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Two-way interplays between capital buffers and credit growth: Evidence from French banks

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We assess the extent to which bank capital buffers exacerbate the cyclical behavior of credit. We empirically study the relationships between GDP growth, capital buffers and loan growth with firm-level data for French banks over the period 1993–2009. Based on panel data simultaneous equations and Granger causality tests, our findings point to mutually reinforcing mechanisms between capital buffers and loan growth, all the more as better quality capital is considered. Overall, those empirical results lend support to a counter-cyclical financial regulation focused on high-quality capital and loan growth smoothing.

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1. Introduction

The current financial crisis has highlighted a number of weaknesses in the banking regulation. In particular, changes in bank capital may have had pro-cyclical effects amplified by the Basel II risk-sensitive capital rules. Indeed, bank capital requirements are likely to rise with increasing risk in economic downturns, at a time when capital is more difficult to raise, which may result in a credit contraction and finally a deepening of recessions. To address this problem, the Basel Committee has recently advocated that banks should build up countercyclical capital buffers in good times in addition to their minimum capital requirements in order to be able to draw down on these surpluses in bad times (BIS, 2008; BCBS, 2010). This measure is complemented by more forward-looking provisioning requirements as provisioning policies are also likely to be pro-cyclical (Bikker and Metzmakers, 2005; Bouvatier and Lepetit, 2008).

In principle, banks may have private incentives to constitute “capital buffers” i.e. capital in excess of their regulatory requirements (Berger et al., 1995, 2008; Flannery and Rangan, 2008): (i) to avoid adjustment costs in raising equity on short notice or penalties if their capital ratio falls under the regulatory minimum; (ii) to avoid cash shortfall, should attractive investment opportunities (e.g. profitable acquisitions) arise in the future; (iii) through market discipline, to increase confidence from shareholders, depositors, as well as rating agencies; and (iv) to protect their going-concern value, in case the regulatory capital is not seen as sufficient.

The key question is whether these capital buffers are built in a pro-cyclical way, i.e. decreased during good times and increased in bad ones. If so, they would not be able to dampen the losses during recessions and would amplify the risks of credit restrictions, thus contributing to worsening output fluctuations. Empirical results are mixed in the economic literature. Although a majority of studies document this pro-cyclical effect, i.e. a negative link between buffers and the business cycle (see for example Ayuso et al. (2004) for Spanish banks; Lindquist (2004) for Norwegian banks; Marcucci and Quagliariello (2008) for Italian banks; Stolz and Wedow (2011) for German savings and cooperative banks; Alfon et al. (2004), Francis and Osborne (2009) for UK banks, Tabak et al. (2011) for Brazilian banks and Deriantino (2011) for Indonesian banks), the results appear to differ across banks, whether they are small or large, poorly or highly capitalized, and even between commercial and savings banks.

Another strand of the literature provides complementary results based on a cross-country perspective. Using a large panel dataset on OECD commercial banks, Bikker and Metzmakers (2004) find a weak relationship between the observed equity capital ratios and the business cycle, with substantial cross-sectional variation. Jokipii and Milne (2008) evidence a positive relationship between the business cycle and capital for banks in European Union accession countries, as well as for cooperative and smaller European banks. Fonseca and González (2010) find a non-significant effect of the business cycle on capital buffers across countries, for 59 out of 71 countries, while negative for seven of them and positive for five others.

Nonetheless, the studies mentioned above focus on one side of the debate only. Indeed, for capital buffers to have a pro-cyclical effect, two conditions should be met: (i) they should co-move with the business cycle; (ii) they should positively influence bank lending, and hence exacerbate the business cycle. That is why another strand of the pro-cyclicity literature has examined the role of bank capital in explaining fluctuations in loan growth. Here again, the results vary across studies. On the one hand, Sharpe (1995) and Jackson et al. (1999) conclude that negative shocks to capital lead low-capitalized banks to cut back on new lending during recessions, their results being confirmed by more recent papers such as Gambacorta and Mistrulli (2004), Nier and Zicchino (2005), Francis and Osborne (2009), Berrospide and Edge (2010). Moreover, this effect is reinforced by new evidence showing that credit is a key determinant for output growth in European countries (Rondorf, 2012). On the other hand, Berger and Udell (1994) find that the decline in loan growth was actually larger for well-capitalized banks in the particular case of commercial real estate lending and two other credit sub-categories in the US. Stolz and Wedow (2011) show as well as that low-capitalized banks in Germany do not reduce lending during economic downturns. These results relate to those put forward by Frame et al. (2009) who find no significant effect of the public fund injections on the loan supply for the largest US banks.

In this paper, we revisit the issue of pro-cyclicity using a rich and novel dataset on French banks coming from the French Prudential Supervisory Authority. First, we investigate the build-up of capital

buffers throughout the cycle and its impact on bank lending through panel estimations. Second, we investigate the Granger-type causality between bank capital buffers and loan growth at the bank level. In both estimations, we check the robustness of the results to the change in the definition of capital (total capital, Tier 1 and core Tier 1), as banks may rely on different types of capital through the business cycle.

The rest of the paper is organized as follows. Section 2 describes the data and discusses stylized facts resulting from simple descriptive statistics. Section 3 presents the econometric estimations and discusses the results. Section 4 analyzes the causality between bank capital and loan growth. Section 5 concludes.

2. A first look at the data

2.1. The data

To construct our sample of banks, we rely on an internal database provided by the French Prudential Supervisory Authority. 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 buffer figures, as relevant as possible, solvency regulations being applied in France at the consolidated level. Moreover, we decide not to use the quarterly Basel II-compliant data, because they would not give us long enough time series. Indeed, quarterly reports are available for French banks after 2007 only. As we are investigating banking behavior over a full business cycle and for the sake of consistency with the selection rule applied in the subsequent Granger causality tests, we restrict our sample to those banks for which at least eight consecutive observations are recorded for our two dependent variables: capital buffer and loan growth. In addition, as we are interested in the behaviour of banks for which lending is a major activity, we exclude the banks whose loan stock is below 100 million euros, which correspond to the 2nd percentile of the loan stock variable. 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, which represent about 70% of the total bank loan outstanding in 2009. This makes our final dataset representative of the French banking system.

We pay particular attention to the treatment of bank mergers, which may otherwise distort the loan growth series. To this end, we rely on a Prudential Supervisory Authority internal database of 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 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 end up with 1228 bank-year observations. Finally, we remove the remaining outliers by winsorizing the loan growth and risk-weighted assets-to-total assets series.²

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.

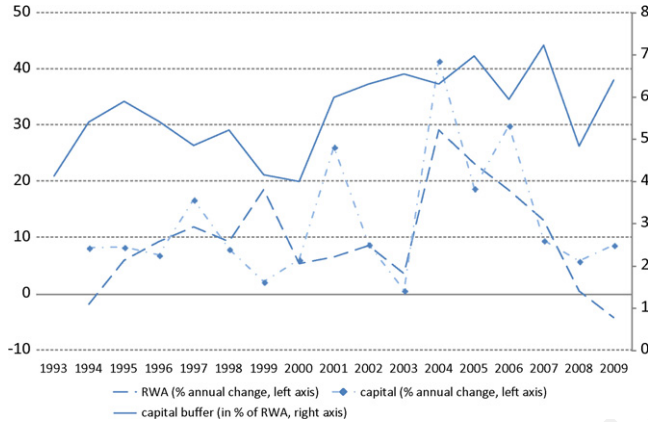
2.2. Descriptive statistics

Table 1 displays some descriptive statistics for the whole panel and the different categories of banks. Two main stylized facts emerge from this table. First, the high levels of the capital buffers show that most banks hold far more capital than the minimum requirements. Second, most of the variations in capital buffer are cross-sectional.

² The winsorization levels are 1% and 99% for the loan growth variable and 99% for the risk series.

Table 1
Descriptive statistics on the variables used in the model.

	1993–2009		All banks		Commercial banks		Mutual, savings and cooperative banks		1993–2009		All banks		Commercial banks		Mutual, savings and cooperative banks		Financial and investment firms	
	Number of institutions	Observations	98	21	21	61	16	16	Number of institutions	Observations	98	21	21	61	16	16	16	
Total capital buffer, in % of regulatory minimum			1228	287	287	745	196	196			1228	287	287	745	196	196		
Mean	71	71.2	71	71.2	71.2	59.4	114.1	114.1	Mean	0	0	0.6	0.6	0.1	0.1	0.1		
Median	46.4	38.1	46.4	38.1	38.1	45.2	71.4	71.4	Median	-0.2	-0.2	0.3	0.3	-0.1	-0.1	-0.1		
Std.	103.4	93.2	103.4	93.2	93.2	78.9	167.4	167.4	Std.	1.6	1.6	2.3	2.3	1.1	1.1	1.6		
Min	-135.4	-31.8	-135.4	-31.8	-31.8	-135.4	-9.6	-9.6	Min	-4	-4	-4	-4	-2.3	-2.3	-3.5		
Max	1286.6	891.1	1286.6	891.1	891.1	1199.1	1286.6	1286.6	Max	5.1	5.1	5.1	5.1	4.8	4.8	3.1		
Tier 1 capital buffer, in % of regulatory minimum			285.6	229.8	229.8	300	311.2	311.2	Liquidity ratio, in % of total assets			20.9	27.6	20.1	14.3	14.3		
Mean	244.1	155.8	244.1	155.8	155.8	276.2	202.6	202.6	Mean	16.2	16.2	23.4	16.1	16.1	7.9	7.9		
Median	214	223.7	214	223.7	223.7	143.4	359.2	359.2	Std.	16.4	16.4	19.8	13.6	13.6	17.1	17.1		
Std.	-63.2	30.3	-63.2	30.3	30.3	-63.2	11.6	11.6	Min	0	0	0.6	1.1	1.1	0.1	0.1		
Min	2555.8	2031	2555.8	2031	2031	997.6	2555.8	2555.8	Max	81.1	81.1	78.7	81.1	81.1	71.4	71.4		
Max			511.2	467.1	467.1	540.4	464.8	464.8	Loan growth, in %			9.1	7.9	8.5	13.1	13.1		
Core Tier 1 capital buffer, in % of regulatory minimum			464	325.2	325.2	526.8	370.2	370.2	Mean	7.1	7.1	5.3	7.4	7.4	5.5	5.5		
Mean	316.9	445.7	316.9	445.7	445.7	243.8	326.7	326.7	Std.	43.8	43.8	34.6	13.4	13.4	96.8	96.8		
Median	-45.9	5.6	-45.9	5.6	5.6	-45.9	-39.7	-39.7	Min	-32.4	-32.4	-54	-10.2	-10.2	-41.2	-41.2		
Std.	4161.6	4161.6	4161.6	4161.6	4161.6	1942.3	1909.4	1909.4	Max	80.5	80.5	95.7	42.3	42.3	103.7	103.7		
Min			9.2	5.6	5.6	9.3	13.7	13.7	Real GDP growth, in %									
Max	8.4	7.2	8.4	7.2	7.2	8.8	8.2	8.2	Mean	1.6	1.6	1.8	1.6	1.6	1.6	1.6		
RoE, in %	3	13.3	3	13.3	13.3	12.4	67.8	67.8	Median	1.4	1.4	1.4	1.4	1.4	1.4	1.4		
Mean	-98.5	-98.5	-98.5	-98.5	-98.5	-68.1	-68.1	-68.1	Std.	-2.7	-2.7	-2.7	-2.7	-2.7	-2.7	-2.7		
Median	995.3	411.1	995.3	411.1	411.1	217.7	995.3	995.3	Min	3.4	3.4	3.4	3.4	3.4	3.4	3.4		
Std.									Max									
Risk-weighted assets, in % of total assets									Inflation, annual percentage change									
Mean	61.6	59.1	61.6	59.1	59.1	60.3	68.7	68.7	Mean	1.7	1.7	1.7	1.7	1.7	1.7	1.7		
Median	62.1	54.3	62.1	54.3	54.3	63.1	67.2	67.2	Median	0.6	0.6	0.6	0.6	0.6	0.6	0.6		
Std.	25.8	29.6	25.8	29.6	29.6	21.6	30.2	30.2	Std.	0.6	0.6	0.6	0.6	0.6	0.6	0.6		
Min	4.6	14.7	4.6	14.7	14.7	15.2	4.6	4.6	Min	2.8	2.8	2.8	2.8	2.8	2.8	2.8		
Max	178	178	178	178	178	157.5	130.4	130.4	Max									



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

Firstly, the high levels of capital buffers are evidenced by the median reaching 46% and the average 71% of the regulatory minimum (set at 8% of risk-weighted assets). This high figure suggests that most of the time the regulatory constraint is not binding. The buffers are especially high for the financial and investment firms (114.1%). Tier 1 and core Tier 1 capital stand largely higher than the regulatory minima (244% and 464% respectively for the whole sample median, the minima being defined at 4% and 2%, respectively). Mutual savings banks and credit cooperatives exhibit a very high level of capitalization for these highest quality forms of capital (median figures reaching 276% for Tier 1 and 527% for core Tier 1), in line with the stylized facts reported in the literature (Stolz and Wedow, 2011; Jokipii and Milne, 2008). This finding may stem from the fact that this kind of banks has a lower access than commercial banks to debt instruments included in the overall capital, such as hybrid securities.

Secondly, the high degree of cross-sectional heterogeneity of buffer levels mainly explains the high standard deviations (103.4% overall). In this respect, with similar levels of within and between standard deviation of the buffer variable, at respectively 73.6 and 76.6%, it is not straightforward to infer how much of the econometric results are driven by cross sectional vs. time series changes in capital buffers. Depending on the drivers of the results, one can better assess the relative roles of differences in banks' risk aversion (if the cross-section dimension is the main driver) or economic cycle (if the time series dimension is the main driver), even though both concepts are intertwined as the risk aversion of a specific bank can change across the business cycle.

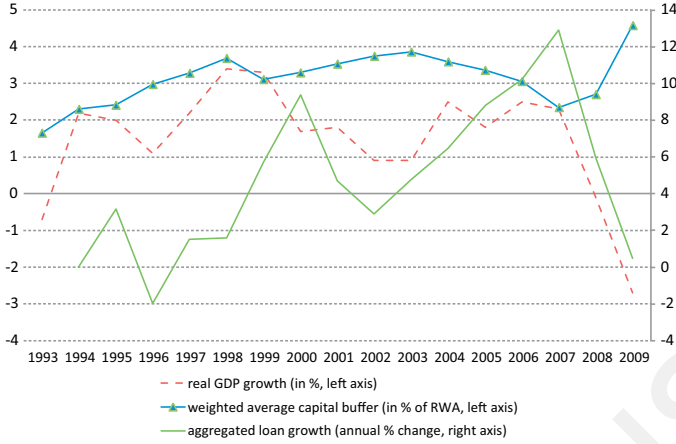
Graph 1 shows the fluctuations of the buffers. It suggests that banks do not target a fixed buffer as the evolution of their capital level is not correlated with the evolution of their risk-weighted assets (RWA). Another interpretation is that changes in the level of capital buffers are driven both by changes in the levels of total capital and of RWAs, i.e. both the numerator and the denominator of the regulatory solvency ratio.

Graph 2 illustrates the evolution of GDP growth, credit growth and the average capital buffer for the French 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 across the periods. At first glance, there seems to be a break in 1999, when the relationship between the average weighted capital buffer on the one hand, GDP growth and loan growth on the other hand, turns to be negative. Hence, our econometric investigation allows for an alternative specification estimated on the period prior to 1999 (see *infra*).

Table 2 provides the correlation coefficients between the variables used in our model. We find a slightly negative correlation between capital buffers and real GDP growth (-0.02). As expected, loan growth and GDP growth are positively correlated (0.1). Interestingly, there is a negative correlation between the different capital buffers and loan growth (between -0.2 and -0.3). These negative correlation coefficients are consistent with the intuition that a declining GDP growth would be associated with larger capital buffers that would in turn slow down loan growth. The ratio of risk-weighted assets

Table 2
Correlation coefficients between the variables.

	Total capital buffer	Tier 1 capital buffer	Core Tier 1 capital buffer	Size	Risk	ROE	Liquidity ratio	Loan growth	GDP growth	Inflation	CB's interest rate
Total capital buffer	1										
Tier 1 capital buffer	0.72	1									
Core Tier 1 capital buffer	0.49	0.84	1								
Size	-0.24	-0.33	-0.41	1							
Risk	-0.06	-0.09	-0.09	-0.18	1						
ROE	0.10	0.13	0.04	-0.06	-0.03	1					
Liquidity ratio	0.11	0.11	0.20	0.06	-0.07	-0.08	1				
Loan growth	-0.26	-0.24	-0.23	0.12	0.01	0.08	-0.09	1			
GDP growth	-0.02	-0.09	-0.08	-0.01	0.05	0.00	0.02	0.10	1		
Inflation	0.02	0.08	0.09	0.01	-0.05	0.03	-0.02	-0.02	-0.86	1	
CB's interest rate	-0.05	-0.17	-0.16	0.02	0.05	-0.07	0.10	-0.07	0.31	0.01	1



Graph 2. Cyclical developments in capital buffers and loan growth.

to total assets is positively correlated to GDP growth (0.05), which suggests that banks turn toward riskier assets in upturns.

3. Model and econometric results

3.1. Total capital buffers, loan growth and the business cycle

We aim at estimating the interactions between capital buffers and loan growth as well as the effect of business cycle on each of the two variables. The hypothesis we want to test is whether capital buffers amplify the real cyclical effects on loan growth. As both capital buffers and loan growth are endogenous, we estimate a *simultaneous* system of two equations in which both variables are *endogenous* to each other.³ The GDP growth rate is introduced as an explanatory variable and as a proxy for the business cycle, alongside with a set of control variables, some of them being bank specific, others, representative of the French economy as a whole. After eliminating non-significant variables, we retain the following simultaneous set of equations:

$$\begin{cases} b_{it} = \alpha_1 b_{it-1} + \alpha_2 c_{it-1} + \alpha_3 g_{t-1} + \alpha_4 r_{it-1} + \alpha_5 s_{it-1} + \alpha_6 \pi_{it-1} + \alpha_{i0} + \varepsilon_{it} \\ c_{it} = \beta_1 c_{it-1} + \alpha_2 b_{it-1} + \beta_3 g_{t-1} + \beta_4 r_{it-1} + \beta_5 s_{it-1} + \beta_6 l_{it-1} + \beta_7 \dot{p}_{t-1} + \beta_8 i_{t-1} + \beta_{i0} + v_{it} \end{cases} \quad (1)$$

where all the variables indexed by i and t are specific to bank i at time t . b_{it} stands for the capital buffer, $b_{it} = (K_{it} - K_{it}^R)/K_{it}^R$ expressed as the relative gap between the observed bank capital K_{it} and the regulatory minimum capital requirement K_{it}^R ; c_{it} is bank i 's loan growth rate (calculated as the first difference in logarithms); r_{it} is the ratio of risk-weighted assets (RWA) to total assets; s_{it} measures the bank relative size, as the bank's total assets minus the average total assets (both in logarithms); π_{it} is the return on equity (ROE); l_{it} is the ratio of liquid assets (cash, interbank loans and securities held in the trading portfolio and available for sales) to total assets. The variables not indexed by i are related to the whole French economy: g_t is the real GDP growth rate; \dot{p}_t is the inflation rate and i_t , the central bank's nominal policy rate. α_k and β_k , $k \geq 1$, are parameters to estimate which are common to all banks; α_{i0} and β_{i0} are fixed effects, ε_{it} and v_{it} are the residuals of each equation.

We estimate the system on panel data by using a three-stage least square estimator. To correct for heteroskedasticity, we use analytical weights, which are inversely proportional to the variances.

³ As a robustness check, we re-estimate the model using a single panel-GMM equation with credit growth on the left-hand side, capital defined as endogenous and other control variables on the right-hand side. The regression is carried out in line with the Arellano–Bover (1995) methodology. The results, available upon request, are consistent with those presented in the paper.

First, concerning the buffer equation, we find a significantly negative effect of the GDP growth on the bank capital buffers (Table 3, column 1). Hence, an economic downturn brings about an increase in bank capital buffers, as banks hoard more precautionary reserves in bad times. This result can be interpreted as capital buffers being prone to generate pro-cyclical effects.

As regards the control variables, the large coefficient on the lagged buffer (0.74) shows the high degree of persistence in these capital buffers. The lagged loan growth has a significantly negative sign, as it is positively correlated to the required capital, which is the denominator for the capital buffer. The significant and negative sign on the risk-weighted assets ratio suggests that risk is an important driver of capital buffers either through volume or compositional effects. The coefficient on the bank's size is significantly negative: large banks tend to hold less capital than others, comforting the *too-big-to-fail* hypothesis. The significant and positive coefficient on the return on equity suggests that a higher profitability may help build up capital buffers through retained earnings.

Second, turning to the loan growth equation, we find a significantly positive coefficient on GDP growth, as expected, whereas the liquidity ratio is not significant.⁴ Concerning the capital buffer, its coefficient is significantly negative (Table 4, column 1). This means that banks reduce their loan supply in the aftermath of an increase in their capital buffer, after controlling for GDP growth as a proxy for loan demand⁵ and other effects. The low size of the coefficient (-0.01) implies a weak economic effect. As an illustration, an average bank eliminating completely its capital buffer (71%) would only increase its loans by 0.7%. However, by its sign, this result contrasts with those commonly reported in the literature on US banks (Berrospide and Edge, 2010).

On the whole, we find a pro-cyclical effect of capital buffers arising from the estimation of both equations: capital buffers are increased by economic downturns (buffer equation), and their increase tends to dampen loan growth (loan equation). From this perspective, capital buffers amplify the effect of the business cycle.

3.2. Testing different forms of capital

To check for robustness, we carry out alternative estimations by retaining different definitions of bank capital. Doing that, we are also able to test whether the relationship between bank's capital buffers, the business cycle, and loan growth is stronger for the purest forms of capital, i.e. Tier 1 and core Tier 1 capital. To this end, we define Tier 1 capital buffers as the relative gap between the observed Tier 1 capital and its regulatory minimum level: $b_{it}^{Tier1} = (K_{Tier1,it} - K_{Tier1,it}^R) / K_{Tier1,it}^R$, and use the corresponding definition for core Tier 1 capital buffer. The regulatory minimum ratio is set at 4% of RWAs for the Tier 1 capital ratio and 2% for core Tier 1. We then successively substitute these high-quality capital buffers for the total capital buffer in both Eq. (1) and re-run the estimation.

The results strengthen the case for pro-cyclicity already found in the previous estimations. More specifically, the coefficient of GDP growth on capital buffers is negative and higher in absolute value with better quality capital (Table 3, columns 2 and 3). Consequently, the highest-quality capital is more sensitive to the business cycle than the total capital.⁶ The autoregressive coefficients of Tier 1 and core Tier 1 buffers are also greater than in the previous estimation, suggesting that banks face higher adjustment costs for the purest forms of capital. The coefficient on bank size for the core Tier 1 buffer is negative and higher (in absolute value) than for the total capital buffer estimation, although it is not significant for Tier 1 capital. This implies that constraints would be looser for large banks, considered as *too big to fail*, especially concerning their core Tier 1 capital. Turning to the loan equation, high-quality capital buffers have a larger and more significantly negative impact than total buffers. Unsurprisingly, the impact of GDP on loan growth remains the same, whatever measure of capital buffer is introduced in the equation.

⁴ We do not have at our disposal a better control variable for loan demand than GDP growth. Jimenez et al. (2010) used credit registers for Spanish banks but we do not have access to such data for French banks.

⁵ This reflects the fact that the variability of capital buffers increases with their quality, core Tier 1 capital buffer being more volatile than the other forms of capital (see Table 1).

⁶ Two banks have been removed from the sample because of an insufficient number of data points.

Table 3
Estimated coefficients for the capital buffer Eq. (1).

Explanatory variables	Exp. sign	Total buffer		Tier 1 buffer		Core Tier 1 buffer		Core Tier 1 buffer		Total buffer		Core Tier 1 buffer		Core Tier 1 buffer		
		All	(1)	All	(2)	All	(3)	All	(4)	Credit cooperatives	(5)	Less capitalized	(6)	1993–2006	(7)	All
Lagged buffer	+	0.67*** (29.48)	0.88*** (52.52)	0.81*** (41.02)	0.82*** (43.38)	0.81*** (43.41)	0.82*** (37.83)	0.23*** (3.90)	0.82*** (37.83)	0.81*** (43.41)	0.23*** (3.90)	0.82*** (37.83)	0.82*** (37.83)	0.82*** (37.83)	1.12*** (43.74)	1.12*** (43.74)
Real GDP growth	?	-0.07*** (-2.98)	-0.19*** (-6.19)	-0.25*** (-4.83)	-0.25*** (-4.86)	-0.16*** (-3.73)	-0.25*** (-4.86)	-0.07*** (-1.52)	-0.16*** (-3.73)	-0.16*** (-3.73)	-0.07*** (-1.52)	-0.25*** (-4.86)	-0.25*** (-4.86)	0.45*** (6.27)	0.45*** (6.27)	
Credit growth in log	-	-0.52*** (-4.10)	-0.75*** (-4.35)	-1.60*** (-5.62)	-1.60*** (-5.62)	-0.92*** (-3.34)	-1.60*** (-5.62)	-0.15*** (-0.72)	-0.92*** (-3.34)	-0.92*** (-3.34)	-0.15*** (-0.72)	-1.52*** (-5.03)	-1.52*** (-5.03)	-1.29*** (-4.72)	-1.29*** (-4.72)	
Return on equity	-	0.26*** (3.50)	0.42*** (4.38)	0.12 (0.75)	0.12 (0.76)	0.54 (1.63)	0.12 (0.76)	0.90 (1.54)	0.54 (1.63)	0.90 (1.54)	0.14 (0.86)	0.14 (0.86)	0.14 (0.86)	0.10 (0.621)	0.10 (0.621)	
Risk	?	-0.26*** (-2.57)	-0.01 (-0.06)	-0.09 (-0.40)	-0.01 (-0.44)	0.01 (0.03)	-0.01 (-0.44)	-0.24 (-1.44)	0.01 (0.03)	-0.24 (-1.44)	-0.11 (-0.45)	-0.11 (-0.45)	0.22 (1.23)	0.22 (1.23)		
Size	-	-0.06*** (-3.98)	-0.04* (-1.70)	-0.13*** (-3.40)	-0.12*** (-3.59)	-0.10*** (-2.16)	-0.12*** (-3.59)	0.02 (0.56)	-0.10*** (-2.16)	0.02 (0.56)	0.02 (0.56)	-0.13*** (-3.00)	-0.13*** (-3.00)	-0.08*** (-2.36)	-0.08*** (-2.36)	
Real GDP growth × core Tier 1 buffer	?	-	-	-	-	-	-	-	-	-	-	-	-	-0.14*** (-9.75)	-0.14*** (-9.75)	
Constant		0.58*** (6.86)	0.88*** (6.97)	1.78*** (7.78)	1.72*** (11.17)	1.47*** (6.82)	1.72*** (11.17)	1.16*** (5.07)	1.47*** (6.82)	1.16*** (5.07)	1.16*** (5.07)	1.76*** (6.77)	1.76*** (6.77)	1228	1228	
Number of observations		1228	1228	1228	1228	745	1228	116	745	116	116	1072	1072	7	7	
Number of estimated coefficients		7	7	7	6	7	6	7	7	7	7	7	7	7	7	
R-square		0.49	0.76	0.68	0.68	0.78	0.68	0.23	0.78	0.23	0.23	0.68	0.68	0.91	0.91	

Note: t-statistics in brackets.

* Significant at the threshold of 10%.

** Significant at the threshold of 5%.

*** Significant at the threshold of 1%.

Table 4
Estimated coefficients for the loan growth Eq. (1).

Explanatory variables	Exp. sign	Loan growth All (1)	Loan growth All (2)	Loan growth All (3)	Loan growth All (4)	Loan growth Credit cooperatives (5)	Loan growth Less capitalized (6)	Loan growth GIMM 1993–2006 (7)	Loan growth All (8)
Lagged credit growth in log	+	0.08*** (4.34)	0.09*** (4.48)	0.08*** (4.38)	0.08*** (4.40)	0.05*** (2.84)	-0.05 (-0.73)	0.10*** (5.05)	0.08*** (4.21)
Total buffer	?	-0.01*** (-2.64)	-0.005*** (-2.81)	-0.004*** (-2.89)	-0.004*** (-3.08)	-0.003*** (-2.56)	0.03 (1.49)	-0.004*** (-3.17)	-0.01*** (-3.29)
Tier 1 buffer	?								
Core Tier 1 buffer	?								
Real GDP growth	+	0.03*** (8.40)	0.03*** (8.85)	0.03*** (8.41)	0.03*** (11.25)	0.03*** (7.44)	0.01 (0.56)	0.03*** (7.71)	0.02*** (4.77)
Inflation	+	0.04*** (6.72)	0.04*** (7.57)	0.04*** (6.87)	0.04*** (8.76)	0.03*** (5.06)	0 (0.02)	0.04*** (6.63)	0.05*** (7.32)
Central bank's policy rate	-	-1.74*** (-5.88)	-2.01*** (-7.03)	-1.89*** (-6.56)	-1.82*** (-6.64)	-1.49*** (-5.95)	-0.82 (-0.72)	-2.05*** (-6.94)	-1.73*** (-6.00)
Risk	+	0.01 (0.46)	0.01 (0.41)	0.01 (0.75)	- (0.02)	0.04*** (2.42)	0.02 (0.33)	0.01 (0.85)	0.01 (0.63)
Liquidity ratio	+	0.003 (0.14)	0 (0.10)	0.02 (1.00)	0.02 (0.79)	0.06*** (2.44)	0.26** (2.07)	0.02 (0.84)	0.03 (1.12)
Size × 100	+	5.05*** (2.21)	4.09*** (1.71)	3.67 (1.49)	0.30 (1.30)	0.003 (0.01)	-0.93 (-0.82)	0.48* (1.78)	0.32 (1.29)
Real GDP growth × core Tier 1 buffer	-								
Number of observations		1228	1228	1228	745	116	1072	1228	(1.83)
Number of estimated coefficients		8	8	7	8	8	8	9	
R-square		0.28	0.28	0.28	0.55	0.22	0.29	0.29	

Note: *t*-statistics in brackets. Constant omitted as found to be insignificant.

* Significant at the threshold of 10%.

** Significant at the threshold of 5%.

*** Significant at the threshold of 1%.

In an alternative specification, we remove the “risk” variable, i.e. the risk-weighted asset to total asset ratio. Indeed, this variable may be correlated with capital buffers, since higher “risk” results in higher minimum capital requirements and hence lower capital buffers, everything else being equal. The results are not sensitive to the exclusion of the risk variable as they remain unchanged in both equations (column 4 of Tables 3 and 4).

On the whole, both the effect of GDP growth on capital buffers and that of capital buffers on loan growth prove more significant for better quality capital. This suggests a stronger pro-cyclical effect of Tier 1 and core Tier 1 capital.

3.3. *Different categories of banks and periods*

We now split the sample into different categories of banks to check whether the legal form has a bearing on how banks build up their capital buffers. More precisely, we focus on the sub-sample restricted to credit cooperatives and savings banks. As these institutions are typically smaller than commercial banks, they are likely to have a lower access to funding markets. We can therefore expect a high persistence of capital buffer and a high sensitivity to the business cycle for precautionary reasons. The estimations show that these financial institutions closely mirror the whole panel’s behavior for the buffer (Table 3, column 5) as well as the loan equation (Table 4, column 5). However, in the loan equation, the coefficients on the risk variable and the liquidity ratio turn to be significant with the expected sign. We interpret this result as reflecting the lower diversification of mutual banks and credit cooperatives in terms of funding sources and thus their greater reliance on liquid assets.

To check whether low-capitalized banks have a different behavior, we re-run the estimation over a sample restricted to the least capitalized banks, i.e. those having a core Tier 1 capital buffer smaller than the 10th percentile of the distribution. Results show that the coefficient on their core Tier 1 capital buffers is not significant and positive, contrarily to the results obtained over the whole sample (Table 4, column 6). This evidences non-linearities in the effect of capital buffers on credit growth. When banks are very low-capitalized, an increase in their capital buffer will not reduce their credit growth as for the other banks, as their situation is relatively improved by higher capital. Moreover, we also find a positive and significant coefficient on the liquidity ratio for those banks, in contrast to the estimation on the whole sample. This highlights the fact that less capitalized banks are more dependent on the availability of liquid assets for their loan supply.

To further check the results, we also exclude the 2007–2009 period in order to verify that the former estimations are not driven by the effects of the financial crisis and/or the simultaneous implementation of Basel II. Results remain broadly unchanged as regards our variables of interest (Table 4, column 6); likewise, if we restrict our sample to the 1993–1998 period, i.e. before the euro cash changeover.

Finally, we test for possible non-linear effects in the loan equation in order to determine if the effect of business cycle on loans could depend on the level of capital buffers. To do this, we introduce an interactive variable equal to capital buffer multiplied by GDP growth (Table 4, column 7). We do find a significant and positive coefficient on this variable when core Tier 1 capital buffer is considered. This result suggests that a higher level of capital makes lending growth more sensitive to GDP growth. Hence, higher capital buffers do not enable banks to insulate their lending supply from the effect of the business cycle. We interpret this as a sign of high risk aversion on the part of banks having large capital buffers. Overall, our results highlight the decisive role of banks’ behavior in terms of risk aversion and how risk aversion changes over a business cycle, as compared to the role of capital regulation.

On the whole, the econometric investigation carried out in Section 3 shows that capital buffers amplify the cyclical behavior of loans caused by the business cycle. This result is all the more relevant as we focus on the Tier 1 and core Tier 1 capital, i.e. the purest components of bank capital. In the current context of discussions around the future Basel III framework and especially the future implementation of a countercyclical capital buffer, these results *qualitatively* support the view that an efficient macroprudential regulation should aim at smoothing credit growth. From that perspective, bank capital is a relevant instrument. Indeed our results suggest that regulating the purest components of capital is the most efficient way to dampen the pro-cyclicality of banks’ capital buffers.

4. Granger causality tests

We now check for the causality between capital buffers and loan growth. If the pro-cyclicality hypothesis of capital buffers holds, the causality would run from the capital buffers to the loan growth.

4.1. 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 . Therefore these tests should be understood as being about statistical instead of economic causality. Under the usual assumptions of stationarity of the series, the standard version model is the following:

$$y_t = \sum_{k=1}^K \alpha_k y_{t-k} + \sum_{k=1}^K \beta_k x_{t-k} + \mu + u_t \quad (2)$$

where α_k , β_k and μ are parameters to estimate, K is the optimal number of lags in the regression, u_t is the residual of the equation. The causality test comes down to estimate Eq. (2), 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 . The null hypothesis H_0 is that of no causality: $H_0: \beta = 0$, where $\beta = (\beta_1, \dots, \beta_K)$ is the vector of the lagged coefficients β_k .

Here, as we deal with panel data with a small time dimension ($T=17$ at maximum), standard individual tests of H_0 are not powerful. Hence, we use a panel-causality test proposed by Hurlin (2005, 2008). We test for causality for each bank individually, running N individual regressions as Eq. (2). Let us re-write Eq. (2) to take into account the cross-section dimension of the panel:

$$y_{it} = \sum_{k=1}^K \alpha_{k,i} y_{i,t-k} + \sum_{k=1}^K \beta_{k,i} x_{i,t-k} + \mu_i + u_{it} \quad (3)$$

where α_{ki} , β_{ki} and μ_i are parameters to estimate, allowed to be different across individuals, K is a given number of lags common to all individuals, and u_{it} is the residuals of the equation. Following Hurlin (2005), we test for the homogeneous non-causality. The null hypothesis H_0 is that there does not exist any individual causality: $\beta_i = 0, \forall i = 1, \dots, N$, where $\beta_i = (\beta_{i1}, \dots, \beta_{iK})$ is the vector of the coefficients β_{ik} . Therefore, the alternative hypothesis (H_1) is that there exists at least one individual for which there is causality.

First, we calculate the individual Wald tests W_i for $\beta_i = 0$ in the N estimations of Eq. (3). Then, we compute the mean Wald test for the panel: $\bar{W}_N = (1/N) \sum_{i=1}^N W_i$. This statistic \bar{W}_N converges toward a $\chi^2(K)$, when the time dimension of the panel T tends toward 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 :

$$\tilde{Z}_N = \sqrt{N} \left[\bar{W}_N - \frac{K}{N} \sum_{i=1}^N \frac{T_i - 2K - 1}{T_i - 2K - 3} \right] \left[2 \frac{K}{N} \sum_{i=1}^N \frac{(T_i - 2K - 1)^2 (T_i - K - 3)}{(T_i - 2K - 3)^2 (T_i - 2K - 5)} \right]^{-1/2} \quad (4)$$

\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).

4.2. Results at the bank level

We test for Granger-causality between banks' capital buffers b_{it} , and the loan growth c_{it} by implementing the test described above. The sample is the same as in the previous section, composed of 96 banks⁷ on period 1994–2009. First, we check that the series are stationary by using the standard

⁷ Two banks have been removed from the sample because of an insufficient number of data points.

Table 5
Granger causality tests on total capital buffer and loan growth – Panel level.

Null hypothesis	Hurlin Z-stat. (1)	Null hypothesis	Hurlin Z-stat. (2)
Whole panel			
Capital buffer does not Granger cause loan growth	1.88* (0.06)	Core Tier 1 capital buffer does not Granger cause loan growth	1.83* (0.07)
Loan growth does not Granger cause Capital buffer	2.85*** (0.00)	Loan growth does not Granger cause core Tier 1 capital buffer	1.81* (0.07)
Mutual, savings banks and credit cooperatives			
Capital buffer does not Granger cause loan growth	1.69* (0.09)	Core Tier 1 capital buffer does not Granger cause loan growth	2.42** (0.02)
Loan growth does not Granger cause Capital buffer	1.45 (0.15)	Loan growth does not Granger cause core Tier 1 capital buffer	0.62 (0.53)
Commercial banks			
Capital buffer does not Granger cause loan growth	1.00 (0.32)	Core Tier 1 capital buffer does not Granger cause loan growth	0.48 (0.63)
Loan growth does not Granger cause Capital buffer	2.19** (0.03)	Loan growth does not Granger cause core Tier 1 capital buffer	2.56*** (0.01)
Financial companies			
Capital buffer does not Granger cause loan growth	0.26 (0.79)	Core Tier 1 capital buffer does not Granger cause loan growth	-0.53 (1.41)
Loan growth does not Granger cause Capital buffer	1.55 (0.12)	Loan growth does not Granger cause core Tier 1 capital buffer	0.16 (0.87)

Note: figures in brackets are *p*-values. Akaike and Schwarz criteria indicate an optimal lag equal to 1 (annual data).

*** We reject the null hypothesis at the confidence threshold of 1%.

** We reject the null hypothesis at the confidence threshold of 5%.

* We reject the null hypothesis at the confidence threshold of 10%.

panel unit root tests: Levin et al. (2002); Im et al. (2003) as well as augmented Dickey–Fuller and Phillips–Perron tests (Table A1). The results of every test show that both series (buffer and loan growth) are stationary.

Second, we estimate the following system of two equations:

$$\begin{cases} b_{it} = \sum_{k=1}^K \alpha_{1ik} b_{it-k} + \sum_{k=1}^K \beta_{1ik} c_{it-k} + \mu_{1i} + u_{1it} \\ c_{it} = \sum_{k=1}^K \alpha_{2ik} b_{it-k} + \sum_{k=1}^K \beta_{2ik} c_{it-k} + \mu_{2i} + u_{2it} \end{cases} \quad (5)$$

we get the same optimal number of lags $K = 1$, for all banks using either the Akaike or Schwarz criteria.

Results show that causality runs in both directions over the whole panel (Table 5, column 1). This finding suggests mutually reinforcing effects between the two variables but might result from a common factor, as the two variables are affected by economic growth. A look at the estimated coefficients confirms the negative relationship between the capital buffer and credit growth found in the previous estimations. An analysis by category of banks reveals that, for the largest category of credit institutions, namely mutual banks, the causality runs only from capital buffers to credit growth, confirming that their lending decisions strongly depend on the level of capital. The lower size of mutual banks compared to commercial banks and their more limited access to funding and debt markets may explain this stronger causality. We carry out the same tests after removing successively one of the 96 banks from the sample and verify that all results are robust to the exclusion of any bank.

We re-run the Granger causality tests after substituting core Tier 1 capital buffer for total capital buffer (Table 5, column 2). The estimations in Section 3 have shown that high quality capital buffers have a stronger impact on loan growth. The causality tests confirm this finding for mutual, saving banks and credit cooperatives. High quality capital levels may be a more important driver of banks' lending decisions than total capital.

5. Conclusion

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 capital buffers, credit growth as well as their relationship with output. The French banking system is characterized by a large number of mutual banks whose lending growth is shown to be more dependent on the size of capital buffers. Our empirical approach is based on both panel data simultaneous equations estimations and Granger causality tests within a unified, integrated, framework. We find evidence of a two-way causality between capital buffers and loan growth, pointing to mutually reinforcing mechanisms. We also find that French banks' capital buffers, as well as loan growth at the individual level, depend on GDP growth in a pro-cyclical manner.

The focus on a single country, where bank lending is by far the prevailing form of external finance, provides an analytical framework in which to examine timely research questions related to the pro-cyclicality 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 massive deleveraging observed since the deepening of the current financial crisis. It is also in line with recent theoretical models such as Valencia (2010). The fact that higher capital buffers may reduce, rather than sustain, the credit supply following a severe recessionary shock should be seriously considered when designing macro-prudential policy.

Appendix A.

See Appendix Table A1.

Table A1
Panel unit root tests.^{a,b}

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

^a Lags are selected by Akaike criterion.

^b The tests include an individual intercept.

^c The null hypothesis is rejected when $p\text{-value} < 0.05$.

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