) by the speed of adjustment λ.

5.2 Estimation Method and Control Variables

To account for the impact of unobservable bank heterogeneity on default risk, we estimate Equation (6) by using a dynamic panel data method. Specifically, we use the Generalized Method of Moments (GMM) estimator which is an instrumental variable estimator proposed by Blundell and Bond (1998). This system GMM estimator uses both the lags of the explanatory variables and the data from the original level specification as instrumental variables.

A critical advantage of any GMM estimator is that it controls for endogeinity in the relationship between risk and managerial incentives. This is because the specification allows variables to be treated as strictly exogenous regressors, endogenous regressors, and predetermined regressors.[14] Thus, our method examines the effect of CEO pay on bank risk while ruling out reverse causality arguments according to which bank CEO who wish to pursue more risky strategies encourage their boards to offer risk-inducing pack packages. Further, our approach controls for dynamic endogeneity (when past compensation structures affect future bank risk-taking which, in turn, affects future CEO compensation structures; see Wintoki et al. 2012).

Additionally, the specific system estimator we employ in this study has two key advantages over other dynamic panel data methods (most notably, the difference-in-difference estimator proposed by Arellano and Bond [1991]). First, as long as the instruments are valid, the GMM estimator exhibits higher levels of both consistency and efficiency. Second, unlike the difference estimator, the system GMM estimator permits the use of time-invariant (or highly persistent) variables in our specifications.

System GMM specifications may be estimated either via a one-step or a two-step approach. While the one-step estimation produces unbiased standard errors, it is not asymptotically efficient in the estimation of the coefficients. The asymptotically more efficient two-step estimator, on the other hand, tends to bias the estimated standard errors downwards (Blundell and Bond, 1998). We

[14] Commonly, in the difference equation, first differences in endogenous variables are instrumented with lags of the specific variables. In the level equation, endogenous variables in levels are instrumented with lags of their first differences.

use the asymptotically efficient two-step GMM system as well as the Windmeijer (2005) procedure to lower the bias and correct the standard errors.

Finally, the validity of our approach rests on two assumptions (which we confirm to be valid for each of our estimations). First, for the instruments to be valid, they need to be uncorrelated with the error term. We use the Hansen J-statistic of over-identifying restrictions to test this assumption (where statistically insignificant values confirm the validity of the instruments). Second, the system GMM estimator requires stationarity in the post-instrumentation error terms. This implies the absence of second-order serial correlation in the first difference residual. We employ the m_2 statistic developed by Arellano and Bond (1991) to test for the lack of second-order serial correlation in the first-difference residual. An insignificant m_2 statistic indicates that the model is correctly specified.

Our models control for several bank characteristics that may affect the level of bank default risk. First, we control for bank leverage defined as 1 minus the ratio of the book value of equity to total assets (**LEVERAGE**). Given the call option properties of equity (which combine unlimited upside with limited downside potential), the benefits to shareholders associated with risk-shifting increase with firm leverage (Jensen and Meckling, 1976). Therefore, higher levels of bank leverage will induce CEOs to shift risk to regulators and debtholders (John and John, 1993). We expect LEVERAGE to increase the target levels of default risk and, hence, a negative relationship between LEVERAGE and DD.

We control for bank performance measured by the ratio of pre-tax profits to total assets (**ROA**). We introduce this variable to the baseline specification because earnings are the most frequent performance measure on which bonus payments are based (Murphy, 2000; Basel Committee, 2011).[15] Also, we wish to control for any possible link between bank profitability and default risk. For instance, higher profitability may make it easier for banks to retain earnings, build up capital and increase DD. On the other hand, higher profitability may equally be due to higher risk taking (and therefore be linked to lower DD). We also control for the future growth prospects of a bank because growth prospects may constrain risk-taking. Commonly, a bank's future growth prospects are captured by charter value. Keeley (1990) and Park and Peristiani (2007) show that management's

[15] A recent report by the Basel Committee (2011) into remuneration in the banking industry states that 'measures of financial performance, such as targets based on revenue, profit or income, cash flow or return on equity, are still frequently used for performance measurement. Measures used are often accounting-based and retrospective' (pg 3).

desire to preserve charter value will help mitigate against managerial risk-shifting incentives. We proxy charter value (**CHARTER**) by the market to book ratio. We expect CHARTER to enter our model with a positive coefficient.

We measure the importance of fee based and trading activities in a bank's business model via the ratio of non-interest income to total operating income (**NONINTSHARE**). Previous studies generally conclude that the revenue streams associated with non-conventional business lines tend to be more volatile than interest based activities (Stiroh, 2006). We, therefore, expect NONINTSHARE to exert a negative influence on distance to default.

The impact of bank **SIZE** (measured by the log transformation of total assets) on managerial risk-taking is ambiguous ex ante. On the one hand, large banks may benefit from a broader investment opportunity set, including improved access to funding and capital markets. On the other hand, large banks enjoy greater protection under the financial safety net. Therefore, implicit and explicit too-big-to-fail policies may exacerbate existing risk-shifting incentives and increase the level of risk (see Benston et al., 1995).

Table 4 near here

The final vector of control variables includes corporate governance and CEO characteristics. Agency theory postulates that conflicts over the riskiness of a firm ensue, because managers are risk-averse and shareholders are risk-neutral. When the board's interests are aligned with those of shareholders, this should be associated with increased risk-taking, while banks with less shareholder-oriented boards should be associated with less bank risk-taking (Pathan, 2009). We measure the degree to which the board's interests are aligned with those of shareholders through a measure of board independence (**INDEP**). INDEP is the share of directors who are not executives, former executives or related to (former) executives to board size. We expect more independent, shareholder-controlled boards to be associated with higher risk. Further, as a proxy for CEO power, we use a dummy variable which is equal to 1 if the CEO also acts as chairman of the board (**DUALITY**). We expect DUALITY to exert a risk-reducing effect. Finally, we also include **AGE** as the log transformation of CEO age in our model. Older CEOs with an established track record may be subject to less regulatory scrutiny than younger CEOs and should, therefore, display more risk-taking behavior.

Finally, when estimating Equation (6), we make the assumption that the bonus variables, LEVERAGE and ROA are endogenous with respect to bank risk. This assumption is strongly supported by previous work. First, Coles et al. (2006) point out that managerial incentives, investment and financial policies are likely to be jointly determined. Thus, treating the bonus variables as strictly exogenous would bias inferences about any causal relationship between firm risk and managerial incentives. Second, banks adjust their equity ratio (through changes in their capital buffers) and their risk simultaneously (e.g., Shrieves and Dahl, 1992; Jokipii and Milne, 2011). In our empirical design, this implies a mutual causal relationship between DD and LEVERAGE.[16] Finally, the endogeneity of ROA could be explained by either a tradeoff between risk and return (i.e. higher profitability is achieved through increased bank risk-taking) or, alternatively, the notion that the default risk of a bank affects the investment strategies which managers pursue in search of maximizing future returns (e.g. poor performance could lead banks to engage in a 'gamble for solvency').

Panel A of Table 4 presents descriptive statistics of the default risk scores, the CEO bonus variables and the control variables described above. For ease of exposure, we replicate the summary statistics on the compensation data which we have presented in Table II. Panel B of Table 4 reports pairwise correlations between the variables.

6 Bonus Compensation and Bank Default Risk

6.1 Multivariate Results

In this section, we analyze the impact of CEO cash bonuses on the target level of default risk. We expect bonus payments to exert a positive (i.e. risk-reducing) influence on the distance to default of a bank. Since cash bonuses can only be received in states of bank solvency, larger cash bonus payments should incentivize CEOs to avoid default. Further, the payoffs from cash bonuses tend to be linear functions of performance which do not reward increased risk-taking (provided that current

[16] The number of instruments may affect the validity of the specification tests. To limit the number of instruments as well as for simplicity's sake, we assume that the lagged values of the remaining bank characteristics are exogenous with respect to bank risk. However, as explained in Section 6.5, our results are confirmed when we model all bank characteristics as endogenous and use their lag values (from time $t-2$) as instruments.

bank performance does not place CEOs below the threshold at which a bonus is payable; see Smith and Stulz, 1985).

The results reported in Table 5 confirm our expectation. Panel A reports the results of the two-step system GMM estimations on bank default risk. The estimated coefficients can be interpreted as the short-term determinants of bank default risk. Both LGBONUS (which measures the absolute dollar contribution of bonus payments to CEO wealth) and RELBONUS (which captures the importance of cash bonuses relative to total cash compensation) enter the models with a positive sign (both highly statistically significant). Accordingly, higher bonus payments lower default risk. This result is robust to the inclusion of bank control variables and country dummies.[17]

In all models reported in Table 5, specification tests confirm the validity of our empirical approach. First, there is no second-order serial correlation in the first difference residual as indicated by an insignificant m_2 statistic. Second, the validity of our instruments is confirmed by an insignificant Hansen J-statistic of over-identifying restrictions. The latter confirms the absence of correlation between the instruments and the error term. While this test offers support for the validity of the instruments *as a group*, Table 5 also reports the results of a difference- in J-test to assess the validity of the specific instruments we employ for the bonus variables, EQUITY and ROA. The results show that these instruments are valid as we cannot reject the null hypothesis of exogeneity of the instruments. Third, as required by the partial adjustment framework, the coefficients on the lagged depended variables (DD_{t-1}) are positive, statistically significant at customary levels and lie within the [0;1] interval. Next to confirming the validity of adopting a dynamic framework, the observed coefficients of previous DD on current DD also indicate the speed at which banks adjust towards their long-run level of risk. Our estimations show that sample banks close between 47% and 46% (i.e. 1 minus the coefficient on DD_{t-1}) of the gap between desired and actual risk in one year.[18]

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[17] To avoid perfect collinearity between any bank-specific effects and the country fixed effects, we only introduce country dummies into the model specifications if we have at least three banks for a country.

[18] The adjustment speed in our sample of U.S. and European banks is higher than the adjustment speed that Jokipii and Milne (2011) estimate for a sample of U.S. banks.This could be explained by several reasons. First, Jokipii and Milne (2011) employ accounting data, while the inclusion of market data in the DD calculations makes our risk measure more forward looking. Second, the banks in our sample are larger and may, hence, have more opportunities

Interesting results also emerge when we examine the coefficients on the control variables. First, we observe a negative relationship between LEVERAGE and distance to default. This is consistent with the argument that risk-taking incentives increase with bank leverage as leverage improves the expected payoff from risky investment strategies (Jensen and Meckling 1976; John and John, 1993). Second, banks that engage in non interest income activities exhibit a lower distance to default as suggested by previous studies on the impact of diversification strategies on banking risk. Third, larger banks are associated with lower risk-taking. While this result runs contrary to too-big-to-fail arguments (i.e. large banks face incentives to increase risk-taking in an effort to maximize the value of safety net guarantees), the positive coefficient on SIZE is consistent with larger banks having access to more diversified and liquid sources of funding as well as perhaps more advanced risk management techniques.

Table 5 near here

In Columns (5) and (6), we re-estimate the models after including corporate governance variables. Due to missing board data for some banks, the sample size for these models is reduced by around 10% compared with previous specifications. However, in spite of the reduction in the sample size, we can confirm the main finding of a positive association between CEO cash bonus compensation and distance to default.

The coefficients reported in Panel A of Table 5 measure the short-term impact of bank characteristics on bank default risk. Our empirical strategy also allows us to explore the long-term marginal effects of bonus payments on the target level of default risk which we report in Panel B (computed as the coefficients reported in Panel A divided by 1 minus the coefficients on lagged DD). For ease of exposition, we only present the long-term marginal effects on the bonus variables.[19] LGBONUS and RELBONUS are both positive and highly significant showing that CEO cash bonuses also exert a risk reducing effect on the level of risk which banks target in the long-term. Further, as shown in Panel C, the magnitude of effect of bonus compensation on the long-term ef-

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to alter their risk level more rapidly than smaller banks (e.g. via easier access to funding and capital markets). Third, while we employ annual data over a recent and relative short time period, Jokipii and Milne (2011) examine bank risk using quarterly data during 1986 – 2008.

[19] The estimated long-term coefficients on the remaining variables are available from the authors upon request.

fects are higher than the short-term effects (significant at below 5%). This finding is particularly interesting, because it demonstrates that the risk effects of bonuses are higher in the long- than in the short-term. Evidently, bonuses are not a short-term compensation tool.

6.2 Additional Analysis: Controlling for Other Pay Variables

Cash compensation is not the only type of incentive compensation which could affect the risk preferences of bank CEOs. Both equity-like compensation such as options and stock grants (Smith and Stulz, 1985; Guay, 1999) as well as debt-based compensation such as pension arrangements (Sundaram and Yermack, 2007; Edmans and Liu, 2011) may equally affect managerial risk-taking. However, data availability issues make it difficult to directly control for the various non-cash related components of CEO compensation. These data issues do not only apply to some European banks (which seldom disclose information such as strike prices and expiry dates which are essential to value CEO stock options), but also to U.S. banks. For instance, SEC rules only mandated the disclosure of CEO pensions and other forms of deferred compensation from spring 2007 (before that, the pension values of U.S. CEOs have to be approximated using elaborate actuarial methods as in Sundaram and Yermack, 2007).

In this section, we demonstrate that the positive relationship between distance to default and bonus payments we report above is not driven by other pay variables. For instance, if bonus payments were a substitute for other forms of pay, CEOs with higher bonus compensation may receive fewer options grants. Given CEO options are associated with higher bank risk-taking (Mehran and Rosenberg, 2007; Hagendorff and Vallascas, 2011), options acting as substitutes for bonuses could provide an explanation for our results.

We use a variety of tests to exclude this and similar explanations for our main finding that bonuses lower bank risk. Using ExecuComp (for U.S. banks) and annual reports (for European banks), we are able to determine for all banks if and when CEOs are given option grants. As a preliminary test for whether options and bonus payments are substitutes, we examine whether bonus compensation is lower when options are granted more frequently. We compare the values of the bonus variables in banks below and above the median of the sample distribution of the frequency at which CEOs are granted options. The results of this simple comparison show that bonuses and option grants act as complements rather than substitutes. In the group of banks above the median of the option distribution, bonuses make up 37% of total CEO cash compensation while this share drops to 29% in the group of banks below the median (difference is significant at 1%).

Further, we run several additional tests to control for the effect of compensation incentives other than bonuses. The results of these tests are reported in Table 6. First, we exclude banks that grant cash bonuses but no stock options to CEOs during the sample period. At a minimum, this exclusion should eliminate any effects stemming from banks employing cash bonuses as a substitute to stock options. Second, we exclude CEOs who do not hold options during the sample period. It is conceivable that these CEOs will attribute more value to cash bonuses than CEOs with more varied pay incentive structures. The results of these additional tests are reported in the first four columns of Table 6 and confirm our finding of a positive link between cash bonuses and distance to default.

Table 6 near here*

Third, in an effort to control for the role of options in shaping the incentive structure of CEOs, Columns 5 and 6 re-estimate the models with a binary variable (**OPTION**) which is equal to 1 if the CEO holds stock options at t-1 and has received an annual option grant at time$t$. This additional control variable, introduced as an endogenous regressor to our baseline specification, does not enter the model with a significant coefficient and, more importantly, we continue to report a positive link between cash bonuses and distance to default.

Finally, we follow a line of research that demonstrates that unobserved managerial heterogeneities are key to explaining the level and variation in executive compensation (Graham et al., 2012; Coles and Li, 2011). According to this work, director fixed effects can largely capture differences in director compensation including its non-cash components. For instance, Coles and Li (2011) report that director fixed effects and firm characteristics jointly explain over 60% of the sensitivity of managerial wealth to stock prices ('delta') and near 90% of the sensitivity of managerial wealth to stock volatility ('vega'). We re-estimate our models after replacing the bank fixed effects with CEO fixed effects. While CEO fixed effects can only capture the time-invariant dimension of unobserved CEO heterogeneity, CEO characteristics such as risk aversion, talent and reputation which are important determinants of CEO compensation are likely to be time-invariant in the short-term. In the last two columns of Table 6, we estimate the models with CEO fixed effects using CEOs with at least three consecutive years of data (yielding a total of 148 CEOs). The results show that we continue to observe a risk reducing effect of cash bonuses.[20]

The next two subsections examine settings in which we expect the risk-reducing effects of CEO bonus compensation to no longer hold. Specifically, we examine the impact of cash bonuses on default risk when banks are highly risky and when the bank regulatory environment is weak.

6.3 Bank Default Risk and the Risk-taking Incentives of Cash Bonus Compensation

Risk-taking incentives embedded in CEO compensation contracts are likely to emerge when banks are close to default. When a bank's growth prospects are poor, shareholders gain more from high-risk activities which maximize the value of the financial safety net than from attempts to increase low levels of charter value (Keeley, 1990; Park and Peristiani, 2007). At distressed banks, shareholder-designed compensation contracts will, therefore, promote risk-shifting to bondholders and regulators. Even though CEO cash bonus payments mitigate conflicts between shareholders and creditors and lead to less risky banks under normal conditions, we expect cash bonus payments do not exert a risk-reducing effect for risky banks and may even increase bank risk at distressed banks (see Noe et al., 1996).

Table 7 near here

To test these expectations, we introduce an interaction term between the lagged distance to default and each of our CEO bonus pay variables such that the target distance to default at time t ($\mathrm{DD}^*_{i,t}$) is now determined as follows:

$$DD^*_{i,t} = \gamma_1 BONUS_{i,t-1} + \gamma_2 BONUS_{i,t-1} \times DD_{i,t-1} + \beta \mathbf{X}_{i,t-1} + B_i \tag{7}$$

The variable descriptions are the same as before. We substitute Equation (7) into Equation (4) and report the findings in Table 7. The coefficients on the compensation variables (LGBONUS, RELBONUS) now capture the effect of CEO cash bonuses on the target level of bank risk when

[20] CEO fixed effects also capture individual characteristics (such as managerial talent) which may correlate with bonus payments. Therefore, it is not surprising that the coefficients on the bonus variables are slightly lower under these specifications compared with the models based on bank fixed effects.

DD_{t-1} equals zero (when a bank is financially distressed). The coefficients on the interaction terms between lagged DD and the bonus variables capture the effect of CEO cash bonus payments on the target level of risk when DD_{t-1} increases (when banks become less risky).

In all model specifications, the results are in line with our intuition. The coefficients on LGBONUS and RELBONUS are negative and statistically significant at customary levels. For cash bonuses, this means that the risk-reducing impact which we observed in the general sample gives way to a risk-increasing impact at highly-risky banks. We, therefore, conclude that when banks are financially-distressed, CEO bonus pay increases default risk.

Table 8 near here

Table 8 reports the short-term and long-term marginal effects for each of the bonus variables on DD. We distinguish between high-risk (banks located in the lowest 5% percentile of the sample distribution of DD), medium-risk (median DD), and low-risk banks (banks located in the top 5% percentile of the sample distribution of DD) and compute the marginal effects of bonus payments for each of the resulting groups. As previously, short-term marginal effects capture the change in observed default risk associated with a marginal change in the bonus variables, while long-term effects capture the effect that marginal changes in the bonus variable have on the unobservable target level of risk.

The results of Table 8 confirm the main finding in Table VII that the impact of CEO cash bonuses on managerial risk preferences is different for financially distressed banks compared with non-distressed banks. Both long- and the short-term marginal effects indicate that, at the least risky banks, CEO cash bonuses reduce risk, while cash bonuses are risk-increasing at the riskiest banks in our sample.

This is further demonstrated by Figure 2 which plots the long-term marginal effects of LGBONUS, RELBONUS for different values of DD (risk).[21] The Figure shows that the DD threshold at which the risk-shifting incentives exerted by LGBONUS and RELBONUS are completely constrained (i.e., where the marginal effect of bonuses on DD is ≥ 0) is equal to 2.667 and 3.414 standard deviations from the default point, respectively. However, if we stipulate that any negative

[21] To facilitate comparisons between the long-term effects associated with the different compensation variables in Figure 1, we have rescaled the long-term marginal effects so that DD=0 corresponds to -100.

impact of LGBONUS (RELBONUS) on DD is statistically significant (at the 5% level), the value declines to 1.64 (2.32) standard deviations. Notably, only about 9% (17%) of banks in our sample show DD values below this value. Therefore, Figure 2 demonstrates that any risk increasing effect of bonus payments is limited to the most risky banks in our sample which approach the default point.

6.4 Cash Bonus Compensation, Bank Default Risk and Regulatory Power

Our sample covers countries where supervisory regimes vary both in terms of the specific powers available to regulators as well as in terms of the enforcement of these powers. Noe et al. (1996) argue that the prospect of regulatory intervention by regulators will minimize the risk-shifting incentives posed by executive compensation contracts. More generally, Webb (2008) argues that in environments where regulatory scrutiny is higher, bank CEOs become more risk-averse and any potential risk-shifting incentives posed by executive compensation will be less effective. We, therefore, expect that stronger bank regulatory regimes constrain pay-induced incentives for bank CEOs to shift risk and that the risk-reducing effect of bank bonuses is not observable under weaker regulatory regimes.

Table 9 near here

We test if bank regulatory regimes moderate the effect of bonus pay on bank risk-taking by constructing two indicators of regulatory power. First, we use the Barth et al. (2004) database on financial regulation to build an index of regulatory power based on 14 regulatory powers.[22] The

[22] The index measures bank regulatory power as the equal-weighted sum (incl. sub-questions) of the following questions (yes=1; no=0): (1) Does the supervisory agency have the right to meet with external auditors to discuss reports without the approval of the bank? (2) Are the auditors required to communicate misconduct by managers/directors to the supervisory agency? (3) Can legal action against external auditors be taken by supervisors for negligence? (4) Can supervisors force banks to change internal organizational structure? (5) Are off-balance sheet items disclosed to supervisors? (6) Can the supervisory agency order directors/management to constitute provisions to cover actual/potential losses? (7) Can the supervisory agency suspend director's decision to distribute: a) Dividends? b) Bonuses? c) Management fees? (8) Can the supervisory agency supersede bank shareholder rights and declare a bank insolvent? (9) Does banking law allow supervisory agency to suspend some or all ownership rights of a problem bank? (10) Regarding bank restructuring and reorganization, can the supervisory agency or any other government agency do the

resulting index captures the extent to which the supervisory environment is sensitive to bank risk-taking, the breadth of disciplinary powers available to regulators, and how well these powers are enforced.[23] Second, because the regulatory index shows that the U.S. exhibits a stronger regulatory environment than Europe (the index value is 13 in the U.S., while the average score for European countries in our sample is 8.5), we construct a dummy variable to indicate if a bank is chartered in the U.S..[24] To avoid collinearity between country dummies and the bank fixed effects, we stipulate that there needs to be a minimum of three banks in a country when we include country dummies in the model. The target level of distance to default ($DD^*_{i,t}$) can now be expressed as follows.

$$DD^*_{i,t} = \gamma_1 BONUS_{i,t-1} + \gamma_2 BONUS_{i,t-1} \times REG + \beta \mathbf{X}_{i,t-1} + B_i \tag{8}$$

where $BONUS_{t-1}$ is the vector of bonus variables, REG is either the index of regulatory power (**REGULAT**) or the U.S. dummy (**US**), $\mathbf{X}_{i,t-1}$ the vector of bank characteristics, and B_i captures unobservable bank characteristics. When we substitute (8) into (4), the coefficient on the interaction term captures the differential impact of cash bonuses on the target level of bank risk in regulatory environments of varying power.

Table 9 presents our findings. In Panel A, the interaction terms between the bonus variables and the regulatory power variables are significant at customary levels in some specifications. The analysis of the marginal short-term (Panel B) and marginal long-term effects of compensation on DD (Panel C) estimates the marginal bank risk implications of CEO bonus pay linked to high and low levels of regulatory power. Regulatory power is either proxied by the maximum (13) or the

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following: a) suspend shareholder rights? b) remove and replace management? c) Remove and replace directors? We obtain updated values on regulatory variables from the Worldbank website (http://go.worldbank.org/SNUSW978P0).

[23] We estimate the models with the regulatory index only for countries for which we have at least three banks over the sample period. This choice ensures a minimum number of observations for each regulatory regime and reduces the risk that the regulatory effect is coincident with the bank fixed effect in countries with only one bank. However, if we conduct the analysis for the full sample of banks, this leads to similar conclusions.

[24] We accept that using a U.S. dummy to indicate a stricter regulatory regime is a basic test only. However, the European subsample is not large enough to implement the instrumental variable approach and to test our hypothesis regarding the effects of executive compensation under regulatory regimes of varying strength for the European sample only.

minimum (5) of the regulatory power index during the sample period (Columns [1], [3], [5], [7]) or, alternatively, by the U.S. dummy (Columns [2], [4], [6], [8]). Overall, Panels B and C show that only in stronger regulatory environments, bonus payments (both when expressed in absolute terms [LGBONUS] and in relative terms [RELBONUS]) lower managerial risk preferences (the marginal effects are significant below the 5%-level). Further, this effect can be observed both on a short-term and long-term basis (apart from the two regressions on RELBONUS which also include CEO fixed effects where the marginal effects are not significantly different from zero).

In sum, our analysis suggests that the effect of CEO cash bonus varies with the power of the regulatory regime under which banks operate. Overall, our findings are consistent with the view that CEOs respond to the risk-reducing incentives posed by higher bonus payments in the U.S. and in powerful regulatory regimes in Europe.

6.5 Additional Robustness Tests

Table 10 shows the results of further tests that we have conduced to assess the robustness of our results.[25] First, the recent financial crisis has affected the levels of default risk prevailing in the banking industry. During 2008, the average default risk in the banking industry was higher than during any year in our sample. It could be possible that these effects in the last year of our sample period influence our main conclusions. We, therefore, repeat the analysis after omitting 2008 from the sample period. The results of this additional analysis, presented in Column 1 and 2 of Table 10, demonstrate that, although the exclusion of 2008 reduces our sample size by around 15%, our main conclusions remain unchanged.

Table 10 near here

Next, we test the robustness of our results to changes in the sample composition. This is motivated by the fact that the European banks are on average larger than the U.S. banks in our sample. Specifically, we examine the influence of bank size on our findings by testing how our

[25] In the interest of brevity, Table 10 only reports the coefficients and significance levels for LGBONUS and RELBONUS as well as their interactions with distress risk and regulation. Furthermore, in this latter case, we only summarize the robustness tests when regulatory power is measured by REGULAT. The complete results are available from the authors upon request.

results change after excluding banks located in the lowest quartile of the sample distribution of total assets.[26] The exclusion of small banks leads to a sharp reduction in sample size to 88 banks (the total number of observations decreases to 604). The results of these tests (reported in Column 3 and 4 of Table 10) show that, despite this substantial reduction in the sample size, our main conclusions remain qualitatively unchanged.

Finally, our models have been estimated under the assumption that the bonus variables, LEVERAGE and ROA are endogenous variables, while the remaining bank variables are strictly exogenous. We re-run the analysis assuming that all bank control variables are endogenous and use only a single lag of these variables as instruments. As shown in Column 5 and 6 of Table 10, these changes to our specification do not affect our main findings.

Overall, the tests discussed in this section show that our findings on the nexus between bonus compensation and banking risk are robust to changes in the sample period, sample composition and model specification.

7 Conclusions and Policy Implications

In this paper, we contribute to the ongoing debate over the optimal design of CEO remuneration policies in the banking industry. We analyze the impact of CEO cash bonuses on the distance to default of a sample of U.S. and European banks. Our results show that banks where CEOs receive large bonus payments (both in absolute terms as well as relative to their total cash compensation) display lower levels of default risk. We interpret this finding as consistent with the shape of the CEO payoff function which underlies a typical cash bonus plan. Thus, CEO cash bonuses are solvency-contingent and, therefore, incentivize CEOs to avoid institutional failure. Additionally, the payoffs from cash bonus plans are typically linear functions of firm performance and are, therefore, not risk-rewarding for CEOs. Finally, our results show that the effect that bonuses exert on risk-taking is not confined to the short-term. Bonus payments also affect the long-term term level of default risk which bank CEOs target.

[26] Unfortunately, we are unable to re-estimate the model separately for each size-based sub-sample (defined, for instance, on the basis of the median value of the size distribution), because we need to maintain a sufficiently large number of observations to implement our instrumental variable approach.

Critically, our analysis indentifies two factors which moderate the link between CEO cash bonuses and bank risk. First, the overall riskiness of an institution determines the effectiveness of risk-taking incentives imbedded in CEO bonus contracts. We find that when banks are highly risky, CEO cash bonuses are associated with higher bank risk-taking. Second, we find that the risk-reducing effect of CEO cash bonus plans only holds when banks operate under powerful bank regulatory regimes (including the U.S.). We argue that CEO incentives to shift risk to regulators counterweigh the risk reducing effects which cash bonus plans exert on CEOs when regulators are not powerful and the prospect of regulatory intervention is low.

Several policy implications derive from our work. First, our results suggest that increased shareholder involvement in setting executive compensation, for instance by giving shareholders a vote on compensation policy ('say on pay') as championed by the Dodd-Frank Wall Street Reform and Consumer Protection Act (Section 951), is unlikely to mitigate risk-taking incentives in the banking industry. We show evidence of increased risk-taking when banks are risky and shareholders have the most to gain from increasing the value of the safety net. When banks are highly risky, regulatory interference (e.g. with the aim of restricting certain pay practices) is the preferred alternative to more shareholder-oriented pay setting.

Second, since bonuses are linked to higher bank risk-taking at the riskiest banks, risk shifting incentives at these institutions are only likely to be constrained if compensation practices more closely align the interests of managers and debt-holders. To achieve this, Bolton et al. (2010) suggest that CEO compensation in banking should be linked to CDS spreads on a bank's outstanding debt. Similarly, Edmans and Liu (2011) argue that increasing the share of deferred compensation and other debt-like forms of executive pay will turn executives into unsecured bondholders with a financial interest in the liquidation value of a firm and, thus, remedy incentives for CEOs to shift risk.

Finally, greater levels of disclosure of the terms under which CEO bonus plans are granted will clearly be beneficial. While our results convey a positive view of the role of CEO bonus payments in affecting bank risk, they can only be interpreted as first evidence. More details on the conditions of executive bonus plans (e.g., performance thresholds, caps, clawback provisions) are essential in order to better understand the link between bonus pay and bank risk. Future work should, hence, focus on the terms of executive bonus plans to derive bonus plan payoff functions for individual

CEOs. A better understanding of the shape of individual CEO bonus payoff functions will be a vital stepping stone in order to understand how bonus plans should be designed to mitigate risk-taking by bank CEOs.

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Table 1: Sample Overview

Panel A: Sample Distribution by Year

	No.	%
2000	82	9.16
2001	90	10.06
2002	98	10.95
2003	109	12.18
2004	117	13.07
2005	111	12.40
2006	108	12.07
2007	95	10.61
2008	85	9.50
Total	895	100.00

Panel B: Sample Distribution by Country

	# banks	# obs.	%banks	%obs.	Average Assets (mln US$)
Austria	1	5	0.85	0.56	
Belgium	1	8	0.85	0.89	
France	4	34	3.42	3.80	
Germany	3	17	2.56	1.90	
Ireland	3	26	2.56	2.91	
Italy	11	88	9.40	9.83	
Netherlands	4	27	3.42	3.02	
Spain	4	29	3.42	3.24	
UK	10	83	8.55	9.27	
Europe	*41*	*317*	*35.04*	*35.42*	*498,070.6*
U.S.	76	578	64.96	64.58	103,651.1
Total	117	895	100.00	100.00	243,350.5

Table 2: Descriptive Statistics: CEO Bonus Pay

The table reports summary statistics for remuneration variables based on a sample of 117 European and U.S. banks for the period 2000-2008. Total Bonus are total bonus payments during the reporting year. Total compensation is the sum of basic salary, bonus payments and other income. LGBONUS is the log transformation of 1+total bonus. RELBONUS are CEO bonus payments over total compensation (%). Small banks are banks located in the lowest quartile of the sample distribution of total assets. Sources: ExecuComp (U.S.) and bank annual reports (Europe).

		N	Mean	Median	St.Dev.	1 Pctile	99 Pctile
Panel A: Full Sample							
	Total bonus	895	1,234.45	460.47	2,468.79	0.00	13,000.00
	Total compensation	895	2,512.62	1,435.96	3,060.79	314.88	16,097.19
	LGBONUS	895	4.85	6.13	3.18	0.00	9.47
	RELBONUS (%)	895	32.73	35.40	25.97	0.00	90.04
Panel B For Europe and the U.S.							
Europe	Total bonus	317	1,278.66	631.83	1,756.77	0.00	8,533.69
	Total compensation	317	2,842.79	1,920.09	2,817.37	185.25	14,038.48
	LGBONUS	317	5.47	6.45	2.89	0.00	9.05
	RELBONUS (%)	317	33.62	34.77	22.54	0.00	75.00
U.S.	Total bonus	578	1,210.20	396.15	2,784.16	0.00	16,986.75
	Total compensation	578	2,331.54	1,274.06	3,174.20	359.03	18,677.29
	LGBONUS	578	4.52	5.98	3.18	0.00	9.74
	RELBONUS (%)	578	32.24	36.03	27.68	0.00	91.30
Panel C: Europe and the U.S. (excl. Small Banks)							
Europe	Total bonus	312	1,294.71	661.59	1,766.14	0.00	8,533.69
	Total compensation	312	2,876.25	1,937.10	2,827.27	185.25	14,038.48
	LGBONUS	312	5.47	6.50	2.91	0.00	9.05
	RELBONUS (%)	312	33.55	34.76	22.61	0.00	75.00
U.S.	Total bonus	380	1,659.90	610.99	3,331.67	0.00	18,484.41
	Total compensation	380	2,961.92	1,706.10	3,674.60	611.60	20,361.20
	LGBONUS	380	4.85	6.42	3.43	0.00	9.824
	RELBONUS (%)	380	35.37	40.82	29.08	0.00	91.45

Table 3: Distance to Default in banks with high and low CEO Bonus Pay

Panel A reports the mean and median of DD for banks that with low and high values of LGBONUS (based on the median of the sampling distribution). Panel B reports mean (median) DD for banks that show low and high values of RELBONUS (relative to the median of the sampling distribution). Column 3 shows the results of a *t*- (z-) test of equality in the mean (median) DD between the two groups.

Panel A: DD in banks with low and high values of LGBONUS			
	LOW LGBONUS	**HIGH LGBONUS**	**(2) *minus* (1)**
	(1)	**(2)**	**(3)**
Mean DD	3.754	4.417	0.663***
			(5.183)
Median DD	3.614	4.175	0.561***
			(5.084)
N	895	895	

Panel B: DD in banks with low and high values of RELBONUS			
	LOW RELBONUS	**HIGH RELBONUS**	**(2) *minus* (1)**
	(1)	**(2)**	**(3)**
Mean DD	3.719	4.452	0.733***
			(5.760)
Median DD	3.563	4.178	0.615***
			(5.677)
N	895	895	

* significant at 10%; ** significant at 5%; *** significant at 1%

Table 4: Summary Statistics and Correlations

Panel A of this table reports summary statistics for the variables employed in the analysis. Panel B shows pairwise correlation coefficients and the related significant values The data refer to a sample of 117 European and U.S. banks from 2000 – 2008. In European countries where two-tier board structures exist, DUALITY is defined as the CEO acting as chairman of the supervisory board.

Panel A: Descriptive Statistics

	Description	**N**	**Mean**	**Median**	**St.Dev.**	**1 Pctile**	**99 Pctile**
DD	Distance to default	895	4.085	3.920	1.939	-0.060	8.814
LGBONUS	Log (1+total bonus)	895	4.854	6.134	3.179	0.000	9.473
RELBONUS	Bonus over total compensation (%)	895	32.732	35.397	25.967	0.000	90.036
LEVERAGE	1- (Equity over total assets) (%)	895	92.062	92.269	3.730	79.405	97.854
ROA	Pre-tax profits over total assets (%)	895	1.352	1.386	1.214	-3.039	4.690
CHARTER	Market value of equity scaled by book value of equity	895	1.908	1.736	1.050	0.250	5.687
NONINTSHARE	Non interest income over total operating income	895	37.805	37.150	19.266	2.224	86.698
ASSETS	Total assets ($ million)	895	243,351	42,927	484,413	1,524	2,459,149
SIZE	Log of total assets ($ million)	895	17.744	17.575	1.854	14.237	21.623
INDEP	Number of independent directors over board size (%)	805	73.186	75.000	14.732	33.333	100.00
DUALITY	Binary variable which is 1 if the CEO also acts as chairman of the board	895	0.521	1.000	0.500	0.000	1.000
AGE	Log of CEO age (years)	876	4.021	4.043	0.126	3.664	4.277

Panel B: Correlation Matrix

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
(1) DD	1										
(2) LGBONUS	0.190***	1									
(3) RELBONUS	0.189***	0.912***	1								
(4) LEVERAGE	-0.122***	0.181***	0.149***	1							
(5) ROA	0.355***	0.270***	0.270***	-0.360***	1						
(6) CHARTER	0.195***	0.287***	0.326***	-0.031	0.525***	1					
(7) NONINTSHARE	0.067*	0.275***	0.319***	0.164***	0.149***	0.203***	1				
(8) SIZE	0.076*	0.223***	0.215***	0.490***	-0.210***	-0.221***	0.382***	1			
(9) INDEP	0.006	-0.078*	-0.036	0.125***	-0.172***	-0.179***	0.104**	0.189***	1		
(10) DUALITY	0.043	-0.012	0.081*	-0.251***	0.148***	0.162***	0.069*	-0.186***	-0.128***	1	
(11) AGE	0.070*	-0.044	0.005	-0.231***	0.056	-0.028	0.056	-0.042	0.048	0.291***	1

* significant at 10%; ** significant at 5%; *** significant at 1%

Table 5: Distance to Default and CEO Bonus Pay

Panel A reports regression results, obtained via the two step system GMM estimator proposed by Blundell and Bond (1998) of the impact of cash bonuses on distance to default (DD) with Windmeijer (2005) corrected standard errors in parentheses. The model controls for bank characteristics, time dummies and country dummies. LGBONUS measures log (1+cash bonus). RELBONUS is the ratio of cash bonuses to total cash compensation. LEVERAGE is equal to total liabilities over total assets, ROA is pre-tax profits scaled by total assets, CHARTER is the ratio between the market value of equity and the book value of equity. NONINTSHARE is the ratio of non interest income over total operating income. SIZE is the log of bank total assets. INDEP is the ratio of independent directors to board size. DUALITY is a dummy which equal to 1 if a top executives is also the chairman of the board at the end of the fiscal year. AGE is the log of CEO age. When country dummies are included, there need to be a minimum of three banks in a country. The difference in *J*-test assesses the validity of the specific instruments we employ for the bonus variables, EQUITY and ROA. Panel B shows the long-term marginal effects computed for the compensation variables. The level of significance has been derived using a non-linear Wald test with χ^2-statistics reported in parentheses. Panel C shows the results of a Wald test of equality between the long-term and the short-term effect. χ^2-Statistics are reported with the associated p-values in parentheses.

	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Panel A: Regression Results						
DD$_{t-1}$	**0.535*** **	**0.531*** **	**0.543*** **	**0.546*** **	**0.538*** **	**0.539*** **
	(9.06)	**(10.54)**	**(9.34)**	**(10.67)**	**(10.15)**	**(10.27)**
LGBONUS$_{t-1}$	**0.073*** **	**0.058*** **			**0.055*** **	
	(3.47)	**(2.83)**			**(2.71)**	
RELBONUS$_{t-1}$			**0.721*** **	**0.527** **		**0.568** **
			(2.89)	**(2.22)**		**(2.15)**
LEVERAGE$_{t-1}$	**-8.897*** **	**-10.246** **	**-7.524** **	**-10.278*** **	-2.289	-2.559
	(2.68)	**(2.31)**	**(2.35)**	**(2.72)**	(0.37)	(0.55)
ROA$_{t-1}$	21.032	**26.161* **	23.544	**27.375* **	19.085	28.603
	(1.27)	**(1.82)**	(1.37)	**(1.75)**	(0.73)	(1.10)
CHARTER$_{t-1}$	**-0.150* **	-0.123	**-0.175** **	-0.133	-0.126	-0.158
	(1.70)	(1.49)	**(1.98)**	(1.64)	(1.07)	(1.44)
NONINTSHARE$_{t-1}$	**-0.754*** **	**-0.802** **	**-0.828*** **	**-0.814** **	**-0.782** **	**-0.878** **
	(2.64)	**(2.47)**	**(2.81)**	**(2.52)**	**(2.40)**	**(2.41)**
SIZE$_{t-1}$	**0.131*** **	**0.134*** **	**0.131*** **	**0.138*** **	**0.112*** **	**0.112*** **
	(3.23)	**(3.79)**	**(3.14)**	**(4.26)**	**(2.78)**	**(3.20)**
INDEP$_{t-1}$					-0.122	-0.106
					(0.85)	(0.68)
DUALITY$_{t-1}$					0.332	0.365
					(0.84)	(0.95)
AGE$_{t-1}$					0.207	0.227
					(0.66)	(0.79)
Constant	**4.892* **	6.223	3.690	**6.185* **	-2.062	-1.798
	(1.88)	(1.61)	(1.46)	**(1.86)**	(0.34)	(0.40)
Observations	778	778	778	778	684	684
Number of banks	117	117	117	117	114	114
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Country dummies	No	Yes	No	Yes	Yes	Yes
m_2- pvalues	0.976	0.916	0.981	0.907	0.968	0.907
J-test pvalues	0.437	0.426	0.317	0.485	0.526	0.572
Difference-in J-test pvalues	0.427	0.392	0.297	0.453	0.575	1.000
Panel B: Long-term Marginal Effects						
LGBONUS$_{t-1}$	**0.157*** **	**0.124*** **			**0.119** **	
	(12.66)	**(7.88)**			**(6,63)**	
RELBONUS$_{t-1}$			**1.578*** **	**1.161** **		**1.232** **
			(8.13)	**(4.87)**		**(4.46)**
Panel C: H_0: Long-term Marginal Effects=Short-term Marginal Effects						
LGBONUS$_{t-1}$	**9.83*** **	**6.74*** **			**5.42** **	
	(0.000)	**(0.009)**			**(0.020)**	
RELBONUS$_{t-1}$			**6.63** **	**4.42** **		**3.98** **
			(0.010)	**(0.036)**		**(0.046)**

* significant at 10%; ** significant at 5%; *** significant at 1%

Table 6: Additional Analysis: Additional Pay Variables

The Table reports regression results, obtained via the two step system GMM estimator proposed by Blundell and Bond (1998) of the impact of cash bonuses on distance to default (DD) with Windmeijer (2005) corrected standard errors in parentheses. The model controls for CEO fixed effects, bank characteristics, time dummies and country dummies. LGBONUS measures log (1+cash bonus). RELBONUS is the ratio of cash bonuses to total cash compensation. OPTIONS is a dummy equal to 1 if a *t*-1 the CEO holds a portfolio of stock options and at time *t* has been granted with further stock options. LEVERAGE is equal to total liabilities over total assets, ROA is pre-tax profits scaled by total assets, CHARTER is the ratio between the market value of equity and the book value of equity. NONINTSHARE is the ratio of non interest income over total operating income. SIZE is the log of bank total assets. The level of significance has been derived using a non linear Wald test with $\chi 2$-statistics reported in parentheses. The difference in *J*-test assesses the validity of the specific instruments we employ for the bonus variables, EQUITY and ROA. Panel C shows the results of a Wald test of equality between the long-term and the short-term effect. χ^2-Statistics are reported with the associated p-values in parentheses.

	Excl. Banks without CEO options		Excl. CEOs without options		Controlling for CEO options		Controlling for CEO fixed effects	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Regression Results								
DD_{t-1}	0.490***	0.483***	0.499***	0.487***	0.491***	0.505***	0.444***	0.432***
	(8.03)	(8.27)	(8.37)	(8.81)	(8.20)	(9.36)	(6.57)	(6.25)
$LGBONUS_{t-1}$	0.065***		0.069***		0.095***		0.060**	
	(3.18)		(2.90)		(4.23)		(2.55)	
$RELBONUS_{t-1}$		0.715**		0.704**		1.080***		0.602**
		(2.55)		(2.46)		(3.93)		(2.07)
$OPTIONS_{t-1}$					-0.129	-0.154		
					(1.10)	(1.37)		
$LEVERAGE_{t-1}$	-7.464**	-6.791**	-10.388***	-9.964***	-9.272***	-6.881*	-12.383***	-11.713***
	(1.99)	(2.02)	(2.92)	(3.04)	(2.82)	(1.81)	(3.84)	(3.80)
ROA_{t-1}	45.283***	50.783***	22.912	26.468	27.921	25.147	22.143	27.126
	(3.76)	(3.22)	(1.46)	(1.60)	(1.48)	(1.13)	(1.26)	(1.51)
$CHARTER_{t-1}$	-0.245***	-0.279***	-0.143	-0.164*	-0.095	-0.135	-0.101	-0.111
	(3.16)	(3.04)	(1.53)	(1.86)	(0.94)	(1.06)	(1.01)	(1.07)
$NONINTSHARE_{t-1}$	-0.859**	-0.981***	-0.704**	-0.848***	-0.820**	-0.961***	-0.702**	-0.849**
	(2.46)	(3.00)	(2.19)	(2.85)	(2.29)	(2.58)	(2.02)	(2.48)
$SIZE_{t-1}$	0.153***	0.164***	0.155***	0.166***	0.149***	0.127***	0.180***	0.199***
	(3.02)	(3.67)	(3.64)	(3.83)	(3.22)	(2.70)	(3.90)	(4.35)
Constant	3.204	2.503	5.886**	5.444**	4.948*	3.329	7.513***	6.677***
	(1.14)	(0.98)	(2.09)	(2.14)	(1.89)	(1.10)	(2.78)	(2.58)
Observations	710	710	713	713	584	584	669	669
Number of Banks	106	106	107	107	117	117		
Number of CEOs							148	148
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country dummies	No	No	No	No	No	No	No	No
CEO fixed effects	No	No	No	No	No	No	Yes	Yes
m_2- pvalues	0.983	0.966	0.844	0.810	0,771	0.713	0.965	0.948
J-test – pvalues	0.646	0.622	0.517	0.553	0.850	0.847	0.275	0.360
Difference-in J-test pvalues	0.637	0.585	0.492	0.497	0.970	0.983	0.243	0.319
Panel B: Long-term Marginal Effects								
$LGBONUS_{t-1}$	0.128***		0.137***		0.186***		0.108**	
	(9.86)		(8.61)		(18.67)		(6.16)	
$RELBONUS_{t-1}$		1.385**		1.370**		2.181***		1.060**
		(6.43)		(6.28)		(14.17)		(4.04)
Panel C: H_0: Long-term Marginal Effects=Short-term Marginal Effects								
$LGBONUS_{t-1}$	7.44***		7.03***		12.45***		4.66**	
	(0.006)		(0.008)		(0.000)		(0.031)	
$RELBONUS_{t-1}$		5.40**		5.64**		10.01***		3.22*
		(0.020)		(0.018)		(0.001)		(0.072)

* significant at 10%; ** significant at 5%; *** significant at 1%

Table 7: Distance to Default, CEO Bonus Pay and Financial Distress

The table reports regression results, obtained via the two step system GMM estimator proposed by Blundell and Bond (1998), of the impact of cash bonuses on distance to default (DD) with Windmeijer (2005) corrected standard errors in parentheses. The model controls for bank characteristics, time dummies and country dummies. LGBONUS measures log (1+cash bonus). RELBONUS is the ratio of cash bonuses to total cash compensation. LEVERAGE is equal to total liabilities over total assets, ROA is pre-tax profits scaled by total assets, CHARTER is the ratio between the market value of equity and the book value of equity. NONINTSHARE is the ratio of non interest income over total operating income. SIZE is the log of bank total assets. The coefficients on the compensation variables capture the effect of CEO compensation on the target level of bank risk when DD_{t-1} equals zero (when the bank is financially distressed). The coefficients on the interaction terms between lagged DD and the compensation variables capture the effect of CEO compensation on the target level of risk when DD_{t-1} increases (when a bank becomes less risky). The difference in J-test assesses the validity of the specific instruments we employ for the bonus variables, EQUITY and ROA. When country dummies are included, there needs to be a minimum of three banks in a country.

	(1)	(2)	(3)	(4)	(5)	(6)
DD_{t-1}	**0.375*** **	**0.376*** **	**0.327*** **	**0.416*** **	**0.390*** **	**0.320*** **
	(6.07)	**(5.94)**	**(3.00)**	**(6.86)**	**(5.90)**	**(3.85)**
$LGBONUS_{t-1}$	**-0.106*** **	**-0.105*** **	**-0.144*** **			
	(3.24)	**(3.03)**	**(3.81)**			
$LGBONUS_{t-1}*DD_{t-1}$	**0.035*** **	**0.035*** **	**0.040*** **			
	(4.99)	**(4.70)**	**(4.33)**			
$RELBONUS_{t-1}$				**-1.519*** **	**-1.614*** **	**-1.870*** **
				(3.92)	**(3.64)**	**(4.35)**
$RELBONUS_{t-1}* DD_{t-1}$				**0.436*** **	**0.473*** **	**0.528*** **
				(5.35)	**(5.10)**	**(5.19)**
$LEVERAGE_{t-1}$	**-7.507** **	**-9.564** **	**-10.225*** **	**-7.049** **	**-12.121*** **	**-11.440*** **
	(2.02)	**(2.17)**	**(2.71)**	**(2.14)**	**(3.02)**	**(3.44)**
ROA_{t-1}	23.203	**29.458***	**24.290***	20.921	28.310	**23.922***
	(1.40)	**(1.73)**	**(1.72)**	(1.21)	(1.47)	**(1.83)**
$CHARTER_{t-1}$	-0.156	-0.141	-0.115	-0.125	-0.100	-0.091
	(1.64)	(1.52)	(1.45)	(1.39)	(1.09)	(1.24)
$NONINTSHARE_{t-1}$	**-0.695** **	**-0.806** **	**-0.603** **	**-0.760*** **	**-0.986*** **	**-0.671** **
	(2.30)	**(2.17)**	**(2.00)**	**(2.82)**	**(2.99)**	**(2.12)**
$SIZE_{t-1}$	**0.129*** **	**0.122*** **	**0.166*** **	**0.125** **	**0.139*** **	**0.176*** **
	(2.85)	**(2.88)**	**(3.17)**	**(2.55)**	**(3.28)**	**(3.50)**
Constant	4.365	**6.511***	**6.329** **	3.934	**8.576** **	**7.277*** **
	(1.48)	**(1.68)**	**(2.08)**	(1.54)	**(2.46)**	**669**
Observations	778	778	669	778	778	669
Number of Banks	117	117		117	117	
Number of CEOs			148			148
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Country dummies	No	Yes	No	No	Yes	No
CEO fixed effects	No	No	Yes	No	No	Yes
m_2- p-values	0.378	0.441	0.646	0.340	0.361	0.465
J-test – p-values	0.791	0.831	0.233	0.891	0.923	0.336
Difference-in J-test pvalues	0.773	0.815	0.214	0.886	0.916	0.315

* significant at 10%; ** significant at 5%; *** significant at 1%

Table 8: Distance to Default, CEO Cash Bonus Pay and Financial Distress: Short-term and Long-term Marginal Effects

Panel A shows the short-term marginal effects of the compensation measures on distance to default estimated in the regressions in Table VII for three types of banks: high-risk (5th percentile of the DD distribution), medium-risk (median DD) and low-risk (95th percentile of the DD distribution). Robust standard errors are reported in parenthesis. Panel B reports the long-term marginal effects computed as the ratio between the short-term effect and 1 minus the coefficient of lagged DD. A non linear Wald test is employed to asses the statistical significance of the long-term effect with χ^2-statistics reported in parentheses. Panel C shows the results on a Wald test of equality between the long-term and the short-term effect and reports the χ^2-statistics with the associated p-values in parentheses.

	(1)	(2)	(3)	(4)	(5)	(6)
	Panel A: Short term marginal effects					
HIGH risk banks	**-0.070****	**-0.070****	**-0.103*****	**-1.077*****	**-1.134*****	**-1.334*****
	(-2.58)	**(-2.43)**	**(-3.45)**	**(-3.32)**	**(-3.10)**	**(-3.90)**
MEDIAN	0.031	0.032	0.013	0.189	0.240	0.198
	(1.58)	(1.63)	(0.491)	(0.83)	(1.05)	(0.325)
LOW risk banks	**0.152*****	**0.151*****	**0.154*****	**1.695*****	**1.875*****	**2.023*****
	(4.74)	**(4.73)**	**(3.80)**	**(4.62)**	**(4.96)**	**(4.80)**
	Panel B: Long-term Marginal Effects					
HIGH risk banks	**-0.113****	**-0.111****	**-0.154*****	**-1.843*****	**-1.859*****	**-1.962*****
	(7.24)	**(6.63)**	**(14.89)**	**(12.32)**	**(11.93)**	**(15.48)**
MEDIAN	0.050	0.052	0.019	0.326	0.394	0.294
	(2.50)	(2.58)	(0.49)	(0.415)	(1.05)	(1.00)
LOW risk banks	**0.244*****	**0.246*****	**0.224*****	**2.907*****	**3.075*****	**2.979*****
	(26.53)	**(24.45)**	**(24.60)**	**(19.38)**	**(24.20)**	**(28,49)**
	Panel C: Long-term marginal effects = Short-term marginal effects					
HIGH risk banks	**6.22****	**6.03****	**5.07****	**9.56*****	**10.05*****	**6.18****
	(0.013)	**(0.014)**	**(0.02)**	**(0.002)**	**(0.001)**	**(0.013)**
MEDIAN	2.25	2.19	0.49	0.62	0.92	0.98
	(0.133)	(0.139)	(0.482)	(0.431)	(0.337)	(0.323)
LOW risk banks	**14.31*****	**12.53*****	**6.82*****	**10.56*****	**11.19*****	**8,56*****
	(0.000)	**(0.000)**	**(0.009)**	**(0.001)**	**(0.001)**	**(0.003)**

* significant at 10%; ** significant at 5%; *** significant at 1%

Table 9: Distance to Default, CEO Bonus Pay and Regulatory Power

Panel A reports regression results, obtained via the two step system GMM estimator proposed by Blundell and Bond (1998) of the impact of cash bonuses on distance to default (DD) with Windmeijer (2005) corrected standard errors in parentheses. The model controls for bank characteristics, time dummies and country dummies. LGBONUS measures log (1+cash bonus). RELBONUS is the ratio of cash bonuses to total cash compensation. LEVERAGE is equal to total liabilities over total assets, ROA is pre-tax profits scaled by total assets, CHARTER is the ratio between the market value of equity and the book value of equity. NONINTSHARE is the ratio of non interest income over total operating income. SIZE is the log of bank total assets. REGULAT is an index of regulatory power and US indicates that the bank is chartered in the U.S.. The difference in J-test assesses the validity of the specific instruments we employ for the bonus variables, EQUITY and ROA.Panel B (Panel C) shows short-term (long-term) marginal effects of executive compensation on default risk. In columns 1 and 3 LOW is the minimum value of the regulatory power index during the sample period (5), while HIGH the maximum value (13). In columns 2 and 4. LOW corresponds to a value of US equal to 0 while HIGH to a value equal to 1. When country dummies are included, there needs to be a minimum of three banks in a country. The level of significance has been derived using a non-linear Wald test with χ^2-statistics reported in parentheses. Panel D shows the results of a Wald test of equality between the long-term and the short-term effect. χ^2-Statistics are reported with the associated p-values in parentheses.

	LGBONUS				**RELBONUS**			
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**
				Panel A: Regression Results				
DD_{t-1}	**0.548*** **	**0.554*** **	**0.460*** **	**0.454*** **	**0.561*** **	**0.566*** **	**0.465*** **	**0.461*** **
	(10.82)	**(10.05)**	**(6.82)**	**(6.33)**	**(11.03)**	**(10.55)**	**(6.85)**	**(6.63)**
$LGBONUS_{t-1}$	-0.067	-0.041	0.105	0.004				
	(0.62)	(0.89)	(0.75)	(0.08)				
$LGBONUS_{t-1}$* $REGULAT_{t-1}$	0.012		-0.004					
	(1.35)		(0.31)					
$LGBONUS_{t-1}$* US		**0.121** **		0.055				
		(2.33)		(0.97)				
$RELBONUS_{t-1}$					-1.600	-0.754	0.392	-0.019
					(1.30)	(1.33)	(0.23)	(0.04)
$RELBONUS_{t-1}$* $REGULAT_{t-1}$					**0.198***		0.011	
					(1.96)		(0.08)	
$RELBONUS_{t-1}$* US						**1.671** **		0.585
						(2.57)		(0.99)
$LEVERAGE_{t-1}$	**-10.127*** **	**-9.226** **	**-14.669*** **	**-13.764*** **	**-9.240** **	**-10.865** **	**-13.903*** **	**-13.028*** **
	(2.88)	**(2.04)**	**(4.95)**	**(4.41)**	**(2.49)**	**(2.57)**	**(4.87)**	**(4.17)**
ROA_{t-1}	15.796	22.924	22.307	23.335	22.235	30.428	**29.070***	26.947
	(1.26)	(1.27)	(1.32)	(1.36)	(1.24)	(1.58)	**(1.75)**	(1.57)
$CHARTER_{t-1}$	-0.116	-0.114	-0.059	-0.042	-0.140	-0.126	-0.069	-0.056
	(1.52)	(1.29)	(0.65)	(0.47)	(1.49)	(1.34)	(0.82)	(0.61)
$NONINTSHARE_{t-1}$	**-0.792*** **	**-0.816** **	**-0.778** **	**-0.814** **	**-0.926** **	**-0.871** **	**-0.793** **	**-0.791** **
	(2.74)	**(2.56)**	**(2.17)**	**(2.12)**	**(2.48)**	**(2.39)**	**(2.41)**	**(2.16)**
$SIZE_{t-1}$	**0.118*** **	**0.133*** **	**0.167*** **	**0.169*** **	**0.128*** **	**0.137*** **	**0.175*** **	**0.160*** **
	(2.98)	**(3.63)**	**(4.12)**	**(4.35)**	**(3.17)**	**(3.38)**	**(4.38)**	**(4.01)**
REGULAT	**-0.103***		-0.056		**-0.115*** **		-0.085	
	(1.86)		(0.69)		**(2.66)**		(1.45)	
USA		**-0.717** **		-0.392		**-0.863*** **		-0.351
		(0.026)		(0.85)		**(2.65)**		(0.80)
Constant	**7.473** **	5.751	**10.350*** **	**9.208*** **	**6.693** **	**7.250** **	**9.891*** **	**8.632*** **
	(2.46)	(1.48)	**(3.57)**	**(3.32)**	**(2.12)**	**(2.03)**	**(3.63)**	**(3.10)**
Observations	767	778	660	669	767	778	660	669
Number of banks	115	117			115	117		
Number of CEOs			145	148			145	148
Bank Fixed effects	Yes	Yes	No	No	Yes	Yes	No	No
CEO Fixed effects	No	No	Yes	Yes	No	No	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country dummies	No	Yes	No	Yes	No	Yes	No	Yes
m_2- p-values	0.895	0.816	0.971	0.890	0.816	0.721	0.872	0.839
J-test – p-values	0.902	0.908	0.323	0.460	0.908	0.942	0.443	0.490
Difference-in J-test pvalues	0.910	0.907	0.311	0.450	0.907	0.936	0.427	0.480

* significant at 10%; ** significant at 5%; *** significant at 1%

Table 9 (*cont'd*)

	Panel B: Short-term marginal effects							
High regulatory power	**0.091*** **	**0.080*** **	**0.058** **	**0.059** **	**0.970*** **	**0.916*** **	0.529	0.566
	(3.79)	**(3.15)**	**(2.27)**	**(2.15)**	**(3.30)**	**(2.99)**	(1.64)	(1.63)
Low regulatory power	-0.006	-0.041	0.086	0.004	-0.612	-0.754	0.445	-0.019
	(0.09)	(0.89)	(1.04)	(0.08)	(-0.82)	(1.33)	(0.43)	(0.04)
	Panel C: Long-term Marginal Effects							
High regulatory power	**0.202*** **	**0.180*** **	**0.098** **	**0.108** **	**2.211*** **	**2.112*** **	1.000	1.050
	(13.37)	**(9.70)**	**(4.49)**	**(3.90)**	**(10.47)**	**(7.90)**	(2.50)	(2.47)
Low regulatory power	-0.013	-0.093	0.157	0.007	-1.395	-1.738	0.836	-0.035
	(0.01)	(0.73)	(1.13)	(0.01)	(0.63)	(1.58)	(0.19)	(0.00)
	Panel D: H0- Long term marginal effects = Short term marginal effects							
High regulatory power	**10.21*** **	**7.86*** **	**3.38* **	**2.94* **	**8.54*** **	**6.29** **	2.11	2.10
	(0.001)	**(0.005)**	**(0.067)**	**(0.086)**	**(0.003)**	**(0.012)**	(14.64)	(14.8)
Low regulatory power	0.01	0.68	1.14	0.01	0.59	1.41	0.20	0.00
	(0.928)	(0.410)	(0.285)	(0.934)	(0.441)	(0.236)	(0.655)	(0.966)

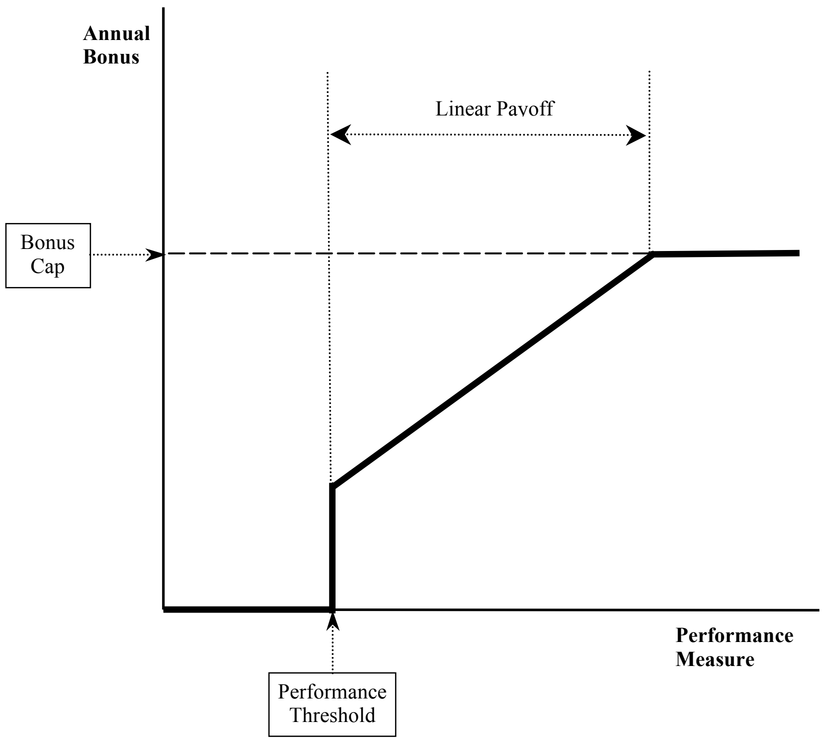
* significant at 10%; ** significant at 5%; *** significant at 1%

Table 10: Robustness Tests

Panel A reports the robustness tests for the models estimated in Table V, obtained via the two step system GMM estimator proposed by Blundell and Bond (1998) with Windmeijer (2005) corrected standard errors in parentheses. Panel B shows the robustness test for the models reported in Table VII while Panel C reports the robustness tests for the models in columns 1 and 3 of Table 9. All models control for bank characteristics, time dummies and country dummies. LGBONUS measures log (1+cash bonus). RELBONUS is the ratio of cash bonuses to total cash compensation.. REGULAT is an index of regulatory power. z statistics in parentheses* significant at 10%; ** significant at 5%; *** significant at 1%

	Excl. Financial Crisis		**Excl. Small Banks**		**All bank characteristics endogenous**	
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Panel A: Basic Specifications						
$LGBONUS_{t-1}$	**0.063*** **		**0.051** **		**0.060*** **	
	(2.97)		**(2.09)**		**(2.67)**	
$RELBONUS_{t-1}$		**0.639** **		**0.554** **		**0.632** **
		(2.36)		**(2.04)**		**(2.00)**
Observations	693	693	604	604	778	778
Number of Banks	117	117	88	88	117	117
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Country dummies	Yes	Yes	Yes	Yes	Yes	Yes
Panel B: Interaction between Bonus Compensation and Bank Distress						
$LGBONUS_{t-1}$	**-0.103** **		**-0.081* **		**-0.118*** **	
	(2.52)		**(1.85)**		**(2.84)**	
$LGBONUS_{t-1}*DD_{t-1}$	**0.035*** **		**0.029*** **		**0.039*** **	
	(4.36)		**(3.06)**		**(4.46)**	
$RELBONUS_{t-1}$		**-1.539*** **		**-1.272*** **		**-1.983*** **
		(3.13)		**(2.69)**		**(3.95)**
$RELBONUS_{t-1}* DD_{t-1}$		**0.459*** **		**0.408*** **		**0.587*** **
		(4.92)		**(3.76)**		**(5.41)**
Observations	693	693	604	604	778	778
Number of Banks	117	117	88	88	117	117
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Country dummies	Yes	Yes	Yes	Yes	Yes	Yes
Panel B: Interaction between Bonus Compensation and Regulation						
$LGBONUS_{t-1}$	-0.055		-0.062		-0.014	
	(0.42)		(0.49)		(0.10)	
$LGBONUS_{t-1}* REGULAT_{t-1}$	0.011		0.012		0.007	
	(1.04)		(1.08)		(0.66)	
$RELBONUS_{t-1}$		-2.309		-1.384		0.673
		(1.43)		(0.80)		(0.41)
$RELBONUS_{t-1}* REGULAT_{t-1}$		0.260**		0.196		0.015
		(1.96)		(1.28)		(0.12)
Observations	684	684	593	593	767	767
Number of Banks	115	115	86	86	115	115
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Country dummies	No	No	No	No	No	No
High regulatory power	**0.087*** **	**1.077*** **	**0.0918*** **	**1.166*** **	**0.083*** **	**0.872*** **
	(3.90)	**(3.56)**	**(3.14)**	**(2.62)**	**(3.52)**	**(2.88)**
Low regulatory power	-0.000	-1.006	-0.03	-0.403	0.023	0.749
	(0.01)	(1.04)	(0.04)	(0.41)	(0.29)	(0.74)

Figure 1: A 'Typical' Bonus Function



Adapted from Murphy (2000)

Figure 2: Distance to Default and the Long-Term Marginal Effects of CEO Bonus Pay Incentives