The Interest Rate Pass–Through in the Euro Area During the Global Financial Crisis *

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Abstract

This paper uses panel vector autoregressive models and simulations of an estimated DSGE model to explore the reaction of Euro–area banks to the global financial crisis. We focus on their interest–rate setting behavior in response to standard macroeconomic shocks. Our main empirical finding is that the pass–through from changes in the money market rate to retail bank rates became significantly less complete during the crisis. Model simulations show that this result can be well explained by a significant increase in the frictions that the banks' business is subject to.

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"Between October 2008 and May 2009 (...) the ECB lowered the interest rate on its main refinancing operations by 325 basis points. Obstacles in the transmission process were, however, threatening to prevent this very accommodative stance of monetary policy from being passed on to lending conditions for households and non–financial corporations." (European Central Bank, 2011, p. 55)

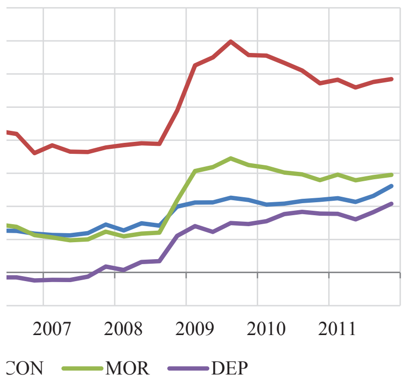
1 Introduction

The European Central Bank (ECB) vigorously cut its policy rates in response to the global financial crisis. However, banks in the Euro Area only partly passed–through the lower refinancing costs to the interest rates charged on loans to their customers and offered on deposits. As a result spreads of retail bank interest rates over money market rates sharply increased from mid–2008 on and remained on this high level since then (see Figure 1). Being concerned by the impediment of the transmission of monetary policy the ECB implemented several additional unconventional monetary policy measures (see European Central Bank, 2011, for a summary).[1]

There are at least three possible explanations for the increase in interest rate spreads. First, the macroeconomic shocks that occurred during the financial crisis could have happened to be much larger than they typically were in the pre–crisis period. Larger shocks would result in a more pronounced response of the economy including a larger increase in interest rate spreads. Second, the financial crisis could have been characterized by an increase in the relative importance of certain shocks that only play a minor role in normal times. A prominent example would be loan supply shocks. In empirical work, they are typically identified as a disturbance inducing an increase in interest rate spreads. Loan supply shocks are found to have played an important role during the global financial crisis (see e.g. Helbling et al., 2011; Hristov et al., 2012). Finally, the

[1] In the Euro area retail rates play an important role in the transmission of monetary policy, since borrowing and lending take place predominantly through the intermediation of the banking sector, contrary to some major economies where securities markets are the main funding source of the real sector. Over the period 2004-08 bank financing constituted around 75% of total external financing by non–financial corporations in the euro area and less than 50% in the United States (European Central Bank, 2010, p. 62).

Figure 1: Spreads of Euro Area retail bank interest rates over money market rates



Notes: Data are taken from the European Central Bank. The money market rate is the EONIA, which is the average of overnight rates for unsecured euro area interbank lending. The retail lending rates refer to interest rates on new business loans to non–financial corporations (‘NFC’), new loans to households for consumption (‘CON’) and new loans to households for house purchases (‘MOR’). The deposit rates are interest rates agreed on new deposits from private households (‘DEP’). The maturities are an average over all maturities.

financial crisis could have led to a fundamental change in the propagation of macroeconomic shocks. Economic agents and in particular banks could have altered their decision–making process, implying a change in the structural parameters of the economy.

This paper provides evidence on the latter explanation for the increase in interest rate spreads. Thus, it contributes to the ongoing debate of whether the transmission of monetary policy through private bank rates has been systematically distorted since the onset of the financial crisis in 2008. We use panel vector autoregressive (VAR) models for the Euro Area member countries to explore how banks adjusted their retail rates in response to changing money market rates. We focus on the interest rate pass–through during the period from 2003–2011 by considering standard macroeconomic shocks, namely a monetary policy shock, an aggregate supply shock and an aggregate demand shock. Following Uhlig (2005), Canova and De Nicolò (2002) and Peersman (2005), we identify these shocks by imposing sign restrictions. We split our sample into a pre–crisis period from 2003–2007 and a crisis period from 2008–2011 and compare the response of the interest rate spreads to shocks of the same (unit) size. Our main finding is that the pass–through from changes in the money market rate to retail bank rates became significantly less complete during the crisis. Banks seemed to be more reluctant in lowering their loan and deposits rates in the period 2008–2011, which explains a large part of the increase in interest rate spreads and provides a rationale for the ECB's unconventional policy measures.

We then provide some theoretical explanations for the decline of the pass–through from money market rates to bank retail rates. Our analysis is inspired by some recent work on New Keynesian dynamic stochastic general equilibrium (DSGE) models, which has emphasized the importance of banks in business cycle fluctuations (Curdia and Woodford, 2010; Dib, 2010; Gerali et al., 2010; Gertler and Karadi, 2011; Kollmann et al., 2011). In particular, we resort to the model of Gerali et al. (2010). The framework incorporates several features of the banking business, which are important for understanding the dynamics of loans and deposits, retail bank interest rates, risk premia and money market rates. In this model banks provide collateralized loans to the private sector and collect deposits from households in an environment of monopolistic competition. Their business is constrained by costs of maintaining an adequate bank capital position and costs related to the adjustment of retail rates. For all these frictions we provide evidence that the financial crisis has increased their altereness. This observation allows us to specify two calibration schemes – the first corresponding to the time prior to the outbreak of the financial crisis and the second, mimicking important features of loan markets thereafter. We simulate the response of the interest rate spread to the same macroeconomic shocks as identified in the VAR analysis for each regime. It turns out that theory replicates our main empirical findings quite well as the model implies a lower degree of policy rate pass–through during the financial crisis. In particular, the more distorted transmission of monetary policy via retail bank rates in the time after 2008 might be attributable to lower average loan–to–value ratios, higher costs associated with restoring the soundness of bank capital, weaker competition among banks in loan markets, and a higher degree of interest rate stickiness.

Our work contributes to the literature in several ways. First, we are among the first to discuss the impact of the global financial crisis on the interest rate pass–through in the Euro Area. We are aware of only one contribution to the literature by the European Central Bank (2009), which attempts to assess the retail bank interest rate pass–through during the global financial crisis by means of a static forecasting exercise. In contrast to our paper the authors come to the conclusion that at least up to mid–2009 the bank interest rate pass-through has worked relatively well.

Second, we use an identified VAR model for the Euro Area to analyze the retail bank interest rate pass–through to macroeconomic shocks. Most of the literature in this area resorts to error correction models, which comprise different bank retail rates next to money market rates and additional explanatory variables. The papers by Cottarelli and Kourelis (1994), Mojon (2000), Toolsema et al. (2001), Sander and Kleimeier (2004), De Bondt (2005), Kok Sørensen and Werner (2006), Kleimeier and Sander (2006) and Kwapil and Scharler (2010) largely confirm that retail bank rates in Europe react sluggishly to changes in money market rates.

Third, by investigating the response of the credit spread to structural shocks we seek to shed light on the manner how banks should be integrated into DSGE models. There seems to be some consensus in the empirical literature that bank interest rate spreads are countercyclical (see e.g. Aliaga-Díaz and Olivero, 2010a), which is also confirmed by the negative correlation between GDP growth and the interest rate spreads in our data. Based on this evidence many theoretical papers incorporate a banking sector into DSGE models that reduces interest rate spreads following a positive technology shock and by this amplifies the real effects of the shock (see e.g. Aliaga-Díaz and Olivero, 2010b; Olivero, 2010). Simple correlations however do not tell anything about the nature of the shocks driving the business cycle. We show that in the Euro Area interest rate spreads only react countercyclically on impact in response to aggregate demand shocks. If aggregate supply or monetary policy shocks hit the economy, spreads are rather procyclical. Given the observable negative correlation between GDP growth and the interest rate spreads in our data, this implies that aggregate demand shocks have been the main drivers of the business cycle in the Euro Area (see De Nicolò and Lucchetta, 2011, for similar results).

Fourth, we use harmonized retail bank interest rate data for 11 Euro Area member countries. In the past many cross–country studies faced the problem that in particular loan rate data was not comparable at all, due to different maturities, different definitions of the borrowers and different use of the loans. Since this data is only available since 2003, the main drawback of our approach is the short sample available for the empirical exercise. For this reason we resort to panel techniques, which is still rather uncommon in the literature on the interest rate pass–through.

The remainder of the paper is organized as follows. In Section 2 we introduce the panel VAR model and provide a detailed discussion on the identification of the structural shocks. Section 3 presents the impulse response analysis and a decomposition of the forecast error variance. Section 4 compares our empirical results with those obtained from simulations of the model of Gerali et al. (2010). Section 5 summarizes our main findings and concludes.

2 Panel VAR models with sign restrictions

2.1 Panel VAR model setup

Consider a panel VAR model in reduced form:

$$Y_{i,t} = c_i + \sum_{j=1}^{p} A_j Y_{i,t-j} + \varepsilon_{i,t}, \tag{1}$$

where $Y_{i,t}$ is a vector of endogenous variables for country i, c_i is a vector of country–specific intercepts, A_j is a matrix of autoregressive coefficients for lag j, p is the number of lags and $\varepsilon_{i,t}$ is a vector of reduced–form residuals. The vector $Y_{i,t}$ consists of four variables:

$$Y_{i,t} = \left[y_{i,t} \;\; p_{i,t} \;\; s_t \;\; r_{i,t} \right]', \tag{2}$$

where $y_{i,t}$ denotes real GDP, $p_{i,t}$ is the overall price level, measured by the GDP deflator, s_t is the nominal short–term interest rate, which serves as the policy instrument of the central bank[2], and $r_{i,t}$ is a retail bank interest rate, which is either a lending rate or a deposit rate. For each variable, we use a pooled set of $M \cdot T$ observations, where M denotes the number of countries and T denotes the number of observations corrected for the number of lags p. The reduced–form residuals $\varepsilon_{i,t}$ are stacked into a vector $\varepsilon_t = [\varepsilon'_{1,t} \ldots \varepsilon'_{M,t}]'$, which is normally–distributed with mean zero and variance–covariance matrix Σ.

We use quarterly data for the EMU member countries covering the period from 2003Q1 to 2011Q4.[3] The set of countries comprises Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Spain. The beginning of the sample is determined by the retail bank interest rates, which are available from the ECBs harmonized MFI interest rate statistics only since 2003. The lending rates refer to interest rates on new business loans to non–financial corporations ('NFC'), new loans to households for consumption ('CON') and new loans to households for house purchases ('MOR'). The deposit rates are interest rates agreed on new deposits from private households ('DEP'). The spectrum of maturities comprises an average over all maturities ('total') and maturities of up to one year ('–1y') as well as over one year ('1y+').

The policy instrument of the ECB is mirrored by the EONIA, which is the average of overnight rates for unsecured euro area interbank lending. As Ciccarelli et al. (2010) point out, the EONIA rate is a sensible measure of the ECB's monetary policy especially during the financial crisis. Since the ECB reacted to the crisis by implementing various non–standard measures in its liquidity management, the EONIA reflects much better the monetary policy stance than the official main refinancing rate.

We express real GDP and the price level in logs, while the interest rates are expressed in percent. All variables are linearly de–trended. We estimate several panel VAR models, which differ with regard to the respective retail bank interest rates $r_{i,t}$. The number of countries M varies across specifications depending on the availability of retail bank interest rates.[4] The advantage of

[2]Since for all member countries of the Euro area the nominal short–term interest rate is identical, we set $s_{i,t} = s_t$ for all i.

[3]The data is taken from Eurostat and the ECB databases. See the Appendix for a detailed description of the data.

[4]See Appendix B about information concerning the country–specific availability of the

the panel approach is that it increases the efficiency of the statistical inference. This holds in particular in cases where the sample is short.[5]

Each panel VAR model is estimated with Bayesian methods using a Normal–inverted Wishart prior, 500 draws and a lag order of $p = 2$. The matrix of constant terms c_i comprises individual country dummies that account for cross–country heterogeneity. The use of the fixed effect estimator in dynamic panels warrants some discussion as it assumes slope homogeneity across countries, which potentially yields biased estimates provided that the degree of heterogeneity is actually pronounced. While the mean–group estimator of Pesaran and Smith (1995) would account for slope heterogeneity across countries, its use however requires that the sample is sufficiently long, i.e. "twice as long as usually recommended in the dynamic panel data literature" (Rebucci, 2010, p. 1183), which in our analysis is clearly violated. Mote Carlo simulations by Rebucci (2010) show in fact that the efficiency of the mean–group estimator is very limited in short samples.[6] Therefore, we follow Born et al. (2012), Hristov et al. (2012) and Tillmann (2012) and adopt the fixed effect estimator.

2.2 Identification of structural shocks

Based on the panel VAR models estimated for the different retail bank interest rates, we generate impulse responses of the variables to structural shocks η_t. As in Canova and De Nicolò (2002), Peersman (2005) and Uhlig (2005) the shocks are identified by imposing sign restrictions. The reduced–form residuals ε_t are related to the structural shocks η_t according to $\eta_t = (U\Omega^{1/2}Q)^{-1}\varepsilon_t$, where $U\Omega^{1/2}$ is the Cholesky factor, $\Sigma = U\Omega U'$, of each draw and Q is an orthogonal matrix, $QQ' = I$, generated from a QR decomposition of some random matrix W, which is drawn from an $N(0, 1)$ density. For each of the 500 Cholesky factors resulting from the Bayesian estimation of the VAR model, the draws of the random matrix

respective retail bank interest rates.

[5]Estimating single models for the individual countries separately would instead suffer from a small number of degrees of freedom if only a limited number of observations is available.

[6]In particular, Rebucci (2010) shows in Monte Carlo simulations that (i) the fixed effect estimator is biased only if slope heterogeneity is sizable, and (ii) in cases where slope heterogeneity is indeed very high, the sample must be remarkably long for the mean–group estimator to outperform the fixed effect estimator.

W are repeated until a matrix Q is found that generates impulse responses to η_t, which satisfy the sign restrictions.

Since each panel VAR model contains four variables, the maximum number of structural shocks that can be identified is four. We identify three of them, a monetary policy shock, an aggregate supply shock and an aggregate demand shock. The sign restrictions we impose are standard (Galí et al., 2003; Peersman, 2005; Straub and Peersman, 2006; Fratzscher et al., 2009; Canova and Paustian, 2010) and summarized in Table 1.

Table 1: Sign restrictions

Shock	Real GDP	GDP deflator	Money market rate	Retail bank rate
Monetary policy	↑	↑	↓	?
Aggregate supply	↑	↓	↓	?
Aggregate demand	↓	↓	↓	?

We assume that an expansionary monetary policy shock has a positive effect on output and prices and a negative effect on the money market rate. A favorable aggregate supply shock is assumed to have a positive impact on output and a negative impact on prices and the money market rate. Finally, we assume that an adverse aggregate demand shock has a negative effect on output, prices and the money market rate. We refrain from imposing restrictions on the retail bank interest rates since the adjustment of these interest rates – especially concerning the adjustment of lending rates – is theoretically ambiguous. Therefore, we let the data determine the sign of the responses of the retail bank rates.

The fourth shock is interpreted as a residual shock, which captures the remaining variation in the data.[7] For all variables we set the time period over which the sign restrictions are binding equal to four quarters. This is in line with Peersman (2005), Farrant and Peersman (2006), Rüffer et al. (2007), Sanchez (2007), Uhlig (2001) and Scholl and Uhlig (2008), who assume that the effects of shocks on economic activity can be quite sustainable.[8] All sign restrictions are imposed as $\leq$ or $\geq$.[9]

[7] See Eickmeier et al. (2009), among others, for a similar approach.

[8] We admit that the choice of the time period over which the sign restrictions are assumed to hold is arbitrary. Therefore, we checked the robustness of our results by imposing sign restrictions that are binding only over two quarters. This modification has no impact on our results.

[9] The estimation of the Bayesian VAR and the identification of the structural shocks is per-

3 Results

Overall, we estimate 12 panel VAR models as we consider four different retail bank interest rates with three different times to maturity. In order to investigate the impact of the financial crisis we split our sample into two periods, the period before the outbreak of the financial crisis 2003Q1–2007Q4 and the period of the financial crisis 2008Q1–2011Q4. Table 2 reports the statistics and the 5% critical values of various tests for instability of the OLS estimates of the panel VAR model at the end of 2007 (see Appendix C for details on the tests).

The second and third column report the results of a Chow–Wald test for the null hypothesis of equal residual variances across both subsamples. For all model specifications it confirms that the two periods are characterized by a significant degree of heteroscedasticity of the reduced–form residuals, with volatilities as measured by the trace of residual covariance matrix being on average six times higher during the crisis than before the crisis. Being aware that a higher residual variance does not necessarily imply a higher variance of the structural shocks, this result can at least be interpreted as a hint on one of the alternative explanation for the observable rise in interest rates spreads mentioned in the Introduction.

Columns 4 to 7 report the results of two tests for the null hypothesis of parameter stability across the two subsamples. For all model specifications they confirm that the financial crisis induced significant changes in the propagation mechanism of structural shocks.

3.1 Baseline model

3.1.1 Impulse response analysis

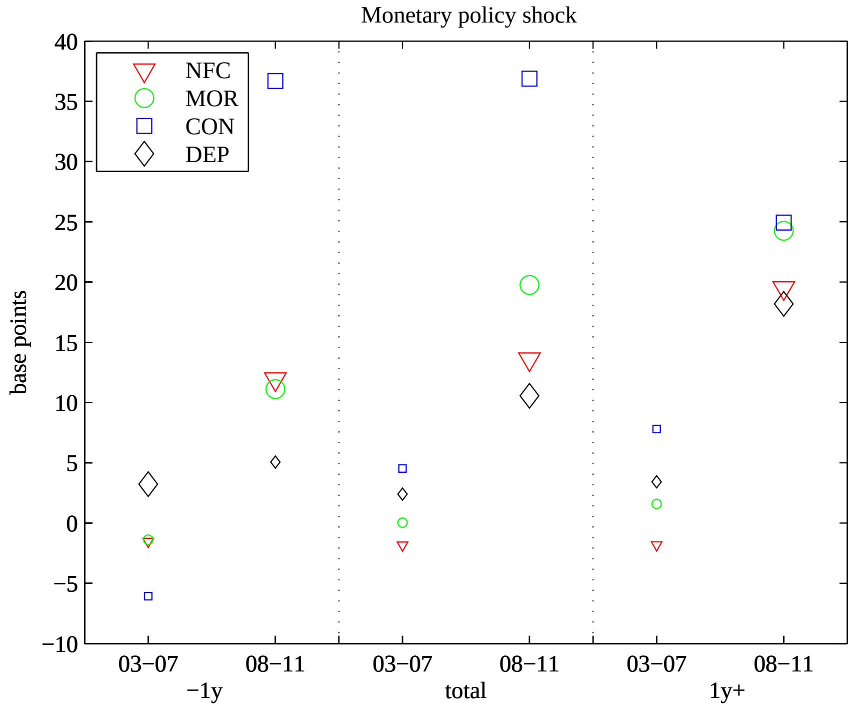
To visualize these changes we compute impulse response functions to the identified structural shocks for both periods. As we are concerned about the adjustment of the retail bank interest rate spreads to the structural shocks, we determine the response of the spreads by calculating the difference between the reaction of the respective retail bank interest rate and the overnight money market rate.

The results are summarized in Figure 2, which displays the impact reaction of the spreads to the shocks.[10] The structural shocks are normalized to a unit shock. The sign of the shocks is chosen such that the money market rate falls

formed in MATLAB, using the codes `bvar.m`, `bvar_chol_impulse.m` and `bvar_sign_ident.m` provided by Fabio Canova (http://www.crei.cat/people/canova/).

[10] The entire impulse responses of the variables to the shocks are not reported here, but are available upon request.

Figure 2: Impact impulse responses of spreads



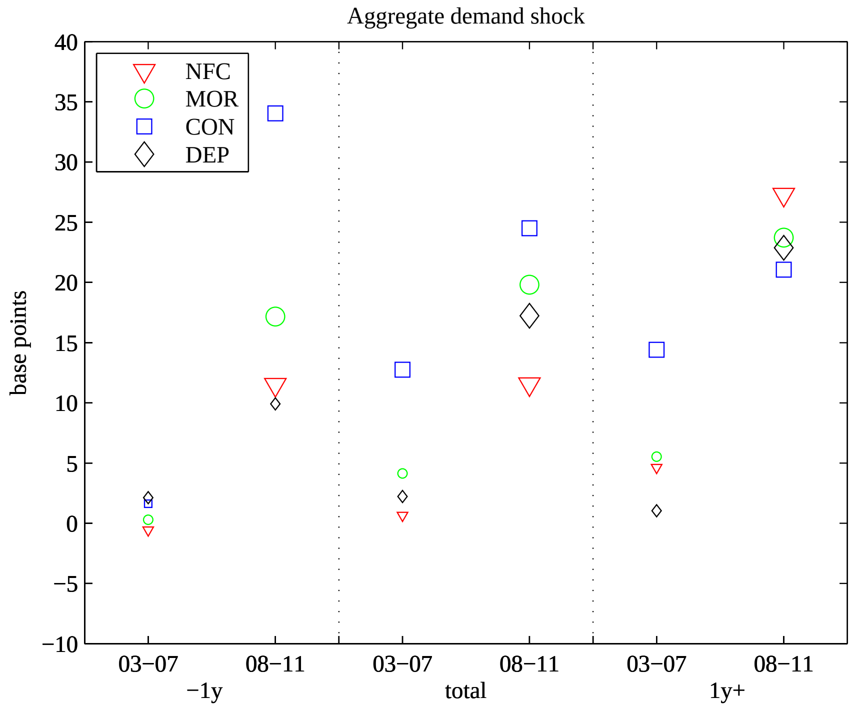
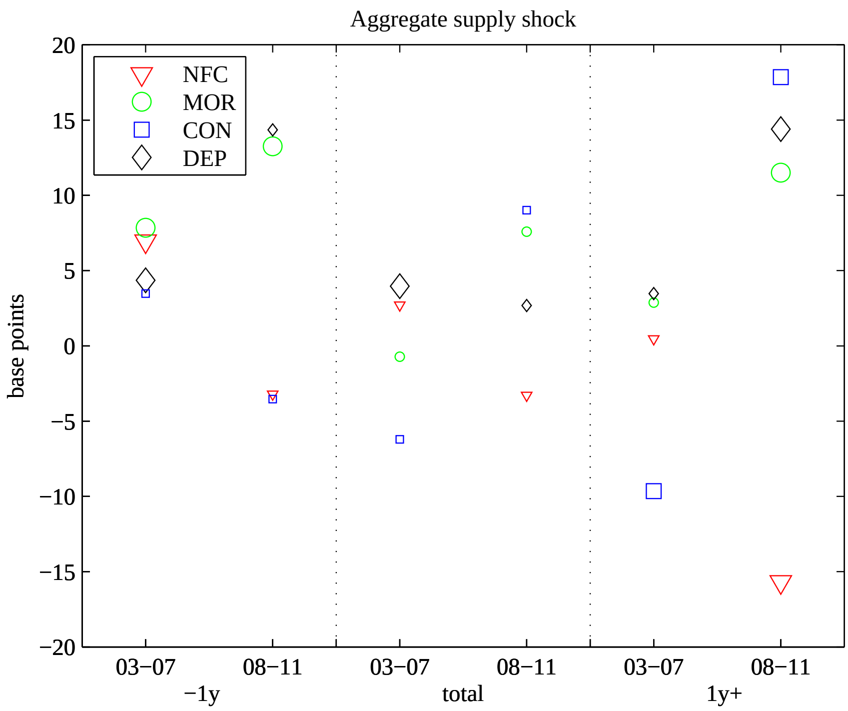


Figure 2 (continued): Impact impulse responses of spreads



Notes: The markers show the impact impulse responses of the spread of retail bank interest rates over the overnight money market rate. The spread is calculated from the medians of the impulse responses, which are estimated from a Bayesian vector–autoregression with 500 draws. The structural shocks are normalized to a unit shock. The sign of the shock is chosen such that the money market rate decreases in response to the structural shock (see Table 1). Large symbols indicate responses of the spread that are significant at the 10% level. The horizontal axis groups the models according to both, sub–sample periods and maturities of retail bank interest rates.

Table 2: Tests for instability in 2007Q4/2008Q1

	Variance statistic	5% critical value	Chow statistic	5% critical value	Wald statistic	5% critical value
NFC total	85.1	9.47	207	0.74	803541	53
MOR total	78.4	9.47	163	0.74	228844	53
CON total	112.1	9.47	192	0.74	2223045	53
DEP total	69.8	9.47	171	0.73	759592	50
NFC -1y	103.0	9.47	256	0.75	222929	57
MOR -1y	76.5	9.47	170	0.74	716606	53
CON -1y	104.9	9.47	158	0.73	2167637	50
DEP -1y	81.3	9.47	123	0.73	438507	50
NFC 1y+	84.3	9.47	174	0.74	1670290	53
MOR 1y+	71.7	9.47	252	0.74	394884	53
CON 1y+	91.1	9.47	227	0.74	1288726	53
DEP 1y+	63.6	9.47	100	0.73	4357290	47

in response to each structural shock. The symbols refer to the adjustment of the respective spreads. If the spread is positive, the decrease in the retail bank interest rates is smaller than the drop in the money market rate and the pass–through is incomplete. If the spread is zero, retail bank interest rates move proportionally with the policy rate and the pass–through is complete. If the spread is negative, the retail bank interest rates are overshooting the drop in the money market rate. Large symbols indicate responses of the spread that are significant at the 10% level. The horizontal axis groups the models according to both, sub–sample periods and maturities of retail bank interest rates.

As a general result the pass–through from changes in the money market rate to retail bank rates became less complete during the crisis. In almost all specifications spreads are more positive on impact following a decrease in the money market rate in the period 2008–2011. By contrast, in the pre–crisis period the pass–through was generally complete and most of the impact spreads were insignificant. This result holds in particular for monetary policy and aggregate demand shocks and for all maturities of the retail bank rates. In the case of aggregate supply shocks the results are less clear–cut. While the crisis–induced reduction of the pass–through is confirmed for longer maturities, the picture is mixed for maturities below one year.

Another finding is that retail bank interest rates with longer maturities are typically stickier than those with short maturities, in particular during the crisis period. Changes in money markets rates are passed–through much faster to loan and deposit rates with maturities below one year, whereas banks seem to be more reluctant concerning loan contracts and deposits with maturities above one year. A comparison of the results across bank products shows that above all interest rates on new loans to households for both, consumption and house purchases, are stickier than interest rates charged on non–financial corporations and deposit rates. The quantitative impact of the unit shocks on the interest rate spreads is more or less the same across shock types. For all shocks spreads rose by between 10 and 15 base points during the crisis period for maturities below one year and between 20 and 25 base points for maturities above one year. A notable exception are the spreads of interest rates on new loans to households for consumption with maturities below one year, which in the case of monetary policy and aggregate demand shocks are significantly larger.

The impulse response analysis also yields some interesting results concerning the cyclicality of interest rate spreads. As already mentioned in the Introduction there seems to be some consensus in the empirical literature that bank interest rate spreads are countercyclical (see e.g. Aliaga-Díaz and Olivero, 2010a). While Table 3 shows that this countercyclicality is confirmed in the case of aggregate demand shocks, interest rate spreads react procyclically in the case of monetary policy and aggregate supply shocks. Given the observable negative correlation between GDP growth and the interest rate spreads in our data, this implies that aggregate demand shocks must have been the main drivers of the business cycle in the Euro Area (see also the variance decomposition presented in the next Section). Since the response of the retail bank interest rates was left unrestricted in our identification approach, this result implies that irrespective of the nature of the shocks the response of retail bank rates has the same sign as the response of the money market rate, but the adjustment is less than proportional to the money market rate and delayed.

Table 3: Cyclicality of interest rate spreads

	Real GDP	Money market rate	Spread
Positive monetary policy shock	$\uparrow$	$\downarrow$	$\Uparrow$
Positive aggregate supply shock	$\uparrow$	$\downarrow$	$\Uparrow$
Positive aggregate demand shock	$\uparrow$	$\uparrow$	$\Downarrow$

Notes: The arrows show the sign of the impact responses which are significant at the 10% level. In the case of real GDP and the money market rate the responses are imposed through the sign restrictions (see Table 1); the reaction of the spread however is determined by the data.

3.1.2 Variance decomposition

For an analysis of the quantitative importance of the structural shocks we compute the forecast error variance decomposition, which in contrast to the impulse response analysis takes into account the estimated magnitude of the shocks. Table 4 reports the forecast error variance decomposition of each variable at the 1– to 5–year forecast horizon as an average over all model specifications. Aggregate demand shocks explain most of the variation of real GDP and interest rates both, before and during the crisis, while monetary policy shocks and to a lesser extent also aggregate supply shocks seem to be most relevant for explaining fluctuations in the aggregate price level. The impact of the financial crisis becomes most obvious with regard to the sources of fluctuations of real GDP. Monetary policy and aggregate supply shocks lost most of their explanatory power in the period 2008–11, whereas aggregate demand shocks seem to be the driving force behind the recession (see De Nicolò and Lucchetta, 2011, for similar results).

The final column, which simply sums up the contribution of all identified shocks to the variation of the endogenous variables, shows that the role of additional (unidentified) shocks seems to have increased during the crisis. This confirms our previous work in which we focussed on the role of loan supply shocks during the financial crisis (Hristov et al., 2012). We found that these shocks, which were identified by an increase in the spread between the money market rate and the loan rate, significantly contributed to the evolution of macroeconomic variables in all Euro Area member countries. Thus, the rise in interest rate spreads documented in the Introduction of this paper may also partly be explained by the emergence of additional (crisis–specific) shocks.

Table 4: Forecast error variance decomposition (in percent)

	Year	Monetary policy shock			Aggregate supply shock			Aggregate demand shock			sum		
		03-07	03-11	08-11	03-07	03-11	08-11	03-07	03-11	08-11	03-07	03-11	08-11
Real GDP	1st	20	3	5	18	4	3	26	56	46	64	63	54
	2nd	21	6	5	15	4	5	32	52	41	68	62	52
	3rd	21	12	6	15	5	6	32	45	40	69	61	51
	4th	21	14	6	15	5	6	32	42	39	69	62	50
	5th	21	15	6	15	5	6	32	42	39	69	62	51
GDP deflator	1st	32	36	15	17	18	25	12	15	11	61	69	50
	2nd	32	32	14	15	12	22	17	29	17	64	73	53
	3rd	32	32	13	16	10	21	18	31	19	65	74	54
	4th	32	33	13	16	10	21	18	30	19	66	73	54
	5th	31	34	13	16	10	21	18	29	19	66	73	53
Money market rate	1st	8	10	4	4	10	10	45	42	45	58	62	58
	2nd	6	5	4	5	5	8	46	50	45	58	60	57
	3rd	7	6	3	5	6	9	46	46	44	59	58	56
	4th	7	8	3	5	6	9	48	43	44	60	57	56
	5th	7	8	3	5	6	9	48	43	44	60	57	56
Retail bank rate	1st	9	10	8	11	11	12	31	21	18	52	42	38
	2nd	8	8	8	9	7	11	40	40	33	57	54	52
	3rd	9	8	8	9	8	11	37	40	35	56	56	54
	4th	9	8	8	9	8	11	37	39	35	56	56	53
	5th	9	9	8	9	8	11	38	39	34	56	56	53

Notes: This Table shows how much of the forecast error variance of each of the variables can be explained by the structural shocks. It is computed as the mean over all 12 model specifications.

3.2 Exclusion of crisis countries

The results of the previous Section could be mainly driven by those countries that were hit most severely by the crisis. In Greece, Ireland, Portugal and Spain the economic downturn was much more persistent and the banking system was significantly more distressed than in the core countries of the Euro Area. In order to see whether the impact impulse responses of the spreads are affected by the crisis countries we excluded Spain, Greece, Ireland and Portugal from our panel and re–estimated the VAR model using only data from Austria, Belgium, Finland, France, Germany, Italy and the Netherlands. Overall, Figure 3 shows that the picture for the core countries is very much the same as if all countries were included in the panel.

3.3 Common monetary policy shock

The identification of the monetary policy shock in our panel VAR requires some words of caution. Since the models are estimated by assuming slope homogeneity, cross–country differences are captured in the regressions either by the country fixed effects c_i or the reduced–form residuals $\varepsilon_{i,t}$. From the latter follows that since the identification matrix $(U\Omega^{1/2}Q)^{-1}$ is common to all countries, the sequence of structural monetary policy shocks $\eta_{i,t}^{MP}$ differs across countries. This is of course at odds with the idea that the member countries of the monetary union should be hit by the same monetary policy shock.

To address this problem we generate a series of union–wide monetary policy shocks and include this common shock as exogenous variable in an otherwise unchanged VAR model. The common monetary policy shock is estimated as the first common factor F_1 from the country–specific monetary policy shock series of the baseline model. For each of the 500 draws the common factors are extracted using

$$\eta_{i,t}^{MP} - \mu_i = l_{i1}F_1 + \ldots + l_{ik}F_k + \zeta_i, \tag{3}$$

where i denotes the country, μ_i is the mean of the country–specific monetary policy shock, $F_1, \ldots, F_k$ are the k unobservable common factors, $l_{i1}, \ldots, l_{ik}$ are the factor loadings, and ζ_i is an i.i.d. white–noise error term. The first common factor F_1 is then standardized to zero mean and unit standard deviation and included as exogenous variable in the panel VAR model, which is of the following form:

$$Y_{i,t} = c_i + \sum_{j=1}^{p} A_j Y_{i,t-j} + B\eta_t^{MP} + \varepsilon_{i,t}. \tag{4}$$

η_t^{MP} denotes the standardized monetary policy shock that is common to all countries in the euro area. The impact of the exogenous monetary policy shock

Figure 3: Impact impulse responses of spreads in core countries

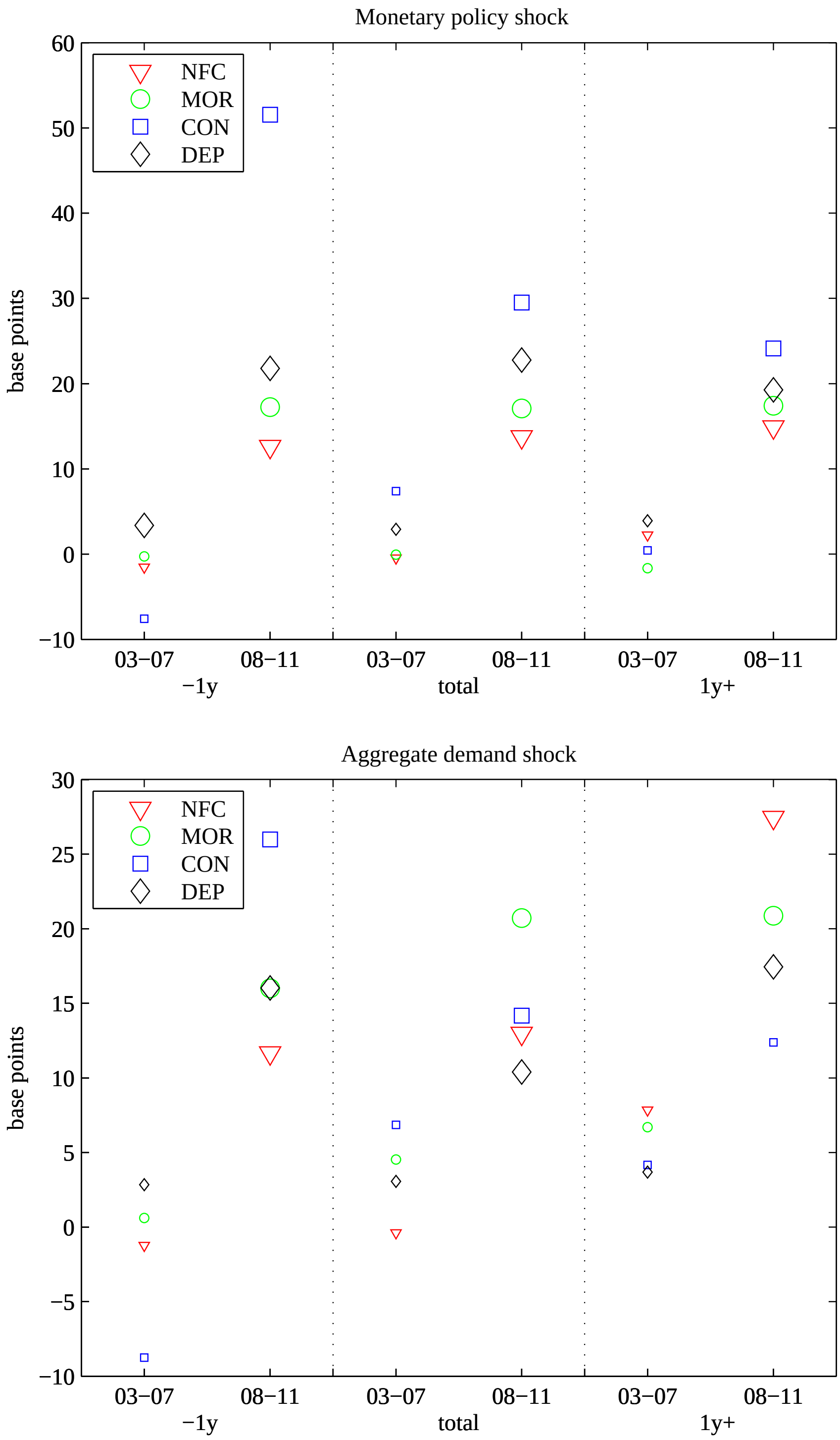
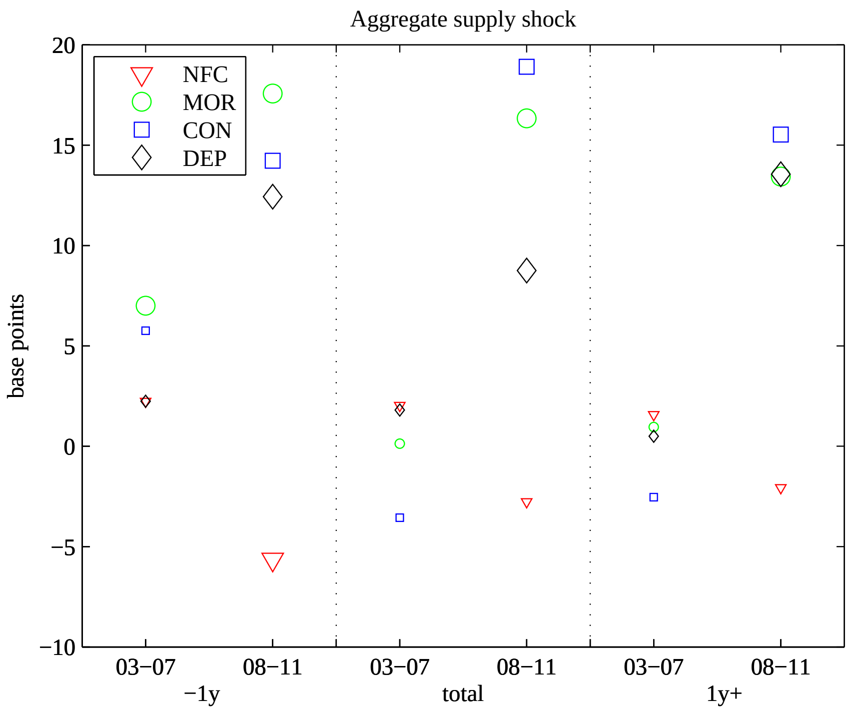


Figure 2 (continued): Impact impulse responses of spreads in core countries



Notes: The markers show the impact impulse responses of the spread of retail bank interest rates over the overnight money market rate. The spread is calculated from the medians of the impulse responses, which are estimated from a Bayesian vector–autoregression with 500 draws. The structural shocks are normalized to a unit shock. The sign of the shock is chosen such that the money market rate decreases in response to the structural shock (see Table 1). Large symbols indicate responses of the spread that are significant at the 10% level. The horizontal axis groups the models according to both, sub–sample periods and maturities of retail bank interest rates.

on the endogenous variables is finally analyzed by calculating the dynamic multipliers of (4).

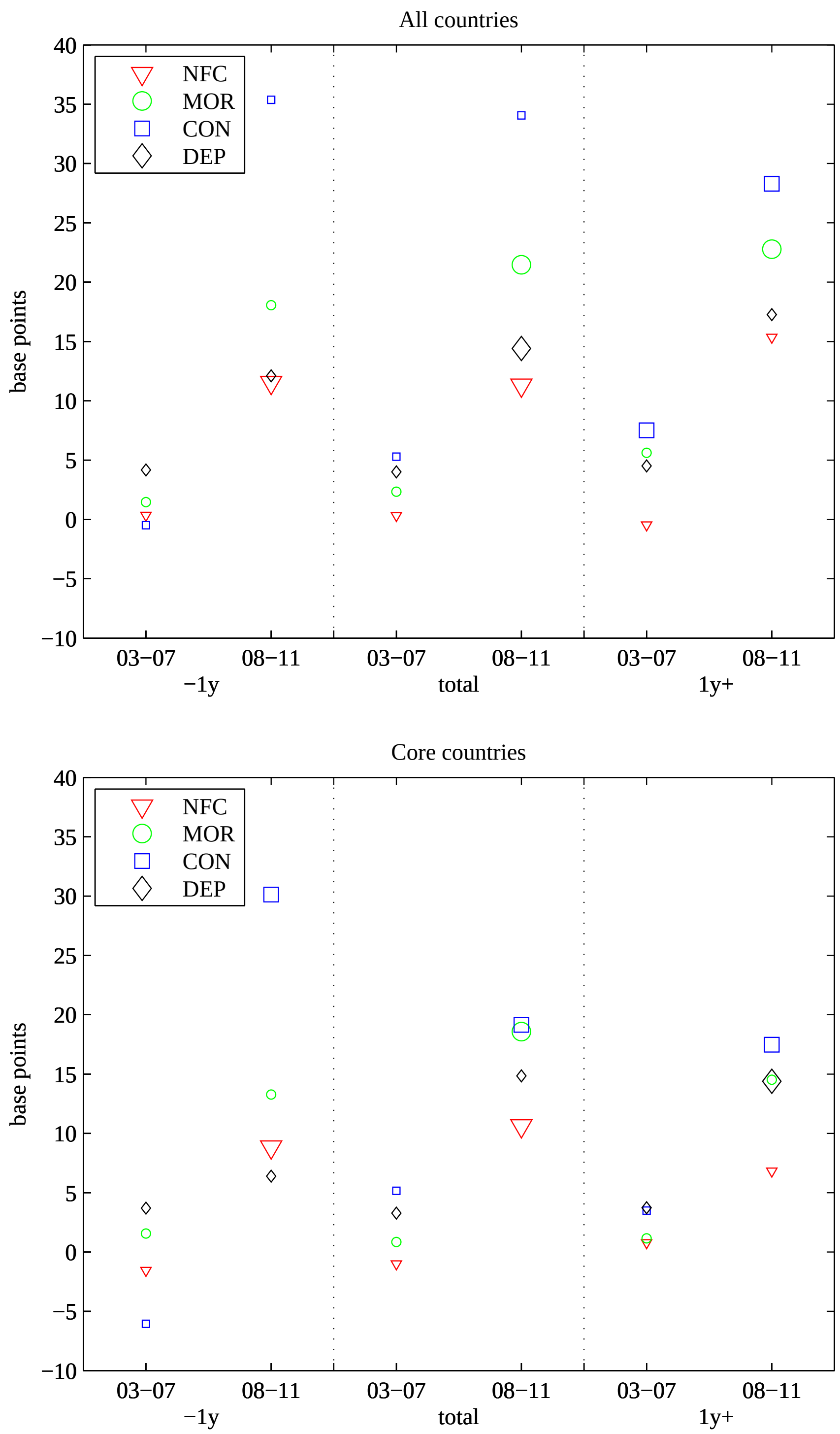
The results of this exercise, which are plotted in Figure 4, are similar to those obtained in the previous Sections. During the financial crisis retail bank interest rates responded more sluggishly to changes in money market interest rates following a monetary policy shock than in the years before the crisis, where the impact reaction of the interest rate spreads was insignificant.

4 Discussion

Our empirical results suggest that the substantial widening of various retail bank spreads during the global financial crisis was to a large extent attributable to a structurally weaker transmission of policy rate changes to loan and deposit rates. To interpret these findings and to develop a sufficiently precise economic intuition about the possible structural changes underlying our empirical results, we employ the DSGE model developed by Gerali et al. (2010). In their framework, banks issue collateralized loans to both entrepreneurs and households, collect deposits, and accumulate capital out of retained earnings. Margins charged on loans and deposits depend on bank capital–to–assets ratios and on interest rate adjustment costs. In addition, balance–sheet constraints establish a link between the business cycle, which affects bank profits and thus capital, and the supply and cost of loans.

We view the model by Gerali et al. (2010) as suitable for our purposes due to several reasons. First, it incorporates various features of retail bank markets as well as private banks' behavior which are important for understanding the dynamics of different types of retail bank interest rates and the related quantities. A number of these features might have changed systematically since the onset of the financial crisis. Second, this theoretical framework belongs to the class of medium scale, estimated DSGE models, a major advantage of which is their rich structure enabling them to replicate the cyclical properties of a several macroeconomic aggregates. Over the past fifteen years these type of models have been increasingly used by central banks and other institutions for economic analysis and forecasting purposes. Third, Gerali et al. (2010) estimate the model with Bayesian techniques using data for the euro area covering the period 1998Q1–2009Q1 and show that their framework fits the euro-area data sufficiently well.

Figure 4: Impact impulse responses of spreads to a common monetary policy shock



4.1 Simulation exercise setup

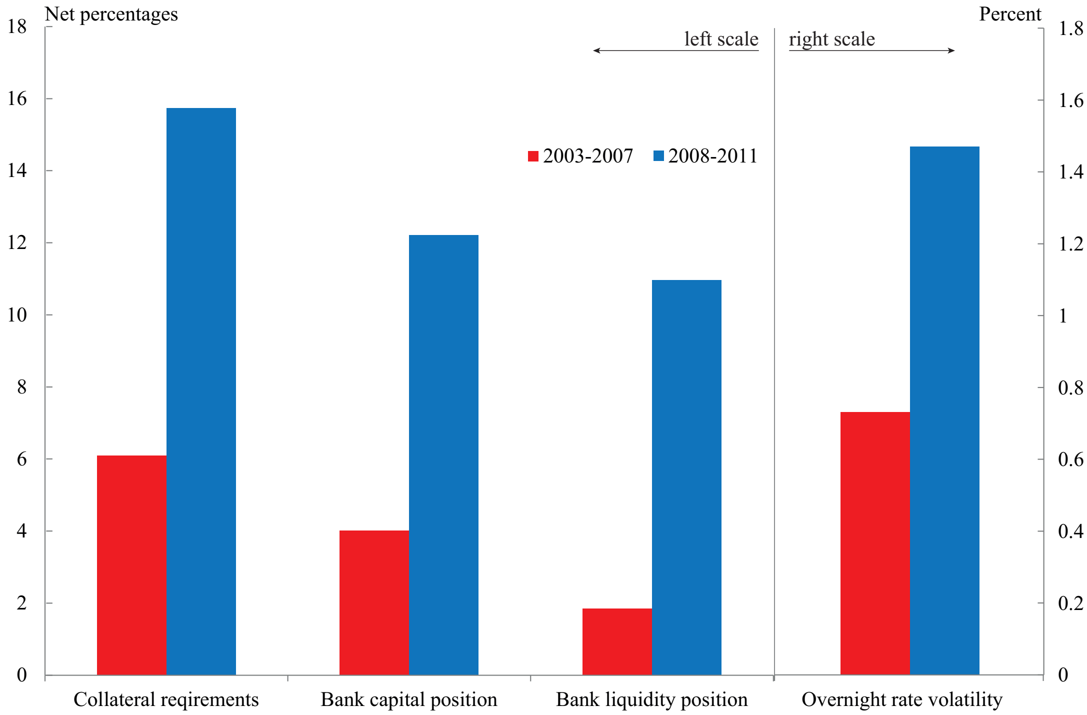
In our simulation exercise we compute the impact reaction of the retail bank rate spreads to the same types of shocks as in the empirical analysis for two different regimes i.e. steady states. The first steady state corresponds to the environment prior to the outbreak of the financial crisis, which is assumed to mimic *normal* conditions. The second steady state corresponds to the environment after the outbreak of the financial crisis. A comparison between the impact responses of the retail bank spreads across regimes reveals changes in the direction and the degree of the pass–through of policy–rate changes to loan and deposit rates.

In particular, we account for possible structural changes between the normal and the crisis regime by assuming that the latter is characterized by less favorable values of several parameters which shape agents' behavior in loan and deposit markets. The choice of parameters is governed by empirical observations as well as theoretical suggestions.

As Sinn (2010) points out, the global financial crisis severely affected banks in the Euro Area. Large loan losses and the increased degree of financial distress caused substantial changes in their ability and willingness to lend as well as in their external funding opportunities (see also European Central Bank, 2010). This is supported by Figure 5 which displays some results of the ECB bank lending survey along with measures for risk premia in interbank money markets. Each pair of bars corresponds to the averages for the periods 2003–2007 and 2008–2011. In the course of the global financial crisis banks have notably tightened their collateral requirements, making it difficult to rule out that since 2008 average loan–to–value ratios have been abnormally low. Furthermore, over the same period the fraction of banks pointing to deteriorations in their capital position as a reason for reducing loan supply was systematically higher than before 2008. This evidence suggests that during the crisis commercial banks might have been confronted with more difficulties and higher costs of raising or adjusting their equity capital position. Likewise the liquidity position of banks became worse.

Cottarelli and Kourelis (1994) argue that during episodes of financial distress banks may face a less elastic demand for loans, as customers asking for loans in incomplete financial markets may find it difficult to tab alternative sources of external funds due to higher problems of informational asymmetries. Thus, the degree of competition between banks may have decreased since the beginning of the financial crisis. Additionally, during such a crisis money markets typically become less liquid. As a result, fluctuations in the money market rate tend to become noisier as reflected by the higher volatility in the EONIA since 2008. The higher degree of noise in money markets aggravates the uncertainty regarding the precise nature of their fluctuations (Cottarelli and Kourelis,

Figure 5: Stylized Facts



Notes: Source: European Central Bank. Bank Lending Survey (BLS), net percentages. Collateral requirements: unbalanced average of the answers to questions 3.B.3, 10.B.1, 10.B.2 and 12.B.1. Bank capital position: answers to question 2.A.1. Bank liquidity position: unbalanced average of the answers to questions 2.A.2 and 2.A.3. Overnight rate volatility: standard deviation of EONIA.

1994). Simple signal extraction arguments suggest that in this environment a larger fraction of the policy signals and cyclical movements in the money market rate are not adequately transmitted to retail bank rates as it is more difficult and costly for banks to disentangle the latter from purely random movements in the interbank rates. This suggests that the costs for adjusting loan and deposit rates might have increased during the financial crises.

We account for the aforementioned effects of the financial crisis on collateral requirements, bank capital positions, competition among banks and interest rate adjustment costs by assuming that the crisis regime in our theoretical model is characterized by lower loan–to–value (LTV) ratios, higher costs for adjusting banks' equity ratio, a lower elasticity of credit demand, and a higher value of the parameter measuring the degree of interest rate stickiness. Table 5 summarizes the calibrated model parameter values.

Table 5: Setup of simulation exercise

	Normal regime	Crisis regime
Collateral requirements		
LTV ratio entrepreneurs	0.35	0.15
LTV ratio households	0.70	0.50
Bank capital adjustment		
Cost parameter	11.07	31.07
Degree of competition		
Loans to firms	3.12	2.12
Loans to households	2.79	1.79
Interest rate stickiness		
Loans to firms	9.36	18.72
Loans to households	10.09	20.18
Deposits	3.50	7.00

Notes: Calibrated model parameters for the normal regime and the crisis regime. Parameter values for the normal regime are taken from Gerali et al. (2010).

We refer to the parameter values reported by Gerali et al. (2010) in order to calibrate the normal regime. The parameter values for the crisis regimes are chosen arbitrarily and solely to illustrate the qualitative, not the quantitative, implications of the model. In fact, the higher the parameter deviation from the normal regime, the stronger the change in the impact response of the loan and deposit rate spreads reported below.

We simulate the impact reaction of the retail bank spreads to the same exogenous shocks as in the empirical analysis, i.e. an expansionary monetary policy shock, a positive aggregate supply shock and a negative aggregate demand shock. The aggregate supply shock is modeled as a temporary change in technology, and the aggregate demand shock as a temporary change in households' preferences. The sign of each shock is chosen such that the central bank lowers the policy rate on impact.

4.2 Model structure

To develop intuition about the impact effects of various types of shocks on the loan and deposit spreads in the model of Gerali et al. (2010), it suffices to take a closer look at the equations describing the optimal loan rate setting in the wholesale and retail banking sector. Since the structure of the first order conditions governing the optimal choice of deposit rate setting is analogous, we do not discuss them explicitly.[11] The representative wholesale bank operates under perfect competition. The wholesale bank provides loans to retail banks at rate R_t^b, which is set according to:

$$R_t^b = r_t - \kappa_{Kb} \left(\frac{K_t^b}{B_t} - \nu^b \right) \left(\frac{K_t^b}{B_t} \right)^2, \tag{5}$$

where B_t is the amount of wholesale loans, K_t^b is the bank capital position and r_t is the policy rate. The dependence on the capital–to–asset ratio K_t^b/B_t is due to the assumption that it is costly to deviate from its desired (or required) value ν^b. The parameter κ_{Kb} governs the severity of the corresponding adjustment cost function. A higher κ_{Kb} implies that it is more costly to change the capital–to–asset ratio.

Retail banks operate in a monopolistically competitive environment. They face quadratic adjustment costs of changing individual loan and deposit rates. The first order condition for the representative loan rate r_t^{bs} charged on loans of type s reads:[12]

$$\begin{aligned} 0 \quad = \quad & 1 - \epsilon^{bs} + \epsilon^{bs} \frac{R_t^b}{r_t^{bs}} - \kappa_{bs} \left(\frac{r_t^{bs}}{r_{t-1}^{bs}} - 1 \right) \frac{r_t^{bs}}{r_{t-1}^{bs}} \\ & + \beta_P E_t \left\{ \frac{\lambda_{t+1}^P}{\lambda_t^P} \kappa_{bs} \left(\frac{r_{t+1}^{bs}}{r_t^{bs}} - 1 \right) \left(\frac{r_{t+1}^{bs}}{r_t^{bs}} \right) \frac{b_{t+1}^s}{b_t^s} \right\}, \end{aligned} \tag{6}$$

where b_t^s is the loan volume, λ_t^P denotes marginal utility of the bank owners, β_P is a discount factor, ϵ^{bs} is the elasticity of substitution between loan products of

[11]The interested reader is referred to the original paper for a more detailed description of the theoretical framework.

[12]Note that due to the assumption of quadratic adjustment costs for loan rates instead of a Calvo–type loan rate staggering retail banks are symmetric with respect to the optimal loan/deposit rate.

type s and κ_{bs} measures the intensity of the loan rate adjustment costs.

The impact reaction of the spread between the loan rate and the policy rate is indirectly affected by the size of the loan–to–value ratio. The link can be derived by considering the linear approximation of (6):

$$\begin{aligned} dr_t^{bs} &= \frac{\kappa_{bs}}{\epsilon^{bs}-1+(1+\beta_P)\kappa_{bs}} dr_{t-1}^{bs} + \frac{\beta_P \kappa_{bs}}{\epsilon^{bs}-1+(1+\beta_P)\kappa_{bs}} E_t dr_{t+1}^{bs} \\ &\quad + \frac{\epsilon^{bs}}{\epsilon^{bs}-1+(1+\beta_P)\kappa_{bs}} dR_t^b, \end{aligned} \tag{7}$$

together with the linear approximation of (5):

$$dR_t^b = dr_t - \kappa_{Kb}\nu^3(\hat{K}_t - \hat{B}_t), \tag{8}$$

where d denotes the absolute deviation from the steady–state value, while hatted variables describe the relative deviation from the steady–state value.

The impact reaction of the spread is given by:

$$dr_t^{bs} - dr_t \approx (\varphi_2 - 1)dr_t + \varphi_2 \kappa_{Kb}\nu^3 \hat{B}_t + \varphi_1 E_t dr_{t+1}^{bs}, \tag{9}$$

where:

$$\begin{aligned} \varphi_1 &= \frac{\kappa_{bs}}{\epsilon^{bs}-1+(1+\beta_P)\kappa_{bs}} \\ \varphi_2 &= \frac{\epsilon^{bs}}{\epsilon^{bs}-1+(1+\beta_P)\kappa_{bs}}. \end{aligned}$$

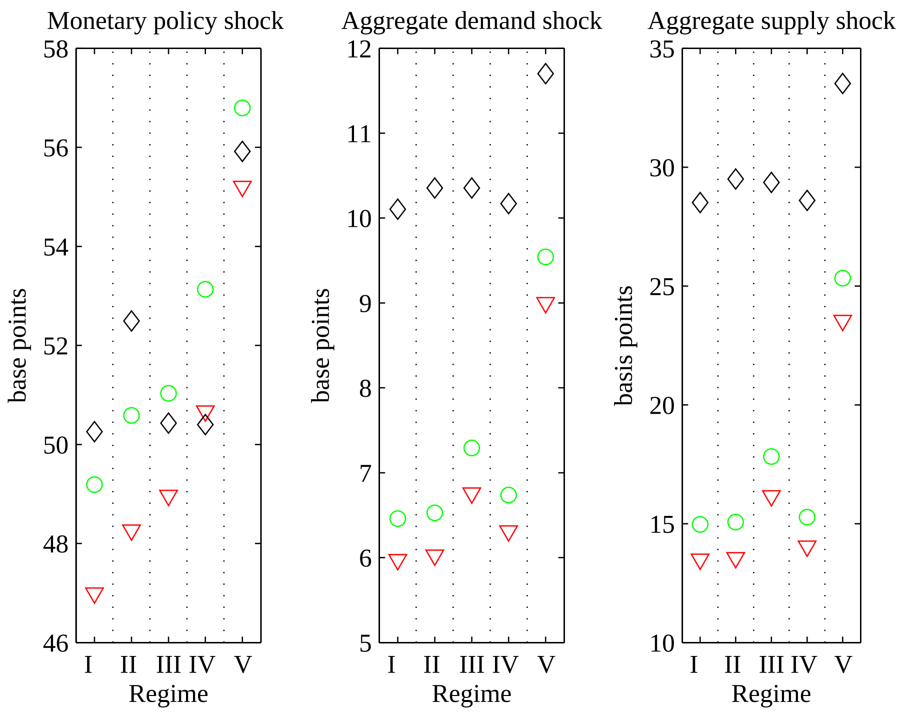
The impact reaction of the spread (9) is derived by inserting (5) into (6) and recognizing that the impact response of bank capital $\widehat{K}_t$ and the previous loan rate $\widehat{r}_{t-1}^{bs}$ is zero as these variables are determined by decisions in the past.

The impact reaction of the spread is positively affected by an increase in the loan volume B_t, while the effect of an increase in the policy rate r_t is ambiguous (see equation 9). On the one hand the spread tends to decline when the policy rate raises, given that $(\varphi_2 - 1) < 0$. On the other hand, the spread tends to widen when the policy rate increases because the expected future loan rate $E_t r_{t+1}^{bs}$ comoves positively with the policy rate provided that the shift of the policy rate is at least to some degree persistent. Overall, however, the impact reaction of the spread is largely driven by the policy rate since the parameter calibration suggested by Gerali et al. (2010) implies that $|\varphi_2 - 1|$ is substantially larger than the sum of φ_1 and $\varphi_2\kappa_{Kb}\nu^3$ (Gerali et al., 2010, p. 124). The deposit rate spread too, depends negatively on the policy rate with movements in the latter being its main driving force.

4.3 Simulation exercise results

Figure 6 compares the impact responses of retail bank spreads in the model economy to macroeconomic shocks under the different regimes. We denote the regime with normal conditions by (I), while the crisis regimes are summarized by (II–V), where (II) refers to lower loan–to value ratios, (III) to higher costs of bank capital adjustment, (IV) to a lower degree of competition and, (V) to higher costs of interest rate adjustment. Circles refer to loans to households, triangles to loans to firms, and diamonds to deposits.

Figure 6: Simulation of impact reaction of retail bank spreads to shocks under different regimes



Notes: Symbols summarize the impact reaction of the respective retail bank spreads to the respective shocks under different regimes. Regime I - normal conditions; Regime II - lower loan–to–value ratios; Regime II - higher costs of bank capital adjustment; Regime IV - lower elasticity of substitution between loan varieties; Regime V - higher costs of interest rate adjustment. Green circles refer to loans to households, red triangles to loans to firms, and black diamonds to deposits.

Overall, our findings indicate a tendency for a widening of the retail bank spreads when the economy moves from the normal regime (I) to one of the crisis regimes (II–V), which indicates that the interest rate pass–through is becoming weaker. Similar results are obtained if the preference shock is replaced by a government spending shock and the technology shock is replaced by a cost push shock which shifts the goods market Phillips curve upwards. Accordingly, our simulation results are consistent with the idea that changes in structural parameters governing commercial banks' behavior during the financial crisis are a possible source for explaining the observable increase in the incompleteness of the interest rate pass–through.

4.4 Intuition of model simulation results

4.4.1 Lower steady state loan–to–value ratio

Generally, the impact reaction of the spread depends indirectly on the size of the loan–to–value ratio, because the latter determines the sensitivity of loan demand to changes in interest rates. A sudden decrease in the policy rate induced by a expansionary monetary policy shock causes a rise in aggregate demand associated with an expansion in loan demand and an acceleration of inflation. However, the increase in loan demand is less pronounced when collateral requirements are tighter, i.e. the case of a lower loan–to–value ratio. Accordingly, the expansion in aggregate demand is comparatively weak and the increase in the inflation rate is dampened with the consequence that the reaction of the policy rate to the economic upturn is less intensive. The weaker increase in the policy rate, in turn, leads to a stronger increase in the spread as the relationship between these two variables is negative (see equation 9).[13]

The aggregate demand shock is assumed to emerge as a preference shock lowering the marginal utility of consumption relative to that of housing and leisure. Such a negative preference shock induces a rebalancing of demand away from consumption of non–durables and towards housing. As a consequence, relative house prices increase which, everything else equal, leads to a loosening of the collateral constraint. This, in turn, enables borrowers to expand borrowing and thus, to maintain a higher level of consumption. However, with a lower loan–to–value ratio, the effect of easing the collateral constraint is comparatively weaker. Accordingly, in the crisis regime a negative preference shock implies a stronger decrease or a less pronounced increase in aggregate consumption leading

[13] Note that the weaker increase in loan demand and the weaker increase in expected future loan rates associated with a lower loan–to–value ratio tend to dampen the widening of the loan–rate spread. However, as noted above, the latter two effects are dominated by the one resulting from changes in r_t.

to a smaller acceleration in inflation. Thus, there is a less aggressive increase of the policy rate. As a result, the loan rate spread increases by more.[14]

A lower loan–to–value ratio weakens the pass–through of policy rate changes to loan rates also in the case of a technology shock. Consider a *positive* technology shock which, as usual, is associated with a decelerating inflation and a corresponding lowering of the policy rate. If banks only accept lower loan–to–value ratios then the increase in loan demand and consumption will be less pronounced. Consequently, the decline in inflation and thus the reduction in the policy rate will be larger than in an economy characterized by normal collateral requirements. Accordingly, the increase in the loan rate spread will be stronger even though the less positive reaction of the loan volume and the more negative response of expected future rates again put a downward pressure on it.

4.4.2 Higher bank capital adjustment costs

A higher κ_{Kb} makes the spread in wholesale banking $dR_t^b - dr_t$ more sensitive to changes in the loan volume B_t. Consequently, the retail banking spread $dr_t^{bs} - dr_t$ also becomes more exposed to changes in B_t since the wholesale rate R_t^b is an important determinant of the loan rate charged by retail banks. For example, an expansionary monetary shock implies an increase in loan volume and hence, everything else equal, widens the wholesale spread by an amount equal to

$$\kappa_{Kb}\nu^3|\widehat{B}_t|$$

on impact. The magnitude of this positive reaction is stronger for higher values of κ_{Kb}. The intuition with respect to a negative aggregate demand and a positive technology shock is similar since they both lead to an increase in loan volume B_t.[15] The reaction of the loan rate spread is further magnified through a self reinforcing loop stemming from the direct link between the policy rate and the spread itself. For any given initial reaction of of the policy rate an increase in the spread (e.g. due to a higher κ_{Kb}) implies a tightening of the collateral constraints and thus depresses borrowers' consumption. This puts a downward pressure on inflation. The monetary authority reacts to it by lowering the policy rate which, in turn, affects positively the loan rate spreads (see equation 9). It is this self reinforcing relationship between the policy rate and the loan rate spread that leads to a higher deposit rate spreads in this type of crisis regime.

[14]In the regime characterized by lower loan–to–value ratios there is a less pronounced increase in the loan volume B_t as well as the expected future loan rate. Both effects work towards lowering the spread. However they are dominated by the effect of r_t on the retail rate spread.

[15]See the discussion in Section 4.4.1 as well as equation (9).

4.4.3 Lower elasticity of substitution between loan varieties and higher interest rate adjustment costs

When setting loan rates retail banks compare the opportunity cost of deviating from the optimal policy (which would prevail under perfectly flexible interest rates) with an increasing marginal costs of interest rate adjustment. Accordingly, for any given shock to marginal costs, the lower the curvature of the profit function, the lower its sensitivity with respect to deviations from the optimal retail bank rate and thus, the smaller the desired adjustment $r_t^{bs} - r_{t-1}^{bs}$. It is straightforward to show that the curvature of banks' profits is positively related to the degree of substitutability between individual bank products measured by ϵ^{bs}. Consequently, lower values of ϵ^{bs} tend to attenuate the response of loan and deposit rates to policy rate changes.[16] Put more technically, a lower degree of competition flattens the bank–rate Phillips Curve (see equation 7) as it weakens the link between the interest rates set by private banks r_t^{bs} and marginal costs represented by R_t^b while, at the same time the relationship between current and expected loan and deposit rates becomes stronger. For standard calibrations the former effect operating via the marginal–cost term in equation (7) dominates (see also the discussion above).

Higher values of the parameters measuring the banks' costs of adjusting retail bank rates directly increase the degree of loan and deposit rate stickiness. Consequently, any given change in the policy rate is passed–through to a smaller extent to the interest rates charged by retail banks.

In both types of crisis regimes, there is a stronger increase of the deposit rate spread as well. This is again the consequence of the self reinforcing link between the spread in loan rates and the interest rate set by the monetary authority described above.

5 Concluding Remarks

We employ panel VAR models for the Euro Area member countries to investigate how banks reacted to the global financial crisis by adjusting their retail rates in response to changing money market rates. We focus on the interest rate pass–through during the period from 2003–2011 by considering three macroeconomic shocks, namely a monetary policy shock, an aggregate supply shock and an aggregate demand shock, which are identified by imposing sign restrictions.

We find that the interest rate pass–through in the Euro Area became significantly less complete during the financial crisis 2008–2011, while it was generally

[16]See Andrés et al. (2008) for a more detailed discussion of this issue with regard to the goods market.

complete before. This result holds for various categories of retail bank interest rates, i.e. lending rates and deposit rates with different maturities. However, retail bank interest rates with longer maturities are typically stickier than those with short maturities, which applies in particular for the period during the financial turmoil.

The increasing incompleteness of the interest rate pass–through in the course of the financial crisis implies that the transmission mechanism of monetary policy was severely distorted. Our results suggest that distress in the banking sector has led to a fundamental change in the propagation of macroeconomic shocks. Simulations of a DSGE model with a banking sector show that the interest rate pass–through becomes more incomplete in response to tighter collateral requirements, higher costs of restoring the bank capital position, weaker competition among banks and higher costs of adjusting interest rates. Empirical evidence supports that these factors have deteriorated substantially since the onset of the financial crisis.

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Appendix

A Data

We use data for 11 European countries that is taken from Eurostat and the ECB covering the period from 2003Q1 to 2011Q4. The panel of countries includes Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Spain. The data comprises:

1. Real GDP (y_t): Gross domestic product at market prices, calendar and seasonally adjusted, in constant 2000 EUR (Eurostat).

2. GDP deflator (p_t): Price index, 2000=100, gross domestic product at market prices, calendar and seasonally adjusted (Eurostat).

3. Money market rate (s_t): Euro Overnight Index Average (EONIA), in percent (ECB).

4. Retail bank rate (r_t): Various retail bank rates with different time to maturity, in percent (ECB, MFI database).

The retail bank interest rates include the following categories: lending rates on loans to non–financial cooperations ('NFC'), lending rates on mortgage loans ('MOR'), lending rates on consumer credit ('CON') and deposit rates on deposits offered to private households ('DEP'). For all retail bank interest rates we consider three different maturity categories: total, i.e. the average over all maturities ('total'), up to 1 year ('–1y') and over 1 year ('1y+').

B VAR model specification with respect to country–data availability

For the different countries the availability of data on the respective retail bank interest rates is mixed. Accordingly, in the panel VAR models for every retail bank interest rate category, the number of countries varies. Table 6 summarizes the countries, which are excluded from the respective estimated models due to the lack of data availability.

C Tests for instability

C.1 Tests for equal residual variances

We employ a multi–equation version of the Chow–Wald test proposed by Stock and Watson (2005) to test for the null hypothesis of equal residual variances

Table 6: Availability of retail bank interest rates

Retail bank rate	Maturity	Missing countries
NFC	total	Greece
	–1y	–
	1y+	Greece
CON	total	Netherlands
	–1y	Belgium, Netherlands
	1y+	Netherlands
MOR	total	Greece
	–1y	Belgium
	1y+	Greece
DEP	total	Greece, Netherlands
	–1y	Belgium, Ireland
	1y+	Ireland, Netherlands

across both subsamples. The test statistic reads

$$\left(\mathrm{diag}\Sigma^{1,\bar{t}} - \mathrm{diag}\Sigma^{\bar{t}+1,T}\right)' \Omega^{-1} \left(\mathrm{diag}\Sigma^{1,\bar{t}} - \mathrm{diag}\Sigma^{\bar{t}+1,T}\right), \tag{C.1}$$

where Σ^{t_1,t_2} denotes the covariance matrix of the residuals of the panel VAR estimated over the subsample $[t_1, t_2]$, diag returns a column vector formed from the main diagonal of a matrix and $\Omega = \left(\mathrm{diag}(\Omega_{1,\bar{t}} + \Omega_{\bar{t}+1,T})\right)^{1/2}$. Ω_{t_1,t_2} is the heteroscedasticity–autocorrelation–adjusted covariance matrix of

$$\begin{pmatrix} \mathrm{vec}\varepsilon'_{t_1}\varepsilon_{t_1} \\ \vdots \\ \mathrm{vec}\varepsilon'_{t_2}\varepsilon_{t_2} \end{pmatrix}, \tag{C.2}$$

where vec stacks the elements of a matrix. The test statistic has an asymptotic χ^2–distribution with degrees of freedom equal to the number of equations $n = 4$ in the panel VAR model.

C.2 Tests for parameter instability

A common approach to test for structural breaks is to use Chow tests. The test statistic can be computed as

$$\frac{\left(\det\Sigma^{1,T} - \det(\Sigma^{1,\bar{t}} + \Sigma^{\bar{t}+1,T})\right)/\nu}{\det(\Sigma^{1,\bar{t}} + \Sigma^{\bar{t}+1,T})/(M \cdot T - 2\nu)}, \tag{C.3}$$

where $\Sigma^{1,T}$ is the covariance matrix of the residuals with no breaks, $\Sigma^{1,\bar{t}}$ and $\Sigma^{\bar{t}+1,T}$ are the covariance matrices for the two subsamples, det denotes the determinant of a matrix, and $\nu = (np + M)n$ is the number of regressors. The Chow statistic has an asymptotic F–distribution with ν and $M \cdot T - 2\nu$ degrees of freedom.

A major drawback of the Chow test statistic is that it only has an F–distribution if the residual variance in the two subsamples is equal. If this assumption is violated it is more appropriate to compute a Wald statistic for the null hypothesis of no structural change. It can be constructed as

$$(A^{1,\bar{t}} - A^{\bar{t}+1,T})'(V^{1,\bar{t}} + V^{\bar{t}+1,T})^{-1}(A^{1,\bar{t}} - A^{\bar{t}+1,T}), \tag{C.4}$$

where A^{t_1,t_2} is a vector of all parameter estimates of the panel VAR model (1) (including the country–fixed effects) estimated over the period $[t_1, t_2]$, V^{t_1,t_2} is the related parameter covariance matrix, and $\bar{t}$ is 2007Q4. The Wald statistic has an asymptotic χ^2–distribution with degrees of freedom equal to the number of estimated parameters ν in the A^{t_1,t_2} vector.[17]

[17] The number of estimated parameters varies across models due to changes in the number of countries for which retail bank rates are available.