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Abstract

We assess the extent to which capital buffers (the capital banks hold in excess of the regulatory minimum) exacerbate rather than reduce the cyclical behavior of credit. We empirically study the relationships between output gap, capital buffers and loan growth with firm-level data for French banks over the period 1993—2009. Our findings reveal that bank capital buffers intensify the cyclical credit fluctuations arising from the output gap developments, all the more as better quality capital is considered. Moreover, by performing Granger causality tests at the bank level, we find evidence of a two-way causality between capital buffers and loan growth, pointing to mutually reinforcing mechanisms. Overall, those empirical results lend support to a countercyclical financial regulation that focuses on highest-quality capital and aims at smoothing loan growth.

Keywords: Bank Capital Regulation, Procyclicality, Capital Buffers, Business Cycle Fluctuations, Basel III

JEL codes: G28, G21

Résumé

Nous évaluons dans quelle mesure les coussins en capital (le capital que les banques détiennent au-dessus du minimum réglementaire) amplifient plutôt que réduisent le comportement cyclique du crédit. Nous étudions empiriquement les relations entre la croissance économique, les coussins en capital et la distribution de crédit à partir de données individuelles sur les banques françaises au cours de la période 1993—2009. Nos résultats montrent que le capital bancaire amplifie les fluctuations du crédit résultant du cycle économique, et ce d'autant plus que le capital considéré est de meilleure qualité. Par ailleurs, en conduisant des tests de causalité de Granger au niveau de chaque banque, nous mettons en évidence une double causalité entre coussins en capital et croissance du crédit, ce qui met en évidence des mécanismes se renforçant mutuellement. Dans l'ensemble, ces résultats empiriques plaident pour une réglementation financière contra-cyclique qui se fonde sur le capital de meilleure qualité et qui tend à lisser la croissance du crédit.

Mots clés : Réglementation du capital bancaire, Procyclicité, Coussins en capital, Fluctuations économiques, Bâle 3

JEL codes: G28, G21

1. Introduction

A puzzling fact about the global financial crisis that broke out in the early Summer 2007 is the disproportion between the restrained losses from actual defaults of US subprime borrowers, on the one hand, and the huge write-offs reported by financial institutions as well as the large-scale real effects on the world economy, on the other hand. This sharp discrepancy constitutes a clear exemplification of financial intermediation being an inherently pro-cyclical activity. As explained by Borio *et al.* (2001) and Lowe (2002), during economic expansions, financial institutions are more willing to take risks, credit markets are more prone to competition, credit spreads, risk premia and other measures of risk aversion approach low levels and the access to credit becomes easier as collateral values are rising. Conversely, in recessions, banks and other financial institutions are weaker, more conservative, and credit declines as net worth and collateral values erode, thus exacerbating business cycles.

In the aftermath of the current financial crisis, supervisors and public authorities have closely monitored the transmission channels through which bank distress might spill over to the real economy, worsening the downturn and dampening the recovery to come. One of these mechanisms involves the lending capacity of banks and the fear of a ‘credit crunch’, in which a sharp decrease in bank capital would result in banks squeezing credit distribution to maintain their capital ratios, leading to a credit rationing that would harm economic growth.

Undoubtedly, this is the main reason why capital injections by public authorities were conditionally granted, subject to firm commitments by financial institutions to neither cut dramatically their loan distribution, nor tighten too severely their credit conditions. Within the context of imminent exit strategies – i.e. the fact that public authorities would withdraw the capital they injected in banks’ balance-sheets during the crisis – the underlying pro-cyclical mechanisms are gaining wide and renewed momentum. Since the G20 Washington Summit in November 2008, international political and regulatory bodies have focused attention on pro-cyclicality and policies to mitigate its outsized effects.[5] For instance, the then Financial Stability Forum (FSF), which became the Financial Stability Board subsequently, set up in 2008 three working groups that analyzed the various facets of pro-cyclicality (e.g. FSF, 2009). The Basel Committee on Banking Supervision (BCBS) has been developing policy measures to mitigate pro-cyclicality (BIS, 2008; Andritzky *et al.*, 2009). It is also worth noting that the so-called "Capital Reform Proposal" launched by the BCBS in the mid-December 2009 contains a full package of measures to promote a more countercyclical capital adequacy framework, including the build-up of capital buffers, as well as incentives to implement forward-looking loan loss provisioning based on expected losses (BCBS, 2009). At the international level, a broad consensus has been reached that addressing pro-cyclicality is a key component of a sound macro-prudential policy.

In this paper, we assess the two-way interplays between bank capital buffers, lending and economic growth, and examine the extent to which capital buffers might be considered as procyclical. By definition, capital buffers denote the excess capital banks hold above the minimum regulatory level. In contrast to a *cyclical* variable, which follows and is mainly driven by the business cycle fluctuations, a *procyclical* variable is supposed to drive and magnify the fluctuations of economic activity. The results reported in this paper provide analytical and empirical background to the current policy debate on the introduction of countercyclical capital buffers in the future "Basel III" package and, more generally, on the procyclical impact of the Basel regulatory framework.[6]

[5] See also the Position Paper on a countercyclical capital buffer published by the CEBS in July 2009 (CEBS, 2009).

[6] The Basel Committee's response to the global financial crisis includes a proposal that requires banks to hold (countercyclical) capital buffers above the minimum capital requirements imposed by regulators. According to the Committee, such buffers should be built up during economic expansions and drawn down throughout significant sector-wide downturns. Interestingly, if capital falls below some pre-specified "buffer ranges" -- but it is still above the minimum regulatory level -- the bank would be subject to capital distribution constraints, restrictions on dividend payouts or constraints on employee bonus payments.

The present paper contributes to and extends the existing literature on the procyclicality of bank capital in at least two important ways.

First, the paper reports empirical evidence for French banks, based on both panel data econometric estimations and Granger causality tests. The focus on France is relevant to the procyclicality literature because bank lending is by far the prevailing form of external finance in this country. Consequently, reductions in lending when bank capital is eroded are likely to have more harmful economic effects than in other, market-oriented, financial systems, where borrowers may alternatively tap the financial markets or deal with other financial intermediaries. Moreover, the "procyclical leverage" hypothesis documented by Adrian and Shin (2010) on US data does not seem to hold equally for *all* industrialized countries (Panetta et al., 2009).[7] Particularly, while banks and financial institutions headquartered in the US and the UK do exhibit a significant *positive* correlation between asset prices and (marked-to-market) leverage, in France and some other few countries the correlation is *negative*. Consequently, there seems to be considerable scope for examining the procyclicality of bank capital outside the US, particularly in countries where the relationship between changes in total assets and changes in leverage is reversed.

Second, the paper reports empirical evidence based on both panel data econometric estimations and Granger causality tests. The primary aim of the panel estimations is to assess the build-up of capital buffers throughout the cycle and their impact on bank lending behavior. Precisely, we estimate two relationships: (*i*) the empirical effect of economic growth on capital buffers and (*ii*) the impact of capital buffers on loan growth. An important feature of the present paper is to look not only at the total capital buffer, but also at the buffers composed of higher-quality capital: Tier 1 capital and core Tier 1 capital. We then investigate the Granger-type causality between bank capital buffers and loan growth at the bank level. To our knowledge, the

[7] Adrian and Shin (2010) lend empirical support to the thesis that banks tend to adjust their capital allocation decisions and balance sheets in order to attain some target levels of leverage. Consequently, a negative shock that erodes capital may reduce bank lending and exacerbate the procyclicality.

present paper is the first one to use Granger causality tests to investigate the two-way interplays between bank capital buffers and credit growth. In our view, this approach is relevant because it sheds new light on the main causal links behind the procyclicality hypothesis.

Our main findings reveal that capital buffers and loan growth at the bank level depend on the output gap in a pro-cyclical manner. In addition, we put forward that bank capital buffers – especially the ones related to the purest forms of capital – exacerbate the cyclical developments of credit. Finally, we find evidence of Granger-causality running from capital buffers to credit growth. Overall, the empirical results lend support to a countercyclical financial regulation aiming at smoothing credit growth and focused on better-quality capital.

The remainder of the paper is organized as follows. We provide an overview of the related literature in Section 2. Section 3 describes the data and discusses some stylized facts resulting from simple descriptive statistics. Section 4 elaborates on the econometric strategy and discusses our main results. Section 5 analyzes the causality between bank capital and loan growth. Finally, section 6 concludes and discusses some policy implications.

2. Banks' capital buffers and loan growth: related literature

The macroeconomic consequences of bank capital requirements have been extensively studied since the adoption of the first Basel Capital Accord at the beginning of the nineties. The empirical literature on the relationships between output, bank capital (buffers), and loan growth, can be classified in two broad categories: the first one investigates the determinants of capital buffers and their potential procyclical effects; the second one studies the role of bank capital and other factors in explaining fluctuations in loan growth.

2.1 Determinants of capital buffers

The conventional starting point for studying capital structure in banks and non-financial firms is the Modigliani and Miller proposition, which states that the capital structure does not affect the value of the firm under the standard assumptions of perfect capital markets and no taxes. In the real world, there are significant departures from Modigliani and Miller's assumptions due to taxes, asymmetric information, agency costs, costly financial distress and, more importantly in the case of banks, regulations. Consequently, targeting 'optimal' levels of capital may be value-enhancing.

More precisely in the case of banks, if the implicit subsidies arising from the mispriced financial safety net (barriers to entry, deposit insurance, implicit guarantees…) were large enough, banks may choose to hold the minimum level of equity capital allowed by their regulators. In that case, one would observe little or no cross-sectional heterogeneity in the reported solvency ratios, as the capital adequacy framework imposes uniform minimum standards. Yet, this simplistic view does not hold, as the reported levels of capital are heterogeneous, and generally higher than the regulatory minimum in developed countries. Flannery and Rangan (2008) and Berger et al. (2008) provide convincing evidence for the largest US banks, while Jokipii and Milne (2008), Gropp and Heider (2010) and Brewer et al. (2008) document the levels and cross-sectional variations in the bank capital ratios of internationally active banks.

Why do banks hold so much costly capital over and above the regulatory minimum? What are the main factors explaining the cross-sectional and time variation in bank capital buffers? The literature provides several competing, albeit not mutually exclusive, answers to these questions (Berger et al. 1995; Berger et al., 2008; Flannery and Rangan, 2008). First, banks may hold excess, "precautionary" , capital in order to avoid adjustment costs in raising equity on short notice or supervisory penalties if they approach the regulatory minimum. Second, if the regulatory capital only imperfectly reflects the risk of losing the bank's charter value, capital buffers act as a cushion that protects its going concern value. Third, banks may also prefer to hold capital buffers because they fear being short of funds, should attractive investment opportunities (e.g. profitable acquisitions) arise in the future. Finally, banks may maintain higher capital ratios as a response to disciplinary pressures exerted by private market forces, to gain access to specific OTC markets (e.g. derivatives) or to obtain a targeted credit rating from external agencies.

Although it seems difficult to disentangle these various determinants of bank capital buffers empirically, a better understanding of the main factors driving the formation of these buffers may help shed light on other relevant policy questions. One of these questions is the procyclicality of capital regulations. Basically, during recessions, the bank capital is likely to be eroded by losses on the loan portfolio, as default probabilities increase with the worsening of the macroeconomic environment; thus banks have to hold more regulatory capital, especially under the Basel II standards. If raising new capital is prohibitively expensive because the whole financial system is under stress, the most cost-effective way to rebuild capital ratios is to cut back on lending, thereby amplifying the initial recessionary shock. If banks *naturally* built up capital buffers during good times, in order to better absorb losses under stressful conditions, the procyclicality concerns would be partially offset. If this is not the case, bank capital buffers may move procyclically. Otherwise stated, banks may target *higher* solvency ratios in bad times in anticipation of future uncertainty and losses on their portfolios, which would eventually reduce the loan supply further and exacerbate the recession.

The empirical literature has not reached consensus regarding the procyclical effects of capital buffers. The common approach used in most papers is to assess the impact of the business cycle on the observed capital or capital buffers. In their seminal paper, Ayuso et al. (2004) report a robust and negative relationship between capital buffers and the business cycle for Spanish commercial and savings banks over the period 1986-2000, i.e.

under the Basel I regime.[8] Some other papers report similar *negative* co-movements between capital buffers and the business cycle using data on banks headquartered in individual countries: Lindquist (2004) for Norwegian banks; Stoltz and Wedow (2010) for German savings and cooperative banks; Alfon et al. (2004) and Francis and Osborne (2009) for UK banks and building societies. However, important asymmetries are reported in these studies between low-and highly-capitalized banks, commercial and savings banks, small and large banks and building societies and commercial banks.

Other studies examine the same question in a broader, cross-country setting. Using a large panel dataset on OECD commercial banks, Bikker and Metzemakers (2004) find a moderate relationship between the observed equity capital ratios and the business cycle. In addition, the procyclicality effect exhibits substantial variations across countries and bank-size classes. Jokipii and Milne (2008) complement and extend these findings by conducting a comprehensive empirical analysis on the determinants of capital buffers of European banks. They confirm the negative co-movement of capital buffers with the cycle but with some important caveats. Particularly, for banks in EU accession countries, as well as for cooperative and smaller European banks, capital buffers move counter-cyclically. Finally, Fonseca and Gonzáles (2010) analyze the bank- and country-specific determinants of capital buffers using a larger panel of banking organizations headquartered in 70 developing and developed countries. They confirm the existence of different patterns of capital buffers across countries, after controlling for the cost of deposits, market power, and other relevant explanatory factors. A significant negative relationship between capital buffers and economic cycle is reported only for seven countries. In

[8] One may be tempted to infer that if bank capital buffers move procyclically under the Basel I capital accord, the procyclical effects should be *a fortiori* stronger after the implementation of the more risk-sensitive Basel II capital accord. However, such an inference is subject to the usual Lucas critique: it would be imprudent to draw policy implications concerning the potential procyclicality of Basel II from the observed cyclical patterns of capital buffers under Basel I. Using a dynamic equilibrium model of relationship banking in which business cycle fluctuations affect the borrowers' default probabilities, Repullo and Suarez (2010) show that capital requirements under Basel II have an ambiguous effect on capital holdings. In the same vein, Heid (2007) proposes a different theoretical model and shows that capital buffers under Basel II may actually move counter-cyclically, because the rise in risk weights will more than compensate the reduction in bank lending.

five other countries, the sign of the relationship is reversed, while in the remaining 59 countries the cycle variable does not enter significantly in the capital buffer regressions.

2.2 The impact of capital buffers on loans and the real economy

The empirical studies mentioned above give some evidence of capital buffers co-varying with the business cycle. However, even if buffers move in a highly procyclical manner, this result is only a *necessary*, not a *sufficient*, condition to observe significant procyclicality in the real economy. Another important causal link in the procyclicality chain has to be confirmed, running from capital buffers to bank lending.

Another strand of the procyclicality literature, which we briefly review in what follows, has examined the role of bank capital in explaining fluctuations in loan growth.

The effects of capital requirements on banks' lending behavior over the business cycle have long been documented, not only because of the implementation of the risk-sensitive Basel II framework. Indeed, concerns about the existence of a so-called "bank capital channel," whereby changes in banks' capitalization influence the transmission of business cycle fluctuations on lending, have been expressed since the observed credit crunches in the late eighties and early nineties.[9,10] The earlier literature, carefully surveyed by Sharpe (1995) and Jackson et al. (1999), conclude that, at least in the short run, negative shocks to capital lead low-capitalized banks to cut back on new lending during recessions.

There are two main conditions for the existence of the bank capital channel. First, the market for bank equity is imperfect, that is, banks cannot easily

[9] Another important question examined in the bank capital channel literature is to what extent the introduction of the Basel I risk-based capital accord at the beginning of the nineties caused or exacerbated the subsequent decline in output observed in several developed countries.

[10] Bank capital may also influence the impact of recessionary shocks on loan growth through the "bank lending channel," which is built on imperfections in the market for bank debt. However, as this channel pertains to monetary shocks (e.g. tighter reserve requirements on demand deposits) and the monetary policy transmission mechanisms, rather than output shocks, we do not discuss here the related literature.

issue new equity to finance profitable lending opportunities due to agency costs, information asymmetries, and tax disadvantages (Kashyap and Stein, 1995). Second, banks are subject to regulatory capital requirements and have no excess capital to absorb output shocks. However, bank capital may affect lending conditions even when capital requirements are not binding, if banks fear the risk of breaching the regulatory minimum in the future or want to maintain high credit ratings (Van den Heuvel, 2002). Some papers add a third condition: banks bear an interest rate risk due to the maturity mismatch between their short-term liabilities and their long-term assets (Van den Heuvel, 2002; Gambacorta and Mistrulli, 2004).

Recent papers confirm the relevance of the bank capital channel and show that capital buffers do influence the response of lending to output shocks. For instance, Gambacorta and Mistrulli (2004) rely on a set of Italian banks representing about 80% of aggregate credit and find that well-capitalized banks are in a better position to preserve lending relationship by absorbing temporary difficulties faced by their borrowers. Interestingly, the introduction of capital requirements higher than the standard minimum of 8% for risky banks resulted in a 20% decline of lending after two years. Using a large sample of listed banks in 31 countries, Nier and Zicchino (2005) estimate standard loan growth equations and confirm that loan losses usually lead to a larger decline in credit for banks having smaller capital buffers. Finally, Francis and Osborne (2009) examine whether a change in individual capital requirements imposed on a sample of UK banks influences banks' internal capital targets and, in turn, the lending supply. They carry out two-step estimations, first identifying the determinants of bank capital ratios and then estimating a model of lending growth. They find a positive relationship between capital requirements and banks' targeted capital ratios and between lending growth and *excess* capital (defined with respect to the unobserved target capital ratio). Their results suggest that lending growth is less constrained for banks which hold surplus capital relative to the internal target. They also simulate the impact of a counter-cyclical measure consisting in raising gradually capital requirements by 3%

over the period 2000--2003; this would have dampened loan growth by 20% over the period 2000--2007, when the credit boom fuelled in the UK. In a recent related contribution, Berrospide and Edge (2010) find a significant and positive relationship between lending and capital ratio at the individual level (better capitalized banks grant more credit), whatever the measure of capital ratio used. However, they infer substantially small effects in magnitude of capital-to-asset ratios on lending after estimating both panel regressions for a sample of large US banks and a vector autoregressive (VAR) model. They attribute this result to the fact that US banks were better capitalized and closer to their target capital ratios in 2008 than they were just before the credit crunch of the early nineties.

The literature has mainly focused on the procyclical effects of capital buffers by assuming a significant decline in economic activity caused by a reduction of the credit supply. Fonseca, Gonzáles, and Da Silva (2010) take a different approach by analyzing and testing the potential expansionary effect of capital buffers through a reduction in interest rate spreads. They find that well-capitalized banks charge *lower* interest spreads to their borrowers and also pay *lower*, but safer, interest rate spreads to depositors. Capital buffers appear to have stronger influence on the economic activity through these *price* channels in developing countries during downturns, partially offsetting the procyclical effect.

3. The dataset

3.1 Description of the data

To construct our sample of banks, we rely on a confidential database provided by the French Prudential Supervisory Authority (PSA). We start from an unbalanced panel dataset covering 231 French banks on a consolidated basis over the period 1993-2009, on a yearly frequency. We prefer to use consolidated rather than solo (unconsolidated) data in order to make the prudential data, especially the capital buffers figures, as relevant as possible. More exactly, we decide not to make use of the quarterly Basel II – compliant data, because they would not give us long enough time series. Indeed, the quarterly reports are available for French banks only after 2007. As we are interested in banks with significantly long time series, and in order to be consistent with the selection rule applied in the subsequent Granger causality tests, we restrict our sample to those banks for which we record at least eight consecutive observations for our two dependent variables: capital buffer and loan growth. In addition, as we are interested in the behavior of banks for which granting loans is one of the main activities, we exclude the banks whose loan stock is below 100 million euros, which is a low threshold and thus not too restrictive. Moreover, we remove bank holding companies for credit cooperatives and mutual banks in order to avoid double counting of loans outstanding, which could stem from the fact that regional credit cooperatives report their prudential and balance sheet data to the supervisor on a "sub-consolidated" basis. After cleaning the initial dataset, we end up with 98 banks that represent about 70% of the total bank loan outstanding in 2009, which makes our final dataset representative of the French banking system.

Particular attention is paid to the treatment of bank mergers, which may otherwise distort loan growth. To that end, we use a Prudential Supervisory Authority internal database listing mergers involving French credit institutions from 1993 onwards. For each merger, we build a fictitious bank the year preceding the merger by summing the loan outstanding of the merging parties. This then allows us to compute a loan growth net of the effect of the merger for the year of this event. In some few cases, we do not have any information on the absorbed entities. This is exclusively the case when the latter are very small banks. In that case, we interpolate the loan growth between the year preceding and the year following the merger. We finally end up with 1,305 bank-year observations.

As far as the aggregated series are concerned, output gaps are extracted from the OECD database and are calculated by using a production function approach to derive estimates of potential output[11]. The main refinancing rates are taken from the Banque de France for the 1993-1998 period and from the European Central Bank databases for the 1999-2009 period.

3.2 Descriptive statistics

The 98 French credit institutions included in our dataset can be split into three categories according to their legal status: (i) 21 commercial banks; (ii) 61 mutual, savings banks and credit cooperatives; (iii) 16 financial and investment firms. Table 1a displays some descriptive statistics for all banks and by decomposing the sample into these three categories. The median capital buffer for the whole French banking system amounts to a high value, namely 46% of the regulatory minimum. This figure suggests that most of the time the regulatory constraint is not binding. The buffers are especially high for the financial and investment firms (92.7%). Concerning Tier 1 and core Tier 1 capital buffers, their median largely overcome the regulatory minimums (238% and 456% respectively for the whole sample). Mutual savings banks and credit cooperatives display a very high level of capitalization for these highest quality forms of capital (median figures reaching 270% for Tier 1 and 523% for core Tier 1), in line with the stylized facts reported in the literature (Stoltz and Wedow, 2010; Jokipii and Milne, 2008). This finding may stem from the fact that this kind of banks may have a lower access than commercial banks to debt instruments included in overall capital, such as hybrid securities. Another interesting feature is the high degree of cross-sectional heterogeneity of these levels of buffer, as reflected by relatively high standard deviations.

[11] Potential output is determined as the level of output that results when all factors of production and total factor productivity are at their potential levels. The output gap is then defined as the difference between actual and potential output, expressed as a share of potential output.

Graph 1 enables us to observe the fluctuations of the buffers. It suggests that banks do not target a fixed buffer as the evolution of their level of capital is not correlated with the evolution of their risk-weighted assets (RWA). Another interpretation is that changes in the level of capital buffers are driven both by changes in the level of total capital and in RWAs.

[Table 1a and Graph 1: Descriptive statistics on main bank variables]

Graph 2 illustrates the evolution of the output gap, credit growth and the average capital buffer for the whole banking system. It suggests the existence of cyclical patterns in the evolution of banks' capital buffers and credit growth. The picture is somewhat mixed as the sign of the relationship between those three variables seems to differ depending on the period. At first glance, there seems to be a break in 2000, when the relationship between the average weighted capital buffer on the one hand, output gap and loan growth on the other hand, turns to be negative. Hence, our econometric investigation will allow for an alternative specification estimated on the period prior to 2000.

[Graph 2: Output gap, capital buffers and bank loan growth]

Table 1b provides the correlation coefficients between the means of the variables in our model. We find a slightly negative correlation between the capital buffer and the output gap (-0.02) i.e. a decrease in the output gap would be coincident to a rise in capital buffer, consistent with the intuition of a precautionary behavior by banks. As expected, loan growth and output gap are positively correlated (0.05). Interestingly, there is a negative correlation between the total capital buffers and the loan growth (-0.06); this result is stronger for Tier 1 or core Tier 1 capital buffers, the correlation coefficients being -0.12 and -0.22, respectively These negative correlations are consistent with the idea that a decreasing output gap would be associated with more capital buffers that would in turn slow down loan growth.

[Table 1b: Correlation coefficients between the means of the variables]

4. Model and results

Our purpose is to understand whether bank capital buffers exacerbate the cyclical behaviour of loans that is to say whether bank capital is a transmission channel from output gap fluctuations to credit developments that behaves pro-cyclically. Hence, we estimate two equations: the first one seeks to assess whether the output gap is a determinant of the capital buffers; the second aims at understanding the effect of capital buffer on the loan growth controlling for the output gap. Should a decrease in the output gap lead *in fine* to a decrease in loan growth through capital buffers, then the procyclical effect of the latter would be demonstrated. This is the hypothesis we are testing in this Section.

4.1 Banks' capital buffers equation

In a first step, we estimate a relationship between banks' capital buffer and a set of explanatory variables. The model is expressed as follows:

$$B_{it} = \alpha_0 + \sum_{m=1}^{M} \alpha_m X_{m,i,t} + \varepsilon_{i,t} \,, \qquad (1)$$

where $B_{it} = 100\frac{K_{it} - K_{it}^R}{K_{it}^R}$ is bank *i*'s capital buffer at time *t*, expressed as the relative gap between the actual amount of bank capital K_{it} and the regulatory minimum capital requirement K_{it}^R; α_0 is the intercept; α_m $m=1,\ldots M$, denote the M coefficients common to all banks on the explanatory variables, $X_{m,i,t}$; $\varepsilon_{i,t}$, the residuals of the equation assumed independent and identically distributed.

As we want to test whether the capital buffer depends on the business cycle, the set of explanatory variables includes a variable capturing the macroeconomic conditions in addition to bank-specific variables. Our explanatory variables are as follows:

- the lagged dependent variable, $B_{i,t-1}$ to account for a possible autoregressive behavior of capital buffer for instance due to adjustment costs of capital. Hence, we expect a positive sign;
- the annual return on equity, ROE_{it}. Considered as a proxy for the cost of capital, it is expected to be negatively correlated with capital buffer;
- the ratio of total provisions for loan to total loans, $Prov_{it}$, as a proxy for the internal measure of risk. The expected sign is ambiguous: it may be positive if the decision of a bank to raise capital signals its risk aversion and/or a better capacity to absorb losses in the future. It may also be negative if losses reduce the level of capital;
- the size of the bank, $Size_{it}$, measured by the total assets of a bank minus mean total assets of all banks, both being taken in logarithm at the end of the year. The ratio of each bank's assets to the mean total assets is meant to avoid spurious correlation stemming from a time trend in banks' assets. We expect a negative sign, as big banks have less incentives to constitute capital buffers due to a lower risk aversion, in line with the *too big to fail* hypothesis and due to their higher ability to diversify risks and access funding;
- the output gap,. $G\breve{D}P_t$ The sign of the coefficient determines whether banks constitute precautionary savings in bad times (if the sign is negative), in which case procyclicality may occur or tend to smooth their activities across the cycle (if the sign is positive).

The model to be estimated over a panel of banks is expressed as follows (expected signs in brackets):

$$B_{it} = 100(\frac{K_{it} - K_{it}^{R}}{K_{it}^{R}}) = \alpha_0 + \alpha_1 B_{it-1} + \alpha_2 G\breve{D}P_t + \alpha_3 ROE_{it} + \alpha_4 \mathrm{Prov}_{it} + \alpha_5 Size_{it} + \varepsilon_{it} \quad (2)$$

(+) (?) (-) (?) (-)

In equation (2), our variable of interest is $G\breve{D}P_t$, the other variables stand for control. We use the Arellano–Bover (1995) Generalized Method of Moments (GMM) estimator to take account of several characteristics of our panel: (i) the possible endogeneity of the explanatory variables, especially the lagged dependent variable; (ii) the presence of fixed effects possibly correlated with the explanatory variables; (iii) the short time dimension (*T*=17) and larger cross-section dimension (*N* =98); (iv) the possible autocorrelation of residuals and heteroskedasticity between banks. As we want to account for a possible endogeneity of the ROE, we use as instruments for the differenced equation: the second and third lags of the dependent variable, the first and second lags of the ROE in level and other explanatory variables in difference; for the level equation, the differenced dependent variable, the ROE and other explanatory variables in level. The choice of lags for the instruments related to the dependent variable is driven by the need to avoid too many instruments compared to the number of individuals. Our post estimation diagnosis includes a Sargan test to check the validity of our instruments and a m2 test checking for the non autocorrelation of order 2 of the differenced residuals. Both tests validate our specification: our instruments are exogenous and not correlated with the error term ε_{it}; and the residuals exhibit the expected characteristics. [12]

Results are presented in Table 2. As far as our variable of interest – the output gap – is concerned, we do find a significant and negative coefficient (Table 2, column 1). This result tends to attest that a worsening real economy situation is translated into an increase in bank capital buffers that can be interpreted as an increase in precautionary reserves in bad times.

[12] Results are not reported for the sake of brevity, but are available from the authors upon request.

Note that this effect would pave the way for pro-cyclicality, if more capital buffers were to amplify the cyclical slowdown of loan growth. As for the control variables, the coefficient of the lagged buffer is positive with a significance close to the 10% threshold. Its small magnitude (0.16) suggests a slight autocorrelation of the total capital buffer's level. The coefficient of the size is significant and has the expected negative sign: large banks hold less capital, in line with the *too-big-to-fail* hypothesis. As regards the other two bank-specific variables, namely the return on equity and the ratio of provisions, they prove significant with signs suggesting the following relationships: a higher profitability, reflecting a higher cost of capital, weighs on the total capital buffer; a higher loan loss provisions rate reflects a more careful behavior of the bank, which increases total capital buffers.

[Table 2: Determinants of banks' capital buffers]

To check for robustness, we carry out several alternative estimations. First, as the relationship between bank's capital buffers, bank-specific and macroeconomic variables might be stronger for the purest forms of capital, we substitute successively the Tier 1 capital buffer, B_{it}^{*}, and the core Tier 1, B_{coreit}^{*}, for the total capital buffer in equation (2). The aim is to check whether the different forms of bank capital react differently to the same set of explanatory variables, especially across the cycle. B_{it}^{*} is thus defined as $\frac{K_{TIER1_{it}} - K_{TIER1_{it}}^{R}}{K_{TIER1_{it}}^{R}}$, the regulatory minimum level for the Tier1 ratio being set at 4% of RWAs. B_{coreit}^{*} is defined as $\frac{K_{core_{it}} - K_{core_{it}}^{R}}{K_{core_{it}}^{R}}$, with a regulatory minimum level for the core Tier1 ratio being set at 2% of RWAs.

The estimations confirm and even reinforce those previously obtained concerning a possible procyclicality effect (Table 2, columns 2 and 3). More specifically, the higher the quality of capital, the higher the coefficient of the output gap, still statistically significant. This suggests that the sensitivity of highest-quality capital to the business cycle is higher than that of the total capital. The autoregressive coefficients of Tier 1 and core Tier 1 buffers are greater than that of the lagged total capital buffer, which suggests that banks face higher adjustment costs for the purest forms of capital. Moreover, the coefficients on the other bank-specific variables except the size are not significant, although of the same sign as for the total capital buffer. The coefficient on size proves significantly negative, larger and more significant than for the total capital buffer, suggesting that the *too-big-to-fail* hypothesis softens the capital constraints on large banks especially for the Tier 1 and core Tier 1 capital. Finally, the lower significance of the coefficients between the bank-specific and the macroeconomic variables suggests that macroeconomic conditions are the main drivers of the level of bank capital buffers.

Second, we restrict our estimations to a sub-sample of credit cooperatives and savings banks in order to analyze the effects of the legal form on the relationship between macroeconomic conditions and capital buffer (Table 2, column 4). As credit cooperatives and savings banks are typically smaller than commercial banks and are likely to have a more limited access to funding markets, we expect a higher autoregressive coefficient and a higher sensitivity to the business cycle. Indeed, the autoregressive coefficient is larger than in the estimation on the whole panel (Table 2, column 4 versus column 1). The coefficient of the output gap is of the same order of magnitude, though much more significant. Lastly, the coefficient on the size is significantly negative, as previously. All in all, the results confirm a negative relationship between the output gap and the capital buffer as well as an autoregressive behavior of the capital buffer. These results are particularly large and significant for the purest forms of capital and for cooperative banks.

4.2 Loan growth equation

In a second step, we estimate the relationship between loan growth - at an individual bank level - and a set of explanatory variables including bank-specific and macro variables. All variables are contemporaneous with the exception of the lagged dependent variable. They are as follows:

- the lagged dependent variable, meant to assess the autocorrelation of credit growth;
- the bank capital buffer, B_{it} which is the key variable in this equation, meant to test for procyclicality, as explained below. It also assesses the bank capital channel, i.e. the impact of the level of capitalization of a bank on its supply of loans,
- the ratio of liquidity of the bank, Liq_{it} measured by the ratio of liquid assets to total assets. Liquid assets are computed as the sum of cash, interbank loans and securities held in the trading portfolio and available for sales. A positive sign is expected as the literature has recently shown that liquidity, in addition to solvency, is an important determinant of loan supply, and that liquidity and solvency have large interactions. However, banks may prefer to hoard liquidity during periods of stress for precautionary reasons. In this case, a negative sign is expected;
- the bank's size, $Size_{it}$, as previously defined, used as a proxy for the magnitude of adverse selection problems faced by banks when raising uninsured finance due to information asymmetry, along the lines of Kashyap and Stein (1995). The latter found evidence that small banks cut loans by more in response to external shocks. The expected sign of this variable is thus positive, as the loan supply of large banks is expected to be more insulated from external shocks across the business cycle;
- the output gap, $G\breve{D}P_t$, for which we expect a positive sign, signaling the banks' risk aversion: banks are thought to increase loans when risks are perceived to be weak, namely in an expansion and reduce them in recessions;

- the refinancing rate of the central bank, r_t, for which we expect a negative sign since this variable represents the cost of bank refinancing.

Therefore, our model is expressed as follows:

$$\Delta \log c_{it} = \beta_0 + \beta_1 \Delta \log c_{it-1} + \beta_2 B_{it} + \beta_3 Liq_{it} + \beta_4 Size_{it} + \beta_5 G\breve{D}P_t + \beta_6 r_t + u_{it} \quad (3)$$

(+) (?) (+) (+) (+) (-)

where β_k are parameters to estimate, β_0 being an intercept and u_{it} is the residuals.

Our variable of interest in this estimation is the bank capital buffers. Its sign will be decisive for testing the hypothesis of procyclicality. Let's assume a negative shock on the output gap. This would result in an increase in the capital buffers according to our previous results in Section 4.1. If after controlling for output, that increase in the capital buffers enhances loan growth (positive sign), lending behavior would act counter-cyclically, mitigating the effects of the initial shock. On the contrary, if that increase in capital buffers results in a decrease in loan growth (negative sign), capital buffers would amplify the initial shock, paving the way for procyclicality.

Note that a positive sign is generally reported in the literature on US banks. It is interpreted as consistent with the bank capital channel hypothesis: well-capitalized banks should be less constrained under stress conditions and would not restrict the credit supply in order to maintain lending relationships. However, the evidence on the impact of bank capitalization on credit growth is somewhat mixed. First, in their seminal paper, Berger and Udell (1994) find that in the particular case of commercial real estate lending and two other credit sub-categories, the decline in growth rates of loans for well-capitalized banks was actually larger than for low-capitalized banks. This finding is inconsistent with the idea that capital constraints were the main driver of the credit crunch during the nineties and implies a negative sign for the coefficient of the capital buffer in the credit growth equation. Second, using micro-level data on German banks, Stolz and Wedow (2010) find that low-capitalized banks actually do not reduce lending during economic downturns. Again, this behavior contrasts with the results commonly reported in the US literature. Finally, anecdotic evidence indicates that despite huge injections of public funds in the largest banks since the beginning of the current financial crisis, bank loans have dried up at a rapid pace in the vast majority of developed countries. Instead of using the public funds to sustain lending to the real sector, the largest banks decided to boost their liquidity buffers and capital ratios perhaps in anticipation of future losses on their asset portfolios. This behavior is consistent with the results reported by Frame et al. (2009), who fail to find a significant effect of the public fund injections on the loan supply of the largest US banks. In addition to a statistical assessment of the effect of bank capital on lending behavior, we also allow for different measures of capital buffer depending on the capital's quality, namely total capital buffer, Tier 1 capital buffer and core Tier 1 capital buffer.

We use the same econometric methodology as previously (a dynamic GMM model *à la* Arellano-Bover). As GMM instruments, we chose the dependent variable (the second to fourth lags in levels for the difference equation and the second and further lags in difference for the level equation) and the buffer that we consider endogenous consistently with the results of Section 4.1 (the first to third lags in levels for the difference equation and the first and further lags in difference for the level equation); as standard instruments we rely on the other explanatory variables in first difference for the difference equation. We also perform an AR(2) test and an m2 Sargan test which both validate our specification.

The main results are presented in columns (1) to (3) of Table 3. The coefficient on the output gap is significantly positive whatever the measure of capital buffer considered: when the output gap increases by one percentage point, loan growth increases by about 3 percent. The liquidity ratio is not found to have a significant impact on loan growth[13].

More importantly, in contrast with the results commonly reported in the literature concerning US banks including the most recent papers (see e.g. Berrospide and Edge, 2010) the coefficient of the buffer is found negative. Consequently, when the buffer increases, banks supply less loans, even after controlling for the output gap. This result is all the more significant as the capital buffer considered is of higher quality, should either the magnitude of the coefficient or the statistical significance of the coefficient be considered. This result is somewhat conflicting with the functioning of the bank capital channel as it has been studied until now and with the findings of most papers on this topic. It is of crucial importance as it signals that an economic downturn that would lead banks to increase their buffers (part 4.1) would result in a decreasing loan growth even after controlling for macroeconomic factors. From this perspective, capital buffers have a procyclical effect. Interestingly, if we consider either the effect of the output gap on the capital buffers or that of capital buffers on the loan growth, results prove much more significant when capital of better quality is looked at. This means that the procyclical effect aforementioned is especially true for Tier 1 and core Tier 1 capital. This results may signal a specific feature of French banks, namely a weak bank capital channel, as stated for instance by Jimborean and Mesonnier (2010). It may also capture more general features of bank lending, as it is consistent with some recent theoretical findings. Indeed, Valencia (2010) shows that banks facing higher uncertainty may prefer to keep higher capital-to-asset ratios and to deleverage, due to precautionary motives that are presumably much stronger during recessions. The magnitude of those effects is not negligible: for example, an increase in the core Tier 1 capital by 1 percentage point (from the lower bound of 2% of

[13] However, as a robustness check, we also used an alternative measure of liquidity given by the deposit-to-credit ratio. In this case, we find a significant and negative coefficient, which tends to confirm that during recessions banks may choose to hoard liquidity for precautionary reasons rather than lending to the private sector.

risk-weighted assets to 3% for instance) would lower loan growth by 5% all other variables kept equal.

We carry out additional estimations to better check the results. First, we restrict our sample to the 1993-2000 period, i.e. before the euro cash changeover, in order to account for potential structural breaks (as observed in Section 3): results remain unchanged as regards our variable of interest both in terms of magnitude and significance (Table 3, column 4), though lagged loan growth appears more important in explaining current loan growth. Likewise, we exclude the 2007-2009 period to check whether our results are driven by the effects of the financial crisis and the simultaneous implementation of Basel 2: results remain unchanged, which suggests that they are robust to changes in the period of observation. Second, as our results might also be driven by the behavior of large and risky banks, we add an interaction term between the capital buffer and the loan loss provision (Table 3, column 5). We do find a negative and slightly significant coefficient for that variable. This result can be interpreted in the following way: riskier banks which exhibit a higher loan loss provisions *or* banks adopting a safer forward-looking behavior do amplify the procyclical effect of bank capital buffers.

[Table 3: Estimation of loan growth]

In conclusion, the econometric investigation carried out in Section 4 shows that capital buffers amplify the cyclical behavior of loans caused by the output gap. This result is all the more relevant as we focus on the Tier 1 and the core Tier 1 capital, that is to say its purest components. In the current context of discussions of the future Basel III framework, those results do support the view that an efficient macroprudential regulation should aim at smoothing credit growth and, having this objective in mind, bank capital is a relevant instrument. In that respect, according to our results, regulating the purest components of capital is the most efficient way to dampen the procyclicality of banks' capital buffers.

5. Granger causality tests

We now check for the causality between capital buffers and loan growth. Should we find that the capital buffers "cause" the credit cycle, this would allow us to validate the procyclicality hypothesis.

5.1 Methodology

Standard Granger causality tests are based on time-series estimations. Variable x_t is said to "cause" variable y_t if the lagged values of x_t improve the forecast of y_t. Under the usual assumptions of stationarity of the series, the standard version model is the following:

$$y_t = \sum_{k=1}^{K} \alpha_k y_{t-k} + \sum_{k=1}^{K} \beta_k x_{t-k} + \mu + u_t \qquad (4)$$

where α_k, β_k and μ are parameters to estimate, K the optimal number of lags in the regression, u_t the residual of the equation. The causality test comes down to estimate Equation (4), on a given time period of length T, and then test for the nullity of all the coefficients on the lagged values of x_t. Generally, the estimation is run as a bivariate vector auto-regression (VAR), as the two senses of causality are searched for simultaneously. The null hypothesis $H0$ is that of no causality: $H0: \beta = 0$, where $\beta = (\beta_1,.., \beta_K)$ is the vector of the lagged coefficients β_k. The nullity of all the coefficients β_k is tested through a Wald test.

As we deal with panel data with a small time dimension (T= 17 at maximum), standard individual tests of $H0$ are not powerful. Hence, we use a panel-causality test proposed by Hurlin (2005, 2008). We will test for causality for each bank individually, running N individual regressions as Equation (4). Let us re-write Equation (4) to take into account the cross-section dimension of the panel:

$$y_{it} = \sum_{k=1}^{K} \alpha_{k,i} y_{i,t-k} + \sum_{k=1}^{K} \beta_{k,i} x_{i,t-k} + \mu_i + u_{it} \tag{5}$$

where α_{ki}, β_{ki} and μ_i are parameters to estimate, allowed to be different across individuals, K a given number of lags common to all individuals, and u_{it} the residuals of the equation. Following Hurlin (2005), we test for the homogeneous non-causality. The null hypothesis is that there does not exist any individual causality.

$$H0 : \beta_i = 0, \forall i = 1,.., N \,. \tag{6}$$

where $\beta_i = (\beta_{i1},..,\beta_{iK})$ is the vector of the coefficients β_{ik} . Therefore, rejecting the null means that there exists at least one individual for which there is causality. The alternative hypothesis can be specified as the following:

$$H(1) \quad \begin{array}{l} \beta_i = 0, \forall i = 1,.., N_1 \\ \beta_i \neq 0, \forall i = N_1,.., N \end{array} \quad 0 \leq N_1 < N \tag{7}$$

First, we calculate the individual Wald tests W_i for $\beta_i = 0$ in the N estimations of Equation (5). Then, we compute the mean Wald test for the panel: $\overline{W_N} = \frac{1}{N}\sum_{i=1}^{N} W_i$. This statistic $\overline{W_N}$ converges towards a χ^2 (K), when the time dimension of the panel T tends towards infinity, whereas it is not appropriate because of its low power for panels with a small time dimension. In the case of small time dimension, Hurlin (2005, 2008) recommends using the following statistics $\tilde{Z}_N$:

$$\tilde{Z}_N = \sqrt{N}\left[\overline{W_N} - \frac{K}{N}\sum_{i=1}^{N}\frac{(T_i - 2K - 1)}{(T_i - 2K - 3)}\right]\left[2\frac{K}{N}\sum_{i=1}^{N}\frac{(T_i - 2K - 1)^2 (T_i - K - 3)}{(T_i - 2K - 3)^2 (T_i - 2K - 5)}\right]^{-\frac{1}{2}} \tag{8}$$

$\tilde{Z}_N$ is shown to follow a normal distribution N(0,1), when the cross-section dimension tends to infinity. Monte-Carlo simulations show that the power of this test is high even for small panels (Hurlin, 2005); it is close to 1 as soon as there are more than 25 individuals in the sample, even for time dimension as small as 10.

5.2 Results at the bank level

We test for Granger-causality between banks' capital buffers B_{it}, and the loan growth Δc_{it} by implementing the panel-causality test proposed by Hurlin (2005, 2008) and described in the previous section. The sample is the same as in the previous section, composed of 96 banks[14] on period 1994-2009. If the procyclicality hypothesis of capital buffers holds, the causality would run from the capital buffers to the loan growth. In this case, we expect the null hypothesis of no causality to be rejected.

We check that the series are stationary. We test the hypothesis of a unit-root by using the standard panel unit root tests: Levin, Lin and Chu (2002); Im, Pesaran and Shin (2003) as well as augmented Dickey-Fuller and Phillips-Perron tests (Table A1 in the Appendix). The results of every test show that both series (buffer and loan growth) are stationary.

The VAR is expressed as follows:

$$\begin{cases} B_{it} = \sum_{k=1}^{K} \alpha_{1ik} B_{it-k} + \sum_{k=1}^{K} \beta_{1ik} \Delta c_{it-k} + \mu_{1i} + u_{1it} \\ \Delta c_{it} = \sum_{k=1}^{K} \alpha_{2ik} B_{it-k} + \sum_{k=1}^{K} \beta_{2ik} \Delta c_{it-k} + \mu_{2i} + u_{2it} \end{cases} \quad (9)$$

When running the individual VARs, we get the same optimal number of lags K=1, for all banks using either the Akaike or Schwarz criteria. Therefore, we fix the common number of lags K to 1.

Results show that causality runs in both directions over the whole panel (Table 4, columns 1 to 2). This finding suggests mutually reinforcing effects between the two variables but might result from a common factor, as the two variables are affected by the output gap. A look at the coefficients in the VAR confirms the negative relationship between the capital buffer and credit growth found in our previous GMM estimations. A deeper analysis by

[14] Two banks have been removed from the sample because of an insufficient number of data points.

category of banks reveals that, for the largest category of credit institutions, namely mutual banks, the causality runs only from capital buffer to the credit growth, suggesting that lending decisions depend more strongly on the level of capital for this category, for previously mentioned reasons (lower size, more limited access to funding and debt markets).

[Table 4: Granger causality tests on capital buffer and loan growth – panel level]

To check the robustness of our results, we first carry out the Granger causality tests dropping successively one of the 96 banks (with replacement) included in the sample. This procedure reveals that the test statistics is robust to the exclusion of any bank.

Second, as we noted in our GMM estimations that the effect of high quality capital buffers on loan growth was higher than the effect of low quality capital buffer, we re-run Granger causality tests by substituting core Tier 1 capital buffer for total capital buffer (Table 4, columns 3 to 4). Results confirm the previous findings while being more significant, in particular when splitting the sample into the different categories of credit institutions. This finding indicates that high quality capital level may be a more important driver of banks' lending decisions than total capital.

6. Conclusion

The current financial crisis has revealed an intrinsic feature of the financial system that bankers and other market participants, as well as policy makers, seem to have forgotten after a long period of unusually stable macroeconomic conditions, suggestively labeled by many "the great moderation." Namely, banks and other financial intermediaries may act in a highly pro-cyclical manner, thereby exacerbating rather than reducing the business cycle fluctuations. One source of pro-cyclicality that has been the focus of intensive debate in the public policy arena since the inception of the crisis is the capital adequacy regulations. In its official response to the financial crisis, the Basel Committee emphasizes that addressing procyclicality should be a key element of a sound macro-prudential policy. Particularly, a great importance is attached to the idea that banks should build up "capital buffers" during expansions to better absorb the shocks throughout significant sector-wide downturns.

The present paper contributes to the post-crisis banking literature on the pro-cyclicality by presenting novel bank-level evidence from France on the two-way interplays between three essential variables at the core of the amplification mechanism: capital buffers, credit growth and output. Our empirical approach is based on both panel data econometric estimations and Granger causality tests within a unified, integrated, framework. We find that French banks' capital buffers, as well as credit growth at the individual level, depend on the output gap in a pro-cyclical manner. The results obtained by performing Granger causality tests strengthen the case for the importance of a countercyclical policy framework as we find evidence of a two-way causality between capital buffers and loan growth, pointing to mutually reinforcing mechanisms.

The focus on a single country, where bank lending is by far the prevailing form of external finance, provides a cleaner analytical context in which to examine timely research questions related to the procyclicality of bank capital. One of the most intriguing results we report is the negative effect of the capital buffer on loan growth. This effect, albeit in contrast with the results reported in the empirical literature, matches the precautionary hoarding of liquidity and massive deleveraging observed since the deepening of the current financial crisis. It is also in line with theoretical models such as Valencia (2010). The idea that capital buffers may reduce, rather than sustain, the credit supply following a severe recessionary shock should be further investigated in prospect of designing a future countercyclical capital adequacy framework.

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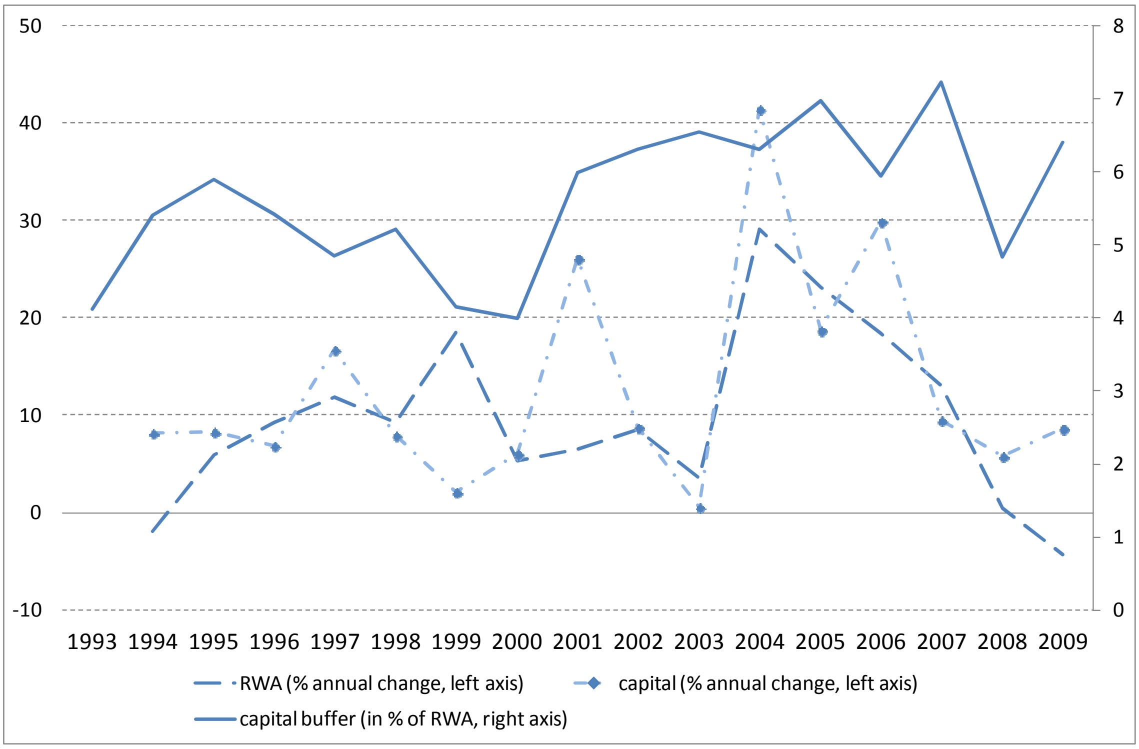
Graph 1: Time profile of the (unweighted) mean of bank variables



Graph 2: Cyclical developments in capital buffers and loan growth

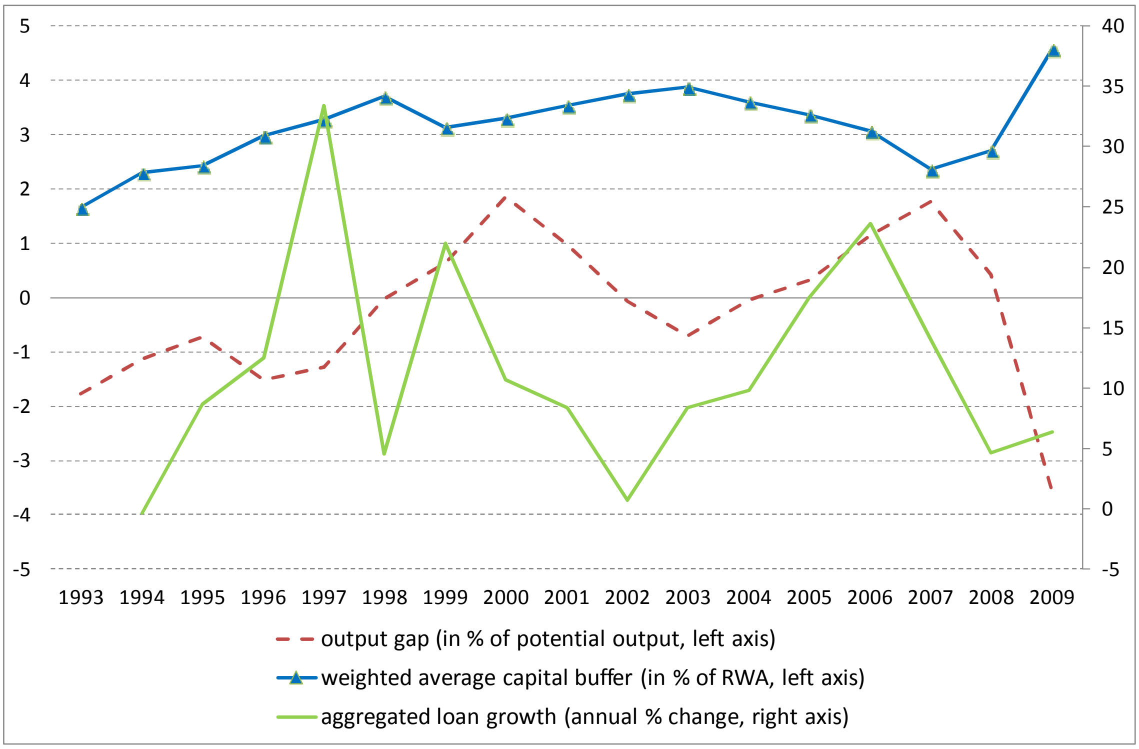


Table 1a: Descriptive statistics on main bank variables

1993-2009	All Banks	Commercial banks	Mutual, savings and cooperative banks	Financial and investment firms
Number of institutions	98	21	61	16
Observations	1,332	310	804	218
Total Capital Buffer, in % of regulatory minimum				
Mean	63.3	63.6	55.5	92.7
Median	45.9	37.3	44.8	74.0
Std.	64.7	68.5	54.6	84.3
Min	1.4	1.5	1.5	1.4
Max	514.3	376.2	514.3	491.9
Tier1 Capital Buffer, in % of regulatory minimum				
Mean	269.2	211.1	293.5	261.1
Median	238.2	152.5	270.2	212.6
Std.	159.2	172.5	139.5	186.6
Min	27.7	30.3	37.7	27.7
Max	989.7	907.2	976.6	989.7
Core Tier1 Capital Buffer, in % of regulatory minimum				
Mean	491.7	421.5	531.1	444.4
Median	456.0	322.1	523.1	370.5
Std.	262.5	324.4	220.6	280.5
Min	49.8	58.9	49.8	82.9
Max	1,732.1	1,627.5	1,732.1	1,435.2
RoE, in %				
Mean	9.2	5.6	9.3	13.7
Median	8.4	7.2	8.8	8.2
Std.	3.0	13.3	12.4	67.8
Min	-98.5	-98.5	-68.1	-46.3
Max	995.3	411.1	217.7	995.3
Ratio of provisions, in % of loans				
Mean	2.1	2.7	1.3	4.2
Median	0.8	1.2	0.7	1.3
Std.	4.0	4.9	1.8	6.7
Min	0	0.0	0	0
Max	46.1	46.1	33.5	42.3
Size (distance to the mean)				
Mean	0	0.6	0.1	-1.1
Median	-0.2	0.3	-0.1	-1.4
Std.	1.6	2.3	1.1	1.6
Min	-4.0	-4.0	-2.3	-3.5
Max	5.1	5.1	4.8	3.1
Liquidity ratio, in % of total assets				
Mean	20.9	27.6	20.1	14.3
Median	16.2	23.4	16.1	7.9
Std.	16.4	19.8	13.6	17.1
Min	0	0.6	1.1	0.1
Max	81.1	78.7	81.1	71.4
Loan growth, in %				
Mean	7.2	7.0	7.8	5.3
Median	7.0	5.7	7.4	5.4
Std.	11.2	14.1	7.1	17.5
Min	-32.4	-32.4	-18.7	-31.8
Max	79.8	79.8	66.9	77.1

Table 1b: Correlation coefficients between the means of the variables

	Total capital buffer	Tier 1 capital buffer	Core Tier 1 capital buffer	Size	Provision ratio	ROE	Liquidity ratio	Loan growth	Output gap	CB's interest rate
Total capital buffer	1	0.71	0.51	-0.24	0.32	0.08	0.14	-0.06	-0.02	-0.04
Tier 1 capital buffer		1	0.84	-0.34	0.25	0.10	0.12	-0.10	-0.01	-0.15
Core Tier 1 capital buffer			1	-0.40	0.26	0.02	0.22	-0.12	-0.01	-0.14
Size				1	-0.08	-0.05	0.07	0.10	-0.02	0.01
Provision ratio					1	0	0.33	-0.08	0.03	0.04
ROE						1	-0.09	0.02	0.01	-0.07
Liquidity ratio							1	0.03	-0.04	0.08
Loan growth								1	0.05	-0.03
Output gap									1	0.47
CB's interest rate										1

Table 2: Determinants of banks' capital buffers

Explanatory variables	Exp. sign	(1) Total buffer GMM All	(2) Tier 1 buffer GMM All	(3) Core Tier 1 buffer GMM All	(4) Total buffer GMM Mutual banks
B_{it-1}	+	0.16	0.67***	0.54***	0.36***
		(1.50)	(5.40)	(3.77)	(3.29)
$G\breve{D}P_t$	?	-3.49**	-9.52***	-11.01***	-3.32***
		(-2.24)	(-3.98)	(-2.67)	(-4.30)
ROE_{it}	-	-0.20***	-0.11	-0.21	-0.12
		(-3.79)	(-0.83)	(-1.05)	(1.01)
$\Pr ov_{it}$	?	6.51**	10.25	24.82	6.97
		(1.70)	(1.38)	(1.53)	(1.27)
$Size_{it}$	-	-0.41**	-0.66***	-1.09***	-0.78***
		(-2.09)	(-3.23)	(-3.50)	(-4.70)
c	+	0.51***	0.83**	1.98***	-3.32***
		(4.41)	(2.26)	(-2.66)	(-4.30)
Number of observations		1,228	1,228	1,228	428
Number of banks		98	98	98	44
Number of estimated coefficients		6	6	6	6
Number of instruments		93	93	93	33
Sargan test (p-value)		0.22	0.23	0.22	0.70
Autocorrelation test AR(2) (p-value)		0.86	0.25	0.50	0.12

Note: *** significant at the threshold of 1 %, ** 5%; * 10 %; t-statistics in brackets

GMM equations estimated as a dynamic panel with orthogonal deviation estimation and White period weights (Arellano-Bover 2-step). All t-statistics use White period robust standard errors. List of instruments differenced equation: 2nd and 3rdlags of buffer, 1st and 2nd lags of ROE in level, other explanatory variables in first difference; level equation: differenced dependent variable and ROE and other explanatory variables in level.

Table 3: Estimation of loan growth

		(1)	(2)	(3)	(4)	(5)
Explanatory variables	Exp. sign	Loan growth GMM All	Loan growth GMM All	Loan growth GMM All	Loan growth GMM 1993-2000	Loan growth GMM All
B_{it}_total	?	-0.03				
		(-1.23)				
B_{it}_Tier1	?		-0.03**			
			(-2.03)			
$B_{it}\_core\_Tier1$	?			-0.05***	-0.02*	-0.04***
				(-4.62)	(-1.86)	(-3.51)
$G\breve{D}P_t$	+	2.73***	2.66***	3.25***	4.18*	3.30***
		(8.39)	(5.58)	(7.99)	(1.73)	(5.93)
r_t	-	-2.32***	-3.36***	-4.10***	-4.26***	-4.02***
		(-3.43)	(-6.38)	(-6.28)	(-3.42)	(-8.07)
$\Delta \log c_{it-1}$	?	-0.12	-0.13	-0.22*	-0.53***	-0.21**
		(-0.74)	(-0.86)	(-1.70)	(-3.47)	(-1.58)
Liq_{it}	+	- 0.14	-0.12	0.17	-0.44	0.22**
		(-0.66)	(-0.60)	(1.01)	(0.81)	(2.05)
$Size_{it}$	+	0.09*	0.05	0.05	0.11	0.06
		(1.84)	(0.93)	(1.17)	(1.60)	(1.27)
$B_{it} * \Pr ov_t$	-					-0.24*
						(-1.66)
constant		0.19***	0.30***	0.44***	0.22**	0.36***
		(4.22)	(4.09)	(6.86)	(1.99)	(5.80)
Number of observations		1,133	1,133	1,133	340	1,133
Number of banks		98	98	98	91	98
Number of estimated coefficients		7	7	7	7	8
Number of instruments		115	115	115	43	115
Sargan test (p-value)		0.77	0.75	0.81	0.16	0.83
Autocorrelation test AR(2)		0.78	0.57	0.05	0.13	0.09

Note: *** significant at the threshold of 1 %, ** 5%; * 10 %; t-statistics in brackets

Table 4: Granger causality tests on total capital buffer and loan growth – Panel level

(1) Null hypothesis	(2) Hurlin Z-stat.	(3) Null hypothesis	(4) Hurlin Z-stat.
WHOLE PANEL			
Capital buffer does not Granger cause loan growth	1.88*	Core Tier 1 capital buffer does not Granger cause loan growth	1.83*
	(0.06)		(0.07)
Loan growth does not Granger cause Capital buffer	2.85***	Loan growth does not Granger cause core Tier1 capital buffer	1.81*
	(0.00)		(0.07)
MUTUAL, SAVINGS BANKS AND CREDIT COOPERATIVES			
Capital buffer does not Granger cause loan growth	1.69*	Core Tier 1 capital buffer does not Granger cause loan growth	2.42**
	(0.09)		(0.02)
Loan growth does not Granger cause Capital buffer	1.45	Loan growth does not Granger cause core Tier1 capital buffer	0.62
	(0.15)		(0.53)
COMMERCIAL BANKS			
Capital buffer does not Granger cause loan growth	1.00	Core Tier 1 capital buffer does not Granger cause loan growth	0.48
	(0.32)		(0.63)
Loan growth does not Granger cause Capital buffer	2.19**	Loan growth does not Granger cause core Tier1 capital buffer	2.56***
	(0.03)		(0.01)
FINANCIAL COMPANIES			
Capital buffer does not Granger cause loan growth	0.26	Core Tier 1 capital buffer does not Granger cause loan growth	-0.53
	(0.79)		(1.41)
Loan growth does not Granger cause Capital buffer	1.55	Loan growth does not Granger cause core Tier1 capital buffer	0.16
	(0.12)		(0.87)

Note: We reject the null hypothesis at the confidence threshold of *** 1 %, ** 5%; * 10 %.
Figures in brackets are p-values. Akaike and Schwarz criteria indicate an optimal lag equal to 1 (annual data).

Appendix

Table A1: Panel unit root tests [1] [2]

Series	Levin, Lin and Chu (t-stat)		Im, Pesaran and Shin (W-stat)		Augmented Dickey-Fuller (Fischer Chi^2)		Phillips-Perron (Fischer Chi^2)	
	H0= common unit root [3]		H0= individual unit root [3]					
	stat.	p-value	stat.	p-value	stat.	p-value	stat.	p-value
Buffer	-6.46	0.00	-4.32	0.00	300.77	0.00	293.33	0.00
Loan growth	-103.57	0.00	-19.54	0.00	451.05	0.00	488.35	0.00

Notes: [1] Lags are selected by Akaike criterion.
[2] The tests include an individual intercept.
[3] The null hypothesis is rejected when p-value < 0.05.

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