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by Stefano Neri and Andrea Nobili

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THE TRANSMISSION OF MONETARY POLICY SHOCKS FROM THE US TO THE EURO AREA

by Stefano Neri* and Andrea Nobili*

Abstract

This paper studies the transmission of monetary policy shocks from the US to the euro area using a two-country structural VAR with no exogeneity assumption. The analysis reveals the following results. First, in response to an unexpected increase in the Federal funds rate, the euro immediately depreciates with respect to the dollar and then appreciates in line with the prediction of the uncovered interest parity condition. Second, there is evidence of a temporary positive spillover to euro-area output in the short run, while a negative effect emerges in the medium run. Third, the contribution of the trade balance channel to the transmission of monetary shocks is negligible. Finally, the degree of pass-through of the exchange rate changes onto euro-area consumer prices is incomplete and small in the short run, while it is close to zero in the medium run.

JEL classification: C32; E52; F42.

Keywords: Structural VAR; Monetary Policy; International transmission.

Contents

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1 Introduction[1]

Over the last ten years there have been a number of important developments, both theoretical and empirical, in open-economy macroeconomics. Since the publication of Obsteld and Rogoff's (1995) Redux model, commonly recognised as the contribution that launched the New Open Economy Macroeconomics (NOEM) research agenda, open-economy dynamic general equilibrium models that incorporate imperfect competition and nominal rigidities have been developed to account for the international transmission of shocks and the dynamics of exchange rates. Lane (2001) and Sarno (2001) provide two excellent surveys. These models are characterised by a higher standard of analytical rigour deriving from fully specified microfoundations than the traditional Mundell-Fleming-Dornbusch (MFD) models of the 1960s and 1970s. Moreover, they allow the welfare effects of different monetary and exchange rate policies to be explicitly evaluated.

After Obsteld and Rogoff's Redux model, the theoretical literature gradually introduced additional features, including sticky prices (see Chari, Kehoe and McGrattan, 2002, among others) and wages (Hau, 2000 and Obstfeld and Rogoff, 2000), pricing-to-market (Betts and Devereux, 2000), non-traded goods (Hau, 2000, and Obstfeld and Rogoff, 2000). The development of Bayesian techniques for general equilibrium models and the availability of fast computers have also allowed researchers to estimate small open-economy and two-country models (Bergin, 2003, Lubik and Schorfheide, 2005, de Walque, Smets and Wouters, 2005, and Rabanal and Tuesta, 2005). Recently, open-economy models have been developed at central banks such as the Riksbank, the Federal Reserve, the European Central Bank and the Bank of Canada.

The existing empirical evidence on the international transmission of shocks has been based on structural vector autoregressions and focused on the US and the G-7 countries. Betts and Devereux (2001) and Kim (2001) found evidence of remarkable spillovers from the former to the latter. Recently, Lubik and Schorfheide (2005), Rabanal and Tuesta (2005) and de Walque, Smets and Wouters (2005) have estimated two-country general equilibrium models of the US and the euro-area using Bayesian techniques. Overall, they show that monetary shocks played only a small role in shaping the dynamics of the real exchange rate between the dollar and the

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euro and that the spillover effects of these shocks have been relatively small.

The objective of this paper is to provide empirical VAR-based evidence on the international transmission of monetary policy shocks between two large economies, such as the US and the euro-area. As far as identification is concerned, we draw from the NOEM literature in defining the set of restrictions that are necessary to identify the shocks. In setting up the VAR we impose the uncovered interest parity (UIP) condition on the exchange rate and the short-term interest rates and we specify the equation for the bilateral trade on the basis of the pricing behaviour for US and euro area exports (Local Currency Pricing, LCP, or Producer Currency Pricing, PCP).

We set up a two-country VAR model that includes both US and euro-area variables, as well as some linking the two economies. Since the data suggest that we can relax the assumption of small open-economy for the euro area as well, in contrast to previous analyses, we do not impose any exogeneity assumption on the VAR model and we assume that the transmission of US monetary policy shocks can work simultaneously through movements in the trade balance, the exchange rate and the short-term interest rates. From a methodological point of view, our approach is an improvement with respect to previous work such as Kim (2001) and Betts and Devereux (2001). The former uses a "marginal method" approach in which the foreign variables of interest are added one at a time to a well-defined closed-economy model for the US. It also assumes that the rest of the world is exogenous to the US economy. Results deriving from this approach may suffer from two related problems. On the one hand an omitted variable problem might arise to the extent that there is significant simultaneity among the foreign variables. On the other hand, results do not also take into account possible feedback effects from the G-7 countries to the US. Betts and Devereux (2001) included both US and foreign variables in their VAR, but they did not consider some relevant variables for the identification of US monetary policy shocks, such as a monetary aggregate and a commodity price index.

Our empirical evidence may be useful in constructing dynamic optimising models for policy analysis. For example, negative evidence on the role of trade may suggest researchers to look for other channels through which shocks are transmitted in a international environment. The behaviour of the nominal exchange rate in response to a monetary policy shock might help us understand whether it is consistent with the predictions of the uncovered interest rate parity. Finally, evidence on the response of the foreign country short-term interest rate to domestic monetary shocks may help us choose the terms to be included in interest rate rules.

Our main findings are the following. After a monetary contraction in the US the euro immediately depreciates and then appreciates with respect to the dollar, consistently with the uncovered interest parity condition. At the same time, we find a strong temporary positive effect on euro-area output in the short run, but a larger and more persistent negative effect in the medium run. Finally, the pass-through of exchange rate changes onto euro-area prices is incomplete and low in the short run and not significantly different from zero in the long run.

The remainder of the paper is organised as follows. Section 2 describes the methodology used for conducting correct inference in our overidentifed two-country VAR. Section 3 presents the two-country VAR and gives details regarding the specification, the identification of monetary policy shocks and the estimation. Section 4 assess the impulse responses and provides some insights into the transmission mechanism. The conclusions are drawn in Section 5.

2 Methodological issues

In this section we briefly describe the structural vector autoregression that will be used in the empirical analysis. More details on the computation of impulse responses are given in Appendix B. We assume that the economy is described by the following system of structural equations

$$A_0 Y_t = A(L) Y_{t-1} + e_t \tag{1}$$

where $A(L)$ is a matrix polynomial in the lag operator L, Y_t is a vector of variables of dimension n and e_t is the vector of structural disturbances, which are assumed to be independent and identically distributed with covariance matrix equal to the identity matrix.

The coefficients of the structural equations in (1) can be recovered from the estimated reduced form by imposing enough restrictions, that is making assumptions on the matrix A_0 that links contemporaneously the variables in Y_t. Under this assumption and if the disturbances u_t are Gaussian, the posterior distribution of the A_0 matrix can be obtained following Sims and Zha (1999). Choosing a flat prior for both $B(L)$ and A_0 and integrating over $B(L)$ give the posterior density of A_0

$$p(A_0) \propto |A_0|^{T-\nu} exp\left[-\frac{1}{2} tr(A_0' S(\widehat{B}) A_0) \right] \tag{2}$$

where ν is the degrees of freedom correction, which is equal to the number of estimated coefficients in each equation.

For overidentified VAR models, expression (2) does not represent the density function of any standard random variable and therefore there is no simple method for generating a random sample for the A_0 coefficients and computing error bands for impulse responses. Several authors have used techniques such as importance sampling. However, Waggoner and Zha (1997, 2003a) pointed out its inefficiency when the shape of the posterior density is highly non-Gaussian and suggested alternative posterior simulators such as Gibbs sampling or Metropolis. In the light of this result, we used 1,500,000 draws from the random walk version of the Metropolis algorithm to obtain a sample from the posterior distribution of the A_0 matrix following Sims and Zha (1999). The first 20 per cent of the draws were discarded as burn-in sample.

In all the non-recursive identifications we implement the likelihood-preserving normalisation suggested by Waggoner and Zha (2003b). This method ensures that the effects of exogenous shocks are in the same direction, thus always maintaining a coherent economic interpretation. Normalization amounts to a rule that determines the sign of all coefficients in each structural equation and it is one of the difficult issues in the context of simultaneous equation models. The choice of normalisation rule is irrelevant for the point estimates of impulse responses but it is crucial for inferential conclusions around the maximum likelihood estimates.

3 A two-country VAR for the US and the euro area

3.1 Speci cation

In this section we first set up a closed economy VAR for the US that allows the identification of monetary policy shocks and then move on to estimating a two-country model to evaluate the effects of policy shocks on the main macroeconomic variables of the euro area and the exchange rate between the dollar and the euro. The estimation of the closed economy is done in order to ensure that the impulse responses that we obtained are in line with those available in the literature. It also helps in assessing the contribution of foreign variables in the transmission of monetary shocks to the US variables.

Following previous studies, such as Leeper, Sims and Zha (1996) and Kim (1999), we include in the vector y_t the Federal funds rate (i_{us}), expressed on an annual basis, the log of the M2 monetary aggregate (m_{us}) in nominal terms, the log of real GDP (y_{us}), the log of the implicit GDP deflator (p_{us}) and the log of the world commodity price index (cp). For a description of the data see Appendix A. Three lags are used in the specification of the model. In the choice of the appropriate number of lags we opted for a parsimonious VAR specification, but at the same time characterised by estimated residuals with good white-noise properties.[2]

3.2 Identi cation

The identification of monetary policy shocks is based on Kim (1999) and it allows the money stock and the short-term interest rate to jointly transmit monetary policy shocks, as suggested

[2] Table C1 in Appendix C shows that the multivariate version of the Lagrange multiplier test suggests that just a lag order of one is sufficient to get uncorrelated VAR residuals. Nevertheless, we opted for a specification with three lags, as it is the optimal lag order chosen for the two-country model, therefore allowing us to make a better comparison of the results.

by Leeper and Roush (2003).[3] Impulse responses to a monetary policy shock that increases the Federal funds rate on impact by 50 basis points are reported in Figure 1. Solid lines represent the median estimates while the dashed lines denote .68 and .90 probability intervals derived from the marginal posterior distributions of the impulse horizon at each step. All figures report impulse responses expressed as deviations from the baseline. Overall, the identification scheme does not imply any particularly puzzling response of the variables and delivers impulse responses that are similar in shape and persistence to those of previous studies, such as Christiano, Eichenbaum and Evans (1999), Leeper, Sims and Zha (1996) and Bernanke and Mihov (1998).

Moving on to the two-country model, the following variables are added: the log of the bilateral exchange rate of the euro vis-á-vis the dollar (e), the euro-area three-month nominal interest rate (i_{emu}), the euro-area real GDP (y_{emu}), the euro-area implicit GDP deflator (p_{emu}) and the nominal trade balance between the two areas (tb). The bilateral exchange rate is expressed as units of US dollars per euro. Therefore, a positive change in this variable measures an appreciation of the euro. The bilateral trade balance is expressed in nominal terms since the appropriate deflator is not available, and it is multiplied by 100 and divided by the level of euro-area nominal GDP. The two-country VAR is estimated in levels with three lags. The sample period used in the estimation is still 1982:3 to 2005:2. Standard tests presented in Appendix C suggested no evidence of serial correlation and heteroskedasticity in the residuals.

The VAR is based on the assumption that the euro area is considered an open economy, as large as the United States. Therefore, we do not impose any block exogeneity restriction in the model, meaning that each correlation between US and euro-area variables is assumed to be bidirectional, going from the US to the euro area and vice versa. In order to check this hypothesis we run a battery of pair-wise Granger causality tests between euro-area and US variables, as well as the multivariate block exogeneity test, in order to examine whether the group of variables used in the benchmark model for the US is block-exogenous with respect to the set comprising all the other variables. The results provide evidence that the null hypothesis that lagged values of euro-area and common variables have zero coefficients in the US block is strongly rejected at any conventional confidence level. Consequently, the strategy of setting up a VAR model for the euro area in which the previously estimated US structural monetary policy shocks enter exogenously in the specification, appears not to be optimal.

The following matrix summarises our proposed identification scheme

[3] The authors pointed out that, when monetary policy shocks generate significant co-movements in money and the short-term interest rate, the price puzzle disappears and the effects on real economic activity and inflation are larger.

$$A_0 = \begin{bmatrix} a_{1,1} & a_{1,2} & 0 & 0 & a_{1,5} & 0 & 0 & 0 & 0 & 0 \\ a_{2,1} & a_{2,2} & a_{2,3} & a_{2,4} & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & a_{3,3} & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & a_{4,3} & a_{4,4} & 0 & 0 & 0 & 0 & 0 & 0 \\ a_{5,1} & a_{5,2} & a_{5,3} & a_{5,4} & a_{5,5} & a_{5,6} & a_{5,7} & a_{5,8} & a_{5,9} & a_{5,10} \\ a_{6,1} & 0 & 0 & 0 & 0 & a_{6,6} & 0 & a_{6,8} & 0 & 0 \\ 0 & 0 & a_{7,3} & a_{7,4} & 0 & a_{7,6} & a_{7,7} & 0 & a_{7,9} & a_{7,10} \\ 0 & 0 & 0 & 0 & a_{8,5} & a_{8,6} & 0 & a_{8,8} & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & a_{9,9} & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & a_{10,9} & a_{10,10} \end{bmatrix} \begin{bmatrix} i_{us} \\ m_{us} \\ y_{us} \\ p_{us} \\ cp \\ e \\ tb \\ i_{emu} \\ y_{emu} \\ p_{emu} \end{bmatrix}$$

which is exclusively based on short-run restrictions.

The above matrix is composed of three blocks: the US economy, the group of euro-area variables and a set of variables linking the two areas, comprising the bilateral exchange rate, the world commodity price index and the bilateral trade balance. The first block of restrictions, represented by the first four rows, allows the identification of US monetary policy shocks and basically reflects the same assumptions used in the closed-economy VAR. All euro-area variables are assumed not to affect the level of US output and prices contemporaneously, and vice versa, reflecting the a priori belief in the existence of lags in the transmission of shocks between the two areas.

By symmetry, the euro-area real sector is modelled like the US one. Real GDP responds to the price level and all the other variables with at least one quarter lag. The equation of the euro-area short-term interest rate (i_{emu}) is considered the reaction function of the "European monetary authority", which sets the short-term interest rate in response to the current value of the commodity price index and the exchange rate, but not the values of the other variables. These two variables control for the systematic response of monetary policy to inflationary pressures stemming from foreign markets. Contrary to the assumptions for the US block, in which the monetary aggregate M2 appears, we do not introduce a measure of money in the equation of the euro area short-term interest rate. This choice can be justified on the basis of the follwing considerations. First, when considering the monetary policy of the Bundesbank, the empirical evidence suggest that developments in the monetary aggregate M3 did not affect in a relevant way the policy decisions. Bernanke and Mihov (1997) conclude that"*holding constant the current forecast of inflation, German monetary policy responds very little to changes in forecasted money growth*". The evidence in Clarida, Galí and Gertler (1998) also suggest that money growth does not enter significantly into a forward-looking monetary policy rule estimated for the Bundebank over the 1979-1993 period.[4] On the other hand, they find that M2 growth enters significantly in a reaction function for the Federal Reserve estimated over the post 82 period.

It is important to underline that although no single monetary authority existed in the euro area until January 1999, the Deutsche Bundesbank was *de facto* setting the stance of monetary policy for the whole area. Our assumption that the euro-area short-term interest rate reacts contemporaneously to the exchange rate is based on the following two considerations. First, Clarida, Galí and Gertler (1998) showed that the exchange rate between the Deutschmark and the dollar enters significantly in an estimated reaction function of the Bundesbank. Second, some countries participating today in the European Monetary Union (EMU), such as Italy, Spain and France, were setting their short-term interest rates in order to stabilise their own exchange rate with respect to the Deutschmark. Again, the evidence in Clarida, Galí and Gertler (1998) supports this statement, at least as far as concerns Italy and France.

We model the two monetary policies as being independently set. Clarida, Galí and Gertler (1998) found some evidence of a response of the Bundesbank interbank lending rate to the Federal funds rate. However, the effect is negligible since the introduction of the US interest rate does not affect significantly the estimation of the benchmark reaction function of the Bundesbank. In addition, we do not include the exchange rate in the information set of the Federal Reserve following Clarida, Galí and Gertler (1998).

In the financial block of the A_0 matrix, the world commodity price equation (cp) is assumed to respond to all the variables, in line with the assumptions in the closed-economy VAR. With respect to the exchange rate, we assume that it depends only on the two short-term interest rates, consistently with some kind of uncovered interest parity. However, we do not take seriously the restrictions imposed by the UIP condition, namely that the coefficients on the interest rates should be equal to minus and plus one. Our assumption is somehow different with respect to that in Kim and Roubini (2000), where all variables included in the system have contemporaneous effects on the exchange rate. However, it would not be possible to allow commodity prices and the exchange rate to respond contemporaneously to all variables at the same time for identification reasons.[5] Finally, we assume that the bilateral trade balance responds to current values of both US and euro-area output and prices and to the bilateral exchange rate. Indeed, if

[4] These results notwithstanding, in Section 4.5 we present evidence obtained with a two-country model including the M3 money stock in the block of euro-area variables and identifying money demand shocks. Results are very similar to those obtained with the benchmark model.

[5] As a robustness check, we also considered the same two-country model in which we allowed the exchange rate to respond contemporaneously to all variables included in the system, while the commodity price index was considered completely exogenous to both the US and the euro area. Results are virtually similar and discussed in more detail in Section 4.2.

US firms price their exports to the euro area according to the local currency (LCP, Local Currency Pricing) and euro area exporting firms price their goods in euros (PCP, Producer Currency Pricing), then the trade balance must also be a function of the contemporaneous value of the bilateral exchange rate. In only one case out of the possible four (two pricing assumptions, PCP and LCP, in the two countries, the US and the euro area) the trade balance should not depend contemporaneously on the exchange rate. The robustness of the results to this alternative identification is tested in Section 4.3.

The world commodity price index and the exchange rate are contemporaneously excluded from this equation. These variables are likely to affect the trade balance only with lags, reflecting previously invoiced trade contracts and advance production planning.

3.3 Estimation

The reduced form VAR is estimated consistently in levels relying on the results in Sims, Stock and Watson (1990). Data are quarterly and the sample period goes from 1982:3 to 2005:2, in order to rule out the non-borrowed regime the Federal Reserve was implementing between 1979 and part of 1982 under Governor Volcker. Moreover, Sims and Zha (2006) find that changes in the variances of structural shocks, between 1979 and 1982, are the major source of instability in a VAR including the main US macroeconomic variables.

The second column of Table 1 reports the summary statistics of the posterior distributions for the coefficients of the Federal fund rate and M2 money equations. The data correctly identify a supply and demand for money, although the estimated elasticities are somehow different from the corresponding values in the closed-economy model. Concerning our proposed identification, it is important to underline that the likelihood ratio (LR) test for overidentified VARs suggests that our short-run restrictions cannot be rejected at 5 per cent confidence level (the statistic is equal to 25.64 and the corresponding *p*-value is 0.78). We view this as a remarkable result since the model is highly overidentified (the number of over-identifying restrictions is equal to 21). The LR test confirms the correctness of the assumption that the real blocks of both the US and the euro area are contemporaneously independent and that the exchange rate depends only on short-term interest rates.

To ensure that our estimated monetary policy shocks do not capture common world demand shocks, negative technology shocks in the US or oil shocks, we estimated a linear regression as in Francis and Ramey (2005) using as dependent variable the time series of monetary shocks and as regressors the following variables: a measure of world demand shocks, obtained as innovations to an AR(1) process for world exports, technology shocks identified using long run restrictions as in Christiano, Eichenbaum and Vigfusson (2003) and the variables in Dedola and

Neri (2006), the measure of oil supply shocks computed by Kilian (2006), Hamilton's (2003) measure of oil shocks and the Hoover and Perez (1994) oil dummies. The policy shocks are computed using the mode of the posterior distribution of the A_0 matrix and the least squares estimates of the residuals. Results for the F-tests that the current and first three lags of each of the shocks are jointly equal to zero suggest that our identified policy shocks are not likely to be contaminated by other demand and supply shocks (see Table 2). Similar results are obtained when using the 16^{th} and 84^{th} percentiles of the posterior distribution of A_0.

4 Results

In this section we study and discuss the set of impulse responses to identified monetary policy shocks in the US.

4.1 The effects of monetary policy shocks on the US variables

The dynamic effects of a 50 basis point increase in the Federal funds rate on the US variables are shown in Figure 2. The usual error bands are also displayed.

Overall, in response to a contractionary monetary policy shock, the US macroeconomic variables follow patterns remarkably similar in shape and timing to those of the closed-economy VAR, although the effects seem to be somehow larger (see Figure 1). More precisely, the liquidity effect is stronger, as the money M2 declines on impact by about 0.8 per cent from the baseline. The fall of the price level is now significant and persistent over the whole horizon, reaching a maximum decline of more than 0.5 per cent after four years.

The forecast error variance decomposition of US variables based on the posterior mean of the A_0 matrix and the OLS estimates of the reduced form coefficients shows that the contribution of monetary policy shocks increases with the forecast horizon. On average, over the entire horizon, these shocks explain around 11 and 13 per cent of the variance of, respectively, real GDP and the GDP deflator. When the mode of the posterior distribution of A_0 is used, this contribution is only slightly larger.

4.2 The effects on the bilateral exchange rate

The upper-right-hand corner of Figure 2 shows the impulse response of the nominal bilateral exchange rate of the euro with respect to the dollar. After the shock the dollar appreciates on impact by slightly more than 2.0 per cent. The appreciation of the dollar is rather persistent with the maximum response occurring two quarters after the shock. In the following quarters, the dollar depreciates slowly and returns to its initial level after three years.

The pattern of the response of the bilateral exchange rate appears to be consistent with the uncovered interest parity condition (UIPC). Conditionally on the monetary policy shock, the UIPC suggests that a positive innovation in the US interest rates relative to euro-area ones should be associated with a persistent depreciation over time of the dollar after the initial appreciation. In addition, the ex-post excess return, defined as the difference in the return between investing in one-quarter US assets and in one-quarter euro-area assets, should be zero. On the basis of our two-country model, we find that the response of the ex-post excess return is never significant over the whole impulse horizon (see Figure 8B). These results are consistent with those of previous studies regarding the effects of a US monetary policy shock on the bilateral exchange rate of the dollar with respect to the currencies of non-US G-7 countries (e.g. Eichenbaum and Evans, 1995; Kim and Roubini, 2000). Using an estimated general equilibrium model, Lubik and Schorfheide (2005) and de Walque, Smets and Wouters (2005) find that a monetary contraction in the US leads to an appreciation of the exchange rate with respect to the euro.

Interestingly, we obtained very similar results to the benchmark model for all the impulse responses when imposing the restriction that the coefficients on the two interest rates were the same but of opposite sign, as predicted by the UIPC. Impulse responses obtained with this alternative scheme are presented in Figure 3. However, in the benchmark model we preferred not to restrict these coefficients and leave them free to differ because the value of the posterior density at the mode was higher.

The same pattern arises looking at the response of the real exchange rate, which is constructed using the responses of the US and euro-area price levels (expressed in the same currency), and is reported in Figure 8B. This result mainly reflects the high correlation between these two variables, which is usually explained by the existence of price rigidities.[6]

Finally, we assess the role of US monetary policy shocks in generating fluctuations in both the nominal and the real exchange rate. To this end, we used the posterior mean of the identification matrix A_0 and the OLS estimate of the reduced form of the VAR. Similar results are obtained using the posterior mode of A_0. Over a four-year horizon, monetary shocks on average account for around 6 and 11 per cent of, respectively, the variance of the forecast error of the nominal and the real exchange rate. Similar results are obtained in the context of an estimated DSGE model by Rabanal and Tuesta (2005), Lubik and Schorfheide (2005) and de Walque, Smets and Wouters (2005).

[6] We obtain very similar results if we derive the response of the real exchange rate using the consumption deflator or a measure of unit labour costs for the two economies instead of the GDP deflator.

As our results may depend on our identifying assumptions, we assessed the robustness of the response of both the bilateral exchange rate and the conditional excess return in a two-country model in which the exchange rate, being a forward-looking asset price, is allowed to respond contemporaneously to all variables included in the VAR model, while the commodity price index is assumed to be completely exogenous to the rest of the system (see Figure 4). This identification is similar to that of Kim and Roubini (2000). The overidentifying restriction test suggests that these assumptions are also accepted by data (the LR statistic is equal to 18.53 and the corresponding p-value to 0.82). All impulse responses are remarkably similar to the benchmark model, including that of the excess return (see Figure 8B).

4.3 The transmission to the euro-area variables

The response of euro-area real GDP is shown in Figure 2. In the short-run we find evidence of a positive spillover, as it increases by around 0.4 per cent after five quarters. Then, there is a significant negative effect starting seven quarters after the shock and remaining persistent and significant at horizons up to four years, when real output falls below the baseline by about 0.6 per cent.

In the context of the recent NOEM literature, theoretical models deliver different predictions regarding the cross-country transmission of monetary policy shocks. The impact on the foreign country depends crucially on the magnitude of the expenditure-switching and income-absorption effects. The relative importance of these channels depends on the type of pricing adopted by firms at the international level, on the value of structural parameters such as the elasticity of substitution between home and foreign goods, between goods within each country and finally on the degree of home bias in consumption.

On the one hand, the *Redux* model by Obstfeld and Rogoff (1995) suggested that a domestic monetary contraction reduces output of both home and foreign country by the same size as a result of bond markets being perfectly integrated and both countries consuming the same basket of goods. On the other hand, Chari, Kehoe and McGrattan (2002) and Corsetti and Pesenti (2001) showed that the signs of the spillover effects strongly depend on parameters such as the elasticity of substitution between home and foreign goods and the elasticity of substitution of goods within each country.

To assess the relevance of the trade channel in the transmission of monetary shocks to the euro area, we study the response of bilateral trade, which is reported in Figure 2. After one quarter it worsens by a small amount (0.05 per cent of euro-area nominal GDP) and then returns to zero. This pattern can be rationalised with the traditional *J*-Curve effect, according to which there should be an initial worsening due to the effect of the exchange rate depreciation on import prices and, consequently, on the nominal value of imports, while traded quantities do not change because of previously invoiced trade contracts. After some lags (usually one or two quarters) the adjustment of quantities to the new prices would lead to an improvement in the trade balance. To the extent that quantities do not respond on impact to the shock and given the fact that the trade balance does not respond on impact, it may be inferred that euro-area export (to the US) and import (from the US) prices move little in response to movements in the exchange rate. One possible rationalisation of this pattern is that euro-area exports to the US are priced in dollars (local curreny pricing, LCP) and US exports to the euro area are also priced in dollars (producer currency pricing, PCP), so that when the value of the dollar changes, prices do not move. Interestingly, Gopinath and Rigobon (2006) find that contrary to standard modelling assumptions there is producer currency pricing in US exports and local currency pricing in US imports. We also experimented with an indentification scheme in which the trade balance does not depend on the exchange rate (see Figure 5). This case corresponds to the combination of pricing behavour according to which euro area exporting firms price in euro (PCP) and US exporting firms price also in euro (LCP) (see the discussion in Section 3 above). The results are remarkably similar to those reported in Figure 2.

An additional result is that monetary policy shocks in the US explain only around 2 per cent of the bilateral trade balance fluctuations in the short run and basically zero in the longer run. These results suggest that trade with the US does not contribute to the transmission of monetary shocks to the euro area.

The positive effect on euro-area real GDP may also reflect a depreciation of the euro with respect to the currencies of other trade partners and consequently larger expenditure-switching effects. To provide some evidence that this channel may be at work, we drop the bilateral exchange rate and include the nominal effective exchange rate of the euro.[7] This variable shows a pattern remarkably similar to that obtained with the bilateral exchange rate with the dollar but the size of the response is smaller.

Precisely, the euro depreciates on impact by about 1.0 per cent and remains below the baseline up to two years after the shock, when the depreciation is slightly more than 2.0 per cent. Nevertheless, if we include in the model the overall extra-area trade balance (as a fraction of real GDP) we find the contribution of external demand to be 0.1 per cent one year after the shock and not significant in the following periods.

[7] Results are available upon request. The nominal effective exchange rate is based on a narrow group of euro-area trade partners, namely United States, United Kingdom, Japan, Switzerland, Sweden, Denmark, Norway, Canada, Mexico, Australia, New Zeland, Korea, Hong Kong and Singapore. This group actually covers about 70 per cent of overall euro-area trade. For further details see the September 2004 issue of the ECB Monthly Bulletin and Buldorini, Makrydakis and Thimann (2002).

Overall it can be argued that the trade balance channel does not help to explain the response of euro-area real GDP either in the short run or in the long run. This result is in line with that of Kim (2001), who also finds that changes in the trade balance play a minor role in the transmission of US monetary shocks to the rest of the G-7 countries. Therefore, other channels must be at work.

In particular, we consider the possibility that the endogenous response of the foreign central bank may contribute to the transmission of external shocks. This feature is present in the analysis of Lubik and Schorfheide (2005). In their model, the monetary authorities of the US and the euro area are allowed to respond contemporaneously to the exchange rate. The estimated model suggests that in response to a monetary contraction in the US, domestic inflation and output decline, the dollar appreciates and, finally, output and inflation increase in the euro area. The monetary authority responds endogenously to these latter developments by increasing the short-term interest rate. In line with the results in Lubik and Schorfheide (2005), we find that output, prices and the short-term interest rate all increase in the euro area. However, the hike in nominal interest rates (less than 20 basis points) is not enough to increase the real short-term interest rate, which actually decreases for the first two quarters (0.2 per cent; see Figure 8B).[8]

In our opinion, this decline in the real rate may explain the initial expansion in output, which as we have shown above cannot be rationalised in terms of the movements in the trade balance with the US and the other trading partners. Similarly, the decline of output in the medium run may also depend on the increase in the real short-term interest rate a few quarters after the shock.

Finally, another channel that might be at work is the increase in disposable income following the decrease in the price of raw materials. Indeed, the magnitude of the response of these international prices is rather large, around 8.0 per cent on impact, and the decline is persistent. This suggest the possibility that the consequent wealth effect might reinforce the positive spillover on real output in the euro area through an increase in aggregate demand.[9]

4.4 Assessing the pass-through onto euro-area consumer prices

In this section we assess the degree of pass-through of the exchange rate onto euro-area prices. Figure 2 shows that after the US monetary contraction, the euro-area GDP deflator increases

[8] The response of the real short-term interest rate, r_t, is obtained by subtracting the quarterly differences in the response of the price level, $logP_t$, (expressed in annual terms) from the response of the nominal short-term interest rate, R_t: $r_t = R_t - 4 \times (logP_{t+1} - logP_t)$.

[9] This view is supported by the positive effect on real private consumption and investment (see Figure 8A). The impulse responses are obtained by replacing real GDP in the euro area with each of the two variables separately.

by around 0.2 per cent in the second year after the shock. This pattern implies a pass-through coefficient (defined as the response of the price level divided by the impact response of the nominal exchange rate) of about 0.1 in the medium run and zero in the long run.

It might be argued that the consumption deflator could be more useful in assessing the effects of exchange rate changes on domestic inflation, as import prices are included in the basket and changes in commodity prices and other imported goods prices could have a more rapid impact on the index than on the GDP deflator, which reflects the price of goods produced domestically. Nevertheless, this variable has also some potential drawbacks, as it includes prices of goods imported solely for consumption purposes, while it excludes prices of important tradeable goods, such as capital and intermediate goods, which also enter the production processes as inputs. In this sense the GDP deflator might better reflect considerations of international cost and price competitiveness, as it has the advantage of shifting the focus from the consumption side to the production side of the economy.

These considerations notwithstanding, in order to give a more exhaustive assessment of the pass-through of exchange rate fluctuations onto euro-area prices, we repeat the analysis substituting both the US and euro-area GDP deflator, respectively, with US personal consumption expenditure and euro-area private consumption deflator, and leaving our identifying assumptions unchanged with respect to the benchmark two-country. The response over time of this alternative measure of prices is similar in timing and persistence.[10]

Our results seem to support some recent empirical work on the euro area. For example, Campa and Minguez (2002) used both cross-country and cross-industry data and found that the average aggregate pass-through onto the total costs of intermediate production after a 10 per cent euro depreciation is 1.2 for the euro area, which declines below 0.5 per cent when the consumer price index is considered.[11] Faruqee (2004) finds that in response to a 10 per cent depreciation of the nominal effective exchange rate of the euro, consumer prices increase by only 0.2 after one and a half years.

[10] We also explored a version of the two-country model that included the US consumer price index (CPI) and the euro-area harmonised index of consumer prices (HICP), finding similar results. They are available upon request.

[11] The authors pointed out that significant differences exist across euro-area countries, mainly due to the distinct degrees of openness to non-euro area imports rather than to the heterogeneity in the structure of imports. Broadly speaking, there is evidence of incomplete and lower pass-through for larger euro-area countries, such as Germany and France.

4.5 Adding the euro-area M3 monetary aggregate

In this sub-section we assess the robustness of our results regarding the transmission mechanism by exploring an alternative two-country model including the M3 money stock in the block of the euro-area variables. The identifying assumptions for the euro-area money demand equation are identical to those used for the US, namely that M3 responds simultaneously to the euro-area real GDP, GDP deflator and nominal short-term interest rate. In addition, we also allowed the euro-area short-term interest rate to respond contemporaneously to the money aggregate in order to rule out the endogenous response of monetary policy to pure money demand shocks. Note that our proposed identification scheme, which now includes 26 restrictions, is still not rejected by the data at the conventional 5 per cent confidence level. The likelihood ratio test statistic is equal to 33.76 and the corresponding p-value to 0.86.

Results are very similar to the benchmark model and are presented in Figure 6. After a positive monetary policy shock the euro-area M3 decreases on impact and remains below the baseline over the whole horizon. Regarding other euro-area variables, we obtain slightly more persistent responses of both the nominal and the real short-term interest rate in the medium run, still suggesting evidence of an endogenous response of European monetary authorities after the US monetary contraction.

4.6 An alternative assumption on monetary policy

The identification of monetary policy shocks has been the subject of debate in academic research, as both recursive and non-recursive identification schemes have been proposed. Differences in the two approaches mainly reflect the assumptions made on the information set available to the Federal Reserve at the time at which the monetary policy instrument is set.

We assess the robustness of our results for the transmission mechanism using an alternative identification scheme based on the idea that the central bank observes current output and prices but not money when setting the short-term interest rate (see for example Christiano, Eichenbaum and Evans, 1999, among others). The use of a scheme that does not allow for simultaneity in the US money market may lead to smaller effects of monetary policy shocks on US real economic activity and to a possibly puzzling response of the US price level (Leeper and Roush, 2003). For consistency, the euro-area short-term interest rate is also assumed to respond contemporaneously to both output and prices. All the remaining identifying restrictions are those of the benchmark VAR. This alternative scheme is still strongly accepted by data, as implied by the overidentifying restriction test. The likelihood ratio test statistic is 21.65 and the corresponding p-value is equal to 0.80.

The responses of the variables are reported in Figure 7. Changing the identifying assumptions in the US block does not affect significantly the qualitative effects of monetary policy shocks on US variables although the magnitude of the effects are smaller, consistently with the results in Leeper and Rush (2003). In particular, the responses of the money M2 and real GDP become sligthly smaller.

The responses of the euro-area variables are also not qualitatively affected. The marginal distribution of the responses of prices seems to be slightly shifted down in the recursive identification. The nominal short-term interest rate increases on impact, while the corresponding real rate rises with a lag of three quarters. Some differences emerge in the response of the commodity price index, which in the case of a recursive scheme declines by a smaller amount compared with the benchmark identification. The error bands of the impact response of the exchange rate are now tighter and the median response is larger. A 50 basis points increase in the Federal funds rate still depreciates the euro by around 2.0 per cent on impact. The response of the exchange rate still shows a dynamic pattern which is consistent with the uncovered interest rate parity condition.

5 Concluding remarks

This paper has studied the transmission of monetary policy shocks from the US to the euro-area using a two-country structural VAR model. Contrary to previous analyses with structural VAR models, no block exogeneity has been imposed on the model.

The analysis has shown that a monetary contraction in the US leads to a temporary positive effect on euro-area output in the short run. A more persistent and negative effect emerges in the medium run. The euro depreciates on impact with respect to the dollar and then appreciates in line with the prediction of the uncovered interest parity condition. The degree of pass-through of the exchange rate onto consumer prices is incomplete and very low in both the short run and the medium run. In assessing the transmission mechanism of monetary shocks we find that the trade balance plays a negligible role, suggesting that researchers should look for other channels through which shocks are transmitted in an international environment. The euro-area short-term interest rate responds endogenously to commodity prices and the nominal exchange rate. Our analysis shows that real interest rates and commodity prices seem to play a more important role.

Tables and figures

Table 1

VAR contemporaneous coefficients

	US closed-economy VAR	two-country VAR
Federal funds rate equation		
a_{11}	188.2	120.2
	(83.7, 289.5)	(32.8, 211.8)
a_{12}	7.8	-75.9
	(-115.3, 131.2)	(-143.6, -8.7)
a_{15}	-14.5	-18.5
	(-21.8, -6.8)	(-24.8, -12.4)
M2 equation		
a_{21}	112.8	332.5
	(-87.4, 294.8)	(274.0, 390.3)
a_{22}	164.8	203.7
	(101.5, 218.5)	(162.5, 246.4)
a_{23}	-4.9	-32.6
	(-42.2, 32.7)	(-69.3, 4.5)
a_{24}	25.7	-11.2
	(-39.6, 50.9)	(-94.5, 72.4)

$Notes$: For each coefficient of the A_0 matrix the table reports the mean and the 0.68 probability interval (in brackets) of the corresponding marginal posterior distribution. Each distribution is obtained from a sample of 1,500,000 draws from the Metropolis algorithm.

Table 2

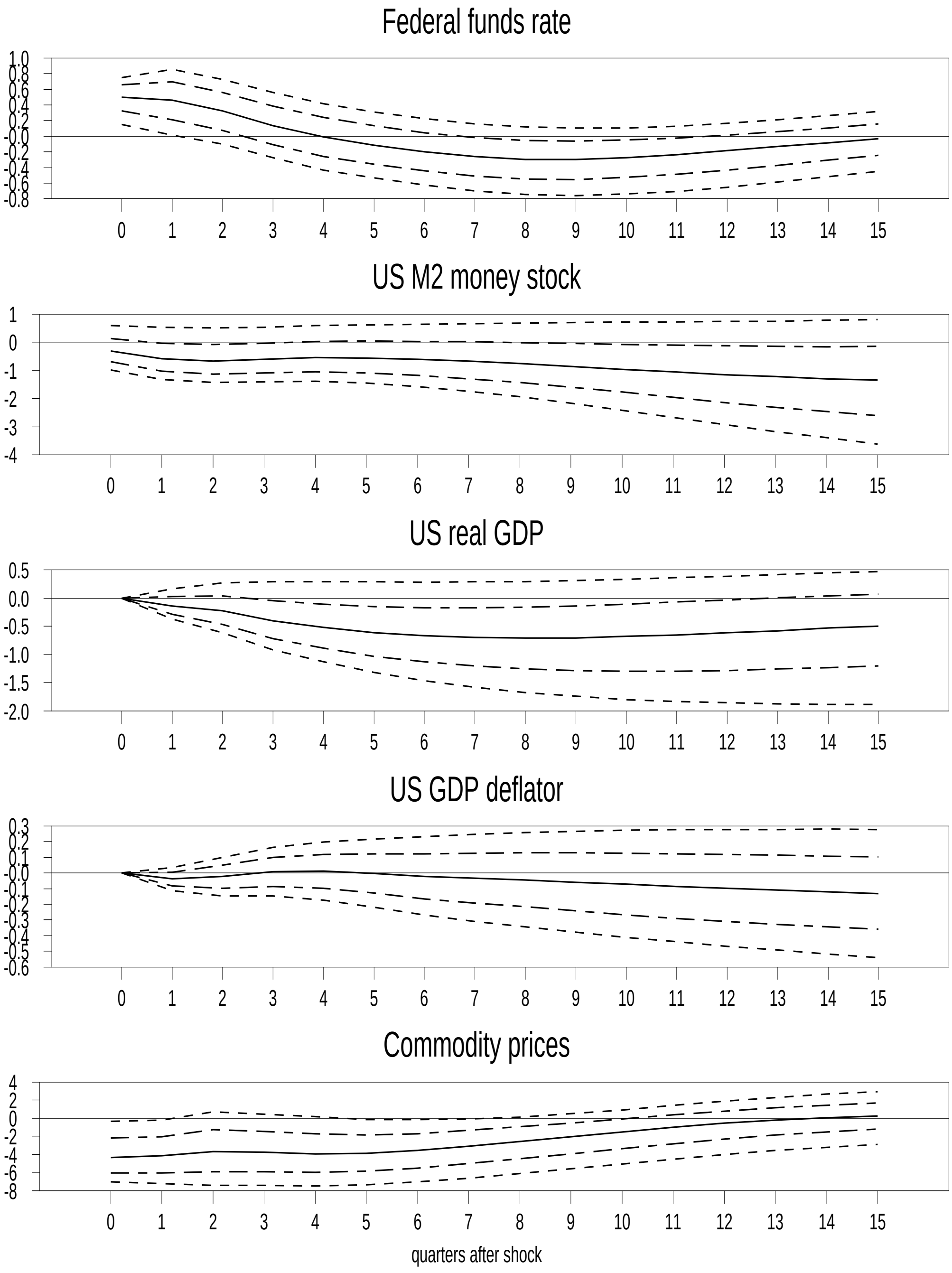
Exogeneity tests for identified monetary policy shocks

Structural global shocks	F-test	p-value
Hamilton's net oil price increase	0.43	0.79
Hoover and Perez oil dummy	1.14	0.35
Kilian's oil supply shock	0.25	0.91
Technology shock	0.43	0.79
World aggregate demand shock	0.35	0.84
All shocks	0.62	0.88

Notes: Monetary policy shocks are computed using the mode of the posterior distribution for the A_0 matrix and the least squares estimates of the reduced form residuals, which is obtained from a sample of 1,500,000 draws from the Metropolis algorithm. The F-test is based on a regression of the monetary policy shocks on a constant and current and three quarterly lags of the variable in question. The null hypothesis is that all of the coefficients on the variable in question are jointly equal to zero.

Figure 1

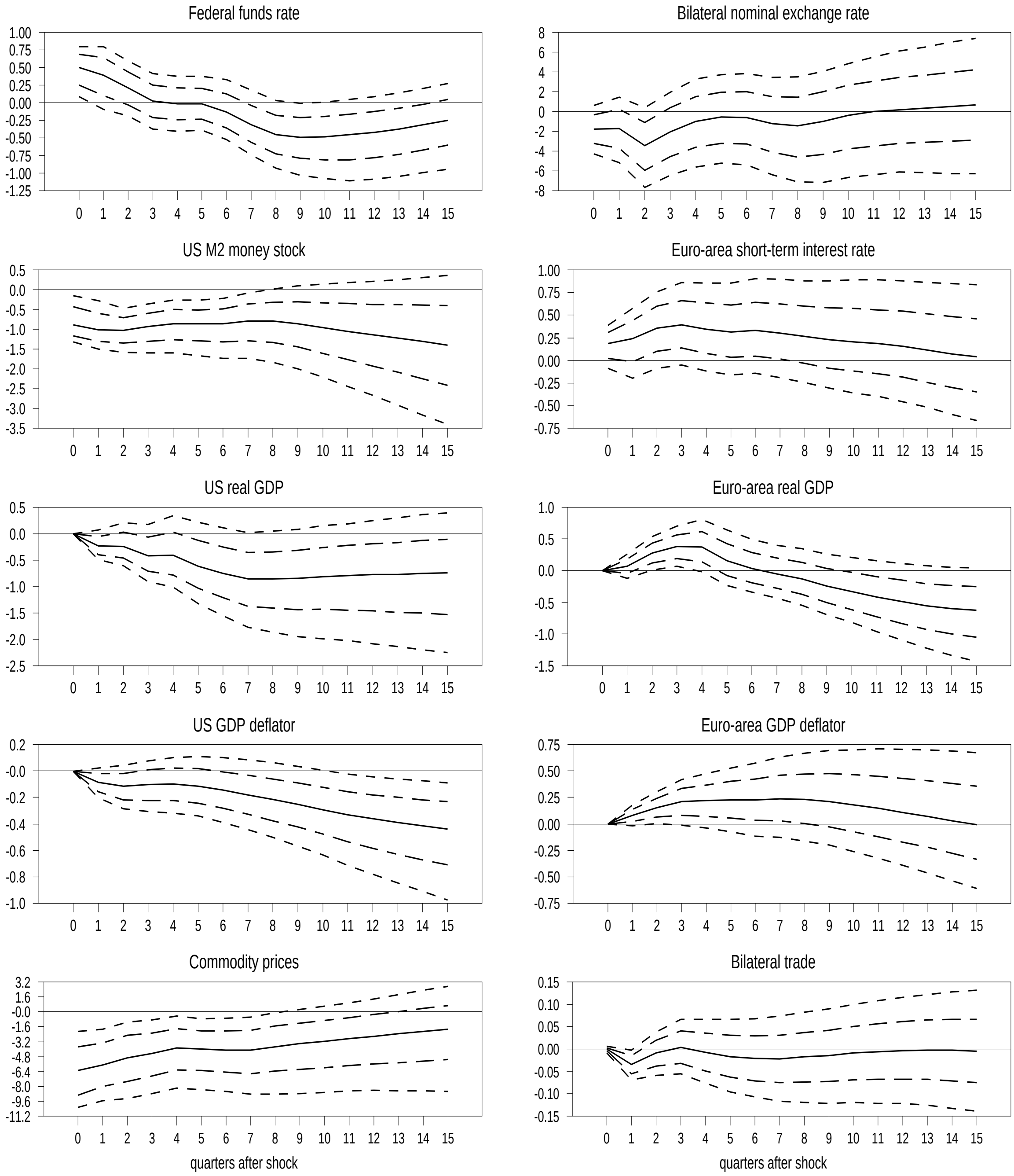
Closed-economy VAR for the US:
responses to a positive monetary policy shock



— median — – — 16, 84 percentiles – – 5, 95 percentiles

Figure 2

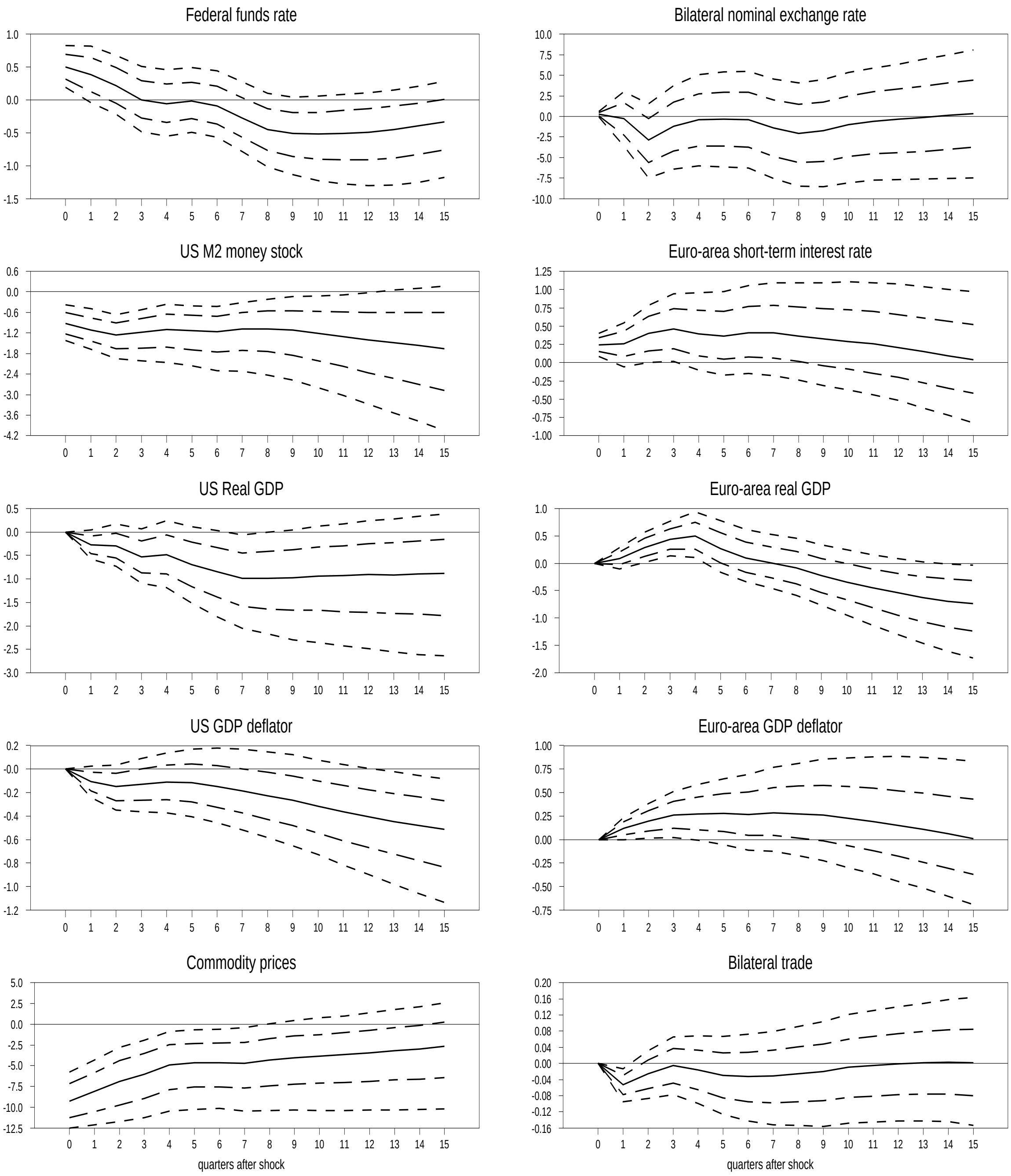
Two-country VAR: responses to a positive monetary policy shock in the US



— median —– –— 16, 84 percentiles – – 5, 95 percentiles

Figure 3

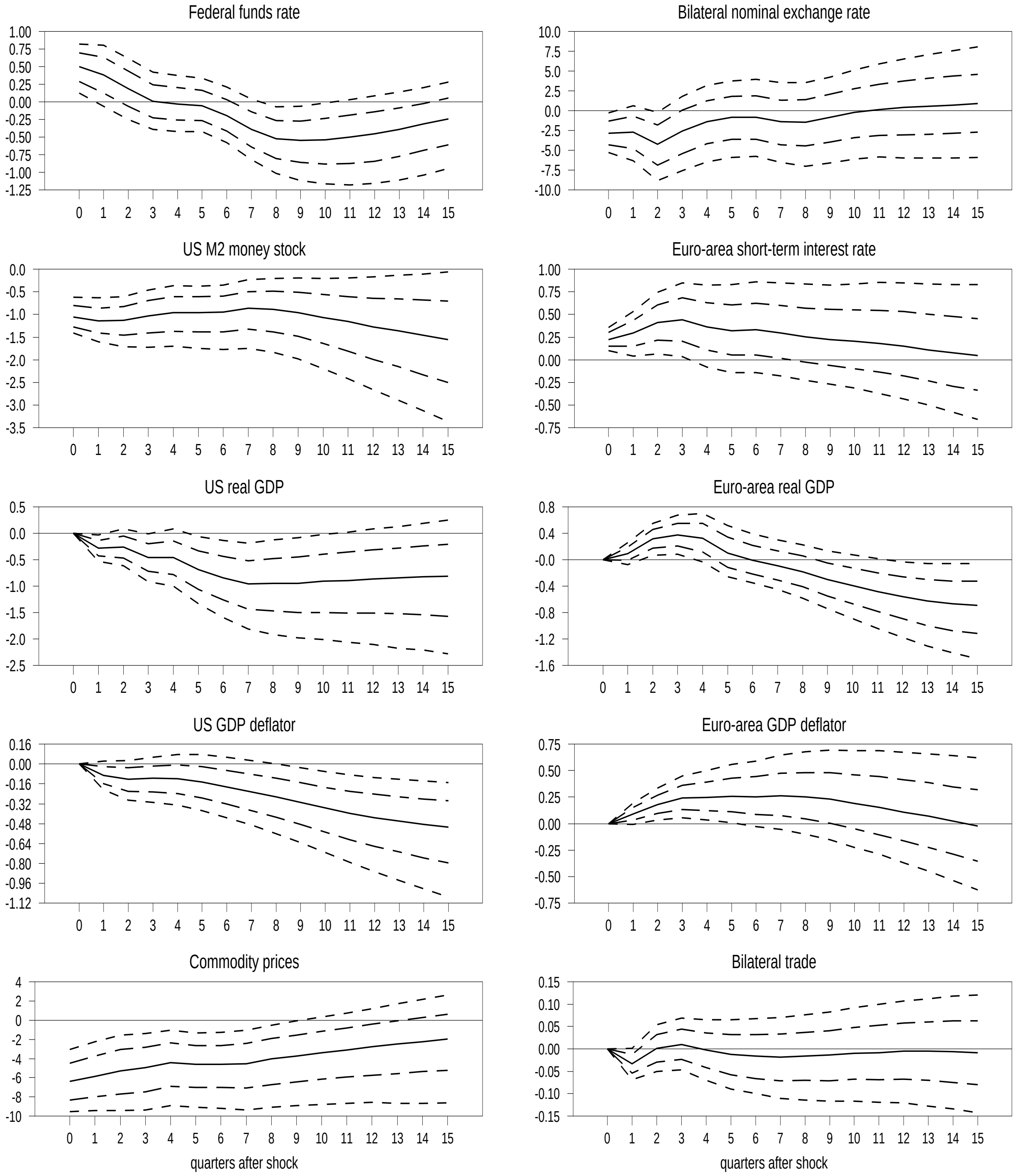
Two-country VAR with uncovered interest parity condition:
responses to a positive monetary policy shock in the US



— median ——— 16, 84 percentiles – – 5, 95 percentiles

Figure 4

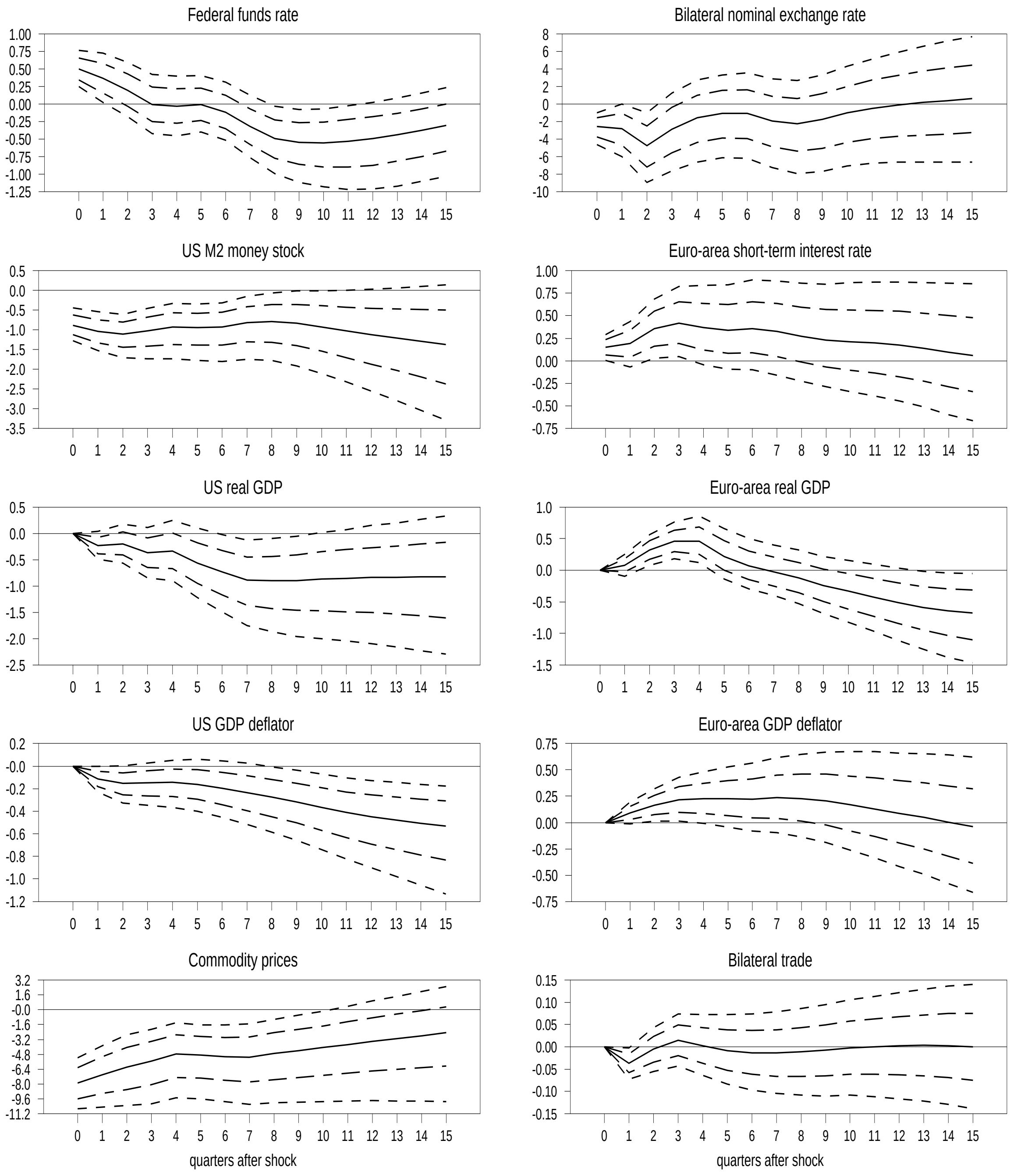
Two-country model with alternative specification of the exchange rate equation:
responses to a positive monetary policy shock in the US



— median ——— 16, 84 percentiles – – 5, 95 percentiles

Figure 5

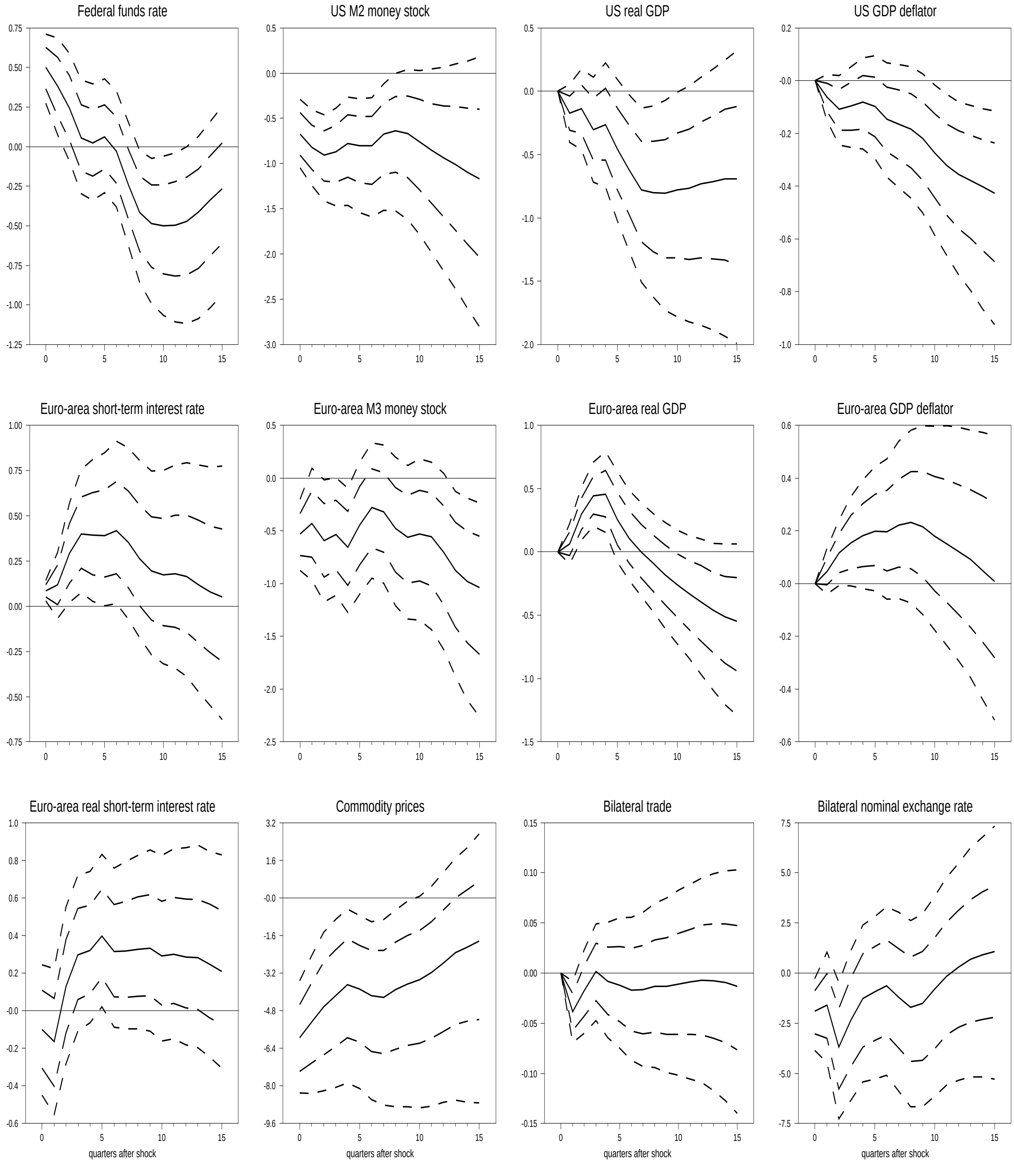
Two-country model with alternative specification of the trade balance equation: responses to a positive monetary policy shock in the US



— median ——— 16, 84 percentiles – – 5, 95 percentiles

Figure 6

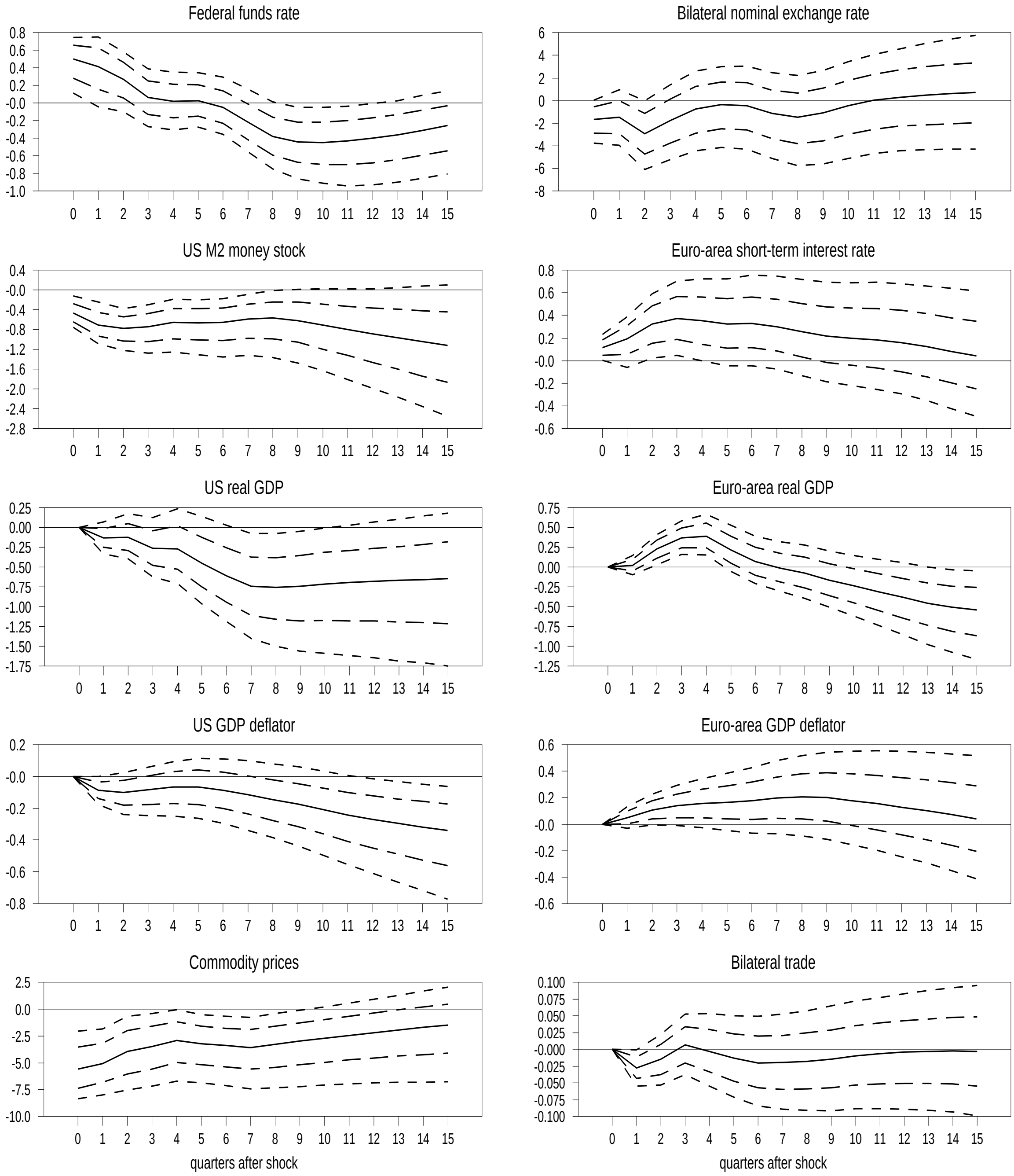
Two-country model with euro-area M3 money stock:
responses to a positive monetary policy shock in the US



— median ——— 16, 84 percentiles – – 5, 95 percentiles

Figure 7

Two-country model with a recursive identification of policy rule:
responses to a positive monetary policy shock in the US



— median ——— 16, 84 percentiles – – 5, 95 percentiles

Figure 8A

Two-country VAR: responses of euro-area real private consumption (left box) and investment (right box) to a positive monetary policy shock in the US

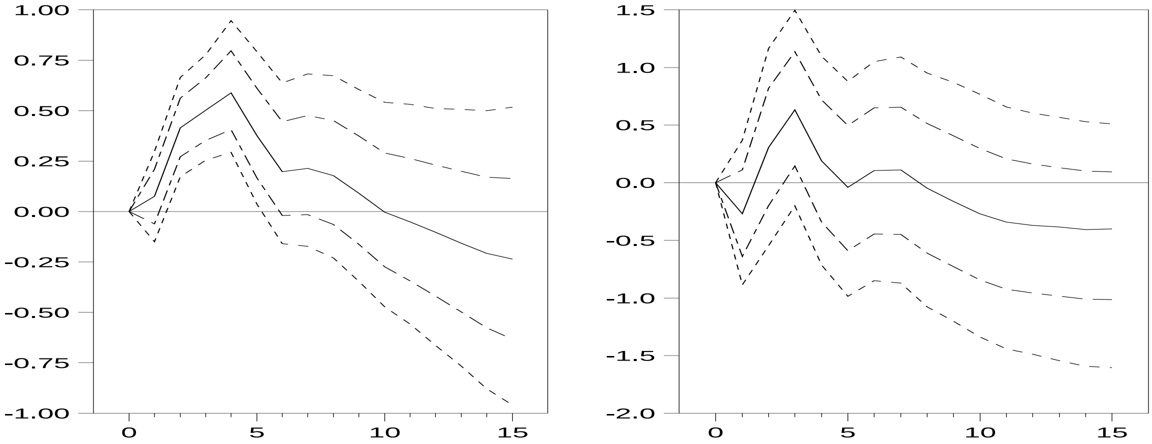
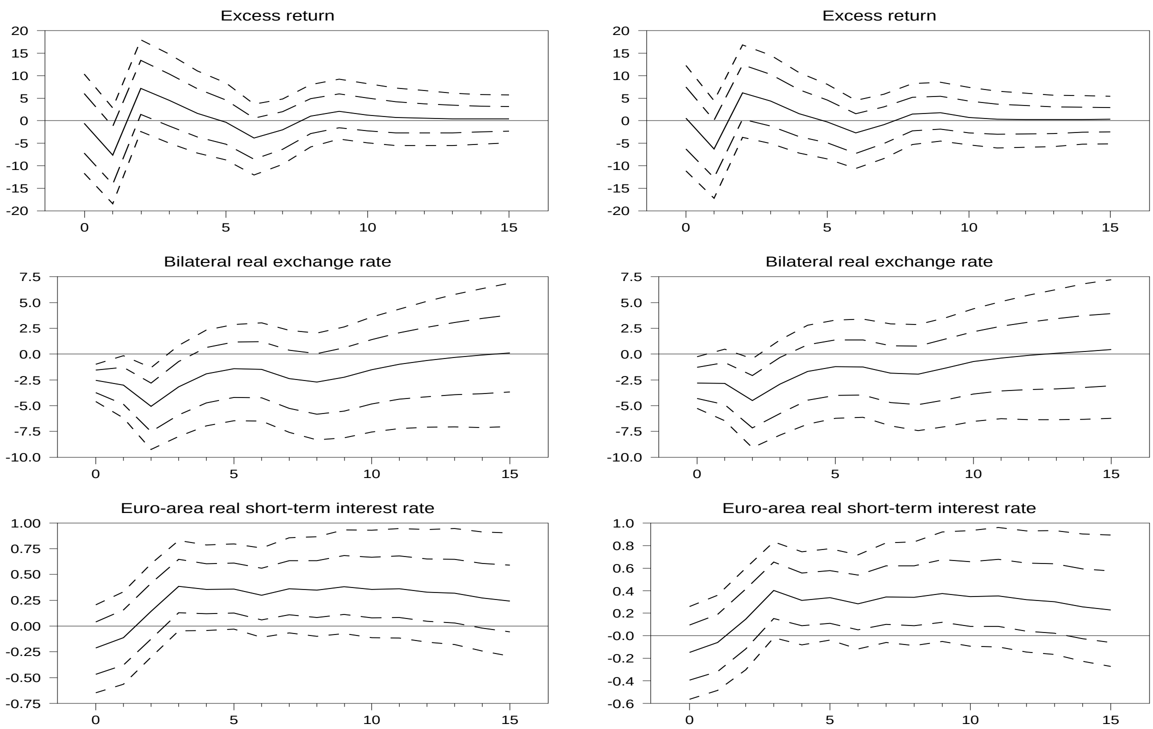


Figure 8B

Two-country VAR: responses of excess return, euro-area real short-term interest rate and real exchange rate to a positive monetary policy shock in the US



— median — – — 16, 84 percentiles – – 5, 95 percentiles

Note: Left column benchmark, right column VAR with the exchange rate responding to all variables.

Appendix A: Description of data and sources

US variables

Real gross private domestic product, billions of chained 2000 dollars, seasonally adjuted annual rate, US Department of Commerce: Bureau of Economic Analysis.

Gross domestic product implicit price deflator, index 2000=100, seasonally adjusted, US Department of Commerce: Bureau of Economic Analysis.

Personal consumption expenditure, chain-type price index, 2000=100, seasonally adjusted, US Department of Commerce: Bureau of Economic Analysis.

Consumer price index for all urban consumers: all items, index 1982-84=100, seasonally adjusted, Federal Reserve Bank of Saint Louis.

Effective federal funds rate, per cent, Federal Reserve Bank of Saint Louis.

Three-month Treasury bill rate, pe rcent, Federal Reserve Bank of Saint Louis.

M1 money stock, billions of dollars, Federal Reserve Bank of Saint Louis.

M2 money stock, billions of dollars, Federal Reserve Bank of Saint Louis.

Unit labour cost, OECD Economic Outlook.

The quarterly bilateral trade balance of the US with respect to the euro area has been computed using monthly international trade statistics of the US Direction of Trade (Bureau of Economic Analysis). We aggregated nominal exports to and nominal imports from the US of each individual euro-area country and seasonally adjusted the resulting time series.

Euro-area variables

Since the beginning of 1991, real gross domestic product, the implicit GDP deflator, real private consumption, real gross fixed capital formation and the consumption deflator are the official time series from Eurostat. Data have been backdated using the quarter-over-quarter growth rates of the corresponding series of the Area Wide Model (AWM) built by Fagan, Henry and Mestre (2001).

Quarterly consumer price index is measured by the quarterly averages of monthly data of the Harmonised Consumer Price Index (HICP).

The short-term interest rate is the three-month money market interest rate. The EU-12 aggregates are taken from the AWM database. After 1998:4, they are updated, respectively, with the three-month EURIBOR as published in the ECB Monthly Bulletin.
M3 money stock, European Central Bank.

Nominal effective exchange rate (PPI and CPI based), Bank of Italy.

Unit labour cost, OECD Economic Outlook.

Other variables

The commodity price index quarterly data are from the Commodity Research Bureau.

Bilateral exchange rate, exchange rate of the dollar vis-á-vis the ECU until 1998:4; exchange rate of the dollar against the euro after 1998:4, Bank of Italy.

Appendix B: Computation of impulse responses and error bands

The economy is described by a system of n structural equations (see Section 2):

$$A_0 Y_t = A(L) Y_{t-1} + e_t \qquad e_t \sim N(0, I) \tag{B-1}$$

which implies the following reduced form:

$$Y_t = A_0^{-1} A(L) Y_{t-1} + A_0^{-1} e_t = B(L) Y_{t-1} + u_t \qquad u_t \sim N(0, \Sigma) \tag{B-2}$$

Given the assumption on the distribution of the structural shocks e_t, the likelihood function has the following representation:

$$p(Y|B(L), \Sigma) \propto |\widehat{\Sigma}|^{-\frac{T}{2}} exp\Big[-\frac{1}{2} trace(S(\widehat{B})\widehat{\Sigma}^{-1})\Big] \qquad S(B) = \sum \widehat{u}_t \widehat{u}_t' \tag{B-3}$$

and assuming a joint diffuse prior distribution

$$p[B(L], A_0) \propto |A_0|^{\nu} \tag{B-4}$$

where ν is the degrees of freedom correction (equal to the number of estimated coefficients in each equation), the posterior distribution of $B(L)$ can be computed:

$$p[B(L)|Y] \sim N\left[\widehat{B}_{ols}, \Sigma(X'X)^{-1}\right] \qquad \Sigma = (A_0^{-1})(A_0^{-1})' \tag{B-5}$$

where $\widehat{B}_{ols}$ is the OLS estimator of $B(L)$. By integrating over $B(L)$ we can obtain the marginal posterior density of A_0

$$p(A_0) \propto |A_0|^{T-\nu} exp\Big[-\frac{1}{2} tr(A_0' S(B) A_0)\Big] \tag{B-6}$$

When the VAR is just-identified we can sample from the distribution of B and then using eq. (B-2) and the assumed A_0 matrix to compute impulse responses. When the VAR is over-identified, we simulate draws from the posterior distribution of A_0 using the the Metropolis algorithm, which is implemented in the following steps:

a) The value of A_0 that maximize the posterior density is used to initialize the algorithm;

b) we generate z from $h(z)$ and u from a uniform distribution over the range $[0, 1]$ where $h(z)$ has a t distribution with ν degrees of freedom and covariance matrix $cS\left(\widehat{B}\right)$, where c is a scaling factor;

c) we compute the n^{th} draw of the Metropolis as $A_0^{(n-1)} + z$;

d) we compute $J(A_0^{(n-1)}, A_0) = min\left[\frac{p(A_0)}{p(A_0^{(n-1)})}, 1\right]$;

e) we set $A_0^{(n)} = A_0$ if $u \leq J(A_0^{(n-1)}, A_0)$ otherwise we set $A_0^{(n)} = A_0^{(n-1)}$;

f) we compute the impulse responses using $Y_t = C(L)Y_{t-1} = [B(L)^{-1}A_0]Y_{t-1}$.

Once a set of N impulse responses has been stored, we compute the percentiles of the marginal distribution of the responses at each step of the impulse horizon.

Appendix C: Diagnostics

Table C1

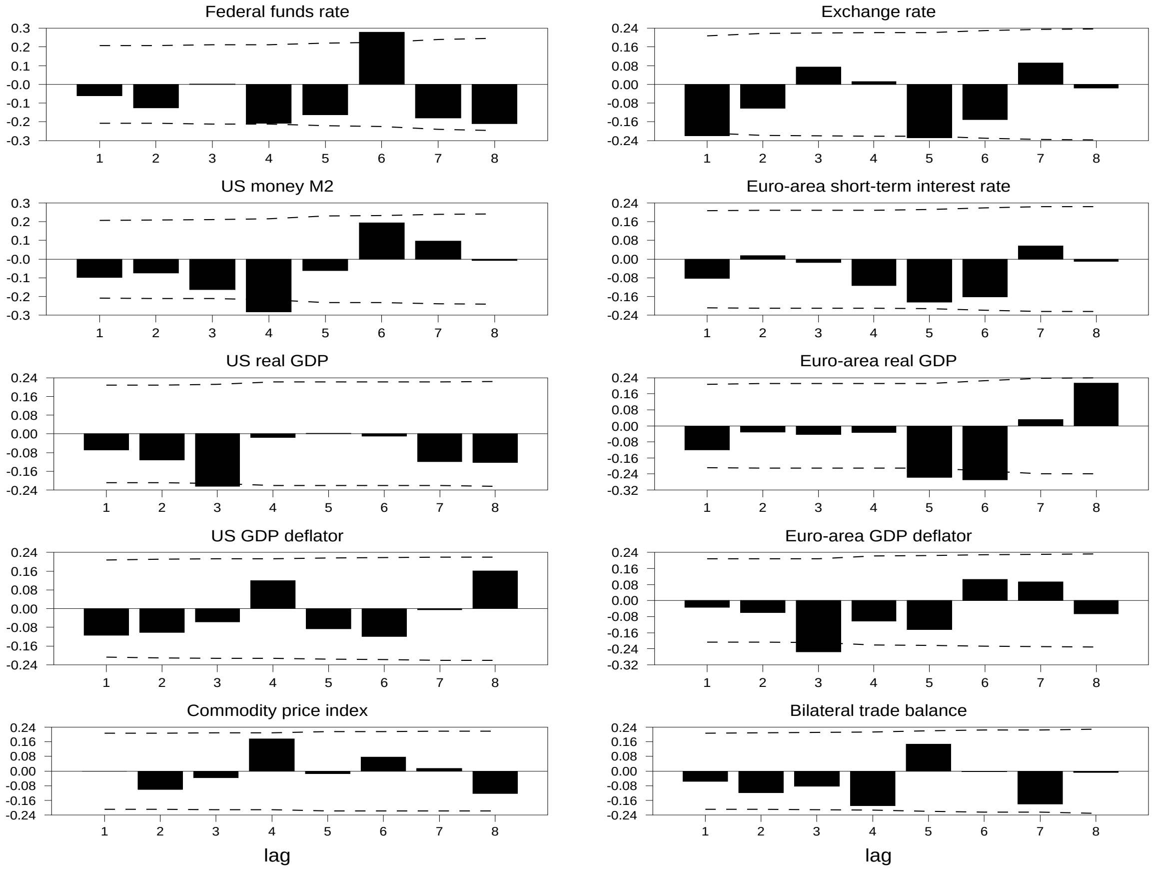
Test for autocorrelation of residuals

Lags	LM test	*p*-value
	two-country VAR	
1	145.4	0.0021
2	155.7	0.0003
3	122.9	0.0597
4	82.7	0.8945
5	117.5	0.1117
6	113.5	0.1685
7	96.3	0.5855
8	91.7	0.7107

$Notes$: H_0: no serial correlation at lag order h.

Figure C1

Two-country VAR: autocorrelation of residuals



References

Bergin, P., R. (2003), “Putting the New Open Economy Macroeconomics to a test”, *Journal of International Economics*, Vol. 60, pp. 3-34.

Bernanke, B., S. and Mihov, I. (1998), “Measuring Monetary Policy”, *Quarterly Journal of Economics*, Vol. 113, pp. 869-902.

Betts, C. and Devereux, M., B. (2000), “Exchange Rate Dynamics in a Model with Pricing-to-Market”, *Journal of International Economics*, Vol. 50, pp. 215-244.

Betts, C. and Devereux, M., B. (2001), “The International Effects of Monetary and Fiscal Policy in a Two-Country Model”, in G. A. Calvo, R. Dornbusch and M. Obstfeld, eds., Money, Capital Mobility: Essays in Honor of Robert A. Mundell, Cambridge, MIT Press.

Buldorini, L., Makrydakis, S. and Thimann, C. (2002), “The Effective Exchange Rate of the Euro”, European Central Bank Occasional Paper, No. 2.

Campa, J. M. and Gonzalez-Minguez, J., M. (2002), “Differences in Exchange Rate Pass-Through in the Euro Area”, Documento de Trabajo, Banco de España, No. 0219.

Chari, V. V., Kehoe, P. J. and McGrattan, E., R. (2002), “Can Sticky Price Models Generate Volatile and Persistent Exchange Rates?”, *Review of Economic Studies*, Vol. 69, pp. 533-563.

Christiano, L. J., Eichenbaum, M. and Evans, C. (1999), “Monetary Policy Shocks: What Have We Learned and to What End?”, In: J. B. Taylor and M. Woodford (eds.), *Handbook of Macroeconomics*, Vol. 1A, pp. 65-148, Elsevier Science, Amsterdam.

Clarida, R., H., Galí, J. and Gertler, M. (1998), “Monetary Policy Rules in Practice: Some International Evidence”, *European Economic Review*, Vol. 42, pp. 1033-1067.

Corsetti, G. and Pesenti, P. (2001), “Welfare and Macroeconomic Interdependence”, *Quarterly Journal of Economics*, Vol. 116, pp. 421-445.

Cushman, D., O. and Zha, T. (1997), “Identifying Monetary Policy in a Small Open Economy under Flexible Exchange Rates”, *Journal of Monetary Economics*, Vol. 39, pp. 433-448.

Dedola, L. and Neri, S. (2005), “What Are the Effects of Technology Shocks? A VAR Analysis with Priors on Impulse Responses”, *Journal of Monetary Economics*, forthcoming.

de Walque, G., Smets, F. and Wouters, R. (2005), "An Open Economy DSGE Model Linking the Euro Area and the US Economy", mimeo, European Central Bank.

Eichenbaum, M. and Evans, C., L. (1995), "Some Empirical Evidence on the Effects of Shocks to Monetary Policy on Exchange Rates", *Quarterly Journal of Economics*, Vol. 110, pp. 975-1010.

Fagan, G., Henry, J. and Mestre, R. (2001), "An Area-Wide Model (AWM) for the Euro Area", ECB Working Paper, No. 42.

Faruqee, A. (2004), "Exchange Rate Pass-Through in the Euro Area: the Role of Asymmetric Pricing Behaviour", IMF Working Paper, No. 14.

Gordon, D., B. and Leeper, E., M. (1994), "The Dynamic Impacts of Monetary Policy: An Exercise in Tentative Identification", *Journal of Political Economy*, Vol. 102, No. 6, pp. 1228-1247.

Francis, N. and Ramey, V. (2005), "Is the technology-driven real business cycle hypothesis dead? Shocks and aggregate fluctuations revisited", *Journal of Monetary Economics*, Vol. 52, No. 8, pp. 1379-1399.

Gopinath, G. and Rigobon, R. (2006), "Sticky Borders", mimeo, Harvard University, MIT and NBER.

Gordon, D., B. and Leeper, E., M. (1994), "The Dynamic Impacts of Monetary Policy: An Exercise in Tentative Identification", *Journal of Political Economy*, Vol. 102, No. 6, pp. 1228-1247.

Hamilton, J. (2003), "What is an Oil Shock?", *Journal of Econometrics*, Vol. 113, pp. 363-398.

Hau, H. (2000), "Exchange Rate Determination: the Role of Factor Price Rigidities and Non-tradeables", *Journal of International Economics*, Vol. 50, No. 2, pp. 421-447.

Hoover, K. D., and S. J. Perez (1994), "Post Hoc Ergo Propter Hoc Once More: An Evaluation of Does Monetary Policy Matter? in the Spirit of James Tobin", *Journal of Monetary Economics*, Vol. 34, pp. 89-99.

Kilian, L. (2006), "Exogenous Oil Supply Shocks: How Big Are They and How Much Do They Matter for the U.S. Economy?", University of Michigan, CEPR, mimeo.

Kim, J. (1999), "Do Monetary Policy Shocks Matter in the G-7 Countries? Using Common Identifying Assumptions About Monetary Policy Across Countries", *Journal of International Economics*, Vol. 48, No. 2, pp. 387-412.

Kim, J. (2001), "International Transmission of US Monetary Policy Shocks: Evidence from VAR's", *Journal of Monetary Economics*, Vol. 48, No. 2, pp. 339-372.

Kim, J. and Roubini, N. (2000), "Exchange Rate Anomalies in the Industrial Countries: A Solution with a Structural VAR Approach", *Journal of Monetary Economics*, Vol. 45, No. 3, pp. 561-586.

Lane, P., R. (2001), "The New Open Economy Macroeconomics: A Survey", *Journal of International Economics*, Vol. 54, No. 2, pp. 235-266.

Leeper, E., M., Sims, A., C. and Zha, T. (1996), "What Does Monetary Policy Do?", *Brookings Papers on Economic Activity*, No. 2, pp. 1-78.

Leeper, E., M. and Roush, J., E. (2003), "Putting "M" Back in Monetary Policy", *Journal of Money, Credit, and Banking*, Vol. 35, pp. 1217-1256.

Lubik, T., A. and Schorfheide, F. (2005), "A Bayesian Look at New Open Economy Macroeconomics", *NBER Macroeconomics Annual 2005.*

Obstfeld, M. and Rogoff, K. (1995), "Exchange Rate Dynamics Redux", *Journal of Political Economy*, Vol. 103, No. 3, pp. 624-660.

Obstfeld, M. and Rogoff, K. (2000), "New Directions for Stochastic Open Economy Models", *Journal of International Economics*, Vol. 50, No. 1, pp. 117-153.

Rabanal, P. and Tuesta, V. (2005), "Euro-Dollar Real Exchange Rate Dynamics in an Estimated Two-Country Model: What is Important and What is Not", mimeo, International Monetary Fund.

Sarno, L. (2001), "Toward a New Paradigm in Open Economy Modelling: Where Do We Stand?", Federal Reserve of Saint Louis, *Review*, Vol. 83, No.3, pp. 21-36.

Sims, A., C. (1992), "Interpreting the Macroeconomic Time Series Facts: the Effects of Monetary Policy", *European Economic Review*, Vol. 36, No. 5, pp. 975-1000.

Sims, A., C., Stock, J. and Watson, M. (1990), "Inference in Linear Time Series Models with Some Unit Roots", *Econometrica*, Vol. 58, No. 1, pp. 113-144.

Sims, A., C. and Zha, T. (1999), "Error Bands for Impulse Responses", *Econometrica*, Vol. 67, No. 5, pp. 1113-1155.

Sims, A., C. and Zha, T. (2006), "Were There Regime Switches in US Monetary Policy?", *American Economic Review*, Vol. 96, No. 1, pp. 54-81.

Waggoner, D., F. and Zha, T. (1997), "Normalization, Probability Distribution and Impulse Responses", Federal Reserve Bank of Atlanta, Working Paper Series, No. 97-11.

Waggoner, D., F. and Zha, T. (2003a), "A Gibbs Sampler for Structural Vector Autoregression", *Journal of Economic Dynamics and Control*, Vol. 28, pp. 349-366.

Waggoner, D., F. and Zha, T. (2003b), "Likelihood-Preserving Normalization for Multiple Equations", *Journal of Econometrics*, Vol. 114, pp. 329-347.

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2000

P. ANGELINI, *Are banks risk-averse? Intraday timing of the operations in the interbank market*, Journal of Money, Credit and Banking, Vol. 32 (1), pp. 54-73, **TD No. 266 (April 1996).**

F. DRUDI and R. GIORDANO, *Default Risk and optimal debt management,* Journal of Banking and Finance, Vol. 24 (6), pp. 861-891, **TD No. 278 (September 1996)**.

F. DRUDI and R. GIORDANO, *Wage indexation, employment and inflation,* Scandinavian Journal of Economics, Vol. 102 (4), pp. 645-668, **TD No. 292 (December 1996)**.

F. DRUDI and A. PRATI, *Signaling fiscal regime sustainability,* European Economic Review, Vol. 44 (10), pp. 1897-1930, **TD No. 335 (September 1998)**.

F. FORNARI and R. VIOLI, *The probability density function of interest rates implied in the price of options,* in: R. Violi, (ed.) , Mercati dei derivati, controllo monetario e stabilità finanziaria, Il Mulino, Bologna, **TD No. 339 (October 1998)**.

D. J. MARCHETTI and G. PARIGI, *Energy consumption, survey data and the prediction of industrial production in Italy,* Journal of Forecasting, Vol. 19 (5), pp. 419-440, **TD No. 342 (December 1998)**.

A. BAFFIGI, M. PAGNINI and F. QUINTILIANI, *Localismo bancario e distretti industriali: assetto dei mercati del credito e finanziamento degli investimenti,* in: L.F. Signorini (ed.), Lo sviluppo locale: un'indagine della Banca d'Italia sui distretti industriali, pp. 237-256, Meridiana Libri, **TD No. 347** (**March 1999)**.

F. LIPPI, *Median voter preferences, central bank independence and conservatism,* Public Choice, v. 105, 3-4, pp. 323-338 **TD No. 351 (April 1999)**.

A. SCALIA and V. VACCA, *Does market transparency matter? A case study,* in: Market Liquidity: Research Findings and Selected Policy Implications, Basel, Bank for International Settlements, **TD No. 359 (October 1999)**.

F. SCHIVARDI, *Rigidità nel mercato del lavoro, disoccupazione e crescita*, Giornale degli economisti e Annali di economia, Vol. 59 (1), pp. 115-141, **TD No. 364 (December 1999)**.

G. BODO, R. GOLINELLI and G. PARIGI, *Forecasting industrial production in the euro area,* Empirical Economics, Vol. 25 (4), pp. 541-561, **TD No. 370 (March 2000)**.

F. ALTISSIMO, D. J. MARCHETTI and G. P. ONETO, *The Italian business cycle: Coincident and leading indicators and some stylized facts,* Giornale degli economisti e Annali di economia, Vol. 60 (2), pp. 147-220, **TD No. 377 (October 2000)**.

C. MICHELACCI and P. ZAFFARONI, *(Fractional) Beta convergence*, Journal of Monetary Economics, Vol. 45 (1), pp. 129-153, **TD No. 383 (October 2000)**.

R. DE BONIS and A. FERRANDO, *The Italian banking structure in the nineties: Testing the multimarket contact hypothesis*, Economic Notes, Vol. 29 (2), pp. 215-241, **TD No. 387 (October 2000)**.

S. SIVIERO and D. TERLIZZESE, *La previsione macroeconomica: alcuni luoghi comuni da sfatare*, Rivista italiana degli economisti, v. 5, 2, pp. 291-322, **TD No. 395 (February 2001)**.

G. DE BLASIO and F. MINI, *Seasonality and capacity: An application to Italy*, IMF Working Paper, 80, **TD No. 403 (June 2001)**.

2001

M. CARUSO, *Stock prices and money velocity: A multi-country analysis,* Empirical Economics, Vol. 26 (4), pp. 651-672, **TD No. 264 (February 1996)**.

P. CIPOLLONE and D. J. MARCHETTI, *Bottlenecks and limits to growth: A multisectoral analysis of Italian industry*, Journal of Policy Modeling, Vol. 23 (6), pp. 601-620, **TD No. 314 (August 1997)**.

P. CASELLI, *Fiscal consolidations under fixed exchange rates*, European Economic Review, Vol. 45 (3), pp. 425-450, **TD No. 336 (October 1998)**.

F. ALTISSIMO and G. L. VIOLANTE, *The non-linear dynamics of output and unemployment in the US,* Journal of Applied Econometrics, Vol. 16 (4), pp. 461-486, **TD No. 338 (October 1998)**.

F. NUCCI and A. F. POZZOLO, *Investment and the exchange rate: An analysis with firm-level panel data,* European Economic Review, Vol. 45 (2), pp. 259-283, **TD No. 344 (December 1998)**.

A. ZAGHINI, *Fiscal adjustments and economic performing: A comparative study*, Applied Economics, Vol. 33 (5), pp. 613-624, **TD No. 355 (June 1999)**.

L. GAMBACORTA, *On the institutional design of the European monetary union: Conservatism, stability pact and economic shocks,* Economic Notes, Vol. 30 (1), pp. 109-143, **TD No. 356 (June 1999)**.

P. FINALDI RUSSO and P. ROSSI, *Credit costraints in italian industrial districts,* Applied Economics, Vol. 33 (11), pp. 1469-1477, **TD No. 360 (December 1999)**.

A. CUKIERMAN and F. LIPPI, *Labor markets and monetary union: A strategic analysis,* Economic Journal, Vol. 111 (473), pp. 541-565, **TD No. 365 (February 2000)**.

G. PARIGI and S. SIVIERO, *An investment-function-based measure of capacity utilisation, potential output and utilised capacity in the Bank of Italy's quarterly model*, Economic Modelling, Vol. 18 (4), pp. 525-550, **TD No. 367 (February 2000)**.

P. CASELLI, P. PAGANO and F. SCHIVARDI, *Investment and growth in Europe and in the United States in the nineties,* Rivista di politica economica, v. 91, 10, pp. 3-35, **TD No. 372 (March 2000)**.

F. BALASSONE and D. MONACELLI, *Emu fiscal rules: Is there a gap?,* in: M. Bordignon and D. Da Empoli (eds.), Politica fiscale, flessibilità dei mercati e crescita, Milano, Franco Angeli, **TD No. 375 (July 2000)**.

A. B. ATKINSON and A. BRANDOLINI, *Promise and pitfalls in the use of "secondary" data-sets: Income inequality in OECD countries as a case study*, Journal of Economic Literature, Vol. 39 (3), pp. 771-799, **TD No. 379 (October 2000)**.

D. FOCARELLI and A. F. POZZOLO, *The patterns of cross-border bank mergers and shareholdings in OECD countries*, Journal of Banking and Finance, Vol. 25 (12), pp. 2305-2337, **TD No. 381 (October 2000)**.

M. SBRACIA and A. ZAGHINI, *Expectations and information in second generation currency crises models*, Economic Modelling, Vol. 18 (2), pp. 203-222, **TD No. 391 (December 2000)**.

F. FORNARI and A. MELE, *Recovering the probability density function of asset prices using GARCH as diffusion approximations,* Journal of Empirical Finance, Vol. 8 (1), pp. 83-110**, TD No. 396 (February 2001)**.

P. CIPOLLONE, *La convergenza dei salari dell'industria manifatturiera in Europa,* Politica economica, Vol. 17 (1), pp. 97-125, **TD No. 398 (February 2001)**.

E. BONACCORSI DI PATTI and G. GOBBI, *The changing structure of local credit markets: Are small businesses special?,* Journal of Banking and Finance, Vol. 25 (12), pp. 2209-2237, **TD No. 404 (June 2001)**.

L. DEDOLA and S. LEDUC, *Why is the business-cycle behaviour of fundamentals alike across exchange-rate regimes?,* International Journal of Finance and Economics, v. 6, 4, pp. 401-419, **TD No. 411 (August 2001)**.

M. PAIELLA, *Limited Financial Market Participation: a Transaction Cost-Based Explanation,* IFS Working Paper, 01/06, **TD No. 415 (August 2001)**.

G. MESSINA, *Per un federalismo equo e solidale: obiettivi e vincoli per la perequazione regionale in Italia,*, Studi economici, Vol. 56 (73), pp. 131-148, **TD No. 416 (August 2001)**.

L GAMBACORTA *Bank-specific characteristics and monetary policy transmission: the case of Italy,* ECB Working Paper, 103, **TD No. 430 (December 2001)**.

F. ALTISSIMO, A. BASSANETTI, R. CRISTADORO, M. FORNI, M. LIPPI, L. REICHLIN and G. VERONESE *A real time coincident indicator of the euro area business cycle,* CEPR Discussion Paper, 3108, **TD No. 436 (December 2001)**.

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L. GUISO and M. PAIELLA, *Risk aversion, wealth and background risk,* CEPR Discussion Paper, 2728, **TD No. 483 (September 2003)**.

2002

R. CESARI and F. PANETTA, *The performance of italian equity fund*, Journal of Banking and Finance, Vol. 26 (1), pp. 99-126, **TD No. 325 (January 1998)**.

F. ALTISSIMO, S. SIVIERO and D. TERLIZZESE, *How deep are the deep parameters?,* Annales d'Economie et de Statistique, (67/68), pp. 207-226, **TD No. 354 (June 1999)**.

F. FORNARI, C. MONTICELLI, M. PERICOLI and M. TIVEGNA, *The impact of news on the exchange rate of the lira and long-term interest rates*, Economic Modelling, Vol. 19 (4), pp. 611-639, **TD No. 358 (October 1999**).

D. FOCARELLI, F. PANETTA and C. SALLEO, *Why do banks merge?*, Journal of Money, Credit and Banking, Vol. 34 (4), pp. 1047-1066, **TD No. 361 (December 1999)**.

D. J. MARCHETTI, *Markup and the business cycle: Evidence from Italian manufacturing branches,* Open Economies Review, Vol. 13 (1), pp. 87-103, **TD No. 362 (December 1999**).

F. BUSETTI, *Testing for (common) stochastic trends in the presence of structural break*, Journal of Forecasting, Vol. 21 (2), pp. 81-105, **TD No. 385 (October 2000)**.

F. LIPPI, *Revisiting the Case for a Populist Central Banker*, European Economic Review, Vol. 46 (3), pp. 601-612, **TD No. 386 (October 2000)**.

F. PANETTA, *The stability of the relation between the stock market and macroeconomic forces*, Economic Notes, Vol. 31 (3), pp. 417-450, **TD No. 393 (February 2001)**.

G. GRANDE and L. VENTURA, *Labor income and risky assets under market incompleteness: Evidence from Italian data*, Journal of Banking and Finance, Vol. 26 (2-3), pp. 597-620, **TD No. 399 (March 2001)**.

A. BRANDOLINI, P. CIPOLLONE and P. SESTITO, *Earnings dispersion, low pay and household poverty in Italy, 1977-1998,* in D. Cohen, T. Piketty and G. Saint-Paul (eds.), The Economics of Rising Inequalities, Oxford, Oxford University Press, **TD No. 427 (November 2001)**.

E. GAIOTTI and A. GENERALE, *Does monetary policy have asymmetric effects? A look at the investment decisions of Italian firms*, Giornale degli economisti e annali di economia, v. 61, 1, pp. 29-60, **TD No. 429 (December 2001)**.

G. M. TOMAT, *Durable goods, price indexes and quality change: An application to automobile prices in Italy, 1988-1998,* ECB Working Paper, 118, **TD No. 439 (March 2002)**.

A. PRATI and M. SBRACIA, *Currency crises and uncertainty about fundamentals*, IMF Working Paper, 3, **TD No. 446 (July 2002)**.

L. CANNARI and G. D'ALESSIO, *La distribuzione del reddito e della ricchezza nelle regioni italiane*, Rivista Economica del Mezzogiorno, Vol. 16 (4), pp. 809-847, Il Mulino, **TD No. 482 (June 2003)**.

2003

L. GAMBACORTA, *Asymmetric bank lending channels and ECB monetary policy,* Economic Modelling, Vol. 20, 1, pp. 25-46, **TD No. 340 (October 1998)**.

F. SCHIVARDI, *Reallocation and learning over the business cycle,* European Economic Review, Vol. 47 (1), pp. 95-111, **TD No. 345 (December 1998)**.

P. CASELLI, P. PAGANO and F. SCHIVARDI, *Uncertainty and slowdown of capital accumulation in Europe,* Applied Economics, Vol. 35 (1), pp. 79-89, **TD No. 372 (March 2000).**

F. LIPPI, *Strategic monetary policy with non-atomistic wage setters*, Review of Economic Studies, v. 70, 4, pp. 909-919, **TD No. 374 (June 2000)**.

P. ANGELINI and N. CETORELLI, *The effect of regulatory reform on competition in the banking industry,* Journal of Money, Credit and Banking, Vol. 35, 5, pp. 663-684, **TD No. 380 (October 2000)**.

P. PAGANO and G. FERRAGUTO, Endogenous growth with intertemporally dependent preferences, Contribution to Macroeconomics, Vol. 3 (1), pp. 1-38, **TD No. 382 (October 2000).**

P. PAGANO and F. SCHIVARDI, *Firm size distribution and growth,* Scandinavian Journal of Economics, Vol. 105 (2), pp. 255-274, **TD No. 394 (February 2001)**.

M. PERICOLI and M. SBRACIA, *A Primer on Financial Contagion*, Journal of Economic Surveys, Vol. 17 (4), pp. 571-608, **TD No. 407 (June 2001)**.

M. SBRACIA and A. ZAGHINI, *The role of the banking system in the international transmission of shocks,* World Economy, Vol. 26 (5), pp. 727-754, **TD No. 409 (June 2001)**.

L. GAMBACORTA, *The Italian banking system and monetary policy transmission: evidence from bank level data*, in: I. Angeloni, A. Kashyap and B. Mojon (eds.), Monetary Policy Transmission in the Euro Area, Cambridge University Press, **TD No. 430 (December 2001).**

M. EHRMANN, L. GAMBACORTA, J. MARTÍNEZ PAGÉS, P. SEVESTRE and A. WORMS, *Financial systems and the role of banks in monetary policy transmission in the euro area,* in: I. Angeloni, A. Kashyap and B. Mojon (eds.), Monetary Policy Transmission in the Euro Area, Cambridge, Cambridge University Press, **TD No. 432 (December 2001)**.

F. SPADAFORA, *Official bailouts, moral hazard and the "Specialtiy" of the international interbank market,* Emerging Markets Review, Vol. 4 (2), pp. 165-196, **TD No. 438 (March 2002)**.

D. FOCARELLI and F. PANETTA, *Are mergers beneficial to consumers? Evidence from the market for bank deposits*, American Economic Review, Vol. 93 (4), pp. 1152-1172, **TD No. 448 (July 2002)**.

E.VIVIANO, *Un'analisi critica delle definizioni di disoccupazione e partecipazione in Italia,* Politica Economica, Vol. 19 (1), pp. 161-190, **TD No. 450 (July 2002)**.

M. PAGNINI**,** *Misura e determinanti dell'agglomerazione spaziale nei comparti industriali in Italia,* Rivista di Politica Economica, Vol. 93 (3-4), pp. 149-196, **TD No. 452 (October 2002)**.

F. PANETTA, *Evoluzione del sistema bancario e finanziamento dell'economia nel Mezzogiorno,* Moneta e credito, v. 56, 222, pp. 127-160, **TD No. 467 (March 2003)**.

F. BUSETTI and A. M. ROBERT TAYLOR, *Testing against stochastic trend and seasonality in the presence of unattended breaks and unit roots*, Journal of Econometrics, Vol. 117 (1), pp. 21-53, **TD No. 470 (March 2003)**.

P. ZAFFARONI, *Testing against stochastic trend and seasonality in the presence of unattended breaks and unit roots,* Journal of Econometrics, v. 115, 2, pp. 199-258, **TD No. 472 (June 2003)**.

E. BONACCORSI DI PATTI, G. GOBBI and P. E. MISTRULLI, *Sportelli e reti telematiche nella distribuzione dei servizi bancari,* Banca impresa società, v. 2, 2, pp. 189-209, **TD No. 508 (July 2004)**.

2004

P. ANGELINI and N. CETORELLI, *Gli effetti delle modifiche normative sulla concorrenza nel mercato creditizio,* in F. Panetta (eds.), Il sistema bancario negli anni novanta: gli effetti di una trasformazione, Bologna, il Mulino, **TD No. 380 (October 2000)**.

P. CHIADES and L. GAMBACORTA, *The Bernanke and Blinder model in an open economy: The Italian case,* German Economic Review, Vol. 5 (1), pp. 1-34, **TD No. 388 (December 2000)**.

M. BUGAMELLI and P. PAGANO, *Barriers to Investment in ICT,* Applied Economics, Vol. 36 (20), pp. 2275-2286, **TD No. 420 (October 2001)**.

F. BUSETTI, *Preliminary data and econometric forecasting: An application with the Bank of Italy quarterly model*, CEPR Discussion Paper, 4382, **TD No. 437 (December 2001)**.

A. BAFFIGI, R. GOLINELLI and G. PARIGI, *Bridge models to forecast the euro area GDP,* International Journal of Forecasting, Vol. 20 (3), pp. 447-460,**TD No. 456 (December 2002)**.

D. AMEL, C. BARNES, F. PANETTA and C. SALLEO, *Consolidation and Efficiency in the Financial Sector: A Review of the International Evidence*, Journal of Banking and Finance, Vol. 28 (10), pp. 2493-2519, **TD No. 464 (December 2002)**.

M. PAIELLA, *Heterogeneity in financial market participation: Appraising its implications for the C-CAPM,* Review of Finance, Vol. 8, 3, pp. 445-480, **TD No. 473 (June 2003)**.

F. CINGANO and F. SCHIVARDI, *Identifying the sources of local productivity growth*, Journal of the European Economic Association, Vol. 2 (4), pp. 720-742, **TD No. 474 (June 2003)**.

E. BARUCCI, C. IMPENNA and R. RENÒ, *Monetary integration, markets and regulation*, Research in Banking and Finance, (4), pp. 319-360, **TD No. 475 (June 2003)**.

G. ARDIZZI, *Cost efficiency in the retail payment networks: first evidence from the Italian credit card system*, Rivista di Politica Economica, Vol. 94, (3), pp. 51-82, **TD No. 480 (June 2003)**.

E. BONACCORSI DI PATTI and G. DELL'ARICCIA, *Bank competition and firm creation*, Journal of Money Credit and Banking, Vol. 36 (2), pp. 225-251, **TD No. 481 (June 2003)**.

R. GOLINELLI and G. PARIGI, *Consumer sentiment and economic activity: a cross country comparison*, Journal of Business Cycle Measurement and Analysis, Vol. 1 (2), pp. 147-170, **TD No. 484 (September 2003)**.

L. GAMBACORTA and P. E. MISTRULLI, *Does bank capital affect lending behavior?*, Journal of Financial Intermediation, Vol. 13 (4), pp. 436-457, **TD No. 486 (September 2003)**.

F. SPADAFORA, *Il pilastro privato del sistema previdenziale: il caso del Regno Unito,* Economia Pubblica, 34, (5), pp. 75-114, **TD No. 503** (**June 2004**).

C. BENTIVOGLI and F. QUINTILIANI, *Tecnologia e dinamica dei vantaggi comparati: un confronto fra quattro regioni italiane*, in C. Conigliani (eds.), *Tra sviluppo e stagnazione: l'economia dell'Emilia-Romagna*, Bologna, Il Mulino, **TD No. 522 (October 2004)**.

G. GOBBI and F. LOTTI, *Entry decisions and adverse selection: an empirical analysis of local credit markets,* Journal of Financial services Research, Vol. 26 (3), pp. 225-244, **TD No. 535 (December 2004)**.

E. GAIOTTI and F. LIPPI, *Pricing behavior and the introduction of the euro:evidence from a panel of restaurants*, Giornale degli Economisti e Annali di Economia, 2004, Vol. 63, (3/4), pp. 491-526, **TD No. 541 (February 2005)**.

2005

L. DEDOLA and F. LIPPI, *The monetary transmission mechanism: Evidence from the industries of 5 OECD countries,* European Economic Review, 2005, Vol. 49, (6), pp. 1543-1569, **TD No. 389 (December 2000)**.

D. J. MARCHETTI and F. NUCCI, *Price stickiness and the contractionary effects of technology shocks*. European Economic Review, v. 49, pp. 1137-1164, **TD No. 392 (February 2001)**.

G. CORSETTI, M. PERICOLI and M. SBRACIA, *Some contagion, some interdependence: More pitfalls in tests of financial contagion,* Journal of International Money and Finance, v. 24, 8, pp. 1177-1199, **TD No. 408 (June 2001)**.

R. CRISTADORO, M. FORNI, L. REICHLIN and G. VERONESE, *A core inflation indicator for the euro area,* Journal of Money, Credit, and Banking, v. 37, 3, pp. 539-560, **TD No. 435 (December 2001)**.

F. ALTISSIMO, E. GAIOTTI and A. LOCARNO, *Is money informative? Evidence from a large model used for policy analysis,* Economic & Financial Modelling, v. 22, 2, pp. 285-304, **TD No. 445 (July 2002)**.

G. DE BLASIO and S. DI ADDARIO, *Do workers benefit from industrial agglomeration?* Journal of regional Science, Vol. 45, (4), pp. 797-827, **TD No. 453 (October 2002).**

R. TORRINI, *Cross-country differences in self-employment rates: The role of institutions,* Labour Economics, V. 12, 5, pp. 661-683, **TD No. 459 (December 2002)**.

A. CUKIERMAN and F. LIPPI, *Endogenous monetary policy with unobserved potential output,* Journal of Economic Dynamics and Control, v. 29, 11, pp. 1951-1983, **TD No. 493 (June 2004)**.

M. OMICCIOLI, *Il credito commerciale: problemi e teorie,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia,* Bologna, Il Mulino, **TD No. 494 (June 2004)**.

L. CANNARI, S. CHIRI and M. OMICCIOLI, *Condizioni di pagamento e differenziazione della clientela,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 495 (June 2004)**.

P. FINALDI RUSSO and L. LEVA, *Il debito commerciale in Italia: quanto contano le motivazioni finanziarie?,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia,* Bologna, Il Mulino, **TD No. 496 (June 2004)**.

A. CARMIGNANI, *Funzionamento della giustizia civile e struttura finanziaria delle imprese: il ruolo del credito commerciale,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 497 (June 2004)**.

G. DE BLASIO, *Credito commerciale e politica monetaria: una verifica basata sull’investimento in scorte,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 498 (June 2004)**.

G. DE BLASIO, *Does trade credit substitute bank credit? Evidence from firm-level data.* Economic notes, Vol. 34 (1), pp. 85-112, **TD No. 498 (June 2004).**

A. DI CESARE, *Estimating Expectations of Shocks Using Option Prices,* The ICFAI Journal of Derivatives Markets, Vol. 2, (1), pp. 42-53, **TD No. 506 (July 2004).**

M. BENVENUTI and M. GALLO, *Il ricorso al "factoring" da parte delle imprese italiane,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 518 (October 2004)**.

F. PANETTA, F. SCHIVARDI and M. SHUM, *Do mergers improve information? Evidence from the loan market,* CEPR Discussion Paper, 4961, **TD No. 521 (October 2004)**.

P. DEL GIOVANE and R. SABBATINI, *La divergenza tra inflazione rilevata e percepita in Italia,* Bologna, Il Mulino, **TD No. 532 (December 2004).**

M. OMICCIOLI, *Il credito commerciale come "collateral"*, in L. Cannari, S. Chiri, M. Omiccioli (eds.), Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia, Bologna, il Mulino, **TD No. 553 (June 2005)**.

S. DI ADDARIO and E. PATACCHINI, *Wages and the city: The italian case,* University of Oxford, Department of Economics. Discussion Paper, 243, **TD No. 570 (January 2006)**.

P. ANGELINI and F. LIPPI, *Did inflation really soar after the euro changeover? Indirect evidence from ATM withdrawals,* CEPR Discussion Paper, 4950, **TD No. 581 (March 2006)**.

2006

C. BIANCOTTI, *A polarization of inequality? The distribution of national Gini coefficients 1970-1996,* Journal of Economic Inequality, v. 4, 1, pp. 1-32, **TD No. 487 (March 2004)**.

M. BOFONDI and G. GOBBI, *Information barriers to entry into credit markets,* Review of Finance, Vol. 10 (1), pp. 39-67, **TD No. 509 (July 2004).**

A. BRANDOLINI, P. CIPOLLONE and E. VIVIANO, *Does the ILO definition capture all unemployment?,* Journal of the European Economic Association, v. 4, 1, pp. 153-179, **TD No. 529 (December 2004)**.

G. M. TOMAT, *Prices product differentiation and quality measurement: A comparison between hedonic and matched model methods*, Research in Economics, No. 60, pp. 54-68, **TD No. 547 (February 2005)**.

M. CARUSO, *Stock market fluctuations and money demand in Italy, 1913 - 2003,* Economic Notes, v. 35, 1, pp. 1-47, **TD No. 576 (February 2006)**.

R. BRONZINI and G. DE BLASIO, *Evaluating the impact of investment incentives: The case of Italy's Law 488/92.* Journal of Urban Economics, vol. 60, n. 2, pag. 327-349, **TD No. 582 (March 2006).**

A. DI CESARE, *Do market-based indicators anticipate rating agencies? Evidence for international banks,* Economic Notes, v. 35, pp. 121-150, **TD No. 593 (May 2006).**

FORTHCOMING

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