Are capital buffers pro-cyclical?
Evidence from Spanish panel data

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1. Introduction

Efforts within the Basel Committee on Banking Supervision to update the 1988 Basel Accord have almost finalised a new accord on banks' capital adequacy, which is already known as "Basel II". There remain, however, some outstanding issues and details. Among these, that of "pro-cyclicality" has, perhaps, produced most debate in the literature[1].

The concept of pro-cyclicality, when applied to the new capital requirements, may in principle be a little confusing. As is well known, one of the primary aims of the new accord is to link capital requirements more closely to risks. Accordingly, in a downturn, for instance, when risks are more likely to materialise, capital requirements might increase. Thus, capital requirements and output growth will move in opposite directions. But if capital requirements increase, banks would have to reduce their loans and the subsequent credit squeeze would add to the downturn. Capital requirements are therefore said to be (likely to be) pro-cyclical because they might amplify the fluctuations of the business cycle.

It has been also argued, however, that if internal risk management models take properly into account the way default probabilities change throughout the business cycle, the effects on credit –and therefore, on output– should not be overstated.

While arguments highlighting or minimising the actual relevance of the pro-cyclicality problem have proliferated, the related empirical evidence is scant. Admittedly, the empirical literature on the impact of capital requirements on bank behaviour is extensive, though mainly confined to the US case[2]. Papers have dealt with issues such as whether the introduction of minimum capital requirements leads banks to hold higher capital; the impact of capital requirements on risk-taking, competitiveness and a level playing field; and whether capital requirements create credit crunches affecting the real economy. Nevertheless, as far as we know, these papers have not analysed the cyclical behaviour of capital requirements, perhaps because the current Basel capital accord ties capital requirements less closely to banks' capital risk.

Against this background, the aim of this paper is to provide some fresh empirical evidence which may prove useful in the debate about the pro-cyclicality of the new capital accord. In particular, we have noticed that most arguments in this debate relate to the cyclicality of the capital *requirements*, thus ignoring the fact that only a few banks hold just the required capital, while most keep capital buffers which, in some

[1] See, among others, BCBS (2001), Borio et al (2001), Danielsson et al (2001), DNB (2001) and ECB (2001).
[2] See, for example, Berger and Udell (1994), Berger (1995), Furfine (2000) and Kwan and Eisenbeis (1997). Jackson et al (1999) offers a detailed review.

cases, are quite significant[3]. The behaviour of these buffers and, in particular, their relationship to the business cycle may thus be an important piece of information. A positive relationship would mean that banks rebuild their capital during upturns, which might be used to fulfil a likely increase in requirements during the next downturn. This might offset, at least partially, the pro-cyclicality of capital requirements under Basel II. A negative relationship, however, would have less encouraging implications in this respect.

In this paper, we analyse the relationship between the Spanish business cycle and the capital buffers (current capital over the minimum capital requirements) held by Spanish commercial and savings banks. Using standard econometric panel data techniques, we build an incomplete panel of Spanish institutions from 1986 to 2000 –thus covering a complete business cycle– and estimate an equation for the behaviour of capital buffers that includes an indicator of the business cycle.

Admittedly, focusing on a single country might reduce the scope of our analysis. While data availability (in particular regarding individual capital requirements) prevents us from a more general approach, it is worth mentioning that the business cycle seems –at least during the period considered– to have been relatively pronounced in Spain, which might render this an interesting case study. Moreover, Spanish banks are highly competitive and efficient, which reduces the probability of idiosyncratic factors biasing the results (see, for instance, ECB, 1999).

Our findings are fairly robust and quite unequivocal. After controlling for other potential determinants of the surplus capital (cost of capital, risk profile of the institution, size…) we find a (robustly significant) negative relationship between the business cycle and the capital buffers that Spanish institutions held throughout the period analysed. This relationship is, moreover, asymmetric, being closer during upturns than during downturns. From a quantitative standpoint, an increase of 1 percentage point in GDP growth might reduce capital buffers by 17%.

According to our results, there is a strong case for seriously taking into account the so-called pro-cyclicality problem in the final design of Basel II. As one of the main objectives of the new accord is to link more directly capital requirements to actual risks, which are cyclical, Pillar 2 seems to be the right place to address the issue.

It should be noted, finally, that our results relate to the behaviour of banks under Basel I. Of course, it cannot be totally ruled out that Basel II may cause a structural change. It is worth noting, however, that Basel II is going to change a quantitative requirement which, for most banks, is not strictly binding. Moreover, on average, capital requirements are not going to be increased. It is therefore an open question whether such a change has necessarily to affect the behaviour of capital buffers. In any event, it

[3] Papers have usually focused on the determinants of the capital ratio or its rate of growth. See the references in Jackson et al (1999).

goes without saying that the policy implications of our results should be viewed with due caution.

The rest of the paper is structured as follows. The theoretical framework as well as the empirical equation we estimate are introduced in section 2, while the data set is described in section 3. The fourth section shows the main results of the basic econometric analysis, and some extensions are considered in section 5, providing more information on the pro-cyclicality of capital buffers. Finally, the conclusions of the paper are drawn in the last section.

2. The theoretical framework and the empirical equation

As explained in the introduction, our contribution to the debate on the pro-cyclicality issue is of an empirical nature. In particular, our main objective is to estimate how capital buffers react to changes in the cyclical position. To consistently estimate such a response, however, we have to take into account the effects of other potential determinants of capital excess, which might have a cyclical behaviour. Yet as commented above, there is as far as we know no model for capital buffer behaviour in the literature. To circumvent this problem, we build a stylised model aimed at providing the basis for an empirical equation for capital buffers.

We start with a simple equation, which is fairly standard in the literature on real investment (see, for instance Blanchard and Fisher, 1989), describing the dynamics for the capital stock of a representative single bank

$$K_t = K_{t-1} + I_t \quad (1)$$

where K_t stands for the capital level at the end of period *t*, and I_t stands for the stock issues or repurchases (with a negative sign), plus retained profits during period *t*. While stock issues or repurchases can be decided in advance, retained profits are unknown at the beginning of each period and therefore the capital level is not known until the end of such a period.

The reasons why banks may wish to hold capital are manifold (see, for instance, Berger et al., 1995). But as in Froot and Stein (1998) or Estrella (2001), we model the decisions a bank makes on capital as the result of a trade-off among three different types of costs related to capital levels.

Firstly, holding capital has a direct cost for banks, as it has to be remunerated. In a context of asymmetric information, capital may even be more costly than alternative bank liabilities such as deposits or debt, as argued, for instance, in Campbell (1979) or Myers and Majluf (1984).

Secondly, holding capital reduces the probability of bankruptcy and therefore the so-called costs of failure, which include the loss of charter value, reputational loss and legal costs of the bankruptcy process (see, for instance, Acharya, 1996). Related to these costs are those associated with the existence of compulsory capital requirements. Higher capital levels also reduce the probability of not complying with those requirements, thus minimising the consequent costs. It is worth noting that including capital requirements in this way (i.e. as an additional cost of -not having enough- capital instead of in the form of a constraint) is not only useful for modelling purposes but also consistent with the way they actually operate. As a matter of fact, before regulatory limits are reached, supervisory authorities usually place some restrictions on the activity of the bank.

Finally, as in the literature on real investment, we also assume that changing the capital level entails adjustment costs. The reasons why these costs emerge in this case are, however, different. Apart from pure transaction costs, the main adjustment costs included those related to the problem of asymmetric information in capital markets. As the issuer has an informational advantage over the potential buyers, issuing (repurchasing) stocks may be seen by the latter as a signal that the company considers that market prices are above (below) the true share value, which would increase the cost of the desired adjustment (see, for instance, Myers and Majluf, 1984; Winter, 1994; or McNally, 1999).

Under the simplest possible approach, all these costs might be gathered in the following equation

$$C_t = (\alpha_t - \gamma_t)K_t + (1/2)\delta_t I_t^2 \tag{2}$$

where α_t represents the cost of remunerating the capital, γ_t represents the costs of failure (and/or the penalties for not complying with the regulatory minimum), and δ_t reflects the existence of adjustment costs. Both linearity (regarding the first two groups of costs) and symmetry (in connection with adjustment costs) are assumed for the sake of simplicity.

In this setting, the representative single bank minimises its intertemporal costs by solving the following problem

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

$$\text{s.t.} \quad C_t = (\alpha_t - \gamma_t)K_t + (1/2)\delta_t I_t^2 \tag{4}$$

$$K_t = K_{t-1} + I_t \tag{5}$$

where the FOC regarding C_t can be written as follows

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

and therefore

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

The regulatory minimum capital may now be subtracted from both sides and the expected capital replaced by the observed capital plus an expectation error to obtain

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

Accordingly, to consistently estimate the effects of the position in the cycle on the capital buffer we have to control for the effects of -i.e. to include in the RHS of our empirical equation- i) the first lag of the dependent variable, which captures the relevance of adjusting costs and should therefore have a positive sign; ii) variables related to the (expected) costs of remunerating capital, which should have a negative sign; and iii) variables capturing the (expected) bank failure costs for the bank, which are linked both to the bank's attitude towards risk (see, for instance, Keeley, 1990 or Salas and Saurina, 2002a) and to the bank's size, as suggested, for instance, by the well-known too-big-to-fail hypothesis.

Looking for the precise variables that best proxy the conceptual ones derived above is an empirical question firmly in line with the nature of this paper. The strategy we have followed is to start with a basic equation built on the basis of the most obvious proxies for α, γ and δ and then expand it in several ways to test the robustness of the results.

In particular, our starting point is the following equation for the capital buffer kept by institution *i* in period *t* (BUF_{it})

$$\begin{aligned} BUF_{it} &= \beta_0 BUF_{i,t-1} + \beta_1 ROE_{it} + \beta_2 NPL_{it} + \beta_3 BIG_{it} + \beta_4 SMA_{it} + \beta_5 GDPG_t \\ &+ \eta_i + \varepsilon_{it}, \quad i = 1,\ 2,\ldots,\ N\ (number\ of\ banks),\ \ t = 1,\ 2,\ldots,\ T \end{aligned} \tag{9}$$

where we first include the endogenous variable with a one-period lag –the expected sign being positive– to test whether adjusting costs are relevant. Direct costs of remunerating the excess capital are approximated by each institution's *ROE* (return on equity). The expected sign for this variable is thus negative. The risk profile of each institution is proxied by *NPL,* which measures the non-performing loans ratio (ratio of non-performing loans to total loans and credits). This is an *ex post* measurement of the risks assumed by the institution and, therefore, its expected sign is negative. In addition, there might be an idiosyncratic time-invariant component in the bank's risk profile, which would be captured, however, by η_i.

BIG and SMA are included to detect differences in the buffer according to the size of each institution. In particular, BIG (SMA) is a dummy variable taking value 1 for banks in the highest (lowest) decile. As commented above, big banks might be thought to keep relatively lower buffers as according to the "too-big-to-fail" hypothesis they believe that in the event of difficulties they will receive support from the regulator.[4]

After including the determinants of the capital buffer suggested by the model, GDP growth (*GDPG*) is added in order to determine whether the business cycle has an *additional* effect on the capital buffer held by institutions. The significance, sign and magnitude of β_5 will allow us to answer the main questions that are the core of this paper.

Finally, η_i is an unobservable variable aimed at capturing all idiosyncratic features of banks that are constant over time but vary from institution to institution –for example, the greater or lesser risk aversion of bank managers commented above– and ε_{it} is a standard random shock.

We then try different extensions of (9), which are explained in more detail in Sections 4 and 5. Beforehand, however, it is worth commenting briefly on our data set.

3. The data set

Our data are drawn from the financial statements regularly and obligatorily sent by institutions to the Banco de España. Consolidated figures have been used (except, of course, for institutions that do not consolidate their data and do not belong to a consolidable group), as capital requirements are imposed at the consolidated group level. The scope of the risks contained in consolidated balance sheets is, moreover, broader, as information about Spanish banking subsidiaries operating outside Spain is included.

[4] Small banks, on the contrary, might have to hold relatively larger buffers if they face more difficulties tapping capital markets. Under this view, SMA might be seen as a proxy for δ rather than for γ.

The way we measure the capital buffer merits some additional comments. In this respect, it is worth briefly reviewing the most significant regulatory changes to the Spanish capital adequacy ratio since 1985. That year saw the introduction in Spain of a capital adequacy ratio replacing the previous legally required ratio (equity/liabilities), whose usefulness was more than questionable. The 1985 capital adequacy ratio was calculated as the higher of a gearing ratio (equity/total assets) of 5% and a risk asset ratio under which a different level of capital was required (from 0.25% to 35%), depending on the risk associated with the different headings.

In 1993 the Spanish regulation was adapted to the Community directive which, broadly speaking, adhered to the 1988 Basel capital accord. The regulation has not changed fundamentally since, and it is based on the definition of the components of the numerator and denominator of the capital adequacy ratio, namely capital and risk-weighted assets. Capital should cover at least 8% of requirements. The new 1993 regulatory framework prompted something of a decline in capital requirements, with a subsequent increase in the capital buffer given that the new regulation was less demanding than the 1985 regulation[5].

In the light of the regulation in place before 1993, *BUF* has been defined as the institution's capital less the requirements to which it was subject, divided by the requirements, thus circumventing the drawback whereby, before 1993, there was no minimum ratio applied across the board to all institutions.

From among all credit institutions a sub-set has been selected made up of domestic institutions (including foreign institutions' subsidiaries operating in Spain) and savings banks. Foreign bank branches and credit co-operatives whose relative significance (in terms of business volume) is scant have been excluded.

The data are annual and span the period from 1986 to 2000. In this manner a full business cycle of the Spanish economy is included, a point of particular importance given that the aim of this paper is, as mentioned, to analyse whether there is a relationship between the business cycle and the capital buffer held by institutions.

Our panel is incomplete since new institutions have started to operate during the period considered while others have ceased to exist. Moreover, the impact of bank mergers during the period has also to be taken into account. Mergers pose an obstacle to calculating averages and, particularly, growth rates. To overcome this drawback so that the least number of observations possible is lost, it has been decided to artificially re-create the merger a period in advance. That is to say, if two banking institutions merge at t, for the purposes solely of calculating averages and growth rates, the

[5] Nonetheless, the Spanish 1993 regulation continues to be more demanding than that currently in force in Basel due mainly to the fact that recognition of unrealised capital gains is not permitted and general and statistical provisions are not considered as Tier 2 capital.

resulting institution is considered to have already existed at *t-1*, reconstructing it on the basis of the data from the individual institutions involved in the merger.

A similar problem arises for institutions which, having belonged at *t-1* to a consolidated group, leave such group at *t*. To calculate both the averages of certain variables and their growth rates, the figure at *t-1* is obtained from their individually reported financial statements.

Under these premises and after eliminating a series of institutions with extremely atypical data (due essentially to the specific nature of their business), an incomplete panel has been obtained comprising up to 142 institutions over a period of up to 15 years, totalling 1309 observations.

Table 1 shows the descriptive statistics of different variables, while Chart 1 plots the aggregate course of the aggregated capital buffer.

[Table 1]
[Chart 1]

4. Econometric results

First of all, it is worth noting that variables in the empirical equation (9) are defined in levels, while some (such as *NPL*) are likely to be correlated with η_i. As usual in panel data analysis, we proceed to transform (9) into first differences, to enable unbiased estimates to be obtained. Further, as the lagged endogenous variable is included among the regressors and other explanatory variables are likely to be endogenous, an estimation procedure based on the generalised method of moments (GMM) seems the most appropriate[6]. In particular, the instruments chosen for the lagged endogenous variable *NPL* and *ROE,* are two-to-four-period lags of the same variables. These lags have been chosen to avoid correlation with the error term ε_{it} (which now appears in first differences) while minimising, at the same time, the number of observations lost. The variables of size and business cycle are considered to be exogenous and therefore used as their own instruments.

Table 2, column 1, shows the main results of the estimation of equation (9). Regarding the significance and sign of the coefficient of output growth, we find that, after controlling for other determinants, there is a clearly significant (at 1%) negative relationship between the capital buffer and the phase of the cycle. Accordingly, in the case of Spain, capital requirements barely sensitive to the cycle (i.e. the 1988 Basle Accord) have translated into relatively pro-cyclical buffers.

[6] We have used the DPD package (Arellano and Bond (1991 and 1988)), the GMM estimator of which is specially designed to obtain unbiased and efficient estimates in dynamic models with lagged endogenous variables as regressors.

[Table 2]

The long-term semi-elasticity of the buffer relative to GDP growth, calculated at the sample buffer average, is 0.17. That is, an increase of 1 percentage point in GDP growth reduces the long-term relative buffer by around 17%. Given that the average sample of the relative buffer has, in the sample considered, been around 40% (in recent years, however, it is around 25%), and given also the usual magnitude of the changes in the rate of growth of GDP, the impact of the cycle, despite being very significant, seems to be moderate in quantitative terms.

The remaining parameter estimates also provide some interesting results. Thus, some relevant persistence in the capital buffer is detected, which, as expected, reveals the existence of non-negligible short-term adjustment costs. The capital cost has, as might be expected, a significant negative effect on the surplus capital that the institutions wish to maintain. Moreover, banks that according to the proxy chosen have a more conservative profile tend to hold higher buffers to meet potential adverse shocks[7]. Kwan and Eisenbeis (1997) also find that the default ratio and the capital ratio are negatively related. And finally, the signs of the dummy variables *BIG* and *SMA* are, respectively, consistent with the too-big-to-fail hypothesis and the relatively greater difficulties for small banks to draw on capital markets. Nevertheless, both variables are only marginally significant (the p-values of the corresponding tests being 0.15 and 0.12).

The estimated equation, on the other hand, passes without any major problem the standard goodness-of-fit tests. All variables have, as commented, the expected sign and most of them are significant even at 1%, and there is significant negative first-order autocorrelation in the residuals (m_1 statistic) and nil second-order correlation (m_2), as should be the case if the error term (in levels) is white noise. The Sargan test for validity of the instruments used is also fully satisfactory, showing a p-value of 0.26.

As commented in section 2, our next step is to test the robustness of these results by trying different extensions of equation (9). To start with, we have extended equation (9) to include variables capturing the specific composition of banks' balance sheets. There are several reasons why such variables might help to explain capital buffers. Thus, Basel I requires different capital ratios depending on the type of exposure. In the same vein, banks themselves might vary their buffers depending on the risk profile of their different portfolios. In particular, we have added three new regressors to equation (9), namely the weight in terms of total assets of i) loans (LOTA), ii) stocks (STTA), and iii) sovereign (SOTA).

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[7] As, for instance, in Wall and Peterson (1995), unexpected shocks might be approximated by means of the standard deviation of the rate of return in previous periods. Yet a standard deviation calculated on the basis of a few observations may not be very significant. For us it has

The second column in Table 2 shows the results of this modified version of our basic equation. Two main results emerge. First and more importantly, the estimated parameter for GDPG remains almost unchanged, thus supporting the robustness of the results in the first column of Table 1. And second, the new regressors are not significant, except for the last one. Significantly, their estimated signs are the expected ones should these variables reflect the risk profile of the bank (i.e. more loans, more stocks and less sovereign exposure are likely to reflect a riskier profile). This might explain why they are not significant in an equation that already includes NPL and a fixed effect.[8]

The next extension we have tried has to do with the possibility that the pro-cyclical behaviour of capital buffers we find is biased as result of the lack of control for other macroeconomic variables. To be relevant, those cyclical macroeconomic variables must affect loan demand –which certainly is clearly pro-cyclical. We have therefore included loan growth (LOANG) as an additional regressor in equation (9). It is worth noting that there might be another "technical" reason for including LOANG. Thus, an increase in loans implies an increase in capital requirements, which, in a context where adjusting capital is costly, is likely to transitorily reduce capital buffers.

The third column of Table 2 shows the results of including loan growth (LOANG) in the regression. The new variable is significant at the 1% level and, as expected, its sign is negative –a contemporaneous increase in loan demand erodes the capital buffer. More importantly, however, the pro-cyclicality of the capital buffer remains almost unchanged, as well as the sign and significance of the other variables.

Similarly, we have also tested whether our results might be influenced by the fact that the non-performing loans ratio is a markedly cyclical variable. Arguably, because of its cyclical behaviour, this variable could influence the sign and significance of GDP growth. The fourth column of Table 2 shows that if non-performing loans are excluded, GDP growth is still significantly negative. The change (i.e. bias) in the point estimate of the parameter of *GDPG* reveals, as expected under the null, that a relevant variable has now been omitted from the equation.

Finally, we have also considered the possibility that capital buffers are maintained not only to withstand contemporary unexpected shocks but also to cover problem loans in future periods. A simple way of controlling for these effects is to use *future values* of such variable as instruments. The last column in Table 2 shows that

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the additional disadvantage that it involves losing the first years of information, substantially reducing the sample size.

[8] Nothing changes if each individual variable –or pair– is considered in turn. We have also tested another potentially relevant balance sheet item (interbank exposures) and obtained the same results.

the main results in column 1 are not altered (either as regards their sign or significance) when future values of non-performing loans ratios are used as instruments.[9]

All in all, we can conclude from Table 2 that the capital buffer held by Spanish institutions has behaved pro-cyclically over the last 15 years. Also, the capital buffer is found to depend, fairly robustly, on the risk profile of the institution, the cost of holding such a surplus and, to a lesser extent, the institution's size.

5. Capital buffer cyclicality: some additional results

After testing the robustness of the results, in this section we include some additional extensions that provide some further evidence related to the capital buffer pro-cyclicality found in the previous section. Table 3 summarises our main findings.

[Table 3]

First, we have used a different measure of the business cycle, which takes into account the possibility of non-constant potential output growth. The first column in Table 3 shows that the pro-cyclicality remains if GDP growth is replaced with the output gap (*OUTGAP*), obtained after applying a standard Hodrick-Prescott filter. The long-term semi-elasticity falls somewhat (12%) while the other properties of the model remain unchanged.

Next, we also investigated whether the pro-cyclicality of capital buffers could depend on specific features of the banks. In this respect, we first tried banks' size, interacting *GDPG* and the dummy variables *BIG* and *SMALL*. As can be seen in the second column in Table 3, we did not find any meaningful difference in the behaviour of big or small banks.

Another interesting feature that might affect the cyclicality of capital buffers is the ownership structure of the different institutions. In particular, the sample analysed in this paper is made up of commercial and savings banks. The former are all in the hands of private shareholders (concentrated to a greater or lesser degree depending on each bank) while the the governance of savings banks is shared among representatives of several stakeholder groups, public authorities (from local and regional government), the founding entity, depositors and workers. There is extensive empirical

[9] We also performed other robustness tests that are not reported here for the sake of conciseness. It might be worth mentioning, however, that we introduced dummy variables to control for the regulatory change that occurred in 1993, without finding any significant effect on the conclusions drawn from the first column in Table 2.

literature on the impact that different ownership structures and distinct corporate governance arrangements may have on the risk profile of institutions[10].

We have, therefore, interacted GDP growth and a dummy variable (*COM*) taking a value equal to one when the institution is a commercial bank and to zero otherwise (i.e. when it is a savings bank). As shown in the third column of Table 3, this variable is only marginally significant, although it has a positive sign, meaning that capital buffers in commercial banks are, if anything, less pro-cyclical than in savings banks. The remaining properties of the model are not affected in any meaningful way.

Finally, we have tested whether the pro-cyclicality we have found is symmetric, i.e. whether it operates in the same way during upturns and downturns. In particular, we have added a new variable to the right-hand side of equation (9): the absolute value of the difference between the output growth and the average output growth in the sample. As shown in the last column in Table 3, this variable is negative and statistically significant (even at 1%). Accordingly, when output growth is above its average, buffers diminish proportionally more than they increase when growth is below its average. The long-term semi-elasticity during upturns increases to 29%, standing at 12% during downturns.

6. Conclusions

The design of a new capital accord (Basel II) has prompted an interesting debate among regulators, supervisors, academics and practitioners. The issue of the potential pro-cyclicality of the new capital requirements is currently playing an important role in the debate.

While most arguments about the cyclicality of the new agreement are of a purely theoretical nature and are centred on the capital requirements themselves, this paper aims to provide some empirical evidence and focus on the behaviour of the capital buffers that most banks hold above the minimum required by domestic regulations. The cyclicality of these buffers might offset or add to the potential cyclicality of the requirements.

Using annual data on Spanish banks from 1986 to 2000, we have built an incomplete panel of 1309 observations and estimated an empirical equation to explain how capital buffers have behaved in that period. After controlling for other potential determinants of the excess of capital –cost of capital, risk profile of the bank, adjustment costs, size and unobservable idiosyncratic features– we have found a fairly robust and significant negative relationship between the capital buffers and the business cycle. In quantitative terms, the pro-cyclicality effect is moderate: an increase of 1 percentage point in GDP growth might reduce the buffer by 17%. This is likely to

[10] See, for example, Esty (1997), Gorton and Rosen (1995), Saunders et al (1990) or, regarding the Spanish case, Salas and Saurina (2002b).

explain why, despite the markedly cyclical behaviour of capital buffers, banks have managed to keep fairly safe levels of capital even at the depths of recessions.

Our results are obtained under the capital accord still in place (Basel I) and, therefore, the conclusions to be drawn regarding Basel II should be taken with due caution. It is still an open question, though, whether the new accord has necessarily to change the behaviour of banks regarding the buffers they maintain over requirements which, on average, are not going to be increased. According to our findings, the issue of pro-cyclicality seems to merit serious attention in the final design of Basel II. It is worth noting that this problem is difficult to address under Pillar 1 without being to the detriment of the objective of more directly linking capital requirements to actual risks. Pillar 2 might thus be a most suitable tool for dealing with this issue. In particular, a closer monitoring of the behaviour of banks' own resources during the expansionary stages of the business cycle would appear to be justified in order to prevent potential negative effects on solvency should a sudden cyclical correction occur.

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Table 1. Summary statistics

Sample period: 1986-2000 (annual data)
Sample size: 1309 observations

Variable	Mean	Standard deviation	Minimum	Maximum
BUF	40.29	40.44	-76.60	240.10
ROE	15.28	10.62	-75.77	57.08
NPL	4.55	3.26	0.00	27.34
LOANG	16.27	16.22	-65.16	190.03
SOTA	12.18	7.86	0.00	48.65
LOTA	49.12	11.11	10.92	98.71
STTA	1.64	1.82	0.00	18.98
GDPG	3.27	1.70	-1.03	5.55
OUTGAP	0.14	1.36	-1.82	2.16

Notes:
- Variables are defined in the main text

Table 2. Estimation of equation (9)

Dependent variable: BUF $_{i,\,t}$

Sample period: 1988-2000 (1029 observations)
Estimation method: GMM, equation in first differences

Explanatory variable	Model 1	Model 2	Model 3	Model 4	Model 5
BUF $_{i,t-1}$	.40 (.00)	.27 (.00)	.37 (.00)	.42 (.00)	.43 (.00)
ROE $_{i,t}$	-.43 (.01)	-.41 (.02)	-.35 (.03)	-.51 (.00)	-.39 (.01)
NPL $_{i,t}$	-1.99 (.00)	-3.82 (.00)	-2.31 (.00)	--	-2.51 (.00)
BIG $_{i,t}$	-14.06 (.15)	-24.57 (.03)	-11.62 (.24)	-13.71 (.13)	-15.97 (.11)
SMA $_{i,t}$	21.74 (.12)	24.09 (.05)	15.12 (.19)	19.64 (.17)	18.20 (.17)
GDPG $_{t}$	-4.09 (.00)	-4.03 (.00)	-3.86 (.00)	-2.14 (.00)	-4.76 (.00)
LOTA $_{it}$	--	-.55 (.11)	--	--	--
STTA $_{it}$	--	-3.53 (0.35)	--	--	--
SOTA $_{it}$	--	.54 (.05)	--	--	--
LOANG $_{it}$	--	--	-.22 (.01)	--	--
m1	-4.65 (.00)	-4.24 (.00)	-4.68 (.00)	-4.52 (.00)	-4.55 (.00)
m2	.16 (.87)	.28 (.78)	.57 (.58)	.23 (.82)	.14 (.89)
Sargan test	114.15 (.26)	120.25 (.48)	125.33 (.31)	105.55 (.13)	116.21 (.35)

Notes:
- See the main text for the definition of the variables
- p-values in brackets
- m1 and m2 stand for first - and second - order residual autocorrelation tests
- In all models, BIG, SMA and GDPG are considered as exogenous
- Instruments for the endogenous variables: lags 2 to 4 in model 1, lags 2 to 3 for BUF, NPL, ROE and LOTA and lag 2 for SOTA and STTA in model 2, lags 2 to 4 for BUF and NPL and 2 to 3 for ROE and LOANG in model 3, lags 2 to 5 in model 4 and leads 1 to 4 for NPL and lags 2 to 4 for BUF and ROE in model 5, as selected in DPD (Arellano & Bond, 1991)

Table 3. Extensions of equation (9)

Dependent variable: $BUF_{i,t}$

Sample period: 1988-2000 (1029 observations)
Estimation method: GMM, equation in first differences

Explanatory variable	Model 1	Model 2	Model 3	Model 4
$BUF_{i,t-1}$	.30 (.00)	.40 (.00)	.41 (.00)	.39 (.00)
$ROE_{i,t}$	-.56 (.00)	-.44 (.01)	-.43 (.01)	-.32 (.04)
$NPL_{i,t}$	-.91 (.04)	-2.04 (.00)	-1.96 (.00)	-2.04 (.00)
$BIG_{i,t}$	-15.75 (.07)	-7.36 (.40)	-13.17 (.18)	-13.04 (.17)
$SMA_{i,t}$	24.84 (.07)	12.21 (.45)	22.16 (.11)	21.77 (.11)
$GDPG_t$	--	-4.18 (.00)	-4.78 (.00)	-4.99 (.00)
$OUTGAP_t$	-3.27 (.00)	--	--	--
$BIG_{i,t}*GDPG_t$	--	-2.40 (.28)	--	--
$SMA_{i,t}*GDPG_t$	--	2.82 (.14)	--	--
$COM_{i,t}*GDPG_t$	--	--	2.21 (.11)	
$\|GDPG\|_t$	--	--	--	-2.08 (.00)
m1	-4.58 (.00)	-4.63 (.00)	-4.57 (.00)	-4.69 (.00)
m2	.15 (.88)	.17 (.87)	.23 (.82)	.13 (.90)
Sargan test	116.83 (.20)	114.43 (.25)	112.02 (.30)	114.12 (.26)

Notes:
- See the main text for the definition of the variables
- p-values in brackets
- m1 and m2 stand for first- and second-order residual autocorrelation tests
- BIG, SMA, GDPG, OUTGAP and COM are considered as exogenous
- Instruments for the endogenous variables: lags 2 to 4 in models as selected in DPD (Arellano & Bond, 1991).

Chart 1. Capital buffer in relative terms (current capital less capital requirements over capital requirements): 1988-2000. Percentage points.