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Bank-Lending Standards, the Cost Channel and Inflation Dynamics

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Editorial

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 the authors 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.

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Bank-Lending Standards, the Cost Channel and Inflation Dynamics

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August 2010

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 Cointegration 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,b) shows that uncertainty about the strength of the cost channel influences the optimal setting of interest rates.

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 via impulse responses obtained from a vector error correction model (VECM) for a system of macroeconomic variables thought to capture the main dynamic features of an economy (Christiano et al., 2005) 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. The VECM is estimated using recent advances in Bayesian cointegration analysis (Koop et al., 2005, 2010). The Bayesian approach has some advantages given that lending standards for the euro area are only available since 2003, which does not allow to obtain precise estimates from observed data only. To circumvent the drawback, we will use the posterior inference about US data to design prior information for the euro area system to obtain first evidence on the interaction between lending standards, interest rates and GDP and inflation in particular.

We find evidence in favor of cost channel effects in the US as well as in the euro area, in line with the existing literature. Lending standards appear to be particularly relevant in the euro area, while their direct effect on financing costs appears to be rather limited in the US.

Only few papers 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 and complements them in the sense that we focus on inflation dynamics and the transmission of policy shocks to inflation including lending standards.

The remainder of the paper is structured as follows: Section 2 describes our data set. In Section 3 we present estimation results for the New Keynesian Phillips Curve augmented with lending standards. The importance of lending standards for the transmission mechanism is assessed in Section 4 Section 5 summarizes and concludes the paper.

2 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. 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 2009. Given that the sample for the euro area is very short, we will estimate the Phillips curve by pooling data of both regions in order to exploit cross-sectional data information. In the unrestricted VECM analysis, posterior moments of the estimated US system will be used as prior information for the euro area 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. To be consistent with the US series, we use the report about net tightening of loans to large enterprizes (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 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 and turned negative (-5.5) at the beginning of 2010. 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

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

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. In the first quarter of 2010, the net tightening of lending standards was, although quite low, still positive (4.4).

The correlation between the lending standards indicator and the interest rate is rather low for the US (see table 2, second column). Contemporaneously, they nearly are uncorrelated (-0.05), and when the Federal Funds rate is lagged by 1 quarter, the correlation is 0.06. The empirical results reported below suggest that the interaction between interest rates and standards has changed after 1999[3], which impacts on the transmission mechanism. Indeed, when the sample is split at the end of 1999, the correlation structure changes significantly. Until 1999, the contemporaneous and the lag correlations are 0.65 and 0.61, respectively. From 2000 onwards, these are -0.11 and 0.12, respectively. In the second sub-sample, the maximum correlation of 0.63 is reached when the interest rate is lagged by 6 periods (e.g. one and a half year). The corresponding correlation between the series for the euro area are positive. The contemporaneous correlation is 0.15, and when the 1 month EURIBOR rate is lagged by 1 quarter it increases to 0.39. The maximum correlation of 0.57 is reached at lag 3.

The bottom panel in figure 1 shows the series for the HP-detrended unit labor costs. For the US, the series corresponds to the unit labor costs of non-financial corporations, and is obtained from the Bureau of Labor Statistics. The series for the euro area represents total unit labor costs. We measure the inflation rate by the GDP-deflator, which, for the US, is retrieved from the IFS database.

As additional controls in the VECM, we will include the producer price index (PPI) and C&I loans for the US and loans to non-financial corporations for the euro area, which

[3] The Financial Services Modernization Act, enacted in November 1999, repealed part of the Glass-Steagall Act (1933), 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.

are depicted in the second panel of figure 1.

3 Lending Standards and the Phillips curve

As a preliminary analysis, we estimate the influence of lending standards on inflation dynamics using a Phillips curve relationship similar to the specifications suggested by Galí et al. (1999) and Galí et al. (2001) among others. Although the sample available for estimation is rather short, especially for the euro area, this analysis allows us to draw some preliminary conclusions concerning the role of lending standards. Interestingly, as we will show in the next section, the results of the VECM analysis largely confirm the findings we report in this section.

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 proportional to the wage bill and financing costs influence marginal production costs. Under the assumption that firms apply mark-up pricing, as it is standard in New Keynesian models, financing costs have a direct influence on the inflation rate.

In particular, we assume that the total cost C_t associated with financing of working capital is $C_t = \kappa_t R_t$, which includes interest costs, R_t as well as non-interest borrowing costs captured by $\kappa_t > 0$. The multiplicative specification captures the idea that financing costs are proportional to interest rate costs, which are magnified when financing conditions become restrictive, $\kappa_t > 1$, and are reduced when conditions are relaxed, $\kappa_t < 1$. We proxy these costs by lending standards, normalized to one when the net balance of banks tightening standards is zero. Then, for a given interest rate level, a rise in lending standards changes the financing costs that firms incur 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. These considerations lead us to the following empirical Phillips curve specification:

$$\pi_t = \gamma_R \kappa_t R_t + \gamma_s s_t + \gamma_f E_t \pi_{t+1} + \gamma_b \pi_{t-1}, \tag{1}$$

where π_t denotes the inflation rate, R_t is the interest rate, s_t are HP-filtered unit labor costs to proxy real marginal cost. The coefficient γ_R in the Phillips curve captures the effect of the financing costs, i.e. the combined effect of the policy rate and standards.

We report results obtained alternatively by two-stage least squares (TSLS, with Newey-West adjusted standards errors) and by the generalized method of moments (GMM). As instruments we use up to four lagged values of the interest rate, the unit labor costs, the inflation rate and of producer price index inflation.

Table 3 presents the results which we obtain with US data for the sample, 1990 - 1999, the sample 2000 - 2009, and for the whole sample. The GMM estimates for the whole sample reported in the last column, are based on two lags of all instruments in contrast to the four lags used in the subsample estimations. We choose to use only two lags in this case to obtain the most significant parameter estimates. Overall, the parameter estimates are robust across estimation methods. The J-statistic of overidentifying restrictions is not rejected for all specifications. To further validate the instruments, we either include or retrieve two lags of the interest rate and the unit labor costs. All J-statistics and also the incremental J-statistic do not reject the overidentifying restrictions.

We see from Table 3 that financing costs are significant in all three samples considered, while real unit labor costs are significant only in the sample 1990 to 1999. The influence of financing costs is somewhat weaker in the second sub-sample. Thus, the strength of the cost channel appears to have declined over time. The lagged inflation rate turns out to be robustly significant only in the first sub-sample and insignificant in the second. The expected inflation rate is significantly estimated by GMM in the second subsample and robustly so over the entire sample.

The BIC at the bottom of the table is reported as an indicative diagnostic measure. If the relationship between the data remains the same over the whole sample, the measure should not deteriorate a lot when we estimate the Phillips curve for the sub-samples. Note that the BIC shows a considerable improvement (around 29%), when the PC curve is estimated for the first sub-sample, while the measure deteriorates significantly (by 21% for the TSLS and by 45% for the GMM estimate) when we estimate it for the second sub-sample. Apparently, inflation dynamics are better captured in the first sample period than in the second one.

Next, we turn to the estimation of the Phillips curve for the euro area. Since the series are quite short, we proceed in the following way. We pool the time series of both regions and obtain an estimate by exploiting cross-sectional information. We use region-specific instruments and include two lags of the interest rate, the unit labor costs and the inflation rate, and four lags of the producer price inflation. We include fixed effects and report period-specific White adjusted standard errors.

Table 4 reports three estimations, one pooling the euro area data with the whole US sample (last column) and two pooling the euro area with, respectively, the first and the second sub-sample (the first two columns). All estimations pass the J-tests of overidentifying restrictions. When pooling with the first US sample, the effect of financing costs is estimated to be lower than the one reported for the respective US sample in Table 3. Unit labor costs, obtaining a positive point estimate, remain insignificant. Finally, future inflation is estimated to have a significant effect. When pooling euro area data with the second US subsample, unit labor costs enter with an implausibly small coefficient and the inflation terms become insignificant. Pooling with the whole US sample deteriorates the importance of the financing costs, the coefficient is only marginally significant, and renders both inflation terms significant. In light of these results, our preferred specification is the pool with the first subsample of US data. Using again the BIC as indicative measure, we observe that it remains the same relative to the estimated pool with the whole sample and that it deteriorates considerably (by 14%) if euro are data would be pooled with the second subsample of US data.

The estimated coefficient on the interest rate is somewhat smaller when we pool US with euro area data, it decreases from 0.03-0.04 to 0.02. This result suggests that cost effects are larger in the US.

What we are ultimately interested in, is the effect of non-interest rate financing costs on the inflation rate. To explore this issue, we estimate the Phillips curve without interacting the interest rate with standards. Table 5 shows the results. While most of the point estimates are of similar orders of magnitude as in Table 3, we see that the coefficients of R_t tend to be larger. Thus, the interest rate appears to exert stronger effects on inflation dynamics when lending standards are not explicitly considered. This result suggests that in empirical studies which do not control for the effect of lending standards, the interest rate is likely to capture the total effect of financing costs. This result may explain the large cost effects associated with interest rate changes reported in Ravenna and Walsh

(2006) and Chowdhury et al. (2006).

4 US and euro area transmission

4.1 Motivation and econometric approach

In the previous section we have seen that there is evidence for a cost channel effect in the US and the euro area. In particular for the US, there is also evidence that the effect decreased over time, partly due to the policy change (abolishment of restrictions on investment and retail banks to merge or combine financial services) that took place at the end of 1999. In this section, we follow Christiano et al. (2005) and estimate a vector error correction model (VECM) for a system of six variables to assess how lending standards enter into the transmission mechanism of monetary policy. We include real GDP, the GDP deflator, the producer price index (PPI), C&I loans (loans to non-financial corporations for the euro area), 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 each cointegrating vector. To this aim, we specify an uninformative prior on the cointegration space. 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 can be used to shape the prior specification of the euro area system, for which we have only very short

[4]Phillips (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.

time series.

Given that the method is based on very 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 N-vector of integrated variables Y_t, we write the vector error correction model (VECM)

$$\begin{aligned} y_t &= \Pi Y_{t-1} + \sum_{j=1}^{p} \Gamma_j y_{t-j} + \varepsilon_t \quad t = 1, \dots T \\ & \varepsilon_t \sim \text{i.i.d}N\left(0, \Sigma\right) \end{aligned} \tag{2}$$

where y_t denotes the vector of variables transformed to stationarity (usually growth rates or differences) and 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. We obtain

$$\Pi = \alpha\beta' \tag{3}$$

$$y_t = \alpha\beta' Y_{t-1} + \sum_{j=1}^{p} \Gamma_j y_{t-j} + \varepsilon_t \tag{4}$$

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 $\text{ec}_t \equiv \beta' Y_{t-1}$.

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

- The matrix

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

  is not identified. Any nonsingular transformation of α and β yields identical Π matrices. We can switch between specifications. 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 $\kappa = (\alpha'\alpha)^{-1/2} = (\beta^{*\prime}\beta^*)^{1/2}$.

- 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 (Strachan and van Dijk, 2004). Finally, Villani (2006) shows that working with the linear normalization may lead to counter-intuitive results on cointegration.

- Conditional on β, the nonlinear VECM (2) 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):

$$\pi(\beta) \propto |P_\tau|^{-r/2} |\beta' P_{1/\tau} \beta|^{-n/2} \tag{5}$$

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 $\tau = 0$, $\tau = 1$ leads to $P_\tau = I_N$, a non-informative prior on the Stiefel manifold.

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

$$vec(\alpha)|\beta, \Sigma, \tau, \nu \sim N\left(0, \nu \left(\beta' P_{1/\tau} \beta\right)^{-1} \otimes G\right) \tag{6}$$

where G may be chosen freely. In the application we set it to I_N. For the rest of the parameters, the dynamics $\Gamma_j,\ j = 1, \dots, p$ and the error covariance matrix Σ, we assume a Minnesota-type prior and an inverse Wishart distribution.[6] Combining these priors with the likelihood, we obtain a posterior normal distribution for α and $\Gamma_j,\ j = 1, \dots, p$, and an inverse Wishart for Σ.

[5] If 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 $H = H^g \left(H^{g\prime} H^g\right)^{-1/2}$.

[6] One might also work with non-informative priors.

The prior for β and α specified in (5) and (6), respectively, implies the following prior distribution for A and β^*:

$$\begin{aligned} vec(\beta^*)|A &\sim N\left(0, \left(A'G^{-1}A\right)^{-1} \otimes \nu P_\tau\right) && (7) \\ \pi(A) &\propto |G|^{-r/2}|A'G^{-1}A|^{-N/2} && (8) \end{aligned}$$

Note that $vec(\beta^*)|A \sim N\left(0, I_r \otimes \nu P_\tau\right)$ 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:

i. Sample $\alpha, \Gamma_j,\ j = 1, \ldots, p$ from $\pi\left(\alpha, \Gamma_j|\beta, \Sigma, y\right)$ and Σ from $\pi\left(\Sigma|\alpha, \beta, \Gamma_j, y\right)$
Transform α to $A = \alpha\left(\alpha'\alpha\right)^{-1/2}$

ii. Sample β^* from $\pi\left(\beta^*|A, \Gamma_j, \Sigma, y\right)$ and use it to transform to $\beta = \beta^*\left(\beta^{*\prime}\beta^*\right)^{-1/2}$ and $\alpha = A\left(\beta^{*\prime}\beta^*\right)^{1/2}$

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

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

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 $r = 1, \ldots, N$ 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 6 and table 7, bottom line). 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 90th percentile interval in the figures discussed below.

4.2 Results

The system for US data is estimated with one lag of endogenous variables and conditional on four cointegrating vectors, given the evidence for the first sample period 1990-1999 (see table 6).[7] Figure 2 depicts the impulse responses for shocks to C&I loans, the Federal funds rate and to lending standards.

A shock to the Federal funds rate decreases GDP significantly. The reaction of the GDP deflator is, although insignificantly so, negative. The PPI reacts significantly positively in the short run and, only after a rather long delay of 15 periods, the shock leads to a significant decrease in the PPI. The positive reaction in loans hints towards a procyclical behavior in the banking sector. The tendency of lending standards to show on impact an average negative reaction to the Fed funds rate shock is consistent with that view. Note however, that the reaction is insignificant over the whole forecast horizon. During this first sample period, a shock in lending standards has no direct effect on all other variables. In the light of these results, we may conclude that there is some evidence for a cost channel effect during the first sample period. This is reflected in the short-run positive effect of an interest rate shock on the PPI, which remains positive for about six to eight periods and turns significantly negative only after fifteen periods. This is also consistent with the Phillips curve estimate reported above, in which we find a positive combined effect of the interest rate and standards. However, standards did not have an additional direct effect on price variables. It appears that the relatively large positive correlation between the interest rate and standards reported in table 2 is due to the reaction of both variables to shocks occurring in the credit market.

Finally, in the first subsample, the impulse responses to a shock in loans show an intuitive pattern, may be except for the response in the producer price index. A shock in loans increases GDP in the short to medium run and the interest rate in the short run.

[7] The same cointegration rank is determined by the Johansen trace statistic. To save space, we do not report the results.

Lending standards have a tendency to increase. The response is insignificant, however.

For the second sub-period, 2000-2009, we find evidence for two cointegrating vectors (table 6). In figure 3 the impulse responses show some noteworthy differences to the first sub-period, which we again interpret as evidence for a significant effect of the policy change in 1999. The reaction of the GDP deflator and the PPI to a shock in the Fed funds rate is now basically insignificant over the forecast horizon, although in the GDP deflator, in the short run, we observe a small positive effect of the Fed funds rate shock and the reaction of the PPI turns negative after 10 periods. In this second sample, the negative reaction of standards to the Fed funds rate shock is more prolonged than in the first sub-period. Shocks in standards now have a direct (positive) effect on the GDP deflator and the PPI, while the effect on loans is negative, significant in the short run. The mean reaction of the Fed funds rate is also negative, in contrast to the first-period reaction. Although standards now have a direct effect on macro variables, there is less evidence for cost channel effects than in the first period, in particular due to the overall nearly insignificant effect of Fed funds rate shocks to GDP and the price indices, and due to the prolonged negative effect on standards. The opposing reactions of the Fed funds rate and standards to a shock in loans mirror the observed negative correlation between these two variables in the second sub-sample.

To analyze the euro area system, we can use the posterior inference on the US system, i.e. the moments characterizing the posterior distribution, to specify the euro area prior distribution. There are two possibilities. First to include information about the parameters characterizing the autoregressive dynamics and second, to additionally include information about the error covariance matrix. The first case is labeled partial US info and the second full US info. To decide on the posterior inference of which sub-sample to use, we again may use as criterion the BIC obtained under the various specifications. Table 7 reports the minimum of -0.68, obtained for the specification using the full US information obtained from the posterior inference for the first subsample. As for the US system, the Bayes factor indicates that four cointegration vectors are appropriate. The impulse responses obtained from this specification are depicted in figure 4.

An interest rate shock has a clear negative, significant effect on GDP. The effect on the deflator is negative in the medium- to long-term, becoming significant after eight to ten periods. After an initial rise, loans decrease significantly and standards rise. In particular this last result is consistent with micro-evidence in Maddaloni and Peydrò (2009), who report that individual banks adjust positively their lending standards after a policy rate increase. In addition, a shock to lending standards has a direct negative effect on loans in the short to medium run, while the other variables are basically unaffected. Overall, the reactions for the euro area are economically more intuitive than for the US. In light of these results, we may conclude also for the euro area that there is evidence for cost channel effects, although less pronounced than in the US, given that prices show a shorter delay in their reaction to policy rate changes. Despite the overall smaller cost channel effect in the euro area, standards react stronger to policy shocks. That is, standards appear to amplify the cost effects of interest rate changes to a larger extent than in the US.

The joint positive reaction of loans and interest rates to a shock in loans and the positive reaction of standards to interest rate shocks contribute to the observed positive correlation between the two variables.

5 Summary and Concluding Remarks

In this paper, we extend the literature on the cost channel transmission of monetary policy by analyzing the role of bank-lending standards as a proxy for non-interest costs associated with the financing of working capital.

We find that monetary policy exerts supply side effects via the cost channel in the US as well as in the euro area, although the cost channel is quantitatively stronger in the US. Turning to the role of lending standards, we find that they do not exert strong additional cost pressure in the US, while they influence financing costs in the euro area.

Our results also indicate 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.

In addition, our paper contributes to the small but growing literature that analyzes the effects of variations in bank lending standards for the business cycle and the monetary transmission mechanism.

Finally, it should be noted that the statistical significance of our results is not always overwhelming which may reflect the limited variation in the rather short series, especially for the euro area. Analyzing longer series, when available, therefore represents an interesting direction for future research.

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A Tables

Table 1: Data sources

	Euro Area[a)]	United States[b)]
Sample	2003:1-2009:4	1990:2-2009:4
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
GDP	euro area 12 (13 from 2009 onwards)	IFS: GDP real and nominal
Deflator	euro area 12, seasonally adjusted	IFS: Deflator
PPI	Monthly commodity price index, import-weighted	IFS: 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)] If not otherwise stated, data from the Federal Reserve Board's website

Table 2: Correlation

Sample US	1990-2009	1990-1999	2000-2009
Corr(Stand,int.rate)	-0.05	0.65	-0.11
Corr(Stand,int.rate(-1))	0.06	0.61	0.11
Max Corr	0.36	0.65	0.63
	(lag 6)	(lag 0)	(lag 6)
Sample EA			2003-2009
Corr(Stand,int.rate)			0.15
Corr(Stand,int.rate(-1))			0.39
Max Corr			0.57
			(lag 3)

Table 3: US: Structural parameter estimates of Phillips curve. Two stage least squares with Newey-West adjusted standard errors and covariances. Standard errors are in squared brackets, P-values in parentheses.

	1990-1999		2000-2009		1990-2009	
	TSLS	GMM	TSLS	GMM	TSLS	GMM
$\kappa_t R_t$	0.03	0.04	0.03	0.02	0.01	0.01
	[0.01]	[0.01]	[0.02]	[0.01]	[0.01]	[0.005])
	(0.02)	(0.00)	(0.09)	(0.03)	(0.34)	(0.02)
s_t	9.28	11.78	-12.54	-1.42	0.66	0.79
	[4.25]	[1.62]	[6.85]	[2.71]	[2.22]	[1.28]
	(0.04)	(0.00)	(0.08)	(0.60)	(0.77)	(0.54)
π_{t-1}	0.29	0.32	-0.19	0.01	0.08	0.17
	[0.14]	[0.07]	[0.19]	[0.09]	[0.18]	[0.09]
	(0.05)	(0.00)	(0.33)	(0.89)	(0.64)	(0.06)
π_{t+1}	0.11	-0.07	0.36	1.01	0.74	0.81
	[0.19]	[0.09]	[0.32]	[0.11]	[0.21]	[0.15]
	(0.56)	(0.45)	(0.27)	(0.00)	(0.00)	(0.00)
Instruments	$J-\chi^2$-test of overidentifying restrictions					
R_{t-j}						
s_{t-l}	$j_{\max}=2$	$j_{\max}=4$	$j_{\max}=4$	$j_{\max}=4$	$j_{\max}=2$	$j_{\max}=2$
π_{t-j}	$l_{\max}=2$	$l_{\max}=4$	$l_{\max}=2$ [a)]	$l_{\max}=4$	$l_{\max}=2$	$l_{\max}=2$
Δppi_{t-m}						
$m_{\max}=4$	χ^2_6 (0.42)	χ^2_{12} (0.69)	χ^2_{10} (0.17)	χ^2_{12} (0.96)	χ^2_6 (0.14)	χ^2_6 (0.68)
	$j_{\max}=4$	$j_{\max}=2$	$j_{\max}=4$	$j_{\max}=2$	$j_{\max}=4$	$j_{\max}=4$
	$l_{\max}=4$	$l_{\max}=2$	$l_{\max}=4$	$l_{\max}=2$	$l_{\max}=4$	$l_{\max}=4$
	χ^2_{12} (0.40)	χ^2_6 (0.86)	χ^2_{12} (0.17)	χ^2_6 (0.95)	χ^2_{12} (0.35)	χ^2_{12} (0.95)
	$\chi^2_{12}-\chi^2_6$	$\chi^2_{12}-\chi^2_6$	$\chi^2_{12}-\chi^2_{10}$	$\chi^2_{12}-\chi^2_6$	$\chi^2_{12}-\chi^2_6$	$\chi^2_{12}-\chi^2_6$
	(0.37)	(0.79)	(0.30)	(0.77)	(0.74)	(0.97)
	Diagnostics					
SSR	0.56	0.61	1.90	3.53	3.56	3.85
BIC	-1.59	-1.54	-0.97	-0.66	-1.23	-1.19

[a)] with $j_{\max}=2,\ l_{\max}=2$, the $J-\chi^2_6$ (0.85) does not reject, while $\chi^2_{10}-\chi^2_6$ (0.02) and $\chi^2_{12}-\chi^2_6$ (0.03) reject the instruments.

Table 4: EA-US: Pooled structural parameter estimates of Phillips curve. Two stage least squares estimates with region-specific instruments. Period-specific White adjusted standard errors and covariances. Sample period for the euro area: 2003-2009. Standard errors are in squared brackets, P-values in parentheses.

Sample period US	1990-1999	2000-2009	1990-2009
$\kappa_t R_t$	0.02	0.04	0.01
	[0.01]	[0.01]	[0.01]
	(0.01)	(0.00)	(0.05)
s_t	1.49	-15.79	-0.49
	[1.52]	[5.22]	[1.39]
	(0.33)	(0.00)	(0.72)
π_{t-1}	0.10	-0.04	0.08
	[0.09])	[0.12]	[0.04]
	(0.28)	(0.27)	(0.00)
π_{t+1}	0.65	-0.21	0.66
	[0.01]	[0.17]	[0.10]
	(0.00)	(0.23)	(0.00)
Instruments	$J-\chi^2$-test of overidentifying restrictions		
(region-specific)			
$R_{t-j},\ s_{t-j}$	$j_{\max}=2$	$j_{\max}=2$	$j_{\max}=2$
$\pi_{t-j},\ \Delta\text{ppi}_{t-m}$			
$m_{\max}=4$	χ^2_{16} (0.33)	χ^2_{16} (0.50)	χ^2_{16} (0.27)
	$j_{\max}=4$	$j_{\max}=4$	$j_{\max}=4$
	χ^2_{28} (0.41)	χ^2_{28} (0.36)	χ^2_{28} (0.45)
	$\chi^2_{28}-\chi^2_{16}$ (0.51)	$\chi^2_{28}-\chi^2_{16}$ (0.26)	$\chi^2_{28}-\chi^2_{16}$ (0.67)
	Diagnostics		
SSR	1.79	2.75	4.56
BIC	-1.07	-0.80	-1.07

Table 5: The role of standards. Phillips curve estimates without standards. Two stage least squares estimates with region-specific instruments. Newey-West and period-specific White adjusted standard errors and covariances. Sample period for the euro area: 2003-2009. Standard errors are in squared brackets, P-values in parentheses.

	US Phillips curve				US-EA pool
Sample period US	1990-1999		2000-2009		1990-1999
	TSLS	GMM	TSLS	GMM	TSLS
R_t	0.07	0.04	0.04	0.03	0.04
	[0.03]	[0.01]	[0.02]	[0.01]	[0.004]
	(0.03)	(0.02)	(0.05)	(0.01)	(0.00)
s_t	11.10	14.17	-10.53	-0.61	2.69
	[5.18]	[2.24]	[6.34]	[2.58]	[1.72]
	(0.04)	(0.00)	(0.11)	(0.82)	(0.12)
π_{t-1}	0.21	0.18	-0.20	-0.02	0.04
	[0.13]	[0.08]	[0.18]	[0.09]	[0.05]
	(0.13)	(0.03)	(0.26)	(0.85)	(0.43)
π_{t+1}	0.28	0.22	0.38	1.02	0.69
	[0.23]	[0.09]	[0.31]	[0.11]	[0.07]
	(0.23)	(0.02)	(0.24)	(0.00)	(0.00)
Instruments	$J-\chi^2$-test of overidentifying restrictions				
(region-specific)					
$R_{t-j},\ s_{t-j}$	$j_{\max}=2$	$j_{\max}=2$	$j_{\max}=2$	$j_{\max}=2$	$j_{\max}=2$
$\pi_{t-j},\ \Delta\text{ppi}_{t-m}$					
$m_{\max}=4$	χ^2_6 (0.27)	χ^2_{12} (0.89)	χ^2_{10} (0.27)	χ^2_{12} (0.95)	χ^2_{16} (0.23)
	Diagnostics				
SSR	0.71	0.99	1.88	3.54	3.14
BIC	-1.49	-1.32	-0.98	-0.66	-1.33

Table 6: 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.

US sample period	1990-1999		2000-2009	
	$\log BF_{0r}$	$Prob(r)$	BF_{0r}	$Prob(r)$
$r = 1$	-26.64	0.00	-47.66	0.00
$r = 2$	-48.79	0.00	-78.77	1.00
$r = 3$	-66.82	0.00	-50.45	0.00
$r = 4$	-72.66	1.00	-72.65	0.00

Table 7: BIC of estimated VARs for the euro area using US posterior moments as euro area prior moments. Partial info: US posterior moments for VAR parameters are used as prior moments for euro area system. Full info: in addition, US posterior scale of covariance matrix is used as prior scale for error covariance of the euro area system. The last line reports the probability of cointegration rank r, $Prob(r)$.

US sample period		1990-1999		2000-2009	
	no info	partial US info	full US info	partial US info	full US info
$\det(\Sigma)$	-1.93	-1.98	-10.46	-2.16	-9.00
BIC	7.85	7.79	-0.68	7.62	0.78
$r,\ Prob(r)$	3, 0.74	2, 1.00	4, 1.00	2, 0.88	2, 0.85

B Figures

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

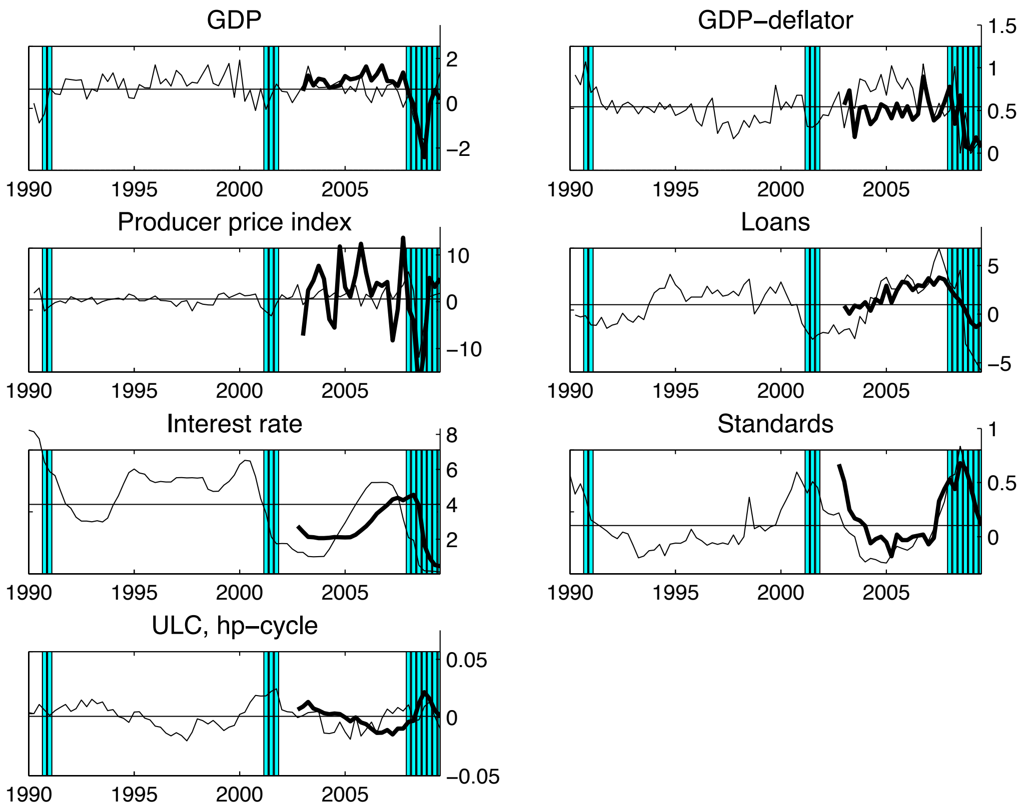


Figure 2: US, sample period 1990-1999: Impulse responses with 90th percentile interval

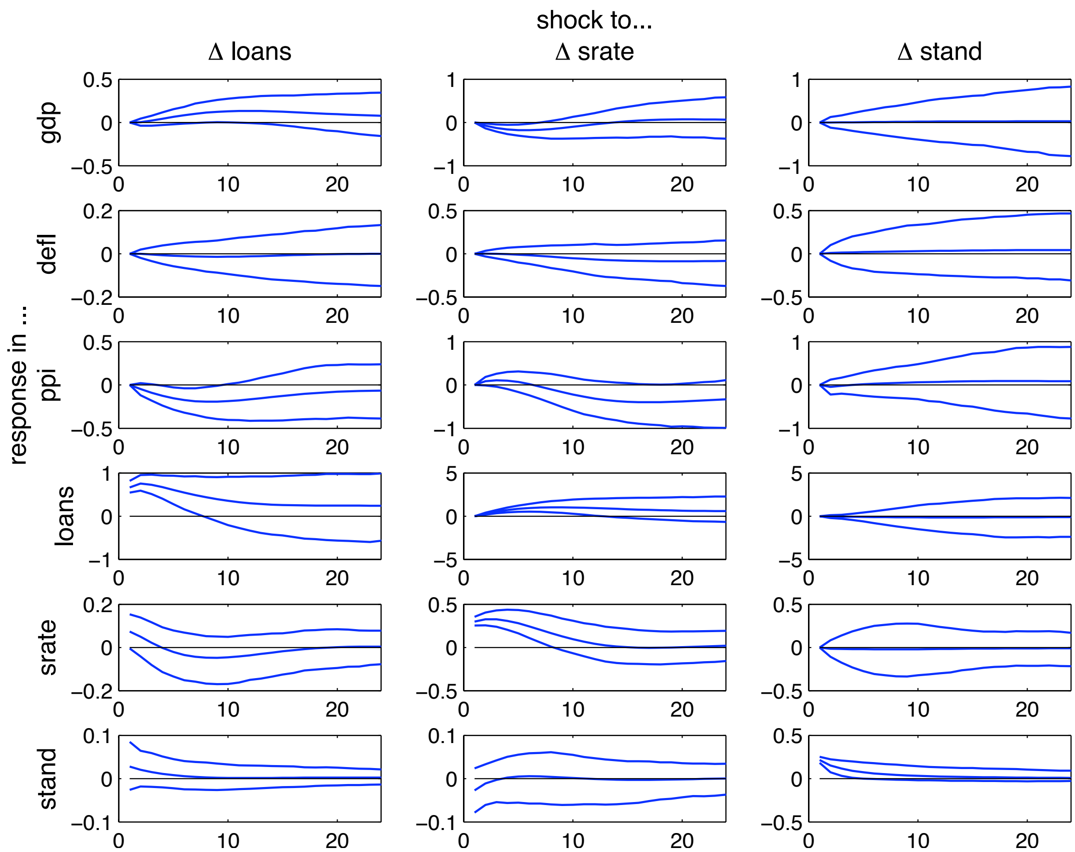


Figure 3: US, sample period 2000-2009: Impulse responses with 90th percentile interval

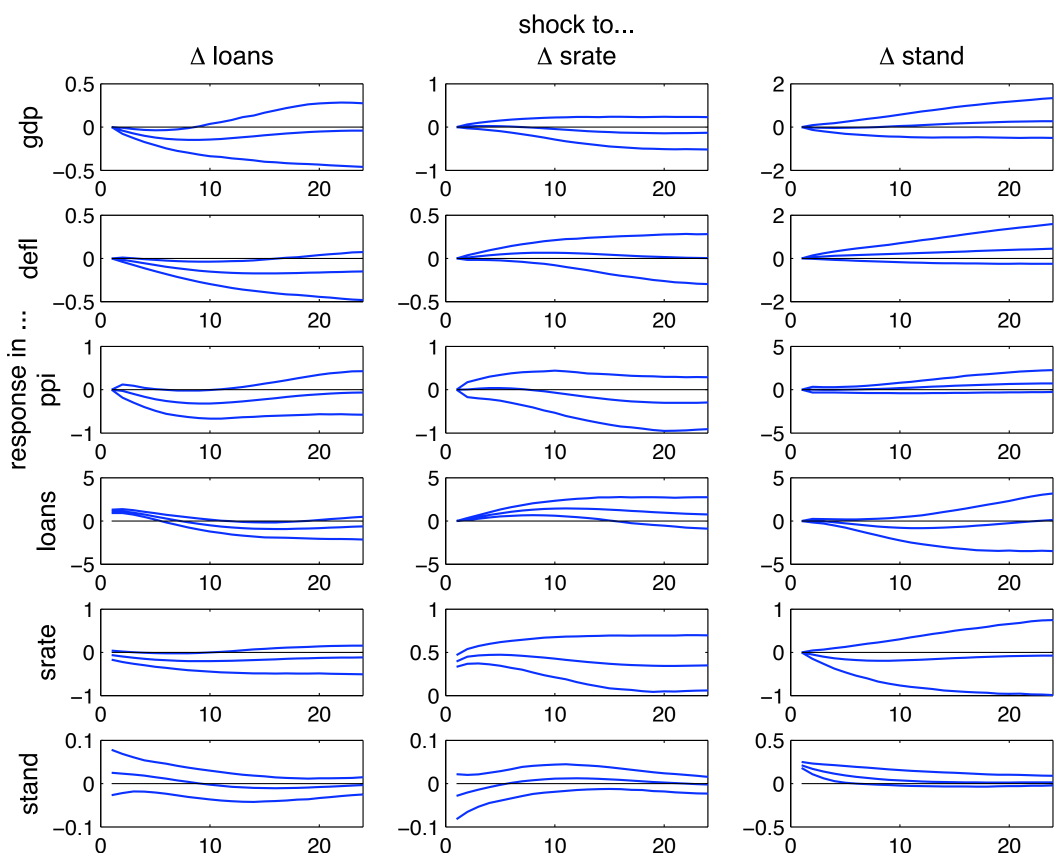
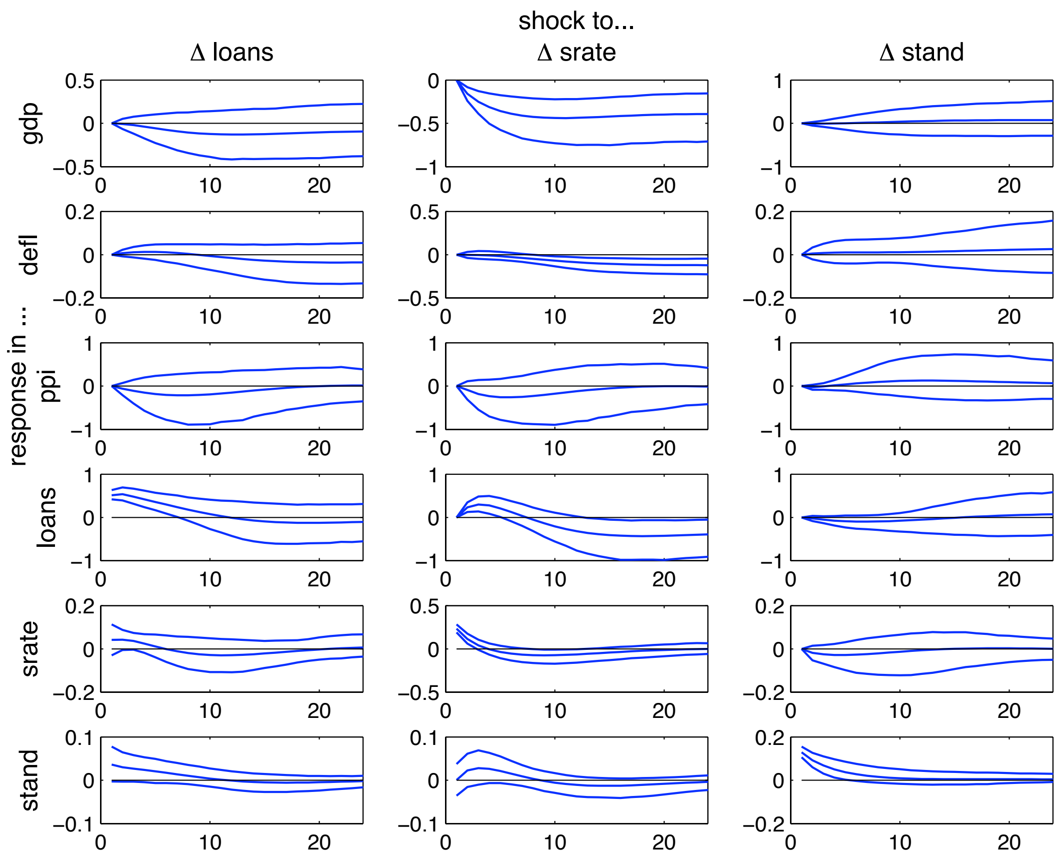


Figure 4: Euro area, with US posterior (sample period 1990-1999) as prior: Impulse responses with 90th percentile interval, full US prior



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