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BANK RISK AND MONETARY POLICY

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Abstract

We find evidence of a bank lending channel for the euro area operating via bank risk. Financial innovation and the new ways to transfer credit risk have tended to diminish the informational content of standard bank balance-sheet indicators. We show that bank risk conditions, as perceived by financial market investors, need to be considered, together with the other indicators (i.e. size, liquidity and capitalization), traditionally used in the bank lending channel literature to assess a bank's ability and willingness to supply new loans. Using a large sample of European banks, we find that banks characterized by lower expected default frequency are able to offer a larger amount of credit and to better insulate their loan supply from monetary policy changes.

JEL classification: E44, E55.
Keywords: bank, risk, bank lending channel, monetary policy.

Contents

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1. Introduction[1]

In contrast to findings for the United States, existing empirical research on the importance of bank conditions in the transmission mechanism of monetary policy provides inconclusive evidence for the euro area. More broadly, the overall judgment concerning the role of financial factors in the transmission mechanism is mixed.[2] This is surprising, since in the euro area banks play a major role as one of the main conduits for the transmission of monetary policy and have a pivotal position in the financial system. The weak evidence for a "bank lending channel" is probably due to two main factors: first, there are significant data limitations, as the bulk of existing evidence was undertaken under the auspices of the Monetary Transmission Network in 2002, which was only a handful of years after the start of monetary union. Second, the role of banks in the transmission mechanism is likely to have changed, mainly because the business of banks has undergone fundamental changes in recent years, owing to financial innovation, financial integration and increases in market funding. In other words, parts of the banking sector have moved away from the traditional "originate-and-hold" to an "originate-and-distribute" model of the banking firm, which is much more reliant on market forces. As a result, it is likely that this new role of banks has an impact on the way they grant credit and react to monetary policy impulses (Loutskina and Strahan, 2006; Hirtle, 2007; Altunbas, Gambacorta and Marqués-Ibáñez, 2009).

Some of the latest literature on the transmission mechanism also underlines the role of banks, by focusing on bank risk and incentive problems arising from/for bank managers. Borio and Zhu (2008) argue that financial innovation, in parallel with changes to the capital regulatory framework (Basel II), is likely to have enhanced the impact of the perception, pricing and management of risk on the behavior of banks. Similarly, Rajan (2005) suggests that more market-based pricing and stronger interaction between banks and financial

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[1] We would like to thank Francesco Columba, Michael Ehrmann, Paolo Del Giovane, Alistair Milne, Fabio Panetta, and participants at the conference "The Transmission of Credit Risk and Bank Stability" (Centre for Banking Studies, Cass Business School, London, 22nd May 2008) for their comments. In particular, we would like to thank two anonymous referees for very insightful comments. This paper was written while Leonardo Gambacorta was at the Economic Outlook and Monetary Policy Department of the Bank of Italy. The opinions expressed in this paper are those of the authors only and in no way involve the responsibility of the Bank of Italy, the ECB or the BIS. Email addresses: Y.Altunbas@bangor.ac.uk, Leonardo.Gambacorta@bis.org and David.Marques@ecb.int.

[2] See Angeloni, Kashyap and Mojon (2003), Ehrmann et al. (2003b), Gambacorta (2005).

markets exacerbates the incentive structures driving banks, potentially leading to stronger links between monetary policy and financial stability effects.

In this paper, we claim that bank risk must be carefully considered, together with other standard bank-specific characteristics, when analyzing the functioning of the bank lending channel of monetary policy. Due to financial innovation, variables capturing bank size, liquidity and capitalisation (the standard indicators used in the bank lending channel literature) may not be adequate for the accurate assessment of banks' ability and willingness to supply additional loans. More broadly, financial innovation has probably changed bank incentives towards risk-taking (Hänsel and Krahnen, 2007; Instefjord, 2005).

In recent years, before the 2007-08 financial turmoil, more lenient credit risk management by banks may have partly contributed to a gradual easing of credit standards applied to loans and credit lines to borrowers. This is supported by the results of the Bank Lending Survey for the euro area and evidence from the United States (Keys at al., 2008 and Dell'Ariccia et al., 2008). The lower pressure on banks' balance sheets was also reflected in a decrease in the expected default frequency, until a reversal in 2007 and more clearly in 2008 (Figure 1).

The 2007-2008 credit turmoil has made it very clear that the perception of risk by financial markets is crucial to the banks' capability to raise new funds. Also, in this respect, the turmoil has affected their balance sheets in different ways. The worsening of risk factors and the process of re-intermediation of assets previously sold by banks to the markets has implied higher actual and expected bank capital requirements At the same time, increased write-offs and the reductions in investment banking activities (M&A and IPOs) have reduced both profitability and capital base. These effects may ultimately imply a restriction of the supply of credit.

According to replies from banks participating in the euro area bank lending survey (BLS), the turbulence in financial markets has significantly affected credit standards and lending supply. The BLS indicated a progressive increase in the net tightening of credit standards for loans to households and firms, especially for large enterprises. A major contribution to the tightening has come not only from tensions in the monetary market, but also from banks' difficulties in obtaining capital or issuing new bonds. Concerning capital needs, banks have made recourse to equity issuance on a large scale to compensate for write-offs. However, due to the higher level of risk, as perceived by the financial markets, and the large amount of capital needed, equity issuance has often relied on new classes of investors, such as sovereign wealth funds. The reassessment of risk has also affected bond issuance: gross issuance of bonds by euro area banks and financial companies declined significantly in the second half of 2007 compared with 2006, and remained very weak in the first part of 2008. All in all, the credit turmoil has vividly demonstrated that the ability of a bank to tap funds on the market and, consequently, to sustain changes in money market conditions is strongly dependent on its specific risk position. It is therefore highly relevant to investigate how the lending supply is influenced by bank risk.

This paper concentrates on the implications of changes described above for the provision of credit supply and the monetary policy transmission mechanism, departing in two ways from the existing literature. First, the paper presents an in-depth analysis of the effects of bank risk on loan supply, using both an ex-post measure of credit risk (loan-loss provisions as a percentage of loans) and an ex-ante measure (the one-year expected default frequency, EDF). The latter is a forward-looking indicator that allows for a more direct assessment of how the markets perceive the effects of a transfer of credit risk impact on bank risk; for instance, due to the use of true-sale securitisation, credit derivatives or synthetic collateralized debt obligations (CDOs). Our second innovation lies in the analysis of the effects of credit risk on the banks' response to both monetary policy and GDP shocks.

We use a unique dataset of bank balance sheet items and asset-backed securities for euro area banks over the period 1999 to 2005. The estimation is performed using an approach similar to that of Altunbas, Gambacorta and Marqués (2009), who analyse the link between securitisation and the bank lending channel. To tackle problems derived from the use of a dynamic panel, all the models have been estimated using the GMM estimator, as suggested by Arellano and Bond (1991).

The results indicate that low-risk banks are able to offer a larger amount of credit and can better shield their lending from monetary policy changes, probably due to easier access to uninsured fund raising, as suggested by the "bank lending channel" hypothesis. Interestingly, this insulation effect is dependent on the business cycle and tends to decline in the case of an economic downturn. Risk also influences the way banks react to GDP shocks. Loan supply from low-risk banks is less affected by economic slowdowns, which probably reflects their ability to absorb temporary financial difficulties on the part of their borrowers and preserve valuable long-term lending relationships.

The remainder of this paper is organised as follows. The next section discusses the econometric model and the data. Section 3 presents our empirical results and robustness checks. The last section summarises the main conclusions.

2. The econometric model and the data

Empirically, it is difficult to measure the effect of bank conditions on the supply of credit by using aggregate data, as it not easy to disentangle demand and supply factors. To date, this "identification problem" has been addressed by assuming that certain bank-specific characteristics (such as size, liquidity and capitalization) influence the supply of loans. At the same time, loan demand is largely independent of bank specific characteristics and mostly dependent on macro factors. The empirical specification used in this paper is similar to that used in Altunbas, Gambacorta and Marqués (2009) and is designed to test whether banks with a different level of credit risk react differently to monetary policy shocks.[3]

The empirical model is given in the following equation:[4]

$$\begin{aligned}
\Delta \ln(Loans)_{i,t} &= \alpha \Delta \ln(Loans)_{i,t-1} + \sum_{j=0}^{1} \delta_j \Delta \ln(GDPN)_{k,t-j} + \sum_{j=0}^{1} \beta_j \Delta i_{M\,t-j} + \sum_{j=0}^{1} \phi_j \Delta i_{M\,t-j} * EDF_{i,t-1} + \\
&+ \sum_{j=0}^{1} \sigma_j \Delta i_{M,t-j} * SIZE_{i,t-1} + \sum_{j=0}^{1} \lambda_j \Delta i_{M,t-j} * LIQ_{i,t-1} + \sum_{j=0}^{1} \chi_j \Delta i_{M\,t-j} * CAP_{i,t-1} + \kappa SIZE_{i,t-1} + \vartheta LIQ_{i,t-1} + \\
&+ \xi CAP_{i,t-1} + \tau LLP_{i,t-1} + \psi EDF_{i,t-1} + \varepsilon_{i,t}
\end{aligned} \qquad (1)$$

with i=1,…, N , k= 1, …,12 and t=1, …, T where N is the number of banks, k is the country and T is the final year.

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[3] For a similar empirical approach, see also, among others, Kashyap and Stein (1995, 2000), Ehrmann et al. (2003a,b) and Ashcraft (2006). A simple theoretical micro-foundation of the econometric model is reported in Ehrmann et al. (2003a) and Gambacorta and Mistrulli (2004).

[4] The model in levels implicitly allows for fixed effects and these are discarded in the first difference representation given in equation (1).

In equation (1) the growth rate in bank lending to residents (excluding interbank positions), $\Delta\ln(Loans)$,[5] is regressed on nominal *GDP* growth rates, $\Delta\ln(GDPN)$, to control for country-specific loan demand shifts. Better economic conditions increase the number of projects becoming profitable in terms of expected net present value, thereby increasing the demand for credit (Kashyap, Stein and Wilcox, 1993). The introduction of this variable captures cyclical macroeconomic movements and serves to isolate the monetary policy component of interest rate changes (Δi_M). The econometric specification also includes interactions between changes in the interest rate, controlled by the monetary policy authority, and bank-specific characteristics. The first three bank-specific characteristics are standard in the literature: *SIZE*, the log of total assets (Kashyap and Stein, 1995), *LIQ*, securities and other liquid assets over total assets (Stein, 1998), *CAP*, the capital-to-asset ratio (Kishan and Opiela, 2000; Van den Heuvel, 2002).

The fourth bank-specific characteristic, which represents the main innovation in this paper, is the bank's risk position, proxied by two variables. The first variable (*LLP*) is loan-loss provisions as a percentage of loans; this is standard in the literature and can be regarded as an ex-post accounting measure of credit risk. The second variable is the one-year ahead expected default frequency (*EDF*), which is commonly used as a measure of credit risk by financial institutions, including central banks and regulators (see, for instance, ECB, 2006, and IMF, 2006).[6] *EDF* is a forward-looking indicator of credit risk computed by Moody's KMV using financial markets data, balance sheet information and Moody's proprietary bankruptcy database.[7] However, *EDF* information is not available for all banks. From 1999 to 2005, the sum of total assets of banks for which Moody's KMV constructs *EDF* figures accounts for around 52% of the total assets of banks in our sample. For banks that do not

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[5] As discussed in Jeffrey (2006), securitisation may dramatically affect bank loans dynamics. Standard statistics do not take into account that fully securitised loans (i.e. those expelled from banks' balance sheets) continue to finance the economy. We aim to tackle this statistical issue by simply re-adding the flows of securitised loans (*SL*) to the change in the stock of loans, to calculate a corrected measure of the growth rate for lending that is independent of the volume of asset securitisation ($\Delta\ln L_t=\ln(L_t+SL_t)-\ln L_{t-1}$). Securitisation data are obtained from the Bondware database combined with other data providers (for more details see Altunbas, Gambacorta and Marqués, 2009).

[6] Furfine and Rosen (2006) use *EDF* to assess the effect of mergers on U.S. banks' risk.

[7] The calculation of *EDF* builds on Vasicek and Kealhofer's extension of the Black-Scholes-Merton option-pricing framework, which makes it suitable for practical analysis, and on the proprietary default database owned by KMV. (For further details on the construction of *EDF*s and applications, see: Crosbie and Bohn, 2003; Kealhofer, 2003; and Garlappi, Shu and Yan, 2007).

have *EDF* figures, we have approximated their default probability in two ways: first, by means of a cluster analysis; second, by estimating the missing *EDF* values using a regression model.

For the first method (cluster analysis), we have grouped banks by year, country, bank size (big, medium, small) and institutional categories (limited companies, mutual banks, cooperative banks). We have then assigned banks with missing EDFs, the value of the more similar group.

For the second method, we used the following model:

$$EDF_{i,t} = \sum_{h=1}^{10} a_h X_{h,i,t} + \sum_{k=1}^{12} b_k C_{k,i,t} + \varepsilon_{i,t} \qquad (2)$$

where the expected default frequency (*EDF*) for bank i at time t is regressed on a vector of 10 banks' balance sheet variables ($X_{i,t}$) and country dummies (C_k) that take the value of 1 if bank i has its main seat in country k and zero elsewhere (these dummies have been inserted in order to capture specific institutional characteristics). The vector of explanatory variables (X) includes: net interest margin over total assets (profitability indicator), other operating income over total assets (earnings diversification), liquid assets over deposits (liquidity management), cost-to-income ratio (efficiency), non-interest expenses over total liabilities (cost structure), equity to total asset ratio (capital adequacy), loan-loss provisions over net interest margin (asset quality), interbank ratio (market based funding), net loans over total asset (weight of traditional intermediation activity) and securities over total assets (weight for investment portfolio activity).[8]

[8] In order to compare the correspondence between the predicted and the observed values of *EDF*, we checked in-sample and out-of-sample performance of the regression. For the in-sample performance, we have computed the mean forecast error and the mean quadratic error for 10 banks randomly excluded from the sample. The two statistics turned out to be 0.012 and 0.002, respectively, two values that seems quite contained. However, this test is not sufficient to test the goodness of the model because the regression has to estimate values of *EDF* for banks that are not in the sample. We, therefore, also computed an out-of-sample test, as follows: the 10 banks' observed *EDF* values were gathered, then we regressed model (2) for the full sample and computed the mean forecast error and the corresponding mean quadratic error for the 10 banks. Also in this case the two statistics turned out to be quite contained (0.033 and 0.008, respectively). To further corroborate the reliability of the *EDF* regression, we tested the difference between the mean of the forecasted *EDF* and the observed one, and were able to accept the null hypothesis of no difference between the two aggregated statistics (the pair t-

Coefficients a_h and b_k are calculated to estimate the value of the *EDF* for those banks (mainly small ones) for which the KMV *EDF* is not available. It is worth noting that the average value for the *EDF* for the whole sample (including estimated values) is higher than that for the subset of banks that have an *EDF* estimated directly by KMV (see Table 1). This captures the fact that by means of the estimation method we attach a probability to go into default to small banks. By including them into the analysis, the average value of the *EDF* increases. The two *EDF* measures are slightly correlated with *LLP* (the correlation if 0.11* when the missing values for EDF are approximated by means of a cluster analysis and 0.03* when *EDF* is approximated by a regression).[9]

Bank-specific characteristics refer to *t-1* in order to avoid endogeneity bias. Following Ehrmann et al. (2003a), all bank-specific characteristics have been normalised with respect to their average across all banks in their respective samples, in order to get indicators that amount to zero over all observations. This means that for model (1) the averages of the interaction terms are also zero and the parameters β_j may be broadly interpreted as the average monetary policy effect on lending for a theoretical average bank.

The sample period is from 1999 to 2005,[10] a period characterised by a homogenous monetary regime for all the banks considered. The interest rate used as one of the monetary policy indicators is the three-month Euribor rate, which captures the effective cost of interbank lending on the monetary market. In the period considered, the dynamic of this variable is the same as that of the policy rate (the correlation between the two monetary policy indicators is above 98%).

The analysis uses annual data obtained from BankSscope, a commercial database maintained by International Bank Credit Analysis Ltd. (IBCA) and the Brussels-based Bureau van Dijk. In particular, we consider balance sheet and income statement data for a sample of around 3,000 euro area banks. Table 1 presents some basic information on the

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test value is 0.58 with p<0.288, df=9).The output of the regressions has not been included in the text for the sake of brevity. All results are available from the authors upon request.

9 In equation (1) we consider only the interaction between the monetary policy indicator and EDF because it allows a more direct assessment of how the markets perceive bank risk as it is a forward-looking indicator.

10 Data for 1998 have also been included to calculate growth rates.

dataset.[11] The sample accounts for around three quarters of bank lending to euro area residents. The average size of banks in the sample is largest in the Netherlands, Finland and Belgium and smallest in Austria, Germany and Italy. The averages of individual bank characteristics differ across countries in terms of capital, loan-loss provisions and liquidity characteristics, reflecting different competitive and institutional conditions, as well as different stages of the business cycle.

In Table 2, banks are grouped depending on their specific risk position, using the estimated *EDFs* (very similar results are obtained using the cluster measure). A "high-risk" bank has the average *EDF* of banks in the fourth quartile (i.e. EDF_H is equal to 1.13%); a "low-risk" bank has the average *EDF* of the banks in the first quartile (EDF_L=0.38%). The first part of the Table shows that high-risk banks are smaller, more liquid and less capitalized. These features fit with the stylized fact that small banks are perceived as more risky by the market and need a larger buffer stock of securities because of their limited ability to raise external finance on the financial market. The lower degree of capitalization appears to be consistent with the higher riskiness of these banks. However, it is worth noting that the standard capital-to-asset ratio used here is not the best measure of the riskiness of bank portfolios, which would be captured more effectively by a measure of capital weighted by risk (Gambacorta and Mistrulli, 2004). Also, low-risk banks make relatively more loans.

3. Results

The results of the study are summarized in Table 3. The models have been estimated using the GMM estimator suggested by Arellano and Bond (1991), which ensures efficiency and consistency, provided the models are not subject to serial correlation of order two and the instruments used are valid (when assessed using the Sargan test). The first two columns present the results for our benchmark equation (1) using the clustered and estimated *EDFs*, which lead to very similar results.

[11] Only euro area banks that have at least four years of consecutive data are included in the sample. Banks that do not report positive figures for total assets, total loans and total capital for any given year are excluded. Investment banks, government financial agencies, special purpose financial institutions and foreign subsidiaries are excluded. Anomalies in loan growth rates are controlled for by checking for possible merger and acquisition activity related to full mergers from 1998 to 2005 in the Thomson SDC Platinum database.

Changes in economic activity have a positive and significant effect on loan demand (Kashyap, Stein and Wilcox, 1993). A 1% increase in nominal *GDP* causes a loan increase of 0.5-0.6%, depending on the model. The response of bank lending to a monetary policy shock has the expected negative sign (see coefficients for $\Delta i_{M\,t}$ and $\Delta i_{M\,t\text{-}1}$).

The riskiness of the credit portfolio has a negative effect on the banks' capacity to provide lending. Other factors being equal, higher loan-loss provisions (*LLP*) reduce profits, bank capital and, therefore, have negative consequences on the lending supply. A similar effect is detected for the *EDF*. The result suggests that banks' risk conditions matter for the supply of loans. As indicated, unlike other bank specific variables, which reflect historical accounting information, *EDF* is a forward looking variable. It reflects "market discipline", including the capability of banks to issue riskier uninsured funds (such as bonds or CDs), which can be easier for less risky banks, as they are more able to absorb future losses.[12] In this respect, there is evidence that euro area investors in banks' debt are quite sensitive to bank risk. More importantly this sensitivity seems to have been increasing in the aftermath of the introduction of the common currency (see Sironi, 2003). As a result, for banks perceived by the market as riskier, it would be difficult to issue uninsured debt or equity funds to finance further lending, for those banks would find it even more difficult to raise public equity in the markets to meet capital requirements (see Shin, 2008 and Stein, 1998).

The effects of liquidity (*LIQ*) and capital (*CAP*) on lending suggest that liquid and well-capitalized banks have more opportunities to expand their loan portfolios. Consistent with Ehrmann et al. (2003b), and contrary to the result for the US, the effect for size is negative, suggesting that small euro area banks are less affected by the adverse implications of informational frictions. This can be explained by the features of banking markets in the euro area: the low number of banking failures, presence of comprehensive deposit insurance schemes, network arrangements in groups, strong relationship lending between small banks and small firms (Ehrmann and Worms, 2004).

[12] For a review of the market discipline literature, see Borio et al. (2004) and Kaufman (2003). Seminal empirical evidence for the US already shows that lower capital levels are associated with higher prices for uninsured liabilities (Flannery and Sorescu, 1996).

As expected, the interaction terms between size, liquidity, capitalization and monetary policy have positive signs. In line with the bank lending channel literature, large, liquid and well-capitalized banks are better able to buffer their lending activity against shocks affecting the availability of external finance (Kashyap and Stein, 1995, 2000; Kishan and Opiela, 2000; Gambacorta and Mistrulli, 2004). The interaction term between EDF and monetary policy has the predicted negative sign, indicating that low-risk banks are more sheltered from the effects of monetary policy shocks.

We also analyse the effect of a monetary policy change on bank lending relative to the level of the intermediary's risk. We therefore estimate the impact on lending of a 1% increase in the short-term monetary rate using the coefficients reported in column II of Table 3. The results of the analysis are summarised in Figure 2, where we compare the effect of monetary policy change on lending for three kinds of financial intermediaries: the average bank for the whole sample (with *EDF*=0.73%), a low-risk bank (whose risk corresponds to the average for the first riskquartile, EDF_L=0.38%) and a high-risk bank (the average bank in the highest riskquartile, EDF_H=1.13%). The aim is not only to verify whether bank risk generates different insulation effects on loan supply, but also to obtain estimates of the size of these effects in relation to specific risk positions. For each bank, both the immediate pass-through (over the first year) and the long-term effect are considered.

Results indicate that, all other factors being equal, a 1% increase in the monetary policy indicator leads to a decline in lending for the average bank of 0.6% in the short term and -1.0% in the long run. Low-risk banks are on average far more insulated from the effects of a monetary policy shock than high-risk banks: the long-term effects are -0.4% and -1.8%, respectively.[13]

We also verify the importance of including bank risk with other standard bank-specific characteristics when analyzing the functioning of the bank lending channel. To do this, we include, in column III of Table 3, the baseline regression (1), excluding the *EDF* measure and its interaction with the interest rate change. In this case the liquidity indicator turns out

[13] Standard errors for the long-term effect have been approximated using the "delta method", which expands a function of a random variable with a one-step Taylor expansion (Rao, 1973).

not to show the expected sign and its interaction with monetary policy is no longer significant. This is probably due to the fact that this simplified regression suffers from omitted variable bias, due to the correlation between the *EDF* measure and the liquidity indicator. Moreover, the correlation between the *EDF* measure and liquidity changes over time: it is negative at the beginning of the sample (-0.2*) and becomes slightly positive at the end (0.1*). This is consistent with the idea that the liquidity indicator captures the probability of a bank default only in the first part of the sample when securitisation is limited. It also suggests that banks hold liquidity not only to decrease the risk of maturity transformation but also as a buffer against contingencies. With securitisation the determinants of liquidity dramatically change and probably relate more to the business model and less to risk management. Splitting the sample into two sub-periods (1999-2002 and 2003-2005), the coefficient of the interaction between the liquidity indicator and monetary policy is positive in the first period and not statistically different from zero in the second (3.28** and 0.38, respectively).

The effect of bank risk on lending supply may be different over the business cycle due to diverse perception of this risk. We have, therefore, introduced an additional interaction term by combining the *EDF* measure with the growth rate in nominal *GDP* in the baseline equation (1):[14]

$$\begin{aligned}
\Delta \ln(Loans)_{i,t} &= \alpha \Delta \ln(Loans)_{i,t-1} + \sum_{j=0}^{1} \delta_j \Delta \ln(GDPN)_{k,t-j} + \sum_{j=0}^{1} \beta_j \Delta i_{M\,t-j} + \sum_{j=0}^{1} \phi_j \Delta i_{M\,t-j} * EDF_{i,t-1} + \\
&+ \sum_{j=0}^{1} \sigma_j \Delta i_{M\,t-j} * SIZE_{i,t-1} + \sum_{j=0}^{1} \lambda_j \Delta i_{M\,t-j} * LIQ_{i,t-1} + \sum_{j=0}^{1} \chi_j \Delta i_{M\,t-j} * CAP_{i,t-1} + \kappa SIZE_{i,t-1} + \vartheta LIQ_{i,t-1} + \\
&+ \xi CAP_{i,t-1} + \tau LLP_{i,t-1} + \psi EDF_{i,t-1} + \sum_{j=0}^{1} \omega_j \Delta \ln(GDPN)_{kt-j} * EDF_{i,t-1} + \varepsilon_{i,t}
\end{aligned} \tag{3}$$

Equation (3) allows us to test for the possible presence of endogeneity between the business cycle and bank risk. The results reported in column IV of Table 3 indicate that the interaction term ϖ is positive and statistically significant, while other coefficients remain broadly unchanged. Hence, the negative effects of an increase in risk on bank loan supply is

[14] From now on, we consider in Table 3 only the models that use the estimated EDF. Results obtained using the clustered EDF are very similar and are not reported for the sake of brevity. These estimations are available from the authors upon request.

reduced in an expansionary phase and vice versa because the market perception of risk is typically reduced in good times and increased in bad times (Borio, Furfine and Lowe, 2001). There are several explanations for such observable fact: myopia and herd-like behavior (Minsky, 1975, Brunnermeier, 2009), perverse incentives in managerial remuneration schemes (Rajan, 2005), widespread use of Value-at-Risk methodologies for economic and regulatory capital purposes (Danielsson et al., 2001, 2004), pro-cyclicality of bank leverage (Adrian and Shin, 2008).[15]

In order to check if the different effects of monetary policy on banks with a diverse risk profile depend on business conditions, we add to the baseline model (1) the triple interaction between monetary policy, *GDP* and the *EDF* measure:

$$(\sum_{j=0}^{1} \varsigma_j \Delta i_{M\,t-j} * \Delta \ln(GDPN)_{k,t-j} * EDF_{i,t-1}).$$

Both the coefficients ς_0 and ς_1 turn out to be positive, with ς_1 significantly different from zero (ς_1=68.1, with a standard error of 19.5). This indicates that the greater exposure of high-risk bank loan portfolios to monetary policy shock is attenuated in good times, consistently with a the reduction of market perception of risk story as described above. All the other coefficients remained basically unchanged.[16]

The reliability of macro variable controls for loan demand shifts are checked by inserting a complete set of time dummies to obtain the following model:

$$\begin{aligned}\Delta \ln(Loans)_{i,t} &= \alpha \Delta \ln(Loans)_{i,t-1} + \theta_t + \sum_{j=0}^{1} \phi_j \Delta i_{M\,i,t-j} * EDF_{i,t-1} + \sum_{j=0}^{1} \sigma_j \Delta i_{M\,t-j} * SIZE_{i,t-1} + \\ &+ \sum_{j=0}^{1} \lambda_j \Delta i_{M\,t-j} * LIQ_{i,t-1} + \sum_{j=0}^{1} \chi_j \Delta i_{M\,t-j} * CAP_{i,t-1} + \kappa SIZE_{i,t-1} + \vartheta LIQ_{i,t-1} + \xi CAP_{i,t-1} + \\ &+ \tau LLP_{i,t-1} + \psi EDF_{i,t-1} + \varepsilon_{i,t}\end{aligned} \quad (4)$$

[15] For a discussion of these issues and a focus on reforms to improve financial stability see de Larosière *et al.* (2009), Volcker *et al.* (2009), Acharya and Richardson (2009), Panetta *et al.* (2009). The Financial Stability Forum (2009) provides a series of recommendations to reduce financial sector pro-cyclicality.

[16] These results are not reported in Table 3 for the sake of brevity.

This model completely eliminates time variation and tests whether the macro variables used in the baseline equation (nominal income and the monetary policy indicator) capture all the relevant time effects. Again, the estimated coefficients on the interaction terms do not vary significantly between the two kinds of model, thereby supporting the reliability of the cross-sectional evidence, as shown above (see column V in Table 3).

Two additional exercises (not reported in Table 3) were also performed. Namely, we introduced a set of geographical country dummies for each model, which are equal to 1 if the head office of the bank is in a given country and to zero if it is elsewhere. This allows controlling for possible country-specific institutional factors that could alter the results. In this case, the interactions between monetary policy and bank-specific characteristics remain basically unchanged.

We also considered a more complete model that also includes a securitisation indicator and its interaction with monetary policy.[17] This model tests whether our results could be affected by the large increase in securitisation activity in the period examined (see equation below):

$$\begin{aligned}\Delta \ln(Loans)_{i,t} &= \alpha \Delta \ln(Loans)_{i,t-1} + \sum_{j=0}^{1} \delta_j \Delta \ln(GDPN)_{k,t-j} + \sum_{j=0}^{1} \beta_j \Delta i_{M\,t-j} + \sum_{j=0}^{1} \phi_j \Delta i_{M\,t-j} * EDF_{i,t-1} \; + \\ &+ \sum_{j=0}^{1} \sigma_j \Delta i_{M\,t-j} * SIZE_{i,t-1} \; + \sum_{j=0}^{1} \lambda_j \Delta i_{M\,t-j} * LIQ_{i,t-1} + \sum_{j=0}^{1} \chi_j \Delta i_{M\,t-j} * CAP_{i,t-1} + \sum_{j=0}^{1} \delta_j \Delta i_{M\,t-j} * SEC_{i,t-1} \; + \\ &\kappa SIZE_{i,t-1} + \vartheta LIQ_{i,t-1} + \xi CAP_{i,t-1} + \varpi SEC_{i,t-1} + \tau LLP_{i,t-1} + \psi EDF_{i,t-1} + \varepsilon_{i,t}\end{aligned} \tag{5}$$

Even in this case no changes occurred to the interaction terms.

Finally, in order to check for potential biases caused by the use of estimated values for a substantial number of banks, we reran all the regressions reported in Table 3, restricting the sample to those banks (mainly large ones) for which the KMV *EDF*s are available. Also in this case, the interactions between monetary policy and bank-specific characteristics

[17] Following Altunbas, Gambacorta and Marqués (2009), the securitisation activity indicator has been constructed as $SEC_{i,t} = \dfrac{SL_{i,t}}{TA_{i,t-1}}$, where *SL* stands for the flow of securitised lending in year *t* and TA_{t-1} represents total assets at the end of the previous year. As for other bank-specific characteristics, the indicator has been normalised with respect to the average across all banks in the respective sample.

remain basically unchanged with the notable exception of size ($\Delta i_{M\,t-j} * SEC_{i,t-1}$) which turned out to be statistically non significant.

4. Conclusions

This paper analyses how bank risk influences bank credit supply and their ability to shelter that supply from the effects of monetary policy changes.

As a result of a very fast process of financial innovation (including the use of credit derivatives, the increase in securitisation activity and the new role of institutional investors), banks have been able to originate new loans and sell them on to the market, thereby obtaining additional liquidity and relaxing capital requirement constraints. We claim that, due to these changes, bank risk needs to be carefully considered together with other standard bank-specific characteristics when analyzing the functioning of the bank lending channel of monetary policy. Indeed focusing on size, liquidity and capitalization may be not be sufficient to accurately assess banks' ability to raise additional funds and supply additional loans. Indeed, the 2007-2008 credit turmoil has shown very clearly that the market's perception of risk is crucial in determining how banks can access capital or issue new bonds.

Using a large sample of European banks, we find that bank risk plays an important role in determining banks' loan supply and in sheltering it from the effects of monetary policy changes. Low-risk banks can better shield their lending from monetary shocks as they have better prospects and an easier access to uninsured fund raising. This is consistent with the "bank lending channel" hypothesis. Interestingly, the greater exposure of high-risk bank loan portfolios to monetary policy shock is attenuated in the expansionary phase, consistently with the hypothesis of a reduction in market perception of risk in good times.

Other interesting avenues remain open to further research. In particular, while this paper analyzes the link between bank risk and monetary policy effects, a reverse relationship may also hold. Namely, monetary policy may affect the risk-taking behaviour of banks and other financial intermediaries via asset prices and collateral values (Jimenez et al, 2008, Maddaloni et al., 2009). Moreover, if banks were to expect some kind of "insurance" from the Central Bank against asset price downturns, this could lead to moral hazard issues in the form of excessive risk taking on average over the business cycle. This calls for a growing need for the Central Bank to be able to anticipate excessive risk-taking by means of careful analysis of the evolution of a number of indicators, including risk premia and credit aggregates.

Tables and figures

Table 1

AVERAGE BANK FEATURES BY COUNTRY [(1)]

(percentages, millions of euros, expected default frequencies and number of banks)

	Lending	Size	Liquidity	Capital	Loan provisions	EDF (1)	Estimated EDF (2)	Securitisation	Number of banks
	(mean annual growth rate)	(EUR mill.)	(% total loans)	(% total assets)	(% total loans)			(% total assets)	
Austria	4.5	3,425	23.7	8.7	3.2	0.4	0.4	0.72	175
Belgium	3.9	23,981	10.8	7.6	1.4	0.1	0.3	0.02	57
Finland	7.4	18,723	11.6	9.4	0.2	0.2	0.2	0.01	4
France	5.2	10,460	13.9	10.0	1.5	0.7	0.8	1.80	250
Germany	2.1	4,699	24.8	5.7	1.0	1.0	0.9	1.66	1,665
Greece	38.4	7,345	13.5	14.2	1.2	1.4	1.3	0.24	8
Ireland	9.3	9,874	17.0	10.4	1.4	0.3	0.3	0.70	24
Italy	12.6	2,058	31.1	13.0	1.0	0.3	0.5	1.22	579
Luxembourg	5.8	6,110	45.2	6.8	4.5	1.2	1.0	5.69	91
Netherlands	6.8	18,803	24.1	9.3	2.7	0.8	1.4	19.36	31
Portugal	11.9	7,362	6.5	12.9	1.9	0.2	0.3	10.18	22
Spain	8.1	15,615	7.5	9.9	1.4	0.1	0.2	1.51	41
Euro area	5.0	5,400	24.9	7.9	1.3	0.5	0.7	1.93	2,948

Sources: Bankscope, Eurostat, KMV-Moody's.
Note: (1) Expected default frequency (EDF) figures are available for 134 banks, representing 52% of the total sample total assets. (2) Data for missing EDF have been estimated by mean of a regression analysis. As a first step, we have regressed the EDF on a number of bank balance sheet variables and country dummies (the latter have been inserted in order to capture specific institutional characteristics). In the second step, we have used the estimated coefficients to calculate the EDF for banks (mainly small ones) for which the KMV EDF are not available.

Table 2

BALANCE SHEET CHARACTERISTICS AND BANK RISK PROFILE (1)

Distribution by banks' risk (estimated EDF)	**Lending**	**Size**	**Liquidity**	**Capitalization**
	(mean annual growth rate)	(EUR mill.)	(% total assets)	(% total assets)
High-risk banks (EDF=1.13%) (a)	2.1	6,310	32.7	5.9
Low-risk banks (EDF=0.38%) (b)	10.2	8,224	25.4	11.4
Δ=(a)-(b)	-8.2	-1,914	7.3	-5.5

(1) A low-risk bank has an average ratio of the EDF in the first quartile of the distribution by bank risk; a high-risk bank an average EDF in the last quartile. Since the characteristics of each bank could change with time, percentiles have been calculated on mean values.

Table 3

REGRESSION RESULTS

Dependent variable: annual growth rate of lending (ΔL_t)	(I) Baseline Model (Cluster analysis)		(II) Baseline Model (Estimated EDF)		(III) Baseline Model (without EDF variables)		(IV) Banks' risk and the business cycle (Estimated EDF)		(V) Time dummies (Estimated EDF)	
	Coeff.	S.Error	Coeff.	S.Error	Coeff.	S.Error	Coeff.	S.Error	Coeff.	S.Error
ΔL_{t-1}	-0.156 ***	0.003	-0.140 ***	0.005	-0.093 ***	0.005	-0.069 ***	0.004	-0.051 ***	0.000
$\Delta GDPN_{t-1}$	0.578 ***	0.093	0.612 ***	0.165	1.479 ***	0.161	0.843 ***	0.121		
$SIZE_{t-1}$	-0.011 ***	0.001	-0.007 ***	0.001	-0.004 ***	0.001	-0.005 ***	0.001	-0.007 ***	0.001
LIQ_{t-1}	0.171 ***	0.008	0.153 ***	0.013	-0.019 *	0.012	0.306 ***	0.011	0.196 ***	0.016
CAP_{t-1}	0.006 ***	0.000	0.006 ***	0.000	0.004 ***	0.000	0.005 ***	0.000	0.014 ***	0.001
EDF_{t-1}	-0.051 ***	0.001	-0.020 ***	0.006			-0.127 ***	0.001	-0.046 ***	0.001
LLP_{t-1}	-0.113 ***	0.002	-0.109 ***	0.003	-0.006 ***	0.001	-0.123 ***	0.003	-0.119 ***	0.004
$\Delta i_{M\,t}$	-0.715 ***	0.105	-0.616 ***	0.184	-0.731 ***	0.199	-0.951 ***	0.154		
$\Delta i_{M\,t-1}$	-0.243 **	0.121	-0.491 ***	0.144	-1.358 ***	0.141	-0.376 ***	0.118		
$\Delta i_{M\,t}$*EDF_{t-1}	-2.090 ***	0.057	-2.343 ***	0.383			-1.202 ***	0.398	-0.855 ***	0.139
$\Delta i_{M\,t}$*$SIZE_{t-1}$	0.705 ***	0.048	0.453 ***	0.082	0.708 ***	0.084	0.342 ***	0.081	0.660 ***	0.091
$\Delta i_{M\,t}$*LIQ_{t-1}	1.321 ***	0.430	2.968 ***	0.932	1.385	0.917	3.423 ***	0.907	0.800 *	0.440
$\Delta i_{M\,t}$*CAP_{t-1}	0.181 ***	0.008	0.068 ***	0.021	0.066 ***	0.018	0.061 ***	0.020	0.236 ***	0.035
ΔGDP_t*EDF_{t-1}							1.190 ***	0.140		
Constant	0.034 ***	0.002	0.030 ***	0.003	0.013 ***	0.003	0.017 ***	0.003	0.035 ***	0.005
Time dummies	NO		NO		NO		NO		YES	
Sample period	1999-2005		1999-2005		1999-2005		1999-2005		1999-2005	
No. of banks, no. of observations	2,947	15,405	2,947	15,405	2,947	15,405	2,947	15,405	2,947	15,405
Sargan test (2nd step; pvalue)		0.383		0.140		0.130		0.200		0.290
MA(1), MA(2) (p-value)	0.000	0.123	0.000	0.150	0.000	0.051	0.000	0.137	0.000	0.147

The model is given by the following equation, which includes interaction terms that are the product of the monetary policy indicator and a bank specific characteristic:

$$\Delta \ln(Loans)_{i,t} = \alpha\Delta \ln(Loans)_{i,t-1} + \sum_{j=0}^{1}\delta_j \Delta \ln(GDPN)_{k\,t-j} + \sum_{j=0}^{1}\beta_j \Delta i_{M\,t-j} + \sum_{j=0}^{1}\phi_j \Delta i_{M\,t-j} * EDF_{i,t-1} + \sum_{j=0}^{1}\sigma_j \Delta i_{M\,t-j} * SIZE_{i,t-1} +$$

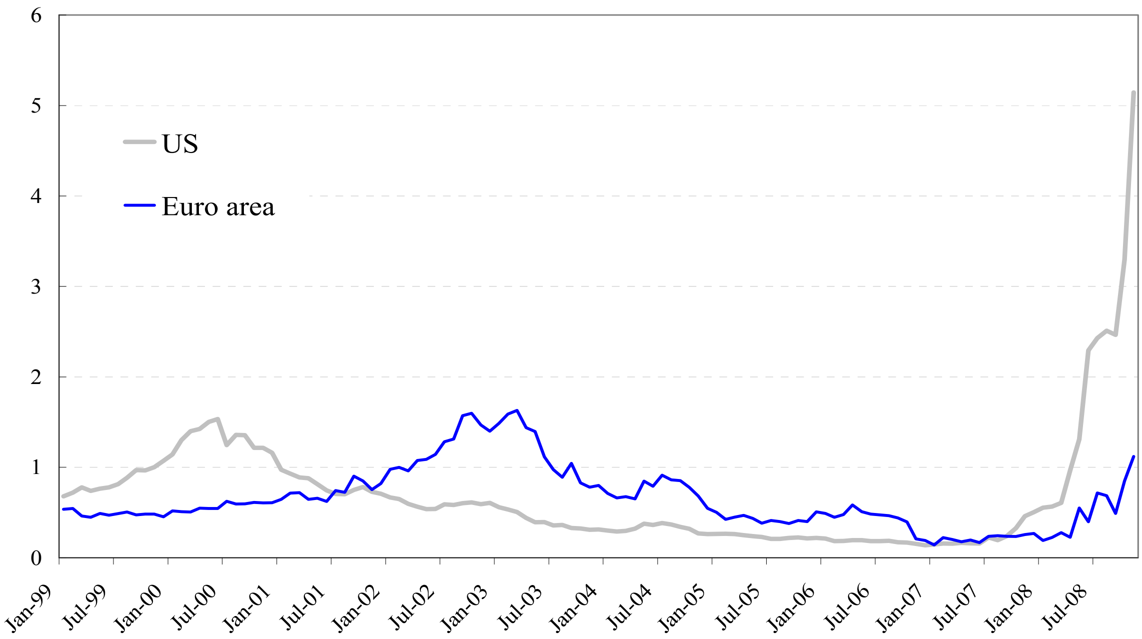
$$+ \sum_{j=0}^{1}\lambda_j \Delta i_{M\,t-j} * LIQ_{i,t-1} + \sum_{j=0}^{1}\chi_j \Delta i_{M\,t-j} * CAP_{i,t-1} + \kappa SIZE_{i,t-1} + \vartheta LIQ_{i,t-1} + \xi CAP_{i,t-1} + \tau LLP_{i,t-1} + \psi EDF_{i,t-1} + \varepsilon_{i,t}$$

with i =1,…, N and t =1, …, T and where: N = number of banks; L_{it}= loans in the balance sheet of bank i in quarter t; i_{Mt} = monetary policy indicator; $GDPN_{it}$ = nominal GDP; $SIZE_{it}$=log of total assets; LIQ_{it}=liquidity ratio; CAP_{it}=capital to asset ratio; LLP_{it}=loan loss provision over total assets; EDF_{it} =Expected default frequency. One lag has been introduced in order to obtain white noise residuals. The interactions terms and control variables that turned out not to be statistically significant in all the models have been removed from the table. The symbols *, **, and *** represent significance levels of 10 per cent, 5 per cent, and 1 per cent respectively.

Figure 1

EXPECTED DEFAULT FREQUENCY

(one year-ahead, averages)

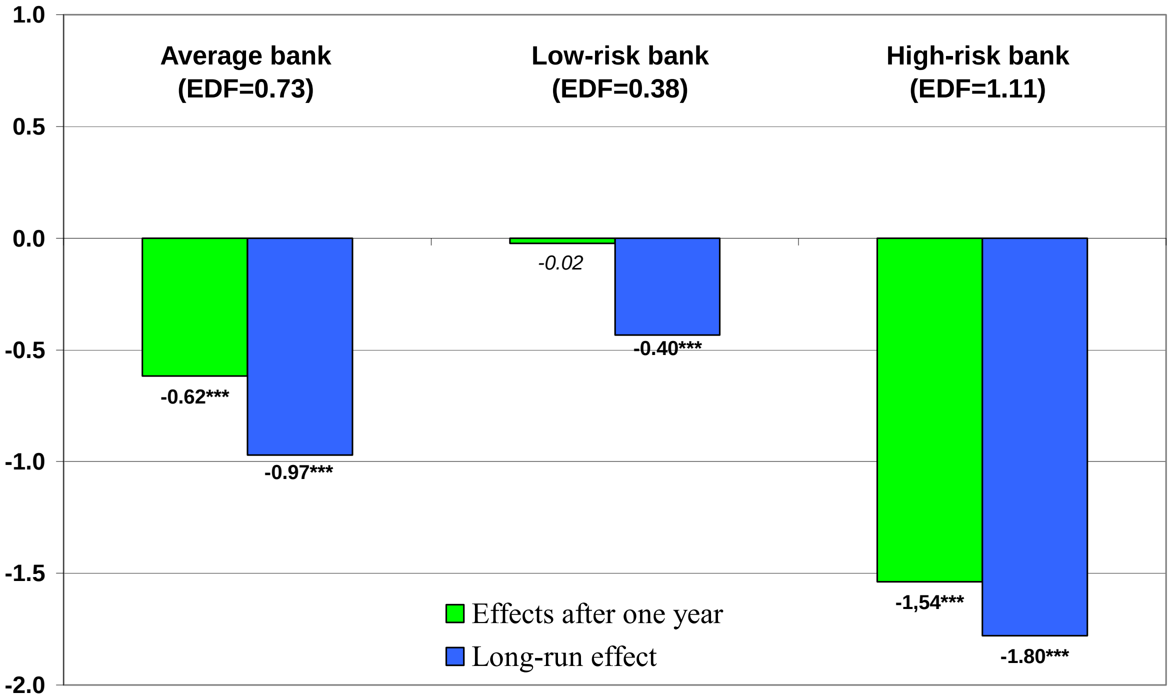


Source: Moody's KMV..

Figure 2

EFFECT OF A ONE PER CENT INCREASE OF THE MONETARY POLICY RATE ON BANK LENDING

(percentage points)



Note: We evaluate the effect of a one per cent increase of the short-term interest rate on bank lending considering banks with a different EDF (Expected Default Frequency). The coefficients are calculated on the base of the benchmark model in Table 3 with estimated EDF. The symbols *, **, and *** represent significance levels of 10 per cent, 5 per cent, and 1 per cent respectively.

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