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Bank-Lending Standards, the Cost Channel and Inflation Dynamics

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Abstract

If firms borrow working capital to finance production, then nominal interest rates have a direct influence on inflation dynamics, which appears to be the case empirically. However, interest rates may only partly mirror the cost of working capital. In this paper we explore the role of bank lending standards as a potential additional cost source and evaluate their empirical importance in explaining inflation dynamics in the US and in the euro area.

Keywords: New Keynesian Phillips Curve, Cost Channel, Bank Lending Standards, Bayesian Analysis

JEL codes: E40, E50

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1 Introduction

If firms have to borrow working capital to finance production, the nominal interest rate represents a cost factor and therefore influences price-setting behavior. These effects have been labeled the cost channel transmission of monetary policy. Several studies find that a cost channel has implications for monetary policy: Ravenna and Walsh (2006) argue that a cost channel limits the scope for monetary stabilization policy. Tillmann (2009a) shows that uncertainty about the strength of the cost channel influences the optimal setting of interest rates by the central bank and Tillmann (2009b) finds that a cost channel dampens the impact of model uncertainty on monetary policy.

Empirical evidence indicates that the cost channel adds substantially to the explanation of inflation dynamics (Tillmann 2008). Moreover, the direct effect of interest rate changes on inflation is typically found to be relatively strong, which is somewhat surprising for a number of reasons: Firms may not have to borrow the entire costs of production in advance Ravenna and Walsh (2006), or alternatively only a part of the firms in the economy may be subject to a cost channel. In either case, the response of the inflation rate should be smaller than the change in the interest rate. In addition, the interest rates relevant for working capital may not respond fully to changes in money market rates. Especially retail interest rates are typically rigid. Hence, banks may shelter firms from large changes in the cost of working capital (Chowdhury et al. 2006; Hülsewig et al. 2006; Kaufmann and Scharler 2009).

These considerations have been reconciled with the empirical evidence by arguing that interest rates do not represent the entire cost of working capital. Chowdhury et al. (2006) argue that broadly defined financial frictions result in additional costs, which are not directly mirrored in interest payment. The purpose of this paper is to explicitly allow for indirect cost channel effects in addition to those directly related to nominal interest rates.

Our analysis is based on the New Keynesian Phillips Curve augmented by the short-term interest rate and bank lending standards as proxy for indirect costs associated with working capital. We assess the role of standards for inflation dynamics using a vector autoregression which we estimate within a Bayesian framework. As the effects of lending standards may depend on the financial system, we estimate systems for the US, as an example for a market-based system, and the euro area which is characterized by a bank-based financial system. Since only relatively short series are available for the euro area, the Bayesian method has some advantages. In particular, we are able to evaluate whether differences between US and euro area estimates are due to a lack of data for the euro area or indicate differences in the transmission mechanism between both regions. Lending standards for the euro area are only available since 2003 which does not allow us to obtain precise estimates. However, we will use the posterior inference about US data to design prior information for the euro area system.

We find that lending standards counteract the interest rate effect on price-setting in the US. That is, a tightening of monetary policy, for instance, is accompanied by looser lending standards, which dampens the higher cost of working capital. For the euro area we find only limited evidence that lending standards matter for inflation dynamics. When we exclude lending standards from the system, we obtain a larger impact of interest rate shocks on inflation. This result suggests that if indirect cost effects associated with bank lending standards are not explicitly accounted for, then the direct influence of the interest rate on price setting appears to be larger. A posterior predictive test on data correlations confirms a well-specified system for the US. Test results for the euro area show that using US posterior information to design the euro area system's prior helps in capturing correlations between the interest rate and standards, and partially between the interest rate and unit labor costs. This suggests that the relation between interest rates and lending standards are similar across both regions. On the other hand, US posterior as prior information does not affect the correlation between inflation and standards, it deteriorates the correlation between inflation and the interest rate. This suggests that, as longer time series will be available, the transmission mechanism may turn out to be different between the regions.

Only few paper analyze empirically the role of bank lending standards. The implications of bank lending standards for the business cycle in the US are explored in Lown and Morgan (2006). Using the confidential euro area country-specific responses to the Bank Lending Survey of the European Central Bank, Maddaloni and Peydrò (2009) study, among other issues, the impact of the overnight interest rate level on lending standards. Our analysis differs from these two papers in the sense that we focus on inflation dynamics and the transmission of policy shocks to inflation via lending standards.

The remainder of the paper is structured as follows: In Section 2 we derive the augmented New Keynesian Phillips Curve which is the basis for our analysis. Section 3 discusses our empirical methodology while Section 4 describes our data set. In Section 5 we represent our estimation results. The importance of lending standards for the transmission mechanism is assessed in Section 6 and in Section 7 we evaluate the usefulness of using the posterior inference about the US to design the prior specification of the euro area's system. Section 8 summarizes and concludes the paper.

2 Theoretical Motivation

Our analysis is based on the New Keynesian Phillips curve augmented by bank lending standards as a factor influencing marginal production cost. To derive the New Keynesian Phillips curve, we closely follow Galí et al. (1999) and Galí et al. (2001). Hence, the discussion will be brief. We assume that the business sector of the economy consists of a continuum of monopolistically competitive firms normalized to have unit mass. Each firm i hires labor, h_{it}, and produces a differentiated good according to: $y_{it} = h_{it}^{1-\alpha}$. Each firm sells its output at a price p_{it} and faces the demand curve $y_{it}^d = (p_{it}/p_t)^{-\epsilon} y_t$, where p_t and y_t denote the aggregate price level and aggregate output. As in Calvo (1983), each period, a fraction $(1 - \theta)$ of the firms is able to adjust its price.

To introduce a cost channel, we follow Ravenna and Walsh (2006) and Chowdhury et al. (2006) and assume that firms have to finance the wage bill in advance of production. Hence, firms have to borrow an amount equal to the wage bill, $w_t h_t$, where w_t is the nominal wage. We assume that the total cost associated with financing of working capital is $\kappa_t R_t^l$, where R_t^l is the interest rate and κ_t captures non-interest borrowing costs. Thus, total expenditure is $\kappa_t R_t^l w_t h_t$.

For $\kappa_t = 1$, borrowing costs consist entirely of interest payments. If $\kappa_t > 1$ firms incur additional borrowing costs beyond the interest rate. If for instance, banks tighten their lending standards during times of rising interest rates, firms may incur additional costs as they may have to provide more collateral. Chowdhury et al. (2006) argue that financial frictions in a broad sense may amplify the cost effects of interest rates to rationalize large cost channel effects.

Cost minimization implies that

$$\kappa_t R_t^l \frac{w_t}{p_t} = mc_{it}(1-\alpha)\frac{y_{it}}{h_{it}}, \tag{1}$$

where mc_{it} denotes marginal cost. Note that (1) implies that $\widehat{mc}_t = \hat{\kappa}_t + \hat{R}_t^l + \hat{s}_t$, where $s_t = (w_t h_t)/(p_t y_t)$ denotes unit labor costs, and hatted variables denote percentage deviations from the steady state.

As in Galí et al. (1999) and Galí et al. (2001) we allow for inflation persistence by introducing firms that follow a backward looking pricing rule. Only a fraction $(1-\omega)$ of the firms which can set prices in the current period, resets prices optimally. The remaining firms follow a backward looking rule.

Combining these assumptions gives rise to the New Keynesian Phillips Curve:

$$\hat{\pi}_t = \lambda \widehat{mc}_t + \gamma_f E_t \hat{\pi}_{t+1} + \gamma_b \hat{\pi}_{t-1}, \tag{2}$$

where $\lambda = \frac{(1-\theta)(1-\theta\beta)(1-\alpha)(1-\omega)}{(1+\alpha(\epsilon-1))}\phi^{-1}$, $\gamma_f = \beta\theta\phi^{-1}$, $\gamma_b = \omega\phi^{-1}$, $\phi = \theta + \omega(1-\theta(1-\beta))$, β is the discount factor, and π_t denotes the inflation rate. The dynamics of marginal costs, $\widehat{mc}_t$, are determined by the borrowing rate, non-interest borrowing costs and wage costs, see equation (1) above.

Moreover, we allow that the interest rate at which firms borrow working capital, R_t^l, tracks the money market interest rate, R_t, only imperfectly. More specifically, we assume that $\hat{R}_t^l = \psi \hat{R}_t$, as it is common in the literature on the cost channel (Chowdhury et al. 2006; Hülsewig et al. 2006). Thus, we obtain

$$\hat{\pi}_t = \lambda\hat{\kappa}_t + \lambda\psi\hat{R}_t + \lambda\hat{s}_t + \gamma_f E_t\hat{\pi}_{t+1} + \gamma_b\hat{\pi}_{t-1}. \tag{3}$$

Hence, in addition to direct interest rate effects, firms may be subject to additional non-interest borrowing costs. If κ_t is constant over time, we obtain the formulation in Chowdhury et al. (2006). If in addition, $\psi = 1$ then (3) reduces to the interest rate augmented New Keynesian Phillips Curve derived in Ravenna and Walsh (2006). In our empirical analysis, we will use bank lending standards, st_t, as a proxy for non-interest borrowing costs. To do so, we assume that $\kappa_t = \kappa_0 + \kappa_1 st_t$, and therefore $\hat{\kappa}_t = \widehat{st}_t$. Thus, the dynamics κ_t mirror fluctuations in standards.

There are several methods to assess the relevance of direct cost effects. One possibility, followed in Chowdhury et al. (2006), is to estimate equation (3) by general methods of moments. The other possibility is to calibrate a model to the countries or regions under investigation (Kaufmann and Scharler 2009). The second approach would require quantifying the direct effect of non-interest borrowing costs on inflation, for which so far reference literature is, to our knowledge, not available. Therefore, we will estimate the Phillips curve with empirical data to obtain the inference on the importance of non-interest borrowing costs.

3 Method and empirical strategy

To assess empirically the relevance of the cost channel and of lending standards in particular, we first extend equation (3) to allow for general effects:

$$\hat{\pi}_t = \lambda_0 \hat{\kappa}_t + \lambda_1 \psi \hat{R}_t + \lambda_2 \hat{s}_t + \gamma_f E_t \hat{\pi}_{t+1} + \gamma_b \hat{\pi}_{t-1}. \tag{4}$$

Under rational expectation agents form unbiased expectations. Hence, realized inflation $\hat{\pi}_{t+1}$ turns out to be expected, $E_t \hat{\pi}_{t+1} = \hat{\pi}_{t+1}$. Solving for $\hat{\pi}_{t+1}$ yields the system

$$\begin{bmatrix} \hat{\pi}_{t+1} \\ \hat{\pi}_t \end{bmatrix} = \begin{bmatrix} \frac{1}{\gamma_f} & -\frac{\gamma_b}{\gamma_f} \\ 1 & 0 \end{bmatrix} \begin{bmatrix} \hat{\pi}_t \\ \hat{\pi}_{t-1} \end{bmatrix} + \begin{bmatrix} -\frac{\lambda_0}{\gamma_f} & -\frac{\lambda_1}{\gamma_f}\psi & -\frac{\lambda_2}{\gamma_f} \\ 0 & 0 & 0 \end{bmatrix} \begin{bmatrix} \hat{\kappa}_t \\ \hat{R}_t \\ \hat{s}_t \end{bmatrix} \tag{5}$$

which, in order to endogenize the cost variables, is extended to a vector autoregressive system

$$\begin{bmatrix} \hat{\pi}_t \\ \hat{s}_t \\ \hat{R}_t \\ \hat{\kappa}_t \end{bmatrix} = \begin{bmatrix} \frac{1}{\gamma_f} \quad -\frac{\lambda_2}{\gamma_f} \quad -\frac{\lambda_1}{\gamma_f}\psi \quad -\frac{\lambda_0}{\gamma_f} \\ \hdashline \\ \mathbf{A_1} \\ \\ \end{bmatrix} \begin{bmatrix} \hat{\pi}_{t-1} \\ \hat{s}_{t-1} \\ \hat{R}_{t-1} \\ \hat{\kappa}_{t-1} \end{bmatrix} + \begin{bmatrix} -\frac{\gamma_b}{\gamma_f} \quad 0 \quad 0 \quad 0 \\ \hdashline \\ \mathbf{A_2} \\ \\ \end{bmatrix} \begin{bmatrix} \hat{\pi}_{t-2} \\ \hat{s}_{t-2} \\ \hat{R}_{t-2} \\ \hat{\kappa}_{t-2} \end{bmatrix} + \varepsilon_t \tag{6}$$

with the error term $\varepsilon_t \sim N(0, \Sigma)$.

This system can be estimated with standard methods. The Bayesian framework we adopt here implies some advantages, in particular to assess the significance of the channel for the euro area and the differences between the US and the euro area. We estimate the system by Bayesian Markov chain Monte Carlo methods. For the interested reader, the sampler is described in Kaufmann and Valderrama (2008).

The influence of unexpected changes in the interest rate, in standards and unit labor costs on inflation is assessed by computing structural impulse response functions identified by a Cholesky decomposition of the error covariance matrix. This decomposition is based on the variable ordering given in (6). The interest rate is ordered second-last as lending standards may well react contemporaneously to interest rate shocks. Interest rates, reflecting monetary policy changes, are less expected to react contemporaneously to a shock in lending standards. A variance decomposition complements the impulse response analysis, with which we can assess the relative importance of shocks in the forecast error variance of the variables.

When interpreting the results, we have to take into account that without any further assumptions on the interest rate smoothing parameter ψ, see equation (6), the effect of the direct interest rate cost, λ_1, is not identified. Moreover, the reaction of inflation to the third structural shock not only contains the cost-push effects of interest rate increases but also the usual demand effect stemming from the interest rate channel. Thus, the impulse response contains the net effect of both channels, conveying evidence about their relative strength.

To assess the importance of standards as non-interest borrowing costs for firms, the system is compared to one in which standards are excluded. If standards are important, we expect that their effect will be picked up by the interest rate and the unit labor costs in the reduced system.

Lending standards for the euro area have been published on a quarterly basis only since 2003. Of course, this is too few data to obtain a precise estimate. However, we will use the posterior inference obtained with US data as prior information for the euro area system. The comparison between this estimate and results obtained with a standard Minnesota-inverse Wishart prior for the VAR-parameters and the error covariance matrix yields first evidence of whether differences in the cost channel transmission are observable between the two regions.

Finally, a posterior predictive test on data correlation complements the analysis and serves as model diagnostic.

4 The data

The data used to estimate the Phillips curve are taken from the ECB's statistical website for the euro area and from the Federal Reserve Board's website and from the International Financial Statistics (IFS) databank for the US (see also Table 1). The beginning of the estimation sample is given by the start of the lending standards series in both regions. By the time of the investigation, for the US the observation sample has been running from the second quarter of 1990 to the third quarter of 2008 and for the euro area the sample begins in 2003 and runs through the third quarter of 2008. Although the sample for the euro area is very short, the Bayesian approach pursued in the paper yields first results to compare whether dynamics are different between regions. The results of the euro area are assessed by using the US posterior inference to shape the prior distribution of the VAR-parameters and the error covariance matrix of the euro area's system. The evaluation yields first evidence of whether differences between both regions are due to lack of data for the euro area or due to the different design of financial systems, i.e. market-based versus bank-based 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 show 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, van Rixtel, Ferrando, de Bondt, and Scope 2005, European Central Bank 2003). Since then, major banks in the euro area have been reporting on the change in their lending standards. To be consistent with the US series, we use the report about net tightening of loans to large enterprizes (Question 1).[2]

[1] The 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.

[2] The 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

The series are depicted in the upper-left 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 recently been exceeded by 83.6 in the fourth quarter of 2008. Lately, the net percentage of lenders tightening standards has come down to 64.2 and 39.6 in the first two quarters of 2009. It is worth noting that the historical low levels around -20 lasted throughout 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 67 in the first quarter of 2003 has been exceeded by 1 percentage point in the fourth quarter of 2008. The net percentage tightening standards has come down to 63 and 48 in the first two quarters of 2009, euro area banks apparently returning more sluggishly – or more cautiously – to less tight lending standards.

The correlation between the lending standards indicator and the interest rate is rather low for the US. Contemporaneously, they nearly are uncorrelated (-0.03), and when the Federal Funds rate is lagged by 1 quarter, the correlation coefficient is 0.1. The corresponding correlation between the series for the euro area are positive. The contemporaneous correlation is 0.16, and when the 1 month EURIBOR rate is lagged by 1 quarter it increases to 0.55.

The bottom panels in figure 1 show the series for the unit labor costs and for the inflation rate. The unit labor costs for the US, which correspond to the unit labor costs of non-financial corporations, are obtained from the Bureau of Labor Statistics. The series for the euro area represents total unit labor costs. Finally, the inflation rate of the US is computed on the basis of the consumer price index (CPI), which is retrieved from the IFS database. The harmonized index of consumer prices (HICP) forms the basis for the euro area inflation rate.

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compiled as the difference between the percentage of respondents who tightened minus the percentage of respondents who eased standards.

The variables enter in the system (6) in deviation from trends, for which we account in the following manner. Given that the Federal Reserve Board and the European Central Bank are devoted to price stability, inflation is not expected to trend but to fluctuate around a certain level rate. Inflation thus enters in levels and a constant accounts for the long-run level rate. The trend in unit labor costs is removed by taking the difference of the logarithmic level, the series enters in growth rates into the system. Interest rates are differenced, given that they usually are borderline non-stationary without a long-term drift, however. Finally lending standards are included in levels, given that the series oscillate around 0 and have an upper and lower bound (100 and -100, respectively). Thus, the variables in equation (6) will have as empirical counterparts the inflation rate and lending standards in levels, real unit labor costs in growth rates and the interest rate in differences.

5 Results

We estimate model (6) by Bayesian Markov chain Monte Carlo (MCMC) methods. We sample 23,000 times from the posterior distribution, discard the first 8,000 to remove dependence on initial conditions, and retain every third draw to remove dependence across the simulations. The US sample is restricted to end in the third quarter of 2007, given that a preliminary analysis identified a regime change in the dynamics of the data in the fourth quarter of 2007, during which the subprime market crisis became virulent and led to the still ongoing deep financial market crisis.[3] There are too few observations to obtain a reliable inference for this period. Therefore, we cap the sample.

According to the cost channel, monetary policy exerts supply side effects on the economy since variations in interest rates influence marginal costs of production. The main purpose of our analysis is to explicitly account for lending standards in this framework. Thus, what we are primarily interested in, is how lending standards influence the transmission of interest rate shocks to the inflation rate.

Figure 2 shows impulse responses for the US along with the 90th percentile interval. We see that a contractionary policy shock, that is a positive innovation to the interest

[3]Although we omit the results, they are available upon request. The estimated posterior probability of a regime change in the third quarter of 2007 is well above 75% and increases to nearly 100% for the rest of the observations.

rate equation, increases prices as well as the unit labor costs, although the increase is insignificant in the second case. Note that the price reaction to the interest rate shock confirms the so-called price puzzle frequently found in the literature and is consistent with the interpretation that monetary policy induces non-negligible supply effects in the short-run. That is, the cost-push effect being larger than the negative demand-side effect of monetary policy shocks.

In contrast, lending standards transitorily decline in response to the policy shock. Thus, the dynamics of lending standards appear to counteract the monetary tightening. Put differently, although interest rates increase, banks loosen lending standards. To the extent that lending standards proxy indirect costs associated with financial intermediation, our results suggest that these costs transitorily decline when monetary policy is tightened.

This result is somewhat surprising, as one would expect an increase in lending standards, resulting in additional costs of working capital and therefore giving rise to additional inflationary pressure. However, it appears that bank lending standards do not amplify but partially mute the cost effects of monetary policy, at least in the US.

Turning to the euro area, the impulse responses in Figure 3 were obtained by using the posterior inference on the US to shape the prior distribution of the model parameters. In the following, we present the results obtained using what we call the full US prior information. This refers to a prior specification which uses the hyperparameters of the posterior distribution of the VAR parameters and of the error covariance matrix to design the prior distribution of the euro area system's model parameters. The resulting conjugate priors are multivariate-normal and inverse Wishart for the VAR parameters and the error covariance matrix, respectively. Using the US posterior as prior information in the euro area system helps in capturing some dynamic features of the data and allows to obtain relevant and significant results, although the euro area time series are quite short. This will be shown in Section 7.

In Figure 3, we see that a contractionary monetary policy shock reduces the price level and increases unit labor costs, although the responses are only marginally significantly different from zero. Note that we do not find a price puzzle with euro area data. The response of standards is also negative, as in the US, but insignificant. Thus, although the role of lending standards is not as pronounced as in the US, we find essentially no evidence in favor of the hypothesis that banks amplify the impact of monetary shocks by adjusting lending standards. This contrasts a bit the results of Maddaloni and Peydrò (2009), who find a positive effect of the lagged overnight interest rate level on lending standards. Their results rely on the analysis of the unweighed panel of euro area country-specific responses to the Bank Lending Survey, which are not publicly available. As longer time series become available, it will be possible to re-assess our aggregate VAR results. Note, however, that shocks to the lending standards have a positive effect on prices which is marginally significant. Thus, although at the aggregate level bank lending standards in the euro area do not appear to influence the cost channel transmission of monetary policy, shocks to the lending standards exert some influence on prices. We conclude that variations in euro area bank lending standards therefore represent a source of inflation dynamics, which contrasts with our results for the US.

Interestingly, the forecast error variance decompositions in Tables 2 and 3 show that the interest rate and standards account for roughly the same fraction of the variance in prices. For the US, after five years the shares increase to 7% and 9%, respectively. Moreover, the interest rate and standards each account for much larger fractions, around four times larger, than unit labor costs. Thus, we find that cost channel effects are more relevant than wage costs in the US. According to Table 3, a similar conclusion emerges for the euro area. At a five-year horizon, interest rate and standards shocks account even for a larger fraction, respectively 10% and 14%, in HICP forecast error variance. The share of unit labor costs increases to 7%.

6 The relevance of lending standards

To get a better picture of how lending standards impact on inflation dynamics, we estimate the systems without standards. Figures 4 and 5 show the responses when we drop standards from the system. We see that in the US as well as in the euro area, the inflation rate now responds substantially stronger to the increase in the interest rate, suggesting the presence of a cost channel in both regions, being twice as large in the US than in the euro area. For the US, the effect on the inflation rate is about twice as large as before. For the euro area, the response is now positive, significantly so. Thus, we find that dropping the lending standards from the system, the direct effect of the interest rate on price setting becomes larger. This result is consistent with the hypothesis that the interest rate picks up some of the effect of non-interest rate cost effects, if those are not explicitly taken into account. Thus, the strong, direct cost channel effect documented in Chowdhury et al. (2006) and Ravenna and Walsh (2006) may at least partially be attributed to neglected effects of indirect, non-interest rate costs of working capital.

Although the change in impulse responses is considerable, the variance decomposition for the US in Table 4 does not document an increased share of interest rate shocks in the CPI forecast error variance. At the five-year horizon the share increases to 6%, compared to 7% when standards are included. Overall, for all variables the variance share accounted for by standards in the larger system is absorbed, at all horizons, in a higher share of inflation shocks in the reduced system.

For the euro area we observe the same consequences. Excluding standards inflates the variance share of inflation shocks in the HICP and unit labor forecast error variances. The error variance of the interest rate is mainly accounted for by own shocks in this specification.

7 US posterior as prior information

In this section we evaluate the influence of using the hyperparameters of the posterior distributions inferred for the US system to design the prior distributions of the euro area system's VAR-parameters and error covariance matrix. We obtained the results presented so far by using what we call the full US prior information, the setting in which the prior distributions of both the VAR-parameters and the error covariance matrix are designed with hyperparameters originating from the US system's posterior. We also show that in general, the US prior information designing the VAR-parameters influences the location of the posteriors, while the additional information on the error covariance matrix helps in increasing the estimation precision. The prior design using only US information on the VAR-parameters is called partial US prior. The results are evaluated against the posterior obtained with a standard Minnesota prior design for the VAR-parameters, with a prior mean of 0 and a prior variance of 0.09 and a shrink factor of 1. An inverse Wishart distribution designs the error covariance matrix with scale $S_0 = \nu I_4$, $\nu = 0.175$, and degrees of freedom $s_0 = 6$, to obtain an expectation of $E(S) = S_0/(s_0 - (N+1)/2) = 0.05I_4$ (with $N = 4$ the number of variables), and a mode mode$(S) = S_0/(s_0 + (N+1)/2) = 0.02$. This prior also designs the US system.

7.1 Influence on posterior distributions

Figure 6 depicts the posterior distributions of the first lag of the VAR-parameters obtained under the different prior designs. We observe that the distributions are mainly affected by adding information of the US posterior distribution on the VAR-parameters (the partial US prior). Adding additionally posterior information on the error covariance matrix (full US prior) does not significantly affect the posterior. This justifies why we previously reported the results for the euro area system obtained under the full US prior design.

The P-values reported in the figure show that the first own autoregressive lag on standards and on the interest rate are significantly affected by including US prior information. The probability of an estimated value larger than the posterior mean obtained under the Minnesota prior design is 97% for standards and even 100% for the interest rate. Thus, including US prior information introduces some persistence is these equations. The posterior distribution of the interest rate coefficient on prices is shifted significantly to the left, the P-value is 0.03. This is interesting, given that for the US system we obtain a positive reaction of prices to interest rates.

The location of the other posterior distributions is not significantly influenced by using US prior information. However, generally, US prior information affects the precision of the estimation. This is reflected in Figure 7, which displays the posterior distributions of the error covariance matrix. The distributions of the error variances of the unit labor cost and of the interest rate equations are shifted to the right. The P-values, reporting the probability of an estimated value larger than the posterior mode obtained under the Minnesota-inverse Wishart prior, are 1.0 and 0.92, respectively. The distribution for the error variance of the standards equation is shifted to the left. In general, the covariance structure is not significantly affected by including US prior information. Again, these results justify the reporting of euro area results obtained with the full US prior specification.

7.2 Influence on impulse responses

The effects of including US prior information on impulse responses are evaluated by comparing the posterior distribution of responses at the 4-, 8- and 16-period horizon, which corresponds to the one-, two- and four-year horizon, respectively. The impulse response distributions of standards and unit labor costs do not significantly change when including US prior information. Therefore, we only depict the impulse responses of prices and the interest rate in Figures 8 and 9, respectively. In Figure 8, second line, we see that including US prior information shifts the reaction of prices to a shock in the interest rate to the left. Using the full US info design accentuates the tendency obtained under the partial US prior design. The P-values, reporting the probability of a larger posterior response estimated under the full US prior than the posterior mean of the distribution estimated under the Minnesota prior, indicate that the shift, relative to the mean, would be marginally significant at the 10% significance level. The reaction of prices to a standards shock is shifted to the right, not as significantly as for the responses to the interest rate shock, however.

The responses of the short-term interest rate to a standards shock (see Figure 9, first line) is shifted to the left. Nevertheless, the precision in the estimates remains quite low and overall, the response is insignificant.

7.3 Posterior predictive test

Finally, we apply a posterior predictive test not only as a model diagnostic tool. It also serves as an additional means of assessing whether the use of US prior information helps in capturing data features. As test features, we choose the correlation structure between the data, contemporaneous and lag/lead correlations.

To perform the test, we simulate N=1000 data replications out of the predictive posterior distribution. The sample length is set to the observed sample T=67 for the US. For the euro area, we simulate data of observed sample length, T=19, and of a relatively large sample, T=100. Generally, the tendency of the test based on the observed data sample length is accentuated when the test is based on the larger data sample length. Therefore, we present test results obtained with data of sample length T=100 for the euro area.[4]

[4]The results based on simulated data of observed data sample length are available upon request.

To quantify the significance, we again report P-values, which summarize the probability of a larger correlation in simulated data than in observed data. For the euro area, we provide three P-values, which correspond to simulated data from the predictive posterior obtained under the different prior designs.

For the US, the test serves as a means of model diagnostic. Figure 10 displays, against observed data correlation, the distribution of the correlations between simulated price data and simulated (leads and lags of) standards, the interest rate and unit labor costs. The P-value of the simulated correlation between prices and unit labor costs lagged by two periods (0.08) indicate marginal significant departure from data features. We also obtain a significant P-value (which is not displayed to save space) for the correlation between standards and unit labor costs lagged by four periods (0.04). Nevertheless, given that all other simulated correlations well capture data features, we conclude that the model for US time series is adequately specified. This also justifies using the US posterior distribution to design the prior distribution of the euro area system.

Figure 11 displays the posterior predictive tests for the correlations with respect to euro area HICP inflation. The correlations between simulated lead values of standards and prices are not well captured by the system, irrespective of the prior distribution design. Nevertheless, in these cases the P-values only indicate marginal significant departure from data features. On the contrary, in the second column we see that the correlations between the simulated interest rate and simulated inflation do not well reproduce data correlations, which are anyway weak in the short observation sample. In our view, this clearly indicates the need for longer time series to obtain corroborating posterior evidence of the results we presented so far.

The P-values in the final column of Figure 11, summarizing the test for the correlation between unit labor costs and prices, yield evidence that including in particular US prior information on the error covariance matrix cancels out observed data correlation, at least up to a lead/lag of unit labor cost of two periods. This gives additional evidence, that both regions may show different dynamic adjustments to shocks, depending on their financial market design. This also will have to be assessed when longer time series will be available.

The P-values in Figure 12, first column, show that US information helps in capturing data correlation between the interest rate and standards, except for the correlation with standards leading by four periods. The test results for the correlations between the interest rate and unit labor costs (second column) are independent of the prior design. Basically, data features are well captured, if we disregard the significant departure from observed data correlation of the simulated correlation with unit labor costs leading by four periods.

In the last column of Figure 12, the results for the correlation between unit labor costs and standards are mixed. The contemporaneous and the correlation with leading standards deteriorate when US information is included. On the other hand, the correlation between lagging standards and unit labor costs are better captured with US prior information. To summarize, US prior helps capturing data correlations between the interest rate and lagging standards, and partly between the interest rate and unit labor costs.

8 Conclusions

In this paper, we analyze the extent to which bank-lending standards as a proxy for non-interest costs of financing working capital give rise to supply side effects of monetary policy and how bank lending standards influence inflation dynamics more generally.

We find that in the US, the cost channel is attenuated by lending standards. Put differently, in response to rising interest rates, banks transitorily lower their lending standards and thereby reduce the cost of working capital.

Since the series available for the euro area are too short for a meaningful analysis, we impose US prior information when estimating the model for the euro area. With US prior information lending standards apparently do not transmit monetary policy shocks, but appear to be a significant source of shocks themselves. That is, shocks to the lending standards result in fluctuations in prices.

Although the model appears to be well specified for the US, there is room for improvement for the euro area specification. We expect that longer data series will improve the model specification. In particular, the fit in the correlation between inflation and the interest rate and unit labor costs is expected to improve.

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A Tables

Table 1: Data sources

	Euro Area[a)]	United States[b)]
Sample	2003:1-2007:4	1990:2-2007:3
Series		
stand	Bank Lending Survey, question 1, net tightening of loans to large enterprises	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
ulc	Euro Area 12, total unit labor cost	Bureau of Labor Statistics, unit labor costs non-financial enterprises
prices	HICP-overall monthly index, seasonally adjusted	IFS: CPI

[a)] All data from the ECB's statistical website, HICP: quarterly data obtained from monthly averages

[b)] If not otherwise stated, data from the Federal Reserve Board's website

Table 2: US: Forecast error variance decomposition

	CPI inflation, attributable to			
horizon	π	s	R	κ
0	1.00	0.00	0.00	0.00
4	0.91	0.02	0.05	0.02
8	0.87	0.02	0.06	0.04
12	0.85	0.02	0.07	0.06
20	0.82	0.02	0.07	0.09
	ULC, attributable to			
horizon	π	s	R	κ
0	0.02	0.98	0.00	0.00
4	0.03	0.93	0.02	0.02
8	0.03	0.90	0.03	0.04
12	0.03	0.87	0.04	0.06
20	0.03	0.84	0.04	0.09
	Interest rate, attributable to			
horizon	π	s	R	κ
0	0.02	0.02	0.97	0.00
4	0.02	0.04	0.77	0.17
8	0.03	0.06	0.57	0.34
12	0.03	0.07	0.47	0.44
20	0.03	0.07	0.39	0.51
	Standards, attributable to			
horizon	π	s	R	κ
0	0.03	0.03	0.08	0.86
4	0.04	0.05	0.04	0.87
8	0.04	0.06	0.04	0.86
12	0.04	0.06	0.04	0.86
20	0.04	0.06	0.04	0.86

Table 3: Euro area: Forecast error variance decomposition, full US prior

	HICP, attributable to			
horizon	π	s	R	κ
0	1.00	0.00	0.00	0.00
4	0.85	0.04	0.05	0.06
8	0.77	0.06	0.07	0.10
12	0.73	0.06	0.08	0.12
20	0.69	0.07	0.10	0.14
	ULC, attributable to			
horizon	π	s	R	κ
0	0.10	0.90	0.00	0.00
4	0.13	0.76	0.06	0.04
8	0.13	0.69	0.11	0.07
12	0.13	0.65	0.13	0.09
20	0.12	0.61	0.15	0.11
	Interest rate, attributable to			
horizon	π	s	R	κ
0	0.05	0.04	0.92	0.00
4	0.08	0.05	0.82	0.04
8	0.10	0.06	0.75	0.08
12	0.11	0.07	0.71	0.10
20	0.12	0.07	0.68	0.13
	Standards, attributable to			
horizon	π	s	R	κ
0	0.13	0.12	0.04	0.71
4	0.17	0.14	0.05	0.64
8	0.18	0.14	0.06	0.62
12	0.18	0.14	0.07	0.61
20	0.18	0.14	0.08	0.60

Table 4: US: Forecast error variance decomposition, without standards

	CPI inflation, attributable to		
horizon	π	s	R
0	1.00	0.00	0.00
4	0.94	0.02	0.04
8	0.92	0.02	0.05
12	0.92	0.02	0.06
20	0.92	0.03	0.06
	ULC, attributable to		
horizon	π	s	R
0	0.01	0.99	0.00
4	0.15	0.83	0.03
8	0.22	0.74	0.04
12	0.25	0.70	0.05
20	0.28	0.67	0.05
	Interest rate, attributable to		
horizon	π	s	R
0	0.08	0.02	0.90
4	0.11	0.03	0.86
8	0.14	0.03	0.83
12	0.15	0.03	0.82
20	0.16	0.03	0.81

Table 5: Euro area: Forecast error variance decomposition, full US prior, without standards

	HICP inflation, attributable to		
horizon	π	s	R
0	1.00	0.00	0.00
4	0.96	0.01	0.03
8	0.95	0.01	0.04
12	0.95	0.01	0.04
20	0.95	0.01	0.04
	ULC, attributable to		
horizon	π	s	R
0	0.08	0.92	0.00
4	0.24	0.74	0.02
8	0.33	0.64	0.03
12	0.37	0.60	0.03
20	0.40	0.57	0.03
	Interest rate, attributable to		
horizon	π	s	R
0	0.03	0.04	0.93
4	0.06	0.08	0.86
8	0.08	0.08	0.84
12	0.09	0.09	0.83
20	0.09	0.09	0.82

B Figures

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

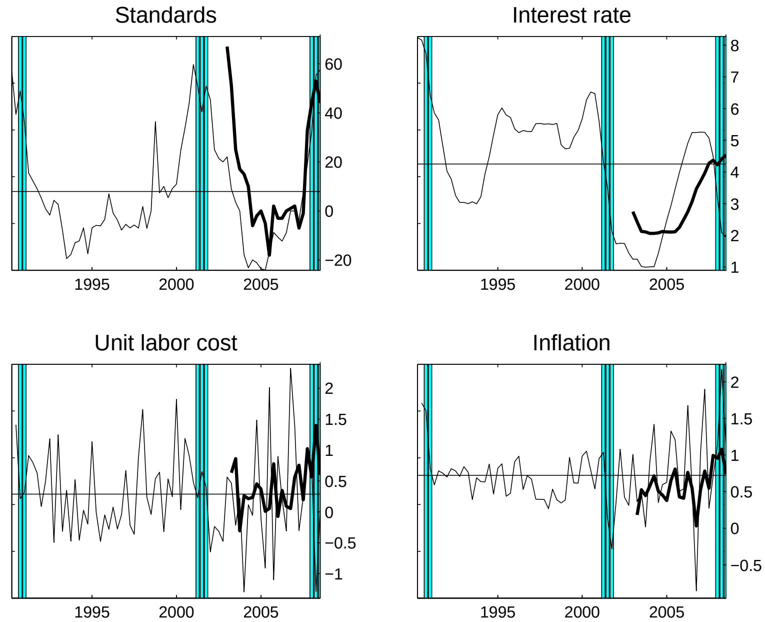


Figure 2: US: Impulse responses with 90th percentile interval

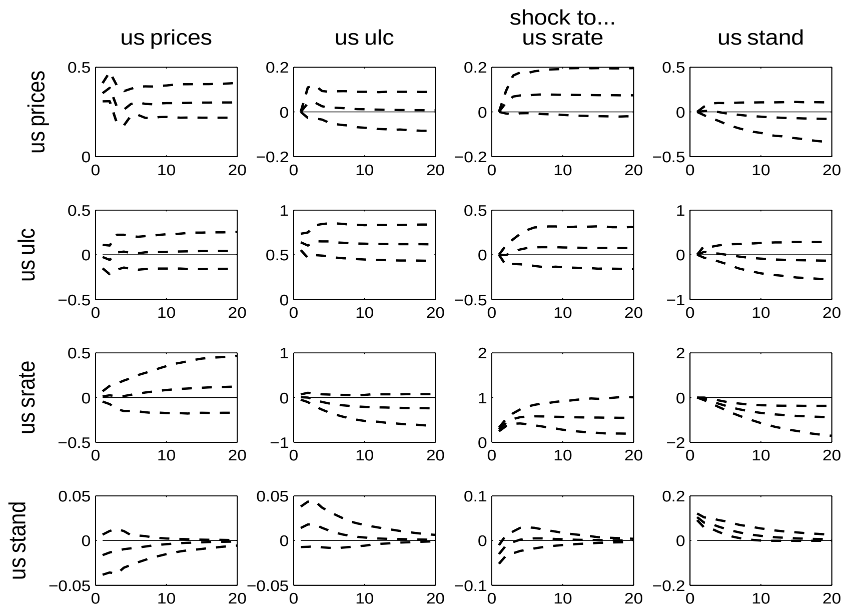


Figure 3: Euro area: Impulse responses with 90th percentile interval, full US prior

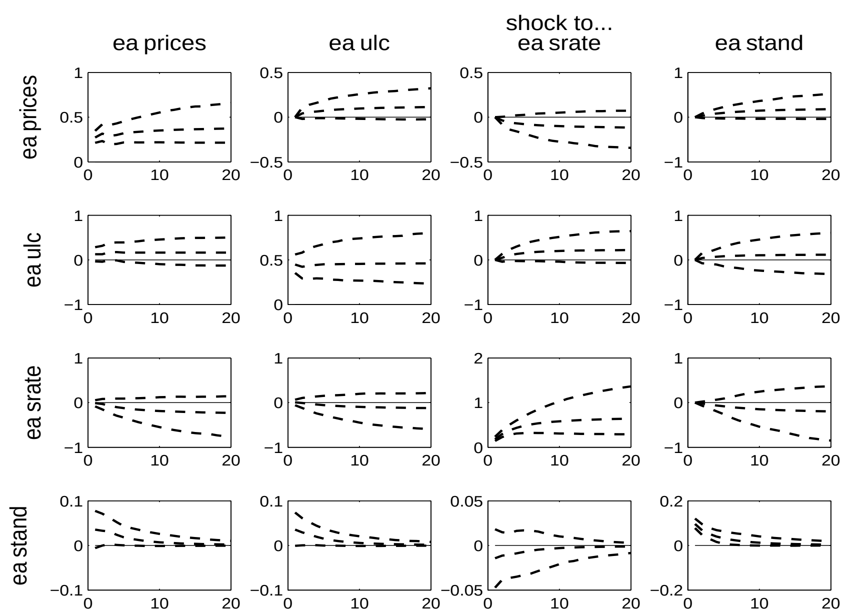


Figure 4: US: Impulse responses with 90th percentile interval, without standards.

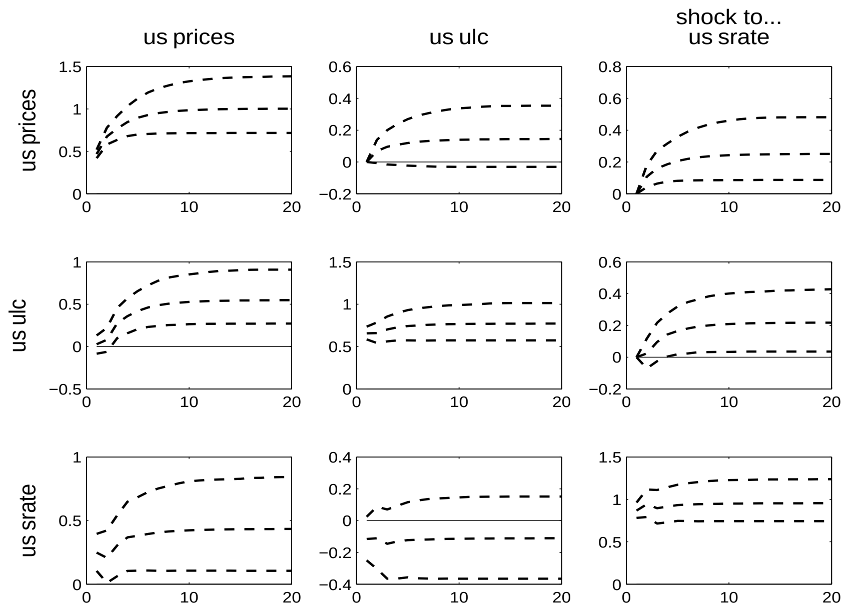


Figure 5: Euro area: Impulse responses with 90th percentile interval, full US prior, without standards.

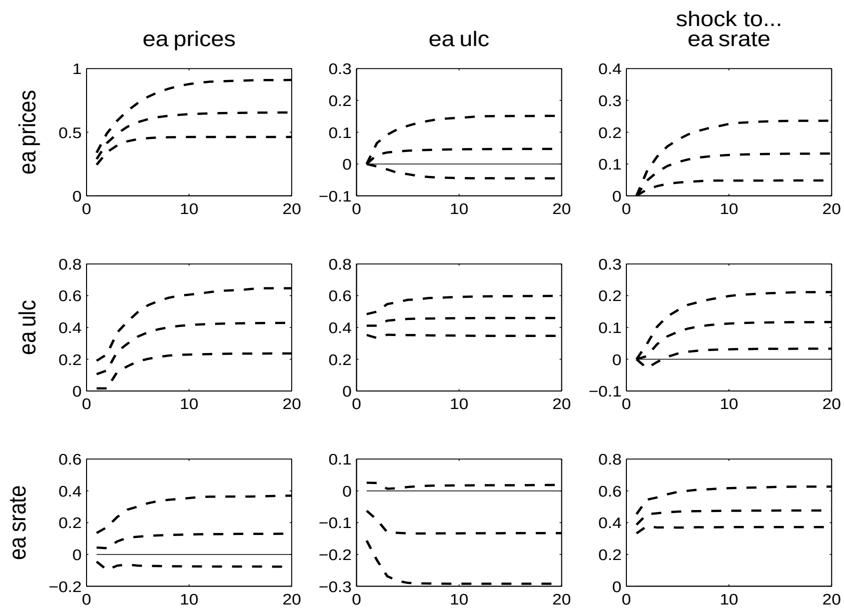
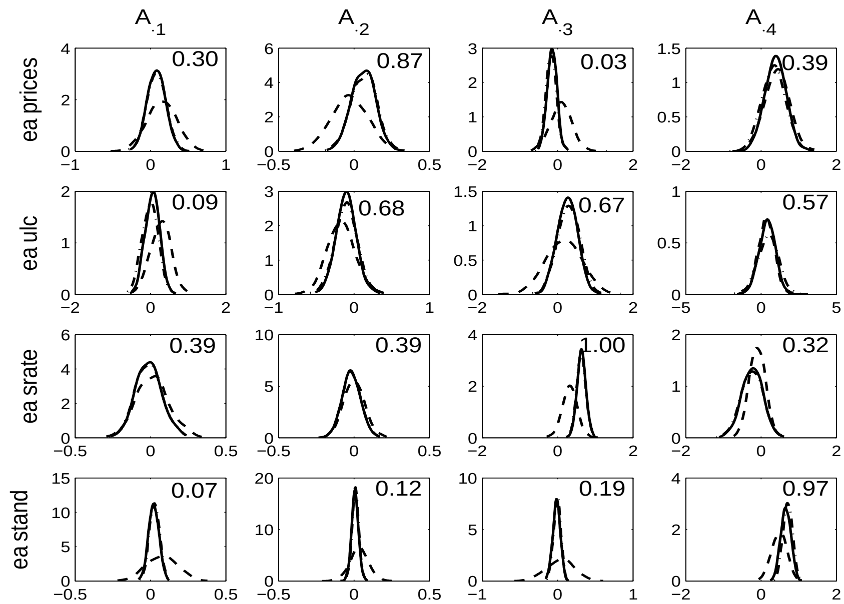


Figure 6: Euro area: Posterior of parameters on the first autoregressive lag, using US posterior as prior. P-value: Probability of a posterior value estimated with full US prior (-.-) exceeding the posterior mean estimated with the Minnesota prior (- -), $P\left(a_{\cdot \text{US Info}} > \overline{a}_{\cdot \text{Minnesota}}\right)$.



(- -) Minnesota prior, (-) Partial US prior, (-.-) Full US prior

Figure 7: Euro area: Posterior of error covariance matrix, using US posterior as prior. P-value: Probability of a posterior value estimated with full US prior (-.-) exceeding the posterior mean estimated with the Minnesota prior (- -), $P\left(\sigma_{\cdot \text{US Info}} > \text{mode}(\sigma_{\cdot \text{Minnesota}})\right)$.

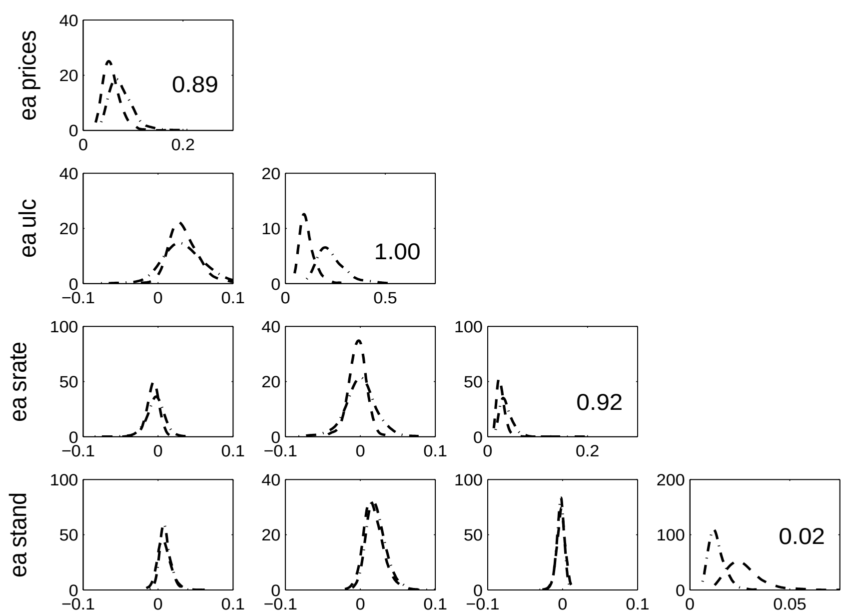
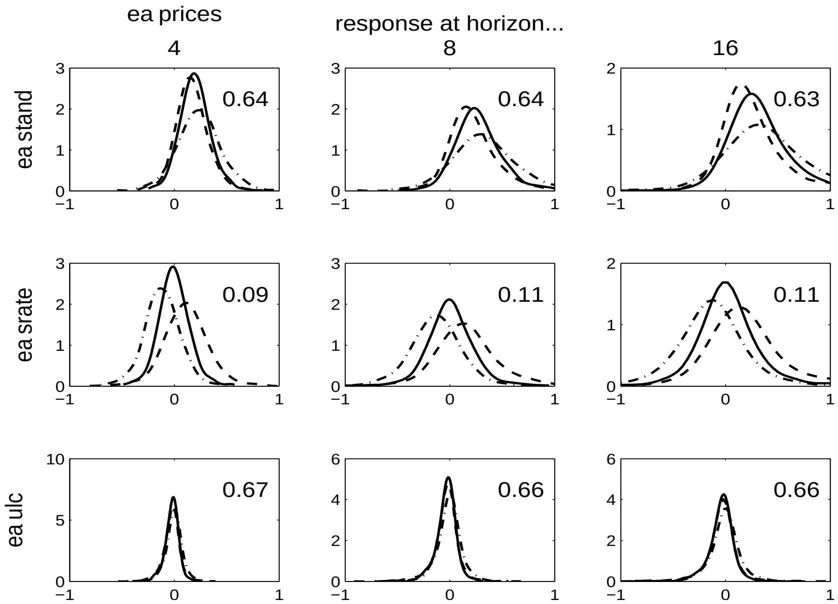
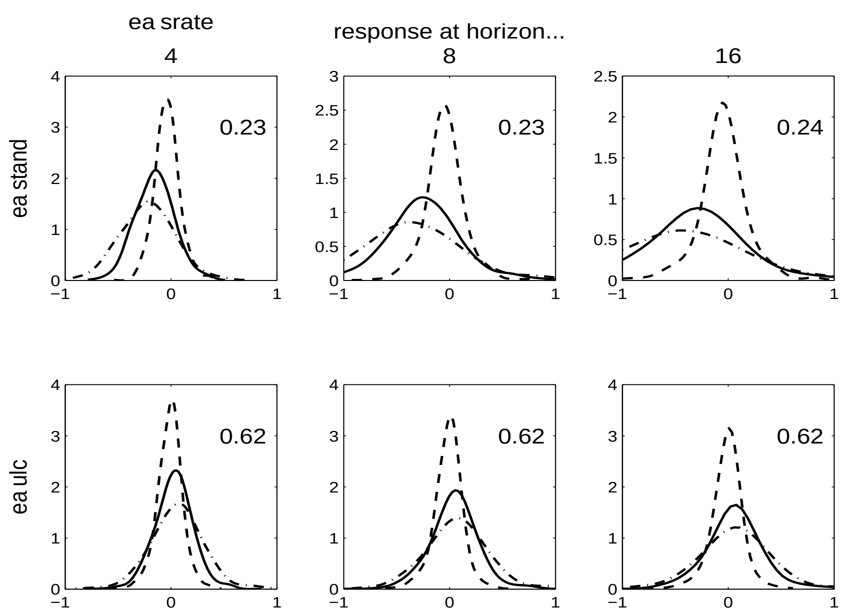


Figure 8: Euro area: Impulse response distribution of prices at various horizons. P-value: Probability of a posterior response estimated with full US prior (-.-) exceeding the posterior mean response estimated with the Minnesota prior (- -), $P\left(\text{resp}_{\text{US Info}} > \overline{\text{resp}}_{\text{Minnesota}}\right)$.



(- -) Minnesota prior, (-) Partial US prior, (-.-) Full US prior

Figure 9: Euro area: Impulse response distribution of the short-term interest rate at various horizons. P-value: Probability of a posterior response estimated with full US prior (-.-) exceeding the posterior mean response estimated with the Minnesota prior (- -), $P\left(\text{resp}_{\text{US Info}} > \overline{\text{resp}}_{\text{Minnesota}}\right)$.



(- -) Minnesota prior, (-) Partial US prior, (-.-) Full US prior

Figure 10: US: Correlation with CPI inflation. P-value: Probability that a simulated correlation exceeds the data correlation, $P\,(\text{corr} > \text{corr}_{\text{data}})$.

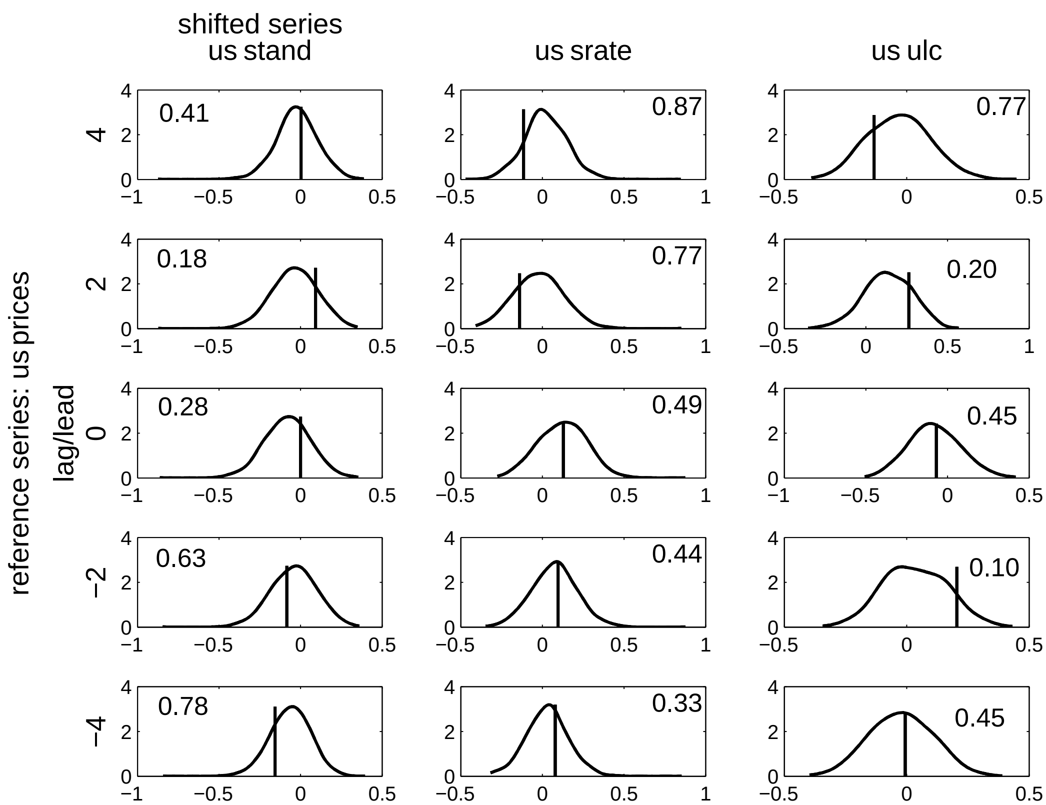
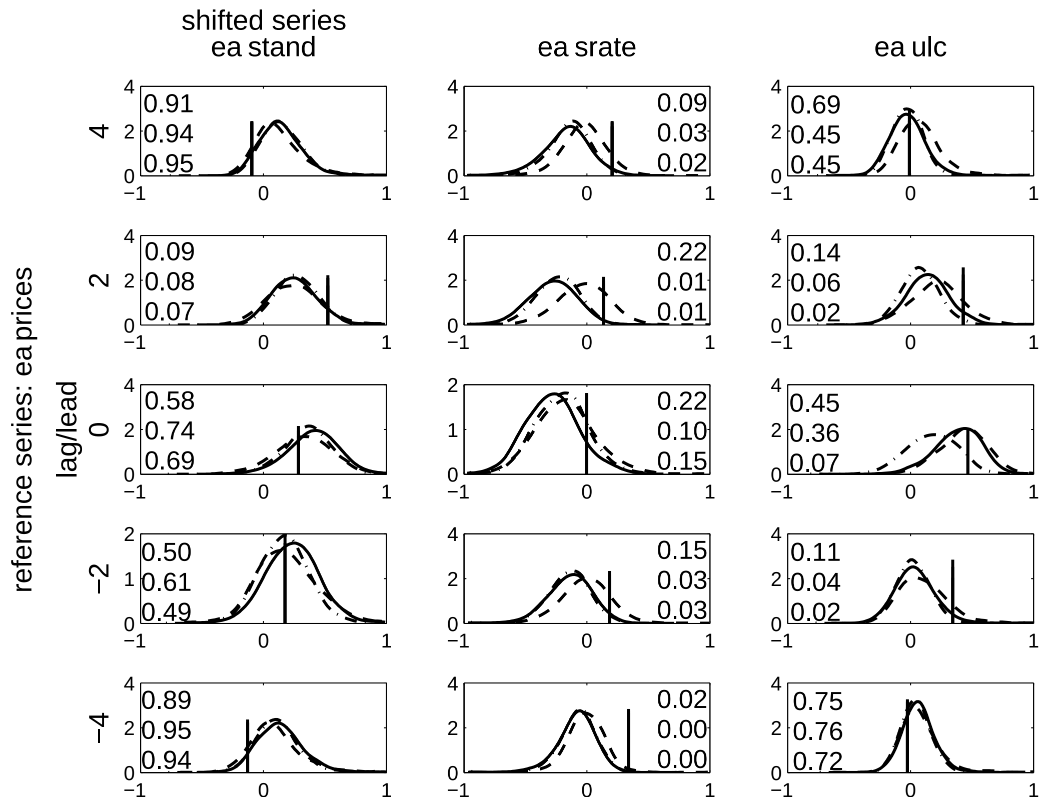
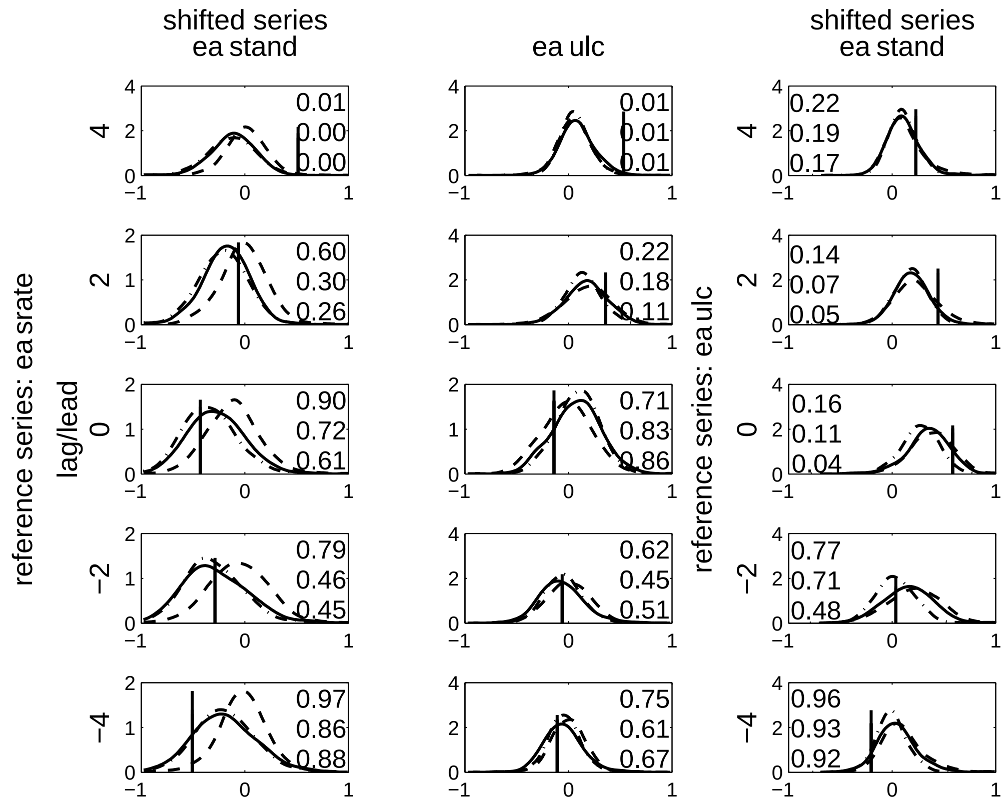


Figure 11: Euro area: Correlation with HICP inflation, $T = 100$. P-value: Probability that a simulated correlation exceeds the data correlation, $P\,(\mathrm{corr} > \mathrm{corr}_{\mathrm{data}})$.



(- -) Minnesota prior (1st), (-) Partial US prior (2nd), (-.-) Full US prior (3rd)

Figure 12: Euro area: Correlation with the 1-month EURIBOR and unit labor costs, $T = 100$. P-value: Probability that a simulated correlation exceeds the data correlation, $P\left(\text{corr} > \text{corr}_{\text{data}}\right)$.



(- -) Minnesota prior (1st), (-) Partial US prior (2nd), (-.-) Full US prior (3rd)