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Abstract

Based on a structural VAR and a dynamic general equilibrium model, we provide evidence of the changes in the monetary transmission mechanism (MTM) in the European Monetary Union after the adoption of the common currency in 1999. The estimation of a Bayesian VAR over the periods before and after 1999 suggests that the effects of a monetary policy shock on output and prices have not significantly changed over time. We claim that this cannot be the final word on the evolution of the MTM as changes in the conduct of monetary policy and the structure of the economy may have offset each other giving rise to similar responses of output and inflation to monetary policy shocks between the two periods. The estimation of a DSGE model with several real and nominal frictions over the two sub-samples shows that monetary policy has become more effective in stabilizing the economy as the result of a decrease in the degree of nominal rigidities and a shift in monetary policy towards inflation stabilization.

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Keywords: monetary policy, transmission mechanism, Bayesian methods.

Contents

* Bank of Italy, Economics, Research and International Relations.

1 Introduction*

Over the last two decades most industrialized countries have experienced a sustained increase in trade, significant changes in the way financial markets operate, reforms toward the liberalization of product and labour markets and a stronger focus of central banks on price stability. In continental Europe, the creation of the European Monetary Union (EMU) in 1999 has been a crucial institutional change that has potentially affected the economies of the member states. The elimination of the exchange rate risk might have spurred trade integration among member countries; the establishment of the European Central Bank (ECB) with a clear mandate to stabilise inflation could have changed the way in which expectations are formed, with potential effects on consumption and investment decisions by households and firms. As more than ten years have passed since the creation of the EMU, there are now sufficient data to allow for a study of the changes that may have occurred in the transmission mechanism of monetary policy (MTM henceforth).

The goals of the paper are to document the changes that may have occurred in the MTM with the creation of the euro and to identify the causes behind these changes. Understanding the evolution of the MTM and disentangling the factors behind are crucial for the assessment of the policy stance and for correctly quantifying the macroeconomic effects of policy decisions. To pursue our objectives we use two approaches: a structural VAR and a dynamic general equilibrium model. We choose to rely also on VAR methods in order to have results that are directly comparable with those of the literature on the subject. The results of the VAR analysis are, however: (*i*) not fully informative on the evolution of the MTM as there could have been changes in more than one of the structural parameters of the data generating process (DGP) of the economy that may have offsetting effects on the VAR representation of the DGP; (*ii*) difficult to interpret as the factors behind an observed change in the response of output and prices to monetary policy shocks cannot be disentangled. For these reasons, we complement the VAR evidence with the estimation of a simplified version of the Smets and Wouters [2007] model over two sample periods, before and after the adoption of the euro. The estimation of a more

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structural model allows us to disentangle the various channels at work, and in particular to understand whether there has been a change in the conduct of monetary policy or in the parameters that characterize the behaviour of the private sector. Furthermore, counterfactual exercises can be performed with a DSGE model, while the reliability of them in the context of structural VAR models is questionable (Benati and Surico [2009] and Benati [2009]).

The monetary transmission mechanism is one of the most largely studied area of monetary economics and this paper is related to a large part of this literature. In the early years of the EMU extensive research has been carried out by the ECB and the national central banks of the Eurosystem (Monetary Transmission Network, MTN) to uncover the main stylized facts of the monetary transmission mechanism both at the aggregate and at the countries level.[1] The main results of the network are the following: (*i*) changes in the monetary policy instrument have temporary effects on aggregate euro area output and long lasting ones on prices; (*ii*) monetary policy affects the economy mainly through the interest rate channel; (*iii*) credit constraints do not play a crucial role at the aggregate level; (*iv*) it is difficult to detect systematic differences across countries. These findings were obtained with a sample period that included only the years prior to the adoption of the euro. Since the MTN provided no assessment on the monetary transmission mechanism after the creation of EMU, our contribution to this literature is to update the analysis at the aggregate level with the additional data that have become available since then.

While there are several studies that investigate the changes in the MTM of the U.S. economy (see for instance Boivin and Giannoni [2006] and Boivin et al. [2009]), few empirical analysis focus on the evolution of the monetary transmission in the euro area after the creation of the EMU and the establishment of the single monetary policy. Among these studies, Weber et al. [2011] provide statistical evidence that a break might have occurred between 1996 and 1999 but they conclude that overall the monetary transmission mechanism in the euro area has not significantly changed. While Weber et al. [2011] adopt an area-wide perspective, Boivin et al. [2008] study the transmission mechanism of common monetary shocks to a subset of euro area countries and conclude that this mechanism has, indeed, changed with the creation of the EMU. The introduction of the euro brought about an overall reduction of the effects of monetary policy on output,

[1] See Angeloni et al. [2003] and the article "Recent findings on monetary policy transmission in the euro area" in the October 2002 Monthly Bulletin of the ECB. See also http://www.ecb.int/home/html/researcher_mtn.en.html at the ECB website.

inflation and the long-term interest rate and an increase in the effects on the exchange rate. The authors rationalize these findings in a stylized and calibrated open-economy DSGE model with an increase in the aggressiveness of monetary policy towards inflation and output and with the disappearance of exchange rate risks. Our paper contributes to the literature by providing a structural interpretation of the changes in the MTM through the estimation of a fully-fledged DSGE before and after the introduction of the euro.

Differently from Boivin et al. [2008], we choose an area-wide approach. There are at least two reasons why we think this is reasonable. First of all, the MTN showed that there cannot be detected significant cross-country differences. Mojon and Peersman [2001], in a country level analysis of the MTM, illustrate that the results are qualitatively similar across countries. The differences in the size of the effects for each countries, while clearly visible on the mean responses, disappear when accounting for uncertainty. Furthermore, if there is some degree of heterogeneity, it has not changed over time (see for instance Ciccarelli and Rebucci [2006] on the effects of a monetary policy shock and Giannone et al. [2008] on the unconditional properties of the business cycle) and it is due mostly to idiosyncratic shocks (see Giannone and Reichlin [2006]). Overall, previous studies seem to suggest that not accounting for the heterogeneity across member countries does not impair the comparison of conditional moments across different periods of time.

The DSGE model we estimate captures the salient features of the macroeconomic time series of the euro area (see Smets and Wouters [2003]). Monetary policy has real effects in the short run because of nominal frictions in wages and prices. The main channel through which it influences the economy is the interest rate channel; price and wage rigidities imply that changes in the nominal interest rate affect the real interest rate on which the decisions on the intertemporal allocation of consumption are based. The euro area is modelled as a closed economy. While acknowledging the importance of the openness dimension of the euro area, we believe it is not a strong assumption to neglect it, taking into account also that the MTN found that the exchange rate channel was not playing an important role at the area-wide level.[2] The model incorporates price and nominal wage rigidities, but it does not include hiring and firing costs in light of what found by Christoffel et al. [2009] on the irrelevance of those labor market frictions to explain the MTM. Finally, we do not include financial frictions in the model for several reasons: first, we want to keep the model simple; second, Gerali et al. [2010] estimate a

[2] The euro area is very similar in size and in the degree of openness to the U.S., which is commonly described as a closed economy. In 2008 the total trade to GDP ratio was around 40% in the euro area and 30% in the U.S.

medium-scale model which incorporates financial frictions and show that these, together with shocks hitting credit markets, played a particularly relevant role only during the latest recession that followed the 2007-08 crisis. Moreover, the financial crisis itself could have determined a change in the private sector behavior and in the conduct of monetary policy. Unfortunately, such hypothesis cannot be tested since few data are available at present.

The empirical evidence from the VAR analysis suggests that the MTM of the euro area has not changed significantly during the last ten years. If anything, monetary policy has become slightly more effective in stabilizing the economy.[3] The results are somewhat different from those obtained by estimating the medium-scale DSGE model. In this case, differences across the two samples emerge more clearly and are due to a reduction in the degree of nominal rigidities and to an increase in the strength of the systematic reaction of monetary policy to inflation. Counterfactual exercises show that changes in the responses of output and prices to monetary and cost-push shocks are mostly explained by a variation in private sector parameters, while the changes in the responses to technology shocks are due to the monetary policy conduct. The observed decline in the volatility of output growth, inflation and the policy rate across the two sub-samples can only be partially explained by a more favourable set of shocks during the EMU period. Furthermore, while the decline in the volatility of inflation is mostly attributable to a change in the monetary policy conduct, the volatility of output growth is explained by changes in private sector parameters.

The remainder of the paper is organized as follows. Section 2 describes the VAR approach, the identification schemes of monetary policy shocks and the results. Section 3 presents the results of the estimation of a medium-scale DSGE model over the pre-1999 and post-1999 periods. Section 4 illustrates the possible explanations for the changes in the MTM using some counterfactual simulations with the estimated model. Section 5 offers some concluding remarks.

2 The VAR approach

In this section we study the transmission mechanism of monetary policy and the possible changes that might have occurred after 1999 using a VAR approach. The VAR model has

[3] By effectiveness of the monetary policy, in this exercise, we mean that both output and prices are more responsive to an exogenous change of the nominal interest rate.

the following representation:

$$y_t = \sum_{\ell=1}^{p} B(\ell) y_{t-\ell} + C x_t + \varepsilon_t \tag{1}$$

where y_t for $t = 1, ..., T$ is a $K \times 1$ vector of endogenous variables, x_t is a $Q \times 1$ vector of exogenous or deterministic variables, ε_t is a $K \times 1$ vector of errors, p is the number of lags and $B(\ell)$ and C, with ℓ being the lag operator, are $K \times K$ and $K \times Q$ matrix of coefficients. We assume ε_t to be independent and identically normally distributed with mean equal to zero and covariance matrix Σ.

All the VARs are estimated with data in levels, so that our results do not depend on some arbitrary data transformation. We have collected data for the euro area economy both at monthly and quarterly frequency.[4] Quarterly data are used to assess the robustness of the results to the frequency of the data and for comparability with those obtained with the estimation of the theoretical model in Section 3. The monthly data include observations from 1994:M1 to 2009:M9 of the following variables: industrial production, our measure of economic activity, the harmonised index of consumer prices (HICP), the overnight interest rate (EONIA), the M2 monetary aggregate, commodities prices and the nominal effective exchange rate. The quarterly data refer to the period 1989:Q1-2009:Q2. Economic activity is measured with real GDP, the price level with the GDP deflator and the short-term nominal interest rate with the 1-month money market rate.[5]

When the number of parameters to estimate is large given the sample information, unrestricted VAR tends to overfit the data. In order to avoid this we resort to Bayesian methods and we combine a priori information with the likelihood function of the data. We define $\alpha = [vec(A)\ vec(C)]'$ a vector of size $(Kp+Q)K$ where we stack the coefficients in $A(\ell)$ and C (p is the number of lags). We choose a normal prior for the coefficients in α and a diffuse one for the variance-covariance matrix of the shocks Σ:

$$\alpha \sim N\left(\bar{\alpha}, \bar{\Sigma}^{\alpha}\right) \tag{2}$$

$$p(\Sigma) \sim |\Sigma|^{-(K+1)/2} \tag{3}$$

where $\bar{\alpha}$ denotes the mean of the prior and $\bar{\Sigma}^{\alpha}$ its variance covariance matrix. We impose the restrictions of the so called Minnesota prior (see Litterman [1986]) on the coefficients

[4] For a detailed description of the data see Appendix A.

[5] The euro area overnight interest rate is not available before 1994. We use the one-month Euribor up to 1999:Q2 and the one-month Eurepo afterwards.

in α (Doan et al. [1984]). This implies that a priori we represent the series included in the VAR as univariate random walks with correlated innovations. All coefficients in $\bar{\alpha}$ are equal to zero except the first own lag of the dependent variable in each equation, which is set to one. Moreover it is assumed that the prior covariance matrix $\bar{\Sigma}^{\alpha}$ is diagonal and that the $\sigma^{\alpha}_{ij,\ell}$ element, corresponding to lag ℓ of variable j in equation i, is equal to

$$\bar{\sigma}^{\alpha}_{ij,\ell} = \begin{cases} \frac{\phi_0}{h(\ell)} & \text{if i = j, } \forall \ell \\ \phi_0 \frac{\phi_1}{h(\ell)} \left(\frac{\sigma_j}{\sigma_i}\right)^2 & \text{if i} \neq \text{j, } \forall \ell \text{, j endogenous} \\ \phi_0 \phi_2 & \text{if j exogenous/deterministic} \end{cases}$$

The hyperparameter ϕ_0 represents the overall tightness of the prior; ϕ_1 the relative tightness of other variables, ϕ_2 the relative tightness of the exogenous variables and $h(\ell)$ the relative tightness of the variance of lags other than the first one (we assume throughout that $h(\ell) = \ell$, that is a linear decay function). The term $(\sigma_j / \sigma_i)^2$ is a scaling factor that accounts for the different scale of the variables of the model. We set $\phi_0 = 0.1$, $\phi_1 = 0.5$ and $\phi_2 = 10^5$ in our benchmark specification (see Canova [2007]), but we perform some robustness exercises on the relevance of the prior tightness to the results. Rewriting the VAR in (1) in companion form as $Y = ZA + u$, the posterior distribution is Normal-Wishart:

$$\alpha|\Sigma, Y \sim N\left(\underline{\alpha}, \left[(\underline{\Sigma}^{\alpha})^{-1} + \Sigma^{-1} \otimes Z'Z\right]^{-1}\right) \tag{4}$$

$$\Sigma^{-1}|\alpha, Y \sim W\left(\left[(Y - Z\hat{A})'(Y - Z\hat{A}) + (A - \hat{A})'Z'Z(A - \hat{A})\right]^{-1}, T\right) \tag{5}$$

where $\underline{\alpha}$ and $\underline{\Sigma}^{\alpha}$ are the mean and covariance matrix of the posterior distribution and $\hat{A}$ is the OLS estimate of the companion matrix A. We draw α and Σ from the posterior using the Gibbs sampling algorithm.

2.1 Identification of monetary policy shock

Isolating exogenous variations in the stance of monetary policy is a difficult task and yet a crucial one as the results on the monetary transmission mechanism may be sensitive to the assumptions for the shock identification. The coefficients of the structural equations below (abstracting for simplicity from the exogenous variables x_t) can be recovered from the estimated reduced form (1)) by imposing enough restrictions on the matrix A_0

$$A_0 y_t = \sum_{\ell=1}^{p} A(\ell) y_{t-\ell} + v_t \tag{6}$$

where v_t are the structural shocks with covariance matrix equal to the identity one. In order to find a set of results that are fairly robust, we proceed following three different strategies for identifying the shock.

The first identification scheme we use is a recursive one (see, among others, Christiano et al. [1999]). We decompose the variance-covariance matrix of the reduced form residuals Σ using a Cholesky factorization. The ordering of the variables is the following: commodity prices (cp), the price level (p), industrial production (y), the EONIA rate (R), the M2 monetary aggregate ($M2$) and the nominal effective exchange rate of the euro (e). Both commodity prices and the exchange rate are used to control for foreign inflationary pressures and to capture the open economy dimension of the euro area. We also consider a recursive identification scheme in a VAR in which both commodity prices and the exchange rate are treated as exogenous variables (i.e. they are included in x_t and not in y_t in equation (1)).

The second identification strategy follows Sims and Zha [1999] and Kim [1999] and assumes that because of information delays monetary policy cannot respond within the month to prices and industrial production. At the same time, we assume that the monetary policy authority observes and reacts to commodity prices, money and the exchange rate. The restrictions of this identification scheme define a money demand and money supply equation; the monetary policy shock influences output and prices only with a lag, while money and the exchange rate are affected contemporaneously. Money demand depends on prices, output and the nominal interest rate. The innovation to commodity prices affects contemporaneously the nominal effective exchange rate. The nominal exchange rate, as an asset price, reacts to all variables in the system. The shocks are exactly identified since this allows to compute probability error bands for the impulse responses using standard Monte Carlo methods.[6] The following matrix:

[6] If our assumptions had implied an overidentified VAR, then standard methods for conducting inference could have not be used. In this case the correct methodology is outlined in Sims and Zha [1998] and Sims and Zha [1999].

$$A_0 = \begin{bmatrix} a_{1,1} & 0 & 0 & 0 & 0 & a_{1,6} \\ a_{2,1} & a_{2,2} & a_{2,3} & 0 & 0 & 0 \\ 0 & a_{3,2} & a_{3,3} & 0 & 0 & 0 \\ a_{4,1} & 0 & 0 & a_{4,4} & a_{4,5} & a_{4,6} \\ 0 & a_{5,2} & a_{5,3} & a_{5,4} & a_{5,5} & 0 \\ a_{6,1} & a_{6,2} & a_{6,3} & a_{6,4} & a_{6,5} & a_{6,6} \end{bmatrix} \begin{bmatrix} cp \\ p \\ y \\ R \\ M2 \\ e \end{bmatrix}$$

summarises our structural identification scheme and allows recovering the structural representation of the VAR (eq. 6) from the reduced form (eq. 1).

The last identification strategy implemented is sign restrictions (see Canova and De Nicolo [2002], Uhlig [2005] and Dedola and Neri [2007]). We impose that prices, output and money respond negatively to a positive monetary policy shock while the interest rate increases. This set of restrictions is imposed only on impact, leaving unrestricted the dynamics of the variables from the second step of the impulse horizon onwards.

2.2 The effects of monetary policy shocks before and after the creation of the EMU

In this section we present and discuss the impulse responses to a monetary policy shock estimated from different VARs under the identification schemes described above. We are interested in documenting the possible changes in the MTM that may have occurred after the creation of the European Monetary Union. To this end we split the sample at 1999:M1 (and 1999:Q1 for quarterly data), the time at which the euro was adopted. Since the econometric methodologies are generally weak in identifying the exact date in which a structural break occurs when the sample is short, as it is in our case, we do not search for it in the data (as it is done by Weber et al. [2011] and Ciccarelli and Rebucci [2006]), but rather choose one a priori. We choose the adoption of the common currency in 1999 as the break date since we are mostly interested in investigating whether there have been changes in the MTM associated with the creation of the EMU.

Several studies have shown that the convergence process across euro area member countries occurred before the adoption of the common currency. Using longer time-series for the main European economies Ciccarelli and Rebucci [2006] found evidence of a structural break in 1991 around the German unification, while Weber et al. [2011] found a break in 1996 and evidence for an other one in 1999. Our data restrict our choices, since splitting the sample before 1999 would imply that the number of data in the pre EMU sample is rather short. Therefore, our pre EMU sample goes from 1994 (in the monthly data exercise) and from 1989 (in the quarterly data exercise) to 1998. We consider then two post EMU samples: the first one spans the period before the financial turmoil, from 1999 to (July or Q3) 2007, while the second one (1999-2009) includes it. This allows us to draw some preliminary considerations on whether the financial crisis and the severe recession of 2008 brought about any visible changes to the monetary transmission mechanism in the euro area. We present results for VARs with the number of lags p equal to 4 in the monthly VAR and to 3 in the quarterly VAR; however, they are robust to different lag specifications.

Overall, the responses of prices and real activity are similar across sample periods and they are in line with the stylized facts on monetary transmission mechanism (see Figure 1). An unexpected rise in the short-term interest rate is followed by a temporary fall in output and a more sluggish and persistent decline in the price level. These results are robust across all identification schemes. The picture obtained from the monthly VAR is confirmed when using quarterly data (see Figure 2)

Looking at the median responses, there is evidence of a small change. We notice that in the EMU sample the decrease of real activity is more pronounced and the price level drops more strongly. In the monthly VAR with Choleski identification, the peak response of industrial production is -1.73 per cent in the pre EMU sample and -1.90 in the post EMU one, while that of HICP prices is -0.23 in the pre EMU sample and -0.29 in the post EMU one. In the quarterly VAR the corresponding figures are - 0.51 and -0.55 for real GDP, in the pre and post EMU samples respectively, and -0.22 and -0.27 for the GDP deflator. However, the uncertainty around the median estimate is high (Figure 3 reports the median of the impulse responses of HICP prices and industrial production of the monthly VAR with Cholesky identification together with the confidence bands). Based on draws from the posterior distribution, the probabilities that the response of output after 18 months is stronger in the post 1999 period and that that of prices larger two years after the shock are both around 60 per cent. These results are somewhat different from those of Boivin et al. [2008], though their econometric methodology and the identification procedure is different and their pre EMU sample include also the last part of the eighties. Weber et al. [2011] find instead that no major changes have occurred in the effects of monetary policy.

The Bayesian approach implies that the posterior distribution on which our results are based comes from the recursive update of the data with the prior information. One may thus want to check how much results are driven by the imposed prior. In the Minnesota restrictions, the hyperparameter ϕ_0 controls the relative weight of sample and prior information (the smaller is ϕ_0, the tighter is the prior and the higher is its importance on the results). Figure 4 displays the impulse responses to a monetary policy shock (identified by Cholesky factorization) under different assumptions for ϕ_0 in the pre and post EMU sample respectively. While a tighter prior smoothes the impulse responses, overall the results are not strongly driven by the prior information.[7]

Concerning the implications of the recent financial crisis for the MTM, comparing the responses of output and prices in the two post EMU samples, there are no significant differences before and after the burst of the turmoil. If anything we observe a larger decline of output when we include the data for 2007-2009. This result is in line with what found in a similar paper by Giannone et al. [2009], in which they found no evidence on a changes in the VAR coefficients after 2008. There is a chance, however, that the few data available are not able to capture the structural changes brought about by the recent crisis. A possible break of the MTM after 2008 could be observed in a VAR only when more data of the new "regime" will be available.

To sum up, the conclusions that we draw from the VAR analysis are that, accounting for the strong uncertainty surrounding the impulse responses, there have been only minor changes in the effects of monetary policy on output and prices over the last 10 years. This VAR evidence, however, cannot tell us the sources of such changes, as modifications of the private sector behaviour cannot be separately identified from changes in the conduct of monetary policy. The estimation of a structural DSGE model in which the various channels at work can be disentangled could be more informative in this respect. The next Section takes the DSGE model to the data and dig deeper into this issue.

3 The DSGE approach

While it is difficult to interpret the impulse responses estimated from a VAR, the estimation of a more structural model can indicate whether there have been offsetting forces that resulted in only minor changes of the monetary policy transmission mechanism as elicited from the VAR or there have been no changes at all.

We illustrate further this point by considering a small-scale DSGE model as the

[7] Furthermore, the prior assumption of modelling the series included in the VAR as random walks seems justified as unit root tests suggest.

data generating process of the time series of inflation, output and nominal interest rate.[8] We simulate the time series of the relevant macroeconomic variables under a baseline calibration of the model. On the simulated data we estimate a VAR and the implied impulse responses to a monetary policy shock, identified through Cholesky factorization. After changing the calibration of the parameters of the reaction of the monetary policy rule to inflation and output gap in the Taylor rule, we generate the new data and estimate the same VAR. We minimize the distance between the VAR impulse response functions obtained with data coming from the DSGE with the baseline calibration and those obtained from the data simulated with the new calibration of the policy rule over the parameter for the slope of the Phillips curve. As shown in figure 5, we found a pretty good match of the impulse responses for reasonable parameterizations of the model. This suggests that, due to the fact that the impulse response functions are a non-linear combination of structural parameters, differences in those parameters may give rise to almost identical impulse responses as estimated from a VAR. This result questions the reliability of a study on the changes in the monetary transmission mechanism based only on the comparison of the impulse responses from a VAR and convinces us to investigate further the MTM by means of a more structural model which fits reasonably well the macroeconomic time series of the Euro area.

In this section we thus estimate a medium-scale DSGE model (Smets and Wouters, 2003 and 2007) for the euro area in the pre and post EMU samples using Bayesian methods. Beyond checking the empirical evidence based on the VAR models and analyze the sources of the observed developments in the MTM, this allows us to perform some counterfactual exercises with more confidence. In fact, recent works on structural VAR convincingly show that, on the one hand, it is impossible to separate the effects of changes in the policy rule and in the variance of the shocks with structural VAR models (see Benati and Surico [2009]) and that, on the other hand, counterfactuals based on SVARs are unreliable, independently of the issue of parameters identification (see Benati [2009]).

[8] The model is a three-equation basic New Keynesian model in which inflation and output gap depend both on a backward and forward-looking term and monetary policy is specified by a Taylor rule in which the nominal interest rate responds to its lagged value and to current inflation and output gap.

3.1 Data and methodology

In order to estimate the model, we use quarterly data for the period 1989:1-2007:2[9] and match the following seven variables: GDP-deflator based inflation, nominal hourly wage inflation, real consumption, real investment, real GDP, employment (matching total hours in the model) and the three-month nominal interest rate. We use linearly detrended data for consumption, investment, GDP and employment and deviations from their respective means for inflation, the interest rate and wage inflation. The linear trends are estimated over the full sample. For a description of the data see Appendix A.

Bayesian methods combine information from the prior distribution of the structural parameters with that contained in the likelihood function of the model. The resulting posterior distribution of the parameters usually does not belong to any standard family and therefore the inference must be based on simulation methods. It has become common practice to use the Metropolis algorithm to generate draws from the posterior distribution. We proceed in two steps. First we maximize the log of the posterior density and compute an approximation of the inverse of the Hessian at the mode. Second, we generate 200,000 draws from the posterior distribution of the parameters using a multivariate normal with covariance matrix proportional to the inverse of the Hessian.[10]

3.2 The model

We estimate a medium scale DSGE model which has been shown to fit reasonably well the macroeconomic time series of the euro area (see Smets and Wouters [2003]). The model features monopolistic competition in product and labour markets as well as nominal rigidities in prices and wages that allow for backward inflation indexation. Various other features such as habit formation, costs of adjustment in capital accumulation and variable capacity utilization are introduced in order to match the data. The main channel through which it influences the economy is the interest rate channel (price and wage rigidities imply that changes in the nominal interest rate affect the real interest rate on which are based the decisions on the intertemporal allocation of consumption of the agents). For the reasons illustrated in the introduction the model disregards a role for the exchange rate

[9] As mentioned in the introduction, our model is simplified in several dimensions; as a consequence it would not capture adequately the macroeconomic developments of the recent financial crisis. For this reason, the estimates are carried out in a sample period ending in the second quarter of 2007.

[10] The estimation is done with Dynare 4.0. The scale factor for the jump distribution has been set in order to obtain acceptance rates around 30 per cent.

and the bank lending channel for the transmission of monetary policy shock.

Our reference model is a slightly simplified version of the one in Smets and Wouters (2007, henceforth SW). We assume separability between consumption and leisure (as in SW, 2003) and we use the standard Dixit-Stiglitz aggregator for prices and wages instead of the Kimball aggregator of Smets and Wouters [2007], set to zero the share of fixed cost in the production function and finally we assume no steady state growth for the economy.

We also modify the interest rate rule followed by the central bank as follows

$$\hat{R}_t = \rho \hat{R}_{t-1} + (1-\rho)\left[\rho_\pi \hat{\pi}_t + \rho_y \hat{y}_t^{GDP}\right] + \epsilon_t \tag{7}$$

where $\hat{y}_t^{GDP}$ is the weighted sum (with weights equal to the steady state shares) of real consumption, real investment and real government spending.

The model has been simplified in order to reduce its parameter space as the length of our time series is limited. For the complete set of equations see Appendix B.

3.3 Prior and posterior distributions

Some of the parameters of the model are calibrated (see Table 1). We set the households' discount factor at 0.995, in order to obtain a steady-state real short-term interest rate of 2 per cent on an annual basis, in line with the historical average for the euro area. The capital share, the depreciation rate and the share of government spending over output are set at 0.25, 0.025 and 0.15. These numbers imply a steady state ratio of consumption to the sum of consumption, investment and government spending of 55 per cent, consistently with the average over the period 1995-2009. The same figure for the share of investment is 22 per cent. Both shares are similar to the values used in Christoffel et al. [2009]. The share of fixed cost in production is set to zero. Allowing for these costs does not affect the shape and magnitude of the impulse responses. The adjustment cost for capacity utilization is set to 0.1. The parameter measuring the mark-up in wage setting is set at 1.5 as in Smets and Wouters [2003] while the inverse of the labour supply elasticity is calibrated at 1.5, in line with the range of available estimates.

The specification and parameterization of the prior distributions are equal across subsamples and reported in Table 2. All the distributions are fairly loose. The mean of the autoregressive coefficient of the shock processes is set at 0.80. The (beta) distribution of the Calvo probabilities for prices (ξ_p) and wages (ξ_w) have a mean of 0.75 which corresponds to an average duration of one year. The means of the beta distribution of the parameters measuring the indexation of prices (ι_p) and wages (ι_p) to past inflation are set 0.50 with a standard deviation of 0.20. The mean of the (beta) distribution of the parameter measuring the degree of habits in consumption (γ) is set at 0.50 in line with Smets and Wouters [2003]. The parameter measuring the risk aversion (σ_c) has a mean of 1.5 while the cost for adjusting investment (φ) has a mean of 5.0, in line with the prior in Smets and Wouters [2003]. For what concerns the prior distribution of the policy parameters we set the mean of the coefficients of the response of past interest rate, current inflation and output respectively equal to 0.75, 1.5 and 0.

3.4 Estimation results and impulse responses

Table 2 reports the summary statistics of the posterior distribution of the model parameters for the two sample periods. The results are based on 250,000 draws generated with the random walk version of the Metropolis algorithm. Several results are worth a comment. First, there is a significant reduction in the Calvo parameters for prices (ξ_p) and nominal wages (ξ_w); the former declines from 0.88 in the pre EMU sample to 0.75 in the post EMU sample while the latter from 0.90 to 0.82. Both findings indicate a decrease in the degree of nominal rigidities in the euro area. Second, the degree of indexation of nominal wage contracts to inflation (ι_w) falls significantly in the post 1999 sample to 0.29 compared with 0.54 in the period before the creation of the EMU. Third, we document an increase in the response of monetary policy to inflation (ρ_π rises from 1.39 to 2.14) and at the same time a decline in the coefficient on output (ρ_y falls from 0.44 to 0.11). These results suggest that the ECB is more focused on inflation stabilization than the joint set of central banks of the countries that have become member of the euro area in 1999. Other, less important findings concern the parameter measuring the cost for adjusting investment (φ), which increase in the post EMU sample and the degree of habit formation in consumption (λ) which also increases. The coefficient of risk aversion (σ_c) is stable across the two periods. The standard deviation of all the structural shocks falls in the post 1999 period. This finding together with the changes in the structural parameters and the policy rule call for a deeper analysis of the role played by these factors in generating the fall in the volatility of real GDP and inflation in the euro area (the first falls from 1.45 to 1.08 per cent and the second from 0.38 to 0.20). We address this issue in section 4.1.

Figure 6 and 7 report the prior and posterior densities of the parameter measuring the degree of nominal rigidities and those of the policy rule. As shown by Figure 6, the post EMU period is characterized by a significantly lower degree of inflation indexation of nominal wages. Figure 7 shows that the post EMU sample is characterized by a policy that responds less to output and more to inflation compared to the one that was in place, in the aggregate of the euro area, before 1999.[11]

To study the differences in the transmission mechanism of monetary policy across the two subperiods we plot the responses to a one per cent standard deviation shock to the interest rate rule after four quarters. Figure 8 (panel (a)) reports the draws of the posterior distribution of the impulse responses of output and inflation at the one year horizon together with the 45 degree lines. In absolute term the response of output after one year is slightly higher in the pre EMU sample. In fact most of the draws are concentrated above the 45 degree line. In the case of inflation, almost all draws from the posterior lies below the 45 degree line. This indicates that the inflation response is stronger in the post EMU sample and confirms what already suggested by the VAR evidence. With the whole posterior distribution of the impulse responses we can compute the probability that the responses of output and inflation are larger in the post EMU period than in the pre one. With respect to output, this probability is equal to 0.16 at the one year horizon and it increases up to 0.36 after three years. Concerning inflation, these probabilities are equal to 0.99 and decreases sharply to 0.09 after 12 quarters, since at that horizon the effect of a monetary policy shock on inflation has died out in both sample. Figure 8 (panel (b)) proposes the same exercise as before but in response to a positive technology shock. In this case the differences across subperiods emerge more sharply. The response of output after 1 year is almost always stronger and positive in the post EMU sample while slightly negative in the pre EMU sample. The response of inflation is closer to zero in the post EMU sample and negative in the pre EMU one.

The impulse responses to a monetary policy shock look different from those obtained in the VAR exercise in section 2. In order to square the evidence coming from the two methodologies, we take the point of view of the econometrician and we estimate a basic VAR using the estimated DSGE model as data generating process. The estimates from the VAR, based on actual data, may suffer from a small sample bias due to lack of longer time series. Our interest, however, is not in quantifying this small sample bias, that would apply also to the DSGE estimation, but in the extent to which the responses of output and inflation to a monetary policy shock obtained from a simple trivariate VAR are able to

[11] A dimension over which the monetary policy conduct differs across the two samples is the degree of forward-lookingness of the central bank reaction to inflation. Our benchmark Taylor rule assumes that in both periods the interest rate responds to the current level of inflation, while one could argue that the ECB responds mainly to expected inflation over a medium run horizons. However, even when we allow for the possibility of a response to future inflation, the data prefer the benchmark specification.

show the differences in some crucial parameters, across the two subperiods, revealed by the structural estimation of the model. We therefore simulate long time series from the DSGE setting the parameters at the median of the posterior distributions in the pre and post EMU samples, therefore taking into account only the uncertainty of the shocks. Figure 9 plots the impulse responses from the VAR(1) on inflation, output and the nominal interest rate to a monetary policy shock identified through Cholesky factorization. We find that in this basic VAR the response of inflation is slightly stronger after 1999, as evidenced in the DSGE estimation, while there are no perceptible differences in the responses of output and the interest rate.[12] The exercise shows that the simple VAR has a hard time in detecting the differences in the transmission mechanism of monetary policy and the in the magnitudes of the responses of output and inflation.

4 Explaining the changes in the MTM

Having estimated the model we proceed with a counterfactual analysis which aims at providing an explanation for the changes in the transmission mechanism of monetary policy and, more generally, for its effectiveness in terms of output and inflation stabilization. Following Boivin and Giannoni [2006] we characterize the behaviour of monetary policy by the set of monetary policy parameters ρ, ρ_π and ρ_y and the behaviour of the private sector by all other parameters.

Figure 10 plots the impulse responses to a monetary policy shock that raises the short-term interest rate by one standard deviation. Each panel contains the impulse responses for the four possible combinations of monetary policy (MP) and private sector (PS) parameters for the two sub-samples, 1989:1 - 1998:4 (pre EMU) and 1999:1 - 2007:2 (post EMU). The figure clearly shows that the observed change in the monetary transmission mechanism is explained both by a change in the systematic conduct of monetary policy and by the private sector behavior. In order to highlight the contribution of monetary policy, we compare the responses obtained setting the private sector parameters at the median of their marginal posterior distributions for the pre EMU sample and those of the policy rule at the median of the posterior of the two sample periods. By comparing the black solid line and the blue dashed lines one can see that the change in the behaviour of monetary policy has increased its effectiveness in stabilizing both output and inflation.

[12]Note that these results are not directly comparable with those of the VARs in section 2 since in that case data were in levels and the VAR were richer including also the commodity prices, the effective nominal exchange rate and the money M2.

The role played by the changes in the structural parameters of the private sector can be gauged by fixing the policy parameters at the median of the posterior of the pre EMU and compute the impulse responses by varying the other parameters. By comparing the black solid lines and the red solid lines with dots one can see that the changes in the degree of nominal rigidities has made monetary policy more effective in controlling inflation. Inflation responds more sharply on impact in the post 1999 sample compared to the pre 1999 one. To sum up, changes in both the behaviour of the central bank and the private sector have contributed to modify the transmission mechanism of monetary policy in the direction of increasing the effectiveness in stabilizing inflation around its target.

Whether a modification of the monetary policy regime had effect on the transmission mechanism of the euro area economy should emerge not only from the analysis of the responses to an unexpected change of the policy rate, but also from that of the responses to other shocks. In fact, monetary policy influences macroeconomic variables mostly by reacting systematically to all shocks that hit the economy. Therefore, we perform the same counterfactual exercise as before, analyzing the impulse responses functions to a transitory technology shock and to a shock in the price markup. In the first case (see Figure 11), the responses of output, inflation and prices are weaker in the post EMU sample, while the impulse responses of the short-term nominal interest rate is similar across sample periods. The changes in the responses of output and prices are attributable almost entirely to a change in the systematic conduct of monetary policy. Indeed, maintaining the private sector parameters constant at the pre EMU level, a change in policy from the pre EMU rule to the post EMU one explains almost all the changes of the price level and output responses across the two subsamples.

In response to a positive shock to the price mark-up (see Figure 12)[13], the price level is less reactive in the post 1999 sample, while there are no major changes in the response of real activity. Most of the changes of the prices impulse responses across periods are due to changes in the private sector parameters.

4.1 Implications for the volatility of output and inflation

As already mentioned in section 3.4 and documented by Canova et al. [2009], there has been a drop in the volatility of the main macroeconomic variables after 1999 (see Table 3, Panel A). The standard deviation of the real GDP declines from 1.45 to 1.08 per cent,

[13] An unexpected increase of the mark-up in the product market increases both prices and output and could be interpreted for instance as a change in the oil prices.

the one of the GDP deflator inflation from 0.38 to 0.20 per cent and the one of the short-term nominal interest rate from 0.67 to 0.22 per cent. In this section we analyze this fall in volatility. In the same spirit as the analysis of the Great Moderation we want to uncover the origins of the generalized decline in the volatility of the economy and see whether it has been "good policy" or "good luck". As we have done for the conditional moments of our estimated model, we use a set of counterfactual experiments that allow us to disentangle the effects on the unconditional moments due to changes in the volatility of the structural shocks from those related to changes in the structure of the economy and in monetary policy. The panel B of Table 3 reports the results of the counterfactual experiments for alternative policy rules, structural parameters and shock processes.

The model replicates the fact that the volatility of output, inflation and the nominal interest rate is lower in the post EMU sample. Only a fraction of the decline in these volatilities is due to a more favourable set of shocks (compare the first and the last line across the left-hand and right-hand side columns in Table 3, panel B). If the monetary policy rule of the EMU period were in place before 1999, the volatility of inflation would have been lower while that of output would have been higher (compare the first two lines of the left-hand columns). Moreover, the volatility of inflation increases when we adopt, in the EMU sample, the pre EMU monetary policy rule. It is thus fair to conclude that the changes in the behaviour of monetary policy are behind most of the decline in the volatility of inflation. Doing the same exercise, we notice instead that output stabilization is mainly due to the changes in private sector parameters (first line vs third line in the left-hand columns). When switching from a pre EMU to the more recent policy rule, independently of the shocks and the private sector parameters, the inflation volatility decreases together with a decline or a substantial stability of the nominal interest rate volatility (compare the first and third rows to the second and last ones in all columns). This suggests that a stronger inflation stabilization is achieved not through a stronger reaction of the nominal interest rate but through the steering of expectations. Overall, the story that emerges is a more interesting one, compared to an all-shocks or an all-policy one explanation for the decline in output and inflation volatility.

5 Conclusions

The creation of the EMU and the establishment of the ECB with a clear-cut mandate for maintaining price stability might have contributed to changing the transmission mechanism of monetary policy impulses. The paper provides a quantitative assessment of the changes in the MTM of the euro area economy based on a structural VAR and a dynamic general equilibrium model. According to our VAR approach there emerge no significant differences in the impulse responses of prices and output to a monetary policy tightening shock before and after the setup of EMU. We argue that, while the VAR methodology is useful to compare our results with those obtained in the early years of the common currency by the Monetary Transmission Network, the analysis of the MTM based solely on VARs is not very informative on the effective changes occurred to the economy. The estimated responses to monetary impulses cannot always detect variations to the MTM, as different changes in the economy might have offsetting effects on the impulse responses. A more structural model, based on stronger assumptions, is thus needed to complement the VAR results.

The estimation of a closed-economy model for the Euro area, similar to Smets and Wouters [2003, 2007], suggests that after 1999 the nominal rigidities became weaker while the coefficient on inflation in the monetary policy rule increased and that on output declined. Counterfactual analysis indicate that changes in the private sector parameters are responsible of the stronger reaction of output and inflation to a monetary policy shock and the milder reaction of prices after a cost-push shock in the post 1999 period, while the modification of the monetary policy conduct influenced the responses of both output and prices to a technology shock. The drop in macroeconomic volatility observed across the two periods is only marginally attributable to a more favorable set of shocks in the EMU sample; the one on inflation is due mostly to changes in the monetary policy rule parameters while that on output to changes in the private sector behavior.

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Appendices

A Data and sources

Monthly data

Eonia rate: Eonia, ECB (FM.M.U2.EUR.4F.MM.EONIA.HSTA).

Nominal effective exchange rate: ECB EER-12 group of currencies, changing composition of the euro area against Euro, ECB (EXR.M.Z08.EUR.EN00.A).

HICP: Harmonised Index of Consumer Prices, overall index, seasonally adjusted, not working day adjusted, ECB (ICP.M.U2.S.000000.3.INX).

Industrial production: Total Industry excluding construction and MIG Energy - NACE Rev2, working day and seasonally adjusted, ECB (STS.M.I5.Y.PROD.NS0021.4.000).

M2 monetary aggregate: Index of Notional Stocks, MFIs, central government and post office giro institutions reporting sector (changing composition), working day and seasonally adjusted, ECB (BSI.M.U2.Y.V.M20.X.I.U2.2300.Z01.E).

Commodity prices: Index of Fuel and Non Fuel Commodities (2005=100), IMF.

Unemployment rate: Standardized unemployment rate, total (all ages), Total (male & female), seasonally adjusted, not working day adjusted, percentage of civilian workforce, Eurostat, (STS.M.U2.S.UNEH.RTT000.4.000).

Quarterly data

Real consumption: Final consumption of households and NPISH's, constant prices, euro
area 15 (fixed composition), seasonally adjusted, not working day adjusted, ECB
(ESA.Q.I4.S.1415.P31000.0000.TTTT.Q.U.A).

Real investment: Gross fixed capital formation, constant prices, euro area 15
(fixed composition), seasonally adjusted, not working day adjusted, ECB
(ESA.Q.I4.S.1000.P51000.0000.TTTT.Q.U.A).

Real GDP: Gross domestic product at market price, Chain linked volumes, reference year 1995, seasonally and partly working day adjusted, mixed method of adjustment, ECB (ESA.Q.I4.S.1000.P51000.0000.TTTT.Q.U.A).

GDP deflator: Gross domestic product at market price deflator (ECB compilation) seasonally and partly working day adjusted, mixed method of adjustment, ECB (ESA.Q.I4.S.0000.B1QG00.1000.TTTT.D.U.I).

Real wages: Wage per head (WRN), Area Wide Model database, deflated with the GDP deflator.

Employment: Total employees, persons (LEN), Area Wide Model database.

B The model equations

The log linearized model has 17 endogenous variables:

$$\left\{ y_t \ c_t \ i_t \ z_t \ r_t \ q_t \ k_t \ \mu_t^p \ l_t \ k_t^s \ w_t \ \mu_t^w \ mpk_t \ \pi_t \ \pi_t^w \ r_t^n \ y_t^{gdp} \right\}$$

and 7 exogenous variables:

$$\left\{ \varepsilon_t^g \ \varepsilon_t^a \ \varepsilon_t^b \ \varepsilon_t^i \ \varepsilon_t^p \ \varepsilon_t^w \ \varepsilon_t^r \right\}$$

$$
\begin{aligned}
y_t &= c_y c_t + i_y i_t + z_y z_t + \varepsilon_t^g && (8)\\
c_t &= \frac{\lambda}{1+\lambda} c_{t-1} + \frac{1}{1+\lambda} E_t\{c_{t+1}\} - \frac{1-\lambda}{\sigma_c(1+\lambda)} \left(r_t + \varepsilon_t^b \right) && (9)\\
i_t &= \frac{1}{1+\beta} i_{t-1} + \frac{\beta}{1+\beta} E_t\{i_{t+1}\} + \frac{1}{1+\beta}\frac{1}{\varphi} q_t + \varepsilon_t^i && (10)\\
k_t &= \delta i_t + (1-\delta) k_{t-1} + \delta(1+\beta)\varphi \varepsilon_t^i && (11)\\
\mu_t^p &= \alpha(k_t^s - l_t) - w_t + \varepsilon_t^a && (12)\\
\mu_t^w &= w_t - \sigma_l l_t - \frac{\sigma_c}{1-\lambda}(c_t - \lambda c_{t-1}) && (13)\\
y_t &= \alpha k_t^s + (1-\alpha) l_t + \varepsilon_t^a && (14)\\
k_t^s &= k_{t-1} + z_t && (15)\\
mpk_t &= -(k_t^s - l_t) + w_t && (16)\\
z_t &= \frac{1-\psi}{\psi} mpk_t && (17)\\
\pi_t &= \frac{\iota_p}{1+\beta}\pi_{t-1} + \frac{\beta}{1+\beta} E_t\{\pi_{t+1}\} - \frac{1}{1+\beta}\frac{(1-\xi_p)(1-\beta\xi_p)}{\xi_p}\mu_t^p + \varepsilon_t^p && (18)\\
w_t &= \frac{1}{1+\beta} w_{t-1} + \frac{\beta}{1+\beta}\left(E_t\{w_{t+1}\} + E_t\{\pi_{t+1}\}\right) - \frac{1+\iota_w\beta}{1+\beta}\pi_t + && (19)\\
&\quad + \frac{\iota_w}{1+\beta}\pi_{t-1} - \frac{1}{1+\beta}\frac{(1-\xi_w)(1-\beta\xi_w)}{\xi_w}\mu_t^w + \varepsilon_t^w \\
\pi_t^w &= w_t - w_{t-1} + \pi_t && (20)\\
q_t &= \frac{R_*^k}{R_*^k + 1 - \delta} E_t\{mpk_{t+1}\} + \frac{1-\delta}{R_*^k + 1 - \delta} E_t\{q_{t+1}\} - r_t + \varepsilon_t^b && (21)\\
r_t^n &= \rho r_{t-1}^n + (1-\rho)\left(\rho_\pi \pi_t + \rho_y y_t^{gdp}\right) + \varepsilon_t^r && (22)\\
r_t &= r_t^n + E_t\{\pi_{t+1}\} && (23)\\
y_t^{gdp} &= \frac{c_y}{c_y + g_y + i_y} c_t + \frac{i_y}{c_y + g_y + i_y} i_t + \frac{g_y}{c_y + g_y + i_y} \varepsilon_t^g && (24)
\end{aligned}
$$

Tables and Figures

Table 1. Calibration

Parameter	value	description
β	0.995	discount factor
δ	0.025	capital depreciation rate
g_y	0.15	steady state share of gov't spending on output
α	0.25	capital share in prod function
i_y	0.22	share of investment
ϕ_w	1.5	steady state mark-up of wage setters
σ_l	1.5	inverse of the labour supply elasticity
ψ	0.1	adj. cost of capital utilization

Table 2. Summary statistics of the posterior distribution of the structural parameters

		prior		posterior: pre EMU		posterior: post EMU	
	distr.	mean	std. dev	0.90 interval	median	0.90 interval	median
σ_a	Inv. gamma	0.01	0.05	[0.003, 0.005]	0.004	[0.002, 0.003]	0.003
σ_b	Inv. gamma	0.01	0.05	[0.003, 0.006]	0.004	[0.003, 0.005]	0.004
σ_i	Inv. gamma	0.01	0.05	[0.004, 0.006]	0.005	[0.002, 0.003]	0.003
σ_g	Inv. gamma	0.01	0.05	[0.019, 0.026]	0.022	[0.010, 0.014]	0.012
σ_p	Inv. gamma	0.01	0.05	[0.002, 0.003]	0.002	[0.002, 0.003]	0.002
σ_w	Inv. gamma	0.01	0.05	[0.002, 0.003]	0.003	[0.002, 0.003]	0.002
σ_R	Inv. gamma	0.01	0.05	[0.001, 0.002]	0.002	[0.002, 0.002]	0.002
ρ_i	Beta	0.80	0.10	[0.450, 0.725]	0.590	[0.480, 0.752]	0.613
ρ_b	Beta	0.80	0.10	[0.832, 0.941]	0.892	[0.731, 0.896]	0.824
ρ_g	Beta	0.80	0.10	[0.738, 0.914]	0.832	[0.738, 0.910]	0.833
ρ_a	Beta	0.80	0.10	[0.853, 0.947]	0.899	[0.876, 0.973]	0.943
φ	Gamma	5.00	2.00	[2.586, 6.145]	4.086	[3.694, 9.232]	6.001
σ_c	Gamma	1.50	0.20	[1.291, 1.809]	1.536	[1.265, 1.749]	1.492
λ	Beta	0.50	0.20	[0.288, 0.548]	0.426	[0.460, 0.680]	0.579
ξ_w	Beta	0.75	0.10	[0.865, 0.923]	0.895	[0.756, 0.938]	0.818
ξ_p	Beta	0.75	0.10	[0.848, 0.915]	0.882	[0.677, 0.838]	0.745
ι_w	Beta	0.50	0.25	[0.210, 0.838]	0.539	[0.078, 0.607]	0.286
ι_p	Beta	0.50	0.25	[0.064, 0.480]	0.218	[0.034, 0.415]	0.146
ρ	Beta	0.75	0.10	[0.795, 0.894]	0.849	[0.736, 0.869]	0.803
ρ_π	Gamma	1.50	0.50	[1.041, 1.843]	1.392	[1.410, 2.779]	2.137
ρ_y	Normal	0.00	0.50	[0.270, 0.679]	0.438	[-0.005, 0.425]	0.108

Note: Results based 200,000 draws obtained with the Metropolis algorithm.

Table 3. The volatility of output, inflation and the nominal interest rate

Panel A. Data

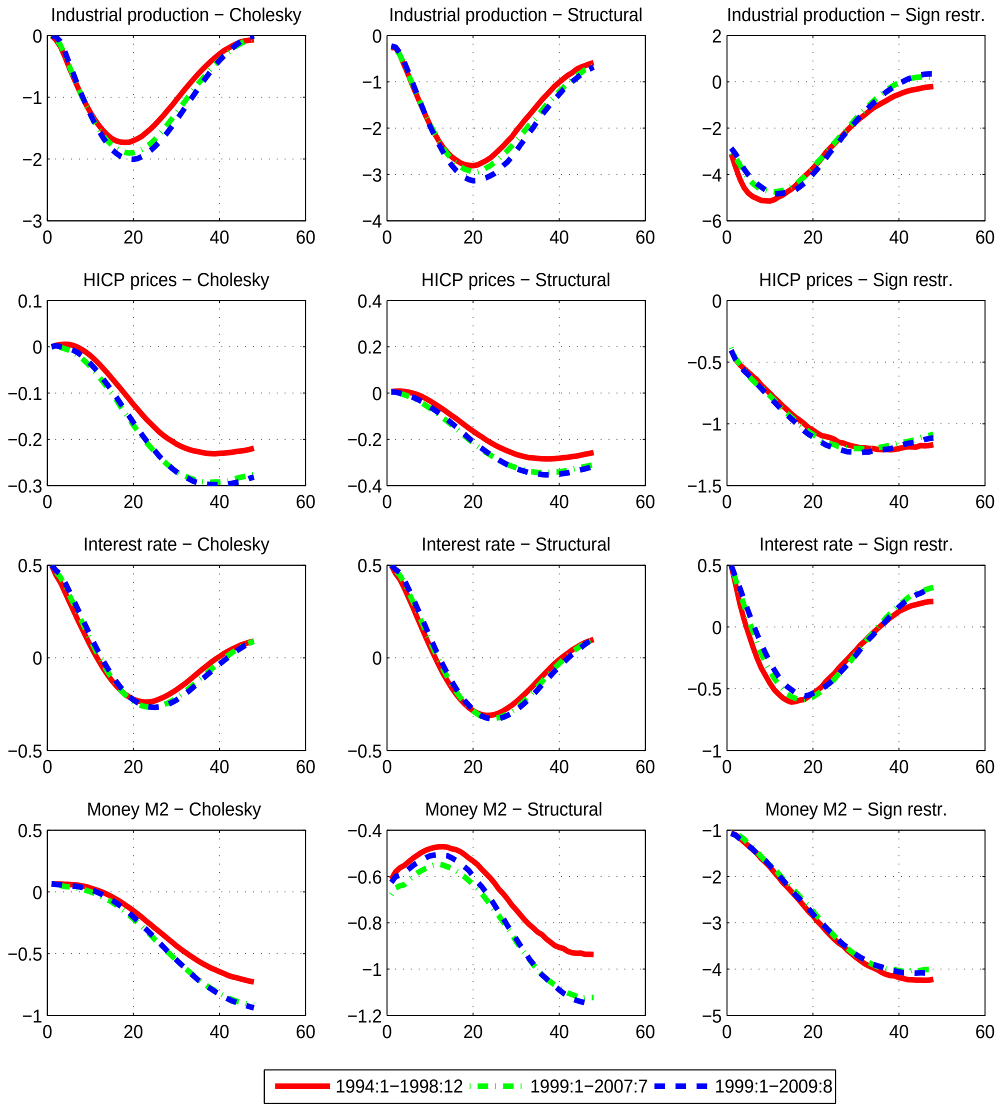
	std(ΔY)	**std(π)**	**std(r)**
pre EMU	0.47	0.38	0.67
post EMU	0.34	0.20	0.22

Panel B. DSGE

		std(ΔY)	**std(π)**	**std(r)**	**std(ΔY)**	**std(π)**	**std(r)**
MP	**PS**	*pre EMU shocks*			*post EMU shocks*		
pre	pre	0.73	0.46	0.62	0.56	0.35	0.53
post	pre	0.99	0.37	0.66	0.77	0.32	0.58
pre	post	0.62	1.43	1.23	0.43	0.90	0.81
post	post	0.75	0.43	0.59	0.50	0.34	0.46

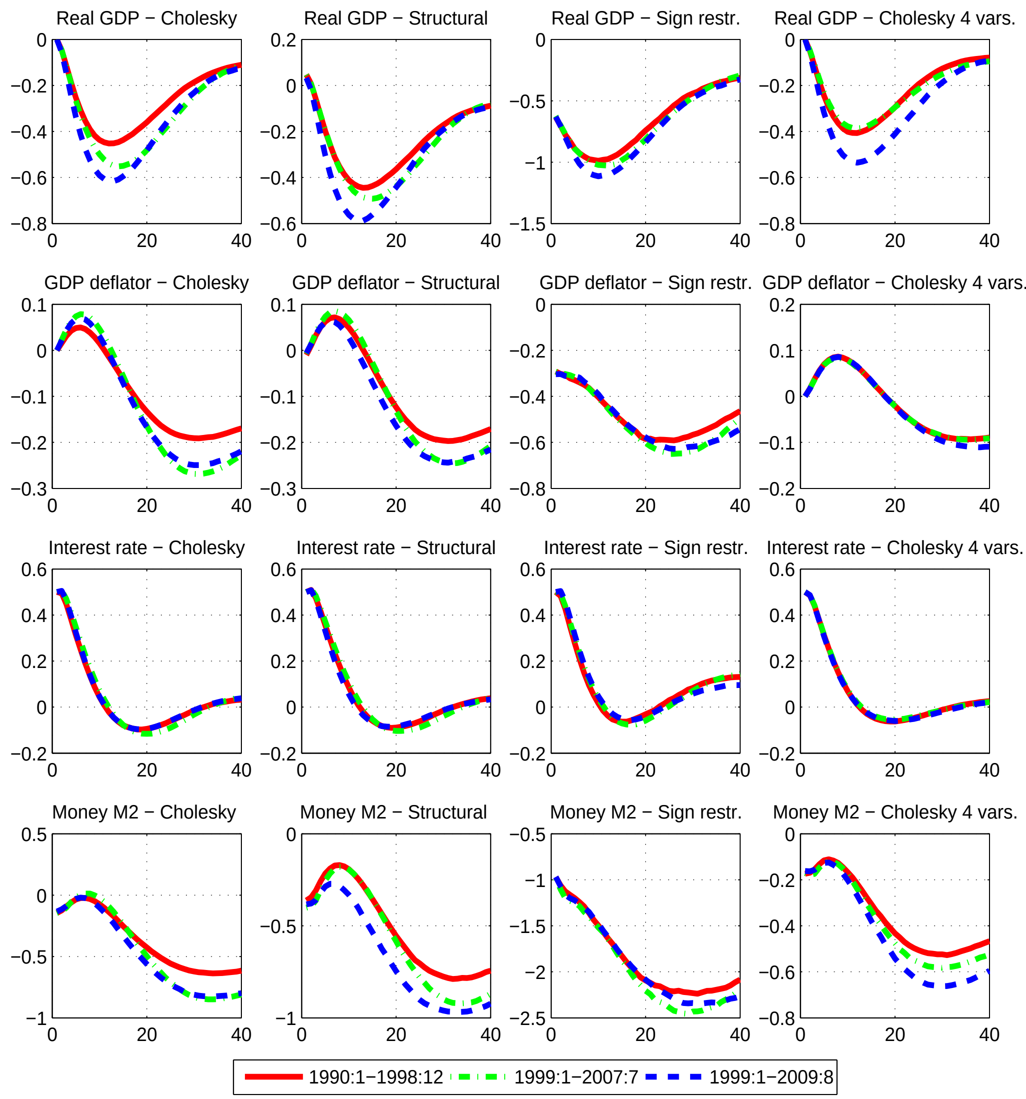
Notes: The standard deviation are theoretically computed from the solution of the model with different combinations of the private sector (PS), monetary policy (MP) and shocks parameters; π is the quarter-on-quarter inflation rate. ΔY is the quarter-on-quarter output growth rate.

Figure 1 - Median impulse responses based on the VAR with monthly data



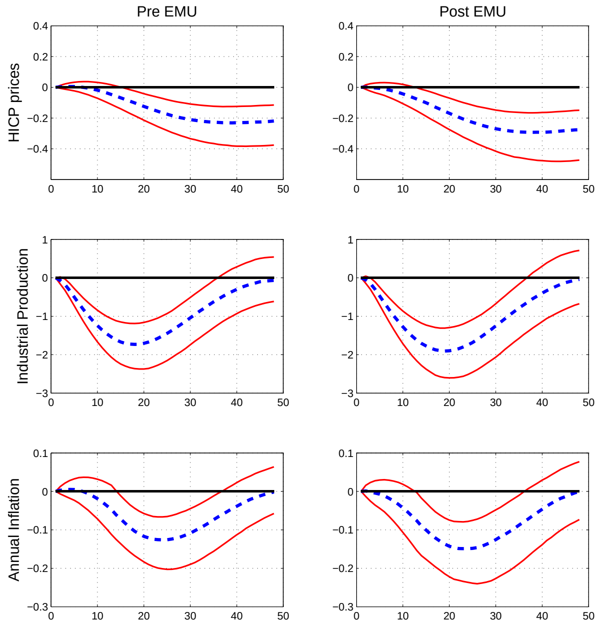
Note: Impulse responses are obtained with a 6-variable VAR.

Figure 2 - Median impulse responses based on the VAR with quarterly data



Note: Impulse responses are obtained with a 6-variable VAR.

Figure 3 - Impulse responses based on the VAR with monthly data and Choleski identification scheme



Note: Impulse responses are obtained with a 6-variable VAR. The confidence bands are the 95 percentiles.

Figure 4 - Robustness check on the prior tightness

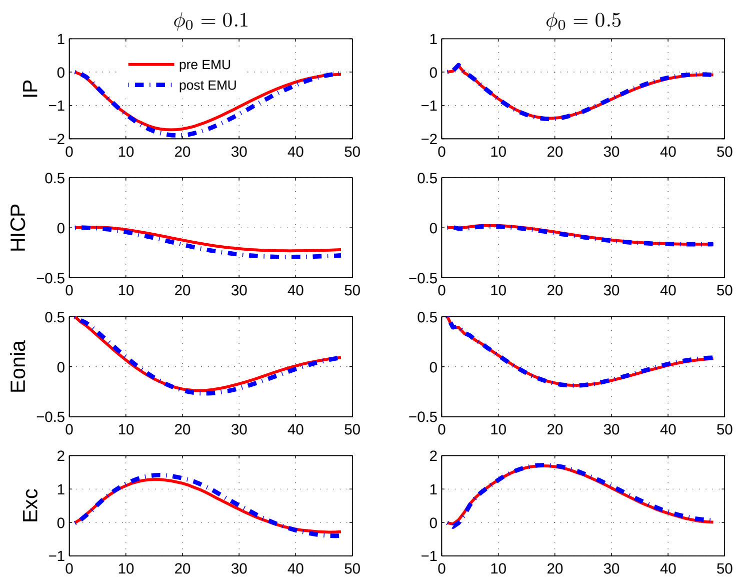
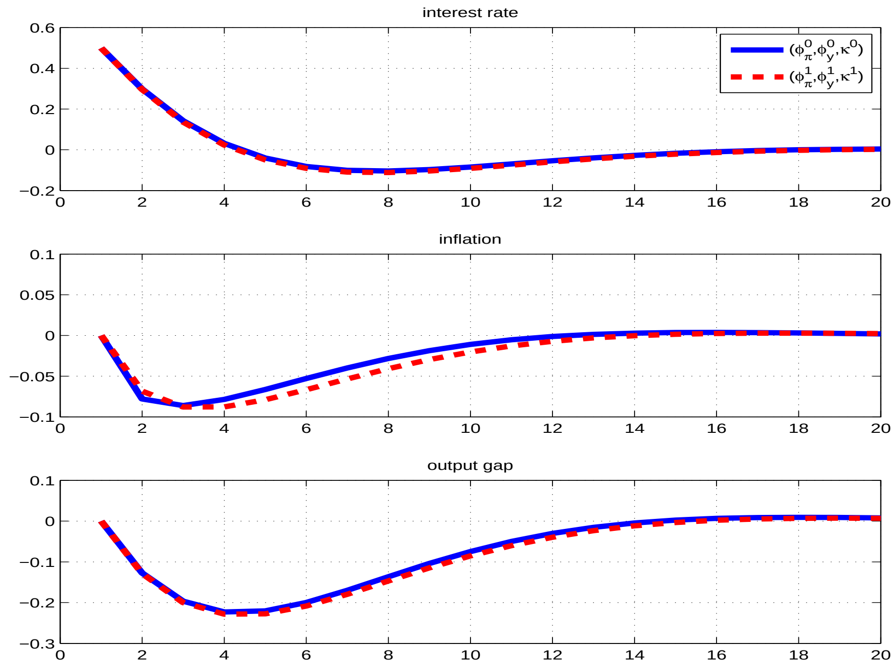


Figure 5 - VAR on simulated data



Notes: The small-scale DSGE model is given by the following equations:

$$\begin{aligned} x_t &= \alpha x_{t-1} + (1-\alpha) E_t x_{t+1} - \theta \left(i_t - E_t \pi_{t+1}\right) + \varepsilon_t^d \\ \pi_t &= \beta \pi_{t-1} + (1-\beta) E_t \pi_{t+1} + \kappa x_t + \varepsilon_t^p \\ i_t &= \rho i_{t-1} + (1-\rho)\left(\phi_\pi \pi_t + \phi_x x_t\right) + \varepsilon_t^i \end{aligned}$$

where x_t is the output gap, π_t is inflation and i_t is the nominal short-term interest rate. The exogenous variables ε_t^d, ε_t^p and ε_t^i are respectively demand, supply and monetary policy shocks. The impulse responses to a monetary policy shock, identified through Cholesky factorization, are obtained from the VAR estimated on the simulated data of the DSGE model above with different structural parameters. The blue solid line is the impulse responses coming from a DSGE with $\kappa^0 = 0.05$, $\phi_\pi^0 = 1.5$ and $\phi_x^0 = 0.5$. The red dashed line is the impulse response coming from a DSGE with $\kappa^1 = 0.04$, $\phi_\pi^1 = 2.5$ and $\phi_x^1 = 0.1$

Figure 6 - Prior and posterior marginal distributions: nominal rigidities

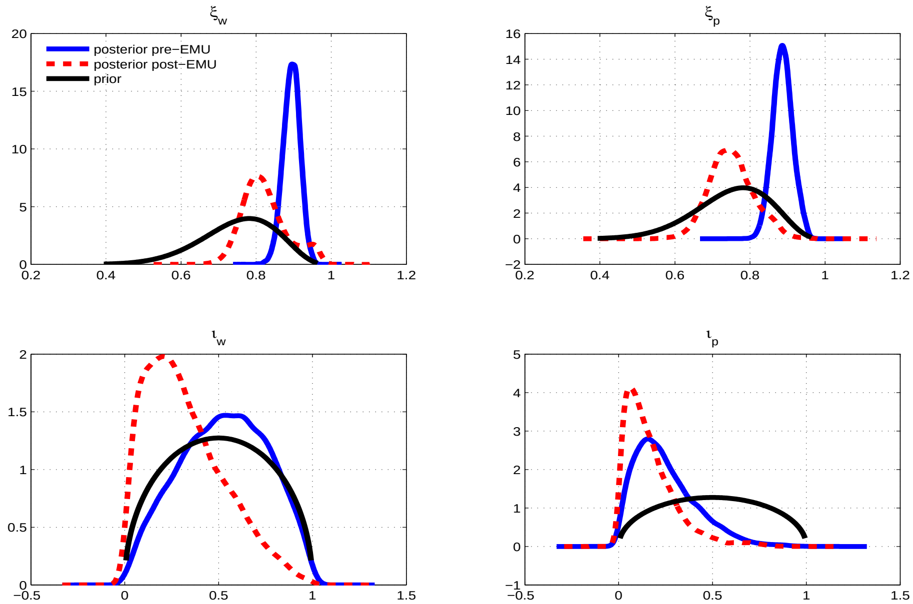


Figure 7 - Prior and posterior marginal distributions: monetary policy rule

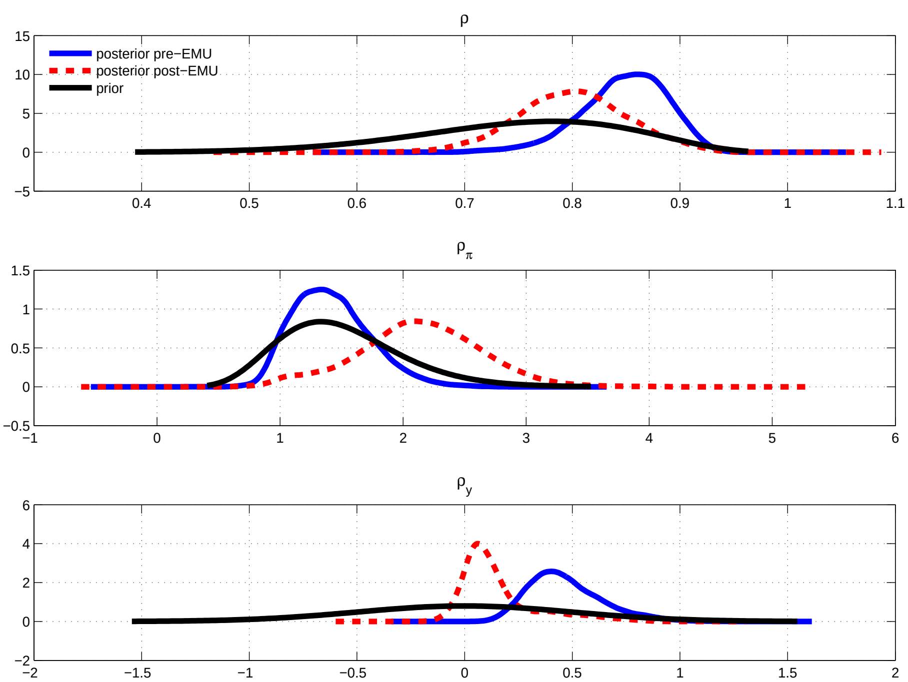
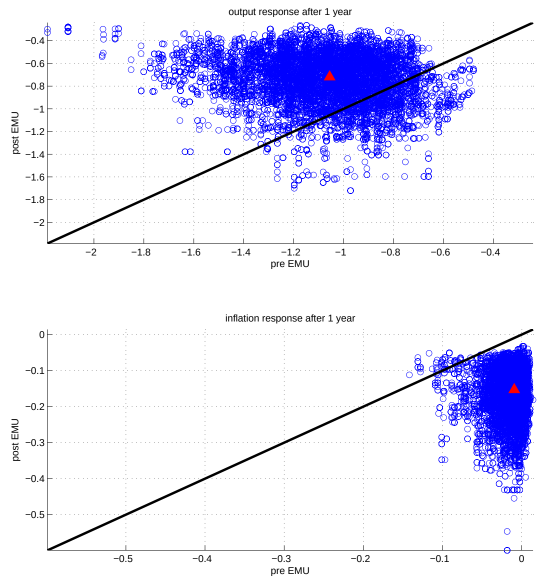
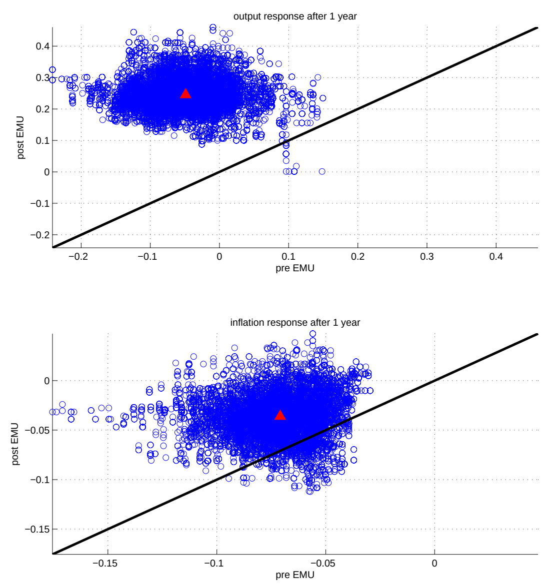


Figure 8 - The impulse responses at the 1 year horizon in the DSGE model

Panel (a) - Impulse responses to a monetary policy shock

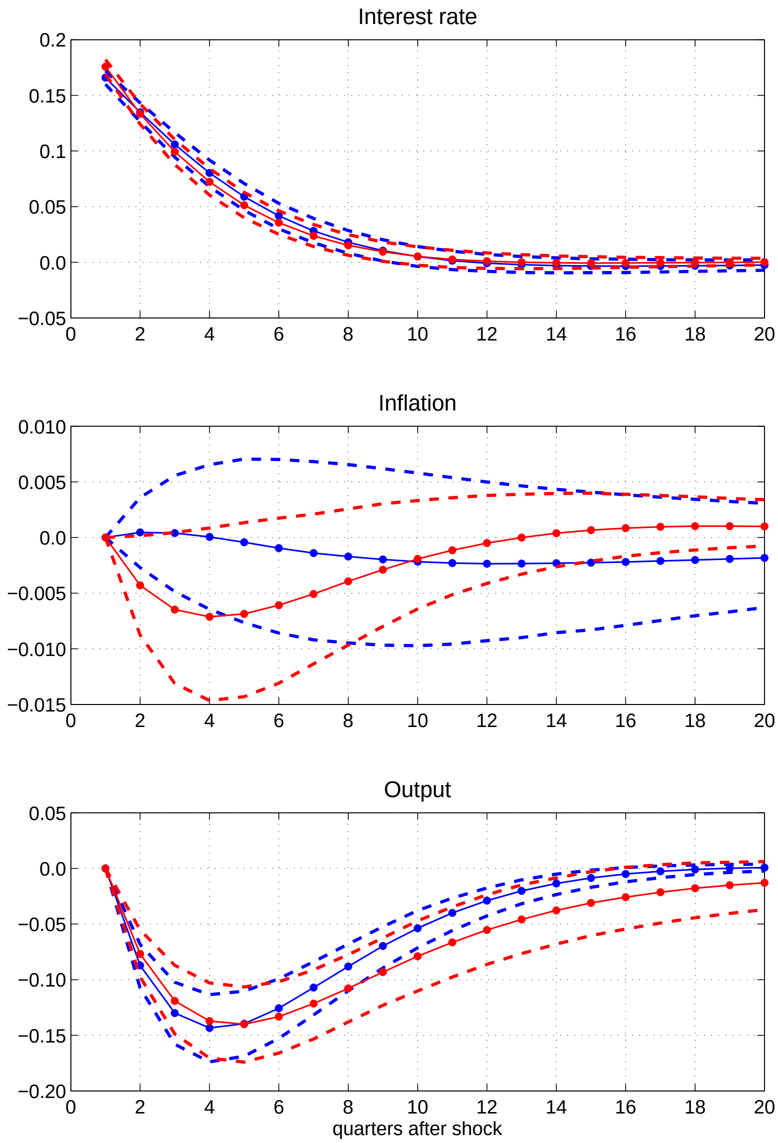


Panel (b) - Impulse responses to a technology shock



Notes: The figures plot the 10000 draws from the posterior distribution of the impulse responses of output and inflation to a monetary policy (panel (a)) and technology shock (panel (b)) at the 1 year horizon. In the horizontal axis the responses of the pre EMU sample; in the vertical axis those of the post EMU sample. The black solid line is the 45 degree line. The red triangle is the median of the posterior draws.

Figure 9 - Impulse responses to a monetary policy shock in a trivariate VAR estimated on simulated data



Notes: The blue lines are for the pre EMU sample; the red ones for the post EMU sample. The impulse responses are obtained from a VAR(1) on inflation, output and the interest rate with Cholesky identification. The VAR is estimated on 2000 time series, each of length 1000 observations, simulated from the DSGE model (described in section 3.2) with the parameters fixed at the median of their posterior distribution. The dotted line is the median in the two subperiods. The dashed lines are the 95% probability intervals which account for uncertainty of VAR estimates and the shocks.

Figure 10 - Impulse responses to a monetary policy shock - counterfactual analysis

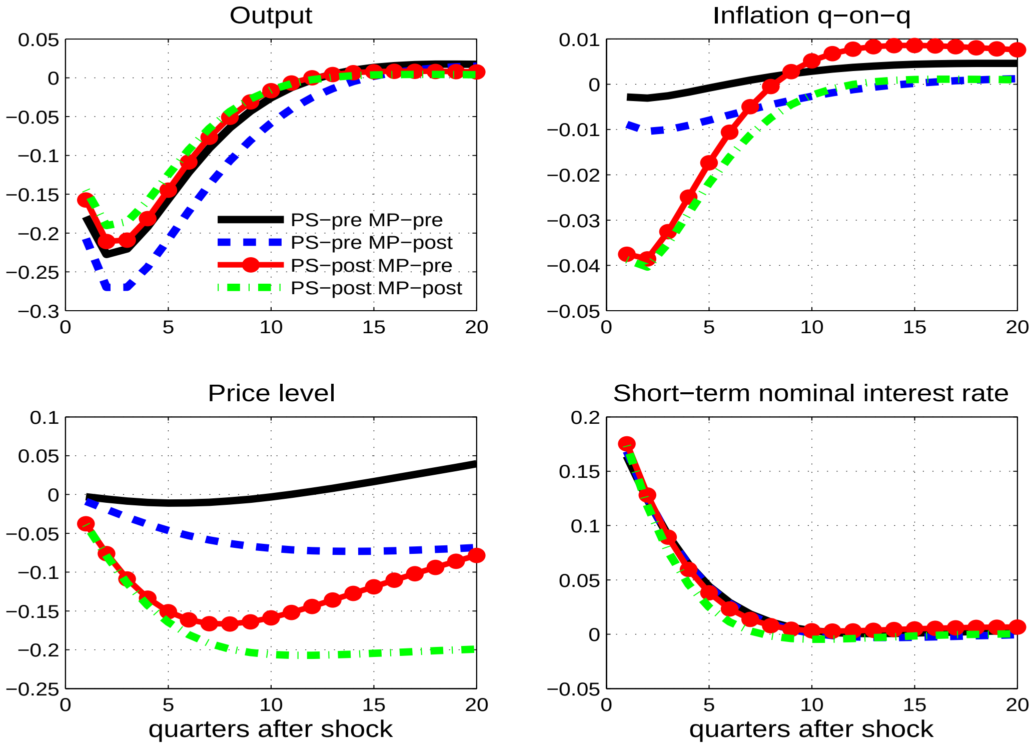


Figure 11 - Impulse responses to a technology shock - counterfactual analysis

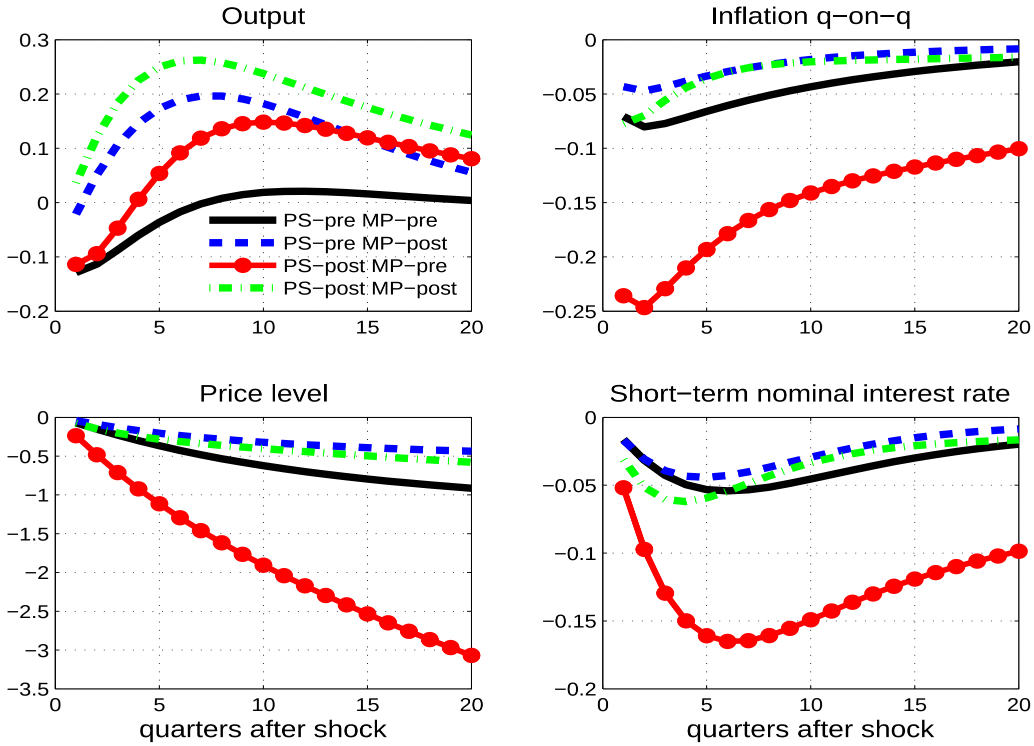
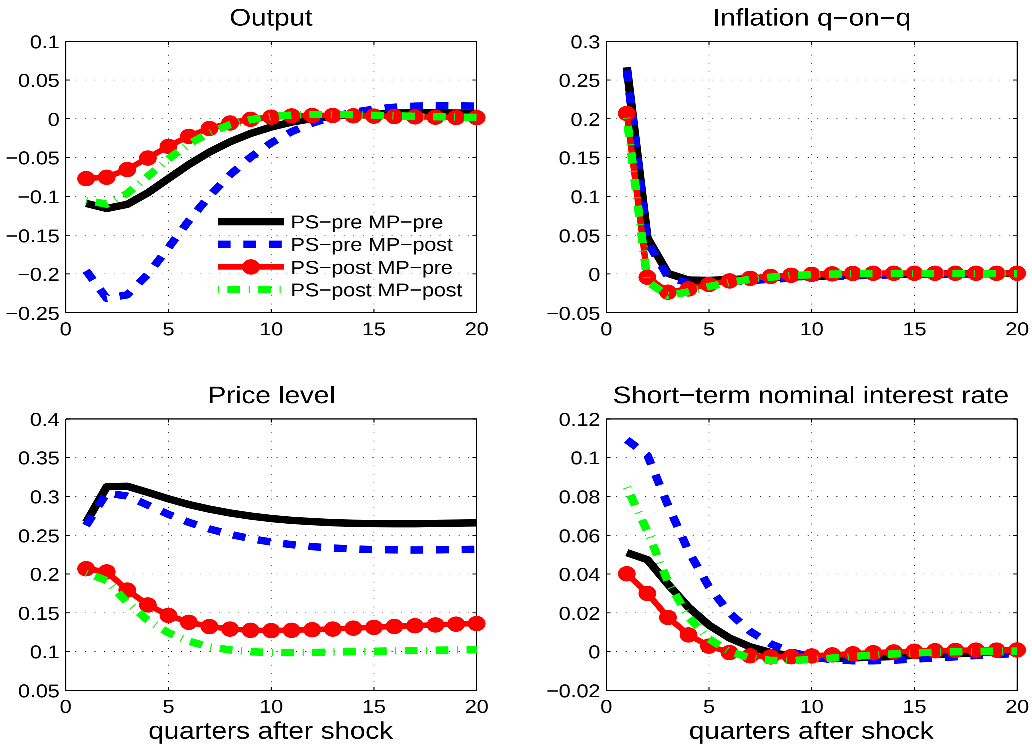


Figure 12 - Impulse responses to a price mark-up shock - counterfactual analysis



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