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Firm failure and relation lending: new evidence from small businesses

1. INTRODUCTION

Financial institutions are a key source of external funding for firms. In the US, borrowing from financial institutions accounts for about 25% of the stock of external finance of the productive sector (Mishkin, 2009) and for almost all the stock of external finance of small businesses (Bodenhorn, 2003). Moreover, as Hawkins (2002) points out, the financial system in emerging economies is centered on banks. For example, the ratio of domestic bank lending to the sum of domestic bank lending and private sector domestic debt securities on issue is around 0.8 for Latin American and East Asian economies. Thus, for small businesses in developed countries and for most firms in developing economies, banks loom large.

Bank relationships have long been valued in financial

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economics theory. In a world in which informational asymmetries exist and matter, banks gather valuable information on firms' creditworthiness through repeated lending. The establishment of long-term relationships allows banks to identify solvent firms and provide them with funding, reducing credit risk. Firms that form relationships with banks obtain several advantages, such as lower interest costs, lower collateral demands, and protection against credit rationing during times of financial distress. Therefore, both banks and firms seeking to finance profitable projects benefit from repeated contracting.

The evidence indicates that relations are favorable to firms in the sense that they increase funds availability and reduce loan rates (Elyasiani and Goldberg, 2004). There is also evidence indicating that the establishment of banking relationships may be more important for small firms which are typically more bank dependent than large firms (Fazzari et al., 1988; Dietsch, 2003; Guiso, 2003; Gómez-González and Reyes, 2010). However, while there is consensus in the literature about the importance of credit relationships, surprisingly there is no study on the effects of bank-firm liaisons on small firms' failure probability.

This study uses a uniquely rich data set comprised of information on individual loans of a large number of small firms in Colombia, and contributes to the literature in two important ways. First, it provides evidence that the establishment of one or more long-term relationships with banks reduces significantly the probability of default of small businesses. This is the first study in providing such evidence. And second, it shows that unlike large firms in which the number of banking relationships does not affect the probability of failure, the choice between unique and multiple relationships significantly affects the hazards of bankruptcy.

Section 2 is a literature review section. Section 3 makes a description of the data used in the empirical analysis. Section 4 presents the empirical model and estimation results. Section 5 concludes.

2. LITERATURE REVIEW

The existence of asymmetric information in financial markets makes bank lending unique. Various theoretical studies claim that financial intermediaries have a comparative advantage in the production of information about borrowers that leads to efficiency gains in the borrowing-lending process (e.g., Boyd and Prescott, 1986; Diamond, 1991). The establishment of client relationships may produce an important flow of information that facilitates the building of a bank-specific informational capital that mitigates informational problems in credit markets. Fama (1985) claims that these endogenous long-term relationships affect the ability firms have to raise capital. In particular, firms involved in a long-term relationship with a bank have a higher probability of raising the required funds to finance an investment opportunity than otherwise identical firms not involved in these kinds of relationships.[1]

Empirical studies have shown that relationship banking has a significant effect on the availability and quantity of credit, particularly for small businesses. To mention just a few examples, Petersen and Rajan (1994) use the 1987 National Survey of Small Business Finance and find that small firms increase their financing ability when they establish a long-term relationship with a bank. Berger and Udell (1996) arrive to the same conclusion using data from Federal Reserve's Survey of the Terms of Bank Lending to Business. Ely and Robinson (2009), using data from the Small Business Administration, report similar findings. Additionally, Boot and Thakor (1994) show that interest rates charged by banks to firms and collateral requirements decrease as the relationship matures.

Banks find it also convenient to establish credit relationships with firms, because their cost of maintaining long-term relationships is normally lower than the cost of repeated direct monitoring (Haubrich, 1989). The process of repeated contracting can produce valuable input for the lender in making decisions on loan issuing, loan pricing, and collateral requirements.

While there seems to be a consensus on the importance of

[1] In a world in which firms' repaying capacity in the future is not directly observable, the firm that obtains a new loan sends a signal to the market that its projects are worth of being financed, increasing its probability of raising the required funds. A firm that has a well established relationship with a bank has a better chance to obtain a new loan, and therefore the lender-borrower relationship can affect firms' ability of raising capital.

the establishment of long-term relationships between banks and non-financial firms, the optimal number of such relationships has been a subject of debate. Diamond (1984) shows that in a one-shot game each firm optimally chooses a unique banking relation. However, in a repeated lending setting the results are mixed. Some studies argue in favor of a unique long-term relation (Haubrich, 1989; Sharpe, 1990) while others argue in favor of multiple relations (Rajan, 1992; Detragiache et al. (2000); Von Thadden, 2004).

In the real world, unique long-term bank relations and multiple bank relations coexist (Degryse et al., 2009). Some recent studies have shown firm heterogeneity can explain this empirical regularity. Bolton and Scharfstein (1996) build a model in which the optimal number of banking relations depends on individual firm characteristics, such as technology, credit rating, and the industry to which the firm belongs. Bris and Welch (2005) show that if borrower quality is not known, the best firms choose to have a lower number of creditors to signal themselves as robust firms. Von Rheinbaben and Ruckes (2004), assume that revealing information can be costly for firms, and show that highly competitive and innovative firms tend to sustain a lower number of banking relations. Using similar arguments, Yosha (1995), and Bhattacharya and Chiesa (1995) obtain the same result.

There are no theoretical studies in which the firm's size matters for the optimal number of relationships. However, papers that have studied the microeconomic determinants of the optimal number of banking relationships using data from different countries have shown that small firms tend to maintain fewer bank relationships than large firms (Ongena and Smith, 2000; Dietsch, 2003; Guiso, 2003; Qian and Strahan, 2007; Gómez-González and Reyes, 2010). This empirical regularity may obey to the fact that while all types of banks lend to large firms, normally only small banks extend credit to small businesses (Berger et al., 1995; Peek and Rosengren, 1996; Strahan and Weston, 1998).

Two questions that have not been addressed in the literature appear: *i)* do banking relationships matter for small businesses survival?; and, *ii)* does the probability of surviving depend on the number of relationships? This study answers these two important questions. We work on two hypotheses:

- H_1: The establishment of at least one long-term banking relationship increases the probability of surviving for a small firm.
- H_2: Moreover, firms involved in multiple relationships have a lower hazard of failing than otherwise identical firms involved in unique long-term relationships.

The intuition behind H_1 is simple. On the one hand, in contrast to the "financial irrelevance theorem" of Modigliani and Miller (1958), which claims that the firm's value is independent of its capital structure, the related empirical literature has shown that the value of the firm depends on its ability to raise external funds (see, for instance, Makhija and Spiro, 2005; Maher et al., 2008; Chan et al., 2009, among others). On the other hand, it has long been recognized in the literature that the probability of failure of a firm depends heavily on the value of the firm (for a literature review see De Giuli et al., 2008). Therefore, the probability a firm fails depends on its ability to raise funds. Fazzari et al. (1988) show a financial hierarchy exists and depends on firm-specific characteristics. They also show that, unlike large firms which are able to raise funds in capital markets, small firms are bank-dependent. It is to expect, then, that small firms' survival depends on their capacity to establish long-term credit relations with commercial banks.

Regarding H_2, Detragiache et al. (2000) claim that relationship banks may be unable to continue funding projects due to internal problems, and firms holding a unique relationship with one of such banks may have to refinance with another bank. However, the latter does not know the quality of the project and may refuse to lend or charge high informational costs to the firm. In that scenario, the establishment of multiple banking relations can reduce the probability of an undesirable liquidation of the project. It is sensible to expect, then, that small firms with multiple relationship creditors have a higher probability of surviving than similar firms with only one committed credit provider.

Hypotheses H_1 and H_2 are formally tested in this study.

3. DESCRIPTION OF THE DATA

This study includes both microeconomic and macroeconomic covariates in order to explain the probability of failure of small businesses. Microeconomic information is collected from two different data sets. On the one hand, we use data reported by commercial banks to the Superintendencia Financiera de Colombia[2] in the Format 341. This data, which is collected in a quarterly basis, contains specific information about each loan, including identification of its holder, amount of the loan, ex ante interest rate, type of guarantee, and credit rating of the debtor. We use data on the number of banking relationships from December 1999 to December 2007.

On the other hand, this study uses use data on non-financial firms' balance sheets reported to the Superintendencia de Sociedades de Colombia.[3] With this information we calculate firm-specific financial ratios used as covariates in the empirical models estimated in this paper. We use data from December 1995 to December 2007. The two microeconomic data bases are matched using the firms' identification code.[4]

Macroeconomic information is collected from the National Department of Statistics of Colombia.

Our interest relies on the failure of small businesses. Therefore, two decisions regarding firm classification were required. First, we had to decide which definition of *failure* we were to follow. Second, we had to decide how to classify firms according to their size.

The definition of *failure* has been subject of debate. In this study we follow the most accepted view, consisting of a juridical definition in which failure is equivalent to bankruptcy[5] (see Charitou et al., 2004).

[2] The to Superintendencia Financiera de Colombia is the regulator of Colombia's financial sector.

[3] The superintendencia de Sociedades is the firm's/corporation's supervisory agency in Colombia.

[4] In Colombia, the firms' identification code is known as the *NIT*.

[5] The juridical definition of failure is arguably the most adequate because it provides an objective criterion that allows firms to be separated easily into two distinct populations.

Regarding size, firms were placed into three size groups based on asset size.[6] For the allocation of firms into these groups we followed the criterion established by the Law 905 of 2004, which considers small firms as those with total assets' value is less than 5,000 minimum monthly Colombian wages. Medium-size firms are those with total assets' value ranging between 5,001 and 30,000 minimum monthly Colombian wages. Large firms are those with total assets valued over 30,000 minimum monthly Colombian.

Table 1 shows the distribution of firms in the sample according to their size. Noteworthy, the proportion of small firms in the sample grew importantly, especially from 1999 on. An important number of small firms that reported information to the Superintendencia de Sociedades between 2000 and 2007 had not done so before. Therefore, the data set we use in this study exhibits left (and also right) censoring. In order to deal efficiently with this peculiarity we use survival analysis techniques in the empirical analysis of this study.

TABLE 1. DISTRIBUTION OF FIRMS ACCORDING TO THEIR SIZE, 1995-2007

	Firm size				*Proportion of small firms in the sample (%)*
Year	*Small*	*Medium*	*Large*	*Total*	
1995	822	5,472	2,990	9,284	8.85
1996	947	5,219	2,993	9,159	10.34
1997	1,271	5,260	3,071	9,602	13.24
1998	1,391	5,043	2,956	9,390	14.81
1999	2,024	4,863	2,868	9,755	20.75
2000	2,787	5,176	2,814	10,777	25.86
2001	2,513	4,833	2,780	10,126	24.82
1002	2,423	4,496	2,575	9,494	25.52
2003	2,392	4,439	2,626	9,457	25.29
2004	3.153	4,421	2,531	10.105	31.20
2005	10,859	5,987	2,882	19,728	55.04
2006	12,278	8,108	3,236	23,622	51.98
2007	10,706	7,630	3,398	21,734	49.26

Following previous studies (e.g., Shumway, 2001; Gómez-González and Hinojosa, 2010) and theoretical expectations, the following financial ratios were considered in the explanation of time to failure (table 2).

[6] Firm size could have been alternately measured as market value, sales or number of employees.

TABLE 2. DESCRIPTION OF FIRM-SPECIFIC COVARIATES INCLUDED IN THE EMPIRICAL MODEL

Variable	*Description*
Liquidity (LIQ)	(Current assets + long-term investments) / (current liabilities + long-term debt)
Leverage (LEV)	Total liabilities / Equity
Profitability (PROF)	ROA
Inefficiency (INEFF)	Operational costs / Total assets
Debt composition (COMP)	Short-term liabilities / Total liabilities
Relationships dummy (RELDUM)	Indicator function taking the value 1 if the firm has at least one banking relationship, and 0 otherwise
Number of relationships (NUMREL)	Number of banking relationships held by the firm

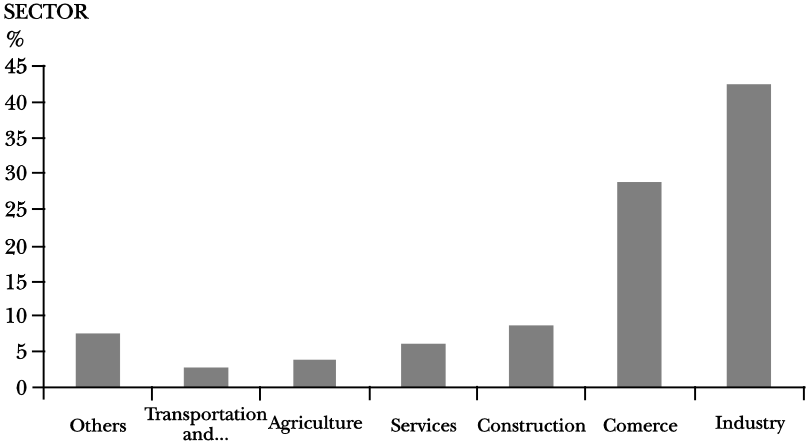
Two macroeconomic variables were also included: the annualized quarterly growth rate of real GDP (GROWTH) and the average real interest rate on loans (INT). These two variables were used to account for factors related to the business cycle that may excerpt influence on the business failure process.

Pair-wise correlations of the included covariates were small and in no case they exceeded 0.42 in absolute value.

More than 40% of the firms included in the sample belong to the industrial sector. The commercial sector is the second most representative of the sample (roughly 30% of the firms in the sample belong to this sector). See figure 1.

In this study we are particularly interested in the effect of relationship lending on the probability of failure of small

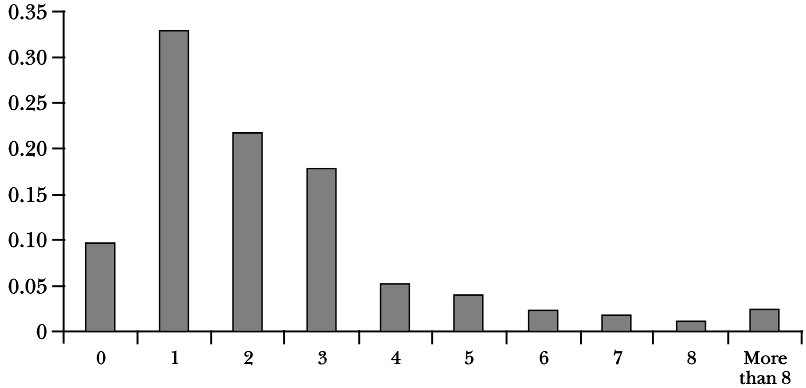
FIGURE 1. EMPIRICAL DISTRIBUTION OF FIRMS ACCORDING TO ECONOMIC SECTOR



firms. Relationship lending is defined as a long-term implicit contract between a bank and its debtor. We follow the empirical literature on relationship lending and use the duration of the bank-borrower relationship as a proxy for relationship lending (see, for instance, Petersen and Rajan, 1994; Berger and Udell, 1995; Ongena and Smith, 2001). We consider this an appropriate proxy, because duration reflects the degree of relationship intensity over time. Figure 2 shows a histogram of the number of relationships for the period 1995-2007.

As shown in figure 2, the density concentrates in a small number of relationships. The empirical probability of having at most two relationships is around 65%. In this sense, firms in Colombia behave similar to firms in the United Kingdom, Norway, and Sweden, which maintain fewer than three relationships on average, while contrast with firms in Italy, Portugal, and Spain, which maintain on average ten or more bank relationships (Degryse et al., 2009).

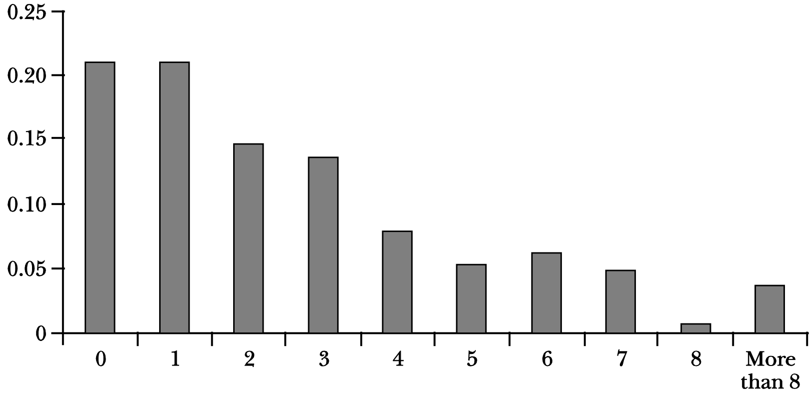
FIGURE 2. HISTOGRAM OF THE VARIABLE "NUMBER OF BANKING RELATIONSHIPS", 1995-2007



There is cross-sectional difference in the number of relationships. Smaller firms exhibit a less concentrated histogram than larger firms, and the empirical probability of having at least one but no more than three relationships is lower for the former than for the latter (figure 3).

Other interesting regularity is that the empirical probability of having more than one but less than three relationships is lower during economic expansions than during economic contractions. This fact provides evidence that the number of

FIGURE 3. HISTOGRAM OF THE VARIABLE "NUMBER OF BANKING RELATIONSHIPS" FOR SMALL FIRMS



bank relationships vary during the business cycle (Gómez-González and Reyes, 2010).

Table 3 shows descriptive statistics for the firms included in our sample. We have a total of 162,242 observations corresponding to 33,576 firms for the sample period. Out of these firms, 3,597 became bankrupt at some point between December 1995 and December 2007. Thus, the overall percentage of failures is 10.7%. The number of periods in risk corresponds to the number of quarters it takes a firm to fail after appearing for first time in the dataset. This time does not correspond necessarily to time to failure, because the data presents right censoring due to, for example, mergers and acquisitions.

TABLE 3. DESCRIPTIVE STATISTICS

Category	*Total*	*Mean*	*Minimum*	*Median*	*Maximum*
Time at risk (quarters)	–	19.328	4	12	52
Failures	3,597	0.107	0	0	1

4. EMPIRICAL MODEL AND ESTIMATION RESULTS

We use a hazard function model to study the time to failure of non-financial firms in Colombia. This approach generalizes the more common binary response approach by modeling not only the occurrence of the event but also the time it takes an individual to change of state (allowing a more efficient use of the available information). Hazard function models applied to this problem can provide answers to questions that are relevant for both supervisors and firm managers, such as: after the occurrence of a negative shock, what is the probability that a firm fails in the following months, given it has survived up to that moment? Or, what is the predicted time to failure for a firm of some given characteristics? In the context of this paper, using the hazard function model we are able to test empirically hypotheses H_1 and H_2 presented above.

Preliminary analysis on the raw data (data not conditioned on covariates) showed that the survival functions of firms belonging to different economic sectors are statistically identical. Therefore, we do not differentiate firms according to their economic sector. Similar tests showed, however, that small firms have a different survival function that the rest of the firms (table 4). Therefore, we estimated separate hazard functions for small firms and for medium and large firms.[7] Figure 4 shows the hazard function of failure of small firms.[8] This function is clearly non-monotonic, showing that the most commonly used parametric models for the distribution of duration do not seem to be appropriate for modeling the baseline hazard considered in this study.[9] Therefore, this paper estimates a proportional hazards model in which no parametric form is assumed for the baseline hazard function,

TABLE 4. TESTS FOR THE EQUALITY OF THE SURVIVOR FUNCTIONS
Null-hypothesis: Small firms and medium and large firms have identical survivor functions

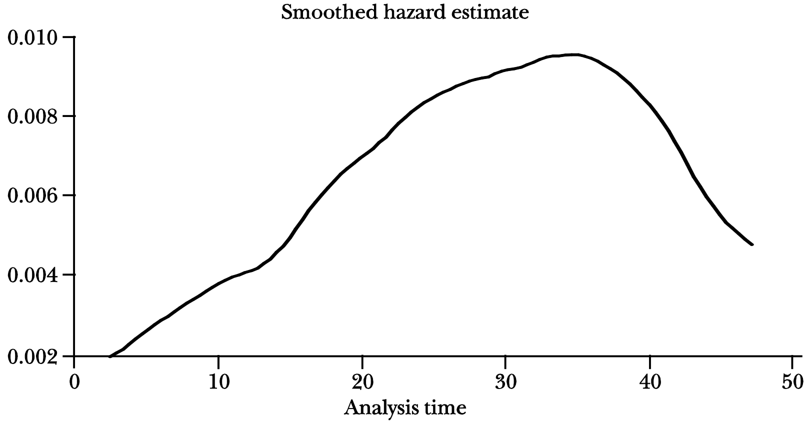
	Test for failure		
Test	*Log-rank*	*Cox*	*Wilcoxon*
χ^2 (1 d.f.)	7.02	7.03	7.12
Prob > χ^2	0.0081	0.0080	0.0076

[7] Firms of medium and large sizes have statistically identical survivor functions according to different tests. Thus, both groups were pooled into a unique group of firms.

[8] Hazard functions are estimated here as a kernel smoothed difference of the Aalen-Hansen estimator of the cumulative hazard functions. We use an asymmetric Epanechnikov kernel function.

[9] The hazard function of medium and large firms exhibited a non-monotonic behavior as well.

FIGURE 4. ESTIMATED SMOOTHED HAZARD FUNCTION – FAILURES OF SMALL FIRMS



following Cox (1972). As shown below using a specification test, this assumption seems to be appropriate for the problem of interest.

Under the proportional hazards specification the hazard rate takes the following multiplicative form (see, for example, Gómez-González and Hinojosa, 2010)

$$\lambda(t, X, \beta, \lambda_0) = \varphi[X(t), \beta]\lambda_0(t) \tag{1}$$

The baseline hazard function, $\lambda_0(t)$, captures the direct effect of time on the hazard rate. We specify $\varphi\ [X(t),\ \beta] = exp\ [X(t)'\ \beta]$, where $X(t)$ is a vector of time-varying covariates and β is the vector of parameters to be estimated. Under this specification, the coefficients can be given a partial derivative interpretation (Kiefer, 1988). In other words, each coefficient represents the constant, proportional effect of the corresponding covariate on the conditional probability of ending a spell.

We estimated the corresponding proportional hazards model using the partial maximum likelihood method proposed by Cox (1972).[10] Ties in duration are handled by applying the method of Efron (1977).[11] Two different models were

[10] The parametric models, estimated for comparison purposes, are estimated by maximum likelihood.

[11] Hertz-Picciotto and Rockhill (1997) show that the Efron method for handling ties is to be the preferred, especially when the sample size is small either due to heavy censoring or from the outset.

estimated. Specification 1 incorporates both macroeconomic covariates and all microeconomic covariates except NUMREL. Specification 2 incorporates the two macroeconomic variables and all microeconomic covariates except RELDUM. Estimation results are shown in table 5.

TABLE 5. ESTIMATION RESULTS FOR DURATION TO FAILURE (PROPORTIONAL HAZARDS SPECIFICATION)

	Specification 1		*Specification 2*	
Covariate	*Small firms*	*All other firms*	*Small firms*	*All other firms*
LIQ	−0.008	−0.051	−0.008	−0.050
	(0.012)	(0.32)	(0.012)	(0.31)
LEV	0.021	0.11	0.021	0.12
	(0.031)	(0.042)	(0.032)	(0.042)
PROF	−0.044[b]	−0.038[b]	−0.044[b]	−0.041[b]
	(0.023)	(0.017)	(0.023)	(0.019)
INEFF	0.003[a]	0.003[a]	0.003[a]	0.003[a]
	(0.002)	(0.002)	(0.002)	(0.002)
COMP	0.002[b]	(0.002)	(0.002)[b]	0.001
	(0.001)	(0.002)	(0.001)	(0.001)
RELDUM	−0.069[c]	−0.009[a]		
	(0.009)	(0.005)		
NUMREL			−0.017[b]	−0.002
			(0.009)	(0.004)
GROWTH	−0.010[b]	−0.011[b]	−0.010[b]	−0.011[b]
	(0.006)	(0.005)	(0.006)	(0.005)
INT	0.001	0.002	0.001	0.002
	(0.002)	(0.005)	(0.003)	(0.005)
Log-likelihood	−135,466.7	−15,323.5	−79,045.2	−16,457.1
LR χ^2 (8 d.f.)	9,429.5	1,315.0	6,282.7	1,848.1
Prob > χ^2	0.000	0.000	0.000	0.000

NOTES: Standard errors in parenthesis. Degrees of freedom in parenthesis.
[a] Significant at the 10% level. [b] Significant at the 5% level. [c] Significant at the 1% level.

Both specifications are globally significant for the two groups of firms. The results show that both firm-specific and business cycle variables are important determinants of time to failure of non-financial firms. For all firms, profitability, inefficiency, and the annualized quarterly growth rate of real GDP appear to be significant explanatory variables of the failure process. A one percentage point increase in profitability leads to a reduction in a firm's probability of failure of around 4%. A one percentage point increase in inefficiency leads to a reduction in a firm's probability of failure of around 0.3%. And a one percentage point increase in the annual growth rate leads to a reduction in a firm's probability of failure of around 1%.

Two other covariates are important explaining the failure process for small businesses but not for medium and large firms. A one percentage point increase in the ratio of short-term debt to total debt increases in 0.2% the hazard of failure of a small firm, but has no effect on the probability of failure of a medium or large firm.

Our more interesting result relates to the effect of relationship lending on the hazard of failure of firms of different size. According to the results of Specification 1, all else constant, a small firm without a partner bank will reduce in almost 7% its hazard rate of failure if it establishes a long-term relationship with a financial institution. In contrast, the effect of relationship lending on the hazard rate of failure of a large firm is much modest and less significant.

According to the results of Specification 2, an increase in the number of banking relationships benefits small firms while has no effect on larger firms. A small firm gains a reduction of 1.7% in its hazard rate of failure when increasing by one its number of relations. This effect is zero for a larger firm.

Altogether, the results show that relationship lending has an effect on firms' probability of failure. However, the effect is not homogeneous across firms of different sizes. The effect is much stronger and significant for small firms, which are more affected by the existence of asymmetric information in credit markets.

Table 6 presents evidence that the proportional hazards assumption is adequate for our sample. The proportional hazards factorization implies that the effect of the covariates on the hazard function is constant over time. This hypothesis can be tested using the Schoenfeld's residual test, which tests for a zero slope in a generalized linear regression of the residuals on time. The null hypothesis of the test is that the slope is zero. A rejection of the null hypothesis indicates that the proportional hazards assumption is unsuitable.

TABLE 6. SCHOENFELD'S RESIDUALS TEST RESULTS – SMALL FIRMS

	Specification 1			*Specification 2*		
Covariate	ρ	χ^2	$Pr > \chi^2$	ρ	χ^2	$Pr > \chi^2$
LIQ	0.248	1.77	0.1837	–0.177	1.41	0.2351
LEV	0.093	0.28	0.5995	–0.136	0.70	0.4045
PROF	0.019	0.01	0.9360	0.195	1.69	0.1932
INEFF	0.323	2.14	0.1434	0.112	0.55	0.4569
COMP	0.180	0.84	0.3582	0.089	0.35	0.5552
RELDUM	0.321	0.83	0.3623			
NUMREL				0.045	0.12	0.729
GROWTH	0.044	0.12	0.728	0.091	0.29	0.590
INT	0.123	0.78	0.3771	0.312	2.32	0.1277
Global test		9.87	0.2743		9.49	0.3027

Note that the null hypothesis of a zero slope cannot be rejected in any case. Therefore, the Schoenfeld's residuals test provides evidence that the proportional hazards specification is adequate in this study.

5. CONCLUSIONS

We study the effect of relationship lending on small firms' failure probability using a uniquely rich data set comprised of information on individual loans of a large number of small firms in Colombia. We control for firm-specific variables and find that small firms involved in long-term liaisons with commercial banks have a significantly lower probability of becoming bankrupt than otherwise identical firms not involved in a long-term credit relationship. We also find that small firms with multiple banking relationships face a lower failure hazard than otherwise identical firms involved in a unique long-term relationship. We thus find evidence that supports the validity of our working hypothesis H_1 and H_2.

We also show that while long-term relationships are beneficial for firms of all sizes, these are significantly more important for small businesses. These results are complementary to those of other studies that suggest that the establishment of banking relationships may be more important for small firms which are typically more bank dependent than large firms (Fazzari et al., 1988; Dietsch, 2003; Guiso, 2003; Gómez-González and Reyes, 2010). However, this study is the first to show that small firms with long-term credit relationships with financial institutions are less prone to fail than otherwise identical firms without these kinds of relationships.

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Stress tests for banking sector: a technical note

1. INTRODUCTION

Financial stability requires that financial institutions hold enough capital to protect against adverse events. In particular we consider those based on portfolio or default management. For that purpose stress tests could provide some light on the robustness of the system to tail-events that deteriorate the balance sheet of banks.

This paper provides a technical description of the models considered by the Banco Central de Chile for conducting stress tests to Chilean commercial banks. The framework is hybrid given that bottom-up and top-down approaches are mixed. First, market risk is computed at bank-level based on gap analysis. Second, credit risk is calculated by a dynamic

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model for the system as a whole. In a second stage this risk is distributed among banks.

The paper is organized as follows. Section 2 presents the market risk in which the model used for interest rate risk is discussed in detail. Section 3 discusses the credit risk model which is a non-lineal VAR between the system aggregates (loan loss provisions, credit growth, and write-offs) and the macroeconomics variables (output growth, short and long term interest rates, terms of trade, and unemployment). Finally, section 4 contains implementation of those risks and how the results are presented in the Financial Stability Report.

2. MARKET RISK

Market risk can be decomposed into two main sources: interest rate and exchange rate. In both cases there are several financial instruments that can reduce the risk associated to these variables. An exchange rate shock leads to an instantaneous event in which the assets or liabilities indexed in that currency change at the time of the shock having a well defined result in profits. In our case we consider two possible scenarios as shocks for the exchange rate risk. Both are computed as percentile values of the distribution of exchange rate variation with in fifteen working days. We consider that span since it is the observed frequency in which the weights in portfolios change significantly. The percentiles considered are the median for baseline scenario, which implies a variation of 2% of the exchange rate, while the risk scenario uses a tail-event (percentile 99) of 20% variation.

Interest rate risk is a little more complex since almost all the financial instruments provide cash flows in future periods, for that a movement in the shape of the yield curve generates many changes along the maturity structure.

One standard approach to deal with interest rate risk is using the Value at Risk (VaR) approach. In that case the interest rate risk of the portfolio is computed as the sensitivity of the portfolio to the volatility of the short term interest rate. This approach assumes that the yield curve is determined by the dynamic of the short term interest rate, and for that the only element that is necessary for the computation of the VaR is the duration of the portfolio.

Here we propose a more accurate approach improving the VaR in two ways: *i)* we consider each cash flow in the portfolio; and *ii)* we model the yield curve using factors. The second improvement relies on the model proposed by Nelson and Siegel (1987). This model is very popular among practitioners and central banking analysts. Even more, Christensen et al. (2009) provides a non-arbitrage proof for the existence of the model.

In our case, we follow a discrete-time version of the model, which is close related to Diebold and Li (2006) in which the parameters of the model are dynamic factors.[1] We note that standard factor models as Vasicek (1977) and Cox et al. (1985) rely on a single factor which is the short term interest rate. Extensions of these models to have several factors can be found in many papers and technical textbooks (Shreve, 2004). However, our goal is to provide a simple discrete-time version of the model that is similar to the VAR specification in time-series econometrics.

Because our goal is to compute the risk associated with the interest rate, we take a strong assumption which is the log pure expectation hypothesis. This assumption allows us to keep a simple relationship between the n maturity discount rate and the short term interest rate. It is important to note that by construction, the model provides a consistent relation between the yield curve and the forward curve. Using that, we are able to compute the two risks associated to the interest rate: valuation and repricing.

2.1 Building the dynamic Nelson-Siegel Model

In this section we provide a pure econometrical derivation of the dynamic Nelson-Siegel model. We consider Z_{nt} as the discount factor a time t for a payment at time $t + n$. In short, that interest rate is the yield of a n maturity zero-coupon bond.

Following Campbell et al. (1997), we consider the log yield defined as $z_{nt} = \log\,(1 + Z_{nt})$, and the validity of the log pure

[1] Alfaro (2011) shows that the Dynamic-Nelson-Siegel (DNS) model belongs to the affine class of models. Using the Stochastic Discount Factor and assuming that the Jensen terms are ignorable then the DNS is obtained following the procedure proposed by Campbell et al. (1997).

expectations hypothesis. The latter implies that the long term interest rates are equal to the average of the expected value of the short term interest rates:

$$z_{nt} = \frac{1}{n}\sum_{i=0}^{n-1} E_t\left(z_{1,t+i}\right). \tag{1}$$

Given (1) we need only to define the stochastic process for the short term interest rate, which in our case is z_{it}. This interest rate is defined over a set of factors which are serially correlated. In particular, we consider three factors collected in the vector Λ_t such as $z_{it} = b'\Lambda_t$ where

$$\Lambda_t \equiv \begin{pmatrix} \lambda_{1t} \\ \lambda_{2t} \\ \lambda_{3t} \end{pmatrix} = \begin{pmatrix} \phi_1 & 0 & 0 \\ 0 & \phi_2 & \theta \\ 0 & 0 & \phi_3 \end{pmatrix}\begin{pmatrix} \lambda_{1t-1} \\ \lambda_{2t-1} \\ \lambda_{3t-1} \end{pmatrix} + \begin{pmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \end{pmatrix},$$

and $b = (1, 1, 1)'$ or $b = (1, 1, 0)'$. The first setting implies three factors Vasicek (1977) model, meanwhile the second calibration is based on Balduzzi et al. (1998), which includes the case of the Nelson and Siegel (1987) model.

It is important to note that the transitional matrix is constrained for the first factor. We could easily remove this constraint but that implies several additional degrees of freedom. Still under $\theta \neq 0$ we have some degrees to calibrate the model.

Vasicek (1977) provides a single factor model that can be easily extended with many factors (Shreve, 2004). In particular under our setting $z_{it} = \lambda_{1t} + \lambda_{2t} + \lambda_{3t}$.

For simplicity we consider $\theta = 0$, using that we could compute the effect of each factor separately. For example, for any factor we have $E_t\left(\lambda_{t+i}\right) = \phi^i \lambda_t$. Similar results hold for the other two factors. In addition to that, we note that

$$S \equiv \sum_{i=0}^{n-1}\phi^i = \left(\frac{1-\phi^n}{1-\phi}\right) \quad \text{and} \quad \lim_{\phi\to 1} S = \lim_{\phi\to 1} n\phi^{n-1} = n,$$

where L'Hopital rule was used for the last term. As overall we can consider a model in which $\phi_1 = 1$ but both ϕ_2 and ϕ_3 are less than one. In that case the model for the n maturity log yield is as follows:[2]

[2] RiskAmerica is a Chilean center for financial research. They provide daily estimates of the yield curve using this model in a continuous-time

$$(2) \qquad z_{nt} = \lambda_{1t} + \frac{\lambda_{2t}}{n}\left(\frac{1-\phi_2^n}{1-\phi_2}\right) + \frac{\lambda_{3t}}{n}\left(\frac{1-\phi_3^n}{1-\phi_3}\right).$$

The use of a random walk component is for empirical reason only. We will apply the same criteria for the case of the Nelson-Siegel model.[3] Note that we can take the dynamic of the factor in short as $\Lambda_t = F\Lambda_{t-1} + U_t$, where F is the transitional matrix and U_t collects the error terms. We need F^i for which we have:

$$F_2 = \begin{pmatrix} \phi_1^2 & 0 & 0 \\ 0 & \phi_2^2 & \theta\left(\phi_2 + \phi_3\right) \\ 0 & 0 & \phi_3^2 \end{pmatrix} \text{ and } F_3 = \begin{pmatrix} \phi_1^3 & 0 & 0 \\ 0 & \phi_2^3 & \theta\left(\phi_2^2 + \phi_2\phi_3 + \phi_3^2\right) \\ 0 & 0 & \phi_3^3 \end{pmatrix}.$$

It is clear that the coefficient located in the second row third column is the relevant for our analysis. In particular that expression for F^4 is $F^4[2,\ 3] = \theta(\phi_2^3 + \phi_2^2\phi_3 + \phi_2\phi_3^2 + \phi_3^2) =$ $= \theta\sum_{k=0}^{3}\phi_2^{3-k}\phi_3^k$. From here it is simple to generalize the result as follow

$$\pi_i \equiv F^i[2,\ 3] = \theta\phi_2^{i-1}\sum_{k=0}^{i-1}\left(\frac{\phi_3}{\phi_2}\right)^k = \frac{\theta\phi_2^i}{\phi_2}\left[\frac{1-\left(\phi_3/\phi_2\right)^i}{1-\left(\phi_3/\phi_2\right)}\right] = \theta\left(\frac{\phi_2^i - \phi_3^i}{\phi_2 - \phi_3}\right).$$

Recalling that $z_{1t} = \lambda_{1t} + \lambda_{2t}$ we have $E_t\left(z_{1,t+i}\right) = b'E_t\left(\Lambda_{t+i}\right) =$ $= b'F^i\Lambda_t$. Noting that

$$b'F^i = \begin{pmatrix} 1 & 1 & 0 \end{pmatrix}\begin{pmatrix} \phi_1^i & 0 & 0 \\ 0 & \phi_2^i & \pi_i \\ 0 & 0 & \phi_3^1 \end{pmatrix} = \begin{pmatrix} \phi_1^i & \phi_2^i & \pi_i \end{pmatrix},$$

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setting. Following the results of Tapia (2008) is possible to match the parameters of that model with the discrete-time version presented here. For example the estimates for persistence of each factor are 0.9947, 0.6516, and 0.001, meanwhile the estimated volatilities are 0.00657, 0.01750, and 0.03423. With these results we note that the volatility of the third factor is above five times the volatility of the highly persistent factor, we conclude that the first factor is random walk, and the third one is just noise.

[3] It is clear that assuming that one factor is random walk implies that interest rates are non stationary. We use this assumption only to fit the empirical evidence.

we can compute a closed-form solution for z_{nt} using (1). As we discussed above we consider $\phi_1 = 1$ in order to accommodate the empirical findings. Again ϕ_2 and ϕ_3 are less than one for which we could find the coefficient for the second factor as we did it in the case of Vasicek model: $\left(1-\phi_2^n\right)/\left(1-\phi_2\right)$. In the case of the third factor we have

$$T \equiv \sum_{i=0}^{n-1} \pi_i = \frac{\theta}{\phi_2 - \phi_3} \sum_{i=0}^{n-1} \left(\phi_2^i - \phi_3^i\right) = \frac{\theta}{\phi_2 - \phi_3} \left[\frac{\left(1-\phi_2^n\right)}{\left(1-\phi_2\right)} - \frac{\left(1-\phi_3^n\right)}{\left(1-\phi_3\right)} \right].$$

This implies a n maturity interest rate as

$$z_{nt} = \lambda_{1t} + \frac{\lambda_{2t}}{n} \left(\frac{1-\phi_2^n}{1-\phi_2} \right) + \frac{\lambda_{3t}}{n} \left(\frac{\theta}{\phi_2 - \phi_3} \right) \left[\left(\frac{1-\phi_2^n}{1-\phi_2} \right) - \left(\frac{1-\phi_3^n}{1-\phi_3} \right) \right]. \tag{3}$$

The model (3) is an extension of Balduzzi et al. (1998), in which we add a random walk factor. Since $\left(1-\phi^n\right)/\left(1-\phi\right)$ is the coefficient of the factor in Vasicek model, then the last expression in the model above is a pseudo derivative of that coefficient.

Using previous results we can see that the discrete-time version of Nelson-Siegel is based on (3) in the case of $\phi_2 = \phi_3 = \phi$. We will take this duty by two steps. First we replace ϕ_3 by ϕ, having T as follows

$$T = \frac{\theta}{\phi_2 - \phi} \left[\frac{\left(1-\phi_2^n\right)}{\left(1-\phi_2\right)} - \frac{\left(1-\phi^n\right)}{\left(1-\phi\right)} \right].$$

Second we take $\phi_2 = \phi$ in limit using L'Hopital rule.

$$\begin{aligned}
\lim_{\phi_2 \to \phi} T &= \theta \lim_{\phi_2 \to \phi} \left[\frac{d}{d\phi_2} \left(\frac{1-\phi_2^n}{1-\phi_2} \right) \right] \\
&= \theta \lim_{\phi_2 \to \phi} \left[\frac{-n\phi_2^{n-1} + \left(1-\phi_2^n\right)}{\left(1-\phi_2\right)^2} \right] \\
&= \frac{\theta}{1-\phi} \left[\frac{\left(1-\phi^n\right)}{1-\phi} - n\phi^{n-1} \right].
\end{aligned}$$

Also, Nelson-Siegel model assumes that the second factor is a weighted average of its own past and the past of the tendency or the unobserved factor, which implies that $\theta = (1-\phi)$. With these assumptions we could write the n maturity interest rate in closed-form as:

$$z_{nt} = \lambda_{1t} + \frac{\lambda_{2t}}{n}\left(\frac{1-\phi^n}{1-\phi}\right) + \frac{\lambda_{3t}}{n}\left[\left(\frac{1-\phi^n}{1-\phi_2}\right) - n\phi^{n-1}\right].$$

We note that the model has a random walk factor. In practice, this means that the Nelson-Siegel assumes that interest rates for any maturity have unit root. This is empirically valid even when theoretically interest rate should be mean reverting. For the case of US, Hoti et al. (2009) provides newly evidence of the presence of unit roots for all US debt instruments studied: three and six months Treasury Bills; 1, 2, 3, 5, 7, and 10 years Treasury Notes, and corporate and mortgages bonds.

For a fixed ϕ the model (4) is linear in the factors which can be estimated by using cross-sectional regressions or the Kalman filter. Diebold and Li (2006) show minor differences of these two approaches for the case of DNS model. Based on that we estimate the factors (λs) using cross-sectional regression of the yield curve.

2.2 Stressing the Yield Curve

In this section we discuss about the factors of the Nelson-Siegel model and how they can be used for stressing the yield curve. We argue than only the second factor should be stressed in order to account for a temporary shock in the curve. In particular this is similar to stress the short term interest rate leading to a some sort of liquidity shock.

Recalling that Nelson-Siegel model can be replicated by factor model we could represent the model as follows

$$\begin{pmatrix} \lambda_{1t} \\ \lambda_{2t} \\ \lambda_{3t} \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 \\ 0 & \phi & 1-\phi \\ 0 & 0 & \phi \end{pmatrix} \begin{pmatrix} \lambda_{1t-1} \\ \lambda_{2t-1} \\ \lambda_{3t-1} \end{pmatrix} + \begin{pmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \end{pmatrix}.$$

Since the first factor is a random walk we note that $z_{\infty t} \equiv \lim_{n\to\infty} z_{nt} = \lambda_{1t}$ which means that the first factor estimates the very long term interest rate. Diebold and Li (2006) call it the level because the entire yield curve reverts toward that number.

When n = 1 we have $z_{it} = \lambda_{1t} + \lambda_{2t}$ which means that the second factor is an estimator of the difference between the short term and very long interest rates. This is the negative of the slope of the yield curve.

Finally, for the third factor we consider m such that $m/2 = \left(1-\phi^{m}\right)/\left(1-\phi\right)$, for a fixed ϕ. Using (4) we have $z_{mt} =$ $= \lambda_{1t} + \lambda_{2t}/2 + \left(1/2 - \phi^{m-1}\right)\lambda_{3t}$ then a linear combination provides an estimate of the third factor as $\lambda_{3t} = \left(2z_{mt} - z_{1t} - z_{\infty t}\right)/\left(1-2\phi^{m-1}\right)$.

It is interesting to note that m implies that the coefficient of the second factor should be equal to half of the maturity. Given that the coefficient is a function of ϕ we could interpret this results as persistence of the second factor. Diebold and Li (2006) call this factor as curvature, since it is estimated by a linear combination of a middle interest rate and the two extreme values of the curve. For US, the authors found $\phi = 0.97$ which implies m = 27.1, meanwhile in the case of Chile we have m = 16.5 using $\phi = 0.9$.

It should be noted that we have three possible shocks to accommodate a stressed yield curve. In particular shocking the error of the first factor provides a permanent effect on the very long term interest rate. This can be interpreted as well as a parallel shift of the entire curve. A shock in the second factor implies that the short term interest rate increases. This can be interpreted as a liquidity shock since the second factor is stationary. In other words, the shock will be diluted along the yield curve having no effect in the very long term interest rate. Finally, a shock in the third factor affects the speed in which the interest rate converges to its long term value.

From the discussion above we conclude that shocking the second factor is more suitable for the purpose of stress test. In particular, this shock puts a liquidity risk into the portfolio analysis. We agree that moving the third factor could be also interesting given that this will affect more to there pricing risk than the valuation. This, because the first is associated with the renegotiation of current assets and liabilities.

Given that we compute monthly yield curves from October 2002 to May 2009, getting from there estimates of the factors λs.

From table 1 we estimate on the average a 5% long term interest rate and 1.71% of short term interest rate. In the case of the error terms associated with the dynamic factor model we define σ_k as the standard deviation of e_k. From the data we could get an estimator for the standard deviation of the error terms $\hat{\sigma}_1 = 0.005467$, $\hat{\sigma}_2 = 0.012385$, and $\hat{\sigma}_3 = 0.0087042$.

TABLE 1. DESCRIPTIVE STATISTICS OF YIELD CURVE FACTORS

	Average	*Standard deviation*	*Minimum*	*Maximum*
λ_1	0.048948	0.016090	0.022055	0.082500
λ_2	−0.033626	0.028709	−0.099340	0.030162
λ_3	−0.013621	0.016450	−0.053356	0.028337

In this sense a shock of 250 basis points to the second factor is equivalent to two standard deviation of this factor. This means a monthly event with probability 2.3% which is also consistent with the empirical movement in the yield curve observed when Lehman Brothers filed for chapter eleven.

Based on a stressed yield curve $\left(z_{nt}^*\right)$ we need an algorithm for both type of risk: valuation and repricing. We consider the effect of a change in the value of the portfolio using the yield curve, meanwhile for repricing we use the consistent forward curve for changes in the financing across time.

For the first risk we take $\alpha_{jt} = \exp\left(-z_{jt} j\right) - \exp\left(-z_{jt}^* j\right)$ as the change in the discount factors, and $N_{jt} = \left(A_{jt} - L_{jt}\right)$ as the net cash-flow to be received in time $t + j$. The total effect is $V = \sum_{j=1}^{n} \alpha_j N_{jt}$ practice the information of the cash flows is available only in ranges of maturities. In that case we replace the j by the midpoint of each range.

For repricing $\beta_{it} = \exp\left[E_t\left(z_{1,t+i}\right)\right] - \exp\left[E_t\left(z_{1,t+i}^*\right)\right]$ is the change in the one period forward rate for the time $t + i$. It is important to note that the expected value of the short term interest rate is the forward rate under the log pure expectations hypothesis. Moreover, under the Nelson-Siegel model we have an explicit expression: $E_t\left(z_{1,t+i}^*\right) = \lambda_{1t} + \phi^i \lambda_{2t} + (1 - \phi) i \phi^{i-1} \lambda_{3t}$. In addition to that we should note that risk is based on the rolling of the cash-flows for that we use $\Gamma_j = \sum_{i=j}^{n} \beta_{it}$ such that the total effect is $R = \sum_{j=1}^{n} \Gamma_j N_{jt}$. In the case that cash flows are collapsed in ranges we need to adjust Γ_j appropriately.

3. CREDIT RISK

By law, banks are required to have a minimum level of capital in order to cover default or credit risk. Indirect measures of defaults are available in the market, such as ratings or credit spread. However, the business of a representative bank involves several sectors and exposures which makes it hard to standardize a default measure in order to evaluate risk for the whole system. Moreover, many business activities do not have appropriate default measures forcing the researcher to make ad hoc assumptions in both levels: banks and the system.

In this section we choose an aggregate model in which the banking system is taken as a whole. We consider that the model is good enough for policy makers in the sense that captures the relation between macroeconomics variables and banking system aggregates such as loans, written-downs, and provisions. Given that aggregate risk is directly affected by macroeconomics variables through a non-linear system, our model could be considered as a non-linear vector autoregressive model (VAR).

3.1 Models and Assumptions

A standard measure in the industry of risk exposure is the ratio loan loss provision stock (S) over total loans (L). This measure –usually reported by loans categories– is considered a good estimate of expected loss. However, statistical properties of this measure, and contagion channels are not provided in financial reports.

In order to understand the relation between these components we define their dynamics.

First, the dynamics of loan scan be decomposed into previous loans, plus new ones, minus paid, and minus no-paid loans (write-offs): $L_t = L_{t-1} + L_{n,t} - P_t - W_t$. Also, the dynamic of loan loss provision stock can be written as previous stock, plus provision expenditure, minus write-downs, and minus recovered write-off loans: $S_t = S_{t-1} + E_t - W_t - R_t$. We note that these equations are simple accounting relationships. A structural equation requires the following assumption:

Assumption 3.1. The stock of loan loss provisions is proportional to the level of loans, and recovered write-offs are zero.

We denote the first statement as $S_t = \varphi_t L_t$, where factor φ means systemic risk. Later, we make further assumptions on this factor relating it with macroeconomics variables.[4] The second statement is the assumption used to simplify the problem.

Combining the new equations with the accounting relations we can write provision expenditure as follows: $E_t = W_t + \varphi_t \Delta L_t + L_{t-1}(\varphi_t - \varphi_{t-1})$. Based on this equation we want to have a measurement of risk. Following Matus (2007), we will focus our analysis in the following variables: $X_t \equiv \sum_{i=0}^{n-1} E_{t-i} / L_t$ and $Y_t \equiv (E_t - W_t) / L_{t-1}$. Statistical properties of these variables rely on the distribution assumption of loans, write-offs, and systemic risk. In practice Y_t is stationary but X_t is not. That could be explained by the low variation of write-offs related to the credit growth such that the persistence on L implies that X_t is also persistent.[5]

Assumption 3.2. The factor φ_t is a linear function of economic variables. However, those variables are known with lags.

This assumption implies $\varphi_t = \alpha + z'_{t-1}\beta$, where z stands for a vector of macroeconomic variables. Note that φ_t has two effects in loan loss provisions: *1)* direct, as a change in the macroeconomics variables; and *2)* indirect, as a factor of credit growth.

3.2 Empirical Application

We use balance sheet information for banks available at the

[4] The ratio between stock of provisions over total loans is a widespread metrics of risk in the industry.

[5] Alfaro et al. (2009) provides a theoretical result for this finding which proves that X_t is a martingale meanwhile Y_t is centered at zero. The result is based on a simple binomial model for loans and keeping other variables fixed; however the authors also provide empirical support showing that unit root tests cannot be rejected for X_t meanwhile for Y_t the results are mix but biased toward stationarity.

Chilean Superintendency of Banks and Financial Institutions (SBIF). We have monthly data from 1997 to 2010 by type of loans: consumer, commercial, and mortgage. The series, used at the system level, are loan loss provisions, write-offs and loans.

As it was discussed in the previous section we use expenditure which is proportional to the change in the stock. The dependent variable is the difference between provision expenditure and write-offs, that could be considered as a measure of the difference between expected and realized credit risk. It should be noted that provision expenditure should increase sufficiently the stock of provision to satisfy the regulatory requirements of provisions. The latter are computed based on the portfolio quality of each bank using statistical models.

On the exogenous variables we have a monthly measure of output growth which is the annual change of Economic Activity Index (IMACEC). Alfaro et al. (2009) use a monthly measure of output gap having similar results than using output growth. However the use of the former variable for forecasting implies additional assumptions on future path of both the trend and cycle of the output growth. A second key variable is the short-term interest rate (STIR). For this purpose we use an indexed rate for loans with maturity between one and three years. This interest rate is a good measure of funding cost for commercial business and also reflects the dynamics of that part of the yield curve. The unemployment rate has additional information relative to the output growth and interest rate and accounts for uncertainty related to future income. The benchmark model considers these three variables. However, the current state of analysis includes two additional variables that improves the fit of the equations: terms of trade and long-term interest rate. Terms of trade (TOT) is used in commercial aggregates as well for improving the fit of the rate of unemployment equation. The variable is generated as the ratio between the price of copper and the price of oil. Finally, long-term interest rate (LTIR) is measured by mortgage interest rate and it has a marginal impact on the growth of this kind of loans.

The forecast of the macroeconomic variables is obtained from the dynamic model, which is a structural VAR. It should be noted that banking aggregates are not included in the dynamic specification of macroeconomic variables. Therefore, the macroeconomic module is exogenous for the financial module. As the forecasts could be different than the one published in the latest Monetary Policy Report of the Central Bank the model is calibrated under the baseline scenario in order to match the dynamic of the output growth implicit in that report.

Based on the latest estimation of the model we consider the sample from January 1997 to August 2010. Tables 2 to 4 show the main results for the model. For the case of provisions we can see that the signs of the macroeconomics variables are as expected, and also the overall fit is about 50%. The short-term interest rate has a positive impact meanwhile output growth has a negative one. It is interesting to note that interactions between credit growth and macroeconomics variables are significant as we stated in the theoretical framework. This finding will be relevant for the simulation part as non-linear effects will be generated with this interactions.

TABLE 2. RESULTS FOR PROVISION MODEL

	Consumption	*Commercial*	*Mortgage*
Lag	0.2577		
	(0.1040)		
Credit growth (CG)	–0.0881	0.0483	
	(0.0280)	(0.0295)	
IMACEC		–0.0036	–0.0019
		(0.0014)	(0.0005)
STIR	0.0129	0.0097	
	(0.0028)	(0.0010)	
TOT		–0.0008	
		(0.0004)	
CG*IMACEC		–0.3097	
		(0.1984)	
CG*STIR	1.4590		
	(0.4291)		
CG*Unemployment		–0.4287	–0.0685
		(0.2524)	(0.0302)
R^2	0.501	0.405	0.384
Observations	164	164	164

NOTES: Robust standard errors in parentheses. Dummies are not reported.

TABLE 3. RESULTS FOR WRITE-OFF MODEL

	Consumption	*Commercial*	*Mortgage*
Lag 1	0.4699 (0.0787)	0.1808 (0.0689)	0.1806 (0.0925)
Lag 2			0.2717 (0.0740)
IMACEC	−0.0114 (0.0051)	0.0026 (0.0010)	
STIR		0.0027 (0.0011)	
Unemployment		0.0122 (0.0017)	0.0012 (0.0002)
CG*IMACEC	−0.5275 (0.3045)		
R^2	0.353	0.397	0.648
Observations	164	163	162

NOTES: Robust standard errors in parentheses. Dummies are not reported.

TABLE 4. RESULTS FOR CREDIT GROWTH MODEL

	Consumption	*Commercial*	*Mortgage*
Lag 1	0.4063 (0.0624)	0.1376 (0.1039)	
Lag 3	0.1276 (0.0445)	0.0548 (0.0659)	
Lag 6			0.3286 (0.0746)
IMACEC	0.0885 (0.0256)	0.0877 (0.0221)	
IMACEC (−1)	−0.0525 (0.0253)		
STIR	−0.0720 (0.0288)		
LTIR			−0.0794 (0.0322)
Write offs	−0.5385 (0.2461)	−4.1385 (2.2898)	
R^2	0.604	0.187	0.201
Observations	164	163	164

NOTES: Robust standard errors in parentheses. Dummies are not reported.

The auxiliary equations for write-offs and credit growth show the persistent of those aggregates and also the relationship with macroeconomics variables included in the analysis.

The fit for consumer credit growth seems reasonable, but most of them, is due to the seasonal effects and the persistence of the variable. Indeed, removing seasonal dummies the fit is 49% and removing this and the lag structure the fit is only 30%. In the cases of commercial and mortgage credit growth we can see that the fit is relative small (around 20%). This will lead to the fact that in simulations we will have large confidence intervals. Those could overlap under different scenarios. In particular that could imply no difference between baseline and risk scenarios. This problem does not happen in our analysis given that risk scenario is constructed on tail-events and the computed confidence interval does not overlap.

The macroeconomic model is reported in table 5, there we can see that the sign of the variable are as expected and the overall fit is reasonable for the sample size. It is important to stress that this model is considered only for the dynamic of macroeconomics variables and for that it does not represent the core model used by the bank for conducting monetary policy. The latter is based on several dynamics models as VAR and DSGE, and the macroeconomics model presented here is calibrated to match the dynamic published as baseline scenario in the Monetary Policy Report.

Following the steps summarized above in this section we are able to forecast loan loss provisions and credit growth for consumer, commercial and mortgage loans (see figure 1). Passing those forecasts to the bank level implies to use the current level of loan loss provision and loans as the initial value of the forecast of each bank. For example a particular bank has total loans for L_t with an initial composition of 70% on consumer, 20% on commercial, and 10% on mortgage. Given that the forecast for L_{t+i} will be the weighted average of the forecasts for credit growth of each type of loan. In addition to that the cumulated expenditures between t and $t+i$ are subtracted from earnings which implies that both the return on equity (ROE) and the capital adequacy ratio (CAR) moves under the risk scenario. At this point is relevant to note that the model always implies a reduction on ROE but the final effect on CAR depends on the initial position of the bank. Indeed, the data shows that under the Asian crisis in 1998 the CAR for the system exhibited

TABLE 5. RESULTS FOR MACROECONOMIC MODEL

	IMACEC	*STIR*	*LTIR*	*Unemployment (rate)*	*TOT*
Lag 1	0.9454	0.9020	0.7784	1.2718	0.0180
	(0.0375)	(0.0558)	(0.0553)	(0.1064)	(0.0580)
Lag 2		−0.1825		−0.2503	
		(0.0948)		(0.1656)	
Lag 3		0.1721		−0.1143	
		(0.0782)		(0.3060)	
Lag 4				0.0492	
				(0.1720)	
Lag 6	−0.0996				
	(0.0404)				
MA 8					−0.2234
					(0.0667)
MA 12			0.2020		
			(0.0783)		
MA 14					−0.2054
					(0.0679)
IMACEC				−0.0145	
				(0.0127)	
IMACEC (−1)				−0.0100	
				(0.0160)	
IMACEC (−2)				−0.0078	
				(0.0183)	
STIR	−0.0821				
	(0.0456)				
TOT	0.0160			−0.0041	
	(0.0084)			(0.0021)	
R^2	0.842	0.864	0.641	0.978	0.081
Observations	164	164	164	164	164

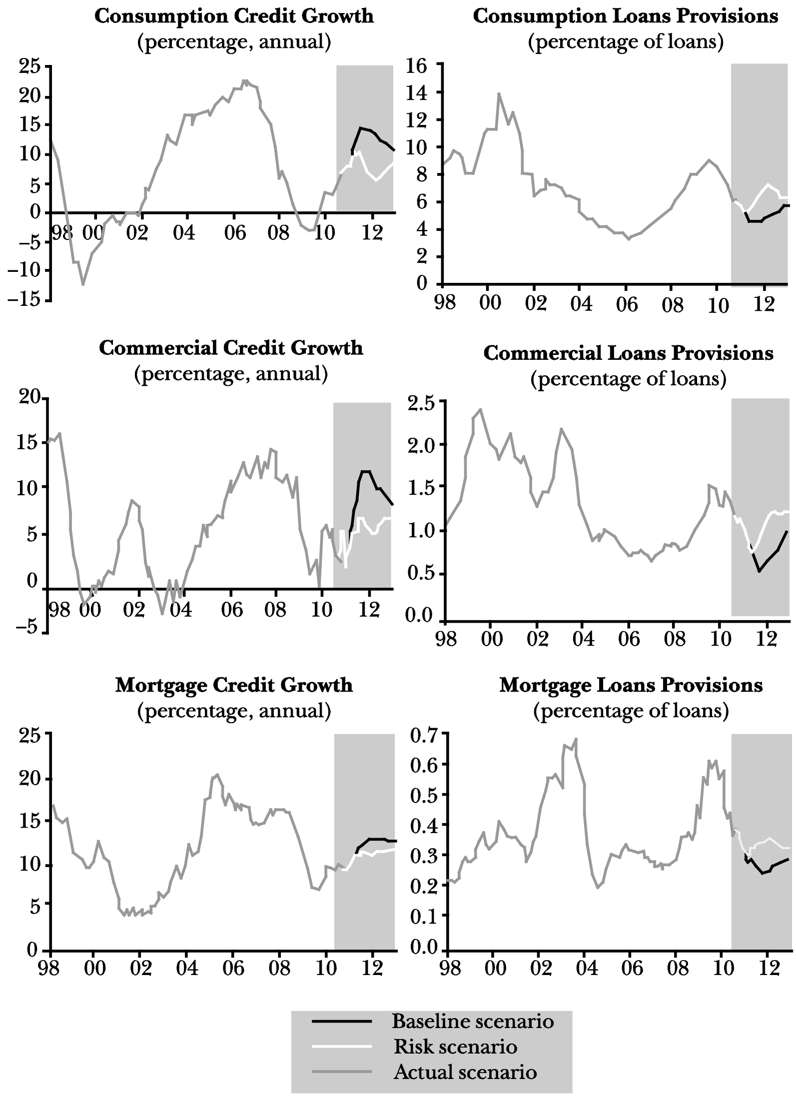
NOTES: Robust standard errors in parentheses. Dummies are not reported.

a large increment as result of the strong contraction of credit growth.

4. FINANCIAL STABILITY REPORT

In the previous sections we discussed how the market and credit risks are computed for banks. It should be noted that the latter implies a path of losses which are added on a yearly basis. For that we add the effects up to one year after the

FIGURE 1. CREDIT GROWTH AND LOAN LOSS PROVISIONS FORECASTS, 1998-2010



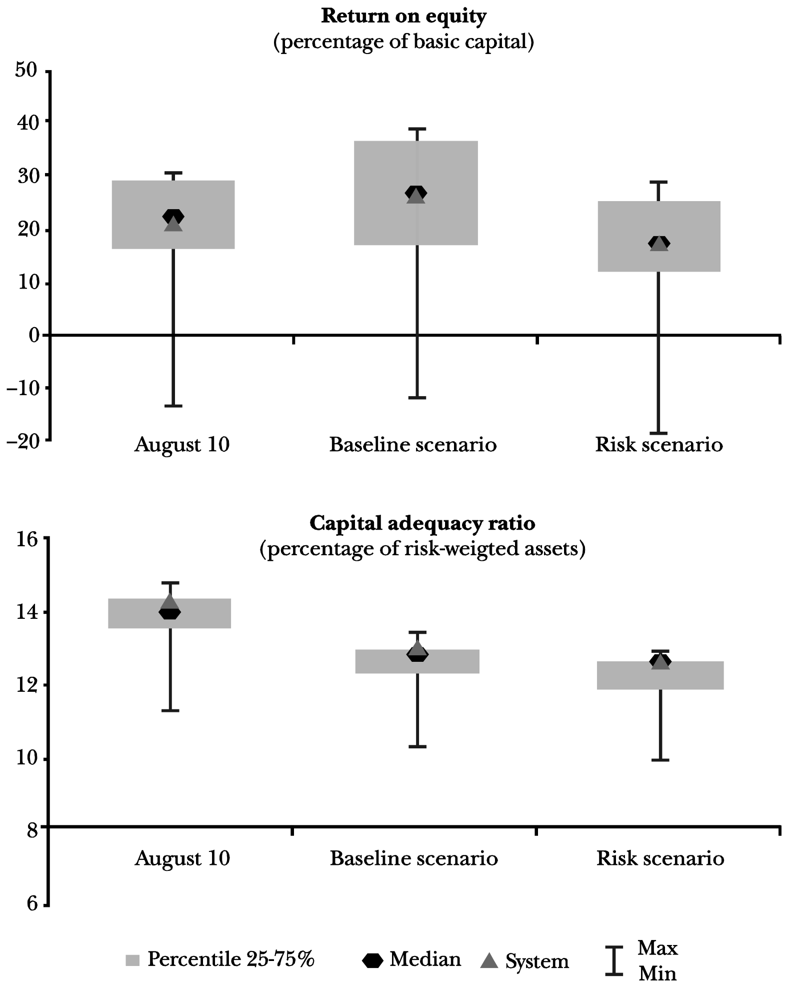
shock. The sum of market and credit risks is simple given that both risks are computed in monetary terms: change in the value of the portfolio in the first case and increasing in the loan loss provision in the second. Both can be considered as an increment on expenditures and for that reducing directly there turn on equity. Given that the initial value of return on equity is the first buffer to support the increment on expenditures generated by the simulated risk scenario (table 6).

TABLE 6. RISK SUMMARY

Market Risk		*Market Risk*	
Interest rate	−1.38	Consumption	−2.74
Exchange rate	0.30	Commercial	−4.63
		Mortgage	0.19
Total	*−1.08*	*Total*	*−7.19*

NOTE: % of basic capital.

FIGURE 2. IMPACT OF DIFFERENT RISK SCENARIOS



The results obtained are at the bank level and that information is discussed directly with a committee from the SBIF. This is a regular biannual meeting previous to the release of the Financial Stability Report (FSR). The goal of this meeting is to discuss the results obtained in the simulations with the regulator, and for that giving more sense of the numbers. After that qualitative analysis the results are adjusted accordingly and a new set of bank level results are available for the board of the Central Bank. Comments received from the board are discussed again with SBIF team and a final round of results are obtained.

Finally, for the FSR we consider the results for the whole system. Indeed, the details of each risks are summarized at the level of the system and the distribution of banks is presented in a graph where the amount of capital is used as a weight (see figure2). In terms of ROE banks exhibiting losses remains in a similar situation under the base line and risk scenarios, also over-exposed banks change this indicator dramatically. However, CAR is usually well-above the norm given that most of the bank are well-capitalized. It should be noted that the norm is 8% but in the case of Chile there are financial incentives to be over 10% due to the regulation of the pension funds administrators.

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Sherene A. Bailey-Tapper

Investigating the link between bank capital and economic activity: evidence on Jamaican panel data

1. INTRODUCTION

Business cycle fluctuations and increased asset risk may lead to large swings in bank capital, as has been reflected in the current economic and financial crisis. If these swings in capital are procyclical or result in amplifying business cycle fluctuations, then this has the potential to fuel further economic uncertainty. The global financial crisis has highlighted that financial systems which are excessively procyclical can lead to adverse consequences by reinforcing the momentum of economic cycles. For instance, in a recession, when raising capital is costly, profits are decreasing and risks are likely to materialize, banks may be forced to reduce their loan portfolio in order to meet capital requirements or increase capital holdings. The resulting contraction in available credit is likely to deepen and prolong the recession. In addition, in a downturn, banks are likely to pass on increased costs of capital to

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borrowers in the form of higher interest rates, given the prevailing low sources of finance. This may result in manufacturing and other industries cutting back on investment spending and aggravating the downturn. This can lead to a breakdown in the normal linkages between savers and investors, thereby compromising the effectiveness of monetary policy and further undermining financial stability. Alternatively, during an expansion, procyclicality may be manifested in the form of banks' lowering capital holdings and increasing bank lending, given the greater willingness of institutions to take on risks and to compete more aggressively for new business during this phase of the cycle. This is likely to fuel the pace of economic acceleration, which may ultimately impair financial stability when the cycle bursts. Therefore, procyclicality of the financial system raises challenges for policymakers in maintaining macro stability.

The global financial crisis has also reignited debate on the potential consequences of strongly risk sensitive regulatory regimes, such as the Basel II, for macro financial stability. A key objective of the new Basel Accord is to increase the risk sensitivity of minimum capital requirements of banks, but this has stimulated strong debate on the procyclical effects such risk-sensitive requirements might have on the economy. For instance, under Basel II, capital requirements will be dependent on the current risk assessment of borrowers, resulting in increased capital requirements if borrowers are downgraded during a downturn. However, this would result in the wide scale increase in the capital of banks during a downturn, which could further jeopardize macroeconomic stability. This could offset the intended goal of capital regulation, which is to enhance stability of individual banks and the entire financial system. In this respect, the Basel II Accord may make it harder for policymakers to maintain macroeconomic stability. In this context, proposals for reform of financial system regulation stress the need to make the financial system less procyclical.

Against this background, some studies have focused on the link between the business cycle and capital requirements. However, few banks hold just the minimum capital required by regulators. As such, a number of authors have investigated the relationship between excess capital and the business cycle.[1] In some literature, capital buffers are seen as a potential solution for mitigating procyclicality if it emerges under Basel II. This would mean that banks accumulate capital during upturns which might be used to satisfy a likely increase in capital requirements during a next downturn (Ayuso et al., 2002). For instance, banks build reserves during an expansion in order to dampen exuberance in good times e.g. when there is a sharp increase in house prices, banks should build buffers to ensure their lending practices are robust against rapid increases in housing prices which could quickly reverse.

Banks may hold capital buffers to avoid possible undesirable regulatory and market sanctions if capital sharply and unexpectedly declines below the minimum. In these instances, buffers help banks avoid costs related to market discipline and supervisory intervention. Furthermore, banks hold capital buffers as a signal to the market of their soundness (and satisfy the expectations of the rating agencies) and for competitive reasons and in order to facilitate borrowing funds at lower interest rates. Higher portfolio volatility also leads institutions to increase their capital buffers, given that some institutions differ in the amount of capital they hold on the basis of their risk aversion. Banks may also hold buffer capital as a way of positioning to exploit potential investment opportunities by increasing their capital ratio above the Basel requirement.

Economic cycles cannot be avoided; therefore it is critical for supervisors to develop macro prudential approaches and appropriate regulatory measures to reduce the impact of future economic cycles and diminish procyclical behavior. Against this background, it is important for local regulators to examine the role of bank capital in influencing economic credit cycles. If there is increased procyclicality under Basel II, can this impact be offset, at least partially, by banks' capital buffers? Against this background, the paper assesses the relationship between economic activity and bank capital. This is accomplished, by estimating an equation for the determinants of bank capital, which incorporates, as one of

[1] Throughout the paper, excess capital is used synonymously with buffer capital. Both terms refers to the amount of capital banks' hold in excess of that required of them by regulators.

the determinants, an indicative measure of growth in economic activity.

The paper is organized as follows: section 2 presents the literature review while section 3 outlines the framework employed to investigate the determinants of bank capital. Section 4 gives a brief description of the data and the estimation technique employed. Section 5 presents the findings of the model, while the policy implications of the results and the conclusion are outlined in section 6.

2. LITERATURE REVIEW

A number of studies on the procyclicality of bank capital have focused on the sensitivity of capital requirements to economic activity or business cycle fluctuations. In one study, using data across 120 countries, Bikker and Metzemakers (2004) investigated the determinants of commercial banks' own internal capital targets and the potential sensitivity of these levels to the business cycle based on the Basel I accord. As expected, results showed that minimum requirements do not fluctuate substantially over the business cycle. However, smaller banks combined a relatively risky portfolio with limited buffer capital, which could induce procyclicality of these institutions' capital holdings under Basel II.

Bikker and Metzemakers (2004) also found that banks tend to hold substantial capital buffers on top of minimum requirements, reflecting that they hold capital for other reasons than strictly meeting the capital requirements. More recent studies have examined how this excess capital behaves over the business cycle. Some authors also examine whether there is a positive relationship between capital buffers and growth in economic activity, in order to assess whether this could help in offsetting the greater procyclicality in capital requirements anticipated under Basel II. Some findings show that buffers will not be sufficient to prevent procyclicality of bank capital and lending, therefore strong regulatory reform is needed.[2]

[2] Banks' lending position is a function of historically determined capital positions and capital requirements imposed by regulation. Risk sensitivity

Using annual data on Spanish banks over the period 1986 to 2000, Ayuso et al. constructed an equation for capital buffers over time and across institutions that reflected the cost of capital, non-performing loan rations, size-specific dummy variables and the annual growth in GDP. They found that capital buffers are negatively related to the growth rate in GDP under the Basel I framework. That is, capital buffers tend to fall in periods of rising GDP and rise when GDP falls. The results also showed that capital buffers are negatively related to the cost of capital, the level of non-performing loans and to a dummy variable which accounts for banks in the largest 10% of the sample. The findings for Spanish bank raise concerns as to whether capital buffers would be useful in mitigating the anticipated procyclical impact expected under Basel II.[3]

Stolz and Wedow (2005) investigated the effect of the business cycle on the regulatory capital buffer of German savings and cooperative banks over the period 1993 to 2003. They found that capital buffers fluctuate countercyclically over the business cycle. The study also found that banks with low capital buffers reacted differently to the business cycle than banks with relatively higher capital buffers. For instance, in business cycle downturns, low-capitalized banks dampen the increase in capital, while well capitalized banks boost the increase in capital. In addition, low capitalized banks do not decrease risk weighted assets in a business cycle downturn by more than well-capitalized banks. The authors found that while this issue may raise some supervisory concerns, it also implies that low capitalized banks do not cut back on lending, as these institutions did not reduce risk-weighted assets in a downturn. As such, the results do not support the widely held view that banks with low capital buffers cut back on lending in order to increase capital buffers in a downturn, thereby further aggravating the contraction in economic activity.

Jokipii and Milne (2006) investigated the cyclical behavior of bank capital buffers on capital regulation of European

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of capital requirements may imply a substantial increase in the procyclicality of bank lending.

[3] Credit standards need to be raised and credit extensions restricted during an upturn in order to minimize bad debt losses that erode capital during a downturn.

banks, under the old Basel 1988 Accord, using panel data over the period 1997 to 2004. Their objective was to determine the extent of the co-movement between this buffer and the cycle, and to determine whether such co-movement is country, bank type or bank size specific. They found that, for EU15 countries, and controlling for individual bank costs and risks, there was a negative co-movement between capital buffers and the business cycle.[4] A main conclusion from their study is that, the negative co-movement of capital buffers, after implementation of Basel II, will exacerbate its pro-cyclical impact.

Boucinha and Ribeiro (2007) investigated the determinants of Portuguese banks' capital buffers using data from 1994 to 2004. The key determinants included in the study were risk measures including the ratio of provisions to non-performing loans (NPLs), a default ratio and stock holdings as a share of total assets. A ROA variable and its variance were included to measure the capacity of the institutions to absorb losses and a output variable was included to capture the impact of the business cycle on the bank's holdings of excess capital. The results showed a negative impact of the output gap on excess capital, not only suggesting that banks tend to cover the higher risks that arise in cycle downturns with higher capital reserves, but also that the lending behavior may be procyclical, in that it will tend to amplify economic cycles. The findings confirm the theory that banks adjust their capital reserves in response to changes in the risks they face. That is, both those directly resulting from changes in the macroeconomic environment throughout the cycle and those resulting from banks' own decisions.

3. THE THEORETICAL FRAMEWORK

3.1 How is Bank Capital Determined?

The framework employed to evaluate the determinants of

[4] EU15 represents the number of member countries in the European Union prior to the accession of ten candidate countries on 1 May 2004. The EU15 is comprised the following 15 countries: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden, and United Kingdom.

bank capital buffers is based on the model by Ayuso et al. (2002). This model starts with an equation, which is based on the literature on real investment which describes the dynamics of the capital stock of a representative single bank. Based on equation (1), k_t stands for the capital level at the end of period t and I_t stands for stock issues or repurchases plus retained profits during period t.

$$k_t = k_{t-1} + I_t \,. \tag{1}$$

In addition, the model captures the decision a bank makes on capital as a result of a tradeoff among three different types of costs related to capital levels (see Froot and Stein, 1998).

These costs are outlined in equation (2):

$$C_t = (\alpha_t - \gamma_t)k_t + (1/2)\delta_t I^2{}_t \,. \tag{2}$$

Where α_t represents the costs of remunerating capital; γ_t represents the cost of failure (and/or penalties for not complying with the regulatory minimum); and δ_t reflects adjustment costs.[5] The costs of remunerating capital involve direct costs to the bank of holding capital. The opportunity cost of bank capital or the cost of remunerating capital may even be more costly than alternative bank liabilities such as deposits or debt [see Campbell (1979) and Majluf (1984)].

Secondly, holding sufficient capital reduces the probability for the bank to face costs related to not complying with compulsory capital requirements. Additionally, holding capital minimizes the probability of bankruptcy as well as costs related to loss of charter value, reputational costs and legal costs of the bankruptcy process (see Acharya, 1996).

The final cost represented in equation (2), has to do with adjustment costs as a result of changing the capital level. These costs are related to transaction costs as well as costs due to the presence of asymmetric information between buyers and sellers of stocks in the capital markets, which can increase or reduce adjustment costs.

Against this background, a typical bank minimizes its intertemporal costs by solving the following problem:

[5] Assumptions include linearity between the first two groups of costs and symmetry in relation to adjustment costs.

(3) $$\underset{\{I_{t+i}\}_0^\infty}{Min}\, E_t \sum_{i=0}^{\infty} \beta^i C_{t+i}\ ,$$

such to

(4) $$C_t = (\alpha_t - \gamma_t)k_t + (1/2)\delta_t I^2{}_t$$

(5) $$k_t = k_{t-1} + I_t\ .$$

Based on first order conditions, equation (4) can be re-written as follows:

(6) $$I_t = E_t\left(\frac{1}{\delta}\sum_{i=0}^{\infty}\beta^i(\gamma_{t+i} - \alpha_{t+i})\right).$$

And therefore:

(7) $$E_t(K_t) = K_{t-1} + E_t\left(\frac{1}{\delta}\sum_{i=0}^{\infty}\beta^i(\gamma_{t+i} - \alpha_{t+i})\right).$$

Subtracting the regulatory minimum from both sides of equation (7), and replacing expected capital by observed capital and including an expectation error term yielded the expression outlined in equation (8):

(8) $$(K - \bar{K})_t = (K - \bar{K})_{t-1} + E_t\left(\frac{1}{\delta}\sum_{i=0}^{\infty}\beta^i\gamma_{t+i}\right) - E_t\left(\frac{1}{\delta_t}E_t\sum_{i=0}^{\infty}\beta^i\alpha_{t+i}\right) + \varepsilon_t\ .$$

A more specific empirical model is outlined in equation (9), for the capital buffer $\left(BUF_{it}\right)$ held by institution i in period t:

(9) $$\begin{aligned} BUF_{it} = \beta_0 BUF_{i,t-1+\beta_1} ROE_{it} + \beta_2 NPL_{it} + \beta_3 BIG_{it} + \\ + \beta_4 SMA_{it} + \beta_5 GDPG_{t+\eta_t} + \varepsilon_{it,} \end{aligned}$$

i = 10.5 1, 2, ..., N (number of banks), t=1, 2, ..., T.

Where $BUF_{i,t-1}$ captures adjustment costs as a result of the bank increasing its buffer capital, while ROE represents the cost of remunerating excess capital and is expected to have a negative coefficient. The NPL variable captures the risk profile of the institution. After including the determinants of capital buffer based on the model described above, a GDP growth variable was also included in order to determine whether the business cycle has an additional effect on the capital buffer held by banking institutions. And BIG and SMA represent dummy variables to capture institution size, where BIG (SMA) take the value of 1 for the largest (smallest) banks based on asset size.

4. EMPIRICAL ANALYSIS

4.1 Data

Two equations for bank capital were estimated for each sector in the banking system, which includes commercial banks, merchant banks and building societies, over the period February 2002 to March 2009 based on the model outlined in equation (9).[6] The data used in this study is an unbalanced panel of monthly balance sheet data for each of the three sectors, an indicative measure of economic activity and a proxy for cost of capital. A summary of the data statistics are presented in tables 4 to 6 in the appendix.[7]

Cost of capital is proxied by the average monthly 30-day private market rate, risk by the ratio of non-performing loans to total loans and changes in economic activity by growth in the money supply. The dependent variable for the excess capital equation (excess capital ratio) is measured as excess capital as a ratio of risk-weighted assets, while for the dependent variable for the total capital equation (total capital ratio) is measured as total capital as ratio of risk-weighted assets. Additionally, government of Jamaica (GOJ) sovereign bonds as a share of total assets and shares as a proportion of total assets are also included in the analysis.[8] Dummy variables were also used to capture the impact of small and large banks, respectively.

[6] An excess capital equation and total capital equation were estimated.

[7] The study utilizes panel data, because of the advantage of capturing both differences across banks and time-series variation, as well as of allowing for meaningful statistical inferences even using a sample with a relatively small number of banks observed over an equally short time period. The explicit treatment of the model's dynamic is relevant not only to infer on the persistence of the dependent variable, but also to ensure that estimates for other parameters of the model are consistent.

[8] This variable was included given potential for increased volatility in GOJ bond yields and likely marked to market losses to negatively impact earnings and impair capital.

The generalized method of moments (GMM) estimator was employed in the empirical assessment and is very useful in obtaining unbiased and efficient estimates in dynamic models with lagged endogenous variables as regressors. Against this background, the paper utilizes the GMM estimator developed by Arellano and Bond (1991). The procedure was applied to both equations.

The main advantages of this methodology consist in the possibility of obtaining consistent estimates for the parameters of interest when the persistence of the dependent variable needs to be explicitly modeled and not requiring strong hypotheses about the exogeneity of the regressors.

4.2 Results: Merchant Banks[9]

One of the key findings from the excess capital equation is that overall results show a significant inverse relationship between growth in the money supply and the excess capital ratio (see table 1, panel A).[10] This result indicates that during an expansion there are likely to be declines in the excess capital ratio, while the reverse is expected to occur during a contraction in economic activity. An implication of this result is that, for instance, during a downturn, to the extent that an increase in the excess capital ratio may be reflective of growth in excess capital or a decline in risk weighted assets, this may be the result of a reduction in loan supply; which is likely to further aggravate the downturn. In other words, the policy implication of this is that institutions build up capital buffers when it is too late, only to further aggravate prevailing economic conditions and further jeopardize financial stability. Findings by bank size also confirm a negative and significant relationship between bank capital and money supply growth for the smaller merchant banks. This result reflects the fact that smaller institutions are more likely to increase buffer capital to avoid market and regulatory sanctions and to facilitate borrowing funds at lower interest rates. However, for the

[9] In general, the instruments chosen in the empirical assessment were one lag of the loans to asset ratio, the ratio of non-performing loans to total loans and the cost of capital variable.

[10] Overall results relate to findings for *All banks*.

larger merchant banks, there is positive and significant relationship between growth in the money supply and excess capital holdings. This outcome suggests that larger merchant banks have stronger risk management practices, enabling them to identify and account for risks when economic activity accelerates by building capital buffers. As such, the buildup in capital buffers during the period of expansion can then be utilized in the event of a downturn. In addition, the coefficient associated with the lagged dependent variable is positive and significant for both large and small merchant banks, presenting evidence in favor of the adjustment cost hypothesis. For all merchant banks, in particular smaller commercial banks, there is a positive and significant relationship between the cost of capital variable and excess capital. This is consistent with a priori expectations, in that, when there are increases in this variable, these institutions are likely to retain capital for satisfying future funding needs rather than substituting alternative liabilities.

Also, for merchant banks, overall findings show that an increase in GOJ sovereign bonds as a share of total assets results in higher capital buffers, suggesting that banks with higher exposure to market risk hold higher capital reserves in order to cover for the additional risk. However, overall results show that a higher weight of loans and stocks and shares as a proportion of total assets is associated with declines in the excess capital ratio. This impact is largely reflective of the resulting expansion in risk weighted assets.

The results showed similar findings for the total capital equation (see table 1, panel B). Most notably, findings for all merchant banks and for smaller merchant banks show a significant and inverse relationship between growth in the money supply and the total capital ratio. This suggests that for the sector, and in particular for the smaller merchant banks, excess capital holdings may have influenced the performance of bank capital. Additionally, these results imply that there may be increased procyclicality under Basel II, given the anticipated increased sensitivity of capital requirements to the business cycle under the new Accord. Additionally, for the larger merchant banks, there is a positive and significant relationship between growth in the money supply and capital holdings.

TABLE 1. GMM MODEL: MERCHANT BANKS. DYNAMIC PANEL RESULTS

	All banks		*Large banks*		*Small banks*	
	GMM	*t-statistic*	*GMM*	*t-statistic*	*GMM*	*t-statistic*
		A				
Excess capital ratio (–1)	0.908059	80.636352	0.68443	10.17773	0.850883	20.4051
Growth M2	–0.146169	–7.819126	0.191802	1.721919	–0.102489	–1.799846
COC	0.033532	2.121068	–0.030149	–0.264763	0.008247	2.43927
Loans to TA	–0.028615	–2.702348	–0.251972	–31.66202	–0.146012	
NPLTOTL	0.104796	1.14207	–0.010441	–0.08548		
GOJ sovereign to TA	0.130222	9.057384	0.034307	0.567282		
Shares to TA	–0.455722	–2.297734	–1.510423	–9.63021		
@LEV(@ISPERIOD("2002"))	–0.009415	–2.052762	–0.018258	–1.345218	0.006508	0.905746
@LEV(@ISPERIOD("2003"))	0.020436	1.316535	0.037625	3.155015	0.00785	2.55108
@LEV(@ISPERIOD("2004"))	–0.015913	–3.851947	–0.033814	–2.191135	–0.037286	–1.465231
@LEV(@ISPERIOD("2005"))	0.123299	1.168345	0.108073	1.761509	–0.03816	–2.248006
@LEV(@ISPERIOD("2006"))	0.024048	2.409209	0.00542	0.984749	0.00972	0.278034
@LEV(@ISPERIOD("2007"))	0.009467	0.55572	–0.004603	–0.229135	0.013529	2.079039
@LEV(@ISPERIOD("2008"))	–0.018139	–1.790206	–0.0184458	–8.033932	0.004735	0.293282
		Effects specification				
R^2		0.8208		0.8301		0.8122
J-statistic		78.0883		57.015		29.3612
Sum squared residuals		0.6423		0.5851		1.425
Instrument rank		87		86		86

B

Total capital ratio (–1)	0.945077	66.15944	0.673392	11.29773	0.850883	20.4051
Growth M2	–0.063238	–1.0684433	0.22618	2.666005	–0.102489	–1.799846
COC	0.014433	0.53154	–0.31977	–0.22427	0.008247	2.43927
Loans to TA	–0.020609	–0.98904	–0.265742	–21.97891	–0.146012	–4.020318
NPLTOTL	0.018821	0.22323	–0.017123	–0.205157		
GOJ sovereign to TA	0.129616	9.218918	0.033136	0.638439		
Shares to TA	–0.362374	–3.065043	–1.537832	–10.76731		
@LEV(@ISPERIOD("2002"))	–0.004978	–0.690523	–0.020645	–1.756495	0.005434	0.600278
@LEV(@ISPERIOD("2003"))	0.017027	1.254004	0.03693	3.5959	0.005222	13.12304
@LEV(@ISPERIOD("2004"))	–0.021132	–3.00378	–0.035584	–2.020909	–0.018702	–11.52061
@LEV(@ISPERIOD("2005"))	0.123308	1.650168	0.106916	1.752447	–0.022416	–157.8733
@LEV(@ISPERIOD("2006"))	0.021158	2.865869	0.003903	0.649991	0.013614	1.771663
@LEV(@ISPERIOD("2007"))	0.007752	0.419642	–0.005694	–0.27152	0.014711	2.348312
@LEV(@ISPERIOD("2008"))	–0.016216	–1.538077	–0.019217	–9.362625	0.012633	8.795275
Effects specification						
R^2		0.8144		0.8292		0.8122
J-statistic		72.4457		58.4378		293612
Sum squared residuals		0.6649		0.5884		1.425
Instrument rank		87		87		86

4.3 Results: Commercial Banks

In contrast to the merchant banks, overall results for the commercial banking sector show that there is no significant relationship between growth in the money supply and excess capital (see table 2, panel A). However, findings by bank size show a significant inverse relationship between growth in the money supply and excess capital holdings for the smaller commercial banks. For the larger commercial banks, there is an insignificant relationship between growth in the money supply and excess capital holdings. The results in relation to cost of capital are similar to those obtained for the merchant banks. Findings for the overall sector and for the smaller commercial banks show a positive and significant relationship between cost of capital and excess capital holdings.

Results for the overall sector show an insignificant relationship between deterioration in loan quality and excess capital holdings. There were similar findings for the larger commercial banks. However, for smaller commercial banks, deterioration in loan quality is associated with declines in the excess capital ratio, and may reflect the impact of a greater level of provisioning by these institutions in response to the increased default risk.[11]

Similar to the merchant banks, results for all commercial banks show a positive and significant relationship between the ratio of GOJ sovereign bond holdings to total assets and the excess capital ratio. This suggests that a higher weight of GOJ sovereign bonds as a proportion of assets results in these institutions holding higher capital buffers as a means of covering additional exposure related to market risk. In addition, the overall findings show that increases in the loan to asset ratio is associated with declines in the excess capital ratio, largely reflecting the impact on risk weighted assets as a result of the growth in loan holdings.

For the total capital equation, the results also show an insignificant relationship between growth in the money supply and bank capital for all commercial banks and for larger commercial banks (see table 2, panel B). For the smaller commercial banks, similar to the finding from the excess capital

[11] Increased provisioning is a substitute for holding higher capital.

equation, there is a significant inverse relationship between growth in the money supply and bank capital. Nonetheless, the findings for other determinants of bank capital are largely consistent with the findings from the excess capital equation across all banks.

4.4 Results: Building Societies

Results for the building societies' sector show that there is no significant relationship between growth in the money supply and excess capital (see table 3, panel A). Regarding the other determinants of excess capital, as in the case of the merchant banks and commercial banks, there is a negative and significant relationship between the loan to asset ratio and excess capital, and this result holds regardless of bank size.

Of importance, unlike for the other sectors, the overall results show an insignificant relationship between the ratio of GOJ sovereigns to total assets and the excess capital ratio. This finding is not surprising for the building societies given the relatively lower of share of GOJ sovereigns as a ratio of total assets for this sector.

For the total capital equation, the results also show an insignificant relationship between growth in the money and bank capital, regardless of bank size (see table 3, panel B). In addition, the findings for other determinants of bank capital are largely consistent with results from the excess capital equation.

5. CONCLUSION AND POLICY IMPLICATIONS

The main purpose of the study was to investigate the impact of economic activity on bank capital in Jamaica during the period 2002 to 2009. This is accomplished through the estimation of a dynamic panel framework, which includes as one of the determinants, an indicative measure of growth in economic activity. The motivation for the study stems from ongoing debate that risk based capital requirements, in particular the Basel II accord, is anticipated to result in bank capital reinforcing economic cycles. This is anticipated to occur in a

TABLE 2. GMM MODEL: COMMERCIAL BANKS. DYNAMIC PANEL RESULTS

	All banks		*Large banks*		*Small banks*	
	GMM	*t-statistic*	*GMM*	*t-statistic*	*GMM*	*t-statistic*
		A				
Excess capital ratio (–1)	0.927955	42.95414	0.856253	42.26902	0.965676	70.38426
Growth M2	–0.151202	–1.445029	–0.157352	–1.482519	–0.245477	–2.906133
COC	0.045507	2.241119	–0.099074	–0.382257	0.033173	3.130909
Loans to TA	–0.008648	–1.545952	–0.082088	–3.912763	–0.02188	–2.652898
NPLTOTL	–0.17377	–1.212287	–0.150771	–1.157714	–0.097404	–2.01774
GOJ sovereign to TA	0.048734	2.130923	–0.058883	–0.737292		
@LEV(@ISPERIOD("2002"))	–0.001023	–0.170198	–9.31E–5	–0.032332	–0.003871	–0.29192
@LEV(@ISPERIOD("2003"))	0.003764	0.382574	–0.000959	–0.484359	0.006367	0.258843
@LEV(@ISPERIOD("2004"))	–0.003484	–1.14822	–0.004271	–1.539021	–0.005622	–1.909526
@LEV(@ISPERIOD("2005"))	0.015743	1.050041	0.004169	2.396324	0.02631	0.855415
@LEV(@ISPERIOD("2006"))	–0.001099	–0.54113	–0.00083	–0.629506	–0.006773	–1.008837
@LEV(@ISPERIOD("2007"))	0.010146	1.645642	0.004982	0.787192	0.013389	1.364473
@LEV(@ISPERIOD("2008"))	–4.80E–5	-0.006881	0.012862	2.407693	–0.008941	–2.154441
		Effects specification				
R^2		0.9292		0.9683		0.9021
J-statistic		69.005		60.713		71.7034
Sum squared residuals		0.3323		0.0648		0.2693
Instrument rank		87		87		88

B

Total capital ratio (–1)	0.951962	43.66013	0.855914	43.44087	0.953948	63.16448
Growth M2	–0.16726	–1.353204	–0.156744	–1.491351	–0.361517	–2.643453
COC	0.043017	2.153063	–0.097281	–0.371641	0.145187	1.555919
Loans to TA	–0.005128	–0.971154	–0.082029	–3.902724	–0.028283	–2.801881
NPLTOTL	–0.08537	–1.414777	–0.151329	–1.180911	–0.284759	–2.044721
GOJ sovereign to TA	0.04926	1.81117	–0.058524	–0.729559		
@LEV(@ISPERIOD("2002"))	–0.001517	–0.251104	–7.41E–5	–0.025417	–0.000904	–0.070737
@LEV(@ISPERIOD("2003"))	0.003088	0.3039	–0.000974	–0.485045	0.007943	0.304388
@LEV(@ISPERIOD("2004"))	–0.003048	–1.059126	–0.004264	–1.556829	–0.006113	–3.53046
@LEV(@ISPERIOD("2005"))	0.015682	0.97298	0.004162	2.367349	0.028183	0.953082
@LEV(@ISPERIOD("2006"))	–0.001672	–0.574562	–0.000844	–0.64662	-0.004439	–0.834994
@LEV(@ISPERIOD("2007"))	0.010154	1.625518	0.004967	0.785175	0.013745	1.407175
@LEV(@ISPERIOD("2008"))	-0.000169	–0.020574	0.012828	2.397794	–0.007432	–3.674329
Effects specification						
R^2		0.9271		0.9683		0.9010
J-statistic		68.4824		60.6837		73.269
Sum squared residuals		0.3447		0.0648		0.2693
Instrument rank		87		87		87

TABLE 3. GMM MODEL: BUILDING SOCIETIES. DYNAMIC PANEL RESULTS

	All banks		*Large banks*		*Small banks*	
	GMM	*t-statistic*	*GMM*	*t-statistic*	*GMM*	*t-statistic*
		A				
Excess capital ratio (–1)	1.012613	71.23733	1.014542	61.38285	1.008353	63.86466
Growth M2	–0.028348	–0.849255	–0.032485	–0.963728	–0.028939	–0.810322
COC	0.013953	1.375687	0.012957	0.012957	0.012436	1.099655
Loans to TA	–0.006801	–2.592217	–0.00704	–2.532119	–0.010161	–2.590507
NPLTOTL	0.018315	0.045806	0.025087	0.468351	0.01771	0.36179
GOJ sovereign to TA	–0.022275	–0.627209	0.009825	0.187037	0.006958	0.158431
@LEV(@ISPERIOD("2002"))	0.011684	5.23706	0.011683	4.632709	0.011153	4.894417
@LEV(@ISPERIOD("2003"))	0.002759	0.718446	0.002443	0.543159	0.002712	0.706885
@LEV(@ISPERIOD("2004"))	0.006127	1.828824	0.006364	1.654198	0.005946	1.914018
@LEV(@ISPERIOD("2005"))	0.001346	0.437019	0.001357	0.439971	0.001165	0.301579
@LEV(@ISPERIOD("2006"))	0.001149	0.278404	0.002371	0.562861	0.001223	0.255671
@LEV(@ISPERIOD("2007"))	0.004706	1.530192	0.005321	1.819147	0.005336	1.468505
@LEV(@ISPERIOD("2008"))	0.003564	0.4883	0.00365	0.51276	0.004034	0.527046
		Effects specification				
R^2		0.9917		0.9916		0.9917
J-statistic		60.201		57.7368		53.8703
Sum squared residuals		0.0354		0.0356		0.0351
Instrument rank		87		87		87

B

Total capital ratio (–1)	1.011983	90.95128	1.013094	78.35096	1.007938	81.59394
Growth M2	–0.028878	–0.892728	–0.033407	–1.012699	–0.030292	–0.890709
COC	0.012395	1.275937	0.011897	1.125513	0.011976	1.099804
Loans to TA	–0.00789	–2.42487	–0.008202	–2.320956	–0.010422	–2.432261
NPLTOTL	0.018193	0.438048	0.023003	0.480169	0.015932	0.359958
GOJ sovereign to TA	–0.032227	–0.855477	–0.003535	–0.068594	–0.002458	–0.052376
@LEV(@ISPERIOD("2002"))	0.011596	5.202734	0.011605	4.785088	0.011233	4.899401
@LEV(@ISPERIOD("2003"))	0.002838	0.736971	0.002876	0.73487	0.002806	0.72991
@LEV(@ISPERIOD("2004"))	0.006105	1.788598	0.006198	1.920043	0.005998	1.924855
@LEV(@ISPERIOD("2005"))	0.001408	0.486621	0.001382	0.40211	0.001279	0.345705
@LEV(@ISPERIOD("2006"))	0.001123	0.276463	0.001372	0.288884	0.001267	0.265314
@LEV(@ISPERIOD("2007"))	0.004612	1.650409	0.005123	1.75511	0.005266	1.574791
@LEV(@ISPERIOD("2008"))	0.003729	0.509448	0.003931	0.51002	0.00412	0.536788
Effects specification						
R^2		0.9917		0.9916		0.9917
J-statistic		62.6693		59.769		56.2294
Sum squared residuals		0.0353		0.0355		0.035
Instrument rank		87		87		87

context where there is expected to be increased sensitivity of capital charges to changes in the economic cycle. This is expected to materialize because capital requirements or charges will be revised to take into account a more dynamic assessment of the credit rating of borrowers which is based on the stage of the economic cycle. While this will give a better assessment of the true capital charges facing the institution, it could jeopardize macroeconomic and financial stability by reinforcing the state of an economic cycle. As such, in anticipation of the introduction of the Basel II accord in some economies, many authors have explored the relationship between economic activity and bank capital buffers, particularly in a context where most institutions hold capital well in excess of the minimum regulatory requirement.

Against this background, this study is intended to provide local regulators and policymakers with evidence on the relationship between bank capital and economic activity under the existing Basel accord. Evidence from the study is useful in understanding whether capital buffers can be a useful tool in mitigating procyclicality, particularly in a context where this may increase with the introduction of the new Accord. The study also examines the importance of other determinants of bank capital, which is also important in understanding the risk motives of these institutions.

Based on the results of the model, there is evidence for the commercial banks and merchant banks that bank capital is likely to reinforce economic cycles, driven by the smaller institutions. One implication of this result is that banks are unlikely to build up buffers during expansions and, as such, are more likely to have difficulties meeting capital requirements and offsetting losses when there is a downturn in the business cycle. This is of concern for regulators, given that based on the findings, capital buffers would not help in offsetting any increased procyclicality of risk sensitive capital requirements under Basel II.

Based on an April 2009 report by the Financial Stability Forum, there are various approaches regulators can employ to address increased procyclicality in the financial system. Some of the recommendations from the report are that regulators should employ techniques to strengthen the regulatory framework so that the quality and level of capital in the banking system increase during strong economic conditions, which can be drawn down during periods of economic and financial stress. In addition to this, regulators should maintain close monitoring and surveillance of the financial system during periods of economic downturn. Secondly, regulators should employ enhanced stress testing practices to inform the buildup of capital buffers above the regulatory minimum during periods of economic expansion. This is in an effort to fully capture potential areas of vulnerabilities, as well as risks which may materialize in the event of a downturn. This would involve regulators continually revising stress testing in relation to financial developments and the banks' evolving risk profile. Additionally, similar to what has been done by the Bank of Spain; a dynamic provisioning can be employed, which is also useful in dampening procyclicality. Under this technique, banks make provisions based on the losses expected when loans are originated. This would result in a rising stock of provisions when actual losses are low, which would help to protect banks in periods when actual losses are high.

An important finding is that merchant banks cover additional market risk associated with holding an increasing share of GOJ sovereigns by holding higher capital buffers. This is also the case for the commercial banking sector. For the larger merchant banks, there is positive and significant relationship between growth in the money supply and excess capital holdings. The implication of this is that these institutions are more likely to identify and account for risks when economic activity accelerates, by building capital buffers, which can be utilized in the event of a downturn. While findings show that for smaller commercial banks, there is a significant inverse relationship between growth in economic activity and excess capital holdings, providing evidence of procyclicality or the likelihood for banking activity to reinforce economic credit cycles.

Annex

FIGURE A.1. AVERAGE EXCESS CAPITAL RATIO FOR COMMERCIAL BANKS (EXCESS CAPITAL/REQUIRED CAPITAL), 2002-2009

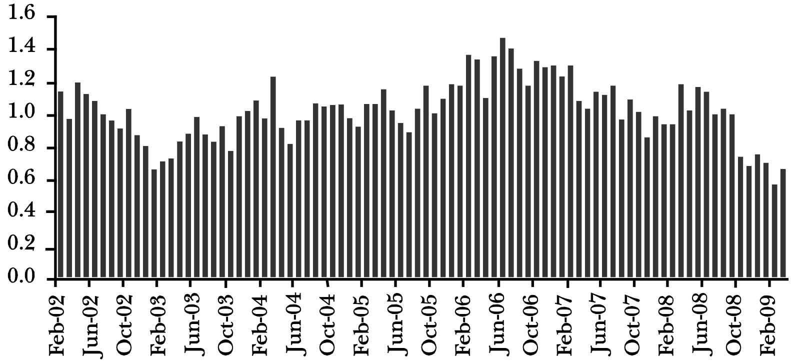

FIGURE A. 2. AVERAGE EXCESS CAPITAL RATIO FOR MERCHANT BANKS (EXCESS CAPITAL/REQUIRED CAPITAL), 2002-2009

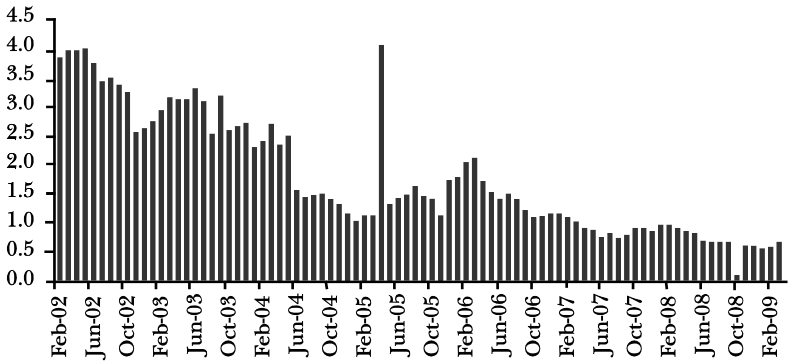

FIGURE A. 3. AVERAGE EXCESS CAPITAL RATIO FOR BUILDING SOCIETIES (EXCESS CAPITAL/REQUIRED CAPITAL), 2002-2009

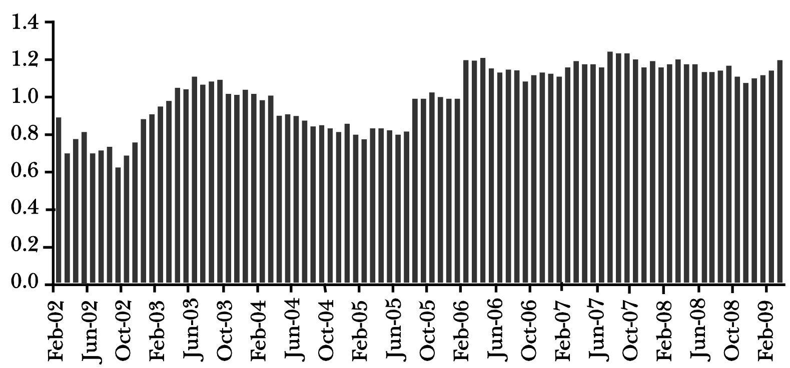

FIGURE A 4. EXCESS CAPITAL RATIO BY BANK SIZE: MERCHANT BANKS (EXCESS CAPITAL/REQUIRED CAPITAL), 2002-2009

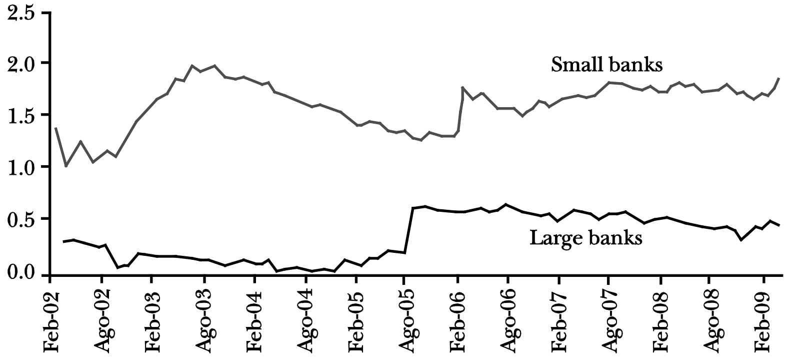

FIGURE A. 5. EXCESS CAPITAL RATIO BY BANK SIZE: COMMERCIAL BANKS, 2002-2009

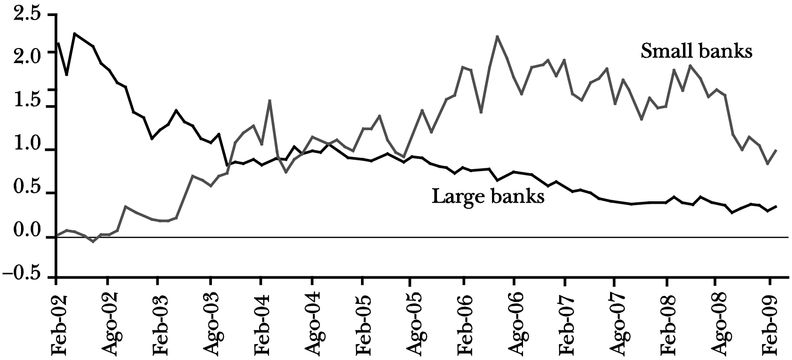

FIGURE A. 6. EXCESS CAPITAL RATIO BY BANK SIZE: FIAS, 2002-2009

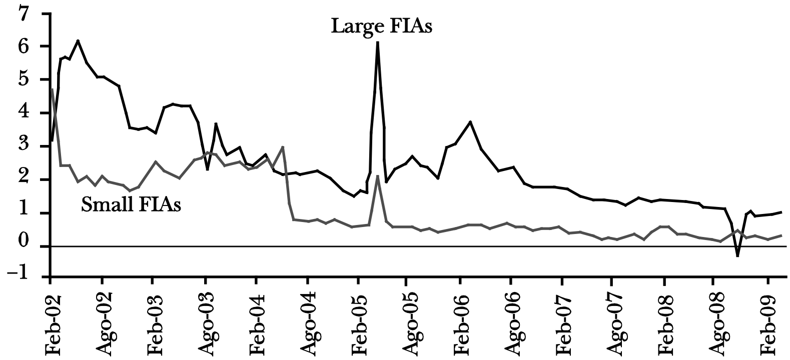

TABLE A. 1. SUMMARY STATISTICS: COMMERCIAL BANKS

	EC	*GM2*	*COC*	*LOANSTA*	*NPLTOTL*	*GOJSTA*
Mean	0.097943	0.000709	0.158393	0.315000	0.026964	0.097612
Median	0.059198	0.000200	0.140000	0.320000	0.020000	0.080000
Standard deviation	0.105931	0.001167	0.050249	0.140081	0.025562	0.086647
Skewness	1.366869	2.623706	2.242343	0.146656	4.468303	1.028531
Sum	43.878470	0.317520	70.960000	141.120000	12.080000	43.730000
Sum squared deviation	5.016001	0.000609	1.128643	8.771400	0.292071	3.355944
Observations	448	448	448	448	448	448

TABLE A. 2. SUMMARY STATISTICS : MERCHANT BANKS

	EC	*GM2*	*COC*	*LOANSTA*	*NPLTOTL*	*SHARESTA*	*GOJSTA*
Mean	0.253434	0.000742	0.158976	0.259578	0.051747	0.018434	0.201325
Median	0.250000	0.000205	0.140000	0.235000	0.030000	0.020000	0.175000
Standard deviation	0.147536	0.001180	0.048550	0.177161	0.046891	0.015491	0.139430
Sum	42.070000	0.123130	26.390000	43.090000	8.590000	3.060000	33.420000
Sum squared deviation	3.591543	0.000230	0.388926	5.178670	0.362793	0.039593	3.207708
Observations	166	166	166	166	166	166	166

TABLE A. 3. SUMMARY STATISTICS: BUILDING SOCIETIES

	EC	*GM2*	*NPLTOTL*	*LOANSTA*	*GOJSTA*	*COC*
Mean	0.100809	0.000707	0.036424	0.539622	0.027122	0.156279
Median	0.044762	0.000200	0.030000	0.450000	0.010000	0.140000
Standard deviation	0.111588	0.001163	0.022800	0.206517	0.030480	0.047835
Sum	34.67828	0.243360	12.53000	185.6300	9.330000	53.76000
Sum squared deviation	4.270995	0.000464	0.178302	14.62865	0.318651	0.784837
Observations	344	344	344	344	344	344

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