THE IMPACT OF THE INTEREST RATE LEVEL ON BANK PROFITABILITY AND BALANCE SHEET STRUCTURE

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Abstract

We study the sensitivity of bank profits and balance sheet structure to changes in the level of interest rates in Spain during the 2000-2016 period. Autoregressive distributed lag (ARDL) models with controls for the business cycle and interest rate levels are estimated for the time series of key asset and liability categories (credit, financial securities, time deposits, etc.) and profit components (returns on asset and liabilities, provision charges, etc.). We find a non-linear relation between interest rates and net interest income, which is positive at low interest rate levels. This relation is driven by the effect of interest rates on asset and liability returns, and also on credit growth, and on the bank mix of credit, deposits and financial securities. Broader profit measures also present a non-linear relation with interest rates, which can be negative even for low interest rate levels if provisioning charges are high enough.

1 Introduction

The observation of low interest rates and low bank profitability in the years after the financial crisis initiated in 2008 has renewed the interest on the relation between interest rate levels and bank profits. The question of whether low interest rates actually erode bank profits has been well present in the current public debate between financial market participants and monetary authorities.[1] However, there is a relative scarcity of empirical studies of the link between interest rate conditions and bank profits.

Furthermore, existing work has focused on studying the effect of monetary policy on profit measures relative to total assets, that is, bank returns. The size and composition of the balance sheet of banks is however also affected by the level of interest rates, and these balance sheet changes can contribute to better explain bank profit variations. For example, there exists long-standing evidence that contractionary monetary policy is linked to lower aggregate credit, e. g., Bernanke and Blinder (1992), Kayshap et al. (1993), which in turn will contribute to lower interest income. Banks can also alter the composition of their assets (e. g., bank credit relative to debt securities holdings) and liabilities (e. g., term deposits relative to wholesale funding) in response to changes in interest rates, creating a further channel for interest rates to affect bank income. We contribute to the literature with a combined study of the effect of interest rate levels on bank income and balance sheet evolution, examining the growth in the level of these variables in addition to the bank returns that constitute the focus of the existing work on bank profitability and interest rates. This comprehensive approach can help to understand better the transmission of monetary policy to bank profitability and, ultimately, solvency.

We study the effect of the interest rate level on different components of the Net Interest Income (NII henceforth) including both volumes and average interest rates of different categories of assets and liabilities (credit, financial securities holdings, tem deposits, etc.). This breakdown of the NII provides a comprehensive view of the different impact of interest rate changes on different bank stakeholders (loan borrowers, depositors, securities issuers, etc.). Additionally, we evaluate the effect of interest rates on provision charges and other financial income (e.g., income from trading activities), allowing us to form a measure

1 There are multiple examples of association of low interest rates with weak bank profits among financial market participants, especially in the European context, e. g., Dobbs et al. (2013) and S&P Global (2016). The possible negative effect of low interest rates on bank profitability is recognized in public statements of regulators, e.g., Constancio (2016), Fischer (2016) and Rajan (2013), though there is no consensus on its quantitative significance, and low interest rates generate broader macroeconomic concerns beyond their effect on bank profits.

of net interest related income (NIRI henceforth) that adds up the effects of interest rates on different lines of P&L accounts. A priori, the conclusions on the impact of interest rates on NII do not necessarily apply to NIRI due to possible countervailing effects in other income components such as provision charges. In periods of economic stress, the negative weight of provision charges on bank profits can be substantial and the effect of interest rate levels on borrower defaults can have a greater impact on bank profits than fluctuations in interest income.[2]

In this article, we estimate with time series data of the Spanish banking system Autoregressive Distributed Lag (ARDL henceforth) models for the average interest rate and volume growth of each component of NII and for the growth rates of the rest of income sources, e.g., provision charges, in NIRI. The models include controls for the state of the business cycle (GDP Growth, Unemployment, etc.) and the interest rate level (12 month Euribor). Following the work of Borio et al. (2017), we consider a quadratic relation between the bank profitability and the interest rate level, but we introduce a more granular decomposition of profit in volume and return components, we use more general temporal dynamics, and consider systematic specification selection based on economic and statistical criteria. For those models for which the quadratic interest rate term is significant, the response of bank profitability to rate changes will depend on the level of interest rates. Dynamic sensitivity analysis to interest rate levels is carried out to evaluate the effect of interest rate shocks on the different components of bank profitability.

We find that the response of bank profits to interest rate changes is a function of interest rate levels. For periods of high rates such as 2007-2009 in Spain, the estimated models reveal that interest rate increases are associated to a sizeable contraction of credit and rapid growth of provision charges and financing costs. These negative effects on bank profits dominate mitigating factors, such as the rise of interest rates earned on assets and the partial substitution of credit with debt securities, leading to lower bank profitability. On periods of very low interest rates such as 2013-2015 in Spain, interest rate hikes contribute to NII growth, as financing cost increase at a slower pace than interest rates earned on assets and the volume of activity is not too adversely affected. However, the impact of the interest rate increase on provision charges can still lead to a net reduction of bank profits in a low interest rate environment.

The rest of the article is organized as follows. Section 2 discusses the related literature. We present theoretical considerations that motivate the analysis in Section 3 and describe the available dataset in Section 4. The methodological approach is detailed in Section 5 and results are presented in Table 6. Section 7 adds final considerations.

2 Literature Review

As pointed out in Borio et al. (2017), the literature studying the relation of monetary policy and bank profitability is relatively small, but there is still some evidence that serves as a reference for our work. The authors in that article use data for an international panel of banks in the period 1995-2012 and explore how monetary policy, through its effect on interbank rates and the yield curve, impacts several measures of profitability relative to total assets (NII over total assets, ROA, etc.). The dynamic panel data models in Borio et al. (2017) incorporate a single lag of dependent variables and a contemporaneous

2 Interest rate levels and business cycle conditions are found to be significant explanatory factors of credit risk in multiple studies, e. g., Duffie et al. (2007) for the United States, Pesaran et al. (2006) for an international sample, and Jiménez and Saurina (2006) and Jiménez and Mencía (2009) for Spain. The effects of the economic cycle on defaults have also been found to be stronger during recessions, as in the study of Italian defaults of Marcucci and Quagliariello (2009).

quadratic term for interest rates, allowing for a varying effect of interest rate changes as a function of their level. That article finds a positive relation between bank returns and interest rates, which is more significant when these rates are low. In terms of profit components, they find a positive relationship of interest rates with NII and loan loss provisions, and a negative relation with non-interest income (all accounting variables measured relative to total assets).

Related to the work in Borio et al. (2017), the present article studies the possibly non-linear effect of interest rate levels on bank profits, but the contributions of the two articles differ in several respects. As mentioned in the introduction, we consider not only average returns, but also levels of bank profits and variations in the volume and composition of bank balance sheets, decomposing in greater detail than other works the channels through which interest rates impact bank profitability. We generalize the model dynamics in Borio et al. (2017), who use a single lag of the dependent variable as dynamic control, allowing for an ARDL model with up to second order lags of all the explanatory variables. Models are selected after an exhaustive specification process, and we calculate dynamic responses to examine explicitly the effect of interest rate shocks over time.

The work in Borio et al. (2017) and most other references measure the average relation between interest rates and bank profits across a panel of multiple countries. On the contrary, our work is focused on aggregate time series data for Spain. The particular relation and dynamics of bank variables in a given country can differ from the average effect observed internationally, and there is potential value in identifying this country specific information. Spain is an interesting target for study as it is a large European economy, whose banking sector was materially affected by the financial crisis initiated in 2008, as documented for example in Banco de España (2017).

Our work is also related to recent contributions by Claessens et al. (2017) and Altavilla et al. (2017). Claessens et al. (2017) examine a large panel of banks from 2005 to 2013 and focus on Net Interest Margin over total earning assets (NIM) and ROA. They find that a reduction in interest rates harms both NIM and ROA, and that this effect is more pronounced if the initial level is low. Moreover, that article finds that a prolonged period of low interest rates further deteriorates NII and ROA. Altavilla et al. (2017) study profitability in a panel of European banks in the period 2000-2016. They do not find evidence of a significant effect of interest rates on ROA if current and expected macro conditions are included as controls in panel regressions. They find though a significant effect of interest rates on profit components (relative to total assets) such as NII or provisions.[3] The relative contribution of the current work relative to these articles are in line with the commented differences with respect to Borio et al. (2017): more granular decomposition of the impact of interest rates on bank profit components, time series analysis of a specific crisis-hit country and more general dynamics of model variables.

The earlier literature contains additional examples of international studies of bank profitability. Demirgüç-Kunt and Huizinga (1999) examine for the 1988-1995 period the relation of NII and profits (relative to total assets) with bank level characteristics and macro variables, finding a positive effect of interest rates on both profit measures, which is attenuated or even nullified in countries with higher income. Saunders and Schumacher (2000) study a sample of European and American banks in the 1988-1995 period

3 Altavilla et al. (2017) contains complementary analysis of the aggregate time series of the European sample, in line with their panel data results, and stock return analysis, which is less related to the current work.

decomposing NII in several factors and finding that lower interest rate volatility can reduce bank margins. Beckmann (2007) examines a sample of Western European banks in the period 1979-2003 and finds that both market structure and business cycle variables are significant explanatory variables for bank profitability (relative to total assets). Albertazzi and Gambacorta (2009) examine the effect of banking sector structure and macro variables on bank profitability with a country-level panel of Euro Area and anglo-saxon countries. They find that short term market rates affect provisioning ratios whereas long term rates are positively associated with the ratio of NII over total assets. Bolt et al. (2012) study over a long 1979-2007 sample the relation between macro variables and several bank profit measures relative to total assets. The main finding in this article is an asymmetric (stronger) relation of bank profits with cycle measures such as GDP growth during recessions.

The literature also includes some examples of country specific studies such as Lehmann and Manz (2006), which identify significant effects of business cycle and interest rates on the profitability of Swiss banks, and Alessandri and Nelson (2015) for the United Kingdom. This latter article introduces a model of monopolistic competition and analyzes empirically the determinants of different profit ratios of British banks. Alessandri and Nelson (2015) find a long run positive effect of the level and slope of the yield curve on bank interest margins, and that NII variations are not fully hedged by other income components such as trading income.

As the current article also considers other interest related components of bank income (fees, trading income, etc.) in addition to NII, the literature on hedging of banking income is a relevant reference. Gorton and Rosen (1995) find that interest rate swap positions of US banks are exposed to rate increases, but that banks hedge most of those exposures. Purnandam (2007) finds that US banks with higher probability of distress use more intensely derivatives to cover interest rate risk and that banks that do not use derivatives follow more conservative balance sheet policies. Respecting the possible hedge of interest income through diversification of bank activities, the early literature was positive on this hypothesis, as seen for example in the survey by Saunders and Walter (1994), but later work has not found convincing evidence, suggesting that this form of diversification might even increase income risk. Examples of this later work include DeYoung and Roland (2001) and Stiroh (2004) in the US, Lepetit et al. (2008) in Europe and Williams (2016) in Australia.

Finally, our work is also related to the literature examining the transmission and amplification of monetary policy through the financial sector. Changes in interest rates affect the balance sheet strength of borrowers (balance sheet channel) and the volume of lending activity of banks (lending channel), creating a credit channel for monetary policy amplification, as considered in Bernanke and Gertler (1995). Banks with different balance sheet characteristics can be affected differently by monetary shocks creating a bank balance sheet channel. For example, Kashyap and Stein (2000) identify that US banks with higher proportion of securities are less affected by monetary contractions, a result in line with the earlier finding in Kashyap et al. (1993) that credit is substituted with commercial paper after a negative monetary shock.[4] The working of these different channels of monetary policy transmission impacts the profitability and solvency of banks, potentially generating a bank

4 The transmission channels considered affect both the demand and supply of credit, generating a challenging identification problem for those studies trying to separate exactly the supply effects. Jiménez et al. (2014) contribute to this literature with use of a granular dataset to separate demand, supply volume and risk composition factors, identifying that low capital banks take more risks with lower short term rates.

balance sheet channel that amplifies further policy shocks, e. g., Gambacorta and Mistrulli (2004) and Van den Heuvel (2007). The current work provides evidence on the aggregate effects of the fluctuation of monetary conditions on key Spanish macrofinancial magnitudes related to the aforementioned transmission channels. This analysis at the system level can help macroprudential policy and serve as a guide for work with more granular datasets.

3 Theoretical Considerations

The empirical exercise in this article is motivated by some key features of workhorse theoretical models, which suggest that an exclusive focus on average returns offers an incomplete evaluation of the effects of monetary policy on bank profitability. The indeterminacy of theoretical results on the net effect of monetary policy on bank profits, due to the presence of multiple profit components through which this policy can operate, also highlights the importance of quantifying empirically the effects on each of these components.

Theoretical bank competition models, as the well-known Monti-Klein model,[5] show that the amounts of loan demand and deposit supply in the banking sector are not independent of the average returns earned and paid on assets and liabilities. An increase in reference interest rate levels might be associated with higher returns, but it will also typically be related to lower bank balances, leading to welfare losses for bank borrowers and potentially reducing the level of bank profits. A small investor solving a portfolio allocation problem can safely assume that she can scale up and down her position in a given security without affecting the expected return on that security. The expected return of this small investor on a given security is thus independent of the volume invested. This assumption is not expected to hold when studying the banking sector as a whole, where the feasible volume and composition of bank activity at a given time depends reasonably on prevailing interest rates and average bank returns.

An increase in the reference rate is expected to contract the volume of total assets, and in particular credit, and also affect the relative volumes of the different categories of assets and liabilities: substitution of loans with debt securities on the assets side, or sight deposits with term deposits on the liability side, etc. The amount of non performing exposures is also expected to increase with a rise in the reference rate, with a detrimental effect on bank income. Financial and market structure characteristics of the banking sector (varying degree market power in different segments, different mix of variable and fixed rate contracts on asset and liability sides of the balance sheet, differing demand and supply conditions for existing and new loans and deposits, maturity structure, etc.) lead us too to presume that the elasticity of the bank rates to the policy rate varies across assets and liabilities categories. For example, market power in retail segments could make loan rates more responsive to raises in the policy rate than deposit rates.

These reasonable *a priori* conjectures on the effects of changes in reference interest rates highlight the convenience of modelling NII in terms of the volumes and rates of the different categories of assets and liabilities that comprise it, rather than directly at the aggregate level. The question of the final effect that a change in the reference rate may have in NII is complex to answer based exclusively on theoretical arguments, since different components may respond in opposite directions or with different velocity, and then we advocate to assess this question empirically.

5 The Monti-Klein model is originally derived in a monopoly framework, Klein (1971) and Monti (1972), but it has been since extended to an oligopoly setting. See Freixas and Rochet (2008) for more extensive treatment of this matter.

Regarding the broader profit measure NIRI, which includes commissions and fees and provision charges, the magnitude and direction of the response to a change in the policy rate are also unclear on purely theoretical grounds. For example, existing literature clearly shows that and increase in the reference rate produces, ceteris paribus, an increase in default rates and therefore an increase in provision charges, although the intensity and velocity of the impact has to be measured empirically. The response of NIRI, which is the sum of NII, Other Financial and Banking Income (OFBI) and provision charges, to an increase in the reference rate will depend on whether a potential increase in NII and OFBI can compensate the reduction of profits via the increase in provision charges. This possibility depends, in turn, on multiple factors (asset and liability positions, maturity structure, history of interest rates, etc.), and therefore empirical analysis can be a priori expected to find a differing impact of interest rates on bank profits for different periods and financial conditions. In particular, we would expect that loan demand and default rates respond more negatively to rate increases in periods of high interest rate levels.

4 Dataset

We use aggregate system-level information for all deposit institutions in Spain for the NIRI variables. This dataset originates from the regulatory reports of these institutions to Banco de España. In order to account only for the exposures in Spain and exclude exposures abroad, the system-level series are built through the aggregation of individual level statements instead of consolidated statements.[6] Time series are measured at quarterly frequency and cover 16 years from 2000Q1 to 2015Q4, which allow us to study a full economic cycle including both expansive and recessive years. Mergers and acquisitions can generate unbalanced individual bank profit series, but this issue does not affect the current study as we use aggregate data and focus on the systemic evolution of bank profitability.

For the components of NII, six volume series are obtained from balance sheet reports and six average interest rates series are constructed from the ratio of P&L income or expense items and balance sheet stocks for the corresponding asset and liability categories. For example, the series of average interest rate on credit is obtained by dividing the series of interest income from credit exposures by the volume of interest producing credit. Asset and liability categories include credit, debt securities holdings, rest of assets (derived as difference of total interest producing assets and loan credit and debt securities, reflecting mostly cash and interbank positions),[7] sight deposits, term deposits and rest of liabilities (derived as difference of total interest bearing liabilities and sight and term deposits, reflecting mostly wholesale funding). The variable NII can be recovered from the formula

$$NII = \sum_a Vol_a \times Rate_a - \sum_l Vol_l \times Rate_l \quad [1]$$

where Vol denotes balance sheet stock, Rate denotes average interest rate, a indexes the three categories of assets and l indexes the three categories of liabilities.

6 This approach is reasonable given that we use as potential profit determinants macroeconomic variables specific to Spain (GDP Growth, Unemployment, etc.) that would not be connected with business abroad, and that foreign exposures are heavily concentrated in the largest entities. Therefore, consolidated level data would be less representative of bank profitability in Spain.

7 All asset side categories used (credit, debt securities and rest of assets) refer to performing or interest producing assets (non performing positions are not computed). This definition of volume variables facilitates their interpretation as a proxy of value generation in the banking sector, which would be less adequate if non-performing assets where included, as many non-performing exposures plausibly reduce social surplus.

Provision charges capture the flow of new provisions for asset value deterioration recognized in the P&L in each quarter, rather than the stock of provisions in the balance sheet, and it is obtained directly from the regulatory database. Other financial and banking income (OFBI henceforth) is derived as the difference between the gross income series (income excluding provisioning charges, operating expenses and taxes) and the NII series, which are both in the regulatory report database. OFBI captures mostly earnings coming from banking fees and financial operations (e.g. securities trading), whose profitability can be affected by interest rate levels. We combine the different income items (NII, OFBI and Provision Charges) into the single measure of net profitability NIRI, i. e.,

$$\text{NIRI} = \text{NII} + \text{OFBI} - \text{Provision Charges} \quad [2]$$

For the interbank rate, we use the 12-month Euribor series *(Euribor),* obtained from the Statistical Bulletin of Banco de España. We use the 12-month maturity instead of the 3-month rate, which is often considered in the literature, because most credit products in Spain use the 12-month rate as reference. Additionally, correlation between 3-month and 12-month Euribor rates is very high (99.01% over the sample period), so this choice is not critical. We consider as measure of the slope of the interest rate curve the difference *(Slope)* of the 10-year Spanish bond rate, also obtained from the Statistical Bulletin, minus the 12-month Euribor. As controls for the state of the business cycle, we consider house price growth, unemployment, and real GDP growth data obtained from the Spanish Ministry of Public Works and the National Statistical Institute.[8] These macroeconomic variables are measured quarterly, but growth rates for house prices and real GDP are calculated in inter-annual terms.

Table 1 presents the main descriptive statistics of the interest rate and macro variables, the components of NII and the components of NIRI (NII, OFBI, and Provision Charges). This table shows the wide range of values of the macro variables over the sample period (e. g., inter-annual real GDP growth varies from –4.2% in 2009Q2 to 6.5% in 2000Q1). Regarding bank income variables, NII growth is more stable along the cycle than the growth of other components of NIRI, presenting a lower standard deviation (12.5%) than OFBI (19.6%) and provisioning charges (108.8%). The high volatility of provision charges growth is due to the big differences in provision charges between periods of economic expansion and recession. Concerning the components of NII, sight deposit growth is much more stable (6.0% Std. Dev.) along the cycle than term deposit growth (16.8% Std. Dev.) or rest of liabilities growth (11.0%). The average interest rate paid for sight deposits (0.7%) is, as expected, lower than for term deposits (2.9%) and other liabilities (2.6%). On the asset side, the interest rate earned on loan credit (4.1%) is approximately 1% higher than that paid for term deposits (2.9%), and it is the highest rate of all the three asset components of NII, given average rates of 3.8% and 2.7% for debt securities and rest of assets. Similarly, the term deposit rate is the highest average rate paid on liabilities.

Chart 1 presents graphically the evolution of the main components of NIRI. NII and OFBI follow a mostly positive growth trend in the period 2000-2008. OFBI peaks in year 2007 whereas NII peaks in year 2009, which is a year of maximum interest rates in which the effects of recession had not still materialized fully in the bank balance sheets. In 2010, we see how NII and OFBI are clearly below their peak values, and they remain relatively stable

8 The links to the sources of macroeconomic and interest rate variables are the following: Statistical Bulletin of Banco de España (www.bde.es/bde/en/areas/estadis/), National Institute of Statistics (www.ine.es) and Spanish Ministry of Public Works (www.fomento.gob.es).

DESCRIPTIVE STATISTICS OF BANK INCOME AND MACROECONOMIC VARIABLES

TABLE 1

	Mean	Std. Dev.	Min.	Max.
Macro Variables				
House Index Growth (%)	4.21	9.18	-10.05	18.45
Unemployment (%)	15.74	6.32	7.93	26.94
Slope (%)	1.84	1.48	-0.80	5.48
Euribor (%)	2.45	1.55	0.09	5.37
Euribor Square (%)	8.38	8.39	0.01	28.80
Real GDP Growth (%)	1.69	2.70	-4.22	6.50
Interest Related Income Variables				
Growth (%)				
Net Interest Income	3.10	12.54	-23.99	27.21
Other Financial and Banking Income	8.58	19.59	-36.07	65.55
Provision Charge	43.22	108.84	-75.20	361.62
Net Interest Income Variables				
Growth (%)				
Credit	7.11	12.11	-15.10	29.05
Debt Securities	7.99	12.76	-17.91	41.48
Rest of Assets	4.31	10.46	-17.97	24.94
Sight Deposits	7.23	6.01	-4.18	17.37
Term Deposits	9.90	16.82	-16.86	50.13
Rest of Liabilities	4.87	10.99	-23.98	21.46
Interest rate (%)				
Loan Credit	4.15	1.03	2.43	6.24
Debt Securities	3.78	0.92	2.19	5.54
Rest of Assets	2.67	1.37	0.82	5.41
Sight Deposits	0.67	0.33	0.16	1.49
Term Deposits	2.95	0.63	1.77	4.23
Rest of Liabilities	2.59	1.25	0.70	4.81

NOTES: Data series are available at quarterly frequency and cover the sample period 2000 Q1 – 2015 Q4. Euribor is the 12-month Euribor rate. Slope is the difference “10-year Spanish bond rate” – “12-month Euribor”. Growth variables represent inter-annual growth. Interest rates for the Net Interest Income variables represent average values over the quarter. Statistics for Net Interest Income are presented for completion, although this element is not modelled directly and therefore this time series is not used in the empirical exercise.

in later years, with a mild decline of NII. Provision charges represent the most volatile component of NIRI, presenting a quick rise during the double-dip recession.[9]

5 Methodology

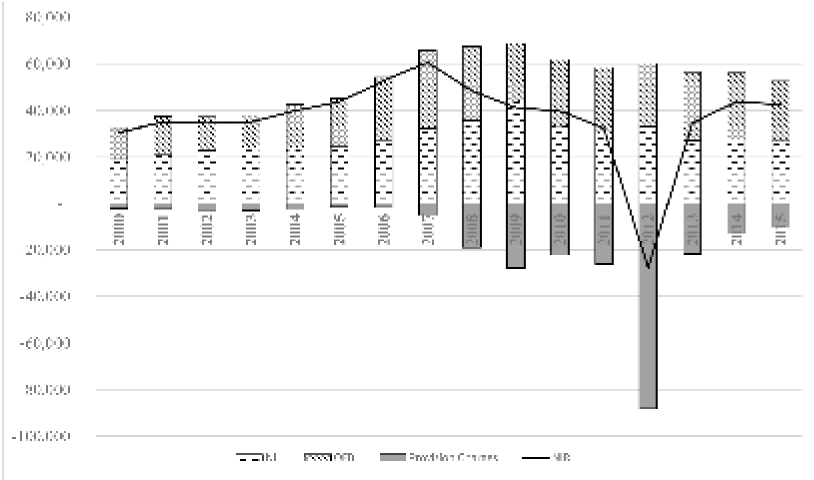
5.1 ESTIMATION FRAMEWORK

As mentioned in the introduction, we adopt an ARDL model for each of the components of NIRI, which add up to a total of 14: one model for OFBI and other for provision Charges, which are modelled directly, and one model for each of the 12 components of NII, which is not modelled directly but through the aggregation defined in formula (1). Formally, we estimate through OLS the following ARDL equation for the variable of interest y_t:

$$y_t = \rho_0 + \sum_{j=1}^{J} \rho_j \cdot y_{t-j} + \sum_{m=1}^{M} \sum_{s=0}^{S(m)} \beta_{m,s} \cdot x(m)_{t-s} + \varepsilon_t \qquad [3]$$

9 Spanish Banks were required to charge additional provisions in 2012 following Royal Decrees 2/2012 and 18/2012, and coinciding with the second dip of the recession. The large increase in absolute value of provision charges in 2012 is related to these factors. Results are found to be robust to the exclusion or treatment via dummies of these quarters with high provisioning levels. It must also be taken into account that the highest growth in provision charges takes place in 2006-2007, as credit quality began deteriorating, rather than in 2012.

CHART 1



NOTE: This figure depicts the evolution of year-end NIRI (Net Interest Related Income) and its P&L components: NII (Net Interest Income), OFBI (Other Financial and Banking Income) and Provision Charges.

where ε_t is the model error, parameters ($\rho_0, \rho_1, \ldots, \rho_J$) determine the autoregressive dynamics of y_t up to lag order J, and parameters ($\beta_{1,1}, \ldots, \beta_{1,S(1)}, \ldots, \beta_{m,s}, \ldots, \beta_{M,1}, \ldots, \beta_{M,S(M)}$) determine the effect of explanatory variables $x(m)_{t-S}$ for $m = 1, \ldots, M$ on y_t, with the lag order of each variable S(m) being potentially different. This model can be recast in an error correction model (ECM) form:

$$\Delta y_t = \rho_0 + \alpha \cdot \left(y_{t-1} - \theta \cdot x_{t-1}\right) + \sum_{j=1}^{J-1} \delta_{yj} \cdot \Delta y_{tj} + \sum_{m=1}^{M} \sum_{s=0}^{S(m)-1} \delta_{m,s} \cdot \Delta x(m)_{t-s} + \omega_t \quad [4]$$

where the long term relationship between dependent variable and the set of all explanatory variables ($x_t \equiv x_t(1), \ldots, x_t(M)$) is governed by θ, correction of short term deviations is governed by α, and short run dynamics in y_t are given by parameters $\left\{\delta_{yj}\right\}_{j=1}^{J-1}$ and, for each variable $x_m, \left\{\delta_{m,s}\right\}_{s=0}^{S(m)-1}$. Pesaran and Shin (1999) establish that OLS estimates of models of this form (and more general ARDL specifications with trends) are consistent and allow using normal asymptotic theory not only when variables (y_t, x_t) are I(0), that is, integrated of order zero and thus stationary, but also when they are pure I(1) process and there is a single long-run cointegrating relation between dependent and explanatory variables.[10]

The estimation framework and tests in Pesaran and Shin (1999) and Pesaran, Shin and Smith (2001) admits either pure I(0), pure I(1) or a combination of both types of variables in the set (y_t, x_t). However, the framework would cease to be applicable if the variables are

10 Furthermore, Pesaran, Shin and Smith (2001) derive the non-standard distribution for F-tests and t-tests of the hypothesis that there is a cointegrating relation between the variables (y_t, x_t). The F-test tests the null hypothesis $H_0^F \equiv \alpha = 0 \cap \alpha \cdot \theta = 0$ against the alternative that $H_A^F \equiv \alpha \neq 0 \cup \alpha \cdot \theta \neq 0$ whereas the t-test exclusively tests the hypothesis $H_0^t \equiv \alpha = 0$ against the alternative $H_A^t \equiv \alpha \neq 0$. The result of rejection of both tests is interpreted as evidence in favor of existence of a long run cointegrating relationship between variables (y_t, x_t). Both the F-test and the t-test have associated upper and lower bounds for the test statistic, with an indetermination region. We utilize the implementation of the tests in Kripfganz and Schneider (2016) to carry out these bound tests in Pesaran, Shin and Smith (2001).

integrated of higher order. We perform standard Dickey Fuller tests (with one lag of the variable examined in the supporting regression) to examine evidence of unit root behavior of dependent and explanatory variables. For robustness, we also perform the KPSS test for level-stationarity in Kwiatkowski et al. (1992).[11] These tests are also applied to the differences of the original series to test for I(2) behavior. The time horizon of the sample (16 years) might limit the power of unit root tests against a true alternative hypothesis of stationarity, leading to the wrong conclusion of existence of unit root behavior when data is indeed stationary. For example, Cochrane (1991) points out the limited power of unit root tests. Actually, the explosive dynamics implied by I(1) processes for interest rate and growth variables are troublesome from a theoretical perspective, as these variables would need to diverge infinitely. In any case, the wrong conclusion of the presence of an I(1) process when data are I(0) would not affect the validity of the ARDL framework employed, as presented in Pesaran and Shin (1999).

The validity of the estimation exercise also requires the absence of serial correlation in the model residuals, and we thus examine this condition with the autocorrelation test of Arellano-Bond (1991), AB test henceforth. Although this test is originally developed in panel data context, it can be applied to time series data and it is equivalent to Cumby and Huizinga (1992) time series test when checking the existence of autocorrelation at a particular lag, as it is done in the current article.[12] These tests are valid under more general assumptions than earlier tests for autocorrelation in times series context, in particular, they do not require normality and conditional heterocedasticity is allowed.

5.2 MODEL SPECIFICATION

We select the specification for the ARDL equation for each component of NIRI based on an exhaustive search over a wide set of potential specifications. We firstly screen specifications based on statistical and economic requirements, and then choose the final specification to be implemented based on an information criterion. The procedure is related to the approach of the European Central Bank (ECB) top-down stress test framework, Henry and Kok (2013), with the main departure being that we select a single optimal model, whereas the ECB approach implements a Bayesian average of several admissible specifications.

In order to have a certain degree of precision in inferences, we require for each explanatory variable that the set of all its lags included in the model is jointly significant.[13] For example, a model with GDP growth and its first two lags as explanatory variables will be admissible if these three variables are jointly significant. Additionally, we evaluate the coefficients on lagged dependent variables to ensure the specification is stationary, and require that at least the first order coefficient is statistically significant.[14] We also require that admissible specifications do not reject the null of absence of residual autocorrelation based on the AB test, and that the specifications conform to a suitable ARDL structure. In particular, if lag

11 In Dickey-Fuller tests, the null hypothesis is unit root behavior and the alternative is generation through a stationary AR(1) process. On the contrary, the KPSS test has stationarity as null hypothesis. For KPSS, we choose maximum lag order with the Schwert criterion and empirical autocorrelation estimated with the Barlett kernel.

12 We use Roodman (2006) implementation in Stata of AB tests. Baum and Schaffer (2013) implement in Stata a general autocorrelation test command for times series data that can be used to verify Arellano-Bond test results are equivalent to those of Cumby-Huizinga (1992).

13 We use as benchmark a joint significance level of 10% and find non-empty lists of admissible specifications for all equations except term deposit growth. For this variable, we relax the significance level requirement to 15% to explore whether it actually lacks any significant relation with macro variables, or just some macro control is marginally insignificant. The latter case applies and we find specifications with generally significant macro controls also for term deposit growth.

14 For an AR(2) specification, we would verify stationarity by checking that: (i) $\rho_1 + \rho_2 \leq 1$, (ii) $\rho_2 - \rho_1 \leq 1$ and (iii) $|\rho_2| < 1$. The requirement of significance in at least the first order coefficient is weak. Given the persistence in the data, lagged values of the dependent variable are easily found to be significant.

S(m) of a variable x(m) is included, all lags in $\{t, t-1, \ldots, t-S(m)\}$ must be incorporated in the model. The F-test and the t-test in Pesaran, Shin and Smith (2001) must reject the absence of a long term relation.

In addition to statistical requirements, we impose sign restrictions on the coefficients of some explanatory variables based on economic considerations. For example, we require that the long run effect of GDP growth (determined by the sum of the coefficients on contemporaneous and lagged values of the variable) on credit growth is positive. As another example, 12 month Euribor is required to have a positive relation with average interest rates on bank balance sheet items (from credit loans to rest of liabilities) for all levels of this interbank rate. Annex A details the full-set of restrictions imposed, which are fairly general.[15]

We choose the final specification to be implemented as that with the lowest value of the Bayes-Schwartz Information (BIC) criterion among those specifications that are no screened out by the statistical and economic restrictions described in the above paragraphs. We aim to pick with this criterion a parsimonious specification among those that satisfy admissibility criteria.

Even with the relatively parsimonious set of explanatory variables used for this study, the number of possible specifications rapidly grows into hundreds of thousands of variants. In general, for a set of N explanatory variables, the number of potential specifications is $2^N - 1$. For example, this would yield approximately 1 million possible specifications given a set of 20 potential explanatory variables. This high number of potential specifications makes the search for an optimal specification very costly in terms of computing time. We impose several simplifying assumptions to make the search process feasible. Firstly, we limit the maximum lags of any explanatory variable to two (J = 2 and S(m) = 2), leaving a maximum of 18 exogenous explanatory variables[16] and a maximum of two lags of the dependent variable. Preliminary trials reveal that relatively few explanatory variables suffice to obtain a high fit to the data. We limit to 9 the maximum number of exogenous explanatory variables in a given model. These assumptions leave $2 \times \left(\sum_{r=1}^{9} C(18,r) - 1 \right) \approx 310{,}000$ possible specifications for each of the 14 models, making the specification selection process feasible.

6 Estimation Results

We present first in subsections 6.1 to 6.3 the estimates resulting from the application of the methodology presented in Section 5 to the dataset described in Section 4. As NII is modelled through the aggregation of volumes and average rates ($NII = \sum_a Vol_a \times Rate_a - \sum_l Vol_l \times Rate_l$ for asset and liability categories a and l), sections 6.1 and 6.2 present the models for these components rather than a single model for NII. Section 6.3 presents the models for OFBI and provision charges. In order to better gauge the dynamic response of bank income to changes in the levels of interest rate variables, we compute and present in subsection 6.4 the effect on bank income of 100 basis points shock to 12 month Euribor in different time periods of the sample horizon.

6.1 ESTIMATED MODELS FOR BALANCE SHEET GROWTH

We display the estimated ARDL models for the growth of the different balance sheet elements of banks in Table 2. We observe that house price growth is the most common control for the state of the business cycle, being present in all models except those for the debt securities and other liabilities, where the relevant macro control is real GDP. On

15 The sign restrictions ensure that we do not use specifications with potential omitted-variable bias inducing inconsistent signs with economic theory and well established previous evidence.

16 For six original exogenous variables (House Price Growth, Real GDP growth, Unemployment, Slope, Euribor and Euribor Squared), each of them entering contemporaneously and with lags t-1 and t-2.

ESTIMATED MODELS FOR NII COMPONENTS: BALANCE SHEET GROWTH

TABLE 2

	Credit	Debt Securities	Other Assets	Sight Deposits	Term Deposits	Other Liabilities
ARDL Coefficients						
Lag(1)	0.9030 (0.0442)***	1.0708 (0.1004)***	0.5282 (0.0892)***	1.0147 (0.1126)***	0.7935 (0.1111)***	1.1545 (0.1050)***
Lag(2)		-0.4188 (0.1037)***		-0.2245 (0.1289)*		-0.4353 (0.1057)***
House Price Growth	0.1431 (0.0298)***		-2.4256 (0.7814)***	0.1135 (0.0454)**	-1.4320 (0.6389)**	
House Price Growth (t-1)			2.5204 (0.8213)***		1.3964 (0.5933)**	
Unemp.					-0.0118 (0.0055)**	-0.0044 (0.0019)**
Slope	-0.0059 (0.0029)**				-0.0102 (0.0143)	
Slope (t – 1)					0.0038 (0.0176)	
Slope (t – 2)					0.0382 (0.0129)***	
Euribor		0.0146 (0.0054)***	0.0427 (0.0248)*	-0.0071 (0.0019)***	-0.0318 (0.0211)	
Euribor Sq.	0.0001 (0.0007)		-0.0076 (0.0041)		0.0053 (0.0030)*	-0.0010 (0.0005)**
Euribor Sq. (t – 1)	-0.0015 (0.0005)***					
Real GDP Growth		-1.1525 (0.3713)***				-0.1662 (0.4929)
Real GDP Growth (t – 1)						-0.9461 (0.8631)
Real GDP Growth (t – 2)						1.2368 (0.5672)**
Constant	0.0216 (0.0098)**	0.0103 (0.0153)	-0.0279 (0.0232)	0.0284 (0.0070)***	0.1790 (0.1082)*	0.0887 (0.0367)**
ARDL metrics						
R-squared	0.98	0.76	0.72	0.87	0.95	0.94
BIC	-331.8	-154.2	-164.6	-287.7	-198.5	-252.2
AB test p-value	0.38	0.20	0.22	0.70	0.75	0.72
ECM coefficients						
Correction term	-0.0961 (0.0353)***	-0.3432 (0.0782)***	-0.4942 (0.0962)***	-0.2420 (0.0754)***	-0.1955 (0.0721)***	-0.2714 (0.0683)***
LT House Index	1.4403 (0.4114)***		0.1774 (0.2237)	0.5281 (0.1436)***	-0.5055 (0.6939)	
LT Unemp.					-0.0711 (0.0231)***	-0.0152 (0.0038)***
LT Slope	-0.0678 (0.0280)**				0.1710 (0.0756)**	
LT Euribor		0.0424 (0.0199)**	0.0972 (0.0470)**	-0.0303 (0.0107)***	-0.2222 (0.1618)	
LT Euribor Sq.	-0.0155 (0.0076)**		-0.0167 (0.0079)**		0.0326 (0.0263)	-0.0036 (0.0025)
LT Real GDP		-3.2543 (1.1116)***				0.6276 (1.0789)
ECM metrics						
ECM R-square	0.4	0.3	0.4	0.3	0.4	0.4
Bounds F-test estat.	10.9	7.2	7.3	4.9	3.9	5.8
Bounds F-test 10% LB	2.7	3.2	2.7	3.2	2.3	2.7
Bounds F-test 10% UB	3.8	4.1	3.8	4.1	3.3	3.8
Bounds t-test estat.	-2.7	-4.4	-5.1	-3.2	-2.7	-4.0
Bounds t-test 10% LB	-3.5	-3.2	-3.5	-3.2	-3.9	-3.5
Bounds t-test 10% UB	-2.6	-2.6	-2.6	-2.6	-2.6	-2.6

NOTES: For each of the asset and liability items indicated in the cols., the panel *ARDL Coefficients* reports OLS estimates for ARDL models in levels as in equation (3) with the year-on-year growth of the stock of the corresponding balance sheet item as dependent variable. These balance sheet items are the volume components of NII in equation (1) ($NII = \sum_a Vol_a \times Rate_a - \sum_l Vol_l \times Rate_l$). For a given explanatory variable, the coefficient is provided with standard error (in parentheses) below it. Reported standard errors are robust to heterocedasticity of arbitrary form. Coefficients for the first and second lag of the dependent variable are provided in rows Lag(1) and Lag(2). When an explanatory variable is not included in any model, it is removed from the table for clarity. Panel *ARDL Metrics* includes the p-value for a first order autocorrelation test of the form given in Arellano-Bond (1991) applied to the residuals of the ARDL models. The panel ECM coefficients reports OLS estimates of correction term α and long term (LT) parameters θ for the Error Correction Model reformulation (ECM) of ARDL models as in equation (4). In the panel *ECM metrics, Bounds F-test estat* and *Bounds t-test estat* provide statistic values for the test for the presence of an integration relation as in Pesaran, Shin and Smith (2001). The null hypothesis is absence of an integration relation for both the F-test and the t-test. For the F-test, the null is (i) accepted if *Bounds F-test estat* is below the lower bound *Bounds F-test 10% LB* and (ii) rejected if *Bounds F-test estat* is above upper bound *Bounds F-test 10% UB.* For the t-test, the null is accepted if *Bounds t-test estat* is above the upper bound *Bounds t-test 10% UB* and (ii) rejected if *Bounds t-test estat* is below lower bound *Bounds l-test estat 10% LB.* A statistic value between the two bounds is inconclusive for any of the tests. *, **, *** denote significance at the 10%, 5% and 1%.

the other hand, unemployment is present in the models for term deposits and other liabilities, while the interest rate slope measure is present in credit and term deposits. First order autoregressive dynamics are found sufficient to fit the data in credit, other assets and term deposits, with AR(2) specifications chosen for the models of debt securities, sight deposits and other liabilities.

Focusing on the controls for the levels of market interest rates, the 12 month Euribor enters exclusively through non-linear terms in the models for credit and other liabilities, it enters only through linear terms in the model for debt securities and sight deposits, and it presents both linear and non-linear effects in the model for the rest of assets and term deposits. The net effect of the 12 month Euribor on credit is negative and non-linear, with interest rate hikes reducing credit more at higher interest rate levels.[17] On the contrary, 12 month Euribor receives a positive and linear coefficient in the model for debt securities holdings. This means that banks tend to invest more in marketable debt securities relative to bank credit products when the level of interest rates is higher, altering the product mix of their asset side. On the liability side, we observe that sight deposits present a negative and linear (thus independent of interest rate levels) relation with the Euribor 12 month, whereas other liabilities present a negative but non-linear relation with the interbank rate. The case of term deposits is mixed: the linear coefficient is negative but the non-linear one is positive, which means a positive relation with the interbank rate at higher levels. At lower levels, the linear and non-linear effects cancel each other and the effect is expected to be non-significant. Thus, given these estimates, banks will substitute sight deposits and other liabilities with term deposits as interest rates increase if the starting interest rate level is sufficiently high. These patterns can be interpreted as consistent with traditional predictions of competition models of the banking sector, e. g., Monti-Klein, with banks reducing credit, increasing deposit funding and adopting a longer position in financial markets as result of an interest rate increase.

The reparametrization of the ARDL models in error correction form also offers relevant information. The error correction term α measures the speed of adjustment of growth rates to their long term value, with a lower absolute value of this coefficient indicating a slower adjustment given a deviation from this long term benchmark. We observe that sight deposits and term deposits present a slower adjustment than rest of liabilities, whereas credit is the asset category with the slowest speed of adjustment. The volumes related to the traditional activities of deposit taking and granting of bank credit adjust more slowly than the volumes associated to investment and funding in wholesale financial markets. Regarding the long term effects θ of the 12 month Euribor in the different volume growth series, these are significant for credit, debt securities, other assets and sight deposits, but not on term deposits and rest of assets. Additionally, we find that both house price growth and the slope measure have a long term effect on credit growth. Finally, the Pesaran-Shin-Smith bound tests in Table 2 are supportive of the presence of a long run integration relation between dependent and explanatory variables for all balance sheet categories, as required in the admissibility criteria.

6.2 ESTIMATED MODELS FOR AVERAGE INTEREST RATES

The estimation results for the ARDL models of average bank interest rates are presented in Table 3. For this set of models, the most relevant control for the state of the business

17 The negative effect of interbank rates on credit growth is imposed as a requisite on the set of admissible specifications, but we do not impose linear or non-linear specifications, with the final estimated model produced as result of the model selection methodology in section 5.2. The signs of interbank rate coefficients on the rest of models for balance sheet growth are not constrained. See Annex A.

ESTIMATED MODELS FOR NII COMPONENTS: AVERAGE BANK INTEREST RATES

TABLE 3

	Credit	Debt Securities	Other Assets	Sight Deposits	Term Deposits	Other Liabilities
ARDL Coefficients						
Lag(1)	0.4362 (0.1448)***	0.3860 (0.1297)***	0.6946 (0.0778)***	0.6392 (0.1097)***	0.5583 (0.1058)***	0.6264 (0.0615)***
Lag(2)	0.1635 (0.0725)**					
House Price Growth				-0.0024 (0.0008)***		
Unemp.	-0.0006 (0.0004)	-0.0009 (0.0003)***	-0.0010 (0.0004)**			-0.0009 (0.0004)**
Unemp. (t – 1)	0.0004 (0.0005)	0.0011 (0.0002)***	0.0000 (0.0006)			0.0009 (0.0004)***
Unemp. (t – 2)	0.0004 (0.0003)		0.0010 (0.0003)***			
Slope		0.0010 (0.0003)***				
Euribor	0.0004 (0.0007)	0.0026 (0.0006)***	0.0028 (0.0006)***	0.0009 (0.0002)***		0.0036 (0.0005)***
Euribor (t – 1)	0.0012 (0.0011)	0.0025 (0.0011)**				
Euribor (t – 2)	0.0023 (0.0011)**					
Euribor Sq.					0.00003 (0.0001)	
Euribor Sq. (t – 1)					0.0003 (0.0001)***	
Real GDP Growth					-0.0705 (0.0199)***	
Constant	0.0028 (0.0024)	0.0050 (0.0020)**	0.0009 (0.0019)	0.0002 (0.0003)	0.0109 (0.0028)***	-0.0013 (0.0015)
ARDL metrics						
R-squared	0.97	0.95	0.97	0.95	0.91	0.97
BIC	-585.8	-588.0	-572.8	-725.5	-598.4	-583.7
AB test p-value	0.19	0.49	0.28	0.49	0.67	0.45
ECM coefficients						
Correction term	-0.4048 (0.0635)***	-0.6266 (0.0838)***	-0.3159 (0.0535)***	-0.3706 (0.0516)***	-0.4419 (0.0776)***	-0.3717 (0.0538)***
LT House Index				-0.0065 (0.0032)**		
LT Unemp.	0.0006 (0.0002)***	0.0003 (0.0001)***	-0.0001 (0.0003)			0.0003 (0.0002)
LT Slope		0.0017 (0.0005)***				
LT Euribor	0.0096 (0.0010)***	0.0082 (0.0005)***	0.0083 (0.0011)***	0.0025 (0.0002)***		0.0095 (0.0009)***
LT Euribor Sq.					0.0009 (0.0001)***	
LT Real GDP					-0.1598 (0.0266)***	
ECM metrics						
ECM R-square	0.6	0.6	0.5	0.5	0.5	0.6
Bounds F-test estat.	17.0	15.0	12.6	22.1	16.0	20.3
Bounds F-test 10% LB	3.2	2.7	3.2	3.2	3.2	3.2
Bounds F-test 10% UB	4.1	3.8	4.1	4.1	4.1	4.1
Bounds t-test estat.	-6.4	-7.5	-5.9	-7.2	-5.7	-6.9
Bounds t-test 10% LB	-3.2	-3.5	-3.2	-3.2	-3.2	-3.2
Bounds t-test 10% UB	-2.6	-2.6	-2.6	-2.6	-2.6	-2.6

NOTES: For each of the asset and liability items indicated in the cols., the panel *ARDL Coefficients* reports OLS estimates for ARDL models in levels as in equation (3) with the average interest rate on the corresponding balance sheet item as dependent variable. These balance sheet items are the interest rate components of NII in equation (1) ($NII = \Sigma_a Vol_a \times Rate_a - \Sigma_l Vol_l \times Rate_l$). For a given explanatory variable, the coefficient is provided with standard error (in parentheses) below it. Reported standard errors are robust to heterocedasticity of arbitrary form. Coefficients for the first and second lag of the dependent variable are provided in rows Lag(1) and Lag(2). When an explanatory variable is not included in any model, it is removed from the table for clarity. Panel *ARDL Metrics* includes the p-value for a first order autocorrelation test of the form given in Arellano-Bond (1991) applied to the residuals of the ARDL models. The panel ECM coefficients reports OLS estimates of correction term α and long term (LT) parameters $\boldsymbol{\theta}$ for the Error Correction Model reformulation (ECM) of ARDL models as in equation (4). In the panel *ECM metrics, Bounds F-test estat* and *Bounds t-test estat* provide statistic values for the test for the presence of an integration relation as in Pesaran, Shin and Smith (2001). The null hypothesis is absence of an integration relation for both the F-test and the t-test. For the F-test, the null is (i) accepted if *Bounds F-test estat* is below the lower bound *Bounds F-test 10% LB* and (ii) rejected if *Bounds F-test estat* is above upper bound *Bounds F-test 10% UB.* For the t-test, the null is accepted if *Bounds t-test estat* is above the upper bound B*ounds t-test 10% UB* and (ii) rejected if *Bounds t-test estat* is below lower bound *Bounds l-test estat 10% LB.* A statistic value between the two bounds is inconclusive for any of the tests. *, **, *** denote significance at the 10%, 5% and 1%.

cycle is unemployment, which enters the specifications for the interest rates of credit, debt securities holdings, rest of assets and rest of liabilities. The use of house price growth and GDP growth provides parsimonious specifications for the interest rates on sight and term deposits. As in the models for volume growth, the AR(1) specification prevail over the AR(2) alternative, which applies only to the model for average interest rate earned on bank credit.

The 12 month Euribor enters linearly all the interest rate models except that of the interest rate on term deposits, where it presents a non-linear lagged positive effect. The higher cost of market financing implied by a higher 12 month Euribor is translated more strongly to the cost of term deposits when the interest rate level is high. This result is consistent with the compression of profit margins on term deposit funds as interest rates near the zero level.

For the remaining categories of assets and liabilities, changes in reference interest rate are translated linearly to their corresponding average interest rates. For traditional bank credit products, this translation is lagged with a high and significant coefficient being applied to the second lag of the 12 month Euribor. There is also a lagged reaction in the debt securities holding category, whereas the relation between 12 month Euribor and the rest of assets and liabilities (interbank positions, wholesale financing products, etc.) is contemporaneous, plausibly reflecting the shorter maturities in these categories.

Examining the ECM reparametrization of the bank average interest rate models, we see that the estimated speed of adjustment α for different interest rates is quite comparable across balance sheet categories, as opposed to the more heterogeneous pattern found for the balance sheet growth models. The exception to this homogenous pattern is the interest rate on debt securities holdings, which presents a faster speed of adjustment than the rest of the models. The long run coefficients on the 12 month Euribor, or the squared term of 12 month Euribor in the model for interest rate on term deposits, are significant and point to the existence of a long term relation between the interbank rate and the average interest rates on different bank balance sheet categories. As regards cointegration, the Pesaran-Shin-Smith bound tests in Table 3 strongly point to the presence of an integration relation. Both the F-test and the t-test reject the null hypothesis of no cointegration relation at the 10% level in all cases.

6.3 ESTIMATED MODELS FOR OTHER BANK INCOME: OFBI AND PROVISIONING CHARGES

The models implemented for OFBI and provisioning charges are presented in Table 4. The two models present some common elements, such as the inclusion of AR(1) dynamics, a purely non-linear effect of the 12 month Euribor and the use of a single macro variable to control for the state of the business cycle. Additionally, the Pesaran-Shin-Smith bound tests support the presence of an integration relation with macro and interest rate controls for both of these variables, with both tests rejecting the null hypothesis. However, there are also significant differences between the two models. The effect of changes in 12 month Euribor on provisioning charges are more persistent due to both a higher AR(1) coefficient and the absence of compensating lagged terms (the first and second lag of 12 month Euribor have opposing signs and comparable magnitude in the model for OFBI). The effect of the business cycle is controlled with the unemployment variable in the model for OFBI, whereas house price growth (a variable related to general economic conditions, but also specifically to the value of real estate collateral) is applied in the model for provisioning charges.

6.4 BANK INCOME DYNAMICS AND INTEREST RATE SHOCKS

We perform in this subsection a dynamic sensitivity analysis of bank income components to market interest rates by introducing a temporary 100bp one-period shock to the 12 month Euribor at the start date of the 3 year study horizon. This is a pure sensitivity analysis

ESTIMATED MODELS FOR ADDITIONAL BANK INCOME COMPONENTS

TABLE 4

	OFB Income	Provision Charge
ARDL Coefficients		
Lag(1)	0.3154 (0.1171)***	0.6010 (0.1430)***
House Price Growth		-2.0561 (0.6482)***
Unemp.	-0.0082 (0.0047)*	
Euribor Sq.	0.0027 (0.0082)	0.0599 (0.0183)***
Euribor Sq. (t – 1)	0.0226 (0.0143)	
Euribor Sq. (t – 2)	-0.0290 (0.0095)***	
Constant	0.2251 (0.1115)**	-0.1593 (0.1054)
ARDL metrics		
R-squared	0.50	0.82
BIC	-47.3	87.0
AB test p-value	0.57	0.74
ECM coefficients		
Correction term	-0.6556 (0.1095)***	-0.3990 (0.0834)***
LT House Index		-5.1537 (1.7662)***
LT Unemp.	-0.0132 (0.0062)**	
LT Slope		
LT Euribor		
LT Euribor Sq.	-0.0071 (0.0052)	0.1502 (0.0250)***
LT Real GDP		
ECM metrics		
ECM R-square	0.4	0.4
Bounds F-test estat.	13.4	9.6
Bounds F-test 10% LB	3.2	3.2
Bounds F-test 10% UB	4.1	4.1
Bounds t-test estat.	-6.0	-4.8
Bounds t-test 10% LB	-3.2	-3.2
Bounds t-test 10% UB	-2.6	-2.6

NOTES: For each of the P&L income categories indicated in the cols., the panel *ARDL Coefficients* reports OLS estimates for ARDL models in levels as in equation (3) with the year-on-year growth rate of the corresponding P&L category as dependent variable. For a given explanatory variable, the coefficient is provided with standard error (in parentheses) below it. Reported standard errors are robust to heterocedasticity of arbitrary form. Coefficients for the first and second lag of the dependent variable are provided in rows Lag(1) and Lag(2). When an explanatory variable is not included in any model, it is removed from the table for clarity. Panel *ARDL Metrics* includes the p-value for a first order autocorrelation test of the form given in Arellano-Bond (1991) applied to the residuals of the ARDL models. The panel ECM coefficients reports OLS estimates of correction term α and long term (LT) parameters $\boldsymbol{\theta}$ for the Error Correction Model reformulation (ECM) of ARDL models as in equation (4). In the panel *ECM metrics, Bounds F-test estat* and *Bounds t-test estat* provide statistic values for the test for the presence of an integration relation as in Pesaran, Shin and Smith (2001). The null hypothesis is absence of an integration relation for both the F-test and the t-test. For the F-test, the null is (i) accepted if *Bounds F-test estat* is below the lower bound *Bounds F-test 10% LB* and (ii) rejected if *Bounds F-test estat* is above upper bound *Bounds F-test 10% UB.* For the t-test, the null is accepted if *Bounds t-test estat* is above the upper bound *Bounds t-test 10% UB* and (ii) rejected if *Bounds t-test estat* is below lower bound *Bounds l-test estat 10% LB.* A statistic value between the two bounds is inconclusive for any of the tests. *, **, *** denote significance at the 10%, 5% and 1%.

under ceteris paribus conditions rather than a scenario analysis, as we keep constant the remaining factors of the model (macro variables and error term) when we introduce the one-period shock on interest rates. This choice of analytical method seeks exclusively to isolate the sensitivity of bank income to a perturbation of interest rates.[18]

The sample covering 2000-2015 presents periods with very different interest rate levels and business cycle conditions. Given this historical sample and the finding of non-linear terms for interest rates in several of the estimated models in subsections 6.1-6.3, the sensitivity of bank income to interest rates can be expected to differ over different subperiods. Thus, we perform the exercise for both the periods 2007-2009 (including a severe economic downturn and the highest interest rate level in sample) and 2013-2015 (a period of economic recovery combined with the lowest level of interest rates in the sample).

For each modelled variable, we compute first the effect of the 100bp one-period shock on the initial quarter. Then, we measure the effect of the shock propagation through the autoregressive terms of the modelled variable and, when applicable, the lagged terms of the Euribor variables. Finally, we take the difference between the generated path induced by the shock and the historical path observed in the data[19]. We measure in this way the impact of the Euribor shock on each of the 12 quarters of the 3 year period of analysis. For NII, we present both the impact of the temporary interest rate shock on each of its volume and interest rate components, and the figure of NII itself that results from aggregating these components: $NII = \sum_a Vol_a \times Rate_a - \sum_l Vol_l \times Rate_l$ (for asset and liability categories α and l). In order to infer a confidence interval around the estimated effect of the shock, we employ a bootstrap methodology. We take a 1,000 draws of the coefficients of the different estimated models and recompute the effect of the 100bp shock on all the variables of interest.[20] This provides a bootstrapped distribution of the effect of the 12 month Euribor shock on all the components of bank income.

6.4.1 Dynamic effect of interest rate shocks in 2007-2009

In year 2007, the 12 month Euribor was already above the 4% level, a high level relative to the sample average value of 2.45%, and bank provisioning charges were escalating quickly. Thus, a 100bp shock to 12 month Euribor can be a priori expected to put pressure on the debt servicing capacity of bank borrowers and push further the cost of financing for banks. The estimated models allow to quantify more precisely the effect of this Euribor shock.

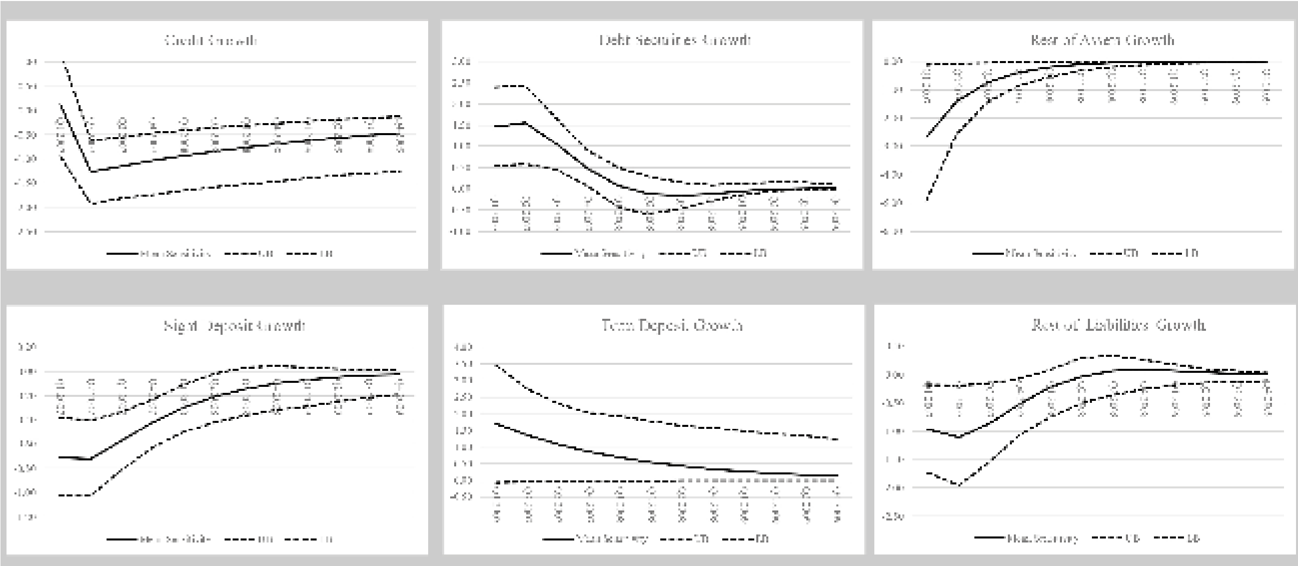
Chart 2 presents the effect of the transitory shock on the year-on-year growth of different assets and liabilities, measured as difference of counterfactual and actual growth rates. Credit growth is slowed down significantly as result of the 100bp shock to 12 month Euribor. Despite the absence of significant initial response on Q1 2007, the decline of 1.3 pp on Q2 2007 is sizeable and confidence intervals stay in the negative territory for the rest of the horizon of analysis. The effect on Q4 2009 is still –0.5 pp with negative confidence

18 Alternative analysis could be constructed, with a staggered calendar of rate changes and permanent changes to the level of the reference rate. Additionally, a consistent macro scenario (with macro variables adapted to the alternative interest rate path) could be applied to obtain the full net effect on bank profitability beyond the pure interest rate effect. In this article, we limit the analysis to this pure rate effect and we use the one period 100bp change as a natural unit of reference.

19 Macro variables other than the 12 month Euribor take the same values in the shock-generated path and in the historical path, so the difference between both paths only reflects the effect of the shock to the 12 month Euribor on the initial quarter, which propagate over several quarters because of the autoregressive structure on the dependent variable and the 12 month Euribor itself.

20 We use the estimated coefficients and their associated variance covariance matrix for each ARDL model in levels to obtain normal random draws.

BALANCE SHEET GROWTH DYNAMICS AFTER 100BP SHOCK TO 12 MONTH EURIBOR (2007-2009) CHART 2



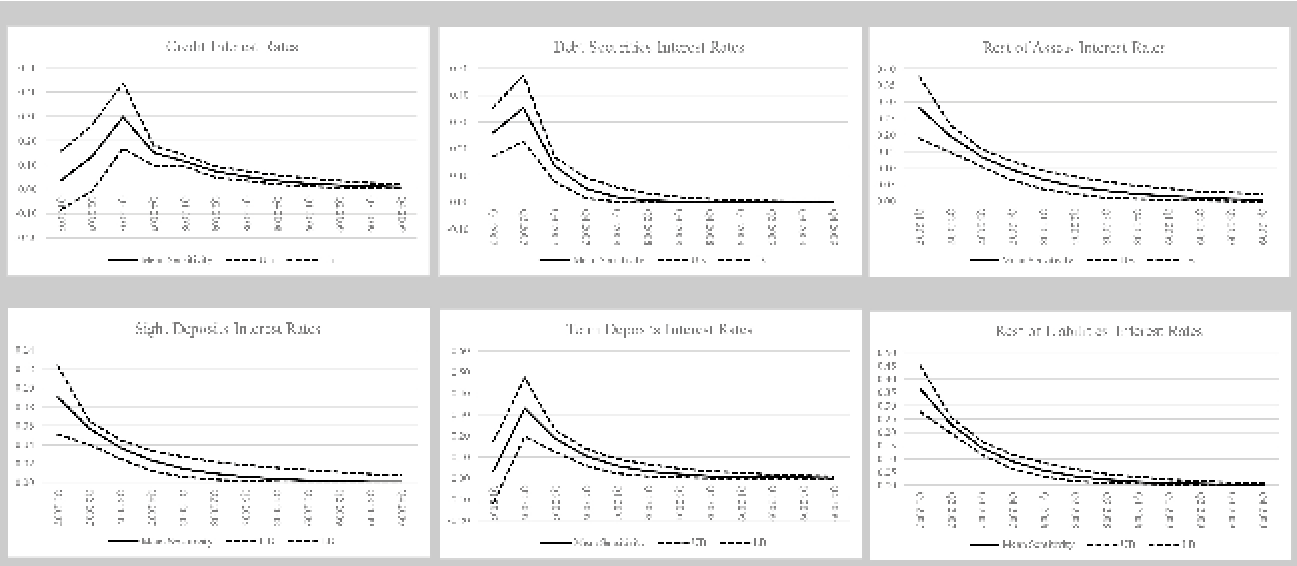
NOTES: For a transitory 100bp shock to 12 month Euribor on Q1 2007, each panel of this figure shows the difference ($y^*_t - y_t$) between counterfactual (y^*_t) and actual (y_t) year-on-year growth rates for the corresponding balance sheet category (e.g., the left-most top panel shows this difference for bank credit growth rate) for each quarter t in the period Q1 2007 – Q4 2009. The models for balance sheet categories are reported in Table 2. The lower (LB) and upper (UB) bounds of the confidence interval correspond with the 5th and 95th percentiles of the distribution of a 1,000 bootstrapped evaluations of the impact of the 12 month Euribor shock.

interval, indicating that the effect on credit does not dissipate quickly. The effect on the growth rate of the rest of assets category (covering for example interbank exposures) is also negative, with a big initial decline of –2.7 pp on Q1 2007 that dissipates rapidly (–0.2 pp on Q1 2008 and –0.002 pp on Q4 2009). On the contrary, the volume of debt securities holdings increases as result of the shock, even though the increase is only significant over the first four quarters of analysis (Q1 2007 – Q4 2007). As the 12 month Euribor goes up, we observe a substitution of traditional bank credit towards debt securities holdings.

On the liability side, we observe substitution from sight to term deposits, which is natural given the rise in their reference interest rate. The confidence intervals for the reduction in growth of sight deposits stay in the negative region from Q1 2007 to Q2 2008, whereas the confidence intervals for the impact on term deposit growth are in the positive region, but they marginally contact zero. The effect on the rest of liabilities is clearly negative, with an initial decline of 0.95 pp that persists as significant for four quarters. The additional market tension introduced by the interest rate shock, plausibly leads to a decline of market financing and greater reliance on term deposits.

Chart 3 displays the effect of the transitory 100bp shock to 12 month Euribor on the average interest rates corresponding to the different asset and liability categories. The rise in the 12 month Euribor increases the levels of all bank interest rates, with a maximum effect of the shock in a given quarter in the 0.3 pp-0.4 pp range for all the series, except sight deposits (with maximum initial reaction of 0.1 pp that dissipates quickly). The differences in the speed of adjustment of different categories are relevant to understand the net changes in NII growth over time. On the asset side, it takes up to three quarters to observe a significant positive effect on credit interest rates, whereas the effect is immediately positive on the debt securities and rest of assets categories, which are more directly linked to wholesale financial markets. On the liability side, the interest rate

AVERAGE BANK INTEREST RATE DYNAMICS AFTER 100BP SHOCK TO 12 MONTH EURIBOR (2007-2009) CHART 3



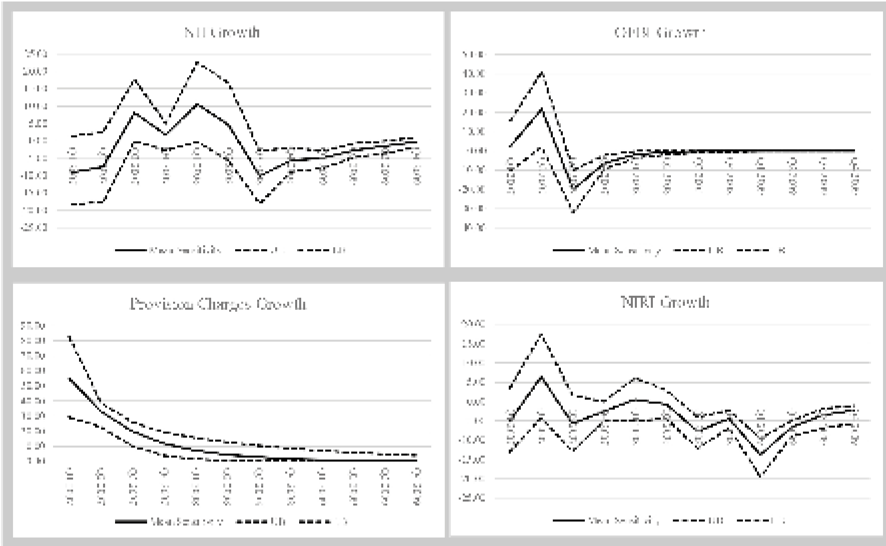
NOTES: For a transitory 100bp shock to 12 month Euribor on Q1 2007, each panel of this figure shows the difference ($y_t^* - y_t$) between counterfactual (y_t^*) and actual (y_t) average interest rates, in pp, for the corresponding balance sheet category (e.g., the left-most top panel shows this difference for the average interest rate on bank credit) for each quarter in the period Q1 2007 – Q4 2009. The models for interest rates are reported in Table 3. The lower (LB) and upper (UB) bounds of the confidence interval correspond with the 5th and 95th percentiles of a 1,000 bootstrapped evaluations of the impact of the 12 month Euribor shock.

cost of the rest of liabilities experiences the strongest effect on the first quarter of analysis (Q1 2007), but the effects of the shock remain sizeable for the whole first year. The response of the cost of term deposits is more lagged (with a peak on Q2 2007), but it is still sizeable.

Chart 4 presents the effects of the 12 month Euribor transitory shock on year-on-year growth rates of NIRI and its components (NII, OFBI and provisioning charges). Firstly, the changes in volume growth and interest rates displayed in Charts 2 and 3 lead to an initial decline of NII year-on-year growth (–9.1 pp on Q1 2007) that is then reversed for some periods (increases in 5 pp-10 pp range from Q3 2007 to Q2 2008) as some of the initial shocks on volumes dissipate and the interest rate on credit picks up, boosting interest rate income. Once all effects have peaked, we observe however a negative, albeit declining, negative effect on NII growth for the last six quarters of analysis (Q4 2008 to Q4 2009). The net effect over the three years of analysis is a lower NII growth, as decline in volume of activity and higher cost of funding dominate over the higher interest rates earned on bank credit and securities holdings.

For OFBI, we observe a sizeable though short-lived effect, with positive increases in year-on-year growth on the first two initial quarters (2.5 pp on Q1 2007 and 21.5 pp on Q2 2007). The effect turns however negative on Q3 2007 (–19.8 pp on Q3 2007), compensating the initial positive effect, and declines quickly afterwards. This pattern is consistent with the use of OFBI to hedge variations in NII, as we observe opposite signs to the effects over NII and the response of OFBI is mostly short term.

The effect of the interest rate shock on provision charges is clearly positive (over 50 pp on Q1 2007) and it remains sizeable until the second half of 2008. Finally, examining the net effect on the year-on-year growth of NIRI, we observe that it is negative on all quarters except on Q2 2007, where the hedging effect of OFBI compensates the negative impact



NOTES: For a 100bp shock to 12 month Euribor on Q1 2007, each panel of this figure shows the difference ($y^*_t - y_t$) between counterfactual (y^*_t) and actual (y_t) year-on-year growth rates for the corresponding P&L component (e.g., the left-most top panel shows this difference for NII growth rate) for each quarter t in the period Q1 2007 – Q4 2009. The lower (LB) and upper (UB) bounds of the confidence interval correspond with the 5th and 95th percentiles of the distribution of a 1,000 bootstrapped evaluations of the impact of the 12 month Euribor shock. The following variable definitions apply: NII (Net Interest Income), OFBI (Other Financial and Banking Income) and NIRI (Net Interest Related Income). NII is modelled as a function of balance sheet growth and interest rates according to equation (1), OFBI and provision charges are modelled directly, and NIRI is modelled as function of the other variables based on equation (2).

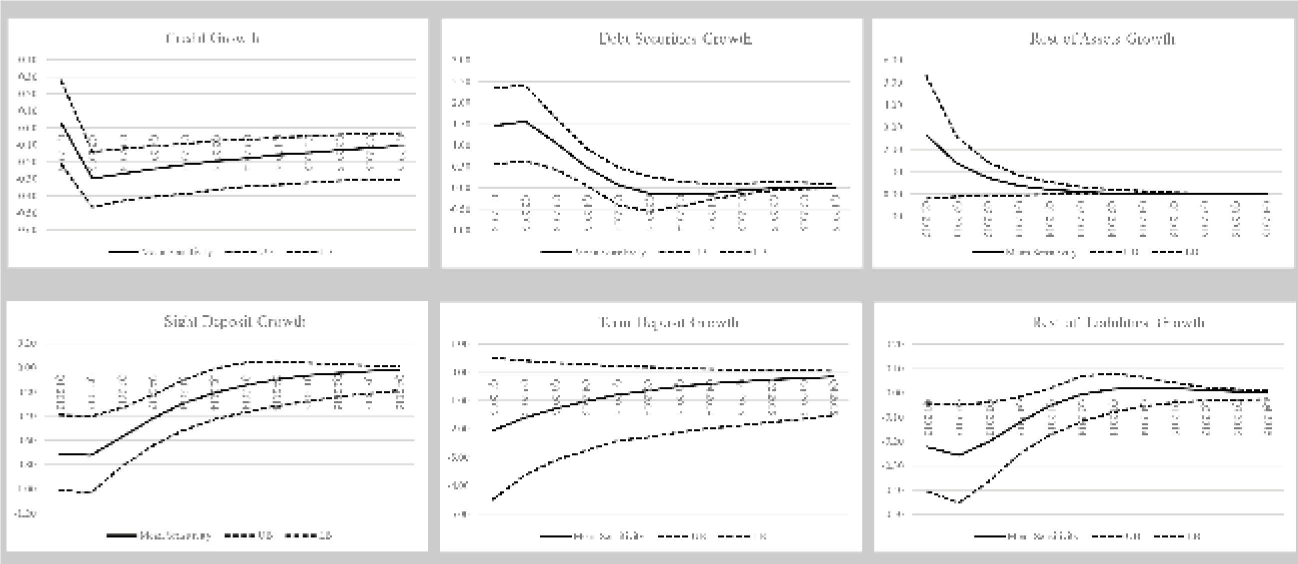
on other profit components. It must be noted however that the confidence interval for quarters in the first half of the study period crosses zero, reducing the significance of the effect of the 12 month Euribor on NIRI growth. Starting on Q3 2008, the confidence intervals stay consistently below zero, indicating a significant effect of the shock in these latter quarters. As the positive effects on NII growth disappear, the negative effects on profitability through lower NII and higher provision charges become dominant and bring down NIRI growth.

6.4.2 Dynamic effect of interest rate shocks in 2013-2015

In year 2013, the Spanish economy was in recession and the provisioning charges of banks remained at high historical levels, but the 12 month Euribor was at a relatively low level (approx. 1%) and the following two years would present a path of positive GDP growth and declining interest rates. Under these different conditions relative to the 2007-2009 period, the 100bp Euribor shock would introduce a priori less pressure on the profit margins of banks. We verify whether this is the case with the model projections.

Chart 5 presents the effect on the growth of different balance sheet categories of a transitory 100bp shock to 12 month Euribor on Q1 2013. As in subsection 6.4.1, the impact on credit growth is negative and persistent, but the effect is weaker. For example, the initial effect (Q1 2013) is again insignificant and the slowing of credit growth begins in the second quarter (Q2 2013) with a shock of –0.3 pp, which is smaller than the shock of –1.3 pp on Q2 2007 measured in the previous experiment. Given that the model implemented for debt securities is linear in 12 month Euribor, the projected path of this variable is positive and coincides with that of the previous experiment in subsection 6.4.1. Interestingly, the effect on the growth of the rest of assets is now found to be positive, though the wide confidence intervals do not allow to preclude a null effect. Based on the results of the two experiments, the increase in 12 month Euribor contracts (expands) this form of exposure when this reference rate is at high (low) levels.

BALANCE SHEET GROWTH AND 12 MONTH EURIBOR DYNAMICS (2013-2015) CHART 5



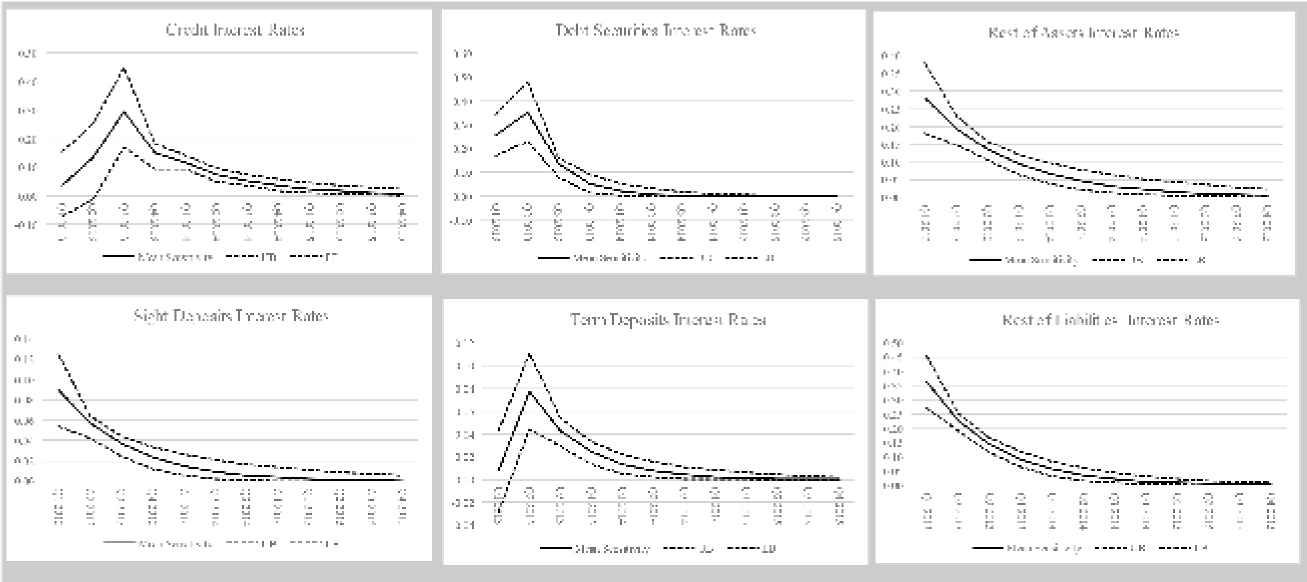
NOTES: For a transitory 100bp shock to 12 month Euribor on Q1 2013, each panel of this figure shows the difference ($y^*_t - y_t$) between counterfactual (y^*_t) and actual (y_t) year-on-year growth rates for the corresponding balance sheet category (e.g., the left-most top panel shows this difference for bank credit growth rate) for each quarter t in the period Q1 2013 – Q4 2015. The models for balance sheet categories are reported in Table 2. The lower (LB) and upper (UB) bounds of the confidence interval correspond with the 5th and 95th percentiles of the distribution of a 1,000 bootstrapped evaluations of the impact of the 12 month Euribor shock.

The lower decline on asset growth as result of the shock to 12 month Euribor is also translated into a more stable liability structure. The positive effect on the growth of term deposits in the previous experiment now disappears, and we observe instead a contraction also in this term of financing. However, the confidence intervals for the impact on term deposit growth are wide and cross zero, so we conclude there is no significant effect on this variable. The negative effect on sight deposits stays the same as the model for this variable is linear on 12 month Euribor. The negative effect on the growth of the rest of liabilities is also smaller (e.g., a reduction of approx. –0.22 pp on Q1 2013 relative to approx. –1 pp. on Q1 2007 on the previous experiment). We still observe in the liability mix a relative substitution towards term deposits, as this form of financing stays approx. constant whereas sight deposits and rest of liabilities decline.

The effect of the transitory shock to the 12 month Euribor is presented in Chart 6. As before, a higher 12 month Euribor implies higher term deposit rates, but the effect of a 100bp hike to the interbank rate has now a much weaker effect, with a peak effect of 0.08 pp rather than 0.33 pp in the previous experiment. For the rest of bank interest rates, their projected reactions coincide with those presented in the previous experiment, since their models are linear in 12 month Euribor.

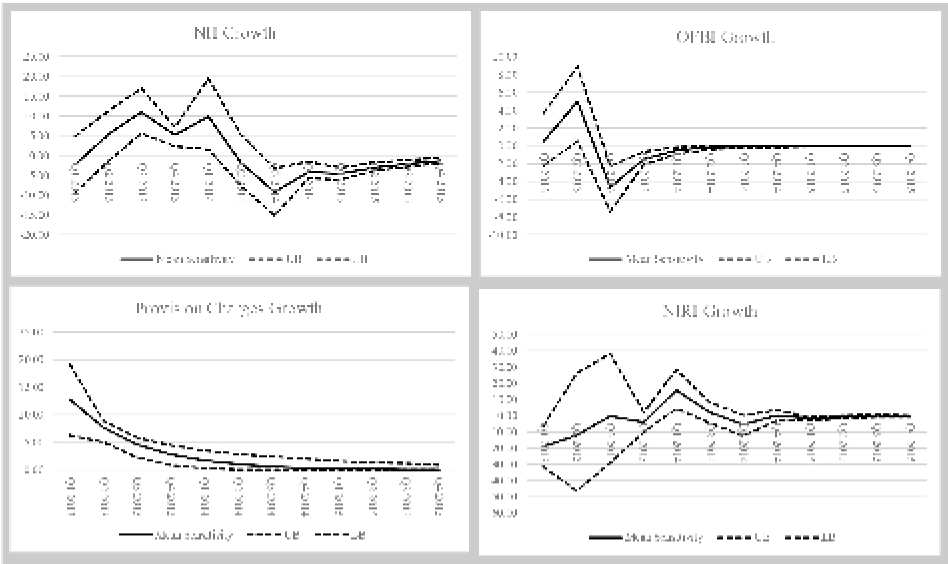
Chart 7 collects the impact of the transitory 100bp shock to 12 month Euribor on NIRI growth and its components. The patterns observed are comparable to the previous experiment, but the levels of the effects and the quantitative conclusions differ. Firstly, the weaker contraction of credit volume and the more moderate increase of the cost of term deposits lead to a lower initial contraction of NII (–2.3 pp on Q1 2013 relative to –9.1 pp on Q1 2007) and a significant acceleration of NII growth from Q2 2013 to Q2 2014 (this effect on NII results from the combination of effects on its components displayed in Charts 3, 5 and 6). In the low interest rate environment of the latter period of the sample, the increase

AVERAGE BANK INTEREST RATE DYNAMICS AFTER 100BP SHOCK TO 12 MONTH EURIBOR (2013-2015) CHART 6



NOTES: For a transitory 100bp shock to 12 month Euribor on Q1 2013, this figure shows the difference ($y_t - y_t^*$) between counterfactual (y_t) and actual (y_t^*) average term deposit rates, in pp, for each quarter t in the period Q1 2013 – Q4 2015. The models for interest rates are reported in Table 3. The lower (LB) and upper (UB) bounds of the confidence interval correspond with the 5th and 95th percentiles of a 1,000 bootstrapped evaluations of the impact of the 12 month Euribor shock.

BANK PROFITABILITY DYNAMICS AFTER 100BP SHOCK TO 12 MONTH EURIBOR (2013-2015) CHART 7



NOTES: For a 100bp shock to 12 month Euribor on Q1 2013, each panel of this figure shows the difference ($y_t^* - y_t$) between counterfactual (y_t^*) and actual (y_t) year-on-year growth rates for the corresponding P&L component (e.g., the left-most top panel shows this difference for NII growth rate) for each quarter t in the period Q1 2013 – Q4 2015. The lower (LB) and upper (UB) bounds of the confidence interval correspond with the 5th and 95th percentiles of the distribution of a 1,000 bootstrapped evaluations of the impact of the 12 month Euribor shock. The following variable definitions apply: NII (Net Interest Income), OFBI (Other Financial and Banking Income) and NIRI (Net Interest Related Income). NII is modelled as a function of balance sheet growth and interest rates according to equation (1), OFBI and provision charges are modelled directly, and NIRI is modelled as function of the other variables based on equation (2).

of 12 month Euribor allows banks to earn higher interest rates on their assets, suffering a smaller penalty in terms of lower volume of activity or increased cost of funds. The impact on OFBI growth is again observed to be short term and to offset to some extent the fluctuations on NII year-on-year growth, but the effect is now smaller in absolute terms (±4 pp range as compared to ±20 pp range in the previous experiment).

CUMULATIVE EFFECT OF INTEREST RATE SHOCKS

TABLE 5

	2007-2009 (a)			2013-2015 (b)		
	Δ%	LB	UB	Δ%	LB	UB
NII	-2.3	-6.1	0.4	3.0	1.1	5.0
NIRI	-6.1	-9.7	-3.2	-3.2	-6.5	-0.1
Total Assets	-1.4	-2.8	-0.5	-0.2	-0.5	0.1

NOTES: Each panel of this table displays the cumulative effect of a 100bp shock to 12 month Euribor on key bank profit and volume of activity variables. For a given variable, we provide the relative variation $\Delta\% = (y^* - y) / y$ between the counterfactual level of the variable y^* implied by the Euribor shock and the actual value y. The lower (LB) and upper (UB) bounds of the confidence interval correspond with the 5th and 95th percentiles of the distribution of a 1,000 bootstrapped evaluations of the impact of the 12 month Euribor shock. Panel (a) displays results for the period 2007-2009 with Euribor shock on Q1 2007 and Panel (b) displays results for the period 2013-2015 with Euribor shock on Q1 2013. The following variable definitions apply: NII (3 year cumulative sum of net interest income), NIRI (3 year cumulative sum of net interest related income) and Total Assets (sum of credit, debt securities holdings and rest of assets on the last quarter of the horizon of analysis). For example, the table displays in panel (a) the change Δ% in cumulative NII and NIRI for 2007-2009, and the change Δ% in the standing volume of total assets on final quarter Q4 2009. Neither NII nor NIRI are modelled directly. NII is modelled in terms of equation (1) ($NII = \sum_a Vol_a \times Rate_a - \sum_l Vol_l \times Rate_l$) and it is thus a function of the 6 models for balance sheet items (reported in Table 2) and the six models for bank rates (reported in Table 3). NIRI is computed with equation (2) and it is thus a function of NII and the models for OFBI and provisions in Table 4.

The increase in provision charges growth is also less marked, with an initial effect of approx. 12 pp on Q1 2013 that has largely dissipated by Q2 2014. It must be noted however that the level of provision charges at the beginning of 2013 was much higher than on the year 2007 (as shown in Chart 1) so a smaller acceleration of the growth of provision charges can have a greater impact on NIRI than in the experiment for the 2007-2009 period. When we observe the net impact on NIRI growth, we effectively observe a higher initial decline (–20 pp on Q1 2013 relative to –5.2 pp on Q1 2007) despite the lower negative impact on NII growth. As the effect on provision charges growth dissipates and the higher NII growth takes hold during 2014, the initial negative effect is reversed and we observe higher NIRI growth due to the shock to the Euribor. For the last year of analysis, 2015, the effect on NIRI growth is almost nil as opposed to the negative effect that we found on the previous experiment.

6.4.3 Cumulative effects of interest rate shocks

In this subsection, we measure the cumulative effect of the interest rate shocks on bank profits and volume of activity over the complete horizon of the experiments for 2007-2009 and 2013-2015. This information complements the quarter by quarter analysis of subsections 6.4.1 and 6.4.2 and allows for a more precise comparison of the two periods. Table 5 shows that the effect of the transitory 100bp Euribor shock on the total NII (sum of NII for the corresponding three year study period) is positive (+3%) during the latter low interest rate period of 2013-2015, but negative (–2.3%) during the high interest rate period of 2007-2009. The confidence interval for this earlier period is wide and crosses zero so a nil effect can not be ruled out. This aggregate result is driven by factors already commented, such as the very different impact of interest rate hikes on the volume of activity of banks and cost of funds for high and low interest rate levels (NII is not modelled with a single aggregate model, but with the combination of 12 models for balance sheet items and bank rates). It is important to notice that the higher NII attained after a positive Euribor shock at low interest rate levels is made possible partly through a substitution of credit towards debt securities, potentially hurting some bank borrowers.

Table 5 also displays the net effect on total NIRI produced over the different study periods, which is significantly negative for both experiments, but greater in magnitude for the 2007-2009 period (–6.1% relative to –3.2%). The shock to 12 month Euribor reduces NII more and contributes to faster growth of provision charges during high interest rate periods. The volume of interest producing assets, which proxies for the total value creation of the banking sector beyond bank profits, is negatively and significantly affected (–1.4%) in the 2007-2009 period, but is associated to an almost nil effect for 2013-2015. The interest rate increase from a very low level would not affect significantly the size of productive assets of the banking sector.

7 Final Considerations

A careful analysis of the relation of bank profitability, and more generally value creation in the banking sector, with the level of interest rates must recognize the multiple channels through which interest rates affect bank profits. The changes in the level of interest rates do not affect only the return on funds through its impact on the average interest rates corresponding to bank assets and liabilities, but also the maximum volume of activity that the banking sector can attain in a given period. The analysis of the return on funds can be insufficient to determine whether an interest rate change will lead to higher bank profits, because the return on funds and volume of activity are interconnected.

Breaking down the components of profit variation can be useful to evaluate the impact of these changes on different bank stakeholders. We find that higher interest rates lead to substitution from bank credit to debt securities on the asset side, and, to some extent, from wholesale financing to term deposits on the liability side. These shifts in asset and liability composition will plausibly hurt bank borrowers and benefit bank depositors. Bank shareholders (through lower profits) are generally negatively affected by interest rate increases in the studied environment. Even though higher rates can boost NII growth at low interest rate levels, they lead to higher provisioning charges, dragging net interest related profitability. At high interest rate levels, further interest rate increases are found to deteriorate profitability both through NII and provisioning channels. The relation between the interest rate level and bank profitability and balance sheet structure is therefore nonlinear, varying as a function of the level of interest rates and bank balance sheet composition.

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Annex A – Additional Elements of the Model Selection Framework

The restrictions on the signs of the model coefficients impose a positive (negative) relation of the business cycle with credit loan growth (provision charges). Restrictions also require a negative (positive) relation of the interbank rate with credit loan growth (provision charges). For interest rate models, we impose generally a positive relation with the interbank rate and a negative relation with the business cycle. We refrain from imposing restrictions over variables for which there is weaker a priori information, such as most volume growth models. For example, OFBI might increase or decrease with the economic cycle as a function of the financial market position of banks.

As the interbank rate potentially enters the models with a quadratic term, the restrictions are imposed on the derivatives at the extremes of the range of possible interbank rate values. For example, if we take Euribor as the interbank rate measure and assume that it enters a model just contemporaneously as $\beta_0 \cdot \text{Euribor} + \beta_1 \cdot \text{Euribor}^2$, the derivative with respect to Euribor is then given by $\beta_0 + 2\beta_1 \cdot \text{Euribor}$, which is a monotonically increasing (decreasing) function if $\beta_1 > 0$ ($\beta_1 < 0$). Given this monotonicity of the derivative, restricting its sign to be positive (or negative) at both extremes of the range of interbank values is sufficient to restrict the sign in all that range. Table B1 sums up all the restrictions imposed on the models.

DEPENDENT VARIABLES WITH SIGN RESTRICTIONS **TABLE A.1**

Variable	Type	House price growth	Unemployment	Slope	Real GDP growth	Euribor
Credit	Volume Growth	≥0	≤0	—	≥0	≤0
Credit	Rate	≤0	≥0	—	≤0	≥0
Debt Security	Rate	≤0	≥0	—	≤0	≥0
Rest of Assets	Rate	≤0	≥0	—	≤0	≥0
Sight Deposits	Rate	≤0	≥0	—	≤0	≥0
Term Deposits	Rate	≤0	≥0	—	≤0	≥0
Rest of Liabilities	Rate	≤0	≥0	—	≤0	≥0
Provision Charge	Aggregate	≤0	≥0	—	≤0	≥0

NOTES: For macroeconomic variables different from Euribor, the sign restriction applies to the sum of the three potential coefficients in the model (β_0, and the two lags β_1 and β_2): $\beta_0 + \beta_1 + \beta_2$. For example, the sign restriction for the effect of house price growth on credit growth is $\beta_0 + \beta_1 + \beta_2 \geq 0$. For the relation of Euribor with credit growth, we consider linear (β_0, β_1, β_2) and quadratic effects (β'_0, β'_1, β'_2), and we require a negative long term derivative, i.e., $\beta_0 + \beta_1 + \beta_2 + 2 \cdot \text{Euribor}\,(\beta'_0 + \beta'_1 + \beta'_2)$ for all the range of Euribor values. The role of Euribor as reference rate connects it more directly with average interest rates on bank assets and liabilities and with provision charge growth (the prevalence of variable rate contracts in Spain links financial burden to Euribor). Thus, we apply a stricter sign restriction for these variables, requiring each lag component ($\beta_0 + 2 \cdot \text{Euribor} \cdot \beta'_0$, $\beta_1 + 2 \cdot \text{Euribor} \cdot \beta'_1$, $\beta_2 + 2 \cdot \text{Euribor} \cdot \beta'_2$) of the long term derivative to satisfy the corresponding restriction. If a particular coefficient does not appear in a specification, then it is taken as zero for the verification of the sign conditions.

Annex B – Unit Root Tests

As described in subsection 5.1., we use Augmented Dickey-Fuller and KPSS tests to examine whether there is evidence of I(1) and I(2) behavior of the variables in the data set. The results of these unit root tests in Table B1 are consistent with I(1) dynamics in the data,

ADF AND KPSS TESTS **TABLE B.1**

	ADF	KPSS		
		Min.	Max.	Lag Order
A Macro Variables				
Euribor	-2.01	0.51**	1.91***	10
Euribor Sq.	-2.73*	0.41*	1.37***	10
House Price Growth	-1.26	0.50**	2.43***	10
Unemp.	-0.83	0.51**	2.45***	10
Real GDP	-2.02	0.42*	1.61***	10
Slope	-2.08	0.31	1.23***	10
Dif. - Euribor	-5.32***	0.06	0.08	10
Dif. - Euribor Sq.	-6.56***	0.05	0.06	10
Dif. - House Price Growth	-3.07**	0.15	0.34	10
Dif. - Unemp.	-4.31***	0.18	0.39*	10
Dif. - Real GDP	-4.11***	0.20	0.28	10
Dif. - Slope	-5.02***	0.09	0.12	10
B Bank Variables				
Prov. Charges Growth	-2.18	0.11	0.28	10
OFBI Growth	-3.20**	0.30	0.51**	10
Credit Growth	-1.01	0.46*	2.16***	10
Debt Securities Growth	-3.49***	0.13	0.22	10
Rest of Assets Growth	-2.14	0.23	0.73**	10
Sight Deposits Growth	-2.10	0.21	0.61**	10
Term Deposit Growth	-1.09	0.37*	1.52***	10
Rest of Liabilities Growth	-1.81	0.46**	1.83***	10
Credit Int. Rate	-1.01	0.49**	1.81***	10
Deb Securities Int. Rate	-1.52	0.51**	1.73***	10
Rest of Assets Int. Rate	-1.42	0.58**	2.35***	10
Sight Deposits Int. Rate	-1.42	0.34	1.18***	10
Term Deposit Int. Rate	-1.41	0.13	0.37*	10
Rest of Liabilities Int. Rate	-1.47	0.50**	1.85***	10
Dif. - Prov. Charges Growth	-4.42***	0.06	0.09	10
Dif. - OFBI Growth	-7.45***	0.03	0.09	10
Dif. - Credit Growth	-2.80*	0.10	0.23	10
Dif. - Debt Securities Growth	-4.90***	0.06	0.13	10
Dif. - Rest of Assets Growth	-6.84***	0.10	0.19	10
Dif. - Sight Deposits Growth	-4.17***	0.15	0.22	10
Dif. - Term Deposit Growth	-4.13***	0.15	0.20	10
Dif. - Rest of Liabilities Growth	-3.90***	0.07	0.15	10
Dif. - Credit Int. Rate	-4.06***	0.06	0.09	10
Dif. - Debt Securities Int. Rate	-4.43***	0.06	0.07	10
Dif. - Rest of Assets Int. Rate	-4.65***	0.05	0.07	10
Dif. - Sight Deposits Int. Rate	-4.09***	0.09	0.14	10
Dif. - Term Deposit Int. Rate	-4.32***	0.14	0.19	10
Dif. - Rest of Liabilities Int. Rate	-4.27***	0.05	0.08	10

NOTES: *Growth* indicates the year-on-year growth series and *Int. Rate* the average interest rate series of the corresponding variable. *Dif.* - indicates the first differenced series of the corresponding variable. *ADF* denotes Augmented Dickey Fuller test statistic (Null: Unit Root). For KPSS test (Null: Stationarity), Schwert criterium indicates a 10 lag order, we report the minimum and maximum KPSS test statistic for lag orders 1 to 10. *, **, *** indicates rejection at 10%, 5%, 1% significance level.

but not with I(2) dynamics. It must also be noted that unit root tests can have low power against the alternative of a stationary and persistent process, especially in a limited 16 year time span in the available sample, so I(0) dynamics for the variables of interest cannot be completely ruled out. The ARDL approach used is valid with either I(0) or I(1) dynamics, but not with I(2) dynamics. The results for the unit root tests in this annex thus confirm the adequacy of applying ARDL to the estimation of the models in the main text.