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Bank-Lending Standards, Loan Growth and the Business Cycle in the Euro Area

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Abstract

We study the relationship between bank lending standards, loan growth and the business cycle in the euro area and the US within a vector error correction model using Bayesian estimation methods. To deal with the short data series available for the euro area, we exploit information from the estimated US system to improve the estimation of the euro area system. We find that tighter bank lending standards are associated with lower loan growth as well as lower output growth in both areas. Differences in reactions appear in the strength and the persistence of responses.

Keywords: Bank Lending Standards, Bayesian Cointegration Analysis

JEL codes: E40, E50

1 Introduction

How important are non-price lending terms, in addition to the interest rate, for loan growth and the business cycle? Using data from the U.S. Senior Loan Officer Survey, Lown and Morgan (2006) show that bank lending standards are significantly related to loan growth and real output. Recently, a number of papers (see e.g. Maddaloni and Peydró, 2013; de Bondt et al., 2010; Ciccarelli et al., 2010; Cappiello et al., 2010) have investigated the role of credit standards in the euro area using the euro area bank lending survey which documents the changes in lending standards as reported by major banks in the euro area. These authors also report a significant impact of monetary policy on lending standards and a significant impact of lending standards on economic conditions like GDP growth and inflation.

All the studies for the euro area cope with the issue of the short sample size (the survey has been introduced in 2003 and is conducted on a quarterly basis) using a panel framework based on the disaggregated country-specific survey responses reported from the National Central Banks to the European Central Bank. These responses are not fully publicly available, however, and the European Central Bank discloses on its website only a weighted average of these country-specific responses. Recently, some of the countries

agreed to publish aggregated country-specific responses on the ECB's website. However, net percentage changes are not reported for all major euro area countries, which renders a representative panel analysis with aggregated country-specific responses unfeasible. Therefore, we work with the euro area aggregate series. To the best of our knowledge, analyses with the aggregate lending survey data are scarce, if existent at all, because the available sample is rather short. The published series start in 2003, leaving us with 40 data points up to the end of 2012.

In this paper, in the same line of research as the cited papers, we present first results for the euro area obtained with the published aggregate bank lending survey data. We estimate a vector error correction model (VECM) including, besides lending standards, GDP, producer prices, consumer prices, loans to non-financial corporations and a short-term interest rate. The specification closely follows the one in Lown and Morgan (2006) for the US as we want to compare results between the two regions. We pursue a Bayesian approach which allows us to incorporate prior information into the euro area system to increase estimation efficiency. We use the posterior inference on the US system to design the prior specification of the euro area system. Specifically, the posterior moments of the coefficients and the error covari-

ance matrix will shape the prior distributions of the euro area parameters and error covariance, respectively.

We find that business cycle dynamics, loan growth and the dynamics of bank lending standards are closely interrelated in the euro area as well as in the US. A tightening of standards is associated with decreasing loans as well as lower output in both economies. In fact, we find that these relationships are remarkably similar in the euro area and in the US, although quantitatively, the influence of standards is more pronounced in the euro area. We also find that standards react to output and loan growth developments and interest rate shocks in the US, while the influences of these variables are similar in the euro area, although not as significant as in the US.

The remainder of the paper is structured as follows: In Section 2 we describe our data set and provide some information on the euro area bank lending survey and in Section 3, we present the econometric approach, which we adopt. Section 4 presents the results and Section 5 concludes the paper.

2 The data

The data used are taken from the ECB's statistical website for the euro area and from FRED[®], the statistical data available on the website of the Federal

Reserve Bank of St. Louis (see Table 1). The beginning of the estimation sample is given by the start of the lending standards series in both regions. For the US the observation sample starts in the second quarter of 1990 and for the euro area in the first quarter of 2003. Both samples end in the fourth quarter of 2012. Given that the sample for the euro area is very short, we will incorporate prior information stemming from the posterior inference on the US system.

Lending standards for the US are taken from the Senior Loan Officer (SLO) Opinion Survey on Bank Lending Practices, a quarterly survey of major banks around the US. As in Lown and Morgan (2006), we use the responses of lenders to the question about lending standards to large firms (Question 1). These report on a quarterly basis how their lending standards have changed over the past three months and the indicator we use is the net percentage of respondents reporting tightening standards in loans.¹ In the euro area, the bank lending survey has been introduced in 2001, see Berg et al. (2005) and European Central Bank (2003). Since then, major banks in the euro area have been reporting on the change in their lending standards.

¹The respondents characterize the changes in lending standards as “tightened considerably”, “tightened somewhat”, “basically remained unchanged”, “eased somewhat” and “eased considerably”. The indicator is compiled as the difference between the number of respondents reporting tightened standards and those reporting eased standards expressed as a percentage of all respondents.

To be consistent with the US series, we use the report about net tightening of loans to large enterprises (Question 1).²

The series are depicted in the lower-right panel of figure 1, in which the bold line represents the euro area series. The shaded areas refer to NBER dated recession periods. The correspondence between a high share of lenders tightening standards and recessions is obvious. There is a high correspondence between the US and the euro area time series, the correlation coefficient being 0.83. For the US, the historical high of 59.7 reached in the first quarter of 2001, has been exceeded during the financial crisis by the level of 83.6 reached in the fourth quarter of 2008, which means that virtually all banks tightened lending standards during this quarter. The net percentage of lenders tightening standards remained persistently considerably high for the subsequent three quarters before turning negative in 2010. Since then, with the exception of one quarter, the share of banks decreasing lending standards exceeded the share of the ones increasing or leaving them unaltered.

It is worth noting that the historical low levels around -20 lasted throughout

²The categories to report changes in lending standards are the same as in the SLO survey, see footnote 1. To take into account that a country's weight does not correspond to the country's lending share in the euro area, the responses are weighted by the country's lending share in total euro area lending when compiling the euro area figures. The net percentage of respondents tightening lending standards is then compiled as the difference between the percentage of respondents who tightened minus the percentage of respondents who eased standards.

2004 until the third quarter of 2005. The percentage of lenders easing lending standards exceeded the percentage of those tightening standards even until the third quarter of 2006. Thus, the majority of lenders eased lending standards consecutively for two and a half years, undoubtedly a consequence of the lasting period of low interest rate levels, decreasing below 2% from 2002 throughout 2004. In the euro area, the historical high of 64 in the first quarter of 2003 has been reached again 2008Q4 and 2009Q1. The net percentage of banks tightening standards came subsequently down to 43 and 21 in the second and third quarter of 2009, respectively, euro area banks apparently returning more sluggishly – or more cautiously – to less tight lending standards. The decrease might have been favoured by the first announcement of the ECB in 2009Q2 to perform a refinancing operation with a 12-month maturity. From 2010 onwards, the index remained at remarkably low levels, but levels in contrast to those of the US remaining in positive territory. Since then, the index has increased abruptly twice, first in 2010Q3 to a level increasing from 3 to 11 and second in 2011Q4 to a level increasing from 2 to 16 and further to 35 in 2012Q1. On the first occasion, the renewed decrease in the index correlates with the ECB's announcement in 2009Q3 about fine-tuning operations when the 12-month refinancing operation, announced

for the first time in 2009Q2, would phase out. On the second occasion, the decrease in the index is correlated to the settlement of two additional longer-term refinancing operations, one with a 12-month and one with a 13-month maturity, in the last quarter of 2010. So, besides reacting to interest rates it seems that, in particular during periods when low interest rates become binding, lending standards may be affected by the conditions characterizing the central bank's liquidity provision.

The correlation between the lending standards indicator and the interest rate is rather low (0.1) for the US, see table 2. When the Federal Funds rate is lagged by 1 quarter, the correlation increases to 0.2, and reaches a maximum of 0.5 when the interest is lagged by 6 periods. The repeal of parts of the Glass-Steagall Act (1933) in 1999 might have changed the interaction between interest rates and standards.³ When we split the sample at the end of 1999, the correlation structure changes significantly. Until 1999, the contemporaneous and the lag correlations are 0.65 and 0.70, respectively. From 2000 onwards, they decrease to 0.14 and 0.31, respectively. In the second sub-sample, the maximum correlation of 0.83 is reached when the

³The Financial Services Modernization Act, enacted in November 1999, repealed in particular the restrictions that prohibited any institution from combining any of the services provided by commercial banks, investment banks and insurance companies. It is often brought forward, that this financial deregulation is mainly to blame for the subsequent exuberance in financial markets and the crisis in the subprime mortgage market.

interest rate is lagged by 7 periods.

The corresponding correlations between the series for the euro area are moderate. The contemporaneous correlation is 0.25, and when the 1 month EURIBOR rate is lagged by 1 quarter it increases to 0.48. The maximum correlation of 0.66 is reached when the interest rate is lagged by 3 periods.

As additional variables in the VECM, we will include GDP growth, the producer price index (PPI) and C&I loans for the US and loans to non-financial corporations for the euro area, which are depicted in the remaining panels of figure 1.

3 Econometric approach

We estimate a vector error correction model (VECM) for a system of six variables to assess the role of lending standards. We include real GDP, the producer price index (PPI), the GDP deflator (the HICP for the euro area), C&I loans (loans to non-financial corporations), the Federal Funds Rate (the EURIBOR) and standards. Standards are ordered last, because the banking system is thought to adjust its lending standards within a quarter to changes in the policy rate, while policy is thought to react, if at all, only with a lag to changes in lending standards.

Given that we combine a non-integrated (lending standards) variable with integrated variables (all others) and that we want to take into account potential cointegration between the integrated variables, we estimate an unrestricted VECM using the Bayesian approach of Koop et al. (2010). Here, unrestricted means that we take into account cointegration without being specific about the cointegrating vectors. To this aim, we work with a uniform prior on the cointegration space, for details see below. The cointegration rank is determined by means of the Bayes factor, which is estimated by the Savage-Dickey density ratio (see Koop et al. (2008)). Working with a VECM rather than a level VAR also circumvents the problems raised in Phillips (1991) and Koop et al. (1995).⁴ The Bayesian approach provides the advantage that the posterior inference obtained for the US, in particular the moments of the posterior distributions of the parameters and the error covariance matrix, can be used to shape the prior specification of the euro area system. In this way, we are able to increase the estimation efficiency, which would be quite low because of the short sample size even if working

⁴Phillips (1991) stresses the need for ignorance priors (Jeffrey's prior) in order to remove the bias towards stationarity which is obtained for integrated series in posterior inference based on flat priors. Further, Koop et al. (1995) show that, even if the ignorance prior yields proper posterior distributions, it prevents the existence of one-period ahead predictive moments. This issue is relevant and has to be addressed, given that we will compute impulse response functions.

with a usual Minnesota type prior specification.

Since the method is based on recent developments in the Bayesian cointegration literature, we give a condensed motivation and description of it in the following, in particular the specification of the prior distribution and the sampling scheme. For a vector of N integrated variables Y_t , we write the vector error correction model (VECM)

$$y_t = \Pi Y_{t-1} + \sum_{j=1}^p \Gamma_j y_{t-j} + \varepsilon_t \quad t = 1, \dots, T \quad (1)$$

$$\varepsilon_t \sim \text{i.i.d}N(0, \Sigma)$$

where y_t denotes the vector of variables transformed to stationarity (usually growth rates or differences) and where deterministic terms are omitted for convenience. Under the assumption of cointegration, i.e. if r linear combinations of Y_t turn out to be stationary, the matrix Π is of reduced rank and can be spanned by two $N \times r$ matrices Johansen (1995). We obtain

$$\Pi = \alpha\beta' \quad (2)$$

$$y_t = \alpha\beta' Y_{t-1} + \sum_{j=1}^p \Gamma_j y_{t-j} + \varepsilon_t \quad (3)$$

where the columns of β contain the cointegration vectors and the rows of α contain the adjustment of each variable to past departures from the long-run relationship prevailing between the series, the so-called error term $ec_t \equiv$

$\beta'Y_{t-1}$.

The approach pursued in Koop et al. (2010) is motivated by the following observations:

- The matrix

$$\Pi = \alpha\beta' = \alpha\kappa(\beta\kappa^{-1})' = A\beta^{*'}$$

is not identified. Any nonsingular transformation of α and β yields identical Π matrices and we can switch between representations. The first in which β is assumed orthonormal, $\beta'\beta = I_r$, while α is unrestricted, and the second in which A is orthonormal while β^* is unrestricted. Alternatively, we specify $\kappa = (\alpha'\alpha)^{-1/2} = (\beta^{*'}\beta^*)^{1/2}$.

- Identification is typically achieved by using the linear normalization $\beta = \begin{pmatrix} I_r \\ b \end{pmatrix}$. However, Strachan and Inder (2004) show that such a normalization puts restrictions on the estimable region of the cointegration space. Moreover, a non-informative prior on b in fact favors regions in the cointegration space where the normalization is not valid. Finally, Villani (2006) shows that working with the linear normalization may lead to counter-intuitive results on cointegration.
- Conditional on β , the nonlinear VECM (1) becomes linear and for

appropriately specified priors, standard multivariate posterior inference is applicable. The strategy is to define a suitable prior for β , which will have implications for the prior on α , in order to obtain its posterior and draw from it.

Given the disadvantage of the linear normalization, Strachan and Inder (2004) propose to specify a prior on the cointegration space $\wp = sp(\beta)$ rather than on cointegrating vectors. This is achieved by introducing a semi-orthogonal $N \times r$ matrix H , which spans the same space as β , $sp(\beta) = sp(H)$.⁵ The prior on the cointegration space takes then the form of a matrix angular central Gaussian distribution with parameter P_τ , $MACG(P_\tau)$, (Chikuse, 1990):

$$\pi(\beta) \propto |P_\tau|^{-r/2} |\beta' P_{1/\tau} \beta|^{-n/2} \quad (4)$$

where the $N \times N$ matrix $P_\tau = HH' + \tau H_\perp H'_\perp$, determines the central location of $sp(\beta) = sp(H)$. The dispersion is controlled by $\tau \in (0, 1)$, which determines the departure from the cointegration space. A very dogmatic prior would set $\tau = 0$. On the other hand, setting $\tau = 1$ leads to $P_\tau = I_N$, a

⁵If the researcher has specific expectations in mind about the cointegrating vectors like e.g. the great ratios in a threevariate system of real GDP, consumption and investment, she could define $H^g = \begin{bmatrix} 1 & 1 \\ -1 & 0 \\ 0 & -1 \end{bmatrix}$. The matrix H^g is then transformed into the semi-orthogonal matrix H by $H = H^g (H^{g'} H^g)^{-1/2}$.

uniform prior on the Stiefel manifold.

For α , a standard normal prior with shrinkage parameter ν may be specified:

$$vec(\alpha)|\beta, \Sigma, \tau, \nu \sim N\left(0, \nu (\beta' P_{1/\tau} \beta)^{-1} \otimes G\right) \quad (5)$$

where G may be chosen freely. In the application we set it to I_N . For the rest of the parameters, the dynamics Γ_j , $j = 1, \dots, p$ and the error covariance matrix Σ , we assume, respectively, a Minnesota-type prior with a prior variance of 0.09 on autoregressive coefficients and a shrinkage factor of 0.25 for the prior variance of off-diagonal coefficients, and an inverse Wishart distribution with $N + 2$ degrees of freedom and scale matrix $S_0 = \kappa I_N$, specifically $\kappa = 0.5$. Combining these priors with the likelihood, we obtain a posterior normal distribution for α and Γ_j , $j = 1, \dots, p$, and an inverse Wishart for Σ .

The prior for β and α specified in (4) and (5), respectively, implies the following prior distribution for A and β^* :

$$vec(\beta^*)|A \sim N\left(0, (A'G^{-1}A)^{-1} \otimes \nu P_\tau\right) \quad (6)$$

$$\pi(A) \propto |G|^{-\tau/2} |A'G^{-1}A|^{-N/2} \quad (7)$$

Note that $vec(\beta^*)|A \sim N(0, I_r \otimes \nu P_\tau)$ and $\pi(A) \propto 1$ if $G = I_N$, given that

$A'A = I_r$. Combining this prior distribution with the likelihood yields again a normal posterior distribution for β^* from which we may sample.

The (collapsed) Gibbs sampler thus iterates over the following steps (Koop et al., 2010):

- i. Sample $\alpha, \Gamma_j, j = 1, \dots, p$ from $\pi(\alpha, \Gamma_j | \beta, \Sigma, y)$ and Σ from $\pi(\Sigma | \alpha, \beta, \Gamma_j, y)$
 Transform α to $A = \alpha(\alpha'\alpha)^{-1/2}$
- ii. Sample β^* from $\pi(\beta^* | A, \Gamma_j, \Sigma, y)$ and use it to transform to $\beta = \beta^*(\beta^*\beta^*)^{-1/2}$
 and $\alpha = A(\beta^*\beta^*)^{1/2}$
- iii. Eventually update the hyperparameters τ and ν in case a hierarchical prior was specified

Obviously, the outstanding advantage of the approach is the ability to sample parameters, α, β , which depend nonlinearly on each other, from normal posterior conditional distributions.

To decide on the cointegration rank we use the Savage-Dickey density ratio

$$B_{0r} = \frac{\pi(\alpha | M_r, y) |_{\alpha=0}}{\pi(\alpha | M_r)_{\alpha=0}}$$

where B_{0r} represents the Bayes factor to evaluate the model with cointegration rank r , M_r , against a model with no cointegration. The Bayes factors

obtained for various $r = 1, \dots, N$ can subsequently be used as weights in Bayesian model averaging or for probability evaluation on the number of cointegrating vectors. For details the reader may refer to Koop et al. (2008).

In the empirical investigation, we determine the cointegration rank for each system by choosing the specification obtaining the highest posterior odds ratio among all possible choices for the cointegration rank, see table 3. The distributions of impulse responses are available from the draws of the posterior distribution. For each posterior parameter draw we compute impulse responses, obtained by a Cholesky decomposition of the error covariance matrix. We depict the mean and the 95th percentile interval in the figures discussed below.

4 Results

Before we present the results for the euro area, which we are ultimately interested in, we first discuss the role of standards in the US sample. As discussed above we will then use this information to obtain more precise estimates for the euro area.

Panel (a) of Figure 2 displays the responses of the variables in the system to one standard deviation increase in lending standards for the US sample

obtained with the model estimated over the sample from 1990 to 2012. We see that real GDP and also loans decline significantly, although transitorily, after a tightening of lending standards, while neither of the two price variables responds significantly. These results are highly similar to those reported in Lown and Morgan (2006). We also see that the Federal Funds rate transitorily reacts negatively, thus potentially offsetting the effect that tighter standards may have on the financing conditions of firms. Although Lown and Morgan (2006) also find that the Federal Funds rate declines, it does so only insignificantly. Turning to the responses of standards to shocks in the other variables, reported in Panel (b), we see that positive innovations to GDP growth tend to significantly ease standards in the short run. After 8 quarters, standards tighten again, but the mean-reversion is insignificant. We also see that an increase in loan growth leads to a tightening of standards. Monetary policy contractions, in contrast, are first associated with significantly looser standards. Within one year however, standards significantly tighten again. Again, these results are largely in line with those reported in Lown and Morgan (2006), with the exception of the response to the interest rate shock, which, albeit also negative, is not significant in Lown and Morgan (2006).

Table 4 shows forecast error variance decompositions (see also the corresponding plots in Figure 3). We see from Panel (a) that shocks to lending standards account for 20 percent of the error variance in GDP at a horizon of 4 quarters and for roughly a third at a horizon of 12 quarters. For loans, shocks to standards account for almost two thirds of the error variance in the long run. While the shares are somewhat lower for the price variables and for the federal funds, we still find that standards account for about 20 percent of the error variances in these variables at longer horizons. Turning to Panel (b), we see that the error variance of standards is dominated by its own innovations, and also driven, to some extent, by the short term interest rate. Overall, these results are in line with the results in Lown and Morgan (2006) and confirm that bank lending standards are important in accounting for business cycle dynamics in the US.

Next, we turn to the analysis of the euro area sample. To estimate the system we design the prior distribution using the moments of the posterior distributions obtained for the US system. To be as consistent as possible for the sample period, we estimate the US system for the period 2000 to 2012. We do not display the responses to save space and because it turns out that the impulse responses do not substantially differ from those obtained for the

whole sample period, except for the responses corresponding to the producer price index. Producer prices significantly decline after a shock in standards. A shock in producer prices, on the other hand, puts upward pressure on standards, although insignificantly so. The responses for the euro area are displayed in Figure 4. Responses to innovations in standards are again shown in Panel (a), while Panel (b) shows the response of standards to shocks in the other variables.

As in the US, output and loans decline significantly following a tightening of standards. The response of loans is not as strong as in the US, but more persistent. Nevertheless, the negative output response is as strong as in the US, but less persistent. Overall these findings for the euro area qualitatively confirm the results reported by Maddaloni and Peydró (2013), de Bondt et al. (2010), Ciccarelli et al. (2010), and Cappiello et al. (2010), and suggest that non-price credit conditions matter for the business cycle in the euro area. While the response of producer prices is marginally negative in the short run, consumer prices do not react to a shock in standards over the whole time horizon considered. As in the US, the interest rate declines transitorily after a positive innovation to standards.

Panel (b) of the figure displays the reaction of standards to shocks in the

other variables. Overall, the responses, except the one to loan growth, which is insignificant, are stronger and quicker than in the US. A positive shock to the output growth rate leads to a significant hump-shaped response of standards. On impact, they are loosened significantly before being tightened again within one year. We can also document that standards (marginally) significantly tighten within one year after an interest shock occurred. This result is in line with Maddaloni and Peydró (2013) who find that low monetary policy rates in the euro area soften lending conditions, which they interpret as evidence in favor of a risk taking channel (see also de Bondt et al., 2010). In contrast to the US, where standards do not react to price shocks, a shock to producer prices leads to tighter lending standards within one year in the euro area, although insignificantly so. The reaction of consumer prices is negative on average, but also insignificant. The insignificance of these responses may also be due to the small number of observations.

To sum up, we find that the responses to shocks in standards are qualitatively and quantitatively similar in the US and in the euro area. In terms of persistence, there are slight differences. The response of GDP is less persistent in the euro area, whereas the response of loans is more persistent in the euro area. However, the reaction of standards to shocks in the other vari-

ables in general are stronger and quicker than in the US. Nevertheless, we conclude that the interrelationship between the business cycle, loan growth and standards show quite a few similarities in the two regions we analyze.

Table 5 shows the variance decomposition for the euro area system, Figure 6 shows it graphically. We see from Panel (a) that although shocks to lending standards play an important role for GDP growth and loans in the euro area, the respective shares of the error variance of those variables are generally smaller than in the US. For the short term interest rate, in contrast, the share of the error variance that can be attributed to standards is larger in the euro area, at any horizon considered. Concerning the error variance of standards, Panel (b) shows that, in contrast to our results for the US, the business cycle seems to play a non-negligible role since GDP growth accounts for 22 percent of the error variance at the 12 and 20 quarters horizons. Nevertheless, as in the US, it is primarily innovations in standards themselves that account for the bulk of error variance in this series.

Although we have shown that the role of standards is largely similar across the euro area and the US, one could argue that these results simply mirror the fact that we impose information from the US system on the estimation of the euro area system. To see if this is indeed the case, we re-estimate the

euro area system but without imposing U.S. information. We see from the impulse response functions in Figure 5 that the tendency of the responses goes towards the ones obtained when using the information from the US system, i.e. the ones depicted in figure 4. However, the standard error bands become considerably wider. Hence, we are confident that imposing U.S. information does not excessively bias the euro area responses towards those found for the US system.

5 Summary

The availability of survey data on bank lending behavior allows researchers to obtain more detailed descriptions of the role of the banking sector for the business cycle. Unfortunately, the euro area bank lending standards survey is relatively young, which complicates the empirical analysis due to the low number of observations. In this paper, we deal with this issue using Bayesian techniques. We obtain first evidence by using the posterior moments of the parameters' posterior distributions estimated for a VECM for US data to design informative prior distributions for the VECM for euro area data.

We find that tighter bank lending standards are associated with lower loan growth as well as lower output growth in the US and in the euro area.

The response of output in the euro area is as strong as in the US, but less persistent. On the other hand, the response of loans are weaker, but more persistent in the euro area than in the US. Standards in the euro area respond stronger and quicker to output shocks than in the US. The positive response to interest rate shocks occurs also sooner in the euro area than in the US, and is not negative on impact. Finally, we also find that in the euro area, in contrast to the US, bank lending standards are largely unresponsive to shocks in loan growth.

A Tables

Table 1: Data sources

	Euro Area ^{a)}	United States ^{b)}
Sample	2003:1-2012:3	1990:2-2012:3
Series		
Stand	Bank Lending Survey, question 1, net tightening of loans to large enterprises	Federal Reserve Board, Senior Loan Officer Opinion Survey on Bank Lending Practices, panel 1, net percentage of domestic respondents tightening standards for C&I loans to large and medium enterprises
Rate	1-month EURIBOR	Federal Funds Rate
GDP	euro area 12 (13 from 2009 onwards)	GDP real and nominal
Prices	HICP, euro area 12, seasonally adjusted	GDP Deflator
PPI	Monthly commodity price index, import-weighted	Producer prices
loans	Loans to non-financial corporations	C&I loans

^{a)} All data from the ECB's statistical website, Prices: quarterly data obtained from monthly averages

^{b)} Data from FRED[®], Federal Reserve Bank of St. Louis

Table 2: Correlation

Sample US	1990-2012	1990-1999	2000-2012
Corr(Stand,int.rate)	0.11	0.65	0.14
Corr(Stand,int.rate(-1))	0.20	0.70	0.31
Max Corr	0.50	0.71	0.83
	(lag 6)	(lag 2)	(lag 7)
Sample EA	2003-2012		
Corr(Stand,int.rate)	0.25		
Corr(Stand,int.rate(-1))	0.48		
Max Corr	0.66		
	(lag 3)		

Table 3: Bayes factor and probability of cointegration rank. In the six variable system there are five integrated variables. $\log BF_{0r}$ is the log of the Bayes factor of a model with zero cointegration against a model with cointegration rank r .

Sample period	US		EA with US prior info	
	1990-2012		2003-2012	
	$\log BF_{0r}$	$Prob(r)$	BF_{0r}	$Prob(r)$
$r = 1$	-5.00	0.00	-15.10	0.00
$r = 2$	-12.30	0.00	-23.71	0.00
$r = 3$	-25.93	0.00	-21.97	0.00
$r = 4$	-37.03	1.00	-30.38	1.00

Table 4: US: Forecast error variance decomposition

(a) Variance share attributable to a shock in standards						
horizon	GDP	PPI	Deflator	Loans	Short rate	Standards
0	0.00	0.00	0.00	0.00	0.00	0.73
4	0.19	0.05	0.03	0.38	0.09	0.66
8	0.29	0.11	0.09	0.60	0.16	0.60
12	0.32	0.16	0.14	0.65	0.19	0.58
20	0.34	0.23	0.22	0.67	0.26	0.60
(b) Variance share of standards attributable to a shock in						
horizon	Δ GDP	Δ PPI	Δ Deflator	Δ Loans	Δ Short rate	Standards
0	0.07	0.02	0.01	0.09	0.08	0.73
4	0.07	0.02	0.03	0.16	0.05	0.66
8	0.07	0.03	0.04	0.18	0.09	0.60
12	0.07	0.03	0.04	0.15	0.13	0.58
20	0.07	0.04	0.03	0.11	0.16	0.60

Table 5: EA: Forecast error variance decomposition

(a) Variance share attributable to a shock in standards						
horizon	GDP	PPI	Deflator	Loans	Short rate	Standards
0	0.00	0.00	0.00	0.00	0.00	0.73
4	0.21	0.02	0.04	0.17	0.19	0.69
8	0.27	0.05	0.08	0.31	0.31	0.56
12	0.26	0.11	0.11	0.36	0.34	0.52
20	0.24	0.19	0.15	0.38	0.35	0.50
(b) Variance share of standards attributable to a shock in						
horizon	Δ GDP	Δ PPI	Δ Deflator	Δ Loans	Δ Short rate	Standards
0	0.17	0.04	0.02	0.02	0.02	0.73
4	0.14	0.08	0.02	0.02	0.05	0.69
8	0.19	0.11	0.03	0.02	0.09	0.56
12	0.22	0.11	0.04	0.02	0.09	0.52
20	0.22	0.11	0.05	0.03	0.10	0.50

B Figures

Figure 1: US (long) and Euro area (short) time series. The shaded areas are NBER-identified recession periods.

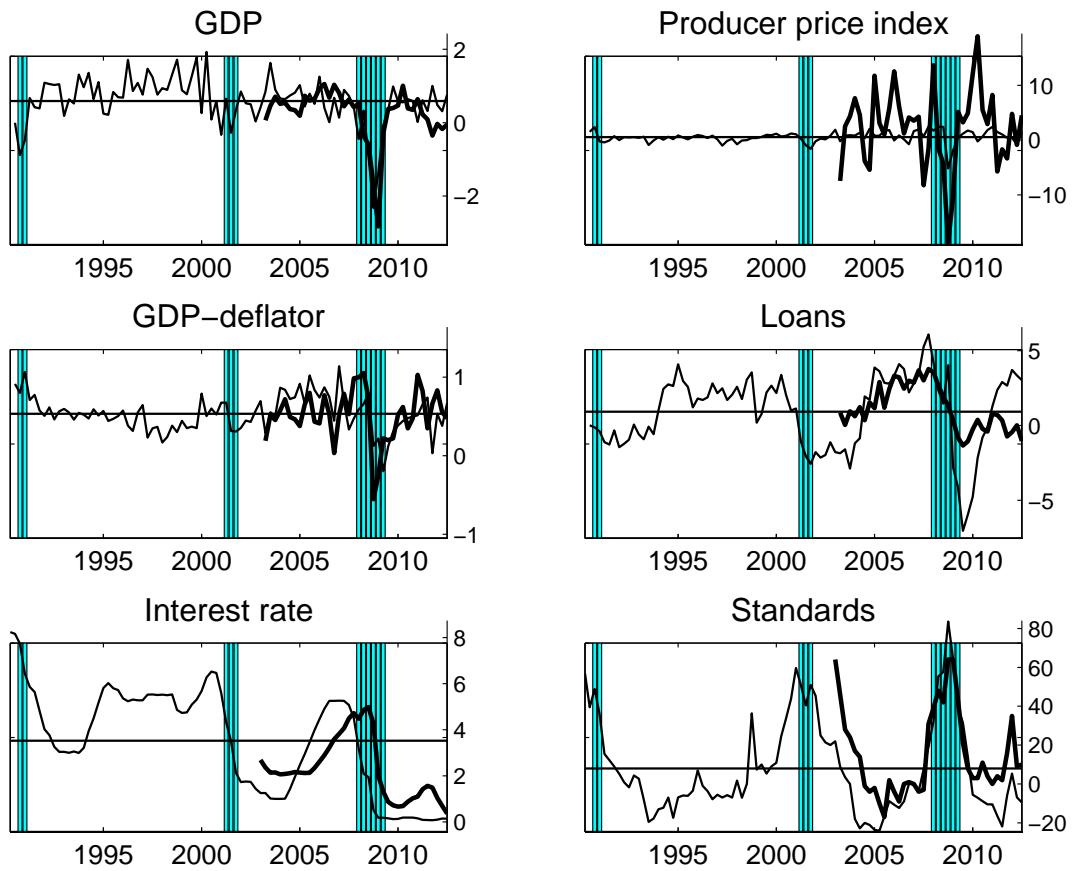


Figure 2: Impulse responses for the US: Sample period 1990-2012

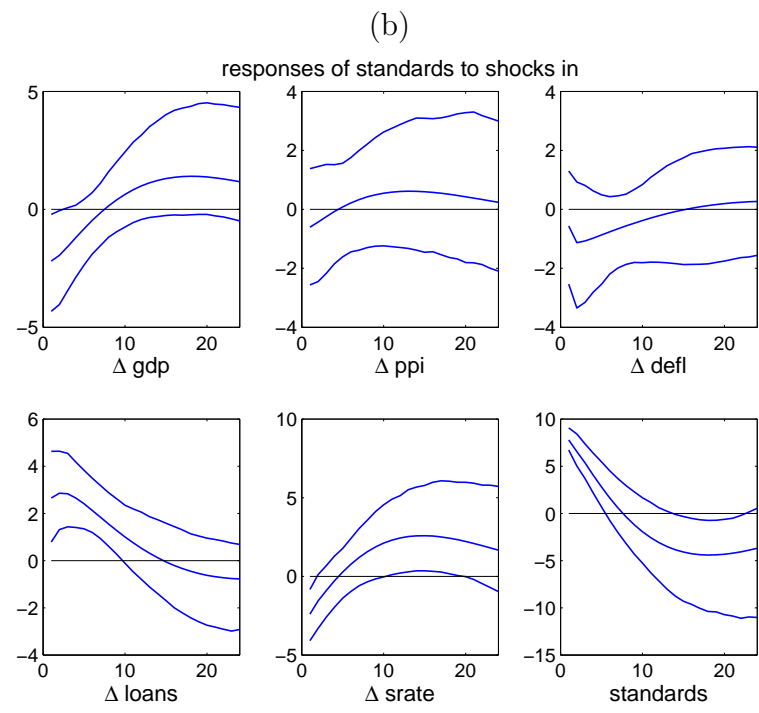
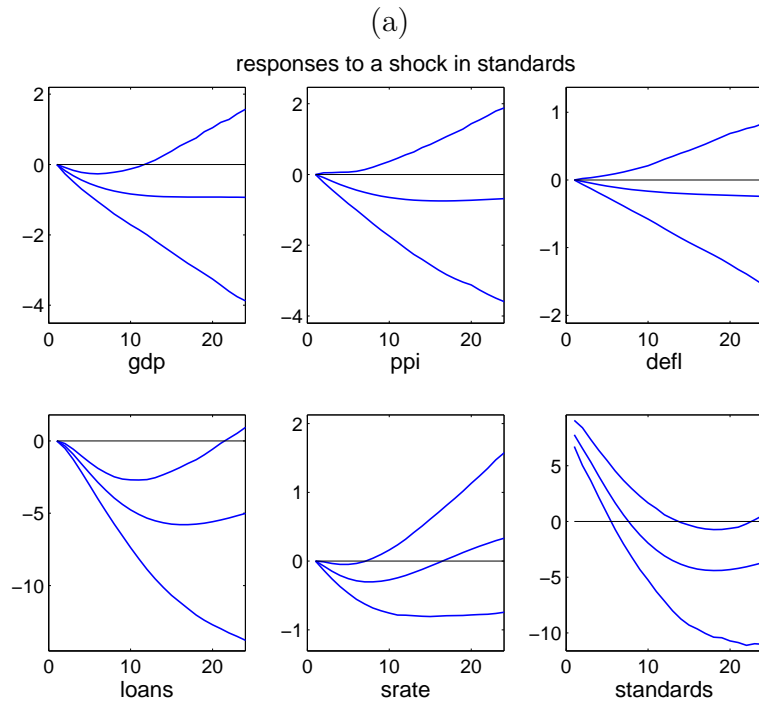


Figure 3: Forecast error variance decomposition for the US: Sample period 1990-2012

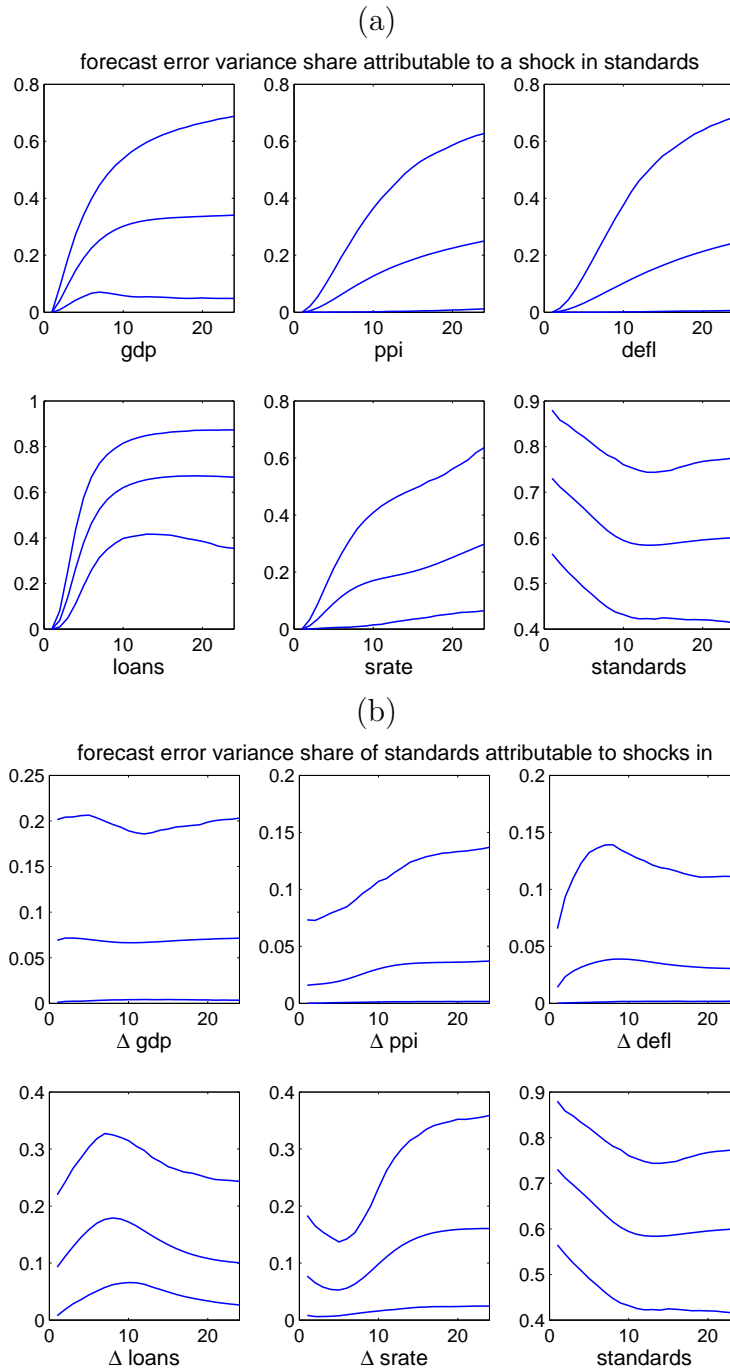


Figure 4: Impulse responses for the EA: Sample period 2003-2012, with prior information from the US system (Sample 2000-2012)

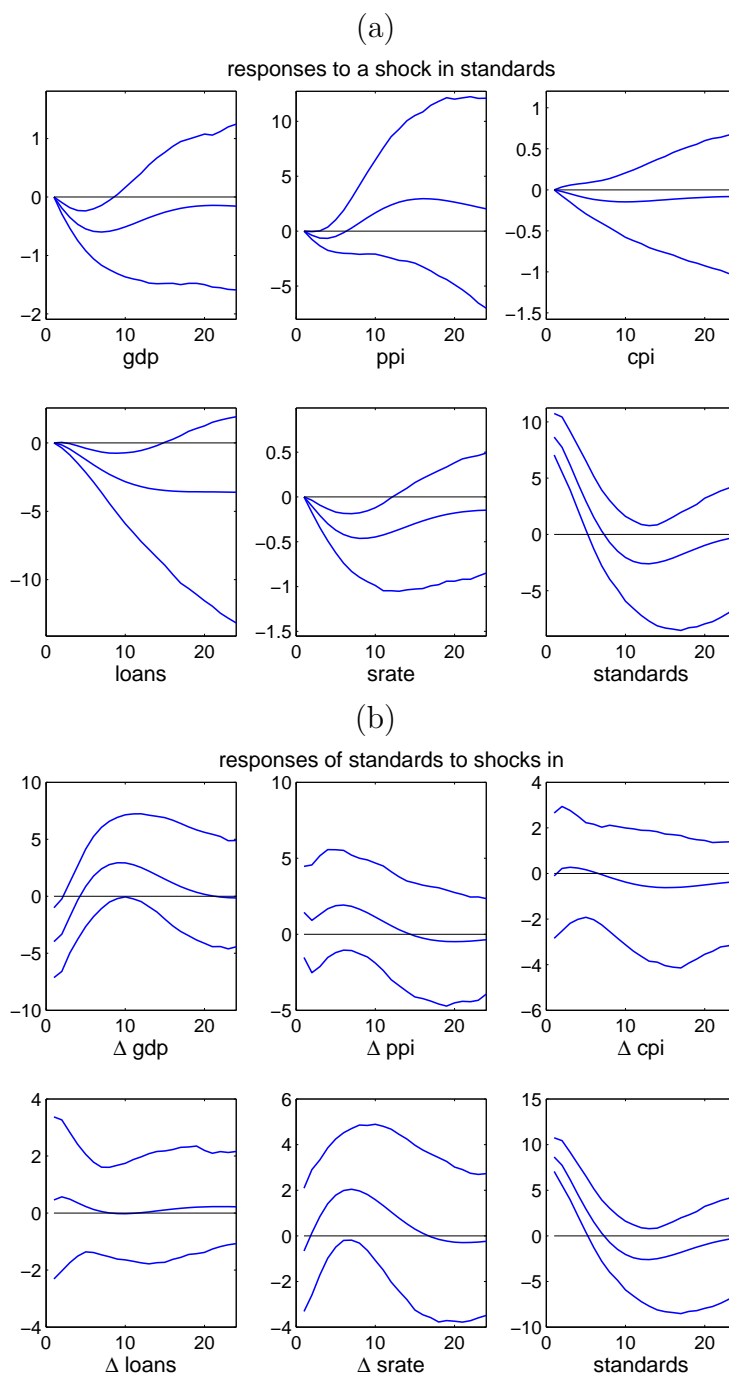


Figure 5: Impulse responses for the EA: Sample period 2003-2012, without prior information from the US system

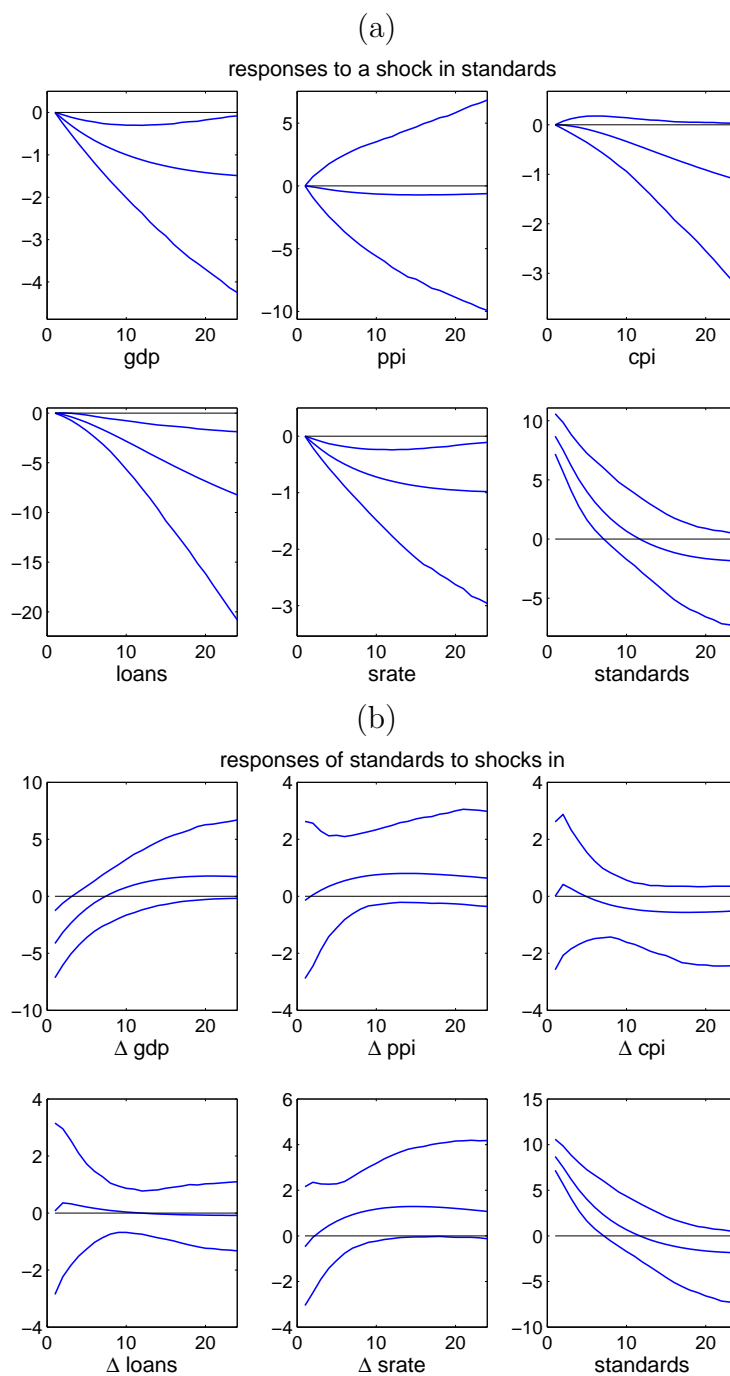
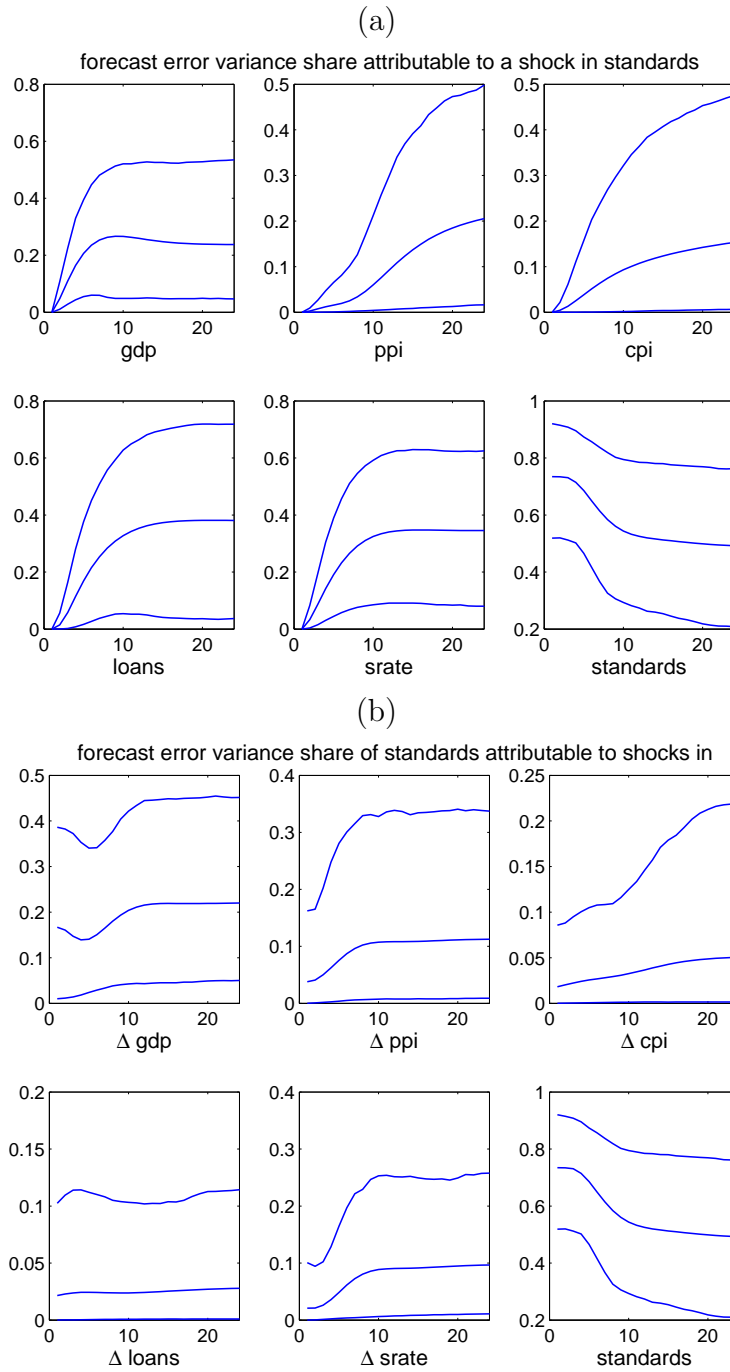


Figure 6: Forecast error variance decomposition for the EA, with prior information from the US system (Sample 2000-2012): Sample period 2003-2012



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Bank-lending standards, loan growth and the business cycle in the Euro area

Abstract

We study the relationship between bank lending standards, loan growth and the business cycle in the euro area and the US within a vector error correction model using Bayesian estimation methods. To deal with the short data series available for the euro area, we exploit information from the estimated US system to improve the estimation of the euro area system. We find that tighter bank lending standards are associated with lower loan growth as well as lower output growth in both areas. Differences in reactions appear in the strength and the persistence of responses.

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