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The relation between banks' funding costs, retail rates and loan volumes: An analysis of Norwegian bank micro data*

Arvid Raknerud[†] and Bjørn Helge Vatne[‡]

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ABSTRACT: We use a dynamic factor model and a detailed panel data set for six Norwegian bank groups to analyze i) how funding costs affect retail loan rates and ii) how retail rate differences between banks affect market shares. The data set consist of quarterly data for 2002Q1-2011Q3 and include information on loan volumes and retail (interest) rates for loans to firms and households. The cost of market funding is represented in our analysis by the three-month money market rate and a proxy for market risk – the credit spread on unsecured senior bonds issued by Norwegian banks. Our estimates clearly suggest incomplete pass-through: a 10 basis points increase in the market rate leads to an approximately 8 basis points increase in retail loan rates. We also find that credit demand from households is more elastic with regard to the loan rate than demand from businesses.

JEL classification: C33, E27, E43

Keywords: monopolistic competition, credit spread, pass-through, funding costs, bank micro data, dynamic factor model

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1 Introduction

In this paper we investigate empirically i) how changes in the funding cost of banks affect loan rates to households and businesses and ii) how retail lending rate differences between banks affect their market shares, i.e., their share of total loans. While the transmission mechanism – how changes in market rates affect retail rates – have been extensively studied both in the theoretical and empirical literature¹, much less is known about the elasticity of credit demand with respect to loan rates. We investigate both issues within a simultaneous system of equations framework. The system is based on a theoretical model of monopolistic competition, where the banks are price setters in the loan markets (Cournot competitors), but face a common funding rate. According to the theoretical model, each bank’s market share (i.e., share of total loans) becomes a function of the ratio of its loan rate to the market loan rate, where the latter is a price index constructed from all the individual banks’ loan rates.

In our econometric implementation of the model, we utilize quarterly panel data on Norwegian banks from 2002Q1 to 2011Q3, which we aggregate up to six bank groups, such that all banks in the same group have a common covered bond company (see Table 1). We also investigate the impact of market risk on retail rates and market shares, where market risk is measured as the average spread between the interest on three-year senior unsecured bank bonds and the three-month Norwegian interbank rate. This credit spread can be interpreted as the compensation required by investors for credit and liquidity risk.

Traditionally the relationship between retail lending rates, loan volumes, funding

¹See e.g. Allen (1988), Hannan and Berger (1991), Angbazo (1997), De Bondt (2002) and De Graeve et al. (2007).

costs and other (macro economic) variables have been examined using time-series econometric models of bank interest rate and credit growth. Typically, the focus has been on aggregate demand and supply of credit. An example is the cointegrated vector autoregressive SMM model of Norges Bank (see Hammersland and Tr  e, 2012). However, the problem of separating supply side and demand side effects has not been solved within this empirical framework. An alternative approach towards resolving the identification problem has been to utilize the heterogeneity between different types of credit (e.g. bank loans vs bonds) and different types of agents (e.g. large vs small firms) regarding how they respond to liquidity shocks. This approach aims to identify exogenous liquidity shocks that affect the supply side of lending through the so-called bank lending channel – but not the demand side. See for example Kashyap and Stein (2000) and Ashcraft (2006). Some background and discussion about the bank lending channel is given by Kashyap and Stein (1994).

In this paper we rely on a structural model with monopolistic competition between banks to separate between demand side and supply side effects. We restrict our attention to microeconomic aspects of banking, by analyzing individual bank group’s market shares of loans, rather than their loan volumes in absolute terms. From our theoretical model, we derive “exclusion restrictions”, i.e., variables that affect banks’ retail rates but not the demand for credit. In this way we are able to estimate the elasticity of demand with respect to loan rates, as well as investigating the impact of changes in funding costs – including risk premiums – on retail rates.

When market risk (credit and/or liquidity risk) increases, banks may restrict loan supply for given interest rates by changing the non-price terms of loans and/or enforce a stricter screening of loan applicants. Norges Bank’s Survey of bank lending² confirms that this has indeed been the case in Norway after 2007Q4. Thus,

²See <http://www.norges-bank.no/en/about/published/publications/norges-banks-survey-of-bank-lending/>

there may be a *direct* effect from changes in market risk to loan supply, especially for unsecured loans.

The period analyzed in this paper – 2002Q1 to 2011Q3 – includes a period of financial distress and is also characterized by increased competition and productivity growth due to rapid increase in Internet-based payment services. One effect of the latter is that the interest margin between loan rates and deposit rates has decreased steadily over the period (or at least until the financial crisis), as documented in Raknerud et al (2011).

In the data, average volumes and interest rates over a quarter are specified for each bank group and for various types of loans, according to sector. We separate between loans to households and loans to corporations in the non-financial sector. The corresponding interest rates and loan volumes are analyzed within the framework of a dynamic factor model. The use of common dynamic factors is a parsimonious way of capturing the comovements among variables, advocated e.g. by Bernanke et al. (2005) and Forni et al. (2000). As a result, we are able to separate between the effect on retail rates of common observed variables (such as interbank market rates) and the effect of unobserved common variables (reflecting, for example, changes in bank regulations, competition and productivity). In accordance with most empirical literature on bank interest rates (e.g., Saunders and Schumacher, 2000), our model includes an interbank market rate, i.e., the three-month Norwegian Inter Bank Offered Rate (NIBOR), as a key exogenous variable.

Our empirical framework allows us to test particular hypotheses about both the short-run and the long-run (“steady state”) relationship between market rates (marginal funding costs) and retail rates, and to estimate the elasticity of credit demand from households and corporations. If banks have market power, they are faced with a trade-off between conflicting goals: a high interest rate on loans on the

one hand and a high market share (loan volume) on the other. If there is incomplete pass-through, the spread between the retail rates and the price of market funding will decrease as a result of an increase in the former.

The main novelty of this paper is to consider the determination of retail lending rates and loan volumes within a simultaneous system of equations which encompasses an underlying theoretical model of monopolistic competition between banks. Exclusion restrictions derived from the theoretical model are essential in order to solve the classical identification problem related to the parameters of the demand equation: retail lending rates are determined simultaneously with loan volumes.

The remainder of the paper is organized as follows. Section 2 describes the theoretical model of monopolistic competition between banks. The data and the empirical model are presented in Sections 3 and 4 , respectively. Finally, Section 5 gives the results and Section 6 concludes.

2 The theoretical framework

We take as a starting point a simple model with heterogeneous banks, and derive explicit demand functions for loans under the assumption of a representative agent with constant elasticity of scale (CES)-preferences over loans from different banks. Our model differs from the models of credit demand under monopolistic competition considered in Freixas and Rochet (2008), by the assumption of a representative consumer with CES-preferences. Thus we do not derive – or describe – the heterogeneity between banks from primary assumptions about e.g. the location of banks and customers (or the distance between them), but rather resort to a rather stylized representation of product differentiation. Of course, the assumption of a representative consumer with CES-preferences is standard in the industrial organization literature since the classical paper by Dixit and Stiglitz (1977).

First, we assume a representative agent which uses loans to finance investments or to purchases durable consumption goods. Total loans equal,

$$L = \sum_{i=1}^N L_i,$$

where L_i is loans from bank i . Total interest payments equal $\sum_{i=1}^N r_i L_i$, where r_i is the loan rate of bank i . L_1, \dots, L_N enter the agent's utility function through the CES-quantity index:

$$\left(\sum_{i=1}^N (a_i L_i)^\rho \right)^{\frac{1}{\rho}}, \quad \rho < 1.$$

Hence, the agent's choice of total amount of loans (L) and each bank's market share, $x_i = L_i/L$, are the results of separable decisions. In particular, the market share x_i follows from cost minimization:

$$\{x_1, \dots, x_N\} = \arg \min_{x_1, \dots, x_N} \sum_{i=1}^N r_i x_i \quad \text{s.t.} \quad \left(\sum_{i=1}^N (a_i x_i)^\rho \right)^{\frac{1}{\rho}} = y.$$

The well-known solution is

$$x_i = y a_i^s \left(\frac{r_i}{R} \right)^{-s}, \quad (1)$$

where

$$s = \frac{1}{1 - \rho}, \quad \text{with } \varrho = \rho / (\rho - 1),$$

and

$$R = \left(\sum_{i=1}^N (r_i / a_i)^\varrho \right)^{\frac{1}{\varrho}}.$$

By allowing the parameters a_1, \dots, a_N to have different values, the demand for loans from different banks will differ even if their loan rates are the same: $r_1 = \dots = r_N$. Since we consider a representative agent, the a_i -parameters cannot be given a direct interpretation in terms of, say, transaction costs or market segmentation, but reflect the combined effect of all non-price factors that affect the preferences for loans towards individual banks, including frictions in the customers' adjustment of their portfolios.

For any variable z_i , $i = 1, \dots, N$, define \bar{z} as the geometric average:

$$\bar{z} = \prod_{j=1}^N z_j^{\frac{1}{N}}. \quad (2)$$

It follows from (1) that

$$\ln(x_i) = -s \ln(r_i/\bar{r}) + \alpha_i, \quad (3)$$

where

$$\alpha_i = \ln(\bar{x}) + s(\ln(a_i) - \ln(\bar{a})).$$

Thus demand depends on the relative price r_i/\bar{r} .

To provide loans, banks need to raise funds. Assume that the wholesale market is the marginal source of funding and that banks face a constant marginal funding costs equal to c , i.e. regardless of the quantity of market funding. Decisions regarding loans and deposits are assumed separable, as in the Monti-Klein model (see Freixas and Rochet, 2008, Section 3.2). Thus deposits are not considered a marginal source of financing. Assume furthermore that each bank has constant operating costs equal to f_i per unit of loans (i.e., costs of labor, intermediary inputs and physical capital). These costs differ across banks and are therefore indexed i .

Next, similar to Japelli (1993) and Corvoisier et al. (2002), we incorporate credit risk through a bank-specific parameter μ_i – the default probability on any loan granted by bank i . The bank's choice of loan rate is then given by the solution to the expected profit maximization problem

$$\max_{r_i} \{(1 - \mu_i)r_i - c - f_i\}Q(r_i),$$

where $Q(r_i) = yw_i^s \left(\frac{r_i}{R}\right)^{-s}$ expresses the bank's market share, x_i , as a function of the retail loan rate, r_i . We assume that bank i takes R and y as given. The solution is then:

$$r_i = \frac{s}{(1 - \mu_i)(s - 1)}(c + f_i). \quad (4)$$

In the limiting case when $s \rightarrow \infty$, the coefficient of c in (4), $s/(1 - \mu_i)(s - 1)$, tends to $1/(1 - \mu_i)$.

Due to the multiplicative form of the demand function (1), the factor yw_i^s does not enter (4). The assumption of monopolistic competition implies that there exists no supply curve from the individual banks. The banks' adjustment is given solely by the mark-up rule (4). For a given (endogenous) interest rate r_i , realized demand is determined by the demand function (1).

If the mark-up coefficient in (4), i.e., the coefficient of $c + f_i$, is less than one, we have incomplete pass-through from market rates to loan rates. The more elastic demand (less market power), the smaller is this coefficient. In the (monopolistic competition) model of Hannan and Berger (1991), incomplete pass-through is a result of market power. However, as seen from (4), market power (inelastic supply of deposits or demand for loans) does not necessarily translate into incomplete pass-through (the mark-up coefficient being *less* than one) in the case of loan rates. The mark-up coefficient will depend both on the functional form of the demand function and on the degree of compensation for market risk – the factor $1/(1 - \mu_i)$. A more than one-to-one adjustment of retail loan rates to changes in market rates are theoretically possible and also sometimes reported in the empirical literature (see e.g. De Bondt, 2002; Table 1). However, the main bulk of empirical results support the view that pass-through is incomplete with regard to loan rates. Thus we will now consider a modification of our model.

So far we have not taken bank regulation into account, but assumed that the banks' marginal source of funding is wholesale funding, regardless of their level of equity. However, during our observation period all banks were subject to the capital requirements of the Basel II Accords. A stylized version of these capital requirements may be formulated as follows (ignoring risk weighting for simplicity): Assume that

$E/Q \geq \alpha$, where E is total equity, Q is total loans, and α is a lower threshold determined by regulation. If this constraint is binding, the marginal cost of funding is a weighted sum of the marginal cost of market funding, c , and the cost of new equity, \bar{c} . The marginal funding cost is now given by $(1 - \alpha)c + \alpha\bar{c}$. If banks set marginal cost equal to marginal revenue, (4) must be modified accordingly:

$$r_i = \frac{s(1 - \alpha)}{(1 - \mu_i)(s - 1)}c + \frac{s\alpha}{(1 - \mu_i)(s - 1)}\bar{c} + \frac{s}{(1 - \mu_i)(s - 1)}f_i. \quad (5)$$

Even if the capital requirement is not binding in a given period, the bank must take into account the possibility that it could become so in the future. In any case the marginal funding cost will depend on the cost of new equity. A discussion of the importance of the cost of equity for banks' funding costs is given in Fabbro and Hack (2011), who find evidence that in Australia there has been an increasing contribution from equity costs to the total funding costs of banks during the last years, especially with regard to loans to businesses.

An important consequence of equation (5) is that the mark-up coefficient may be either less than or larger than 1 also when demand is infinitely elastic – in the latter case the coefficient becomes $(1 - \alpha)/(1 - \mu_i)$. Thus we cannot from the degree of pass-through infer anything about the elasticity of demand.

By focusing exclusively on funding costs and by incorporating market risk through a fixed parameter, μ_i , our formal model offers a simplistic view of the transmission mechanism. Obviously, other factor may affect retail rates.

First, the presence of adverse selection: an increase in the retail rate will attract riskier borrowers and increase the risk of default (thus μ_i will depend on r_i). Thus banks are facing a trade-off : they have the incentive to raise lending rate, as a risk premium, while they cannot do that drastically because of the rising probability of default. In the model of Stiglitz and Weiss (1981), banks do not fully pass all the increase in the market rate over to their retail loan rates. Rather, loan rates

are sticky upwards and credit supply rationed. On the other hand, if banks can discriminate between borrowers through screening and collateral requirements, they may adjust their loan rates more than one-to-one for risky borrowers to compensate for increased risk due to adverse selection.

Second, other types of risk, like interest rate and liquidity risk, may be taken into consideration. Interest rate risk is the least important one: This occurs if a bank issues a loan with a fixed rate, while its funding has a variable rate (see Freixas and Rochet, 2008). To alleviate this risk, banks typically enter into interest rate swaps to achieve a level of variable-rate exposure that matches the variable-rate loans. On the other hand, liquidity risk occurs because of reduced liquidity in wholesale markets. According to Bank of England, during the financial crisis a substantial part of the spread on senior unsecured bonds was compensation for reduced liquidity in funding markets.³

Third, increased risk (as measured e.g. by indicative spreads) may lead to a tightening of credit standards to better screen the high quality borrowers. Riskier projects may face higher collateral requirements, shorter contractual maturity or loan applications may be turned down. While it is difficult to measure (and disentangle) the different types of risk involved – and the effects on retail rates and loan volumes – the above reasoning suggests that increased risk may affect both spreads (between retail rates and the market rate) and loan volumes directly. If the non-price terms of different banks react differently market risk increases, their market shares will also change.

Given the stylized character of the theoretical model, we will not formally test the assumptions behind it below, but rather use it as guidance for operationalization, interpretation of results and choice of functional forms.

³See Chapter 3 (especially Figure 3.16) in Bank of England's Financial Stability Report, Issue 27, June 2010: <http://www.bankofengland.co.uk/publications/fsr/2010/fsrfull1006.pdf>

3 Data

Our sample consists of balance sheets (accounts) data from Norwegian banks from 2002Q1 until 2011Q3 assembled by Statistics Norway⁴. The bank-level data were aggregated into seven bank groups, as listed in Table 1 (see the note to the table for a detailed definition of the bank groups). The grouping was done according to ownership, nationality and common covered bond mortgage (OMF) companies. Covered bonds (OMFs) were introduced in Norway in June 2007 and have become an important source of funding for Norwegian financial services groups and banking alliances.⁵ Key statistics for the seven bank groups are given in Table 1. The last group is a residual group and will not be included in the econometric analysis. During the estimation period, there have been entries, exits, mergers and acquisitions that affect the bank groups. An example is the acquisition of Fokus bank by Danske Bank of March 2007. The sample is constructed on the basis of the bank structure at the end of the estimation period. For example, the time series for the group DNB includes all banks that were included in this bank group at the end of the estimation period.

There is considerable heterogeneity in the funding sources of banks. Small national banks tend to have more deposits than foreign or large national banks, while the latter banks rely more on market funding. For example, Terra Gruppen, which is a group of small banks, had the highest average ratio of household deposits over total loans during 2001-2010: 42 percent. The two foreign bank groups had the lowest ratio – 18 percent – while the largest bank group, DNB, had a ratio of 29 percent.

The log of the market shares for each of the six first bank groups are shown in

⁴See <http://www.ssb.no/skjema/finmark/rapport/orbof/> (in Norwegian).

⁵See the following article by Rakkestad and Dahl in *Penger og Kredit* 1/2010 (in Norwegian): http://www.norges-bank.no/Upload/80111/OMF_marked_i_vekst_PK_1_10_nov.pdf

Figure 1. The corresponding graphs showing the log of the relative loan rate for each bank group (i.e., relative to the market loan rate index) are shown in Figure 2. We see that there is considerable persistence both in the market shares and interest rate differentials between the bank groups over time. However, we see that Bank group 1 has had a declining market share for loans to households while the opposite is the case for Bank groups 2 and 3. Regarding loans to businesses, Bank group 1 seems to have lost a considerable share of their initial market position to Bank group 3. We also observe considerable interest rate differences between Bank group 1 and 3 with regard to household loans, with Bank group 3 generally having lower rates until 2007, but higher thereafter. From Figures 1 and 2 we see no clear connection between (changes in) market shares and relative loan rates.

Table 1: Descriptive statistics for seven bank groups (in 2011).

Bank group	Percentage of market				Percentage of bank loans to:	
	Total assets	Loans to:		Deposits	Households	Businesses
		Households	Businesses			
1. DNB ¹	41	32	30	34	65	26
2. Subsidiaries of foreign banks ²	13	13	18	12	57	34
3. Branches of foreign banks ³	14	11	19	10	54	39
4. SpareBank1-alliansen ⁴	14	19	16	18	68	24
5. Terra Gruppen ⁵	5	9	4	6	76	14
6. Other savings banks ⁶	10	13	11	13	70	24
7. Other commercial banks ⁷	2	3	3	4	69	19

Source: Norges Bank

¹DNB Bank, Nordlandsbanken, DNB Boligkreditt and DNB Næringskreditt

²Nordea Bank Norge, Santander Consumer Bank, SEB Privatbanken and Nordea Eiendomskreditt

³Fokus Bank (branch of Danske Bank), Handelsbanken, SEB, Swedbank, Handelsbanken Eiendomskreditt, Skandiabanken and seven other branches

⁴SpareBank 1 SR-Bank, SpareBank 1 SMN, SpareBank 1 Nord-Norge, Sparebanken Hedmark, the 11 other savings banks in SpareBank 1-alliansen, SpareBank 1 Boligkreditt, BN Bank, Bank 1 Oslo Akershus, 1 commercial mortgage company and 1 other residential mortgage company

⁵Terra BoligKreditt, Terra Finans og Kredittbank, 77 savings banks and 1 commercial bank which are owners of Terra-Gruppen AS + 1 other residential mortgage company

⁶Sparebanken Vest, Sparebanken Møre, Sparebanken Sør, Sparebanken Pluss and Sparebanken Sogn og Fjordane, 14 other savings banks, 10 residential mortgage companies and 1 hybrid covered bond mortgage company

⁷Storebrand Bank, Storebrand Boligkreditt, Landkreditt Bank, Gjensidige Bank, 7 other commercial banks and 2 other residential mortgage companies

Since 2001, Norwegian banks have been obliged to report end of quarter interest rates. We calculate the average interest rate of the banks in a group as a value-weighted average of the reported interest rates. From the bank statistics we get interest rates and volumes of various loans in each bank. The interest rates are weighted by the corresponding nominal book values to obtain a value-weighted average interest rate.

The three-month effective Norwegian Inter Bank Offered Rate (NIBOR) reported by Norges Bank is a proxy for the cost of long- and medium-term market financing. Illustrations of some key rates are provided in Figure 3. The graphs labeled “Loans to households” and “Loans to businesses” are geometric averages based on bank-group specific loan rates. Throughout the observation period, the retail loan rates for businesses lie slightly above that of loans to households.

As discussed above, banks cannot raise more funds solely by increasing the rates on deposits, because bank customers – households and firms – typically do not react quickly to changes in deposit rates. Thus, we interpret the cost of raising senior unsecured bonds from institutional investors in the wholesale market as the marginal funding cost. An unsecured bond may be issued with a fixed or variable interest rate. In the first case, a Norwegian bank typically enters into an interest rate swap to achieve a level of variable rate exposure that matches the variable rate loans. The banks’ costs may be expressed by two components: the variable rate cash flows paid in the interest rate swap (normally three-month NIBOR) and the fixed cash flow due to the issuer-specific credit spread over the swap rate ⁶.

We include both the three-month NIBOR, henceforth denoted r_t , and the spread of unsecured senior bonds issued by Norwegian banks as measures of the cost of mar-

⁶For examples of bank bonds with varying maturity and with interest payments equal to the three-month NIBOR plus a fixed credit spread, see <http://investor.sparebank1.no/obligasjonslan-sparebank1/>

ket funding. As a measure of the latter, we use an index consisting of indicative bid spreads based on average trading levels over the swap rate (three-year fixed/three-month NIBOR) for senior bonds issued by a range of Norwegian banks since 2001, including DNB, Nordea Bank Norge and a representative selection of banks of various sizes and ratings. Both series are shown in Figure 3.

4 The empirical model

We now formulate an empirical model that encompasses the main features of the theoretical model presented in Section 2. As mentioned earlier, we distinguish between loans to businesses (B) and the households (H). We denote the corresponding loan rates for bank group i at time t by r_{it}^B and r_{it}^H , respectively, where $i = 1, \dots, 6$, and t refers to the end of a particular quarter in a given year. As mentioned in Section 3, r_{it}^B and r_{it}^H are calculated as weighted averages of more disaggregate interest rates, where the weights are taken from the outgoing balance in the bank accounts. The corresponding logarithm of the loan market shares are denoted by X_{it}^B and X_{it}^H , respectively.

Retail loan rates We first consider an econometric specification of the equations for the retail loan rates, r_{it}^B and r_{it}^H . Our explanatory variables are proxies for the exogenous funding costs of banks. The main variable is the three-month NIBOR rate, r_t , which is a key determinant of external funding costs. For the individual banks, it is reasonable to assume that r_t is exogenous; that is, the individual bank cannot influence NIBOR through its demand or supply of credit in the interbank market. The rationale behind this assumption is that (major) banks can borrow and lend NOK through the foreign exchange rate markets such as the NOK–USD exchange swap market. Covered interest rate parity implies that the NIBOR rate is determined by international lending and swap exchange rates, which are exoge-

nous to individual Norwegian banks.⁷ We also include the credit spread, s_t , as an explanatory variable.

We now specify a stochastic relation between the retail loan rates (r_{it}^B, r_{it}^H) and the market rates (r_t, s_t). Our model accommodates flexible short-term dynamics, where the different types of retail rates and the retail rates of different banks, are allowed to react differently to exogenous shocks. Moreover, the econometric model incorporates bank group-specific parameters to allow heterogeneity with regard to banks' responses to exogenous variables. Finally, the model incorporates common shocks to account for comovements in the different rates due to unobserved (common) factors.

Conditional on the common explanatory variables, we model the individual retail rates as univariate autoregressive processes, augmented with common dynamic factors. Our approach can be seen as being in the tradition of multivariate structural time series models⁸. Specifically, we assume that, for $L = B, H$ (businesses and households):

$$r_{it}^L = \mu_i^L + \alpha_{i,0}^L r_t + \alpha_{i,1}^L r_{t-1} + \gamma_i^L s_t + \sum_{j=1}^{p_i} \phi_{ij}^L r_{i,t-j}^L + \sum_{k=1}^m \theta_{ik}^L f_{kt} + e_{it}^L, \quad (6)$$

where μ_i^L is a bank group- and interest rate-specific fixed effect, the α -parameters capture the effects of the NIBOR rate by allowing both the current NIBOR rate, r_t (through $\alpha_{i,0}^L$), and the lagged NIBOR rate, r_{t-1} (through $\alpha_{i,1}^L$), to affect r_{it}^L . One lag is allowed to capture the effect of notification rules that restrict the speed at which banks are allowed to increase their loan rates. The credit spread measure, s_t , is assumed to affect bank group i through the parameters γ_i^L .

The autoregressive parameters ϕ_{ij}^L , $j = 1, \dots, p_i$, determine how the effects of a shock in any of the exogenous variables evolve over time. The number of lags, p_i ,

⁷For an example, see equation (1) in Akram and Christophersen (2011): http://www.norges-bank.no/upload/publikasjoner/staff%20memo/2011/staff_memo.0111.pdf

⁸See Harvey (1989) for a general exposition of structural time series models and Stock and Watson (2002) for dynamic factor models.

is allowed to differ from bank group to bank group. Using the Akaike information criterion, we find that $p_i = 2$ or 3 is adequate. Finally, the unobserved stochastic terms consist of m dynamic factors, f_{kt} , $k = 1, \dots, m$, which pick up the dependencies across banks due to common, unobserved variables (e.g., effects of the business cycle, credit market regulations and competition) and the idiosyncratic error term, e_{it}^L , assumed to be independent across banks (i) and over time (t).

Market shares of total loans Analogously to (6), we assume

$$X_{it}^L = \nu_i^L + \beta_{i,0}^L \tilde{r}_{i,t}^L + \beta_{i,1}^L \tilde{r}_{i,t-1}^L + \kappa_i^L s_t + \sum_{j=1}^{q_i} \psi_{ij}^L X_{i,t-j}^L + \sum_{k=1}^m \zeta_{ik}^L f_{kt} + \varepsilon_{it}^L, \quad (7)$$

where the dependent variable is $X_{it}^L = \ln x_{it}^L$ – the log of bank i 's market share (share of total loans in sector L) and $\tilde{r}_{it}^L = \ln(r_{it}/\bar{r}_t)$ – the log of bank i 's relative loan rate, where \bar{r}_t^L is the (market) loan rate index for sector L . Moreover, ν_i^L is a fixed effect, and $\beta_{i,0}$ and $\beta_{i,1}$ capture the direct effects of the current and lagged value of $\tilde{r}_{i,t}^L$ on the dependent variable, cf. (3). The credit spread measure, s_t , is allowed to affect X_{it}^L through the parameters κ_i^L . Thus we allow a direct effect of the credit spread on loan volumes (and thus market shares) through the non-price terms of loans, as explained above. Note that (7) is a dynamic equation, with q_i lags of the dependent variable, $X_{i,t-j}^L$, entering on the right hand side of (7), with corresponding autoregressive parameters ψ_{ij}^L . Finally, the loading coefficients ζ_{ik}^L have the same interpretation as the θ_{ik}^L in (6).

For each bank group the vector of dependent variables consists of $(r_{it}^B, r_{it}^H, X_{it}^B, X_{it}^H)$. The corresponding vector of error terms $(e_{it}^B, e_{it}^H, \varepsilon_{it}^B, \varepsilon_{it}^H)$ is assumed to be independent across different i and t , and normally distributed with unrestricted covariance matrix Σ . Finally, the common dynamic factors, f_{kt} , are assumed to be independent, Gaussian AR(1) processes:

$$f_{kt} = \varpi_k f_{k,t-1} + \eta_{kt}, \quad \eta_{kt} \sim \mathcal{IN}(0, 1); \quad k = 1, \dots, m. \quad (8)$$

Thus, f_{1t}, \dots, f_{mt} are latent stochastic processes that capture the comovements between interest rates and market shares of the different bank groups *not* accounted for by the observed explanatory variables. The impact of the dynamic factors on an individual bank group is determined by the bank group-specific impact coefficients, θ_{ik}^L and ζ_{ik}^L . In our model these factors play a similar role to that of the “risk factor contributions” of Rosen and Saunders (2010), in the context of portfolio risk analysis. Our model is estimated by a version of the maximum likelihood algorithm described in Raknerud *et al.* (2010).

For identification, it is a crucial exclusion restriction that the NIBOR rate, r_t , enters (6), but not (7). This restriction is motivated by the theoretical model in Section 2. Another restriction is that the vector of error terms are assumed uncorrelated across bank groups. The rationale for the latter assumption is that common shock across banks are captured by the dynamic factors. Both these restrictions contribute to exogenous variation in the endogenous explanatory variable $\tilde{r}_{i,t}^L$ – and hence to identification.

Partial effects Our econometric framework allows us to disentangle both short-run and long-run partial effects of changes in exogenous variables on the dependent variables. First, we are interested in the effects of changes in the market rate on retail rates. Assume that the system is in a steady state at t defined by $r_{t-j} = r$ and $s_{t-j} = s$ (r and s are arbitrary fixed values), then

$$r_{it}^L = \frac{\mu_i^L}{1 - \sum_{j=1}^{p_i} \phi_{ij}^L} + \left(\frac{\alpha_{i,0}^L + \alpha_{i,1}^L}{1 - \sum_{j=1}^{p_i} \phi_{ij}^L} \right) r + \left(\frac{\gamma_0^L}{1 - \sum_{j=1}^{p_i} \phi_{ij}^L} \right) s + d_t^L + \tilde{e}_{it}^L \quad (9)$$

is a steady-state equation. Here d_t^L captures the effects of the present and lagged dynamic factors, f_{js} , $s \leq t$, and \tilde{e}_{it}^L is a moving average of the error terms e_{is}^L , for $s \leq t$. Equation (9) determines the long-run relation between retail rates and permanent (or persistent) levels of the exogenous variables. We interpret equation

(9) as the empirical counterpart of (5), with r taking the place of the marginal funding cost, c . Because of its lack of dynamics, it is reasonable to consider the structural model in Section 3 as expressing long-run (equilibrium) relations.

A similar steady state equation with respect to the log market share X_{it}^L , given a permanent value of the retail rate $\tilde{r}_{it}^L = \tilde{r}_i^L$, is given by

$$X_{it}^L = \frac{\nu_i^L}{1 - \sum_{j=1}^{q_i} \psi_{ij}^L} + \left(\frac{\beta_{i,0}^L + \beta_{i,0}^L}{1 - \sum_{j=1}^{q_i} \psi_{ij}^L} \right) \tilde{r}_i^L + \left(\frac{\kappa^L}{1 - \sum_{j=1}^{q_i} \psi_{ij}^L} \right) s + \delta_t^L + \tilde{\varepsilon}_{it}^L, \quad (10)$$

where δ_t^L and $\tilde{\varepsilon}_{it}^L$ are derived in a similar way as d_t^L and \tilde{e}_{it}^L . Equation (10) is the empirical counterpart of (3), with \tilde{r}_i^L taking the place of $\ln(r_i/\bar{r})$. Thus the coefficient of \tilde{r}_i^L can be interpreted as the elasticity of substitution in (3): $-s$. According to the theoretical model in Section 2, this coefficient should be negative and equal across different bank groups.

5 Results

Dynamic specifications To perform statistical tests, assess estimation uncertainty and interpret results, it is important to know whether the variables of interest are stationary or not. Our main assumption is that the vector of dependent variables, $(r_{it}^H, r_{it}^B, X_{it}^H, X_{it}^B)$, as well as the NIBOR rate, r_t , are $I(0)$. These assumptions are formally tested below – and *not* rejected. Consistent with this, all the estimated lag polynomials $1 - \phi_{i1}^L L - \phi_{i2}^L L^2 - \phi_{i3}^L L^3$ and $1 - \psi_{i1}^L L - \psi_{i2}^L L^2 - \psi_{i3}^L L^3$ (where L is the lag operator) have roots outside the unit circle. Moreover, the dynamic factors, f_{kt} , are estimated to be stationary AR(1) processes. The number of factors was set equal to four, while the number of lags in the AR(p_i) and AR(q_i) equations is equal to two in 22 of the 24 equations, and three in the remaining ones. These choices were made by applying the Akaike information criterion⁹.

⁹See Raknerud *et al.* (2010) for details regarding model selection in a similar model.

Table 2 displays the estimated sum of the autoregressive parameters $\sum_{j=1}^{p_i} \phi_{ij}^L$ and $\sum_{j=1}^{q_i} \psi_{ij}^L$, which appear, respectively, in the denominators in the long-run equations (9) and (10). If any such sum is close to one, the corresponding retail rate, r_{it}^L , or log market share, X_{it}^L , is nearly a unit root (integrated) processes. The main impression of these estimates is that the X_{it}^L – processes are highly autocorrelated. Some of these processes (Bank groups 1 and 3) are even close to being unit-root processes, i.e. $\sum_j \psi_{ij}^L = 1$. Market shares thus adjust slowly to changes in relative loan rates, and much more slowly than changes in retail rates do to changes in the NIBOR rate. The retail rates are clearly not unit root processes, but adjust very quickly to exogenous shocks. In fact, almost all adjustment is completed within the same and next quarter of the shock. The Wald tests reported in Table 2 reveal significant bank specific heterogeneity in the autoregressive dynamics, as we get clear rejections of the hypothesis that the sum of the autoregressive coefficients are equal across bank groups.

Table 2: **Estimates of sum of autoregressive parameters for each bank group.** Standard errors in parentheses are obtained by the delta method.

Equation: (X)	$\sum_{j=1}^{p_i} \phi_{ij}^L$		$\sum_{j=1}^{q_i} \psi_{ij}^L$	
	H	B	H	B
Bank group 1	.20 (.06)	.13 (.08)	.94 (.14)	.45 (.15)
Bank group 2	.20 (.06)	.13 (.06)	.43 (.14)	.45 (.14)
Bank group 3	.13 (.06)	.23 (.06)	.94 (.14)	.94 (.14)
Bank group 4	.25 (.06)	.11 (.14)	.69 (.13)	.72 (.13)
Bank group 5	.24 (.05)	.10 (.08)	.53 (.14)	.70 (.14)
Bank group 6	.20 (.05)	.03 (.05)	.37 (.14)	.77 (.14)
P-value Wald-test*	.04	.03	.007	.002

*Wald test of the restriction that all 6 banks groups have equal sum (5 d.f.)

Estimates for the retail rate equations Our focus will be on the estimated long-run relations. Table 3 exhibits the estimates of the coefficients of the long-run retail rate equations (9) for each individual bank group as well as for the representa-

tive bank – defined as the value-weighted average of the six bank groups with weights equal to the average market share of each group (the average of the second and third column in Table 1, respectively). We see that for the representative bank, the coefficient of r in the steady state is close to 0.8, and is significantly below one both in the household and business sector. Thus the hypothesis of complete pass-through in the long run is clearly rejected. If we examine the bank group-specific estimates in Table 3, they are all remarkably close to 0.8 – although somewhat smaller for Bank group 1. A formal test of whether all the steady-state coefficients of r are equal across all bank groups is provided by the Wald test reported in the last row of Table 3. Evidently, we cannot reject the hypothesis of homogeneous long-run parameters. According to our theoretical model, a low coefficient of the NIBOR rate indicate that loans from different banks are considered close substitutes.

We now turn to the coefficients of the indicative spread, s , in the steady-state equations for the retail rates. Table 3 shows that the bank group-specific parameters vary a great deal across bank groups, and that the estimation uncertainty is considerably larger than for the steady-state coefficients of r . However, for both sectors we clearly reject that the common coefficients are equal to zero. Our estimates say that a permanent unit increase in the credit spread leads to about one-third increase in the business loan rate in the long-run. For households, this estimate is somewhat lower, 0.23.

The estimates of the main coefficients of the aggregate equilibrium retail rate equations are depicted in (11):

$$\begin{aligned} \sum_{i=1}^6 w_i r_{it}^H &= d_t + \underset{(0.03)}{0.77}r + \underset{(0.06)}{0.23}s + \text{residual} \\ \sum_{i=1}^6 w_i r_{it}^B &= d_t + \underset{(0.03)}{0.81}r + \underset{(0.08)}{0.30}s + \text{residual}. \end{aligned} \tag{11}$$

Note that the degree of pass-through is much smaller for the spread, s_t , than for the NIBOR rate, r_t . Thus the marginal cost of market funding cannot be written simply as a sum of r_t and s_t . One explanation of this finding may be as follows: As seen from Figure 2, until 2008 the variation in funding costs was dominated by the NIBOR rate. However, from 2008Q1 to 2008Q4, the spread, s_t , increased dramatically, and was at the end of 2011 still higher than its pre-2008 level. Data for issuance indicate that Norwegian banks reduced their ordinary funding activity dramatically during the period 2008Q1 to 2008Q4, when the credit spread soared. At the same time, several authority measures to support banks' funding took effect which e.g. enabled the banks to fund mortgage loans through covered bonds. Moreover, a marked reduction in the policy rate led to a sharp fall in the NIBOR rate. The combined effect is that from 2008Q2 we observe a distinct fall in deposit margins relative to NIBOR (not displayed in the figure) and an (offsetting) increase in the margins of loans to households (relative to NIBOR). The latter effect is clearly visible in Figure 3. To conclude, the estimated effects of variations in s_t are identified mainly by events immediately before and after the onset of the financial crisis in 2008Q3 and it is difficult to separate pass-through effects from the effects of other events that took place simultaneously.

Table 3: Estimates of key parameters in the steady-state equations for retail rates. Standard errors in parentheses.

Equation (r_{it}^L):	Coefficients of r		Coefficients of s	
	H	B	H	B
Bank group 1	.67 (.04)	.81 (.03)	.31 (.09)	.46 (.12)
Bank group 2	.83 (.04)	.82 (.03)	.34 (.09)	.21 (.09)
Bank group 3	.81 (.04)	.82 (.06)	.14 (.11)	.26 (.08)
Bank group 4	.83 (.04)	.82 (.05)	.21 (.08)	.31 (.09)
Bank group 5	.81 (.04)	.74 (.05)	.11 (.07)	.21 (.13)
Bank group 6	.82 (.04)	.77 (.03)	.11 (.03)	.13 (.08)
Common estimate*	.77 (.03)	.81 (.03)	.23 (.06)	.30 (.08)
P-value of Wald-test**	.53	.14	.33	.24

*Value-weighted average across six bank groups

**Wald test of the restriction that all parameters are equal (5 d.f.).

Estimates of demand elasticities The estimates of the value-weighted average elasticity of demand, the coefficient of \tilde{r}^L in (12), show that there is an overall negative relation between the loan rates and market shares in both sectors. This is confirmed by the estimates of the individual demand elasticities in Table 4. For the representative bank, a one percent partial increase in the loan rate to households reduces its market share of total loans with 1.44 percent. In contrast, the demand elasticity is estimated to only -0.65 on average for loans to businesses. Both average demand elasticities are significantly different from zero at the five percent level, although less clearly so for business loans than household loans. In the business sector, some of the estimated elasticities are even positive, but insignificant. It appears that demand from businesses is less elastic than demand from households. This conclusion should, however, be interpreted with some care. As discussed in Section 2, banks may raise their lending standards when they face higher funding costs. Moreover, Maddaloni and Peydro (2011) find that banks raise their lending standards more to households than to businesses. Thus some of the estimated difference in elasticity could be a (supply-side) effect of tighter lending standards.

We do not find a significant negative effect of the risk measure s . This is not surprising, since higher risk is more likely to affect the aggregate supply of credit than individual banks' market shares – which necessarily sum to one over all bank groups (including the residual group).

$$\begin{aligned}
\sum_{i=1}^6 w_i X_{it}^H &= d_t + 0.00s - 1.44\tilde{r}^H \\
&\quad (0.15) \quad (0.43) \\
\sum_{i=1}^6 w_i X_{it}^B &= d_t + 0.05s - 0.65\tilde{r}^B \\
&\quad (0.17) \quad (0.35)
\end{aligned} \tag{12}$$

Table 4: Estimates of key parameters in the steady-state equations for market shares. Standard errors in parentheses.

Equation (X_{it}^L):	Coefficients of \tilde{r}^L		Coefficients of s	
	H	B	H	B
Bank group 1	-1.54 (.34)	-1.19 (.59)	.03 (.17)	-.11 (.24)
Bank group 2	-1.03 (.42)	.50 (.47)	.32 (.14)	-.12 (.21)
Bank group 3	-.49 (.48)	-1.50 (.54)	-.34 (.06)	-.34 (.21)
Bank group 4	-2.26 (.40)	-.76 (.38)	-.17 (.12)	.27 (.15)
Bank group 5	-1.09 (.41)	.30 (.28)	.07 (.14)	.47 (.16)
Bank group 6	-1.48 (.40)	.22 (.25)	.40 (.19)	.60 (.18)
Common estimate*	-1.44 (.43)	-.65 (.35)	.00 (.15)	.05 (.17)
P-value of Wald-test**	.07	.12	.01	.02

*Value-weighted average of six bank group-specific coefficients

**Wald test of the restriction that all parameters are equal (5 d.f.)

Examining the stationarity of the dependent and exogenous variables

The hypothesis that r_t is *not* a unit root process was considered in Raknerud et al. (2011), using both daily data and quarterly data, applying the test proposed by Choi (1994). The null hypothesis of stationarity against the alternative that r_t is a unit root process was not rejected. This result is consistent with Choi and Ahn (1999), who did *not* reject that the real interest rate is stationary, using monthly data for several countries for 1980–1991 (Norway not included). On the other hand, using an Augmented Dickey-Fuller test on quarterly NOK real interest rate data for 1986–2008, Anundsen and Jansen (2011) find evidence that both the nominal interest rate and the inflation rate are $I(1)$, but that the *real* interest rate is $I(0)$. Although we use nominal interest rates, not real interest rates, our data come from a period with inflation targeting and a low and stable inflation rate.

Next, we tested the joint stationarity of the dependent variables against the alternative that any of these time series are unit root processes (possibly cointegrated). To do so, we applied the multivariate test proposed by Choi and Ahn (1999) on the vector $(r_{it}^H, r_{it}^B, X_{it}^H, X_{it}^B)$ for each of the 6 bank groups. We used their proposed LM_I test statistic. The value of LM_I varied from 0.37 to 1.96 in our sample. Since the number of time series in each vector is four ($n = 4$) the value of the test statistic is below any of the critical values reported in Choi and Ahn (1999)¹⁰. For example, the 90% percentile of LM_I , with $n = 4$, is 2.52 (see their Table 1 b).

Goodness-of-fit We shall now assess the goodness-of-fit of our model by reporting diagnostic tests. Test statistics for skewness and kurtosis are shown in Table 5,

¹⁰We used the automatic lag truncation procedure proposed by Andrews (1991), which led to $10 \leq l \leq 14$.

Table 5: Test statistic for kurtosis and skewness

Bank Group	Test-statistic skewness (S)				Test-statistic kurtosis (K)			
	r_{it}^H	r_{it}^B	X_{it}^H	X_{it}^B	r_{it}^H	r_{it}^B	X_{it}^H	X_{it}^B
1	.50	.80	.46	-.01	1.71	.36	.45	1.51
2	.20	.58	.24	-.47	.28	.34	.79	-.29
3	.59	.88	2.15	.78	.29	1.34	.80	.16
4	2.96	1.77	1.65	.60	1.98	5.24	.32	-.07
5	.88	-.03	.27	.01	.04	.23	-.64	-.90
6	.36	1.98	-2.64	.89	.23	.04	2.61	-.66

Table 6: R-squared and Portmanteau test statistic (Q) for serial correlation in the innovations.

	r_{it}^H	r_{it}^B	X_{it}^H	X_{it}^B
R^2	.981	.989	.986	.977
Q	321	296	341	322
sd^2	360	360	360	360
n^*	65	65	62	62
$d.f.$	295	295	298	298
P-value	.14	.47	.03	.16

whereas results of portmanteau tests for serial correlation (based on the Q -statistic) are shown in Table 6. When all parameters are known, the asymptotic distribution of Q is known to be χ^2 with d^2s degrees of freedom, where d is the number of equations and s is the number of lags used in the calculation of Q (see Reinsel, 1993). To use these tests in our situation, certain adjustments of standard procedures are necessary¹¹.

The R^2 reported in the first row in Table 6 is defined as $1 - \text{tr}(RSS) / \text{tr}(TSS)$, where RSS is the matrix of sum of squares of the (one-step-ahead) prediction errors, TSS is the matrix of total sum of squares and $\text{tr}(\cdot)$ denotes the trace. The results in Table 5 show that only r_{it}^H in Bank group 4 is problematic for the assumption of

¹¹The degrees of freedom must be adjusted for dependence among residuals caused by the replacement of true parameters by estimated ones. It is known in some special cases that $Q \sim \chi^2(d^2s - n^*)$, where n^* is the number of estimated parameters, except the parameters of Σ . This result holds in the case of the homogeneous SUTSE model discussed in Harvey (1989), and also in the VARMA(p, q) models, where $n^* = d^2(p + q)$. The degrees of freedom (d.f.) in Table 6 is based on the conjecture that this result is valid also in our case (with $s = 10$ - chosen based on the automatic lag truncation procedure mentioned in footnote 10).

normality, while the results in Table 6 indicate that the vector $(X_{1t}^H, \dots, X_{6t}^H)$ may violate the assumption of no serial correlation in the innovations. However, the rejection is not clear – the lowest P-value in Table 6 being 0.03.

6 Conclusion

We have used a dynamic factor model and a detailed panel data set with quarterly accounts data for all Norwegian banks to study how banks' funding costs affect their interest rates and how changes in an individual bank group's loan rate relative to the market loan rate affect its market share. In our analysis the cost of market funding was estimated by the three-month Norwegian Inter Bank Offered Rate (NIBOR). We found clear evidence of incomplete pass-through from the NIBOR rate to retail loan rates, with the loan rates increasing less than the NIBOR rate. Our estimates show that a unit increase in NIBOR leads to an approximately 0.8 increase in banks' loan rates in the long run. We also find a significant positive relation between the indicative credit spread of uncovered bonds issued by banks and loan rates. The degree of pass-through from the credit spread rate to the loan rates is estimated to be much smaller than for the NIBOR rate. The explanation for this may be that the latter pass-through effect is identified mainly from the huge variations in the credit spread immediately before and after the onset of the financial crisis, and therefore is difficult to separate from the effects of policy measures that were implemented simultaneously.

Finally, we estimate a significant negative credit demand elasticity with respect to loan rates – both for households and businesses. On average a (permanent) one percent increase in a bank's loan rate to households (for a given level of the market loan rate index) reduces its market share by 1.44 percent in the long run. The corresponding demand elasticity is estimated to -0.65 for loans to businesses. This

difference could indicate a higher degree of market segmentation in the business loan sector. However, the finding should be interpreted with some care, as banks may raise their lending standards when they face higher funding costs and the effect may be stronger for households than for businesses, as found by Maddaloni and Peydro (2011). Thus some of the estimated higher elasticity could reflect a (supply-side) effect of changes in lending standards.

References

- [1] Akram, Q. F. and C. Christophersen (2011): Norwegian overnight interbank interest rates. Staff memo 1/2011, Norges Bank.
- [2] Allen, L (1988): The determinants of bank interest margins: A note. *Journal of Financial and Quantitative Analysis*, 23, 231–235.
- [3] Andrews, L. (1991): Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimation. *Econometrica*, 59, 817–858.
- [4] Angbazo, L. (1997): Commercial bank net interest margins, default risk, interest-rate risk and off-balance sheet banking. *Journal of Banking and Finance*, 21, 55-87.
- [5] Anundsen, A.K. and E.S. Jansen (2011): Self-reinforcing effects between housing prices and credit. Evidence from Norway. Discussion papers 651, Statistics Norway.
- [6] Ashcraft, A.B. (2006): New evidence on the lending channel. *Journal of Money, Credit, and Banking*, 38, 751-75.
- [7] Bernanke, B.S., Boivin, J. and P. Elias (2005): Measuring the effects of mon-

- etary policy: A factor-augmented vector autoregressive (FAVAR) approach. *Quarterly Journal of Economics*, 120, 387–422.
- [8] Choi, I.(1994): Residual-based tests for the null of stationarity with applications to U.S. macroeconomic time series. *Econometric Theory*, 10, 720–746.
- [9] Choi, I. and B.C. Ahn (1999): Testing the Null of Stationarity for Multiple Time Series. *Journal of Econometrics*, 88, 41–77.
- [10] Corvoisier, S. and R. Gropp (2002): Bank concentration and retail interest rates. *Journal of Banking and Finance*, 26, 2155–2189.
- [11] De Bondt, G. (2002): Retail bank interest rate pass-through: New evidence at the Euro area level. ECB Working Papers, No 136.
- [12] De Graeve, F., O. De Jonghe and R.V. Vennet (2007): Competition, transmission and bank pricing policies: Evidence from Belgian loan and deposit markets. *Journal of Banking and Finance*, 31, 259–278.
- [13] Dixit, A.K. and J.E. Stiglitz (1977): Competition and optimum product diversity. *American Economic Review*, 67, 297-308.
- [14] Fabbro, D. and M. Hack (2011): The effects of funding costs and risk on banks' lending rates. Reserve Bank of Australia Bulletin, March, 35–41.
- [15] Forni, M., Hallin, M., Lippi, M. and L. Reichlin (2000): The generalized dynamic factor model: Identification and estimation. *Review of Economics and Statistics*, 82, 540–554.
- [16] Freixas, X.F and J.-C. Rochet (2008): *Microeconomics of Banking*. Cambridge, Massachusetts. MIT Press.

- [17] Hammersland, R. and C.B. Tr  e (2012): The financial accelerator and the real economy: A small macroeconomic model for Norway with financial frictions. Staff Memo 2/2012, Norges Bank.
- [18] Hannan, T. and A. Berger (1991): The rigidity of prices: Evidence from banking industry. *American Economic Review*, 81, 938–945.
- [19] Harvey, A.C. (1989): *Forecasting, Structural Time Series Models and the Kalman Filter*. Cambridge: Cambridge University Press.
- [20] Harvey, A.C. and S. J. Koopman (1992): Diagnostic checking of unobserved components time series models. *Journal of Business and Economic Statistics*, 10, 377–389.
- [21] Jappelli, T. (1993): The estimation of the degree of oligopoly power of the Italian banking sector. *Studi Economici*, 49, 47–60.
- [22] Kashyap, A.K. and J.C. Stein (1994): Monetary policy and bank lending. In: *Studies in Business Cycles* (N.G. Mankiw, ed), 29, 221–56. Chicago and London: University of Chicago Press.
- [23] Kashyap, A.K. and J.C. Stein (2000): What do a million observations on banks say about the transmission of monetary policy. *The American Economic Review*, 90, 407–428.
- [24] Maddaloni, A. and J.L. Peydr   (2011): Bank risk-taking, securitization, supervision, and low interest rates: Evidence from the Euro-area and the U.S. lending standards. *Review of Financial Studies*, 24, 2121–65, 2011.
- [25] Raknerud, A., T. Skjerpen and A.R. Swensen (2010): Forecasting key macroeconomic variables from a large number of predictors: A state space approach. *Journal of Forecasting*, 29, 367–387.

- [26] Raknerud, A, B.H. Vatne and K.J. Rakkestad (2011): How do banks' funding costs affect interest margins? Norges Bank Working Paper No. 2011/09
- [27] Reinsel, G.C. (1993): *Elements of Multivariate Time Series Analysis*. Springer, New York
- [28] Rosen, D and D. Saunders (2010): Risk factors contributing in portfolio credit risk models. *Journal of Banking and Finance*, 34, 336–349.
- [29] Saunders, A. and L. Schumacher (2000): The determinants of bank interest rate margins: An international study. *Journal of International Money and Finance*, 19, 813–832.
- [30] Stiglitz, J. E. and A. Weiss (1981): Credit rationing in markets with imperfect information. *American Economic Review*, 71, 393-410.
- [31] Stock, J.H. and M.W. Watson (2002). Forecasting using principal components from a large number of predictors. *Journal of the American Statistical Association*, 97, 1167–1179.

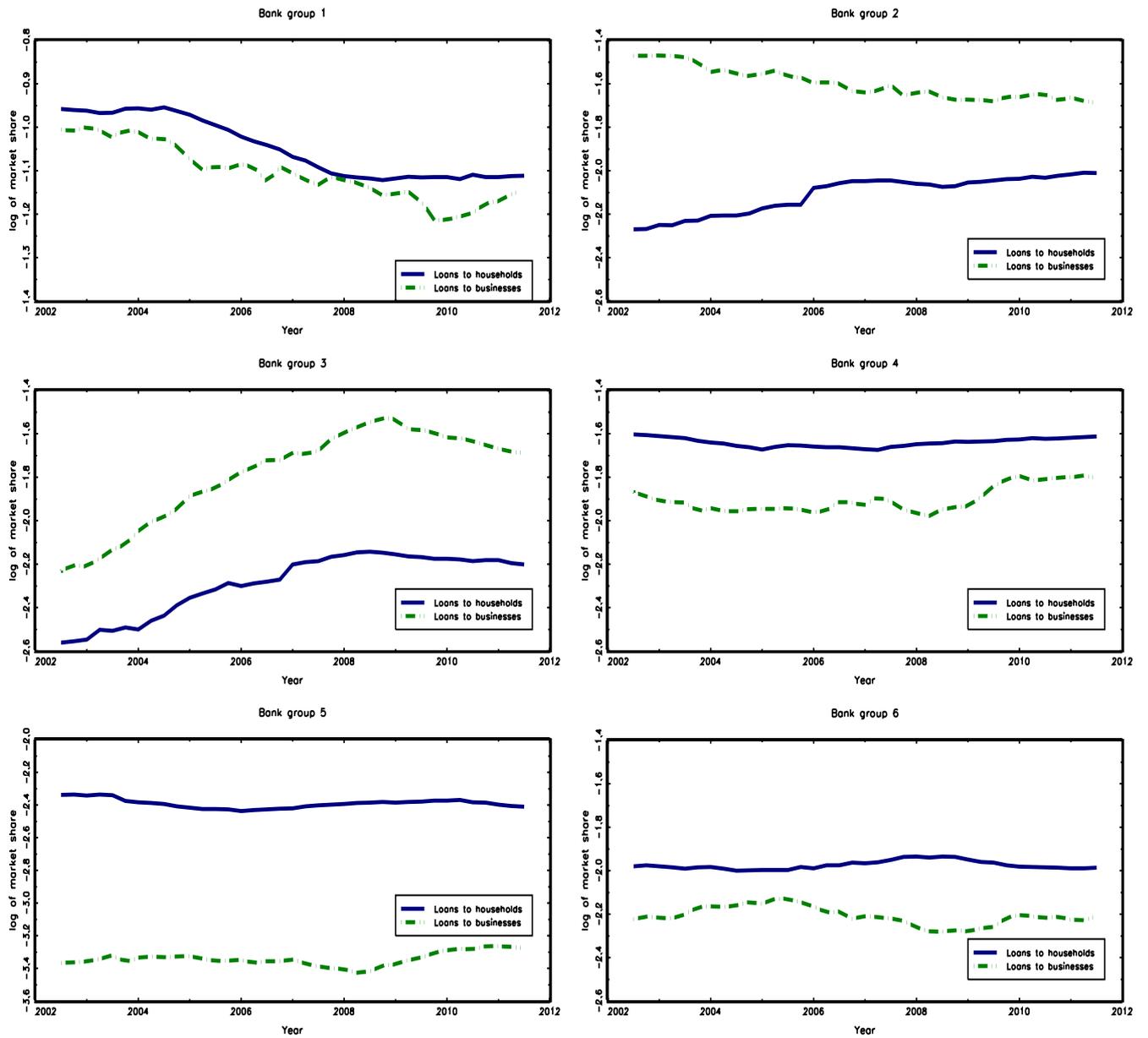


Figure 1: Logarithm of loan market shares for six bank groups: Loans to households and businesses.

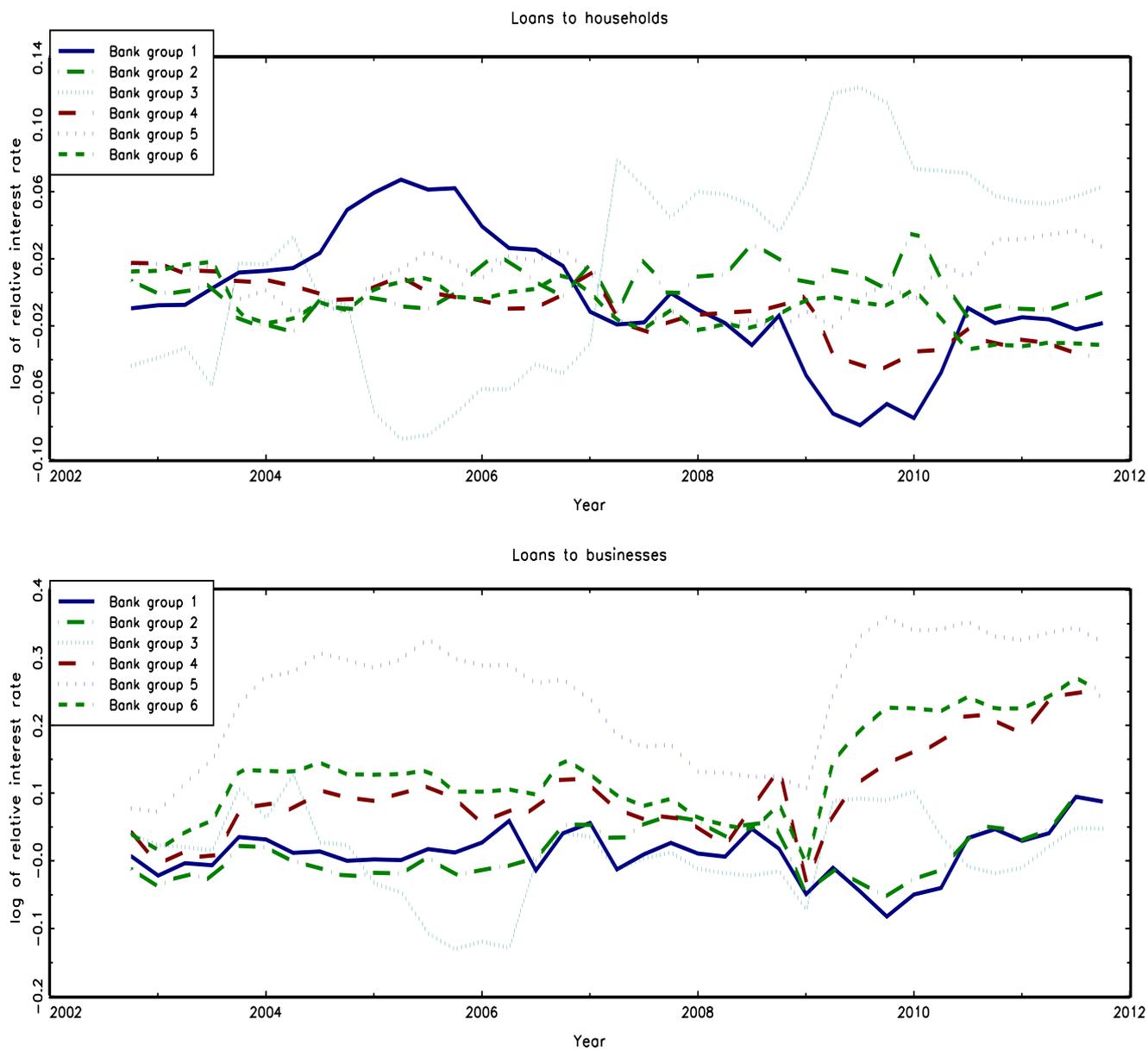


Figure 2: Logarithm of lending rate for each bank group relative to the (market) loan rate index.

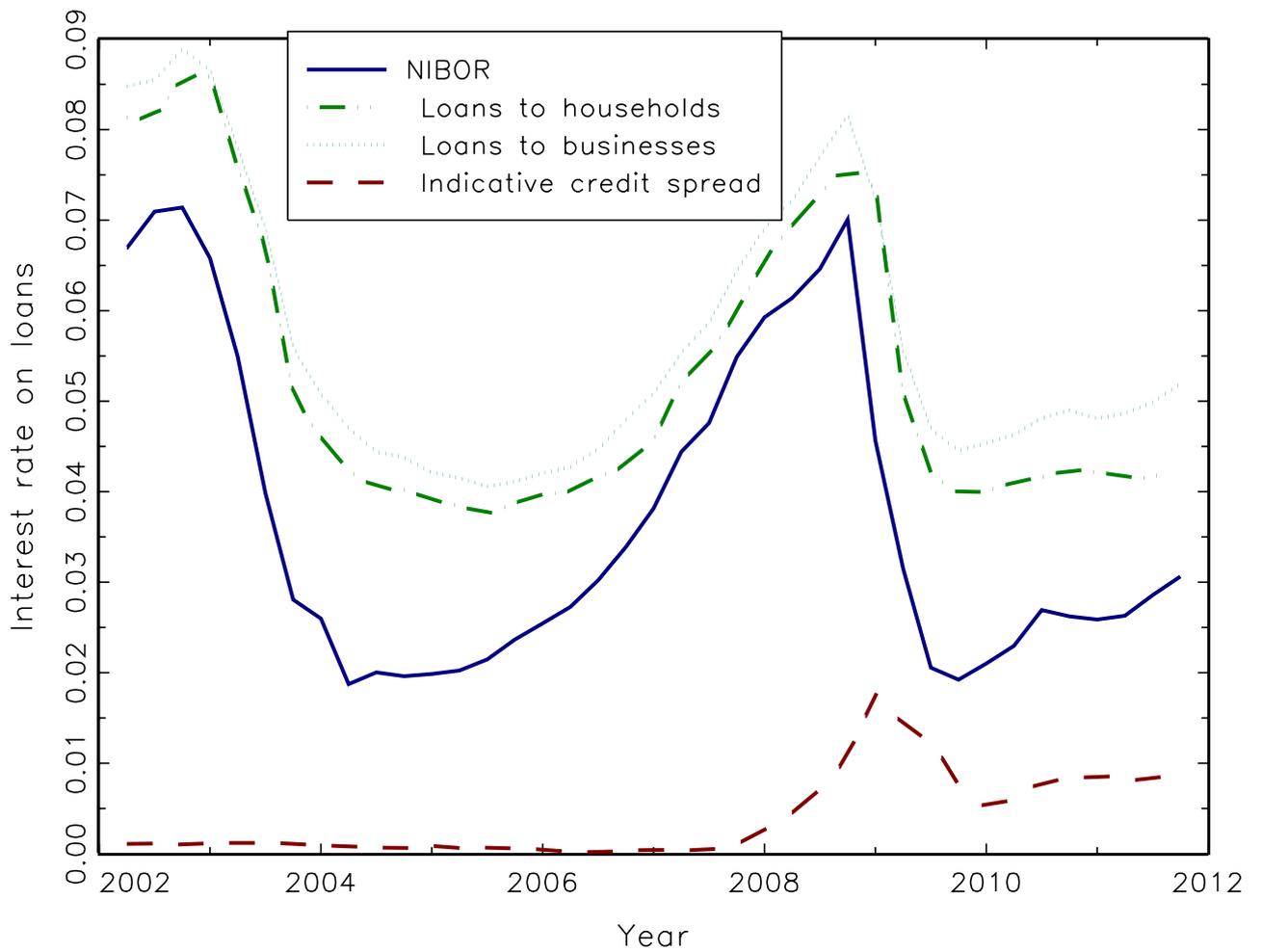


Figure 3: Three month NIBOR rates, market interest rate index for loans to firms and households, and indicative credit spread on senior unsecured bank bonds.