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Bank capital, lending booms, and busts: Evidence from Spain over the last 150 years

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ABSTRACT

In this paper, we analyze the role bank capital played in systemic banking crises and in lending expansion and contraction for nearly 150 years in Spain. We first build a measure of capital ratio (i.e., the capital to assets ratio) for Spain's banking sector, starting in 1880. Then, we analyze if more capital reduces the probability of a banking crisis using a narrative identification of banking crises in Spain. Afterwards, we run a proper econometric test to analyze bank capital levels' impact on lending cycles, controlling for other determinants of credit growth. We find robust empirical evidence that higher levels of capital reduce the probability of a crisis and that increasing the bank capital before loan expansions reduces credit growth and reduces the credit's decline if a systemic crisis materializes. Conversely, overly depleted bank capital when entering a credit contraction period severely impacts lending (i.e., may bring about a deep credit crunch), with negative and lasting effects on the economy and on society's wellbeing as a whole. The paper is particularly useful for supporting the credit cycle smoothing role of recently implemented macroprudential policies (countercyclical capital buffer).

1. Introduction

The Global Financial Crisis (GFC) has shown the lasting impact that bank failures and/or bank recapitalizations may have on the economy and on society as a whole.¹ Soon after the most acute phase of the crisis, policy makers and regulators began soul-searching to understand and amend what went wrong with banks before the crisis. Thus, the G-20 mandated the newly renamed Financial Stability Board (FSB) to examine the GFC's causes and remedies. In subsequent years, the FSB pushed for reforms in banking regulation and beyond.²

One main empirical conclusion of bank regulators, supervisors and policy makers after the crisis was that banks had too little capital when entering the crisis. Capital had expanded—or, even worse, returned to shareholders via buy-backs or higher dividend

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¹ The views expressed in this paper are those of the authors alone and do not necessarily reflect the views of the Banco de España, the Eurosystem or the Single Resolution Board. We thank the Banco de España History Seminar participants, those of the Workshop of Economic History at RIDGE Forum in Montevideo, two anonymous referees and, last but not least, Pablo Martín-Aceña for their very useful comments on this paper.

¹ See, for instance, Chart 1 in Estrada and Saurina (2016).

² See, among others, (Financial Stability Forum, 2009) and G-20 press releases for the reforms' sense of urgency and direction. In particular, the one after the meeting in London in April 2009 (see the (G-20 press release, 2009)) asked for a significant increase in bank capital and for countercyclical buffers and better provisioning requirements.

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streams³—less than in the previous lending expansion, partially based on the notion that credit risk was better measured and managed, so less capital was needed per unit of exposure. The spread and strength of the banking crisis clarified the weaknesses of such an approach that, at a regulatory level, underpinned the Basel II regulatory reforms that started in 1999 and concluded in 2004 (see BCBS (Basel Committee on Banking Supervision 2004) and BCBS (Basel Committee on Banking Supervision 2006) for additional changes before the regulation came into force in 2008). Moreover, the financial crisis underlined the need to move from a pure microprudential view of capital requirements (i.e., focused only on the capital level at each individual bank) to a more comprehensive or holistic view of capital, where the macroprudential approach is a key ingredient.⁴ That is, policy makers must pay attention to the aggregate bank capital ratios and to how they evolve along the business cycle. In this respect, capital requirements need a countercyclical component. When the lending cycle is in full swing, banks must reinforce the level of capital to protect them from the risk expansion, and, more importantly, to build a buffer to cover future losses that may appear when economy trends change. The need for more capital combined with a countercyclical capital buffer (CCyB) in each bank (both proportional to each bank's risks) paved the way to the Basel III agreement (Basel Committee on Banking Supervision 2011).

The empirical analysis shows that higher aggregate capital ratios can lower the probability of a systemic banking crisis (Basel Committee (2010)). Besides, solid and robust evidence shows a countercyclical tool's necessity (Jiménez and Saurina, 2006) and its usefulness to reduce credit crunches during recessions (Jiménez et al., 2017). This research also fits the international literature analysis regarding bank capital's potential role in credit crunches ((Bernanke and Lown, 1991, Rajan, 1994, Peek and Rosengren, 1995, Berger and Bouwman, 2013, Demirgüç-Kunt et al., 2013) and (Vazquez and Federico, 2015)) and in the whole lending cycle (see for instance, (Gambacorta and Mistrulli, 2004) or (Saurina and Shin, 2012)). These previous studies provide crucial support for macroprudential policies.

This paper aims to significantly expand this perspective, so we can learn from the last lending boom and bust associated with a systemic banking crisis in Spain and from a long list of previous ones. We based our analysis on a long period of Spanish history with various banking environments (e.g., different levels of banking concentration and competition, different economic policy environments, different institutional arrangements, etc.). Thus, we add value by focusing on many lending booms and busts in a country particularly prone to them and to banking crises—both systemic and bank idiosyncratic.⁵ Additionally, our approach contributes to the literature by complementing (Elliot et al., 2013) and (Elliot, 2014). Instead of analyzing past policy decisions as potential macroprudential tools, we focused on one regulatory or policy instrument (i.e., bank capital level) and assess its role along various credit cycles.

The paper has a second important contribution. To our knowledge, this is the first elaboration of a long series of bank capital ratios for the Spanish banking system. Thus, we can compare the historical levels of capital ratios in the Spanish banking sector with, for instance, the ones in the U.S. (Berger et al., 1995) and in the UK (Alessandri and Haldane, 2009) for long time periods.

For our methodology, we first estimated a traditional logit model for the banking systemic crises identified via narrative procedures that consider real and financial determinants, including the capital ratio. Afterwards, we showed that these banking crises concur with periods of credit crunches. Finally, we applied a state-of-the-art analysis for estimating the relationship between bank capital levels and lending cycles. Following (Bernanke and Lown, 1991) and (Peek and Rosengren, 1995), we modeled aggregate credit growth as a function of macroeconomic determinants, including the level and evolution of bank capital, using a nonlinear model to endogenously distinguish between periods of high and low credit growth.

We found that the level of bank capital has an inverse relationship with the probability of a systemic banking crisis; in particular, increasing the capital ratio by one percentage point reduced that probability by one percentage point. It also had an asymmetric effect in the credit cycle so that increasing the capital ratio before a credit boom reduced credit growth. On the contrary, in the regime state where credit growth is low or negative, increasing the capital ratio in advance increased the average credit growth or lowered the contraction in credit. These results could indicate that a countercyclical capital buffer reduces the size of an expansion and the magnitude of the subsequent credit crunch.

Therefore, we find support for using a particular macroprudential tool (e.g. CCyB) in various economic and banking environments over nearly a century and a half. Thus, in very different economic development stages and banking structures (with or without foreign banks, with more or less stringent regulation, with more or less oligopolistic structures), we find support for countercyclical tools.

In the next section, we detail how we constructed the capital ratio variable and the other variables. In the third section, we present the methodological approach and, in the fourth section, our main empirical results. We end by summarizing the main conclusions.

2. Databases and a preview of the historical capital ratios

To the best of our knowledge, a long time series for an aggregate capital ratio for the Spanish banking sector does not exist, so we have created one. To fulfill that purpose, we had to combine information from several sources that were not always initially compatible. For instance, we used information from several Spanish Statistical and Bank Bulletins with some historical books to complete our database. We believe this is an important contribution of the paper.

³ See (Acharya et al., 2011) and (Acharya et al., 2017) for estimates of the importance of those buy-backs and dividend payments for U.S. banks.

⁴ To understand the macroprudential approach see, among others, (Crockett, 2000, Borio, 2003, Hanson et al., 2011) and, more recently, (International Monetary Fund 2014, IMF-FSB-BIS, 2016), Mencia and Saurina (2016) and (Bank for International Settlements, 2017).

⁵ An excellent narrative of banking crises in Spain and their driving factors appears in (Martín-Aceña, 2013); we used this narrative to estimate the impact of capital ratios on the probability of this type of crisis occurring in Spain.

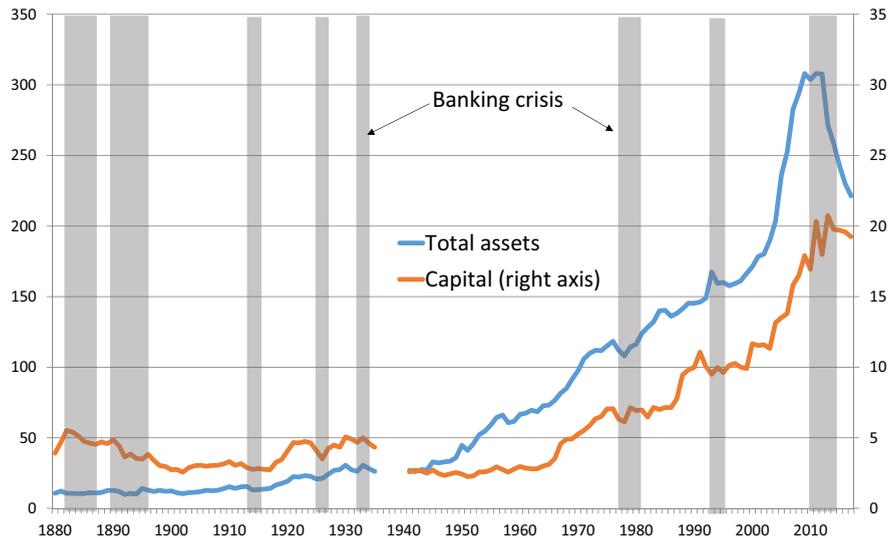


Fig. 1. Total Assets (left-axis) and Capital (right-axis) over GDP (%). From 1880-2017.
Sources: BdE and own calculations.

For the banking system balance sheets, the departure point information was available in Chapter 4 of the Banco de España *Statistical Bulletin*. This source offers the banking system's homogeneous balance sheets from 1962 onwards. These balance sheets show information about banks and savings banks.⁶ We aggregate most of the assets and liabilities entries because the available breakdown was reduced and disparate in previous years.

Data from 1880 to 1914 were obtained from (Anes et al., 1974). To cover from 1915 to 1924, we used the series constructed by Martín-Aceña in a book by (Carreras and Tafunell, 2005). We used several Bank Bulletins of the Bankers Association to obtain Spanish banks' balance sheet information from 1924 to 1935 and from 1941 to 1958. The information from 1952 to 1962 also came from past Banco de España Statistical Bulletins.

However, the aggregated balance sheets for common years in different Statistical Bulletins, Bank Bulletins and books do not exactly match. Thus, we adopted a statistical approach instead of an accounting one to compare them. We considered that a higher aggregation meant a higher compatibility among various databases. Thus, we enlarged the most recent database backwards using the corresponding growth rates of older databases. Afterwards, we adjusted the various asset and liability items to guarantee they added up. In all, we built a consecutive series of values for total assets, total credit and capital.

To scale all these variables and facilitate the empirical analysis presented below, we had to backtrack two macroeconomic variables: nominal GDP and CPI (Consumer Price Index). Both variables came from the Spanish Statistical Bureau (INE, its acronym in Spanish) database and were linked backwards using growth rates from (Carreras and Tafunell, 2005). Due to data constraints, we calculated the real GDP (and the rest of the real variables appearing in this paper) by dividing nominal GDP over CPI. Using the GDP deflator would not significantly modify the results.

Fig. 1 presents the evolution of total assets and capital of the Spanish banking system as a percentage over nominal GDP over the last 137 years. The shaded areas represent the systemic banking crisis identified narratively in (Martín-Aceña, 2013). As can be seen, the total assets over GDP remained relatively stable from 1880 to 1915. Afterwards, the ratio continued an upward trend, only interrupted by the following systemic financial crises. Total banking assets as a percentage of GDP is a rough measure of the leverage or indebtedness of an economy, particularly for economies where banks dominate the financial landscape and market funding is relatively low (such as the Spanish economy). Fig. 1 shows how different Spain's last lending boom and bust was—households' and firms' leverages increased substantially before the full-fledged crisis started, with an impressive correction afterwards that only finished by 2017. The capital over GDP ratio has remained more stable. In fact, at the sample's beginning, it represented around 5%, showing a downward trend until 1920 and a recovery before the Spanish Civil War. Right after that period, it went from a very low level (2.5%) to recovering the previous century's ceiling in 1970. Afterwards, a sustained increase appears. Visually, this ratio has diminished during all recorded systemic banking crises apart from the last one, when it markedly increased.

From a different perspective, the green line in Fig. 2 shows Spain's capital ratio over total assets. After reaching values over 50% at the end of the 19th century, it noticeably declined, reaching values around 15% before the Civil War. The first two systemic banking crises in our sample—in the 19th century—implied a huge relative depletion of capital that was never replenished. In fact, this downward trend continued afterwards, registering a floor of 4% by 1962. The trend then became positive until 1990, where it reached its maximum value (7.6%) since 1946. Then, it declined significantly until the GFC. Contrary to previous ones, this last crisis

⁶ Savings banks were not as prominent before the early 1990s, when they started to open branches nationwide and expanded quite significantly. On the contrary, for most of the period analyzed, commercial banks dominated Spain's banking landscape.

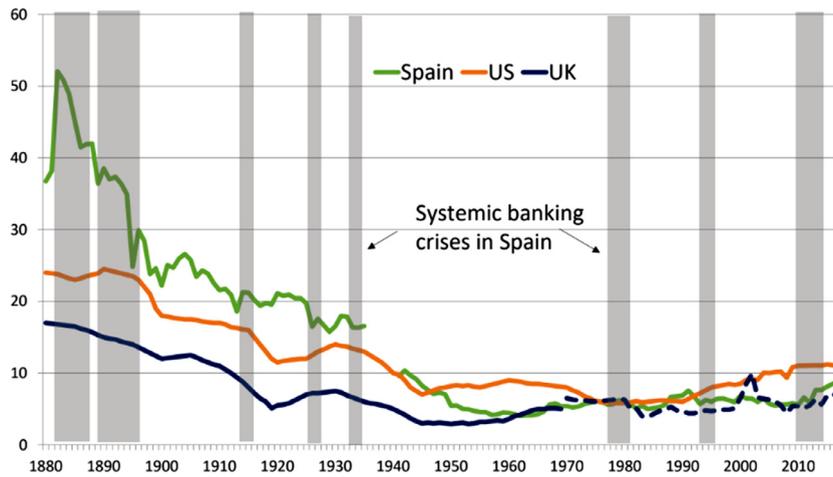


Fig. 2. Capital over total assets (%) for Spain, the U.S. and the UK. From 1880-2017.

Notes: Commercial US banks and UK banks operating in the UK are considered.

Sources: BdE, own calculations, Berger et al. (1995), Sheppard (1971), Federal Reserve Economic Data and World Bank's World Development Indicators. The UK's dotted line includes banks' reserves and not only capital, given that after 1970 published accounts do not separate both entries.

showed an important increase in this ratio, reaching levels not observed since the early forties. The increase resulted from different factors. First, the increase in capital was brought about by stringent new regulations (i.e., the enforcement of Basel III). Secondly, this increasing ratio reflects recapitalization with the public funds of failed savings banks. Last, but not least, it may also reflect the significant deleverage process (e.g., a reduction of private indebtedness) that came after the worst banking and economic crisis since the Spanish Civil War. Therefore, capital is apparently being depleted during banking crises because credit losses must be absorbed. If losses are large enough, given the low bank capital ratios mentioned in the introduction, they may bankrupt the bank, thus forcing its exit from the market. Alternatively, banks can be recapitalized with public funds, as the last three systemic banking crises in Spain and in other developed countries prove.

Fig. 2 also includes the historical evolution of similar ratios in cases in the U.S. (orange line) and in the UK (blue line). As can be seen, Spain had the highest ratio until the Civil War, followed by the U.S. and the UK. All three converge by the mid-fifties and remain less volatile, probably due to the first international initiatives in cooperative banking regulation. After the Civil War, Spain's capital ratio has been quantitatively similar to that of the UK. However, the three ratios shared a diminishing trend until the 1950s. Afterwards, a stable period or a slight increase appeared, followed by a reduction in the mid-nineties and an important increase right after the GFC, as mentioned. (Benink and Benston, 2005) also graphed several countries' average capital ratios for a time period similar to ours. They have common elements with our chart. In particular, the trends share an almost constant path from the mid-fifties onwards. This similarity holds even when considering a different basket of European countries. Further, (Schularic and Taylor, 2012) use a period similar to ours regarding different countries (not exclusively European, as opposed to (Benink and Benston, 2005)), reinforcing the importance of credit cycles in determining the likelihood of financial crises.

3. Methodological approach

3.1. The models

In the empirical analysis, we estimate two kinds of models. The first is a probabilistic model highlighting the main determinants of Spanish systemic banking crises identified via a narrative approach. In this respect, we estimate several logit models of the following form:

$$Pr(y_i = crisis/x_{it-h}, \beta_i) = \frac{e^{x'_{it-h}\beta}}{1 + e^{x'_{it-h}\beta}} \quad (1)$$

The underlying idea is that a latent unobservable continuous variable (y^*) exists that depends linearly on the crisis determinants (x). When that latent variable exceeds a certain threshold, a crisis event happens that makes the model nonlinear. We assume the error term of the underlying linear relationship follows a logistic model to take into account that these events are rare. As explanatory variables, we consider macroeconomic aggregates (GDP growth and current account balances), financial (real interest rates, proxied by the nominal yield of public debt instruments minus the inflation rate and measured via the CPI) and banking sector variables (credit growth and the average loan values in the housing sector, which is the total value of mortgages divided by the total value of the housing stock) as well as the capital ratio. We included all these variables as lagged to avoid endogeneity worries and to check their early-warning properties. The variables' lags were chosen according to their relevance and the best fit of the regression. We expected the capital ratio to be negative, implying that a higher ratio means a lower probability of a banking crisis.

Previous models drew some major criticisms. First, identifying a crisis via narrative techniques uses a combination of qualitative and quantitative information, which can differ from one paper to another (see, for example, (Laeven and Valencia, 2008) or (Reinhart and Rogoff, 2013)). Second, the dependent variable's (the number of systemic crises) number of positive outcomes is usually reduced—our historic sample has only eight systemic crises, but most used samples (luckily) have no more than three episodes. Third, this approach does not allow an analysis of the crises' depth and the duration. Thus, we estimated a second class of nonlinear models that rely on an observed continuous variable (capital ratio) instead of on a binary one. This paper's final objective is to shed light on bank capital's role in preventing or exacerbating systemic financial crises, so we followed the previous literature ((Claessens et al., 2012) and (Jordà et al., 2016)) and opted for estimating a credit to the non-financial private sector equation. As we show below, both approaches have a lot in common. Other widely used possibilities include modeling housing or stock market prices.⁷

In this second approach, we studied the capital ratio's explanatory power for credit after controlling for traditional credit determinants (GDP, real interest rate, housing prices, deviations from its long-run equilibrium, etc.). Bank capital should be lagged to simulate a policy's operational conditions (for example, when the CCyB is activated, banks have one year to meet the new requirement). As in the previous case, the variables' lags are chosen to maximize the fit of the regression's fit.

If we introduce the capital ratio in the equation linearly and find a negative and statistically significant sign, then higher bank capital means lower credit growth in any situation. However, it could be interesting to check if that coefficient differs during stress periods (when credit declines) and during tranquil periods (when credit increases). The easiest way to check this possible nonlinearity is identifying financially stressed periods using a systemic banking crisis identified narratively, with tranquil periods for the remaining observations, and estimating separate coefficients for the capital ratio in both subsamples. Unfortunately, this methodology assumes the identified systemic crises as given.

To solve that problem, use a Markov-switching regime model (Hamilton, 1989). In this class of models, the data identify the periods with a high probability of being in a regime state of reduced or negative credit growth, which is traditionally linked to systemic banking crises. In fact, a Markov-switching regime model allows the economy to reach different states, where the probability of transitioning states only depends on the current state. We analyze filtered probabilities that depend on contemporaneous information at each state.

In any case, both approaches are interlinked. In fact, we can run an autorregressive Markov-switching model for real credit while considering only one explanatory variable, a constant, that can differ in the two regime states:

$$\Delta cr_t = \beta_0(S_t) + \beta_2 \Delta cr_{t-1} + \varepsilon_t; \quad \varepsilon_t \sim N(0, \sigma_\varepsilon^2) \text{ and } Prob(S_t = i / S_{t-1} = j) = p_{ij}$$

where $i, j = 1, 2$; where Δ is the differences operator (that accounts for the variable's integration order). Lowercase letters refer to variables measured in logs, and CR stand for real credit.

In this very simple model, the constant captures the conditional average growth rate of real credit in every regime state while accounting for $\beta_2 \Delta cr_{t-1}$. Then, we calculate the implicit probabilities of being in the regime state where credit grows the least, which should correspond to periods associated with systemic banking crises.

The time series of credit to the non-financial private sector used in this paper for Spain come from the BIS database (1970–2017), for which the original source is the country's financial accounts. We use total credit, which is bank credit plus all fixed-income instruments issued by non-financial private companies.⁸ These time-series are backtracked until 1880 using the credit aggregate from banks' balance sheets, as explained in the previous section.⁹

Fig. 3 shows the growth rate of total credit in real terms (i.e., deflated by the CPI). First, the growth rate shows a notable volatility. Growth rates have fluctuated above 40% in absolute value. Second, credit's real growth declines in all systemic banking crises. This suggests that credit's real growth is promising for capturing this kind of event. As seen in the table accompanying Fig. 3, the descriptive statistics measured by mean and standard deviation are very different before and after the Civil War. This volatility is remarkable; the value before the Civil War was almost double the value afterwards.¹⁰

Table accompanying Fig. 3: Descriptive statistics of total credit growth

	Pre-Civil War 1880–1935	Post-Civil War 1942–2017	Full sample 1880–2017
	Total credit growth	Total credit growth	Total credit growth
Mean	0.043	0.062	0.054
Std. Dev.	0.149	0.075	0.112

Fig. 4 shows the main estimation results of the simple Markov-switching regime model presented above. In all systemic crises identified via narrative techniques, the probability of being in the low credit growth regime is near 100%. However, the episodes'

⁷ See (Dupery and Klaus, 2017) for a recent discussion on several property market variables' effects on the probability of entering a financial stress regime. See (Dembiertmont et al., 2013) for a review of total and bank credit patterns for 40 countries since 1940.

⁸ For robustness, we repeated all estimates using bank credit to the non-financial private sector instead, and the results hold. They are available upon request.

⁹ We realize this is not the best way to proceed because these historical aggregates also include credit to the public sector (i.e., it is impossible to disentangle both as we go back in time), but we found no better alternative.

¹⁰ We conducted a robustness analysis where we divide the full sample into the pre- and post-Civil War periods. Results highly resemble the ones obtained when we used the full sample, which is why the results do not appear in the main paper. They are available upon request.

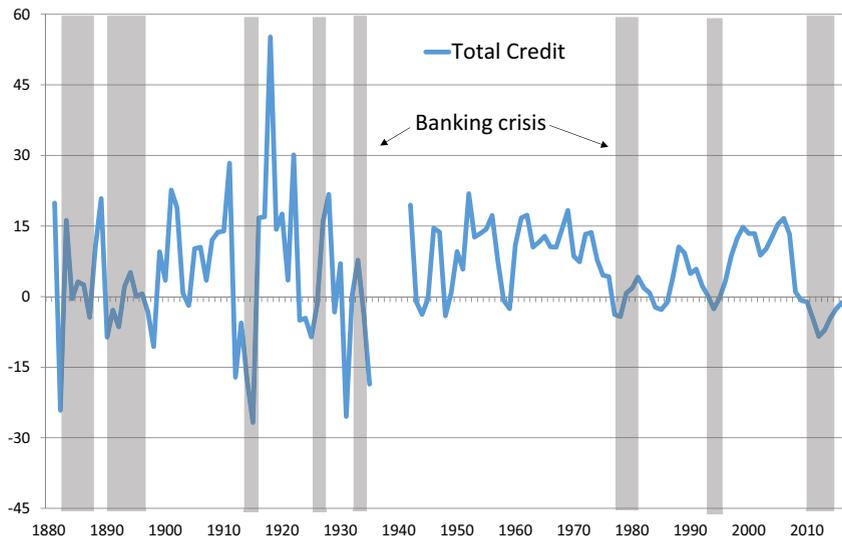


Fig. 3. Real growth of credit to the non-financial private sector. From 1880–2017.
Sources: BIS and own calculations.

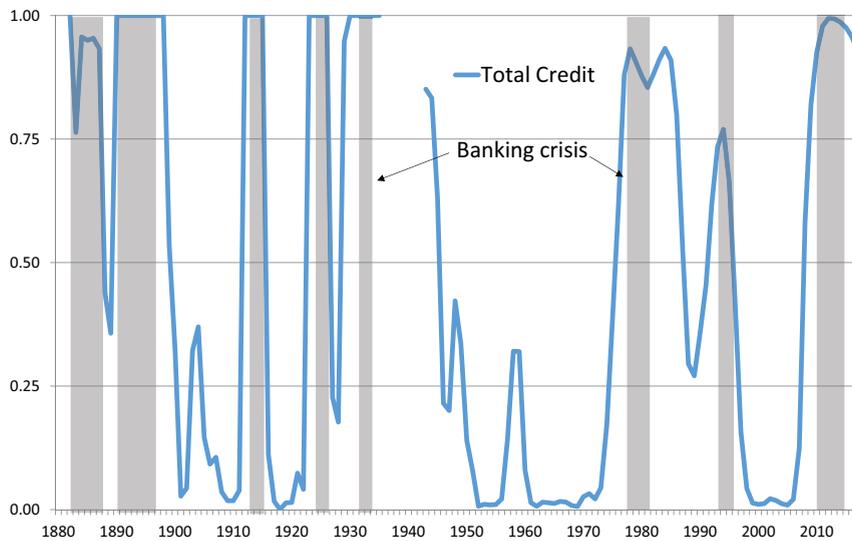


Fig. 4. Systemic banking crisis and probability of being in the low credit growth state. From 1880–2017.
Sources: Own calculations.

duration is not always matched (especially in the mid-seventies), and one period right after the Civil War shows model results compatible with a stress episode that was not identified as a systemic banking crisis via the narrative approach.¹¹

Thus, the model to be estimated will have this more general form:

$$\Delta cr_t = \beta_0(S_t) + \beta_2 \Delta cr_{t-1} + \beta_3(S_t) \Delta CAPR_{t-i} + \beta_4 \Delta y_t + \beta_5 \Delta RR_t + \beta_6 \Delta prv_t + \gamma ECM_{t-1} + \varepsilon_t; \quad \varepsilon_t \sim N(0, \sigma_\varepsilon^2) \text{ and } Prob(S_t = i / S_{t-1} = j) = p_{ij}$$

where $i, j = 1, 2$; where $CAPR$ is the capital ratio, Y is the income, RR is the real cost of funds, PRV is the collateral and ECM are deviations from the long-run equilibrium obtained via linear regression.¹²

¹¹ Regarding the estimates of the autoregressive Markov-switching model of credit growth, the estimated constant implies that total credit declines 2.9% on average yearly in the stress regimes. Besides, that regime's expected duration is 6.9 years. With high credit growth, however, the corresponding constant is 9%.

¹² In this expression we only consider two states of the nature and we only allow two parameters to change with them. We have chosen this specification after several tests, trying to retain the most parsimonious specification.

Regarding explanatory variables, it may be challenging to build the best determinants for this historical period. Thus, we use real GDP as the scale variable to capture income and part of the wealth effects. The cost of lending was measured with real interest rates. We also constructed a relative housing price as our proxy for the collateral (and partly for the wealth) behavior.

3.2. Estimation strategy

We estimated the models using several steps. First, we checked the variables' time-series properties to determine the differencing order for all variables to become stationary. Then, in the logit model, we did not consider observations that occurred after the beginning of each crisis during which the banking sector was still in a stressful situation. In this way, we restricted the model to identify the variables that trigger a crisis event, but not its duration.

Afterward, in the credit model, we assessed whether a cointegration relationship exists among the variables or a subset thereof. If that is the case, to obtain consistent estimates of the parameters in the stationary specification, it would be necessary to include an additional explanatory variable: the deviation of credit from its long-run equilibrium (i.e., an error correction term). This deviation captures the convergence with the long-run equilibrium; thus, the corresponding coefficient should be negative, statistically significant and smaller than one in absolute value. This cointegration relation was obtained using the fully modified least squares (FMLS) estimator, which has good properties for long time series such as we have in this case. We preferred this to the alternative of including the lagged levels of the relevant variables in the stationary specification and jointly estimated the long- and short-run parameters. In this nonlinear environment, it is preferable to estimate the most parsimonious specification possible.

To summarize the cointegration test's results,¹³ we emphasize that credit, real GDP and the capital ratio conform to a cointegrated system. The results are not very robust, but we do consider the estimated coefficients reasonable and in line with previous research. The elasticity with respect to the GDP is positive, and it is higher than the unity, reflecting how wealth effects implied a progressive increase of indebtedness, as shown in Fig. 1. The semielasticity of the equity ratio is negative, thus capturing cost and balance sheet effects. The real interest rate and the housing price are not statistically significant, probably due, in the first case, to the fact that nominal interest rates were regulated during most of the sample period¹⁴ In the second case, the joint consideration of credit extended to households (for which there is evidence of relevance) and non-financial private companies could be the simplest and most straightforward explanation, although it should also be taken into account that, prior to the Civil War, most Spanish households rented the houses in which they lived.

An additional difficulty with the estimation process is that the explanatory variables can be considered endogenous, as they can be determined jointly with the own credit. There is evidence that some GDP losses during financially stressful episodes occur because weak financial institutions impose supply credit constraints on the economy's real sector. In the case of real housing prices, credit conditions also play an important role in determining both the demand for housing services and the supply of new houses. In addition, the real interest rate is the equilibrium price of credit supply and demand, so it is clearly endogenous. Finally, the banking sector's capital ratio should also be considered endogenous, as banks probably choose the capital of the entity based on the volume of loans they would like to supply, as well as the sectoral composition of the credit portfolio, among other factors.

Under these circumstances, it is necessary to instrument these four variables or to use an adequate lag to obtain consistent estimators. The two properties that the instruments fulfill should correlate with the variable to be instrumented rather than with credit growth. Thus, in the case of GDP growth, we considered three instruments that captured demand shocks: world GDP growth, tourist arrivals to Spain and Spanish structural public balance. The first two instruments are determined abroad, so Spanish credit conditions do not affect them; the third is determined by the public sector, so domestic credit constraints should not affect it. For the real housing price, we added a supply side variable: the stock of houses per capita. This variable is predetermined with respect to prices because most houses are built before they are sold, so it should be a valid instrument. In the case of real interest rates, their two-period lagged values could be an adequate instrument, as this price results from the contemporaneous equilibrium between credit demand and supply. Finally, in the case of the capital ratio, the most important variable for this research, we used four instruments. The first two correspond to the banking sector capital ratios in two financial centers: the United States and the United Kingdom. As discussed above, the capital ratios in Spain, the U.S. and the UK shared similar trends, which should be unrelated to Spain's credit cycle. The third instrument is the share of agricultural added value represented by GDP. The evolution of this ratio is mainly determined by real factors (e.g., technological innovation, wealth, preferences) compared to financial ones; furthermore, agricultural added value is much more volatile than that of manufacturing or services, so banks must allocate more capital to loan financing for this activity. The fourth instrument is the Spanish banking sector's price-to-book ratio.

Thus, we ran regressions of GDP, real housing price growth and the capital ratio on these instruments to obtain their adjusted values. We then substituted these adjusted values for the observed values in the credit model to obtain consistent estimators of the parameters of interest. As the R^2 of these auxiliary regressions was relatively high,¹⁵ we concluded that the instruments correlate with the instrumented variables.

3.3. Comparison with previous econometric approaches

As mentioned in the introduction, our primary goal in this paper is to contribute to disentangling the role played by banks' capital during periods of credit expansion and, more importantly, during systemic banking crises. In that respect, several empirical

¹³ These analyses are available upon request.

¹⁴ In fact, this could help explain the sign and statistical significance of the capital ratio, as capital is the most costly funding source for banks.

¹⁵ These analyses are available upon request.

papers have estimated various binary probability models (e.g., logit, probit) in which the variable to be explained was the occurrence of a financial crisis, as in our case. For instance, (Demirgüç-Kunt and Detragiache, 1998), and (Hardy and Pazarbasioglu, 1998) used multinomial logit models to predict systemic banking crises. More recently, (Laina et al., 2015) used univariate signaling and multivariate logit models to study various systemic banking crises in the European Union. They showed that the growth rates of loan-to-deposit ratios and housing prices are the best indicators to predict systemic banking crises, followed by the growth rates of mortgages, household loans and private loans. In addition, (Betz et al., 2014), in focusing on the European banking market, argued that combining bank-specific and country-level macroeconomic and financial data such as house and stock prices greatly improved early-warning models' predictions of financial crises.

In this respect, our second approach, which used the evolution of credit to identify financial stress periods, was based on the banking crisis literature. It focused on how credit evolution of anticipates financial crises. For example, (Kaminsky and Reinhart, 1999) studied the effect of credit (among other variables) on banking and currency crises and concluded that credit extended to the private sector accelerates before a banking crisis emerges. (Borio and Drehmann, 2009) studied the relationship between credit growth and asset prices to determine the likelihood of future banking crises. (Adrian and Shin, 2010, Adrian and Shin, 2014) and (Saurina and Shin, 2012) stressed the importance of analyzing bank leverage to understand bank behavior during expansions and recessions.

In any case, this was not the first time a methodology similar to ours was used for the purpose at hand. For example, (Kaufmann and Valderrama, 2004) estimated equations used to calculate credit extended to households and firms using Markov-switching models, and (Anguren, 2011) built a synthetic indicator of global credit stance using the same methodology. However, to the best of our knowledge, ours is the first study to examine bank capital's role in credit evolution by connecting the traditional bank capital and credit crunch literature (Bernanke and Lown, 1991, Rajan, 1994, Peek and Rosengren, 1995). We also analyzed the role of bank capital and debt on credit from the supply-side perspective (i.e., taking into account the banks' solvency position), following much more recent literature (Jiménez et al., 2012, Jiménez et al., 2014, Jiménez et al., 2017) and (Gambacorta and Shin, 2016)), which used panel data and very granular information at the bank and loan levels.

4. Results

4.1. Predicting systemic financial crisis

The main estimates of the logit model appear in Table 1. The macroeconomic variables considered are GDP growth and current account balance. Both are statistically significant and negative, as expected. GDP growth appears lagged by one period, implying systemic banking crises are rare during expansionary periods. The external balance appears lagged by two periods, highlighting its early warning properties for banking crises. This captures the idea that extended periods of insufficient domestic savings to finance investments increase banks' external indebtedness and risk of experiencing a sudden stop in external financing. To relativize these

Table 1
Logit model for systemic banking crises in Spain for the period 1880–2017.

	(1)	(2)	(3)	(4)	(5)	(6)
Constant	-2.662*** (0.434)	-10.896** (3.113)	-10.105*** (2.252)	-9.982*** (2.188)	-9.313*** (2.122)	-7.091*** (1.566)
GDP growth lagged	-8.392* (4.628)	-23.411** (9.777)	-31.565** (13.795)	-30.983** (13.168)	-28.676** (13.427)	-
Current account balance lagged	-20.905* (11.581)	-67.641*** (16.776)	-72.414*** (22.318)	-76.626** (29.929)	-69.599** (29.171)	-86.040*** (28.858)
Real interest rate lagged	-	52.639*** (17.616)	54.141*** (17.603)	54.012*** (16.698)	46.938*** (17.015)	39.405** (15.467)
Credit growth inertia	-	32.477** (15.383)	25.194*** (8.632)	24.627*** (8.397)	22.514*** (8.097)	-
LTV increase lagged	-	10.093*** (3.121)	8.201* (4.193)	8.044** (3.965)	8.272** (3.894)	6.331* (3.471)
Capital ratio lagged	-	-	-155.319** (77.063)	-155.305** (71.468)	-156.755** (67.242)	-137.180** (56.945)
Deposit guarantee scheme dummy	-	-	-	-0.400 (1.990)	-0.341 (1.932)	-1.489 (1.840)
Global financial crisis dummy lagged	-	-	-	-	1.727 (1.811)	2.744** (1.318)
McFadden R ²	0.068	0.578	0.631	0.631	0.639	0.516
S.E. Regression	0.252	0.172	0.157	0.1593	0.158	0.179
Percent Gain (cutoff = 0.5)	5.09	53.97	61.39	61.17	62.42	50.56
Observations	114	107	107	107	107	108

Notes: The dependent variable is a dummy that indicates whether a systemic banking crisis started in a given year (dummy = 1) or not (dummy = 0). Coefficients are listed on the first row, and standard deviations appear below. The variables' lags were chosen according to their relevance and the best fit of the regression. Credit growth inertia measures the one-period lag of credit growth. *, **, ***, statistically significant at 10%, 5% and 1%, respectively.

outcomes, it should be taken into account that this model has a very reduced explanatory power; in fact, its predictions are only 5% better than one with only a constant on the right-hand side.

When we added macrofinancial variables to the model (real interest rates, credit growth and the variation in the loan-to-value of mortgages), the results improved substantially. In fact, the model's gain increased to 54%. These three variables were negative, and the first two entered with a lag of three and two years, respectively. Thus, periods of high real interest rates, high credit growth and increasing loan-to-values (LTVs) tended to precede systemic banking crises in Spain, confirming the international evidence. Adding the lagged capital ratio¹⁶ to this model did not change the previous results qualitatively, even though this variable was statistically significant; in addition, the model's fit and forecasting ability improved. It was negative, implying that higher capital ratios in the banking sector as a whole reduced the likelihood of a systemic banking crisis. This effect is also economically significant, as it implies that a capital ratio that is 1% higher (above the trend) reduces that probability by 1%, although the effect is smaller than that those found in other studies (see (Basel Committee on Banking Supervision 2010)).

During the lengthy period we considered, one of the most important structural changes in the Spanish banking sector, in terms of its implications for panics and bank runs, was the introduction of a deposit guarantee scheme in 1978. We controlled for this event using a dummy equal to one from that year on. It is negative, which implies a reduction in the probability of having a systemic banking crisis, although it is not statistically significant. Another structural change in the Spanish banking system was the internationalization of some of the biggest Spanish institutions and the Spanish economy's increasing openness to global financial flows. Thus, we constructed a global banking crisis dummy that captured banking crises in two out of three banking sectors: the United States, the United Kingdom and France (see (Reinhart, 2010)). When we introduced this dummy in the regression lagged by two periods, it was positive, implying that banking crises in Spain have a certain relationship with those in major financial centers. However, the dummy is only significant when output and credit growth are not included in the regression. This could imply that some financial risks that cause systemic banking crises are imported through credit and the GDP.

4.2. The cyclical behavior of credit extended to the non-financial private sector

We estimated a more traditional linear credit equation before estimating our nonlinear model. Columns (1) and (2) of Table 2 present this exercise using ordinary least squares (OLS) and two-stage least squares (2SLS). A certain inertia exists in the evolution of credit stock that, along other explanatory variables such as GDP growth, real housing prices and the error correction term, is statistically significant. The use of 2SLS did not change the point estimates much, although the statistical significance of GDP growth and real housing prices growth declined. In all cases, the parameter estimated for the ECM fulfilled all the requirements to confirm the existence of a cointegration relationship, although it implied that the speed of adjustment to the long-run equilibrium was relatively slow. The real interest rate had the expected sign but it was not statistically significant, as it occurred with the equity ratio, which, as in the long-run estimation, had a negative effect on credit growth.

We used columns (3) and (4) to check, in a simple way, whether the relationship between credit and its determinants differs during tranquil and stressful times. We did this introducing a dummy with a value of one when there has been a systemic banking crisis in Spain, according to Martín-Aceña's (2013) classification. Therefore, one minus that dummy represents tranquil times. Two observations are necessary. First, even if the coefficients were not statistically different from zero, indicating that we cannot comment on their magnitude, we noted that the average growth rate of real credit is positive in tranquil times and negative during systemic banking crises. Second, when the dummy interacted with our variable of interest (i.e., the equity ratio), we obtained opposite signs, although, again, the parameters were not statistically significant.¹⁷

To check this asymmetry more formally, we estimated the nonlinear model presented previously, which allowed us to identify the various regime states endogenously. The main results appear in Table 3. Columns (1) and (2) present the specification for which only the conditional mean changed with the state. For that specification, most of the coefficients were statistically significant and had the expected sign. Credit stock growth showed a different mean in both states, presented a certain inertia and depended positively on GDP and real housing price growth (the statistical significance of both variables diminished when instrumented), and the error correction mechanism implied a relatively slow adjustment speed. In the linear specification, the real interest rate seemed to play no role. Again, the change in the capital ratio (our variable of interest) was negative but not statistically significant.

For the second specification, in columns (3) and (4), we allowed the capital ratio (jointly with the constant) to have a nonlinear effect on credit growth. Indeed, this seems to be the case. According to our estimates, increasing the capital ratio in advance of a credit boom reduced credit growth, whereas it moderated the expected decline in credit if it occurred before a credit crunch, although the former effect was estimated more imprecisely. In any case, given the difference in the coefficients' absolute values and short-run statistical significance, even taking into account that the capital ratio had a negative effect on credit in the long run (error correction mechanism), this effect seemed more powerful during credit crunches. We interpreted this as a countercyclical capital buffer that reduced the size of the expansion and, more importantly, the magnitude of the subsequent credit decline. Indeed, this is exactly what a macroprudential tool should do: accumulate during expansions for use during recessions. All the other parameters showed the expected signs, although neither the GDP nor the real interest rate were statistically significant in all specifications.

¹⁶ As the capital ratio shows a unit root during this period, it is measured as deviations from the trend. This definition of the variable jointly with the lag mitigates its potential endogeneity.

¹⁷ We performed a robustness check by considering a different stationary transformation of the equity ratio and repeating the nonlinear regressions presented in the main text. In particular, we eliminated the non-stationary part of the equity ratio by calculating its deviation from its trend, and the results were qualitatively the same. The analysis is available upon request.

Table 2
Linear credit equation for the period 1880–2017.

	Simple model		Systemic banking crisis	
	OLS	2SLS	OLS	2SLS
Constant	0.014 (0.013)	-0.009 (0.014)	-	-
Crisis periods	-	-	-0.013 (0.023)	-0.008 (0.018)
Tranquil periods	-	-	0.021 (0.017)	0.025 (0.018)
Credit growth inertia	0.280*** (0.079)	0.334*** (0.089)	0.246*** (0.092)	0.272*** (0.098)
GDP growth	0.666*** (0.247)	0.574 (0.357)	0.634** (0.257)	0.509 (0.362)
Capital ratio change lagged	0.035 (0.399)	-0.235 (0.600)	-	-
Crisis periods* Capital ratio change lagged	-	-	0.250 (0.527)	0.507 (1.098)
Tranquil periods* Capital ratio change lagged	-	-	-0.196 (0.625)	-0.393 (0.716)
Real interest rate lagged	-0.103 (0.159)	-0.068 (0.154)	-0.101 (0.162)	-0.049 (0.314)
Real housing price growth	0.306*** (0.101)	0.266* (0.152)	0.291*** (0.106)	0.200 (0.158)
Residuals of the cointegrating vector	-0.144*** (0.040)	-0.162*** (0.042)	-0.137*** (0.041)	-0.142*** (0.043)
Adjusted R ²	0.286	0.232	0.276	0.235
Standard deviation	0.088	0.091	0.089	0.091
Durbin-Watson	1.868	1.979	1.854	1.944
Observations	127	127	127	127

Notes: The dependent variable is the real credit growth rate. GDP growth, capital ratio change and housing price growth were instrumented using the variables analyzed in the previous section. Coefficients are listed on the first row, and standard deviations appear below. The lags of the variables were chosen according to their relevance and the best fit of the regression. Credit growth inertia measures the one-period lag of credit growth. *, **, ***, statistically significant at 10%, 5% and 1%, respectively.

From a quantitative perspective, the results are quite significant. If banks increased the capital ratio by 1% in advance of a contractionary period, it would reduce the decline rate of credit growth by more than 8%. This elasticity is relatively high. Nevertheless, if risk-weighted assets (RWA) are around half of total assets, which is the case for Spanish banks (Trucharte et al., 2015), increasing the countercyclical capital buffer (CCyB) by 1% is akin to increasing the capital by 2% in terms of RWA, as current capital requirements specify, which is a very large demand for banks. Given the fact that Basel III developed a CCyB requirement with a maximum value at 2.5% of RWA, this is equivalent to 1.25% of total assets. Therefore, the effect of reaching a requirement at the maximum of the CCyB during the expansion should facilitate a reduction in lending decline (when the CCyB is freed during a recession) of around 11%. This is very significant and demonstrates the potential effects of countercyclical tools. According to the dynamic multipliers (see Fig. 5), the maximum effect is reached three years after the increase in the equity ratio is observed, then decreases progressively during subsequent periods until it becomes negative and stabilizes at the estimated long-run elasticity. In any case, it is important to realize that the lengthy period analyzed includes severe lending booms and busts and significant volatility in credit growth rates, which may help to explain the high value of the elasticity in downturns.

The estimates also show that the states' durations were highly asymmetrical. The expansionary periods lasted more than thirty years on average, but the contractionary ones lasted four and a half years. Therefore, in a lending cycle (boom and bust) the overall elasticity of credit to capital would be much closer to other estimates in the literature (Gambacorta and Shin, 2016).

As the variables used in this paper go back more than one hundred years, full homogeneity is not guaranteed. In addition, the Spanish financial system has changed substantially, and relevant regulation has been modified several times. Thus, to check results' robustness, we repeated the estimations of the Markov-switching regime model using various sample periods to exclude some of systemic banking crisis episodes and increase data homogeneity. Given that we combined various data sources to create a unique and continuous one, subdividing the sample allowed us to focus on shorter time periods. Consequently, we did not need to merge every data set to cover the shorter time periods, which let us use more consistent or homogenous data because we used fewer sources. In fact, after the loss of Cuba, a significant amount of capital returned to Spain from abroad and was partly invested in creating new bank houses that became large Spanish banks during the 20th century (e.g., Banco Hispano Americano, Banco Español de Crédito and Banco de Vizcaya, which were created at the turn of the century). The second exercise consisted of eliminating the information

Table 3
Markov-switching model for credit growth for the period 1880-2017.

	Two regimes, different constants		Two regimes, different capital ratio effects	
	OLS	2SLS	OLS	2SLS
Constant	-0.225*** (0.045) 0.024** (0.011)	-0.248*** (0.041) 0.020* (0.012)	-0.024 (0.024) 0.017 (0.011)	-0.060 (0.051) 0.017* (0.010)
Credit growth inertia	0.256*** (0.071)	0.305*** (0.079)	0.376*** (0.078)	0.382*** (0.105)
GDP growth	0.492** (0.240)	0.368 (0.316)	0.545*** (0.201)	0.230 (0.296)
Capital ratio change lagged	-0.051 (0.406)	-0.751 (0.560)	11.994*** (1.566) -0.545 (0.334)	8.833*** (1.737) -0.938 (0.768)
Real interest rate lagged	-0.013 (0.151)	0.040 (0.138)	-0.030 (0.142)	0.058 (0.161)
Real housing price growth	0.308*** (0.098)	0.269** (0.135)	0.217** (0.085)	0.282** (0.110)
Residuals of the cointegrating vector	-0.128*** (0.036)	-0.138*** (0.036)	-0.111*** (0.035)	-0.125** (0.050)
Log(sigma)	-2.626*** (0.075)	-2.618*** (0.071)	-2.711*** (0.069)	-2.582*** (0.103)
Probability of remaining in state 1	0.222	0.225	0.743	0.787
Probability of remaining in state 2	0.971	0.967	0.940	0.973
Expected duration of state 1	1.29	1.29	3.89	4.69
Expected duration of state 2	33.99	30.05	16.786	37.56
Standard deviation	0.089	0.093	0.142	0.092
Log-likelihood	137.521	134.475	147.287	136.036
Durbin-Watson statistic	1.922	2.056	2.347	1.969
Observations	128	128	127	127

Notes: The dependent variable is the real credit growth rate. When a variable is split into two rows, the values of the first row's parameters refers to the low credit growth episode (numbers in italics), whereas the values of the second row pertain to the high credit growth episode. Coefficients are listed on the first row, and standard deviations appear below. The lags of the variables were chosen according to their relevance and the best fit of the regression. Credit growth inertia measures the one lag of the credit growth. *, **, ***, statistically significant at 10%, 5% and 1%, respectively.

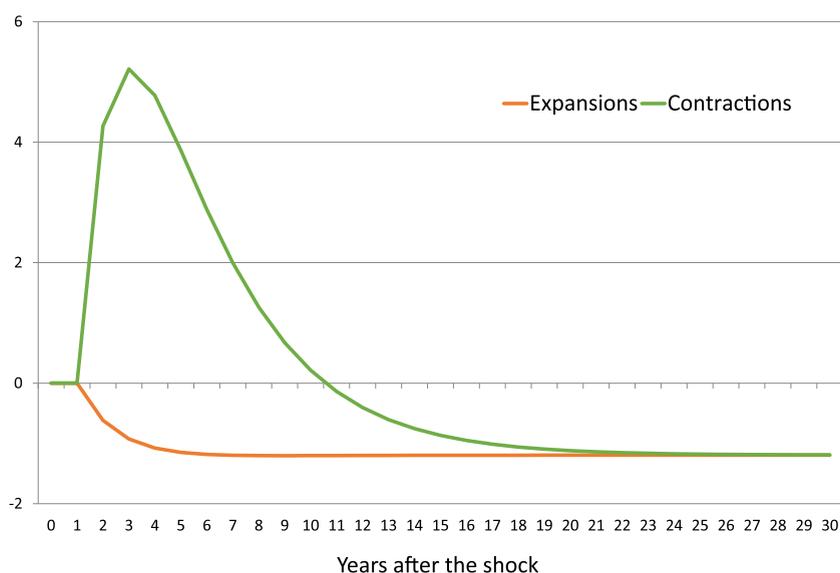


Fig. 5. Credit growth dynamic multipliers of the capital ratio for the period 1880–2017.

Sources: Own calculations.

Table 4
Markov-switching model for credit growth. Robustness analyses for different periods.

Sample period	1898–2017		1880–1998	
	OLS	2SLS	OLS	2SLS
Constant	-0.047 (0.043) 0.011 (0.008)	-0.070 (0.059) 0.013 (0.011)	-0.011 (0.026) 0.023* (0.014)	-0.061 (0.060) 0.017 (0.012)
Credit growth inertia	0.384*** (0.084)	0.389*** (0.099)	0.297*** (0.080)	0.355*** (0.118)
GDP growth	0.678*** (0.221)	0.373 (0.302)	0.629*** (0.229)	0.271 (0.328)
Capital ratio change lagged	11.585*** (1.363) -1.568 (1.049)	8.746*** (1.394) -1.933** (1.005)	13.588*** (2.265) -0.347 (0.351)	8.568*** (1.930) -0.943 (0.820)
Real interest rate lagged	0.121 (0.189)	0.178 (0.179)	-0.037 (0.153)	0.059 (0.165)
Real housing price growth	0.150** (0.068)	0.252*** (0.096)	0.251** (0.101)	0.264** (0.119)
Residuals of the cointegrating vector	-0.104** (0.046)	-0.120* (0.073)	-0.160*** (0.045)	-0.120* (0.065)
Log(σ)	-2.798*** (0.142)	-2.622*** (0.117)	-2.659*** (0.077)	-2.510*** (0.108)
Probability of remaining in state 1	0.763	0.782	0.679	0.782
Probability of remaining in state 2	0.950	0.965	0.922	0.968
Expected duration of state 1	4.22	4.58	3.12	4.58
Expected duration of state 2	20.20	29.23	12.84	31.40
Standard deviation	0.082	0.094	0.154	0.099
Log-likelihood	138.127	123.511	119.521	107.178
Durbin-Watson stat.	2.132	1.957	2.300	1.947
Observations	111	111	109	109

Notes: The dependent variable is the growth rate of real credit. When there a variable is split into two rows, the values of the parameters on the first row refer to the low credit growth episode (numbers in italics), whereas the values on the second row pertain to the high credit growth episode. Coefficients are listed on the first row, and standard deviations appear below. The lags of the variables were chosen according to their relevance and the best fit of the regression. Credit growth inertia measures the one lag of the credit growth. *, **, ***, statistically significant at 10%, 5% and 1%, respectively.

from the new millennium, as during this period, a major structural affected the Spanish economy (Spain joined a monetary union), the banking sector (the Spanish banking system became multinational) and a major GFC occurred.

In Table 4, we show the results of the estimations shown in Table 2, but using these different samples. In other words, we re-estimated the nonlinear model, excluding in the first and second columns the systemic banking crises that occurred in the 19th century. Specifically, we focused on the period 1898–2017, avoiding the two banking crises occurred previously, as their nature might have been quite different from those that occurred during the 20th and 21st centuries. In columns (3) and (4), we excluded the new millennium years (i.e., we focused on the period 1880–1998, disregarding the structural changes occurred in the Spanish economy after 1998 and the GFC).

We obtained similar and comparable results for the two subsamples. The most remarkable difference was in the first subsample, for which we observed that the change in the capital ratio was statistically significant during tranquil periods and had a much higher coefficient. In both cases, the average duration of the expansionary period was shorter, and the coefficient of the error correction mechanism was estimated with much less precision. Other subsamples we considered (pre- and post-Civil War or homogeneous databases) also maintained our main result regarding the capital ratio change's asymmetrical effect on credit growth during turbulent and tranquil periods.

5. Conclusions

One of the lessons of the GFC is that we need better and additional policy tools to smooth the credit cycle and reduce financial crises' costs for taxpayers. In this regard, Spain has had an international advantage, as the countercyclical provisions implemented in its economy could be understood as a macroprudential tool similar to the countercyclical capital buffer.¹⁸ (Jiménez et al., 2017) showed that these provisions effectively smoothed the credit crunch during the last crisis and resulted in slightly less curving during the credit boom. In this paper, we assessed whether these conclusions can be extended to other Spanish financial crises. To do so, we extended the aggregate information from the banking system back to 1880 to include eight systemic banking crises. Providing a nearly 150-year-long series of aggregate capital ratios for the Spanish banking sector a contribution of the paper in itself.

¹⁸ See (Saurina and Trucharte, 2017) for a detailed discussion on the potential complement or substitution between dynamic provisions and countercyclical capital buffers.

The first conclusion from this exercise is that the equity ratio declined substantially from around 40% at the end of the 19th century to almost 4% by 1962 and close to 9% in 2017. This behavior of the Spanish banking system is very similar to that of systems in the United States and the United Kingdom.

When we analyzed the determinants of systemic banking crises, we found that macroeconomic aggregates such as real GDP and external balances were less relevant than macrofinancial variables (real interest rates, credit growth or loan-to-value ratios). By controlling for these variables, we determined that the higher the capital ratio, the lower the likelihood of a systemic banking crisis. In addition, using a nonlinear framework, we showed that the changes in the capital ratio had an asymmetric effect on credit growth. This result is robust to using different subsamples, definitions of credit and transformations of the capital ratio. Our findings complement the results presented by (Gambacorta and Shin, 2016), stressing bank capital's asymmetric role in lending booms and busts, and are also in line with (Saurina and Shin, 2012, Berger and Bouwman, 2013) and Demirgüç-Kunt et al. (2013).

References

- Acharya, V., Gujral, I., Kulkarni, N., Shin, H.S., 2011. Dividends and bank capital in the financial crisis of 2007-2009. NBER Work. Pap. Ser.. Available at <http://www.nber.org/papers/w16896>.
- Acharya, V., Lee, H., Shin, H.S., 2017. Bank capital and dividend externalities. *Rev. Financ. Stud.* 30 (3), 988–1018.
- Adrian, T., Shin, H.S., 2010. "Liquidity and leverage. *J. Financ. Intermed.* 19, 418–437.
- Adrian, T., Shin, H.S., 2014. "Procyclical leverage and value-at-risk. *Rev. Financ. Stud.* 27 (2), 373–403.
- Alessandri, P., Haldane, A.G., 2009. "Banking of the state. Speech. Bank of England.
- Anes, R., Schwartz, G., Tortella, G., 1974. La banca española en la Restauración, datos para una historia económica. Servicio de Estudios. Banco de España.
- Anguren, R., 2011. Credit cycles: evidence based on a non-linear model for developed economies. Documentos de Trabajo del Banco de España No 1113.
- Bank for International Settlements (2017): Range of practices in implementing the countercyclical capital buffer policy, June. Available at <http://www.bis.org/bcbs/publ/d407.htm>
- Basel Committee on Banking Supervision (2004): "International convergence of capital measurement and capital standards. a revised framework," June.
- Basel Committee on Banking Supervision (2006): "International convergence of capital measurement and capital standards. a revised framework. comprehensive version," June.
- Basel Committee on Banking Supervision (2010): "An assessment of the long-term economic impact of stronger capital and liquidity requirements," August.
- Basel Committee on Banking Supervision (2011): "Basel III: A global regulatory framework for more resilient banks and banking systems", June.
- Benink, G., Benston, H., 2005. The Future of Banking Regulation in Developed Countries: Lessons from and for Europe. *Financ. Mark., Inst. Instrum.* 14 (5).
- Berger, A.N., Herrington, R.J., Szego, G.P., 1995. The role of capital in financial institutions. *J. Bank. Finance* 19, 393–430.
- Berger, A.N., Bouwman, C.H.S., 2013. How does capital affect bank performance during financial crises? *J. Financ. Econ.* 109, 146–176.
- Bernanke, B.S., Lown, C.S., 1991. The Credit Crunch. In: *Brookings Papers on Economic Activity*, Economic Studies Program, 22. The Brookings Institution, pp. 205–248.
- Betz, F., Oprică, S., Peltonen, T., Sarlin, P., 2014. Predicting distress in European banks. *J. Bank. Finance* 45, 225–241.
- Borio, C., 2003. "Towards a macroprudential framework for financial supervision and regulation?,". *CESifo Econ. Stud.*, CESifo 49 (2), 181–215.
- Borio, C., Drehmann, M., 2009. Assessing the risk of banking crises. *BIS Q. Rev.* Available at http://www.bis.org/publ/qtrpdf/r_qt0903e.pdf.
- Carreras, A., Tafunell, X., 2005. Estadísticas históricas de España volumen I, II y III, siglos XIX-XX. Fundación BBVA.
- Claessens, S., Kose, M.A., Terrones, M., 2012. How do business and financial cycles interact? *J. Int. Econ.* 87 (1), 178–190.
- Crockett, A., 2000. Marrying the micro-and macro-prudential dimensions of financial stability. Remarks before the 11th International Conference of Banking Supervisors September 20-21.
- Dembiermont, C., Drehmann, M., Muksakunratana, S., 2013. How much does the private sector really borrow? A new database for total credit to the private non-financial sector. *BIS Q. Rev.* March, available at http://www.bis.org/publ/qtrpdf/r_qt1303h.pdf.
- Demirgüç-Kunt, A., Detragiache, E., 1998. The determinants of banking crises: Evidence from industrial and developing countries. *IMF Staff Pap.* 45 (1), 81–109.
- Demirgüç-Kunt, A., Detragiache, E., Merrouche, O., 2013. Bank Capital: Lessons from the Financial Crisis. *J. Money, Credit Bank.* 45, 1147–1164.
- Duprey, T., Klaus, B., 2017. How to predict financial stress? an assessment of Markov switching models. *ECB Work. Pap. Ser.* No 2057. May. Available at <https://www.ecb.europa.eu/pub/pdf/scpwps/ecb.wp2057.en.pdf?1c436759165d9a82eac6fb5d6aeb4402>.
- Elliot, D.J., 2014. "Learning from the history of American macroprudential policy," *Macroprudential Policies. Implementation and Interactions.* In: *Financial Stability Review.* Banque de France, pp. 145–150 No 18April.
- Elliot, D.J., Feldberg, G., Lehnert, A., 2013. The History of Cyclical Macroprudential Policy in the United States. *The History and Economics Discussion Series.* Divisions of Research and Statistics and Monetary Affairs, Federal Reserve Board, Washington D.C. No 29, May 15.
- Gambacorta, L., Mistrelli, P.E., 2004. Does bank capital affect lending behavior? *J. Financ. Intermed.* 13 (4), 436–457.
- Gambacorta, L., Shin, H.S., 2016. Why bank capital matters for monetary policy. *J. Financ. Intermed.* 0, 1–13.
- Financial Stability Forum, 2009. "Report of the Financial stability forum on addressing procyclicality in the financial system," 2 April. Available at http://www.fsb.org/wp-content/uploads/r_0904a.pdf.
- G-20 press release, 2009. Available at https://www.g20.org/Content/DE/StatischeSeiten/Breg/G7G20/Anlagen/G20-erklarung-staerkung-finanzsystem-london-2009-en.pdf?_blob=publicationFile&v=1.
- Hamilton, J.D., 1989. A new approach to the economic analysis on non-stationary time series and the business cycle. *Econometrica* 57 (2), 357–384.
- Hanson, S., Kashyap, A., Stein, J., 2011. A Macroprudential Approach to Financial Regulation. *J. Econ. Perspect.* 25, 3–28 Winter.
- Hardy, D., Pazarbasioglu, C., 1998. Leading Indicators of Banking Crises; Was Asia Different? *IMF Work. Pap.* No 98/91.
- IMF-FSB-BIS, 2016. "Elements of effective macroprudential policies. lessons from international experience," August. Available at <http://www.fsb.org/2016/08/elements-of-effective-macroprudential-policies/>.
- International Monetary Fund, 2014. Staff guidance note on macroprudential policy. *Policy Pap.* November. Available at <http://www.imf.org/en/Publications/Policy-Papers/Issues/2016/12/31/Staff-Guidance-Note-on-Macroprudential-Policy-PP4925>.
- Jiménez, G., Saurina, J., 2006. Credit cycles, credit risk, and prudential regulation. *Int. J. Cent. Bank.* 65–97 June.
- Jiménez, G., Ongena, S., Peydró, J.L., Saurina, J., 2012. Credit supply and monetary policy: identifying the bank balance-sheet channel with loan applications. *Am. Econ. Rev.* 102 (5), 2301–2326.
- Jiménez, G., Ongena, S., Peydró, J.L., Saurina, J., 2014. "Hazardous times for monetary policy: what do twenty-three million bank loans say about the effects of monetary policy on credit risk-taking? *Econometrica* 2 (82), 463–505.
- Jiménez, G., Ongena, S., Peydró, J.L., Saurina, J., 2017. Macroprudential policy, countercyclical bank capital buffers and credit supply: evidence from the Spanish dynamic provisioning experiments. *J. Polit. Econ.* 125 (6).
- Jordà, O., Schularick, M., Taylor, A.M., 2016. The great mortgaging: housing finance, crises and business cycles. *Econ. Policy* 31 (85), 107–152.
- Kaminsky, G., Reinhart, C., 1999. The twin crises: the causes of banking and Balance-of-Payments problems. *Am. Econ. Rev.* 89 (3), 473–500.
- Kaufmann, S., Valderrama, M.T., 2004. Modelling credit aggregates. *Work. Pap. Oesterreichische Natl. Bank.*
- Laina, P., Nyholm, J., Sarlin, P., 2015. Leading indicators of systemic banking crises: Finland in a panel of EU countries. *Working Paper Series.* European Central bank.
- Laeven, L., Valencia, F., 2008. Systemic banking crisis: a new database. *IMF Work. Pap.* (08/224).
- en Martín-Aceña, P., 2013. Crisis bancarias. Nada nuevo bajo el sol. In: Martín-Aceña, P., Martínez-Ruiz, E., Pons, M^a A. (Eds.), *Las crisis financieras en la España Contemporánea, 1850-2012.* Editorial Crítica, p. 2013.

- Mencía, J., Saurina, J., 2016. Macroprudential Policy: Objectives, Instruments and Indicators. Banco de España Occas. Pap. (1601).
- Peek, J., Rosengren, E., 1995. The capital crunch: neither a borrower nor a Lender Be. *J. Money, Credit Bank.* 27 (3), 625–638.
- Rajan, R., 1994. Why bank credit policies fluctuate: a theory and some evidence. *Q. J. Econ.* 109, 399–441.
- Reinhart, C.M., 2010. This time is different chartbook: country histories on debt, default, and financial crisis. NBER Work. Pap. 15815.
- Reinhart, C.M., Rogoff, K.S., 2013. Banking crisis: an equal opportunity menace. *J. Bank. Finance* 37 (11), 4557–4573.
- Saurina, J., Shin, H.S., 2012. Bank capital constraints and lending booms. December, 29. Mimeo.
- Saurina, J., Trucharte, C., 2017. “Las provisiones contracíclicas del Banco de España 2000-2016. English version: The countercyclical provisions of Banco de España 2000-2016. Banco de España.
- Schularic, M., Taylor, A., 2012. Credit booms gone bust: monetary policy, leverage cycles, and financial crises, 1870–2008. *Am. Econ. Rev.* 102 (2), 1029–1061.
- Sheppard, D.K., 1971. The growth and role of UK financial institutions, 1880–1962. Methuen, London.
- Trucharte, C., Pérez, C., Cristófoli, E., Ferrer, A., Lavin, N., 2015. Credit portfolios and risk weighted assets: analysis of European banks. *Financial Stability Journal.* Banco de España.
- Vazquez, F., Federico, P., 2015. Bank funding structures and risk: Evidence from the global financial crisis. *J. Bank. Finance* 61, 1–14 December.