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# Working Paper

# Bank heterogeneity and monetary policy transmission

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AUGUST 2009

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www.bankofgreece.gr

*Printed in Athens, Greece*
*at the Bank of Greece Printing Works.*


ISSN 1109-669

BANK HETEROGENEITY AND MONETARY POLICY TRANSMISSION

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**ABSTRACT**

The heterogeneity in the response of banks to a change in monetary policy is an important element in the transmission of this policy through banks. This paper examines the role of bank liquidity, capitalization and market power as internal factors influencing banks' reaction in terms of lending and risk-taking to monetary policy impulses. The ultimate impact of a monetary policy change on bank performance is also considered. The empirical analysis, using large panel datasets for the United States and the euro area, elucidates the sources of differences in the response of banks to changes in policy interest rates by disaggregating down to the individual bank level. This is achieved by the use of a Local GMM technique that also enables us to quantify the degree of heterogeneity in the transmission mechanism. It is argued that the extensive heterogeneity in banks' response identifies overlooked consequences of bank behavior and highlights potential monetary sources of the current financial distress.

*Keywords:* Monetary policy; Bank heterogeneity; Risk-taking; Bank performance
*JEL classification:* E44, E52, G21, C14

Acknowledgements: We are grateful to Efthymios Tsionas and Heather Gibson for helpful comments. We also thank seminar participants at the Athens University of Economics and Business, Bangor Business School and University of Maastricht for useful comments.

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## 1. Introduction

Understanding the transmission mechanism is crucial for monetary policy. In this respect, the special role of banking institutions in this mechanism has been studied extensively both at a theoretical and empirical level. The existing evidence shows that banks alter their lending behavior in specific ways following a change in monetary policy. But does this reaction involve only changes in lending behavior, as studied in the bulk of the literature? And do all banks in the market respond uniformly to monetary policy changes? This paper tries to answer these questions by analyzing empirically the heterogeneous response of US and euro area banks over the period 1994-2007 in terms of their lending and risk-taking decisions following a monetary policy change. The ultimate impact of a monetary policy change on bank performance is also considered. In our empirical strategy, the distributional effects of monetary policy stem from a number of bank characteristics, namely liquidity, capitalization and market power. We contend that the results obtained can improve our understanding of a number of bank-based channels of monetary transmission.

The aims of this paper are related to a long tradition in the literature of the transmission mechanism that accords banks a special role. In particular, Bernanke and Blinder (1988), among other proponents of the so-called bank lending channel, suggest that the effect of monetary policy on aggregate demand through interest rates may be enhanced by financial market imperfections and the existence of imperfect substitutability between loans and securities in bank portfolios and also as a means of borrowing for firms. One of the identification schemes for the lending channel involves estimation of reduced-form lending equations, where loan supply shifts are traced by using bank-level data on bank characteristics (see Kashyap and Stein, 2000; Ashcraft, 2006).

Yet, while the literature on the impact of monetary policy changes on bank lending is extensive, the corresponding impact on risk-taking and performance of banks is much less investigated. This seems odd since changes in policy rates can affect bank profitability, given that banks can borrow short at lower rates and use these funds to invest in longer-term projects at higher rates. In addition, a decrease in interest rates reduces agency costs or may cause banks to relax their lending standards, raising credit risk and thus non-performing loans (Matsuyama, 2007; Dell'Ariccia and Marquez, 2006). Therefore, especially as bank risk-taking is nowadays considered the *bête noire* of financial stability, central banks are increasingly concerned with avoiding price bubbles and minimizing incentives for banks to take on very high risks. However, it is only very recently that empirical methods have been proposed to examine the impact of monetary policy on bank risk-taking (Jimenez et al., 2007; Ioannidou et al., 2008). The results suggest that in a low interest-rate environment banks have incentives to take on higher risks in search for yield. Still, a major concern for the empirical analysis is the fact that banks respond quite heterogeneously to monetary policy changes and this may also have implications for their risk-taking and profitability, as in the case of lending.[1]

The heterogeneous behavior of banks originates from their different balance sheet characteristics. Theory on the bank lending channel identifies incentive mechanisms that work through the capital structure of banks, their liquidity levels and/or their size and argues that these mechanisms may play an important role in altering bank lending when there is a change in policy interest rates (see e.g. Diamond and Rajan, 2006; Bolton and Freixas, 2006). Empirical evidence on these propositions is restricted to testing whether the interaction of monetary policy variables with bank liquidity and capitalization levels is an important determinant of loan growth (see e.g. Kashyap and Stein, 2000; Gambacorta, 2005), while an analysis of the distributional effects of monetary policy on bank risk and profitability due to bank characteristics is missing. Furthermore, quantifying the level of heterogeneity may be important from a policy perspective as it improves our understanding of the way monetary policy may affect the real economy.

Using a method that identifies parameter heterogeneity at the bank level, this paper aims to (i) add to the limited literature that considers the effect of monetary policy on bank risk-taking and profitability and, more importantly, (ii) analyze and quantify the heterogeneous response of banks following a change in monetary policy. To carry out

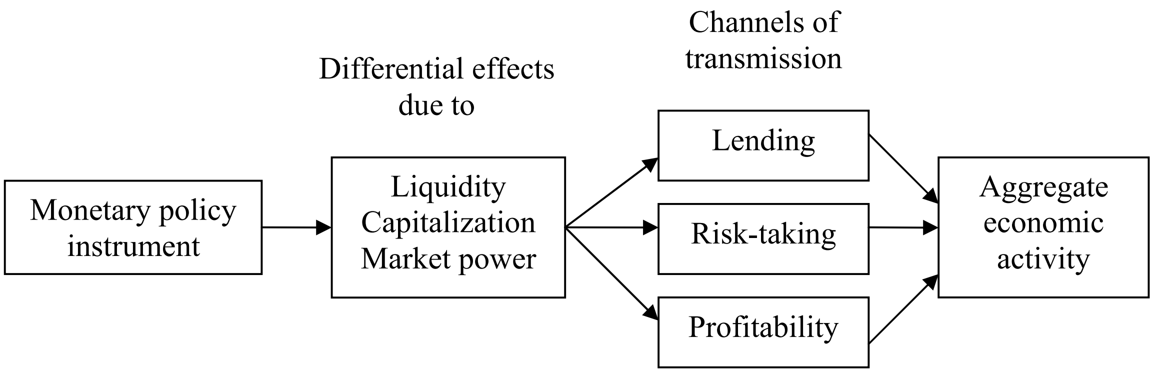
[1] In a recent line of work, Den Haan et al. (2007) examine the impact of monetary policy on the composition of bank loans and find differential responses between different types of loans. Even though this type of heterogeneity is different from the one examined in the present study, it still provides micro-level evidence that the sources of heterogeneity are important and multifaceted.

these tasks we resort to a local generalized method of moments (LGMM) technique. This technique, as developed by Lewbel (2007) and Tran and Tsionas (2008), considers the estimation of a dynamic panel data model, in which it allows the estimated coefficients to directly vary according to the smooth coefficient model (for an introduction to smoothing techniques, see Loader, 1999). The smooth coefficient model lets the marginal effect of the monetary policy variable be an unknown function of observable bank characteristics and hence introduces heterogeneity directly into the marginal effect of a monetary policy change.

Three different equations are considered, corresponding to a loan, a credit risk and a profitability model. The specification of the first equation follows previous studies of the bank lending channel, but its estimation also allows for quantification of heterogeneity. The other two equations provide novel results for the extent of heterogeneity in the impact of monetary policy on bank risk-taking and profitability. The bank characteristics serving as the smoothing variables are chosen to be liquidity, capitalization and market power. For the market power variable a number of reasons are provided below on why it is preferred to size, while its measurement is based on a recent technique that allows estimates to be obtained at the bank level. The choice of the other two bank characteristics is in line with the empirical studies of the lending channel discussed earlier. Given the above, we can explicitly illustrate the plan of the present analysis in terms of a chart (Chart 1). In particular, a change in monetary policy affects lending, risk-taking and profitability of banks in a non-uniform manner because of differential bank characteristics such as liquidity, capitalization and market power. This will have important implications for the real economy. However, analyzing the first two steps is probably already a lengthy task and therefore we leave the overall impact on real output as a future endeavor.

**Chart 1**

**Transmission of monetary policy through banks**




The rest of this paper is organized as follows. Section 2 provides some theoretical considerations about the relationships between monetary policy and (i) bank lending, (ii) bank risk-taking and (iii) bank profitability; it also analyzes the theory behind choosing liquidity, capitalization and market power as the smoothing variables. Section 3 describes the empirical model and the dataset and Section 4 presents the empirical results. Finally, Section 5 summarizes the main conclusions and provides some directions for future research.

## 2. Theoretical considerations

The literature that studies bank behavior and monetary policy is multifaceted and each of the approaches merits a discussion in its own right. Here we focus on the theoretical framework that assigns banks a special role in the monetary policy transmission mechanism via two channels, namely the bank lending and the risk-taking channel. The ultimate impact of monetary policy on banks' profitability is also considered. We further link this literature to the role that a number of bank characteristics such as liquidity, capitalization and market power may play in shaping a heterogeneous response of banks in their lending and risk-taking behavior as well as their performance following a change in monetary policy.

## 2.1. Channels of monetary policy transmission through banks

According to the traditional lending channel view, monetary policy affects bank loan supply and this in turn has an independent and significant effect on aggregate economic activity. In general, two conditions must be fulfilled for a bank lending channel to exist (Bernanke and Blinder, 1988). On the one hand, borrowers are not able to fully insulate their real spending from a decline in the availability of bank loans, i.e. bank loans are imperfect substitutes for other sources of finance. On the other hand, banks are not able to fully insulate their loan supply from a monetary policy-induced change in their reserves, i.e. there are no perfect substitutes for loans in bank portfolios. Both conditions have been subject to considerable debate in the literature. For instance, Romer and Romer (1990) suggested that if banks are able to obtain funds by tapping financial markets, monetary policy would affect banks only through changes in interest rates and, therefore, no bank lending channel would be at work.[2] Empirically the bank lending channel has been explored by many studies with mixed results (for the US case, see Kashyap and Stein, 2000; for the euro area, see the collection of papers in Angeloni et al., 2003). There is consensus, however, that in financial systems that are more market-based the higher degree of asset substitutability makes the bank lending channel less potent.

More recently, the notion of another channel, namely the risk-taking channel, has been put forward. Elements of this channel can be traced in Gibson (1997), who suggested that monetary policy has a greater effect on banks at times when their balance sheets have a riskier composition of assets. Several dimensions of how the risk-taking channel can work have been proposed. Matsuyama (2007) suggests that expansionary monetary policy reinforces the incentives of intermediaries to finance riskier projects. In a similar vein, Dell' Ariccia and Marquez (2006) and Rajan (2006) provide evidence that during lending booms loan quality deteriorates, as both lenders and borrowers are willing to take on higher risks. Furthermore, in addition to this effect working through the risk-taking of banks, it has been argued (see ECB, 2008) that the monetary policy of low interest rates followed in recent years, by affecting asset prices has led some institutional

[2] Unlike what is observed with the Bernanke and Blinder (1988) framework, in Romer and Romer (1990) bonds (securities issued outside the banking system) do not appear in banks' balance sheets and can be perfect substitutes for either certificates of deposit (a bank liability) or loans (a bank asset).

investors to invest increasingly in credit-related assets in search of higher yield. This has allowed banks to increasingly fund themselves by selling loans in the secondary market, thus potentially boosting the supply of new loans. However, this may also have contributed to a higher value of non-performing loans. Empirical evidence to support the above theoretical arguments is scant. Jimenez et al. (2007) use a sample of Spanish banks and a variety of duration models to find that lower short-term interest rates prior to loan origination result in banks granting more risky new loans. Ioannidou et al. (2008) examine the Bolivian case (it has a dollarized banking system) and find that a decrease in the US federal funds rate prior to loan origination raises the monthly probability of default on individual bank loans.

The potential relationship between a change in monetary policy and the risk-taking of banks has interesting implications for the sensitivity of bank profitability to changes in monetary conditions. A steeper term structure of interest rates as a result of looser monetary policy has often been associated with a situation in which banks are more profitable, as they can borrow short at rates that are lower than those at which they can lend long. Moreover, policy changes have implications for the profitability of lending activities because, as discussed above, they can alter the risk-taking of banks and the risk perception of borrowers and thus influence banks' return on credit. Even though there is a very limited body of work that explicitly focuses on the relationship between nominal interest rates and bank profitability, several studies use interest rates as control variables in profitability equations with mixed results (see e.g. Demirguc-Kunt and Huizinga, 1999).[3]

At this point we should probably note that the above discussion of monetary policy transmission through bank lending, risk-taking and profitability does not address the issue of possible heterogeneity in the response of banks following a change in monetary policy.

[3] In fact, there is an extensive literature on the determinants of bank risk-taking that includes control variables for the monetary conditions (see e.g. Jimenez et al., 2006 and references therein). However, the focus of these studies is not shifts in monetary policy and the general practice is to employ only the level (not the change) of various interest-rate variables. The same applies to studies that examine empirical models of bank performance (for a recent review of this literature, see Athanasoglou et al., 2008). An interesting study is that of Albetrazzi and Gambacorta (2006), suggesting that an increase in the money market rate has no separate effect on net interest income, but reduces significantly the return on assets, which is a measure that includes non-interest income.

This heterogeneity, which arises from differential bank balance sheet characteristics, such as liquidity, capitalization and size, has been examined by a number of theoretical and empirical studies mainly in the framework of the bank lending channel. Below we shall examine the implications of bank heterogeneity for the response of bank lending, risk-taking and ultimately profitability to monetary policy impulses.

### 2.2. Distributional effects of monetary policy

Recent empirical studies of the bank lending channel test for distributional effects of monetary policy among banks by controlling for the bank characteristics of liquidity, capitalization and size (see e.g. Gambacorta, 2005; papers in Angeloni et al., 2003). The main contribution of this literature is the use of bank characteristics and hence of panel data to solve the identification problem of the lending channel, i.e. distinguishing between shifts in loan demand and shifts in loan supply.[4] Yet, the heterogeneous impact of a change in monetary policy on bank risk-taking and profitability has not been examined so far. In addition, the extent of heterogeneity on any of the bank characteristics mentioned above has not been systematically analyzed.

In the present study we depart from using size as the third bank balance sheet characteristic influencing banks' reaction to changes in policy rates. The main argument for the use of bank size is that one would expect the largest banks to have an easier time raising uninsured finance, which would make their lending less dependent on monetary policy shocks irrespective of other bank characteristics (Kashyap and Stein, 1997; 2000). However, this implicitly suggests that certain banks have market power in raising uninsured finance, something that may or may not be the result of size. This feature naturally is carried over to the asset side of bank balance sheets causing deviations from perfectly competitive behavior. It is also noteworthy that Lensink and Sterken (2002), in editing a special issue of the Journal of Banking and Finance, suggested that future work should identify whether bank competition plays an important role in the monetary transmission mechanism. Therefore, in this paper we use the market power of banks as a critical element in the pass-through of monetary policy.

[4] This strategy relies on the hypothesis that bank characteristics influence only loan supply movements while loan demand is independent of these characteristics.

In the context of the bank lending channel, Kashyap and Stein (2000), among others, suggest that the impact of monetary policy on lending behavior is stronger for banks with less liquid balance sheets. In a recent theoretical study, Diamond and Rajan (2006) suggest a different means through which the lending channel may be at work. In particular, they suggest that an expansionary open market operation increases financial liquidity, which alleviates the real liquidity pressures on banks, which then allows them to fund longer-term projects and therefore enhance aggregate economic activity. This implies that the effect of monetary policy is further amplified when banks hold a higher value of liquid assets (especially in a low interest rate environment) because banks will use the extra liquidity in their portfolios to make new loans. Thereby one should expect the lending and profitability of these banks to increase above the average.

Turning to the effect that liquidity may have on the relationship between policy rates and bank risk-taking, we would expect banks with higher levels of liquid assets, which reflect higher risk aversion, to have lower incentives to take on more credit risk after an expansionary monetary policy. To the extent that this holds, the impact of monetary policy on bank risk is weakened. Also, if banks with relatively illiquid portfolios face higher risks, due for example to a prolonged period of low interest rates, then these banks may be less able to use part of their liquidity in less risky activities, thus reducing credit risk. However, it may also be the case that excess liquidity of banks in a low interest rate environment provides incentives for higher risk-taking if banks feel particularly safe (a moral hazard problem). This situation, although not representing typical bank behavior, can lead to a deterioration of bank asset quality and a sharp increase in bank risk. On the basis of the above arguments, we may formulate our first hypothesis as follows:

*Hypothesis 1:* We expect that the impact of monetary policy on bank lending, risk-taking and profitability will decrease (increase) with higher (lower) levels of bank liquidity.

In the context of the credit channel, Peek and Rosegren (1995), among others, suggest that, following a contraction in the money growth rate, poorly capitalized banks reduce their lending much more than better capitalized banks, since they usually have less access to markets for uninsured funding. Bolton and Freixas (2006), on the other hand, emphasize the imperfect substitutability of equity capital with other sources of funds for banks and highlight that there is a secondary bank lending channel operating through the bank equity-capital market, even when banks have perfect access to non-deposit liabilities or to the bond market. In particular, a monetary tightening reduces banks' incentives to increase their equity-capital base. If capital regulations are in place, banks are likely to react by decreasing the supply of loans, thus magnifying the contractionary effect of monetary policy.

This is also related to the theory of credit rationing (that originates in Stiglitz and Weiss, 1981), which suggests that the underlying factor of a credit crunch may be a change in the risk perception of banks, which in turn could be triggered by a change in monetary policy, a shortage of bank capital or both. In these situations ─ besides the direct impact of monetary policy on bank lending ─ bank risk-taking and ultimately profitability (through the channels discussed in Section 2.1) are also likely to be heterogeneously affected on the basis of differential levels of equity capital. Especially for risk, more capitalized banks are probably more able to buffer excess loan losses following an expansionary monetary policy, and therefore reduce credit risk, than less capitalized banks. Our second hypothesis then becomes:

*Hypothesis 2*: We expect that the impact of monetary policy on bank lending, risk-taking and profitability will decrease (increase) with higher (lower) levels of bank equity capital.

Having argued that liquidity and capitalization of banks potentially play an important role as regards the extent to which bank balance sheets are sensitive to monetary policy, we turn now to the possible effect of bank market power. Baglioni (2007) investigated a model of the banking industry under both monopolistic competition and a Cournot oligopoly and suggested that under the former the effect of monetary policy is amplified, while under the latter it is weakened. Note that empirical evidence from the European banking industry suggests that most banking systems in euro area countries are characterized by monopolistic competition (see Bikker and Haaf, 2002; Claessens and Laeven, 2004), while some banking systems of newly acceded EU countries feature oligopolistic behavior (see Brissimis et al., 2008).

If, following an expansionary monetary policy, banks on average take on higher credit risk in the search for yield, then banks with high market power will probably not engage in such activities because they already extract rents from the fact that they have market power. However, banks with market power may also see their revenue rise from other activities besides lending (which corresponds to credit risk) and therefore their profits may rise above average following a fall in policy interest rates.[5] On the basis of the above, our third testable hypothesis is formulated along the following lines:

*Hypothesis 3*: We expect that the impact of monetary policy on bank lending, risk-taking and profitability will decrease (increase) with higher (lower) levels of bank market power.

To gain insights into these issues, we build on recent econometric techniques so as to uncover patterns of heterogeneity based on individual bank behavior, thus greatly improving our understanding of the transmission mechanism through banks. Bearing these issues in mind, we now proceed to the discussion of the empirical framework.

## 3. Methodology for estimating the effects of heterogeneity

### 3.1. The empirical model

The starting point in the estimation procedure is the setup of the econometric model. We opt to estimate lending, risk and profitability equations with a view to identifying the degree of heterogeneity in banks' response following a monetary policy change. For this purpose we employ a smoothing technique, and in particular the Local GMM (LGMM) method proposed by Tran and Tsionas (2008).[6]

The main advantage of this approach is that like all smoothing techniques it relaxes the assumption that banks share the same incentives, technology and balance sheet characteristics, and therefore that they have the same response to a shift in monetary

[5] An example of this is higher than average increases in revenue from capital markets or off-balance sheet activities.
[6] In fact, LGMM was first proposed, but not implemented, by Lewbel (2007).

policy. This is accomplished because the parameters of the estimated equations are made observation-specific through localization. LGMM allows the marginal effect of the monetary policy variable to be a function of observable covariates (in particular liquidity, capitalization, and market power) and hence it introduces heterogeneity directly into the marginal effect.[7] By examining the coefficients obtained one can in fact estimate the extent of heterogeneity using simple statistical measures, like the standard deviation or the variance of the coefficients.

In addition, this strategy presents at least three other advantages, compared to conventional methods. First, and related to the above, disaggregating the impact of monetary policy at the bank level allows for the inclusion of different types of banks, and also of banks across different banking sectors in the sample (no uniform production technology is assumed). This advantage is remarkable because the researcher and the policy-maker can gain a more thorough insight into the international banking sector through a single estimation model. Second, LGMM allows for non-parametric heteroskedasticity in the error term, which is important because parameter estimates in panel data models depend critically on the nature of heteroskedasticity. By accommodating heteroskedasticity of an unknown form, the quite restrictive normality assumption is effectively removed and statistical efficiency is improved. Third, this technique (unlike other smoothing techniques) embeds all the properties of the dynamic panel parametric GMM methods, as it allows for a dynamic adjustment of the dependent variable within a panel data context; and it is well-known that bank variables like lending, risk-taking and profitability persist to a large extent (Berger et al., 2000).[8]

The equations to be estimated are directly obtained from the literature. The first equation considers the response of lending to monetary policy and follows from the bank lending channel literature (see e.g. Kashyap and Stein, 2000). The second considers the impact of monetary policy on bank risk-taking and follows from the literature that explains credit risk, which however does not emphasize the effect of monetary policy

[7] This also completely solves the problem of multicollinearity that is present when using interaction terms in parametric models.
[8] For more on these issues and on the derivation of the asymptotic properties of the LGMM estimator, see Tran and Tsionas (2008).

(see e.g. Boyd et al., 2006). Finally, the third equation considers a simple profitability/performance equation (see e.g. Brissimis et al., 2008). Thus, the estimated equations are of the following form:

$$\Delta(l_{it}) = \alpha_1(z_{it})\Delta(l_{i,t-1}) + \alpha_2(z_{it})\Delta M_t + u_{it}^1 \quad (1)$$

$$\Delta(r_{it}) = \beta_1(z_{it})\Delta(r_{i,t-1}) + \beta_2(z_{it})\Delta M_t + u_{it}^2 \quad (2)$$

$$\Delta(\Pi_{it}) = \gamma_1(z_{it})\Delta(\Pi_{i,t-1}) + \gamma_2(z_{it})\Delta M_t + u_{it}^3 \quad (3)$$

where $\Delta$ denotes change, $l$ is lending of bank $i$ in period $t$ (in logarithmic terms), $r$ is a measure of bank risk, $\Pi$ is a measure of bank profitability, $z$ are the bank characteristics of liquidity, capitalization and market power (used alternatively), $M$ is the monetary policy variable and $u$ is the error term. Hence, the empirical model suggests that in all three equations the dynamic adjustment coefficient and the coefficient on the monetary policy variable may vary directly with the bank's liquidity, capitalization, and market power. In other words, the empirical framework follows closely the theoretical considerations discussed in Section 2.

At this point, a number of critical remarks can be made on the estimation procedure. First, it is well-known that non-parametric techniques have to be applied to larger datasets to avoid the so-called "curse of dimensionality". Fortunately, this is not an issue for micro-level studies like the present one, where datasets are quite large. Second, the literature on the non-parametric methods has not yet reached a consensus on the selection of an optimal bandwidth.[9] Recent studies on local techniques (see e.g. Kumbhakar et al., 2007; Park et al., 2007) rely on cross-validation, which involves estimating the average squared prediction or estimation error. Yet, cross-validation methods have been criticized as giving results that are too variable and unreliable (see e.g. Ruppert et al., 1995), and this literature mainly proposes simple plug-in methods. Since cross-validation is a more formal method, it is the main approach followed here. However, robustness of the results is ensured by examining the consequences of small changes in the bandwidth.

[9] A large bandwidth leads to an oversmoothed curve that may inadequately model or completely miss important features of parameter heterogeneity, while a small bandwidth may undersmooth the curve resulting in a very noisy fit.

### 3.2. Data and variables

Large panel datasets for the United States and the euro area (the first 12 EU countries that participated in the euro area) banking sectors are used for the empirical analysis. Bank-level data is obtained from the Bankscope database and covers the period 1994-2007. The final unbalanced datasets are built by applying two selection criteria. First, all types of banks that take deposits (commercial, savings, cooperative and bank holding companies) are included in the sample. Investment banks are not included because they do not take deposits and therefore do not fall within the theoretical discussion provided above. Second, an outlier rule is applied to the main bank-level variables so as to disregard the 2% of both edges of their distribution. This also deletes all banks for which data on the main variables of this study is not available. The final sample consists of 5873 observations for the United States and 6133 observations for the euro area, some 12006 observations in total.[10] The relatively large time dimension (14 years) and the cross-sectional dimension of the country-specific variables allows for considerable variability in the monetary policy variable.

Note here that Bankscope data is annual and an immediate question arises as to why annual data is used to study monetary policy. Most work on the lending channel (see e.g. Kashyap and Stein, 2000) employed quarterly data to study the lending behavior of banks. This is less frequent concerning the risk and profitability equations, for which most of existing literature employs the Bankscope database. In two recent papers, Gambacorta (2005) and Ashcraft (2006) point out that results obtained from lending equations are robust to the use of annual data. In addition, the latter study provides a thorough discussion of the reasons why this holds and therefore it enhances our confidence that a comparison between the US and the euro area on the basis of annual data is worthwhile (in general quarterly bank data on euro area countries are unavailable).

Table 1 provides descriptive statistics for all the variables of the study that are analyzed below. $l$ is measured by total customer loans in real terms and $r$ by the ratio of problem loans to total loans. Let us briefly comment here on our choice for the risk

[10] Note again that the LGMM method allows for the inclusion of banks from different countries in a single estimation since it makes no parametric assumptions on the functional form or the error term.

variable. $r$ represents credit risk, i.e. the potential variation in net income and the market value of equity resulting from a non-payment or delayed payment by borrowers. Whenever a bank grants a loan, it assumes the risk that the borrower will default, that is, he will not repay the principal and interest on a timely basis and thus our measure reflects the quality of bank loans. Since a portion of non-performing loans will probably result in losses for the bank, a high value for this ratio is undesirable. Credit risk represents the major cause of most bank failures and is directly related to the theoretical discussion of Section 2 because it refers to the risk of lending. Thus, it is the one favored in the present analysis.[11]

Finally, bank performance $\Pi$ is measured by the ratio of net profits to total assets (i.e. the return on assets, ROA) so as to capture the income generated from both traditional and non-traditional banking activities. We also employed the net interest margin (i.e. the ratio of net interest income to total assets) and we found qualitatively similar results. The results based on ROA are reported because many banks engage in off-balance sheet activities, which generate non-interest income.[12] The descriptive statistics on the dependent variables show that the average euro area bank is somewhat larger in terms of loans (this is expected given the large number of small local US banks), with the average values for $r$ and $\Pi$ being very close in the two samples.

Following Bernanke and Blinder (1992) and Ashcraft (2006), we utilize the federal funds rate as a measure of monetary policy in the US. For the euro area, we utilize the ECB policy rate from 1999 onwards (2001 for Greece) and before 1999 the official refinancing operation rate for each country separately. Data on these variables show that the average rate has been lower in the euro area; however, the range of their values is approximately equal. In all three equations, we control for economic conditions by including real GDP growth (denoted by GDP) and for financial depth by including the

[11] Additionally we have proxied risk by the Z-index, which represents a more universal measure of bank risk, the ratio of provisions for loan losses to total loans, and the ratio of reserves for loan losses to total loans. All ratios are highly correlated and the estimation results were found to be similar.

[12] In terms of relative importance, interest revenue remains the main component of total bank income. However, its importance has gradually declined over the past two decades as banks sought to diversify their business away from the traditional banking activities. This is the main reason behind our inclination toward ROA, a measure that encompasses both traditional and non-traditional banking activities. Also note that some recent studies estimate bank performance using frontier methods. The aim of this study is to keep this part of the analysis as simple as possible and therefore only accounting ratios are employed.

ratio of stock market capitalization to GDP (CAP). The last measure is obtained from the Beck et al. (2000) database and the very few missing observations at the end of the sample are constructed by own calculations.[13] Given the virtues of the LGMM method in terms of relaxation of the assumptions regarding differences between banks, we do not need to control for other bank characteristics. The descriptive statistics reflect the fact that the US economy has been growing faster during the sample period and that stock market capitalization is double that of euro area countries (even though some euro area countries record very high values for this index).

Turning to the smoothing variables *z*, we use the ratio of the book value of equity to total assets to measure the capitalization of banks and the ratio of liquid assets to total assets to measure liquidity (see also Gambacorta, 2005; studies in Angeloni et al., 2003). The values of the descriptive statistics for these variables are very close in the two samples. As regards the market power variable, making a good choice is a more complicated issue. The literature suggests using either structural measures of competition (i.e. concentration ratios or the Herfindahl index), which, however, have well-known limitations in measuring competitive conditions (see Northcott, 2004 and references therein), or non-structural measures (i.e. the Lerner index or the H-statistic of Panzar and Rosse, 1987).

The non-structural measures reflect the competitive conditions at the industry level, but disaggregating at the bank level is difficult because one has to assume a common production technology for all banks (and across US states or euro area countries in the present study). Recently, Brissimis and Delis (2008) proposed estimating the H-statistic at the bank-level using a simple local regression. Once again, the idea is that local regression allows for observation-specific H-statistics through localization, and in doing so it also relaxes a number of restrictive assumptions regarding the properties of the production function of banks (like in the LGMM case, technology is allowed to differ at a bank-level and in time). Note that the H-statistic does not necessitate defining the location of the market *a priori*, which implies that the potential bias caused by the misspecification of market boundaries is avoided; hence for a bank that operates in more

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[13] We have not been able to compute stock market capitalization for Ireland for the years 1994-1995 because the required data is not available.

than one market, the H-statistic will reflect the average of the bank's conduct in each market. In the words of Shaffer (2004), "because the core model involves only firm-level data, no specific market definition appears in the revenue equation. This is a huge advantage in a sample that contains countries spanning less than a market (euro area countries) up to thousands of local markets (as in the US)." Similarly, inclusion of different types of banks is not an issue (see Claessens and Laeven, 2004; Shaffer, 2004). Finally, note that estimation at the bank level of an alternative to the H-statistic, namely the Lerner index, involves the quite restrictive assumption of imposing a constant marginal cost across banks. This is the main reason why we prefer the H-statistic to the Lerner index in the present analysis.[14]

The empirical Panzar-Rosse model involves estimation of the following revenue equation

$$\ln tr_{it} = \delta_0 + \delta_1 \ln w_{1,it} + \delta_2 \ln w_{2,it} + \delta_3 \ln w_{3,it} + \delta_4 \ln k_{it} + \varepsilon_{it} \qquad (4)$$

where $w_1, w_2$ and $w_3$ are the input prices of deposits, labor and fixed capital respectively, k is a vector of control variables and the H-statistic is H = $\delta_1 + \delta_2 + \delta_3$. In this framework, the H-statistic takes values between -1 and 1; the more negative or close to 0 the H-statistic is, the larger is market power, while the closer the H-statistic is to unity, the more competitive is the bank (see e.g. Barajas et al., 2000). Values between 0 and 1 reflect monopolistically competitive practices and such values correspond to most of the findings of the existing literature on the measurement of bank competition at the industry level (Claessens and Laeven, 2004). In Eq. (4) *tr* is the ratio of gross interest revenue to total assets, $w_1$ is the ratio of interest expenses to total deposits and money market funding, $w_2$ is the ratio of personnel expenses to total assets, $w_3$ is the ratio of other operating and administrative expenses to total fixed assets and *k* corresponds to the ratio of loans to total assets (proxy for the quality of assets) and the logarithm of real total assets (proxy for bank size).[15]

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[14] Note, however, that the Lerner index maps more robustly into a range of oligopoly solutions than the H-statistic, thus one may think that there is a tradeoff between using any of the two methods.

[15] A number of issues come up in estimating the Panzar-Rosse model. First, due to lack of data on total employees for many observations (especially for the early period of our sample), we do not express the unit cost of labor in terms of total employees but in terms of total assets (see Claessens and Laeven, 2004,

Estimation of Eq. (4) is carried out both separately for the US and the euro area and for the combined sample, and the average results are surprisingly close (within the 3% range). This verifies Shaffer's (2004) comments regarding the power of the H-statistic to accommodate location effects and the suitability of the local regression method. The bandwidth required for the local regression method has been calculated by simple cross validation and was found to be equal to 0.673 and 0.707 for the US and euro area samples, respectively, and 0.680 for the combined sample; also, the type of the kernel is the Epanechnikov kernel (for more on these issues see Loader, 1999).[16] Since no significant difference was found in the results, the estimates of market power used in the subsequent analysis are the ones obtained from the combined sample. Table 1 provides some descriptive statistics for the results and Figure 1 shows the frequency distribution of estimated H-statistics. The results reveal that banks in the US behave a little more competitively on average and that the equations for euro area banks have a wider distribution of coefficients. Even though differences are not that large, this finding probably shows that certain euro area banking sectors remain quite segmented, a conclusion which is supported by the fact that M&A activity in the euro area has taken place mainly within each country and not so much across borders. It is interesting to note that some euro area banks are characterized by monopolistic behavior with values of coefficients close to 0 or negative, while such coefficients are very rarely found for the US banking sector. Bikker and Haaf (2002) and Claessens and Laeven (2004) find such

among others). Second, since the H-statistic will be used to smooth the monetary policy variable, we do not include as control variables the liquidity and capitalization ratios, while we verified that these variables exhibit low correlation with the ratio of loans to assets (that is actually employed). Third, a number of alternative specifications were formed on the basis of including additional control variables (i.e. the cost to income ratio as a proxy for bank efficiency, the ratio of interbank deposits to total customer and short-term funding as a proxy for differences in the deposit mix and/or the ratio of cash and due from depository institutions to total deposits as a proxy for correspondent bank activities). Changes in the results were not significant and hence the study proceeded with the simple formulation described in the main text. Finally, the PR model should be estimated on observations that are in long-run equilibrium. The empirical test for equilibrium is justified on the grounds that competitive capital markets will equalize the risk-adjusted rate of return across banks, so that (in equilibrium) the rate of return should not be statistically correlated with input prices. To test for equilibrium, one can calculate another H-statistic ($H_n$) using the rate of return (return on assets), instead of total revenue, as the dependent variable in the regression equation. In this framework, $H_n$=0 indicates that banking systems are in equilibrium. Following exactly the same procedure (in terms of variables and estimation method) with Claessens and Laeven (2004) among others, we find that both the Unites States and the euro area are "in equilibrium". These results are not reported but are available on request.

[16] Specifying a different type of kernel produces negligible differences in the results.

differences between euro area banking sectors when estimating the Panzar-Rosse model at the country level, and this enhances confidence in the results of this paper. Hence, and in line with the theoretical arguments provided in Section 2, these bank-level market power estimates are used as the third smoothing variable $z$ in the main empirical analysis that follows.

# 4. Empirical results

## 4.1. Practical issues in the estimation procedure and general results

The precision of the LGMM estimator depends on the choice of the instruments. We follow Newey (1990) and Tran and Tsionas (2008) in using the kernel estimates of $\{E(dl_{i,t-1}) \mid z_{i,t-1}), E(dl_{i,t-2} \mid z_{i,t-2})\}$ as instruments for $dl_{i,t-1}$, while $M$ is instrumented on its first two lags. The same rule applies for the risk and profitability equations. Again the Epanechnikov kernel is used and the optimal bandwidth was estimated by univariate cross validation and found to be 0.585 for the lending equation, 0.604 for the credit risk equation and 0.621 for the profitability equation. All models include time effects. Finally, it is verified that the results are robust to both the choice of the bandwidth[17] and to adding a third lag in the instruments (however in this case precision deteriorated as the mean square error of the regressions increased).

Inference on the effects of monetary policy is in part based on confidence intervals, which can be constructed using bootstrapping. In our analysis we use nonparametric bootstrap percentile confidence intervals to infer the significance level of the effects. The intervals are estimated at the 95% level using the mean value of our sample (we also used the median with similar results) and the asymptotic normal approximation. We performed 1000 bootstrap replications. If the confidence interval fails to include 0 in the estimations provided below, then for the average estimate the p-value is less than or equal to 0.05 and the effect is statistically significant. Note, however, that in the present study we focus on

[17] Numerous values for the bandwidth have been examined in the ± 20% area of the optimal bandwidth. Indeed, when there is large differentiation from the optimal bandwidth, the resulting distributions reveal oversmoothing or undersmoothing. Hence, we feel that at least in the present panel, cross validation is suitable.

the heterogeneity of the coefficients and not so much on the average effects. Therefore, if bank behavior is far from the average and distributions of coefficients do not approximate the normal distribution, then more than one peak effect may be present in the same relationship. We will elaborate on this point below.

The results on the observation-specific coefficient estimates for the monetary policy variable obtained by LGMM are presented in Figures 2, 3 and 4. In the notes to these figures we report our estimates for the confidence intervals. Only the impact of the monetary policy variable when smoothed by liquidity, capitalization or market power is reported, but the relevant figures illustrating the distribution of the coefficients for the control variables are available and can be provided upon request. In the figures, bars symbolize the frequency of the coefficient estimates and two lines are added, one representing the actual distribution and one the normal distribution. Average values for the mean and the standard deviation of the coefficients for the United States, the euro area and the full sample are provided in Table 2, which also reports the estimated confidence intervals. In addition, the time pattern of the standard deviation (i.e. the degree of heterogeneity) for the combined sample is given in Table 3.

The coefficients on the lagged dependent variable when capitalization is used as the smoothing variable are also reported in the first graph of each figure to reflect the extent of the dynamic nature of the panel. Note that a value of these coefficients between 0 and 1 implies that there is persistence in the dependent variables, but their values will eventually return to their normal (average) level. A value close to 0 means that the dependent variable is characterized by a high speed of adjustment, and a value close to 1 implies very slow adjustment (see Nerlove, 2002). In general, all models indicate significant persistence (i.e. the confidence intervals do not include 0). The coefficients of the lagged variable in the lending equation display a bimodal distribution, with the mode on the left being the dominant one. This distribution of the coefficients is very wide reflecting significant differentiation (heterogeneity) in the level of persistence among banks, a result consistent with variations in the degree of relationship-lending (see Petersen and Rajan, 1995). Interestingly, the distribution of the lagged dependent variable in the credit risk and profitability equations approximates the normal distribution quite well; however, in the credit risk equation heterogeneity seems high as the coefficients take values from a low close to 0 (no persistence) to a high of 0.8-1 (very high persistence). The coefficients on the lagged dependent variable when smoothed by liquidity and market power look very similar and are not presented here.

The findings reported in Table 2 confirm that a bank lending channel is less operative in the United States than in the euro area (smaller average coefficient on the monetary policy variable), a result that backs up the argument suggesting that increased market-based finance (such as exists in the Anglo-Saxon type financial system) reduces the potency of the bank lending channel. However, in both the United States and the euro area the impact of the interest rate variable is statistically significant across all specifications. Notably, the average coefficient decreases over time regardless of the choice of the smoothing variable (-0.407 in 1994 and -0.262 in 2007 when $z$ is capitalization). It is expected that, ceteris paribus, the continuous move toward Anglo-Saxon type of financial systems will further decrease this impact in the long term. Differences in the impact of monetary policy on risk-taking behavior and the profitability of banks are not particularly large between the two sub-samples, although the standard deviation of the distributions presents some differences. The latter result is especially true when market power is used as the smoothing variable, indicating considerable behavioral differences between some of the more segmented euro area banking sectors and the relatively more competitive US market.

Table 3 presents the course over time of the standard deviation of the coefficients obtained from the nine estimated equations. The information that can be extracted from these figures is that there is an apparent increase in the standard deviation (or equally in the variance) of the coefficients during episodes of distress (see figures for 2001 or figures for 2007), while low values occur during episodes of relative stability. This is to be expected as, in general, distress causes different behaviors and stability enhances uniformity.

### 4.2. Monetary policy and lending

The average values reported for the coefficients of the interest rate variable in the lending equations are negative and statistically significant, regardless of which smoothing variable is used. Yet, the standard deviation of estimates reported in Tables 2 and 3 and the distributions of estimates presented in Figure 2 reflect the wide heterogeneity in the bank-level coefficients. For example, Figure 2c shows that when *M* is smoothed by the capitalization of banks, the impact of monetary policy on loan growth ranges substantially and that the distribution of this coefficient is trimodal (with modes at about -0.5, -0.3 and -0.15). This implies that monetary policy changes cause a very different response of banks on the basis of their capital structures, with more capitalized banks (i.e. those on the right end of the distribution) responding less to monetary policy changes. This is also reflected in the high variance of the coefficients (see Tables 2 and 3), while this variance tends to be higher during episodes of financial turmoil or increasing economic fluctuations (see values for 2001 and 2007 in Table 3). Overall, these findings suggest that hypothesis 2 above is confirmed as regards the effect of monetary policy on bank lending.

The results obtained when *M* is smoothed by the market power variable are the same and indeed even more pronounced in terms of variance. The coefficients for banks with relatively high market power are found on the right end of the distribution shown in Figure 2d, and these coefficients are less significant or not significant at all if they fall outside the relevant confidence interval. Therefore, market power of banks and competitive conditions in general prove to be a significant element through which monetary policy affects lending. In particular, as in the case of capitalization, high market power tends to buffer the impact of a shift in policy rates. Thus, the third hypothesis as regards the impact of monetary policy on bank lending is also confirmed. This is a novel result because previous studies considered measures of size or industry-level measures of competition, not bank-level market power.

An approximately normal distribution is illustrated in the figure concerning liquidity (Figure 2b), which implies that a parametric model may capture quite well the "global form" of the transmission through liquidity. Still, the large variance of the coefficients suggests that the impact of monetary policy is not uniform and can be important for certain banks (e.g. those identified with coefficients close to -0.5) and of less importance for other (those identified with coefficients on the right end of the distribution). The fact that different levels of bank liquidity yield wide differences in the response of bank lending following a change in monetary policy suggests that the relevant part of hypothesis 1 is also confirmed.

### 4.3. Monetary policy and risk

The results for the risk equations indicate that credit risk is highly persistent and that the monetary policy variable has on average a negative but marginally significant coefficient. (confidence intervals overlap with 0 when $M$ is smoothed by liquidity and capitalization).[18] Without placing particular emphasis on monetary policy variables (they use them as controls), recent studies on credit risk provide mixed results for the relationship between interest rates and credit risk (see e.g. Chen, 2007; Salas and Saurina, 2003). This is to be expected, since theory suggests that, on the one hand, low interest rates indeed tend to create excessive risk-taking, but, on the other, they also reduce the risk of outstanding bank credit. In the only detailed studies of the risk-interest rate nexus, Ioannidou et al. (2008) find that low interest rates are associated with higher risk and tend to reduce the loan rates of risky vis-à-vis riskless borrowers. Jimenez et al. (2007) suggest that low interest rates reduce credit risk in the very short run but worsen it in the long run. Therefore, even though the present study does not use such detailed data, the general results are in line with those of this very recent literature.

Note, however, that the findings presented in Figure 3 show that the average negative and marginally significant relationship between $\Delta M$ and $\Delta r$ in Eq. (2) is not the same for all banks, with the distributions of the estimated coefficients being bimodal for all three specifications. Banks with higher liquidity and capitalization are on the right end of the distributions shown in Figures 3b and 3c. For these banks the impact of a monetary policy change on credit risk is close to 0, which implies that the relevant parts of hypotheses 1 and 2 are confirmed. However, there also exist some banks in the sample that seem to take on extra risk following an increase in policy rates (i.e. many coefficients are positive). This implies that higher liquidity and, primarily, capitalization induce some banks (those on the right end of the distributions in Figures 3b and 3c) in light e.g. of a contractionary monetary policy to continue promoting risk-taking activities despite the fact that funding becomes more expensive. Nevertheless, such behavior could eventually

[18] $M$ is significant if 90% confidence intervals are constructed (i.e. $M$ is significant at the 10 % level).

lead to the erosion of the liquidity and/or capital base of these banks and in turn to a deterioration in the health of their balance sheets in the face of a credit crunch.

Even when market power is used as the smoothing variable (Figure 3d), more than 20 percent of banks that are identified with comparatively high market power do not change their attitude toward credit risk (see the highest peak of the distribution where coefficients are very close to 0) and only 17 percent of banks with less market power clearly follow risk-averse strategies (see the second pick on the left of the distribution). To expand on the above findings, a moral hazard mechanism may be in place (if one assumes that banks identified with positive coefficients embrace risky borrowers that are rejected from other risk-averse lenders), in order to promote the self-interest of managers or to gain a greater market share. Finally, the fact that in all three graphs the distributions are bimodal with a peak in the neighborhood of zero (indicating no response or a small response) and a peak in the region of negative values, definitely implies that the patterns studied are complicated, the effects are highly heterogeneous and inference cannot be made adequately on the basis of normal, symmetric or unimodal distributions that are associated with parametric models.

### 4.4. Monetary policy and profitability

The estimates of the profitability equations indicate that bank profits persist but they will eventually return to their normal (average) level because, as in the previous two models of bank lending and risk, the coefficients on the lagged variable are between 0 and 1. The average coefficient on the monetary policy variable is not statistically significant and takes a value very close to 0 in all three specifications (see Figures 4b,c,d). Thus, a parametric model would suggest that there is no impact of monetary policy on bank profits.

Yet, heterogeneity of the coefficients in the profitability equations is quite large, especially when the monetary policy variable is smoothed by capitalization and market power. In particular, note that, when $M$ is smoothed by capitalization, the relative coefficients have a mean value of -0.051 (see Table 2). The distribution of the coefficients (see Figure 4c) is trimodal and the highest peak is on the left side (at around -0.21). This implies that for a large number of less capitalized banks (the ones found on the left of the distribution) the impact of monetary policy on bank profits is negative and statistically significant. However, this is not the end of the story illustrated in Figure 4c. While the presence of a second peak implies that a considerable portion of banks do not respond to a monetary policy change (because coefficients around this peak take values close to zero), there is also a third peak on the right end of the distribution showing that for approximately 10 percent of the banks in the sample that are well-capitalized, profitability and monetary policy changes are positively related. Therefore, the part of hypothesis 2 (and the associated literature) that relates to bank profitability does not fully describe the relationship in hand, as it does not predict the positive relationship identified for many banks in our sample. Differently phrased, it seems that banks with higher (lower) levels of equity capital benefit (lose) from a contractionary monetary policy. In contrast, the implications of hypothesis 1 seem to hold as the impact of $M$ on profitability for the banks with higher levels of liquidity is very close to 0.

In the profitability equations, a considerable difference is found between US and euro area banks in that most of the latter are identified with a non-existent (or even a positive) relationship between a policy change and profitability. Two explanations can be considered for this pattern. First, in the euro area the negative relationship may be masked by the considerable changes that occurred in the member countries during the sample period and the comparatively small volatility in interest rates relative to these changes. Second, the maturity structure of bank balance sheets in the euro area is quite different from that in the United States and more detailed data may be needed to generate a significant relationship. Similar effects are found when the monetary policy variable is smoothed by liquidity and most importantly by market power. Indeed banks with higher market power tend to earn higher profits when interest rates rise, while heterogeneity as measured by the variance of the distribution is again considerable (see Figure 4d and relevant columns in Tables 2 and 3).

## 5. Discussion and concluding remarks

It is old news to say that bank behavior nurtured the severe financial distress of 2008, which has unfolded into a recession. The present study seeks to identify whether data from the last 14 years assign a special role to monetary policy in shaping bank profitability, lending and risk-taking. We find that bank-level responses to policy rate changes are often far from the average, normal response identified by the parametric estimates. The validity of this argument is uncovered in the present study with the help of a local generalized method of moments technique, which provides observation-specific estimates of the impact of monetary policy on bank balance sheets. Therefore, heterogeneous bank behavior is identified at the bank level and is actually measured by the variance of the distribution of estimated coefficients.

As banks have a special role in the financing of economic activity (see e.g. Ashcraft, 2005) their heterogeneous behavior is of particular importance to researchers and policy-makers alike. The present analysis showed that banks with healthier balance sheets and market power follow different strategies from the ones with weaker balance sheets, and in some cases they seem to take advantage of the market. However, when it comes to risk-taking and profitability it is well-known that during sudden episodes of financial turmoil (such as the present one) the balance sheet strength of even the healthiest banks quickly deteriorates and those banks exposed to high risks may become insolvent.

The analysis of this paper has already covered a lot of ground and therefore we have chosen to leave some interesting extensions for future research. An obvious immediate extension is to use measures of heterogeneity to assess the effect of monetary policy on real output in empirical models similar to the ones proposed by Driscoll (2004) or Ashcraft (2006). If, in addition, more disaggregate bank-level data on loans is available, an interesting study may involve the effect of heterogeneity on the maturity of loans and the term-structure of interest rates. Other issues concerning differential responses of banks to regulation, with respect to their market power levels and their risk-taking behavior are also important extensions.

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**Table 1**
**Descriptive statistics**

| | United States | | | | Euro area | | | |
|---|---|---|---|---|---|---|---|---|
| Variable | Mean | Standard deviation | Minimum | Maximum | Mean | Standard deviation | Minimum | Maximum |
| Real total assets | 1.63e+07 | 7.67e+07 | 1214 | 1.32e+09 | 1.55e+08 | 6.58e+08 | 11671 | 9.82e+08 |
| Real total loans | 9516037 | 3.98e+07 | 816 | 6.88e+08 | 8.71e+07 | 6.45e+08 | 7312 | 1.13e+08 |
| Problem loans/loans[1] | 0.070 | 0.108 | 0.009 | 0.256 | 0.064 | 0.121 | 0.007 | 0.311 |
| Net profits/total assets | 0.021 | 0.038 | -0.447 | 0.858 | 0.023 | 0.040 | -0.523 | 0.733 |
| Liquid assets/total assets | 0.052 | 0.768 | 0.007 | 0.142 | 0.056 | 0.814 | 0.007 | 0.199 |
| Equity capital/total assets | 0.119 | 0.112 | -0.013 | 0.256 | 0.126 | 0.104 | -0.039 | 0.195 |
| Market power (H-statistic) | 0.673 | 0.175 | -0.199 | 1.134 | 0.498 | 0.387 | -0.236 | 1.095 |
| Interest rate[2] | 4.186 | 1.705 | 1.130 | 6.240 | 3.119 | 1.728 | 1.000 | 6.058 |
| Gross interest revenue/total assets | 0.095 | 0.180 | 0.011 | 0.477 | 0.116 | 0.207 | 0.008 | 0.555 |
| Interest expenses/total deposits and money market funding | 0.039 | 0.220 | 0.008 | 0.126 | 0.046 | 0.398 | 0.012 | 0.241 |
| Personnel expenses/total assets | 0.019 | 0.062 | 0.003 | 0.210 | 0.015 | 0.079 | 0.003 | 0.199 |
| Operating and administrative expenses/total fixed assets | 0.527 | 2.892 | 0.088 | 1.217 | 0.622 | 3.186 | 0.105 | 1.512 |
| Real GDP growth[3] | 3.174 | 1.136 | 0.759 | 4.548 | 2.236 | 0.851 | 0.794 | 3.854 |
| Stock market capitalization/GDP[3] | 1.204 | 0.311 | 0.730 | 1.381 | 0.669 | 0.459 | 0.132 | 2.688 |
| Concentration ratio[3] | 0.305 | 0.027 | 0.263 | 0.342 | 0.672 | 0.198 | 0.195 | 1 |

Notes: All level variables are expressed in thousand dollars.
1. This variable is multiplied by 10 for expositional brevity.
2. This variable corresponds to the federal funds rate for the United States and the European Central Bank policy rate. Before 1999 for the euro area countries (before 2001 for Greece), the euro area policy rate refers to the average of official refinancing rates for each country.
3. Values reported for these variables are euro area averages.

**Table 2**
**Mean and standard deviation of the estimated coefficients on the interest rate variable for the euro area and the United States**

| | Lending (capit.) | Lending (liq.) | Lending (m.p.) | Risk (capit.) | Risk (liq.) | Risk (m.p.) | Profit. (capit.) | Profit. (liq.) | Profit. (m.p.) |
|---|---|---|---|---|---|---|---|---|---|
| Euro area | -0.343 | -0.295 | -0.384 | -0.030 | -0.019 | -0.051 | -0.052 | 0.006 | -0.004 |
| | 0.408 | 0.099 | 1.114 | 0.066 | 0.039 | 0.601 | 0.273 | 0.314 | 0.912 |
| United States | -0.215 | -0.307 | -0.317 | -0.049 | -0.029 | -0.048 | -0.050 | -0.031 | -0.012 |
| | 0.367 | 0.076 | 0.882 | 0.053 | 0.019 | 0.510 | 0.264 | 0.287 | 0.797 |
| Full sample | -0.307 | -0.303 | -0.353 | -0.038 | -0.023 | -0.050 | -0.051 | -0.012 | -0.007 |
| | 0.388 | 0.090 | 1.008 | 0.061 | 0.030 | 0.551 | 0.268 | 0.303 | 0.855 |
| Conf. Intervals | (-0.545)-(-0.101) | (-0.462)-(-0.193) | (-0.567)-(-0.094) | (-0.089)-(0.018) | (-0.070)-(0.019) | (-0.271)-(-0.011) | (-0.200)-(0.247) | (-0.273)-(0.194) | (-0.279)-(0.233) |

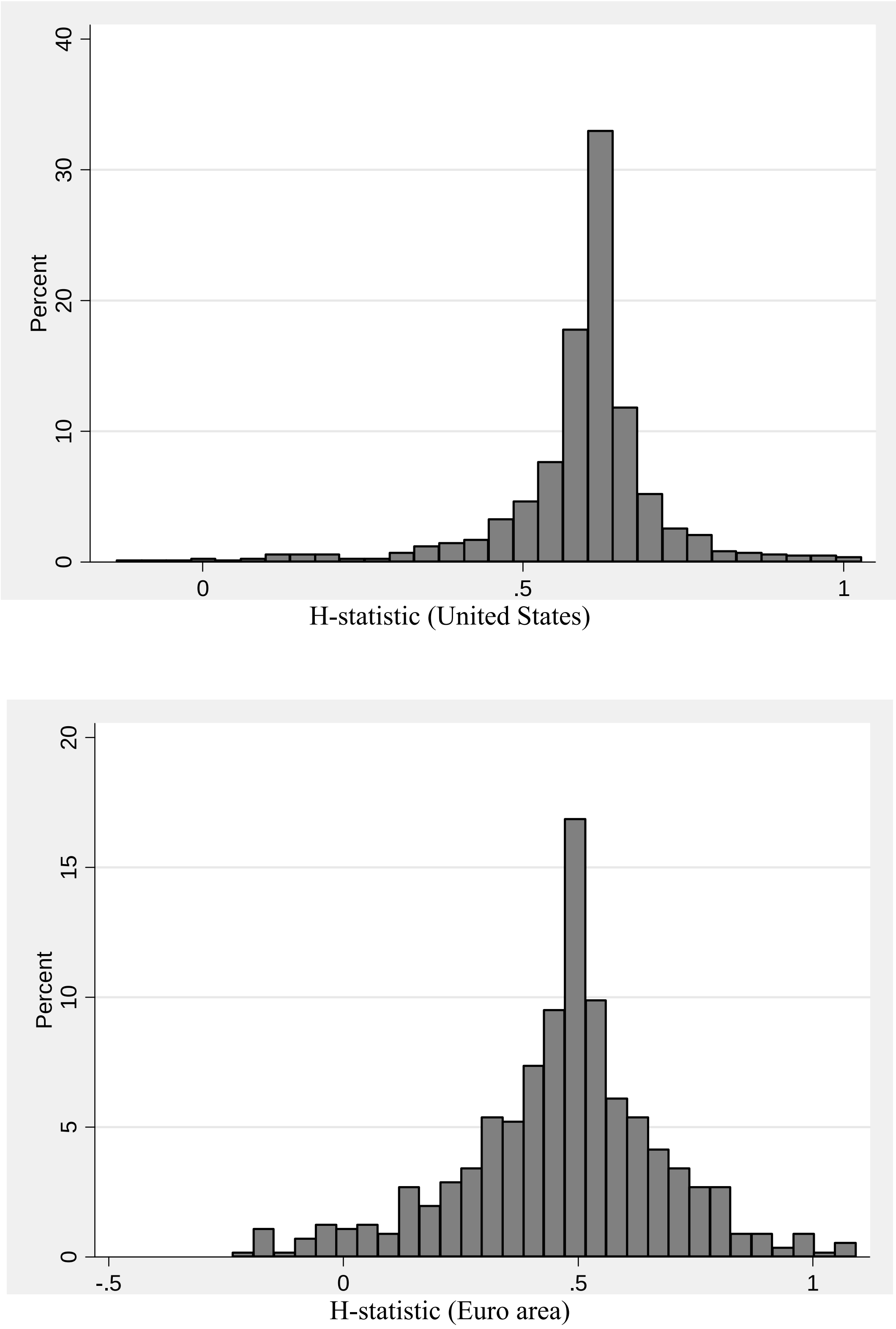
Notes: The table presents the mean and standard deviation (in the first and the second row for each region and the full sample) of the estimated coefficients for the lending, risk and profitability equations. It also presents the values for the constructed confidence intervals (conf. intervals) for the full sample. In each equation capitalization (capit.), liquidity (liq.) or market power (m.p.) are the smoothing variables of the policy rate.

**Table 3**
**Standard deviation of the estimated coefficients on the interest rate variable by year**

| year | Lending (capit.) | Lending (liq.) | Lending (m.p.) | Risk (capit.) | Risk (liq.) | Risk (m.p.) | Profit. (capit.) | Profit. (liq.) | Profit. (m.p.) |
|---|---|---|---|---|---|---|---|---|---|
| 1994 | 0.342 | 0.083 | 0.637 | 0.073 | 0.036 | 0.621 | 0.222 | 0.342 | 0.805 |
| 1995 | 0.421 | 0.085 | 0.604 | 0.062 | 0.035 | 0.618 | 0.235 | 0.364 | 0.878 |
| 1996 | 0.412 | 0.091 | 0.682 | 0.061 | 0.032 | 0.616 | 0.223 | 0.301 | 0.840 |
| 1997 | 0.373 | 0.089 | 0.741 | 0.053 | 0.032 | 0.601 | 0.218 | 0.270 | 0.746 |
| 1998 | 0.351 | 0.084 | 0.807 | 0.054 | 0.034 | 0.568 | 0.211 | 0.285 | 0.713 |
| 1999 | 0.320 | 0.086 | 1.003 | 0.060 | 0.034 | 0.507 | 0.235 | 0.290 | 0.801 |
| 2000 | 0.297 | 0.094 | 1.556 | 0.063 | 0.037 | 0.582 | 0.257 | 0.350 | 0.893 |
| 2001 | 0.585 | 0.099 | 1.971 | 0.078 | 0.047 | 0.577 | 0.341 | 0.387 | 1.121 |
| 2002 | 0.353 | 0.090 | 1.215 | 0.067 | 0.044 | 0.593 | 0.355 | 0.344 | 1.003 |
| 2003 | 0.299 | 0.079 | 0.787 | 0.053 | 0.041 | 0.505 | 0.317 | 0.305 | 0.865 |
| 2004 | 0.363 | 0.080 | 0.783 | 0.056 | 0.031 | 0.518 | 0.327 | 0.207 | 0.788 |
| 2005 | 0.347 | 0.103 | 0.958 | 0.054 | 0.038 | 0.498 | 0.211 | 0.240 | 0.810 |
| 2006 | 0.453 | 0.100 | 0.868 | 0.059 | 0.034 | 0.409 | 0.274 | 0.257 | 0.851 |
| 2007 | 0.512 | 0.107 | 0.622 | 0.069 | 0.044 | 0.501 | 0.320 | 0.305 | 0.857 |
| Average | 0.388 | 0.093 | 0.945 | 0.062 | 0.037 | 0.551 | 0.268 | 0.303 | 0.855 |

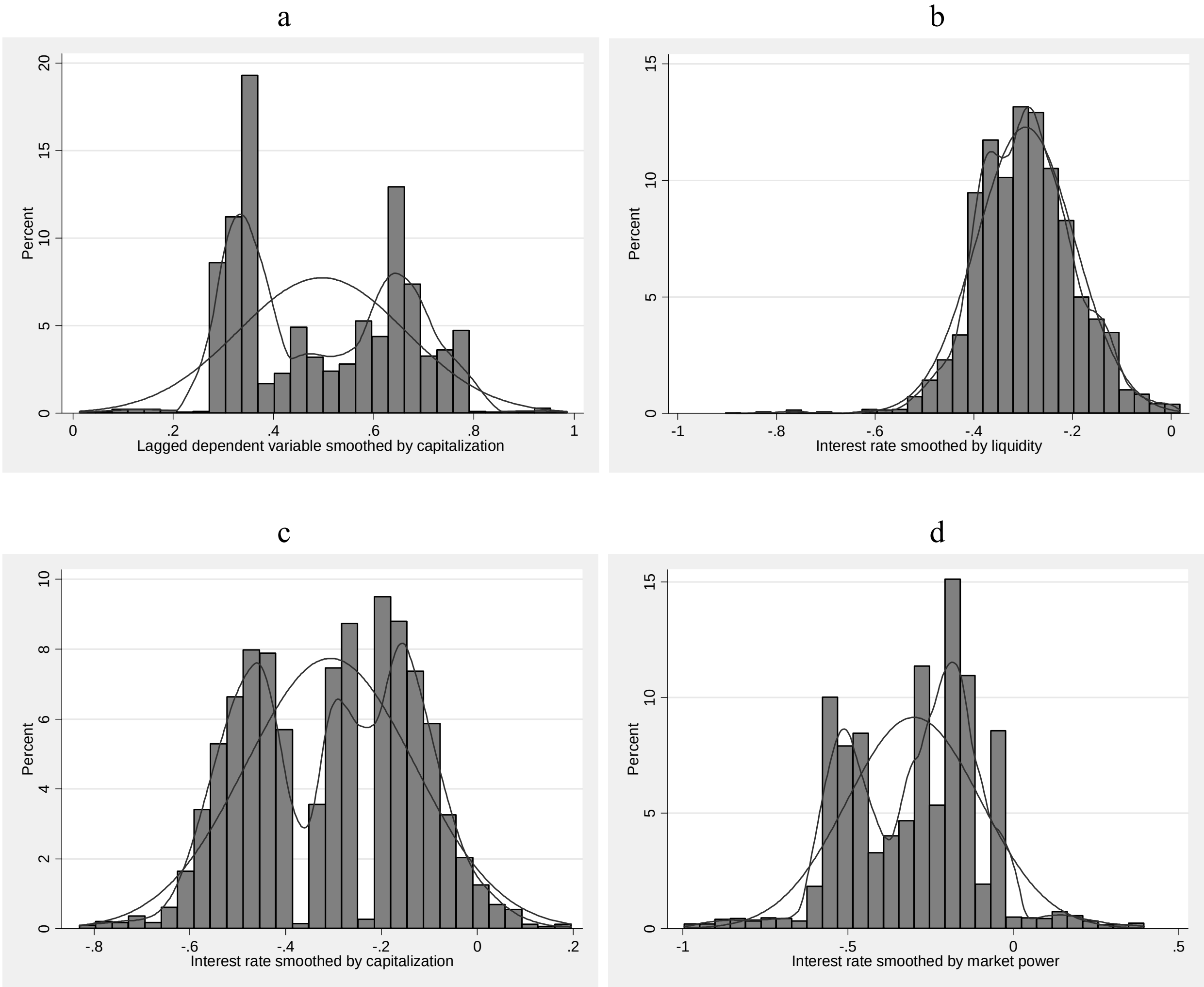
Notes: The table presents year-by-year the standard deviation of the estimated coefficients for the lending, risk and profitability equations. In each equation capitalization (capit.), liquidity (liq.) or market power (m.p.) are the smoothing variables of the policy rate.

**Figure 1**
**Frequency distribution of market power estimates**




Note: The graphs present frequency distributions of coefficient estimates on the H-statistic, as obtained from the estimation of Equation (4), for the United States and the euro area.

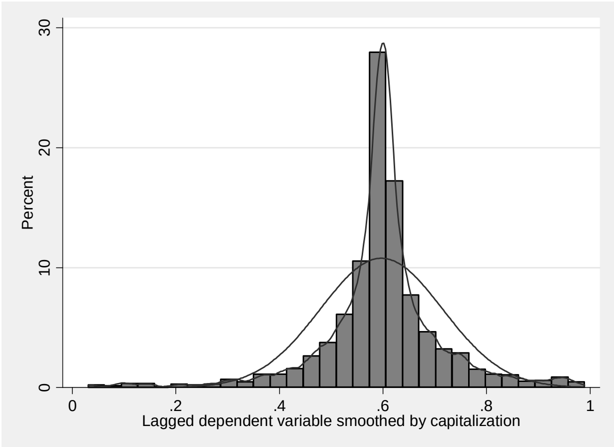
**Figure 2**
**Distribution of coefficients on the monetary policy variable obtained from the lending equation**



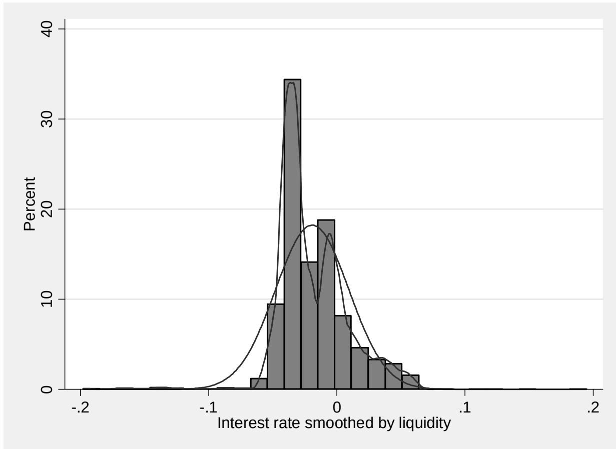

Note: The graphs present frequency distributions of coefficient estimates on the interest rate variable obtained from the lending equation, when the interest rate is smoothed by liquidity, capitalization or market power of banks, respectively. The first graph presents the frequency distribution of the coefficients on the lagged dependent variable when smoothed by capitalization. The two lines present the normal and the actual distribution of the coefficients, respectively. Confidence intervals for the coefficient estimates for the four graphs are (0.280)-(0.747), (-0.462)-(-0.193), (-0.545)-(-0.101) and (-0.567)-(-0.094).

**Figure 3**
**Distribution of coefficients on the monetary policy variable obtained from the credit risk equation**

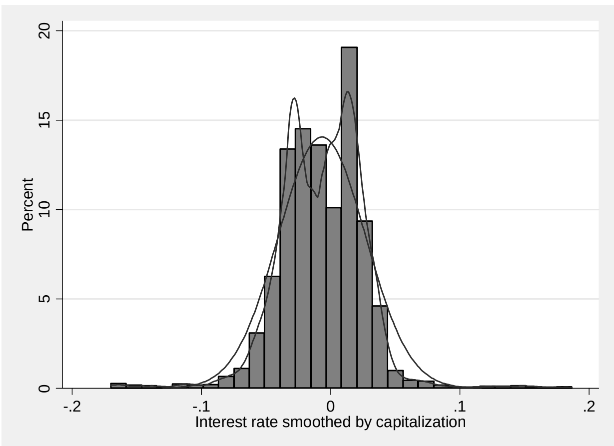
a



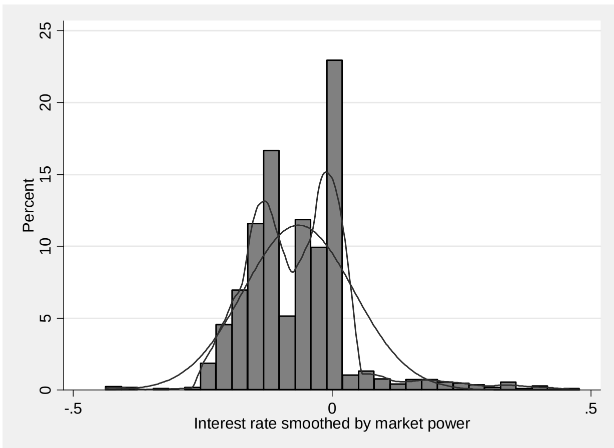

b




c

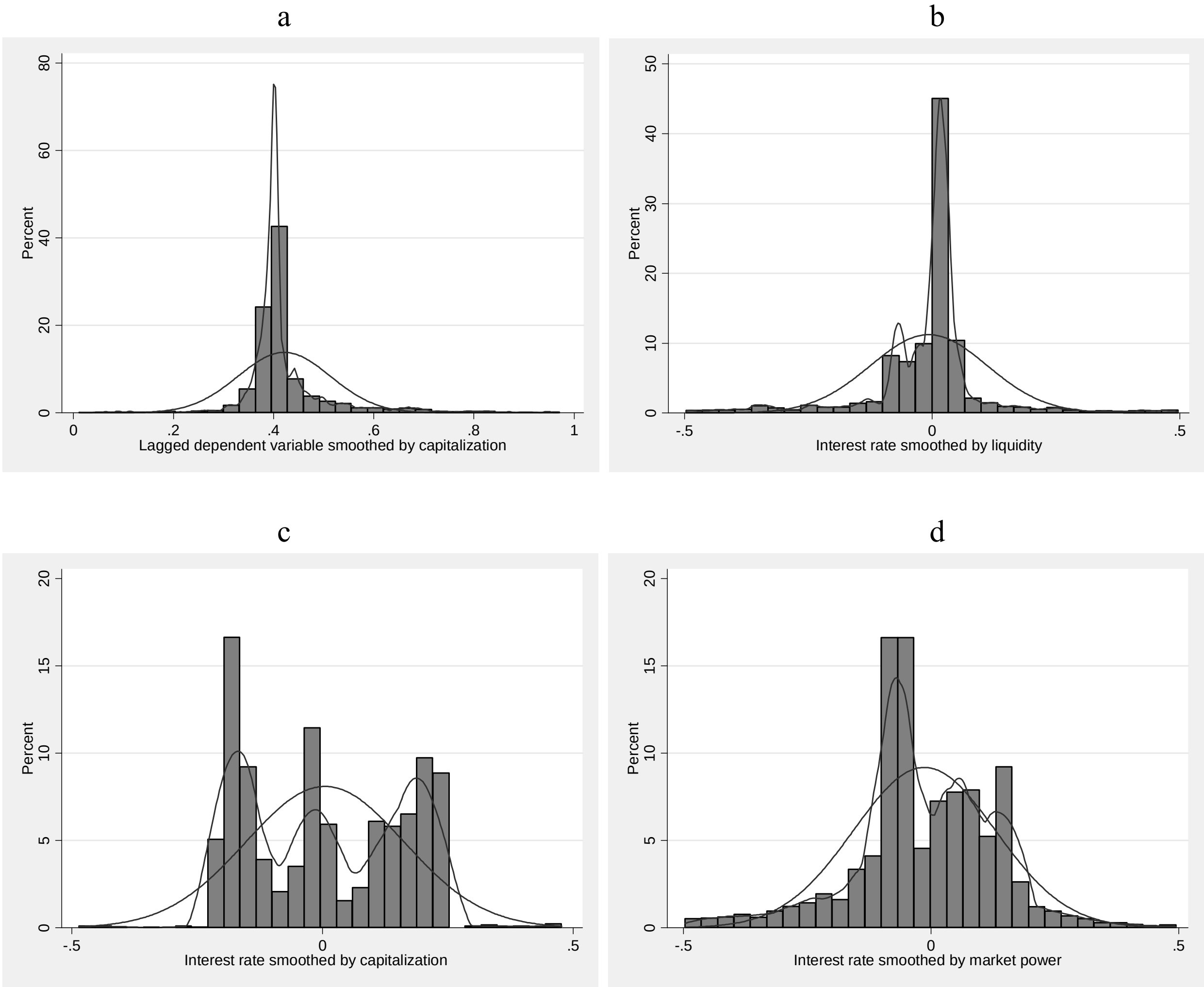



d




Note: The graphs present frequency distributions of coefficient estimates on the interest rate variable obtained from the lending equation, when the interest rate is smoothed by liquidity, capitalization or market power of banks, respectively. The first graph presents the frequency distribution of the coefficients on the lagged dependent variable when smoothed by capitalization. The two lines present the normal and the actual distribution of the coefficients, respectively. Confidence intervals for the coefficient estimates for the four graphs are (0.210)-(0.756), (-0.070)-(0.019), (-0.089)-(0.018), (-0.271)-(-0.011).

**Figure 4**
**Distribution of coefficients on the monetary policy variable obtained from the profitability equation**




Note: The graphs present frequency distributions of coefficient estimates on the interest rate variable obtained from the lending equation, when the interest rate is smoothed by liquidity, capitalization or market power of banks, respectively. The first graph presents the frequency distribution of the coefficients on the lagged dependent variable when smoothed by capitalization. The two lines present the normal and the actual distribution of the coefficients, respectively. Confidence intervals for the coefficient estimates for the four graphs are (0.201)-(0.660), (-0.273)-(0.194), (-0.200)-(0.247), (-0.279)-(0.233).

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