Discussion Papers
Statistics Norway
Research department

No. 742
May 2013

Statistics Norway

Statistisk sentralbyrå

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Evidence form Norwegian microdata

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Abstract:
In this paper, we examine two questions: i) how changes in the funding costs of banks affect retail loan rates and ii) how changes in relative loan rates between banks affect their market shares. To do so, we estimate a simultaneous system of equations model using panel data for six Norwegian bank groups. The data set consists of quarterly data for the period 2002Q1-2011Q3 and includes 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 suggest that 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 credit demand from businesses.

Keywords: credit demand, pass-through, funding costs, monopolistic competition, panel data, dynamic factor model

JEL classification: C33, E27, E43

Acknowledgements: The opinions expressed herein are those of the authors and do not necessarily reflect the views of Statistics Norway or Norges Bank. The authors would especially like to thank Farooq Akram, Sigbjørn Atle Berg, Jin Cao, Terje Skjerpen and Bjørne-Dyre Syversten for useful discussion on this research. The authors also acknowledge the helpful comments and suggestions of seminar participants at the Norges Bank and the International Conference on Computational and Financial Econometrics (CFE) held on December 1–3, 2012 in Oviedo, Spain.

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ISSN 0809-733X
Print: Statistics Norway

Sammendrag

Ved bruk av en økonometrisk modell og paneldata for seks norske bankgrupper analyserer vi to spørsmål: i) hvordan endringer i finansieringskostnader slår ut i endrede utlånsrenter og ii) hvordan endringer i renteforskjeller mellom bankene påvirker deres markedsandeler. Vårt datasett består av kvartalsdata for 2002Q1-2011Q3 og inkluderer informasjon om utlånsvolum og utlånsrenter for enkeltbankers lån til foretak og personer. Kostnaden ved markedsfinansiering er representert i vår anlayse ved 3-månders interbank renten (NIBOR) og en proxy for markedsrisiko: indikativ spread på usikrede 3-års norske bankobligasjoner. Våre resultater viser at 10 basispunkters økning i 3-månders NIBOR leder til omtrent 8 basispunkter økning i utlånsrenten. Vi finner også at etterspørselen etter kreditt i personmarkedet er mer elastisk mhp. lånerenten enn etterspørselen fra foretak.

1 Introduction

In this study, we investigate two related questions: i) how changes in the funding costs of banks affect loan rates to households and businesses and ii) how changes in relative loan rates between banks affect their market shares. While the transmission mechanism, i.e., the pass-through from market rates to retail rates, have been studied extensively in both the theoretical and empirical literature,[1] much less is known about the response of credit demand to changes in loan rates. In this analysis, we investigate both issues within a simultaneous system of equations framework. The system encompasses a theoretical model of monopolistic competition, where banks are price setters in the loan markets (i.e., Cournot competitors), but face a common funding rate. According to our theoretical model, each bank's market share (i.e., share of total loans) is a function of the ratio of its loan rate to the market loan rate, where the latter is a price index constructed from the loan rates of the individual banks.

Conventionally, the relationships between retail lending rates, loan volumes, funding costs and other (macroeconomic) variables have been examined using time-series econometric models. Typically, the focus is on aggregate demand and the supply of credit. An example is the cointegrated vector autoregressive macroeconomic model of Norges Bank (see Hammersland and Træe, 2012). However, the problem of separating the supply and demand side effects has not yet been solved within this empirical framework. An alternative approach to resolving the identification problem is to attempt to identify exogenous liquidity shocks that affect the supply side of lending—through the so-called bank-lending channel—but not the

[1] See e.g. Allen (1988), Hannan and Berger (1991), Angbanzo (1997), De Bondt (2002), De Graeve et al. (2007), and Banerjee et al. (2013).

demand side. See for example Kashyap and Stein (2000) and Ashcraft (2006).[2]

The main novelty of this paper is to consider the determinates of retail lending rates (the interest rate pass-through) and market shares simultaneously. From our theoretical model of monopolistic competition between banks, we derive exclusion restrictions, i.e., variables that affect bank retail rates, but not the demand for credit. Exclusion restrictions 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.

We restrict our attention to the microeconomic aspects of banking by analyzing the market shares of loans of individual banks (or bank groups – see below), not their volume of loans in absolute terms. Nevertheless, we are able to estimate the elasticity of demand with respect to loan rates, as well as investigate the impact of changes in funding costs, including risk premiums, on retail rates. In accordance with most empirical literature on bank interest rates (e.g., Saunders and Schumacher, 2000), our model includes an interbank market rate; the three-month Norwegian Inter Bank Offered Rate (NIBOR), as a key exogenous variable. Moreover, we measure market risk as the indicative spread between the rate on three-year senior unsecured bank bonds and the three-month Norwegian interbank rate. We can interpret this particular credit spread as the compensation required by investors for both credit and liquidity risk.

The period analyzed in this paper, from 2002Q1 to 2011Q3, includes a period of financial distress, with increased market risk premiums and a large fall in the policy rate of the Norwegian central bank. When market risk (credit and/or liquidity risk) increases, banks may restrict the loan supply at given interest rates by changing the nonprice terms for loans and/or enforcing a stricter screening of loan applicants.

[2] Kashyap and Stein (1994) provide some background and discussion of the bank-lending channel.

The Norges Bank's Survey of Bank Lending[3] confirms that this was indeed the case in Norway after 2007Q4. Thus, there may be a *direct* effect from changes in market risk to the loan supply, especially for unsecured loans.

For our empirical analysis, we utilize quarterly panel data on Norwegian banks which we aggregate into six bank groups. In the data, the average volumes and interest rates over the quarter are specified for each bank group and for various types of loans. We distinguish between loans to households and loans to corporations in the nonfinancial sector (business loans). The corresponding interest rates and loan market shares are analyzed using a dynamic factor model. The use of common dynamic factors is a parsimonious way of capturing comovements among variables, as advocated e.g., by Bernanke et al. (2005) and Forni et al. (2000). As a result, we are able to distinguish between the effect on retail rates of commonly observed variables (such as interbank market rates) and the effects of unobserved common variables (reflecting, for example, changes in bank regulations, competition, and productivity).

Our empirical framework allows us to test particular hypotheses about both the short- and long-run "steady-state" relationship between market rates (marginal funding costs) and retail rates. We also estimate the long-run elasticity of credit demand for households and corporations. Our results strongly suggest incomplete pass-through of interest rates. We estimate that a 10 basis points increase in the market rate leads to an approximately 8 basis points increase in retail loan rates. Moreover, we find that credit demand from households is more elastic with regard to the loan rate than credit demand from businesses.

The remainder of the paper is organized as follows. Section 2 describes the theoretical model of monopolistic competition between banks. Sections 3 and 4

[3] See http://www.norges-bank.no/en/about/published/publications/norges-banks-survey-of-bank-lending/

present the data and the empirical model, respectively. Section 5 discusses 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 substitution (CES) preferences over loans from different banks. Thus, we do not derive the heterogeneity between banks from primary assumptions about their location, or the distance between banks and customers, as in the Monti–Klein framework. Instead, we 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 work of Dixit and Stiglitz (1977).

First, we assume a representative agent that uses loans to finance investments or to purchase 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 $\sum_{i=1}^{N} r_i L$, where r_i is the loan rate of bank i. We assume that $L_1, ..., L_N$ enter the agent's utility function, $U(\cdot)$, as follows:

$$U(C_0, L_1, ..., L_N) = u\left(C_0, \left(\sum_{i=1}^{N}(a_i L_i)^{\rho}\right)^{\frac{1}{\rho}}\right), \ \rho < 1, \ a_i \geq 0, \tag{1}$$

where the function $u(\cdot)$ is quasiconcave and increasing in both arguments, where the first argument, C_0, is the numeraire good and the second argument is a CES loan quantity index. According to (1), 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:

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

The well-known solution is

$$x_i = x a_i^{\sigma} \left(\frac{r_i}{R}\right)^{-\sigma}, \tag{2}$$

where $\sigma = 1/(1-\rho)$ and

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

By allowing the parameters $a_1, ..., a_N$ to take different values, the demand for loans from different banks will differ, even if their loan rates are the same: $r_1 = ... = r_N$. As 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 nonprice factors that affect the demand for loans from individual banks.

For any variable z_i, define $\overline{z}$ as the geometric average of $z_1, ..., z_N$:

$$\overline{z} = \prod_{i=1}^{N} z_i^{\frac{1}{N}}. \tag{3}$$

It follows from (2) that

$$\ln(x_i) = -\sigma \ln(r_i/\overline{r}) + \alpha_i, \tag{4}$$

where

$$\alpha_i = \ln(\overline{x}) + \sigma(\ln(a_i) - \ln(\overline{a})).$$

Thus, demand depends on the relative price $r_i/\overline{r}$.

To provide loans, banks need to raise funds. We assume here that the wholesale market is the marginal source of funding and that the banks face constant marginal funding costs equal to c, i.e., regardless of the amount of funding. We assume decisions regarding loans and deposits are 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 may differ across banks and are therefore indexed i. As in Jappelli (1993) and Corvoisier and Gropp (2002), we incorporate credit risk through a fixed bank-specific default probability, μ_i. The bank's choice of loan rate is then given by the solution to an expected profit maximization problem:

$$\max_{r_i} \{(1-\mu_i)r_i - c - f_i)Q(r_i)\}, \tag{5}$$

where $Q(r_i) = xa_i^\sigma \left(\frac{r_i}{R}\right)^{-\sigma}$ expresses the bank's market share, x_i, as a function of the retail loan rate, r_i. We assume that banks take both R and x as given. The first-order condition for solving (5) is then:

$$r_i = \frac{\sigma}{(1-\mu_i)(\sigma-1)}(c+f_i). \tag{6}$$

In the limiting case when $\sigma \to \infty$, the coefficient of c in (6) tends to $1/(1-\mu_i)$.

Due to the multiplicative form of the demand function (2), the factor xa_i^σ does not enter (6). Moreover, the assumption of monopolistic competition implies that no supply curve exists for individual banks, the banks' adjustments being given solely by the markup rule (6). For a given (endogenous) interest rate r_i, the market share is determined by (2).

If the markup coefficient in (6), 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 (the less market power), the smaller the coefficient. In the (monopolistic competition) model in Hannan and Berger (1991), incomplete pass-through is a result of market power. However, as shown from (6), market power does not necessarily translate into incomplete pass-through (the markup coefficient being *less* than one). The markup coefficient will then depend on both the functional form of the demand function and on the degree of compensation for market risk—the factor $1/(1-\mu_i)$. Theoretically, a more than one-to-one adjustment of retail loan rates to changes in market rates is possible and is sometimes reported in the empirical literature (see e.g., De Bondt, 2002; Table 1, and Banerjee et al., 2013; Table 8). However, most empirical results support the view that pass-through is incomplete with regard to loan rates. Thus, we will now consider some modifications of our theoretical model.

Above we assumed that the marginal source of funding for banks 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 as follows (ignoring the risk weighting of Basel II 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, say $\bar{c}$. The marginal funding cost is now $(1-\alpha)c + \alpha\bar{c}$. If banks set marginal cost equal to marginal revenue, (6) must be modified accordingly:

$$r_i = \frac{\sigma(1-\alpha)}{(1-\mu_i)(\sigma-1)}c + \frac{\sigma\alpha}{(1-\mu_i)(\sigma-1)}\bar{c} + \frac{\sigma}{(1-\mu_i)(\sigma-1)}f_i. \tag{7}$$

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, $\bar{c}$. A discussion of the importance of the cost of equity for bank funding costs is given in Fabbro and Hack (2011). Using Australian data, they find evidence that there has been an increase in the contribution from equity costs to the total funding costs of banks during the last few years, especially with regard to business loans.

An important consequence of equation (7) is that the markup coefficient may be either less than or larger than one when demand is infinitely elastic. In the latter case, the coefficient becomes $(1-\alpha)/(1-\mu_i)$. Thus, from the degree of pass-through we cannot 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 an oversimplified view of the transmission mechanism. Obviously, other factors may also affect retail rates.

First, there is the possibility of adverse selection in that an increase in the retail rate will attract riskier borrowers and thereby increase the risk of default (thus μ_i could depend on r_i). In that case, banks are facing a trade-off: they have the incentive to raise the lending rate as a risk premium, but are restrained by the rising probability of default. In the Stiglitz and Weiss (1981) model, banks do not fully pass all of the increase in the market rate to their retail loan rates. Instead, loan rates are sticky upwards and credit supply rationed.

Second, other types of risk, like liquidity and interest rate risk, may also be taken into consideration. Liquidity risk is the most important. According to the 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.[4] Interest rate risk takes place 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 enter into interest rate swaps to achieve a level of variable rate exposure that matches their variable rate loans.

Third, increased risk (as measured e.g., by indicative spreads) may lead to a tightening of credit standards to better screen borrowers. Riskier projects may face higher collateral requirements and shorter contractual maturities, or loan applications may just be turned down. While it is difficult to measure (and disentangle)

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

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.

Given the stylized character of our theoretical model, we will not formally test the assumptions underlying it below. We instead use it as guidance for the operationalization and interpretation of results and the choice of functional form.

3 Data

Our sample consists of the balance sheet (accounts) data of Norwegian banks from 2002Q1 until 2011Q3 as compiled by Statistics Norway.[5] The bank-level data are 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. Introduced into Norway in June 2007, OMFs have become an important source of funding for Norwegian financial services groups and banking alliances.[6] Table 1 provides key statistics for the seven bank groups. The last group is a residual and is not 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 in March 2007. The sample is constructed on the basis of the bank structure prevailing at the end of the estimation period. For example, the time series for the DNB group 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 the banks. Small national banks tend to have more deposits than foreign or large national banks, while the latter tend to rely more on market funding. For example, Terra-Gruppen, which

[5]See http://www.ssb.no/skjema/finmark/rapport/orbof/ (in Norwegian).

[6]See 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

is a group of small banks, had the highest average ratio (42 percent) of household deposits over total loans during the period 2001–2010. In contrast, the two foreign bank groups had the lowest ratio of household deposits to loans (18 percent) while the largest bank group, DNB, had a ratio of 29 percent.

Figure 1 plots the logs of the market shares for each of the first six bank groups. Figure 2 depicts the corresponding graphs for the log of the relative loan rate of each bank group (i.e., relative to the market loan rate index). As shown, there is considerable persistence in both the market shares and interest rate differentials between the bank groups over time. Nonetheless, we observe some striking patterns. For example, Bank Group 1 displays a generally 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 appears to have lost a considerable share of its initial market position to Bank Group 3. We also observe considerable interest rate differences between these bank groups with regard to household loans, with Bank Group 3 generally having lower rates until 2007, but higher rates thereafter. From Figures 1 and 2 we discern no clear connection between market shares and relative loan rates.

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

Table 1: **Descriptive statistics for the 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	35	65	26
2. Subsidiaries of foreign banks[2]	13	13	18	13	57	34
3. Branches of foreign banks[3]	14	11	19	10	54	39
4. SpareBank1-alliansen[4]	14	19	16	19	68	24
5. Terra-Gruppen[5]	5	9	4	6	76	14
6. Other savings banks[6]	9	13	11	13	70	24
7. Other commercial banks[7]	2	3	3	4	69	19

Source: Norges Bank

[1]DNB Bank, Nordlandsbanken, DNB Boligkreditt and DNB Næringskreditt

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

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

[4]SpareBank 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

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

[6]Sparebanken 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

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

The three-month effective NIBOR reported by Norges Bank is a proxy for the cost of long- and medium-term market financing. Figure 3 illustrates the behavior of some of the key rates. 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.

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 case of a fixed rate, a Norwegian bank typically enters into an interest rate swap to achieve a level of variable rate exposure that matches its variable rate loans. The bank 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.[7]

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 market 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.

[7] For examples of bank bonds with varying maturity and with interest payments equal to the three-month NIBOR plus a fixed credit spread, see https://www2.sparebank1.no/portal/1001/3_privat?_nfpb=true& _pageLabel=page_privat_innhold&aId=1201861729341

4 The empirical model

We now formulate an empirical model that encompasses the main features of the theoretical model presented in Section 2. As discussed, we distinguish between loans to businesses (B) and loans to 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 $i = 1, ..., 6$, 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 disaggregated interest rates, where the weights are taken from the outgoing balance in the bank accounts. The corresponding 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, 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 own demand for or supply of credit in the interbank market. The rationale behind this assumption is that (major) banks can borrow and lend Norwegian krone (NOK) through the foreign exchange rate markets; such as the NOK–US dollar (USD) exchange swap market. Covered interest rate parity implies that the NIBOR is determined by international lending and swap exchange rates, which are exogenous to individual Norwegian banks.[8] 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 (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 dif-

[8] For an example, see equation (1) in Akram and Christophersen (2011): http://www.norges-bank.no/upload/publikasjoner/staff%20memo/2011/staff_memo_0111.pdf

ferently to exogenous shocks. Moreover, the econometric model incorporates bank group-specific parameters to allow for heterogeneity with regard to the bank responses to the exogenous variables. Finally, the model incorporates common shocks to account for comovements in the different rates from unobserved (common) factors.

We model the individual retail rates as univariate autoregressive (AR) processes, augmented with common dynamic factors. Our approach then lies in the tradition of multivariate structural time series models.[9] Specifically, we assume that, for $L = B, H$ (businesses and households):

$$r_{it}^{L} = \mu_i^L + \alpha_{i,0}^L r_t + \alpha_{i,1}^L r_{t-1} + \gamma_i^L s_t + \sum_{j=1}^{p_i} \phi_{ij}^L r_{i,t-j}^L + \sum_{k=1}^{m} \theta_{ik}^L f_{kt} + e_{it}^L, \tag{8}$$

where μ_i^L is a bank group- and interest rate-specific fixed effect, the α parameters capture the effects of the NIBOR by allowing both the current NIBOR, r_t (through $\alpha_{i,0}^L$), and the lagged NIBOR, r_{t-1} (through $\alpha_{i,1}^L$), to affect r_{it}^L. We incorporate a single lag 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 AR parameters ϕ_{ij}^L, $j = 1, ..., p_i$, determine how the effects of a shock in any of the exogenous variables evolve over time. The number of lags, p_i, is allowed to differ from bank group to bank group. The unobserved stochastic terms consist of m dynamic factors, $f_{1t}, ..., f_{mt}$, which pick up the dependencies across banks from common unobserved variables (e.g., the effects of the business cycle, credit market regulations, and competition). Both the number of lags, p_i, and the number of factors, m, are chosen by means of Akaike's information criterion (see below). Finally, e_{it}^L is an idiosyncratic error term assumed to be independent across banks (i) and over time (t).

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

Market shares of total loans Analogously to (8), we assume that

$$\ln(x_{it}^L) = \nu_i^L + \beta_{i,0}^L \ln(r_{it}^L/\overline{r}_t^L) + \beta_{i,1}^L \ln(r_{i,t-1}^L/\overline{r}_{t-1}^L) + \kappa_i^L s_t + \sum_{j=1}^{q_i} \psi_{ij}^L \ln(x_{i,t-j}^L) + \sum_{k=1}^{m} \zeta_{ik}^L f_{kt} + \varepsilon_{it}^L, \tag{9}$$

where (for sector $L = H, B$) the dependent variable is $\ln(x_{it}^L)$, that is, the log of bank i's market share (share of total loans in sector L) and $\overline{r}_t^L$ is the (market) loan rate index to sector L. Moreover, ν_i^L is a fixed effect, and $\beta_{i,0}^L$ and $\beta_{i,1}^L$ capture the direct effects of the current and lagged value of $\ln(r_{it}^L/\overline{r}_t^L)$ on the dependent variable, cf. (4). The credit spread measure, s_t, is allowed to affect $\ln(x_{it}^L)$ through the parameters κ_i^L. Thus, we allow for a direct effect of the credit spread on loan volumes (and thus market shares) through the nonprice terms of loans, as explained above. Note that (9) is a dynamic equation, with q_i lags of the dependent variable, $\ln(x_{i,t-j}^L)$, entering on the right-hand side of (9), with the corresponding AR parameters ψ_{ij}^L. Finally, the loading coefficients ζ_{ik}^L have the same interpretation as the θ_{ik}^L in (8).

For each bank group, the vector of dependent variables comprises $(r_{it}^B, r_{it}^H, \ln(x_{it}^B), \ln(x_{it}^H))$. 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:

$$f_{kt} = \varpi_k f_{k,t-1} + \eta_{kt},\ \eta_{kt} \sim \mathcal{IN}(0,1); k = 1, ..., m. \tag{10}$$

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 using a version of the maximum-likelihood algorithm described in Raknerud *et al.* (2010).

For identification, it is a crucial exclusion restriction that the NIBOR, r_t, enters (8), but not (9). This restriction is motivated by the theoretical model in Section 2.

Another restriction is that the vector of error terms is assumed to be uncorrelated across bank groups. The rationale for the latter assumption is that common shocks across banks are captured by the dynamic factors. Both these restrictions contribute to exogenous variation in the endogenous explanatory variable $\ln(r^L_{it}/\overline{r}^L_t)$ and hence to identification.

Partial effects Our econometric framework allows us to disentangle both the short- and long-run partial effects of changes in the exogenous variables on the dependent variables. First, we are most interested in the effects of the changes in the market rate on retail lending 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

$$r^L_{it} = \frac{\mu^L_i}{1-\sum_{j=1}^{p_i}\phi^L_{ij}} + \left(\frac{\alpha^L_{i,0}+\alpha^L_{i,1}}{1-\sum_{j=1}^{p_i}\phi^L_{ij}}\right) r + \left(\frac{\gamma^L_0}{1-\sum_{j=1}^{p_i}\phi^L_{ij}}\right) s + d^L_t \qquad (11)$$

is the corresponding steady-state equation. The coefficients of r and s in (11) determine the long-run relation between retail rates and permanent (or persistent) levels of the exogenous variables r_t and s_t, whereas d^L_t captures the effects on retail rates of the present and lagged dynamic factors, f_{js}, $s \leq t$. We interpret equation (11) as the empirical counterpart of (7), 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 the long-run (equilibrium) relations.

A similar steady-state equation with respect to the log market share $\ln(x^L_{it})$, given a permanent value of the retail rate $r^L_{it} = r^L_i$ and $\overline{r}^L_t = \overline{r}^L$, is given by

$$\ln(x^L_{it}) = \frac{\nu^L_i}{1-\sum_{j=1}^{q_i}\psi^L_{ij}} + \left(\frac{\beta^L_{i,0}+\beta^L_{i,0}}{1-\sum_{j=1}^{q_i}\psi^L_{ij}}\right) \ln(r^L_i/\overline{r}^L) + \left(\frac{\kappa^L}{1-\sum_{j=1}^{q_i}\psi^L_{ij}}\right) s + \delta^L_t, \qquad (12)$$

where δ^L_t is derived in a similar way as d^L_t. Equation (12) is the empirical counterpart of (4). Thus, the coefficient of $\ln(r^L_i/\overline{r}^L)$ can be interpreted as the elasticity of substitution: $-\sigma$. According to the theoretical model in Section 2, this coefficient should be negative and equal across the different bank groups.

5 Results

Dynamic specifications Before performing statistical tests, assessing estimation uncertainty and interpreting results, it is important to verify whether the variables of interest are stationary. Our maintained hypothesis is that the vector of dependent variables, $(r_{it}^H, r_{it}^B, \ln(x_{it}^H), \ln(x_{it}^B))$, as well as the NIBOR, r_t, are $I(0)$ processes. These assumptions are formally tested in the Appendix and *not* rejected. Consistent with this, all the estimated lag polynomials $1 - \sum_{j=1}^{p_i} \phi_{ij}^L \xi^j$ $(L = H, B)$ and $1 - \sum_{j=1}^{q_i} \psi_{i1}^L \xi^j$, where ξ 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, m, was set equal to four, while the number of lags in the AR(p_i) and AR(q_i) equations is either two or three. We made all of these decisions by applying Akaike's information criterion.[10] An assessmen of the goodness-of-fit of our chosen model is provided in the Appendix, which reports R^2 and tests for skewness, kurtosis and serial correlation in the residuals.

Table 2 displays the estimated sum of the AR 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 (11) and (12), respectively. If any such sum is close to one, the corresponding retail rate, r_{it}^L, or log market share, $\ln(x_{it}^L)$, is a near unit-root (integrated) process. The main impression obtained from these estimates is that the $\ln(x_{it}^L)$ processes are highly autocorrelated. In fact, the processes for Bank Groups 1 and 3 appear to be very close to unit-root processes, i.e., to having $\sum_j \psi_{ij}^L = 1$. Market shares thus adjust slowly to changes in relative loan rates, and much more slowly than changes in retail

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

rates to changes in the NIBOR. The retail rates, on the other hand, are clearly not unit root processes, but adjust quickly to exogenous shocks. In fact, almost all adjustment is completed within the same and next quarter of the shock.

The Wald tests in Table 2 reveal significant bank-specific heterogeneity in the AR dynamics. The hypothesis that the sum of the AR coefficients is equal across bank groups is rejected at the 1 percent level for the market shares and at the 5 percent level for retail rates.

Table 2: **Estimates of the sum of the AR parameters for each bank group***

	$\sum_{j=1}^{p_i} \phi_{ij}^L$		$\sum_{j=1}^{q_i} \psi_{ij}^L$	
Equation:	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 for Wald test**	.04	.03	.007	.002

*Standard errors in parentheses are obtained by the delta method
**Wald test of the restriction that all six bank groups have equal sum (5 d.f.)

Estimates for the retail rate equations Our focus is now on the estimated long-run relations. Table 3 provides the estimates of the coefficients of the long-run retail rate equations (11) for each individual bank group as well as for the representative 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 of Table 1, respectively). For the representative bank, the estimated coefficient of r in the steady state is close to 0.8, and is significantly less than one for both 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 than for the other bank groups. A formal test of whether all the steady-state coefficients of r are equal across all of the 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 small magnitude of the estimated coefficient for the NIBOR indicates that loans from different banks are considered close substitutes.

We now turn to the coefficients of the indicative spread, s, in the steady-state retail rates equations. 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 instead suggest that a permanent unit increase in the credit spread leads to about a one-third increase in the business loan rate in the long run. For households, this estimate is somewhat lower at 0.23.

The estimates of the main coefficients of the aggregate equilibrium retail rate equations are depicted in (13):

$$\begin{aligned} \sum_{i=1}^{6} w_i r_{it}^{H} &= \mathrm{d}_t + \underset{(0.03)}{0.77} r + \underset{(0.06)}{0.23} s \\ \sum_{i=1}^{6} w_i r_{it}^{B} &= \mathrm{d}_t + \underset{(0.03)}{0.81} r + \underset{(0.08}{0.30} s. \end{aligned} \tag{13}$$

The estimated degree of pass-through in (13) is much smaller for the spread, s, than for the NIBOR, r. Thus, the marginal cost of market funding cannot be written simply as the sum of r_t and s_t. One explanation for this finding may be that 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 that it is difficult to separate the pass-through effects from the effects of other events that took place at the same time. This is illustrated in Figure 3, which shows that the variation in funding costs prior to 2008 was largely determined by the NIBOR. However, from 2008Q1 to 2008Q4, the spread, s_t, increased dramatically, and by the end of 2011 was still much higher than its pre-2008 level. Moreover, a marked reduction in the policy rate of the Norwegian central bank led to a sharp fall in the NIBOR. The combined effect is that from 2008Q2 we observe a distinct fall in deposit margins relative to NIBOR (not depicted) and an (offsetting) increase in the margins of loans to households (relative to NIBOR). The latter effect is clearly visible in Figure 3.

Table 3: **Estimates of the key parameters in the steady-state equations for retail rates**

	Coefficients of r		Coefficients of s	
Equation:	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 for 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 the demand elasticities The estimates of the value-weighted average elasticity of demand, the coefficient of $\ln(r_i^L/\bar{r}^L)$ in (14), show that there is an overall negative relation between the retail loan rates and market shares in both sectors. We confirm this using the estimates of the individual demand elasticities in Table 4. For the representative bank, the estimates are shown in equation (14):

$$\sum_{i=1}^{6} w_i \ln(x_{it}^H) = \mathrm{d}_t - \underset{(0.43)}{1.44} \ln(r_i^H/\bar{r}^H) + \underset{(0.15)}{0.00} s$$
$$\sum_{i=1}^{6} w_i \ln(x_{it}^B) = \mathrm{d}_t - \underset{(0.35)}{0.65} \ln(r_i^B/\bar{r}^B) + \underset{(0.17)}{0.05} s. \tag{14}$$

We can see that a one percent partial increase in the loan rate to households reduces the market share of total loans by 1.44 percent. In contrast, the demand elasticity is estimated to be 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, the estimated elasticities are even positive for some of the bank groups, albeit statistically insignificant. It thus appears that credit demand from businesses is less elastic than credit 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 Peydró (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, as higher risk is more likely to affect the aggregate supply of credit than the market shares of individual banks, which necessarily sum to one over all bank groups (when we include the residual bank group).

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

	Coefficients of $\ln(r_i^L/\bar{r}^L)$		Coefficients of s	
Equation:	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 the six bank group-specific coefficients

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

6 Conclusion

We have used a dynamic factor model and a detailed panel data set with quarterly accounts data for all Norwegian banks to examine how the funding costs of banks 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, we proxied the cost of market funding using the three-month NIBOR. We find clear evidence of incomplete pass-through from the NIBOR to retail loan rates, with loan rates increasing less than the NIBOR. Our estimates show that a 10 basis point increase in NIBOR leads to an approximately 8 point increase in bank 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. The explanation for this may be that the latter pass-through effect is poorly identified. The credit spread was very low and almost constant until the onset of the financial crisis. It is therefore difficult to separate the effects of increased credit spread from the effects of policy measures that were simultaneously implemented to reduce bank funding costs, e.g., the introduction of covered bonds, which allowed banks to fund mortgage loans more cheaply.

Finally, we estimate a significantly negative credit demand elasticity with respect to loan rates for both 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. We estimated the corresponding demand elasticity to be –0.65 for loans to businesses. This difference could indicate a higher degree of market segmentation in the business loan sector. However, this finding should be interpreted with some care as banks may raise their lending standards when they face higher funding costs and this effect may be stronger for households than for businesses. Thus, some of the higher estimated elasticity for household loans could reflect the (supply-side) effects of changes in lending standards.

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Appendix: Test of stationarity and goodness-of-fit

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 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 over the period 1980–1991 (Norway not included). On the other hand, using an augmented Dickey–Fuller test on quarterly NOK real interest rate data for the period 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 are 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 this, we applied the multivariate test proposed by Choi and Ahn (1999) on the vector $(r_{it}^H, r_{it}^B, \ln(x_{it}^H), \ln(x_{it}^B))$ 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. Given 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).[11] For example, the 90% percentile of LM_I, with $n = 4$, is 2.52 (see their Table 1b).

Goodness-of-fit To assess the goodness-of-fit of our model we now report some diagnostic tests. Table 5 provides the test statistics for skewness and kurtosis, while the results of the portmanteau tests for serial correlation (based on the Q statistic)

[11] We used the automatic lag truncation procedure proposed by Andrews (1991), which led to $10 \leq \#\text{lags} \leq 14$.

Table 5: **Test statistics for skewness and kurtosis**

	Test statistic for skewness (S)				Test statistic for kurtosis (K)			
Bank Group	r_{it}^H	r_{it}^B	$\ln(x_{it}^H)$	$\ln(x_{it}^B)$	r_{it}^H	r_{it}^B	$\ln(x_{it}^H)$	$\ln(x_{it}^B)$
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 (Q) test statistic for serial correlation in the innovations**

	r_{it}^H	r_{it}^B	$\ln(x_{it}^H)$	$\ln(x_{it}^B)$
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

Note: s, d and $n^*$are defined in fn. 11

are shown in Table 6. When all parameters are known, the asymptotic distribution of Q is known to be χ^2 with $d^2 s$ 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 context, certain adjustments to the standard procedures are necessary.[12]

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 the sum of squares of the (one-step-ahead) prediction errors, TSS is the matrix of the 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 normality, while the results in Table 6 indicate that the vec-

[12]The degrees of freedom must be adjusted for dependence among residuals caused by the replacement of the true parameters by the estimated parameters. It is known in some special cases that $Q \sim \chi^2(d^2 s - n^*)$, 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 are based on the conjecture that this result is also valid in our case (with $s = 10$ – chosen using the automatic lag truncation procedure mentioned in Footnote 11.

tor $(\ln(x_{1t}^H), ..., \ln(x_{6t}^H))$ 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.

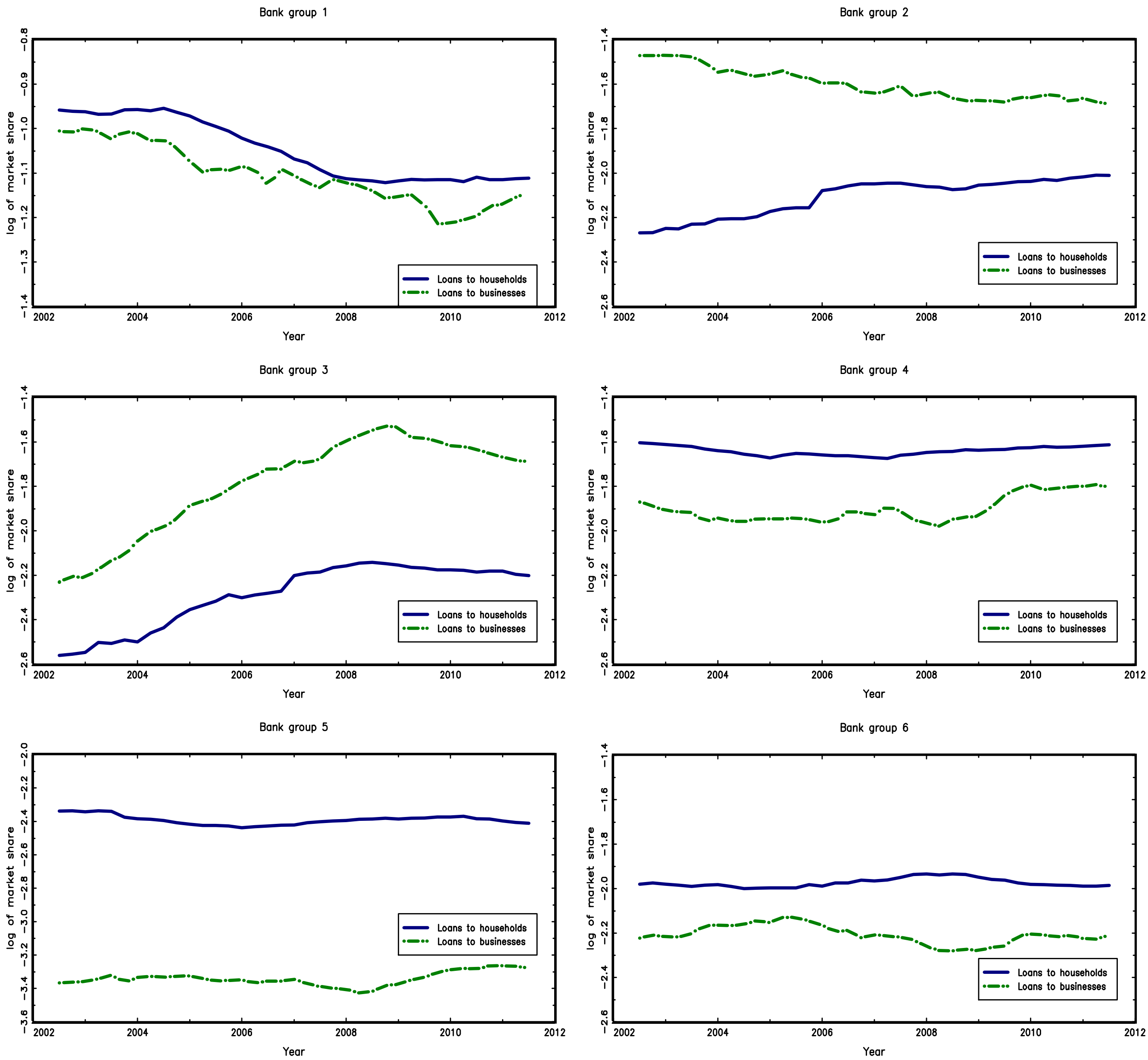

Figure 1: Logarithm of loan market shares for six bank groups: Loans to households and businesses.

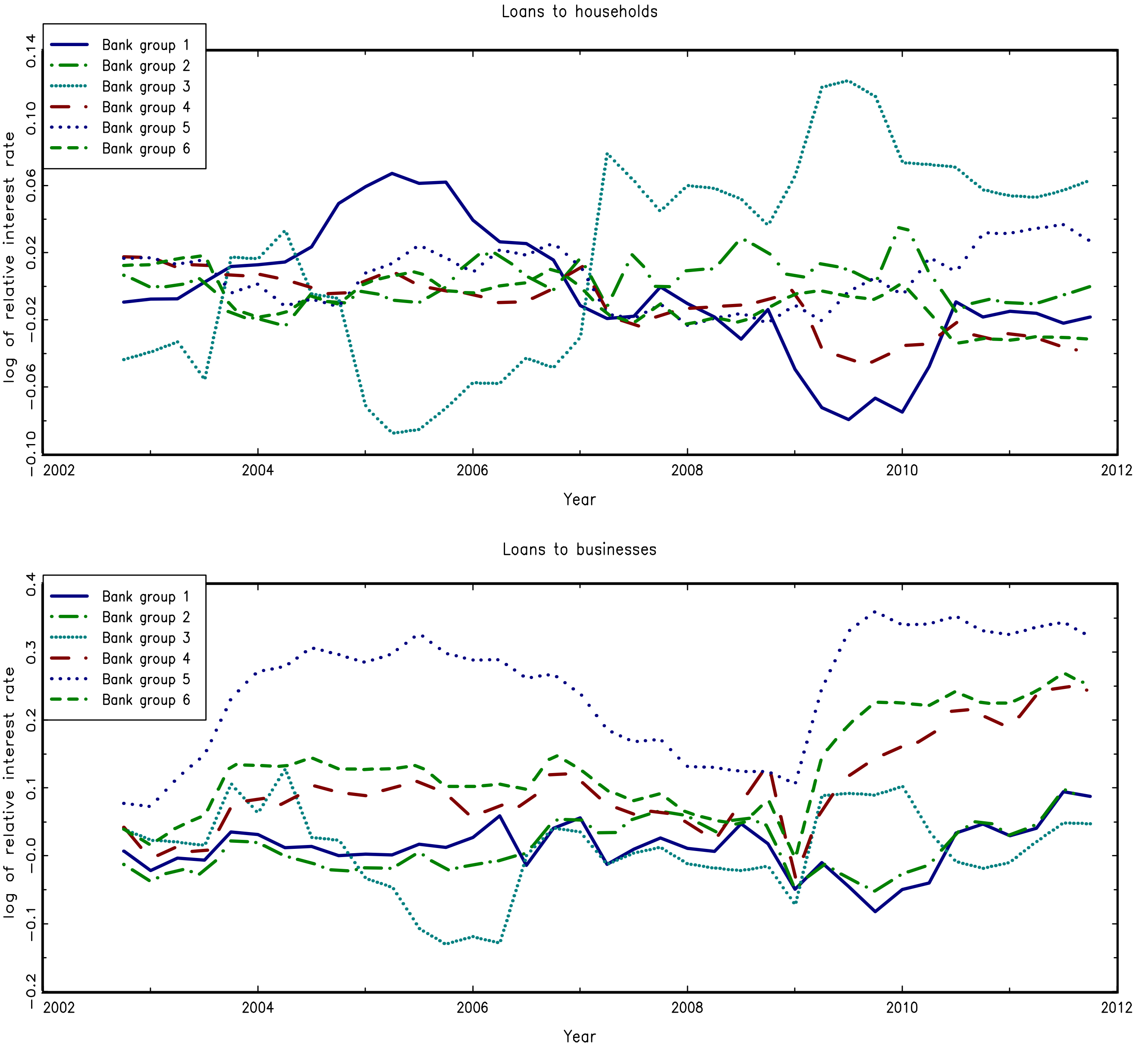

Figure 2: Logarithm of lending rate for each bank group relative to the (market) loan rate index.

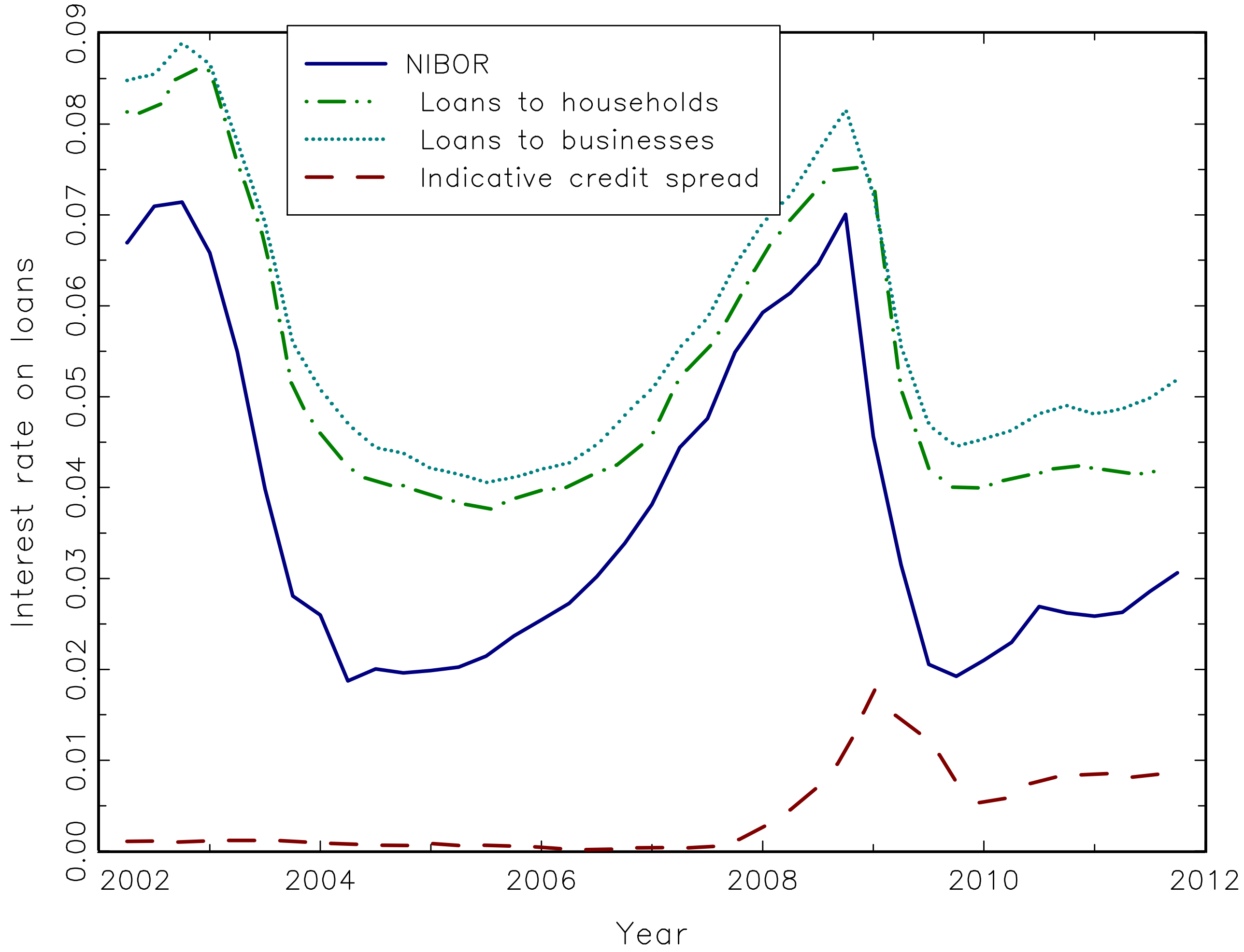

Figure 3: Three month NIBOR rates, market interest rate index for loans to firms and households, and indicative credit spread on senior unsecured bank bonds.

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ISSN 0809-733X

Statistisk sentralbyrå
Statistics Norway

Design: Siri Boquist