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RESULTS FROM AN ESTIMATED EURO AREA DSGE MODEL

by Marcelo Sánchez

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RESULTS FROM AN ESTIMATED EURO AREA DSGE MODEL[1]

Marcelo Sánchez[2]

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Abstract

This paper estimates a linearised DSGE model for the euro area. The model is New Keynesian and allows for a role for oil usage and endogenous price markups. We find that the price markup reacts positively to the ratio of expected discounted profits to current output, which is normally seen to give rise to a "countercyclical" markup. The importance of shocks to monetary policy and oil prices is estimated to have declined in the post-1990 period, in line with the higher predictability of policy and the fall in the persistence and - to a lesser extent - variability of oil disturbances. Counterfactual exercises show that oil efficiency gains would alleviate the inflationary and contractionary consequences of oil shocks, while higher wage flexibility would help ease the impact on real output at the expense of wider fluctuations in inflation. Finally, the rise in price markups induced by an oil disturbance is not found to considerably amplify the inflationary and contractionary effects of the shock.

Keywords: estimated DSGE models, euro area, oil shocks, endogenous markup

JEL Classification: C15, E31, E32, E37

Non-technical summary

In this paper, a linearised DSGE model is estimated for the euro area employing a maximum likelihood approach. The analysis pursued here revisits the questions addressed in the literature on the macroeconomic effects of oil price shocks, focusing on the euro area economy. In this context, we examine what is the magnitude of the effects of oil prices on the euro area economy, and whether the latter have changed over time. In doing so, we consider the role of endogenous price markup changes in the transmission of oil shocks to macroeconomic variables.

The theoretical model used is New Keynesian and draws from that developed by Leduc and Sill (2004). We follow these authors in allowing oil prices to impact the economy by linking oil usage to the variable rate of capital utilisation. Oil prices are modelled as being exogenous to euro area macroeconomic developments. In light of the maintained assumption that no oil income is recycled into domestic spending, oil price hikes act as a tax on the economy. We deviate from Leduc and Sill's (2004) approach in three main points. First, the price markup is not assumed to be exogenous, but is modelled as an endogenous variable instead. More specifically, the markup may depend on the ratio of expected discounted profits to current output in a rather general way. Second, unlike Leduc and Sill (2004), who model nominal rigidities by postulating adjustment cost functions, we adopt price and wage staggered contracts à la Calvo. Third, we do not posit a limited participation environment.

The results reported in this paper look reasonable, in light of the euro area economy's structural characteristics as well as the existing empirical and theoretical literature. The DSGE model estimated here has been found to exhibit reasonably good forecasting properties. With regard to structural parameters, we estimate an intermediate degree of wage rigidity, with the average length of wage contracts slightly exceeding 3 quarters. The price markup is found to react positively to the ratio of expected discounted profits to current output,

which is normally seen to give rise to "countercyclical" markup. This is in line with a number of existing studies for the US. Moreover, the postulated Taylor rule displays higher predictability in the post-1990 period, while there has been a decline in the persistence and - to a lesser degree - volatility of oil prices. The present study has uncovered evidence of significant structural change, as captured by parameter instability between pre- and post-1990 sample periods concerning behavioural parameters and exogenous processes driving structural disturbances.

With respect to the propagation mechanisms captured by the empirical DSGE model, the qualitative features of the responses reported here are broadly in line with those obtained in the literature. In the case of oil price shocks, they are found to generate inflationary and recessionary pressures as well as a reduction in real wages and an increase in the price markup. Overall, there is evidence that impulse responses to shocks have become milder in the post-1990 sample period compared with the pre-1990 period. Variance decomposition analysis points to the smaller macroeconomic importance of shocks affecting monetary policy and oil prices in the post-1990 period, which is in line with the higher predictability of policy and the fall in the persistence and - to a lesser extent - variability of oil disturbances.

Counterfactual exercises show that oil efficiency gains would alleviate the inflationary and contractionary consequences of oil shocks, while higher wage flexibility would help limit the contractionary effects of the disturbance at the expense of wider fluctuations in inflation. Finally, the rise in price markups induced by an oil disturbance is not found to considerably amplify the inflationary and contractionary effects of the shock.

1 Introduction

A large body of research shows that oil price movements have considerable consequences on real economic activity and inflation. An unexpected oil price increase is bad news for oil importing countries, in which the shock induces recessionary and inflationary pressures. The transmission mechanisms through which oil prices have an impact on real economic activity include both supply and demand channels. The supply side effects are related to the fact that crude oil is a basic input to production, and consequently an increase in oil price leads to a rise in production costs that induces firms to lower output. Fluctuations in oil prices also tend to depress aggregate demand. Empirical studies have by and large corroborated these insights for the US economy and other OECD countries.[1]

The present paper revisits the questions addressed in the literature on the macroeconomic effects of oil price shocks, focusing on the euro area economy. More concretely, it examines what is the magnitude of the effects of oil prices on the euro area economy, and whether the latter have changed over time. In doing so, we consider the role of endogenous price markup changes in the transmission of oil shocks to macroeconomic variables. While we do not assume that markup behaviour plays an important role in this regard, our analysis allows us to evaluate the relevance of this channel. Some authors have argued that the macroeconomic consequences of - say - higher oil prices could be amplified if markpus were to rise following the oil shock (as argued by Rotemberg and Woodford, 1996, for the US economy). On the contrary, a reduction of markpus following an oil shock would help mitigate the recessionary/inflationary impact of this disturbance. Price markups are introduced in a context of firms' monopolistic competitive behaviour. As demonstrated by Rotemberg and Woodford (1996), imperfect competition in product markets

[1] For a recent survey of the literature on the US, see Hamilton (2006). For evidence also on economies outside the US, see Mork *et al.* (1994), Bjϕrnland (2000), Cuñado and Pérez de Gracia (2003), and Jiménez-Rodríguez and Sánchez (2005, 2006, 2007).

is needed for unexpected oil price hikes to induce falls in both real output and real wages.[2] In addition, we shall introduce imperfect competition in labour markets.

In this paper, we estimate a linearised DSGE model for the euro area. The theoretical model used draws from that developed by Leduc and Sill (2004), which in turn is an extension of Finn's (1995) real business cycle model to the case where monetary policy plays a role in light of nominal frictions in product and labour markets. As such, the model used here belongs to the New Keynesian tradition in that forward-looking agents solve dynamic optimisation problems with rational expectations in an environment of slow price and wage adjustments.[3] As with Leduc and Sill (2004), we allow oil prices to impact the economy by linking oil usage to the variable rate of capital utilisation.[4] Oil prices are modelled as being exogenous to euro area macroeconomic developments.[5] However, it is worth mentioning that the literature has reported evidence that the direction of influence is not simply from oil prices to the macroeconomy but also the other way around.[6] In light of the maintained assumption that no oil income is recycled into domestic spending, oil price hikes act as a tax on the economy. This convenient simplification features in the related literature and has been rationalised in two ways. First, in connection with the assumption of exogenous oil prices, LS refer to their intention to "capture the impact of OPEC on the supply of oil" (p. 786).[7] Second,

[2] The available evidence is in line with this twofold prediction. See e.g. Jiménez-Rodríguez and Sánchez's (2005) results for the euro area and other major OECD economies.

[3] For the New Keynesian approach, see *e.g.* Clarida *et al.* (1999), and Woodford (2003).

[4] In his calibrated DSGE model, Roeger (2005) instead considers oil as a production factor alongside capital and labour.

[5] Moreover, we follow LS in not introducing any non-linear economic effect of oil prices of the type detected in the empirical literature pioneered by Mork (1989), Lee *et al.* (1995), and Hamilton (1996). For the different economic interpretations of these non-linear models, see Brown and Yücel (2002).

[6] On this, see block exogeneity tests in Jiménez-Rodríguez and Sánchez (2005, 2006). In connection with this, Barsky and Kilian (2004) and Kilian (2007) have warned against assuming on *a priori* grounds the exogeneity of oil prices.

[7] The connection with the economy's resource constraint appears to be more clear in the Appendix of LS' working paper version (Leduc and Sill, 2001). For an empirical assessment of the role of OPEC in driving world oil markets, see Dées *et al.* (2007).

Finn (1995, p. 1256) says that oil expenditure can be seen as value added to the production of final goods not by the domestic economy, but by a foreign country. In this interpretation, the domestic economy exports (imports) final (intermediate) goods to (from) this foreign economy by the amount given by the expenditure on oil; it is the only international trade that occurs and trade balances each period. One way or the other, the assumption of no oil income recycling appears to be a shortcut for modelling a foreign party which produces and sells oil to the domestic economy under study. For simplicity, no other open-economy feature is added to the model.[8] This aspect distinguishes the present paper from Backus and Crucini (2000), who endogenise oil prices in a three-country model that includes two industrial economies and oil producers. An intermediate position is followed by Roeger (2005), in which all the oil used by a single open economy is imported from the rest of the world, entering a balance of payments equation in a context where international trade is not assumed to balance in every period. The approach adopted here is also different from that developed by Rotemberg and Woodford (1996), who assume that income from costless oil production is fully recycled into the domestic economy.

We deviate from Leduc and Sill (2004)'s approach in three main points. First, the price markup is not assumed to be exogenous, but is modelled as an endogenous variable instead. More specifically, the markup may depend on the ratio of expected discounted profits to current output in a rather general way. Second, unlike Leduc and Sill (2004), who model nominal rigidities by postulating adjustment cost functions, we adopt price and wage staggered contracts à la Calvo (1983). While these two approaches are known to be observationally equivalent, the use of staggered contracts makes it more im-

[8] In their closed-economy model, Kim and Loungani (1992) also deduct the cost of energy before obtaining the output available for domestic demand. In the introduction to their paper, the authors refer to the aim of capturing terms of trade effects arising from unpredictable changes in imported energy prices. Unlike Finn (1995) and LS, who link oil usage to capital utilisation, the energy input enters Kim and Loungani's (1992) production function directly alongside standard capital and labour inputs.

mediate to relate the model to evidence available on the frequency of price adjustments in the euro area. Third, we do not posit a limited participation environment.

Our endogenisation of price markups deserves further clarification. As is known since Rotemberg and Woodford (1991, 1995), the three main approaches to explaining markup behaviour can be considered as special cases of a general formulation in which markups are allowed to be a function of the ratio of expected discounted profits to current output. These three approaches are: (i) the standard case where markups are simply assumed to be exogenous;[9] (ii) the case where the markup varies inversely with the ratio of expected discounted profits to current output, which is associated with the so-called "customer market" model of Phelps and Winter (1970),[10] and (iii) the case where the relation between markups and the ratio of expected discounted profits to current real output is positive, which can be rationalised by the "implicit collusion" model of Rotemberg and Saloner (1986).[11] Given that the ratio of expected discounted profits to current output is normally dominated by short-run developments in the latter, cases (ii) and (iii) are often referred to as situations where markups are procyclical and countercyclical, respectively.[12] Our estimated model can assess which of the three types of markup behaviour mentioned above is supported by the data. Evidence for the US appears to show that markups are countercyclical, ranging from the early study of Rotemberg and Woodford (1991) to the more recent work by Banerjee and Russell (2004), Wilson and Reynolds (2005) and Jaimovich (2006).

[9] In our model, markups would in this case be random instead of being simply constant.

[10] According to Phelps and Winter, a higher level of expected discounted profits relative to current output would induce each competing firm to reduce its markup in an attempt to increase its share in future sales to customers.

[11] In Rotemberg and Saloner's view, a rise in the ratio of expected discounted profits to current output allows for higher markups. In this context, implicit collusion can be sustained as an equilibrium given that a firm's deviation to a lower markup would not make it better off as such action would lead to sufficiently large losses from punishment.

[12] Galí (1994a, 1994b) has developed models in which markups depend on the share of investment in total output. This is a case worth exploring, which we leave for further research.

The present study uses a maximum likelihood approach to estimation. DSGE models with price rigidities have already been taken to the data in a maximum likelihood environment. The estimation strategy used involves the evaluation of the likelihood function of the model's log-linear approximation by means of the Kalman filter. This is the case of the studies carried out by Kim (2000) and Ireland (2001, 2004a, 2004b) for the US economy, and by Ambler *et al.* (2003), Dib (2003) and Bouakez and Rebei (2005) for Canada.

Our use of a likelihood approach to estimate the structural parameters of a general equilibrium model relates our work to other analyses conducted in a Bayesian instead of a classical context. For instance, the New Keynesian model extended with staggered wage contracts has recently been estimated using Bayesian techniques by Rabanal and Rubio-Ramírez (2005b) and Smets and Wouters (2007) for the US, and by Smets and Wouters (2003) and Rabanal and Rubio-Ramírez (2005a) the euro area. Smets and Wouters (2005) carry out a comparison between estimated New Keynesian models for the euro area and the US, finding substantially more similarities than differences between the two economies. The body of related literature extends beyond the universe of likelihood-based studies. Regarding the euro area, Galí *et al.* (2001) use a Generalised Method of Moments to estimate the inflation dynamics of a Calvo model incorporating price indexation. Sbordone (2002) employs minimum distance methods to estimate price and/or wage-setting equations separately. Rotemberg and Woodford (1997) and Christiano, Eichenbaum, and Evans (2005) minimised the distance between a structural VAR and the models' predicted impulse responses to a monetary shock. Smets and Wouters (2002) draw from this methodology for fitting a model to synthetic euro area data, while taking the process driving the fundamental factors entering the pricing equations as given.

The remainder of the paper is organised in the following way. Section 2 of the paper develops the DSGE model. Five disturbances – to households' preferences, to firms' technology and desired markups, to monetary policy and to oil prices – are allowed to account for fluctuations in the level of endogenous macroeconomic variables. In Section 3 we explain the data, describe the econometric methodology used, and report and discuss the main empirical results. Section 4 presents our conclusions.

2 The model

This section sets up the model that is at the basis of our econometric analysis. The economy consists of a continuum of households indexed by $i \in [0,1]$, representative final-goods-producing firms and a continuum of intermediate-goods-producing firms indexed by $j \in [0,1]$. Moreover, an oil sector provides the latter type of firms with an oil input which costs nothing to be produced. Oil prices follow an exogenous process, and oil income is not recycled into domestic spending. Each intermediate-goods-producing firm produces a distinct good; each of these goods may also be indexed by $i \in [0,1]$, where firm i produces good i. In addition to private agents, the economy comprises a government sector and a central bank.

2.1 Households

Any given household i maximises the expected utility function

$$\sum_{s=0}^{\infty} \beta^s E_t \left[A_{t+s} \ln C_{t+s}(i) - \frac{N_{t+s}(i)^{1+\gamma}}{1+\gamma} + \ln\left(\frac{M_{t+s}(i)}{P_{t+s}}\right)\right] \tag{1}$$

subject to the budget constraint

$$M_{t-1}(i)+B_{t-1}(i)+K_t(i)+W_t(i)N_t(i)+T_t(i)+D_t^F(i) \geq P_tC_t(i)+\frac{B_t(i)}{R_t}+M_t(i) \tag{2}$$

where $\beta \in (0,1)$ and γ is the inverse elasticity of labour supply with respect to real wages. In (1), each household i uses its resources to consume $P_tC_t(i)$

units of the finished good and save $M_t(i)$ units of money and $B_t(i)/R_t$ units of bonds. The household's resources consist of its previous holdings of money $M_{t-1}(i)$ and bonds $B_{t-1}(i)$, labour income $W_t(i)N_t(i)$, net transfers from the government $T_t(i)$ and nominal profits $D_t^F(i)$ received from the intermediate-goods-producing firms (denoted by F).

The preference shifter shock A_t follows the autoregressive process

$$\ln A_t = \rho_a \ln A_{t-1} + \varepsilon_{at}$$

with $\rho_a \in [0,1)$, where the zero-mean, serially uncorrelated innovation ε_{at} is normally distributed with standard deviation σ_a.

The first-order conditions include:

$$\begin{aligned} \frac{A_t}{C_t(i)} &= \beta R_t E_t \left(\frac{A_{t+1}}{C_{t+1}(i)} \frac{1}{\Pi_{t+1}} \right) \\ \frac{1}{R_t} &= \beta E_t \left(\frac{\lambda_{t+1}(i)}{\lambda_t(i)} \right) \end{aligned} \quad (3)$$

where $\Pi_t \equiv P_t/P_{t-1}$ is the (one-period) gross inflation rate and $\lambda_t(i)$ is the Lagrange-multiplier associated with the households' problem. The intertemporal optimality condition (3) links the real (that is, inflation adjusted) interest rate to the intertemporal marginal rate of substitution.

Each household is assumed to specialise in the supply of different type of labour. Households post the (nominal) wage at which they are willing to supply their labour services to firms demanding them. In a way analogous to firms' price setting, each period t only a fraction $1-\lambda_w$ of households (whose decisions we denote by $|t$), drawn randomly from the population, reoptimise their posted wage. All households resetting their wage in any given period will choose the same wage, and will thus face an identical problem. Next, we lay out and solve this problem.

Let W_t^* denote the wage newly set in period t. The choice of W_t^* must maximise:

$$\sum_{s=0}^{\infty} (\beta\lambda_w)^s E_t \left[A_{t+s} \ln C_{t+s} - \frac{{N_{t+s}}^{1+\gamma}}{1+\gamma} + \ln\left(\frac{M_{t+s}}{P_{t+s}}\right)\right]$$

subject to the flow budget constraint (2) for the relevant household and the sequence of labour demand schedules that are effective while remains W_t^* in place. The latter schedules are given by:

$$N_{t+s|t} = \left(\frac{W_t^*}{W_{t+s}}\right)^{-\Theta_w} N_{t+s}$$

where $N_{t|t}$ is the quantity of labour services provided in period t by a household that last reset its wage in period t, $N_t = \left(\int_0^1 N_t(i)^{(\Theta_w-1)/\Theta_w}\, di\right)^{\Theta_w/(\Theta_w-1)}$ is the index of aggregate employment, $W_t = \left[(1-\lambda_w) W_t^{*1-\Theta_w} + \lambda_w W_{t-1}^{1-\Theta_w}\right]^{1/(1-\Theta_w)}$ is the index of aggregate wages in period t, and Θ_w measures the (fixed) elasticity of labour to its relative wage rate. For convenience, instead of solving for W_t, we will later work with $\Pi_t^w \equiv W_t/W_{t-1}$, the (one-period) gross rate of change in the nominal wage, as the endogenous variable.

The first order condition associated with the problem above is:

$$\sum_{s=0}^{\infty} \lambda_w^s E_t \left\{ Q_{t+s} \left[\frac{W_t^*}{P_t \Pi_{t+s}} - \frac{\Theta_w}{\Theta_w - 1} MRS_{t+s|t} \right] \right\} = 0$$

where $Q_{t+s} \equiv \beta^s (C_t/C_{t+s}) / \Pi_{t+s}$ is the stochastic discount factor between t and $t+s$, $\Pi_{t+s} \equiv P_{t+s}/P_t$ the gross inflation rate for finished goods between t and $t+s$, and $MRS_{t+s|t} \equiv \left(N_{t+s|t}\right)^{\gamma} C_{t+s|t}/A_{t+s}$ is the marginal rate of substitution between leisure and consumption.[13]

[13] The inclusion of staggered wage contracts allows the model to deliver higher inflation persistence by means of a slower adjustment in real marginal costs. As usual, in the linearised symmetric equilibrium, nominal wage growth can be found to be a function of a forward-looking component and a term that measures the gap between the marginal rate of substitution and the real wage.

2.2 Firms

There are two types of non-oil producing firms in the economy: finished- and intermediate-goods-producing firms. The representative finished-goods-producing firm uses $Y_t(j)$ units of each intermediate good j, purchased at the nominal price $P_t(j)$, to manufacture Y_t units of the finished good according to the constant-returns-to-scale technology described by $Y_t = \left(\int_0^1 Y_t(j)^{(\Theta_t-1)/\Theta_t}\, dj\right)^{\Theta_t/(\Theta_t-1)}$. In the previous expressions, Θ_t measures the time-varying elasticity of a firm's output to its relative price, $P_t(j)/P_t$.

The intermediate-goods-producing firm combines $N_t(j)$ units of labour from the representative household and $K_t(j)$ units of capital to manufacture $Y_t(j)$ units of finished good i according to a decreasing-returns-to-scale technology described by

$$[u_t(j)K_t(j)]^{\alpha}\,[Z_t N_t(j)]^{1-\alpha} \geq Y_t(j) \tag{4}$$

where $u_t(j)$ denotes the capital utilisation rate and $\alpha \in (0,1)$ is the capital share.

The aggregate technology shock Z_t follows an autoregressive process with positive drift

$$\ln Z_t = \ln Z + \rho_z \ln Z_{t-1} + \varepsilon_{zt}$$

with $Z > 1$ and $\rho_z \in [0,1)$, where the zero-mean, serially uncorrelated innovation ε_{zt} is normally distributed with standard deviation σ_z.

Capital accumulation obeys the equation

$$K_{t+1}(j) = [1 - \delta(u_t(j))]\,K_t(j) + I_t(j) \tag{5}$$

with $\delta(.) \in (0,1)$, $\delta'(.) > 0$ and $\delta''(.) < 0$. The use of capital involves oil usage according to

$$K_{t+1}(j) = \frac{O_t(j)}{a(u_t(j))} \tag{6}$$

where an increase in the capital utilisation rate leads to higher oil usage, albeit in a decreasing fashion; that is, $a'(.) > 0$ and $a''(.) < 0$.

Let $D_t^F \equiv P_t(j)Y_t(j) - W_t N_t(j) - P_t(j)I_t(j) - P_t^o O_t(j)$ be the intermediate-goods-producing firm's profits, where P_t^o is the relative price of oil (in terms of P_t). The firm maximises $\sum_{s=0}^{\infty} E_t \left\{Q_{t+s} D_{t+s}^F\right\}$ subject to (4), (5), (6) and demand function $Y_t(j) = (P_t(j)/P_t)^{-\Theta_t} Y_t$. The first-order conditions include:

$$\begin{aligned} q_t(j)K_t(j) &= \left[\delta'(u_t(j)) + P_t^o a'(u_t(j))\right] K_t(j) \\ \lambda_t &= \beta R_t E_t \left\{\lambda_{t+1}\left[1 - \delta(u_{t+1}(j)) + P_{t+1}^o a(u_{t+1}(j)) + q_{t+1}(j) u_{t+1}(j)\right]\right\} \\ \Xi_t(j) &\equiv \frac{q_t(j)u_t(j))K_t(j)}{\alpha Y_t(j)} \\ \Xi_t(j) &= \frac{W_t N_t(j)}{(1-\alpha)Y_t(j)} \end{aligned}$$

where $q_t(j)$ denotes the rental rate for capital services $u_t(j)K_t(j)$, and $\Xi_t(j)$ is the nominal marginal cost for the firm in question.

Following Calvo (1983), we assume that each intermediate-goods-producing firm may reset its price only with probability $1-\lambda_p$ each period, independently of the time elapsed since the last adjustment. Thus, each period a measure $1-\lambda_p$ of producers (whose decisions we denote by $|t$) reset their prices, while a fraction λ_p keep their prices unchanged. Let P_t^* denote the price newly set in period t. Moreover, let $Y_{t+s|t}$, $D_{t+s|t}^F \equiv P_t^* Y_{t+s|t} - \Psi_{t+s}(Y_{t+s|t})$ and Ψ_{t+s} be the output, profits and nominal cost in period t, respectively, corresponding to a firm that last reset its price in period t. The choice of P_t^* must maximise the expected stream of profits generated during the period that price is quoted:

$$\sum_{s=0}^{\infty} (\lambda_p)^s E_t \left\{Q_{t+s} D_{t+s|t}^F\right\}$$

subject to the sequence of demand schedules that are effective while P_t^* remains in place. The latter schedules are given by:

$$Y_{t+s|t} = \left(\frac{P_t^*}{P_{t+s}} \right)^{-\Theta_t} Y_{t+s}$$

where $P_t = \left[(1 - \lambda_p) P_t^{*1-\Theta_t} + \lambda_p P_{t-1}^{1-\Theta_t} \right]^{1/(1-\Theta_t)}$ is the index of aggregate prices in period t.

The first order condition associated with the problem above is:

$$\sum_{s=0}^{\infty} \lambda_p^s E_t \left\{ Q_{t+s} Y_{t+s} \left[P_t^* - \frac{\Theta_{t+s}}{\Theta_{t+s} - 1} \xi_{t+s|t} \right] \right\} = 0$$

In the absence of costly price adjustment, the firm sets its markup of price $P_t(i)$ over marginal cost equal to $\Theta_t/(\Theta_t - 1)$. Thus, more generally, Θ_t can be interpreted as a cost-push disturbance, or a shock to the firm's desired markup. In light of costly price adjustments, the firm's actual markup hovers around its desired level over time. This desired level comprises an endogenous component, which is a (log-linear) function of the ratio of expected discounted profits X_t to current output,[14] and an exogenous component which is assumed to follow an autoregressive process. That is,

$$\begin{aligned} \ln \Theta_t &= \ln \Theta + \mu \ln \left(\frac{X_t}{Y_t} \right) + \ln \tilde{\Theta}_t \\ \ln \tilde{\Theta}_t &= \rho_\theta \ln \tilde{\Theta}_{t-1} + \varepsilon_{\theta t} \end{aligned}$$

with $\Theta > 1$ and $\rho_\theta \in [0, 1)$, where the zero-mean, serially uncorrelated innovation $\varepsilon_{\theta t}$ is normally distributed with standard deviation σ_θ.

[14] Expected discounted profits follow the expression $X_t = E_t \sum_{s=1}^{\infty} (\beta\eta)^s (\lambda_{t+s}/\lambda_t) \left[(\Theta_{t+s} - 1)/\Theta_{t+s} \right] Y_{t+s}$. The parameter η may have two different interpretations. In the "implicit collusion" model, it measures the rate at which new products are created as well as the probability that any collusive agreement will survive until the next period. In the "customer market" model, it represents the probability that a firm, for random reasons, be assigned a market share in the next period that is independent of its past pricing behaviour. See Rotemberg and Woodford (1991, 1996) for details.

2.3 The government

The government's deficit is entirely financed by seignorage:

$$M_t - M_{t-1} = T_t + P_t G_t$$

where G_t denotes real government spending, and $T_t = \int_0^1 T_t(i)di$. In the following, we shall assume that G_t grows at the same rate as Y_t, so that the share of government spending in total output remains constant.

2.4 The central bank

In the extension here, monetary policy is assumed to follow a monetary policy rule of the kind proposed by Taylor (1993):

$$\ln\left(\frac{R_t}{R}\right) = \rho_r \ln\left(\frac{R_{t-1}}{R}\right) + (1-\rho_r)\tau_\pi \ln\left(\frac{\Pi_t}{\Pi}\right) + (1-\rho_r)\tau_y \ln\left(\frac{Y_t}{Y}\right) + \varepsilon_{rt} \quad (7)$$

according to which the central bank changes the short-term interest rate R_t in reaction to inflation Π_t, and output Y_t. In this framework, the central bank chooses the steady-state values of the inflation rate Π and output Y as well as the response parameters τ_π, and τ_y. In (7), we also include an interest rate smoothing parameter, ρ_r. Finally, innovation ε_{rt} is assumed to be serially uncorrelated and normally distributed with zero mean and standard deviation σ_r.

2.5 Symmetric Equilibrium

In a symmetric equilibrium, each type of differentiated private agents (be it households or intermediate-goods-producing firms) face identical problems and takes identical decisions, so that $C_t(i) = C_t$, $M_t(i) = M_t$, $B_t(i) = B_t$, $N_t(i) = N_t(j) = N_t$, $T_t(i) = T_t$, $\lambda_t(i) = \lambda_t$, $D_t^F(i) = D_t^F(j) = D_t^F$, $Y_t(j) = Y_t$, $P_t(j) = P_t$, $K_t(j) = K_t$, $I_t(j) = I_t$, $u_t(j) = u_t$, $O_t(j) = O_t$, and $\Xi_t(j) = \Xi_t$

for all t and for all $i \in [0,1]$ and $j \in [0,1]$. In addition, the market-clearing conditions $M_t = M_{t-1} + T_t$ and $B_t = B_{t-1} = 0$ must hold.

Use of domestic agents' budget constraints allows us to derive the aggregate resource constraint:

$$Y_t - P_t^o O_t = C_t + I_t + G_t \tag{8}$$

The autoregressive exogenous processes driving the economy are given by:

$$z_t = \rho_z z_{t-1} + \varepsilon_{zt} \tag{9}$$

$$a_t = \rho_a a_{t-1} + \varepsilon_{at} \tag{10}$$

$$\tilde{\theta}_t = \rho_\theta \tilde{\theta}_{t-1} + \varepsilon_{\theta t} \tag{11}$$

$$p_t^o = \rho_o p_{t-1}^o + \varepsilon_{ot} \tag{12}$$

where lower-case variables z_t, a_t, $\tilde{\theta}_t$ and p_t^o are percentage deviations from steady state of the corresponding upper-case variables. Each of the total five innovations ε_{it} in the model (the four ones in (9)-(12) plus monetary policy shock ε_{rt}) follows a Normal $(0; \sigma_i^2)$ distribution, for $i = z; a; \theta; o; r$.

3 Empirical Analysis

In this section, we describe the data used in the analysis, the econometric strategy pursued and the substantive empirical results.

3.1 The Data

Our theoretical model puts together demand and supply features of the economy to explain the joint behaviour of key euro area macroeconomic variables, including measures of economic activity, inflation, labour market variables and interest rates. In order to estimate the model, we use data measured at a quarterly frequency for four series: real output Y_t (measured by real GDP), real private consumption C_t (as in the national accounts), inflation Π_t (measured by the consumption deflator), and the short-term nominal interest rate R_t. In addition, the exogenous process driving the real oil price P_t^o is characterised by converting the US Brent into euro (data from IMF's *International Financial Statistics*), and then deflating it by the GDP deflator. The sample period is 1990:1 to 2006:1 for the period of interest, with an alternative earlier period of 1970:1 to 1989:4. Unless otherwise noted, information on the euro area comes from the "synthetic" dataset produced at the Econometric Modelling Division of the European Central Bank (see Fagan *et al.*, 2005, for details). Real observable variables (real GDP, real private consumption and the real oil price) are detrended in a quadratic fashion, while nominal variables (inflation and interest rates) are treated as deviations from their unconditional mean.

3.2 Econometric Methodology

The equations listed in the previous section represent agents' behaviour and identities that altogether form a non-linear system. This includes the first order conditions of households and firms, agents' budget constraints, the monetary policy rule and equations describing the exogenous processes driving the economy. In order to estimate the model, we start by focusing on the symmetric equilibrium for prices and quantities. We derive all the log-linearised equations of the model by taking log-linear approximations around the steady state.[15] The resulting linearised DGSE model involve five observable variables (real oil price p_t^o, which is exogenous to macroeconomic developments, and four endogenous variables: real output y_t, real private consumption c_t, inflation π_t and the short-term nominal interest rate r_t); five unobservable shocks (to preferences a_t, technology z_t, desired markups $\tilde{\theta}$, oil prices p_t^o, and monetary policy ε_{rt}), and remaining variables treated as unobservables.

In particular, the linearised DGSE model involves two equations for output

[15] In the present subsection, lower-case variables denote log-deviations of the corresponding upper-case variables from the steady-state values.

inflation and wage inflation (π_t^w), as follows:

$$\pi_t = \beta E_t \pi_{t+1} + \kappa_p (\xi_t + \theta_t) \quad (13)$$

$$\pi_t^w = \beta E_t \pi_{t+1}^w + \kappa_w (mrs_t - \varpi_t)) \quad (14)$$

where $\varpi_t \equiv w_t - p_t$ is the real wage rate, $\kappa_p \equiv (1-\alpha)(1-\beta\lambda_p)(1-\lambda_p)/\left[\lambda_p(1+\alpha(\Theta-1))\right]$, $\kappa_w \equiv (1-\beta\lambda_w)(1-\lambda_w)/\left[\lambda_w(1+\Theta_w\gamma)\right]$, and λ_p and λ_w are the probabilities of keeping prices and wages fixed during the period, respectively. The marginal rate of substitution can be found to equal $mrs_t = \gamma n_t + c_t - a_t$.

Our estimation approach starts by setting up a state-space econometric model from the solution to our system of equations. Once this is done, the Kalman filter is applied to estimate the model's parameters via maximum likelihood and carry out inference (see *e.g.* Hamilton, 1994). Our main objectives are to obtain parameter estimates, compute the contributions made by the various shocks in driving macroeconomic fluctuations and describe the propagation mechanisms in place in the three economies under study.

3.3 Parameter calibrations

There are two types of parameters in the log-linear DSGE model, namely, those parameters or ratios that are calibrated, and those parameters that are estimated.

Calibrated parameters include β, α, λ_p, Θ, Θ_w, and η. Parameters β and α are set to standard values of 0.99 and 0.36, respectively. Rather than estimating λ_p, we set it to 0.77, which is the value implied in a Calvo context by the micro evidence on average duration of price contracts (13 months) reported in Alvarez *et al.* (2006), and Dhyne *et al.* (2006). Coefficients Θ and Θ_w also adopt common values of 6 and 8, respectively, which imply corresponding steady-state markups of 20% in intermediate goods prices (under the benchmark case of purely exogenous price markups) and 14% in wages. Parameter

η is set to 0.89, as in Rotemberg and Woodford (1996).

We follow Finn's (1995) - and Leduc and Sill's (2004) - specifications for the utilisation and depreciation functions $a(u_t)$ and $\delta(u_t)$ as

$$a(u_t) = \frac{u_t^{\gamma_1}}{\gamma_1}; \qquad \delta(u_t) = \frac{u_t^{\gamma_2}}{\gamma_2} \tag{15}$$

Parameters γ_1 and γ_2 are set to the values corresponding to the steady state of the model, that is, 1.97 and 1.30, respectively. These values are consistent with a depreciation rate on capital $\delta = 0.025$ and the ratio of oil usage to real value added equal to 2.7%, as found using the OECD input-output tables available for euro area countries for the most recent year (2000).[16]

We also calibrate the values for the ratios of private consumption and investment to real value added at 0.58 and 0.20, respectively. We ignore changes in either ratio between the two sample periods, as these changes have been rather small.

3.4 Parameter estimates and parameter stability

In addition to the model's calibrated parameters, there are two types of estimated parameters: i) parameters that are estimated separately from the rest of the model; and ii) parameters that are estimated jointly, conditional on those obtained from i) and those that are calibrated. Estimates are reported here for two sample periods, namely, 1970:1-1989:4 (first, or pre-1990, period) and 1990:1-2006:1 (second, or post-1990, period).

The former group of estimated parameters, namely, those that are estimated separately from the rest of the model, involve the autoregressive coefficient and the standard error of the process governing the real oil price. Here, we allow for the estimates (obtained using OLS) to vary across the two sample

[16] Such information is available for Austria, Belgium, Finland, France, Germany, Italy, the Netherlands and Spain. The oil-to-value added ratio is relatively robust to also considering similar data available for other euro area countries (Ireland for 1998, and Greece and Portugal for 1999). The value reported in the text is exclusive of refining products.

periods considered. For the period since 1990, we obtain an autoregressive coefficient of 0.78 and a standard error of 0.14, compared with the corresponding (higher) estimates of 0.93 and 0.19, respectively for the period 1970:1-1989:4.

Table 1 presents the coefficient estimates for the second group of estimated parameters, that is, those that are estimated jointly, conditional on the ones already discussed. More concretely, Table 1 reports the point estimate and the standard deviation of the parameters.[17] Let us start by describing the estimates for the post-1990 period. With regard to behavioural parameters, the estimate of the Calvo parameter for wage rigidity λ_w implies that the average length of wage contracts slightly exceeds 3 quarters, which is in line with the result reported by Smets and Wouters (2005) also for the euro area. Among the other behavioural parameters, our estimate for the labour supply parameter is - at 1.15 - somewhat below the value reported by Rabanal and Rubio-Ramírez (2005a), and Smets and Wouters (2003). The response of the markup to the ratio of expected discounted profits to current output is estimated to equal 0.15, which coincides with the calibration of Rotemberg and Woodford (1996).

Turning to the estimated stochastic processes for the structural shocks, the autocorrelation of technology shocks is somewhat stronger (at 0.60) than that of preference and markup shocks. The standard deviation of preference shocks is the largest of all disturbances (at some 35%), followed by those of markup shocks (23%) and monetary shocks (14%).[18] The standard deviation of technology shocks is estimated to be rather small (around 2%). The low value for σ_z can be rationalised in terms of our introduction of variable capital

[17] Standard errors are computed by taking the square roots of the diagonal elements of -1 times the inverted matrix of second derivatives of the maximised log-likelihood function. As explained by Ireland (2004c), the steps involved in these computations may introduce approximation error into the estimated standard errors.

[18] In the case of σ_θ, the relatively large estimates normally reported in the literature have been attributed to the fact that similarly stylised models may not be successful in matching observed inflation persistence, thereby yielding inflation variability via a rather volatile markup process.

utilisation, which is expected to lower the value of this coefficient as argued in King and Rebelo (2000).

Finally, concerning the estimated policy rule, the long-run response to inflation is estimated to be 1.21 while the long-run response to the output gap is 0.25. Our estimate for the interest rate smoothing coefficient ρ_r is 0.35, which is below the range of values reported by Rabanal and Rubio-Ramírez (2005a), and Smets and Wouters (2003).

Table 1 also presents the coefficient estimates for the pre-1990 period, as well as the stability tests associated with a comparison between the two periods. These tests are Wald tests (see Andrews and Fair, 1988) conducted over several subsamples and various subsets of parameters to test for parameter constancy. The stability tests reported in Table 1 allow us to reject the similarity of estimates between the two periods in the cases of all coefficients as a whole, as well as for the subgroups of structural parameters and shocks' standard errors. The hypothesis of parameter constancy is not rejected for the case of policy parameters, although it must be borne in mind that the imposition of a Taylor rule for the period 1970:1-1989:4 gives rise in this case to a significantly larger error volatility (56%) in the monetary policy equation.

3.5 Impulse responses

Figure 1 presents a set of impulse responses derived from the estimated model. The blue solid lines in this Figure are point estimates for impulse responses over the period 1990:1-2006:1, while the (blue) dotted lines represent the corresponding 95% confidence bands. For comparison purposes, we also report as a red (solid) line the point estimates for the period 1970:1-1989:4. The impulse responses shown in Figure 1 depict the reaction of real output, inflation, real wages and the price markup to exogenous changes in preferences, technology, the price markup itself, monetary policy and real oil prices. While the latter disturbance consists of a unit shock (capturing an unexpected doubling of real oil prices), all other cases represent one-standard deviation disturbances.

The qualitative features of the responses presented in Figure 1 are broadly in line with those obtained in the literature. Under a positive technology shock, the euro area economy experiences deflationary pressures and a boost to real output, while real wages go up and the price markup exhibits some volatility (first rising, then falling). The markup shock naturally induces an increase in the markup, which is associated with an inflation hike as well as a fall in real output and real wages. The monetary shock yields a drop in both real output and inflation, while also leading to lower real wages and a higher markup. The oil price shock generates inflationary and recessionary pressures as well as a reduction in real wages and an increase in the markup.

In the case of the preference shocks, the results are known to depend on the specific parameter values that have been estimated. In this regard, it is worth saying that a - say - positive realisation of this disturbance combines a favourable shift in the demand for consumption and an adverse supply shock development. In our case, both real output and inflation react by rising following such disturbance, while real wages rise and the markup contracts.

In order to assess the plausibility of our DSGE model results for responses to oil price shocks, we have also estimated two other models that are suited for coping with the parameter instability usually reported when estimating linear models since the 1970s. The two models have been estimated for the euro area over the sample 1970:1-2006:1 using the data described above. The first model is a so-called "scaled" structural VAR model in the spirit of Lee *et al.* (1995), which uses a non-linear transformation of the real oil price defined as the latter's unexpected component rescaled by its time-varying volatility. We use as endogenous variables the real oil price, real GDP growth, consumer price inflation, real wage growth, the real effective exchange rate, and real short- and long-term interest rates. Identification of oil shocks is based on a recursive scheme, where oil shocks are allowed to react contemporaneously only to unexpected developments in real GDP growth (see Jiménez-Rodríguez and Sánchez, 2005 and 2006, for details).[19] The second model is one that allows for stochastic volatility in real output and inflation. Unlike Uhlig (1997), who allows for multivariate stochastic volatility in the context of VARs, we simply estimate bivariate models in which either real GDP or inflation is the endogenous variable and real oil prices are treated as the exogenous variable.[20] In this case, our assumption that the regression coefficients are constant (in contrast to time-varying stochastic volatilities) can be justified on the basis that, given the simplicity of the model, there is no attempt at distinguishing between changes in the typical size of the exogenous innovations and changes in the transmission mechanism.

With regard to model comparisons, the (full-sample) scaled model yields responses that lie in between the (larger) ones obtained from the pre-1990 DSGE model and the (lower) ones produced by using the post-1990 DSGE model. The only exception to this is the rather high estimate for the inflation response that is generated by the structural VAR within the first year after the shock. The bivariate stochastic volatility models instead produce full-sample responses that are rather low and comparable to those obtained from the post-1990 DSGE model. These comparative results are overall plausible. While allowing for time-varying variability in the macroeconomic variables, the stochastic volatility model is a linear approach that is expected to deliver milder reactions in real output and inflation compared to our non-linear structural VAR. Taken altogether, the DSGE models estimated for two different sample periods appear to yield results that are comparable to those of the scaled model, which is estimated over the entire sample instead.

Turning to comparisons with earlier DSGE models that analyse the role of

[19] The Akaike criterion indicates that the optimal lag is 4. The impulse responses to oil shocks are scaled down by the sample mean of the so-called scaled oil measure's conditional standard deviation.

[20] In the stochastic volatility models, the optimal lags are 4 for the endogenous variable and 5 for the exogenous variable, according to the Akaike criterion. All variables were detrended using a quadratic trend.

oil shocks, the present paper estimates that the macroeconomic effects of such disturbances are milder than in the calibrated models of Rotemberg and Woodford (1996) and Leduc and Sill (2004) for the US, but somewhat stronger than those reported in the also calibrated work of Roeger (2005) for the EU/OECD. We can also broaden the comparison by referring to simulations obtained using leading macroeconometric models. For the sake of concreteness, let us concentrate on existing results for the euro area (see Table 3). Focusing on our post-1990 estimates, real output (accumulated) responses to the oil shock are closest to the two simulations obtained from the QUEST model. They can also be seen to be stronger than the simulations produced by the AWM, Interlink and Multimod models, and milder than those of NIGEM. Turning to inflation impulse responses, our estimates are in line with NIGEM and QUEST models, which happen to be on the low side. Regarding the remaining models, our pre-1990 (considerably larger) inflation results appear to conform with AWM and Interlink simulations.

In sum, the short-run propagation mechanisms obtained for the euro area look reasonable. In particular, there is evidence that impulse responses to shocks have become milder in the post-1990 period compared with the pre-1990 period. The next subsection turns to variance decomposition analysis, which will allow us to evaluate the mechanisms at play in some more detail.

3.6 Variance decompositions

Table 4 decomposes the forecast error variances in real output, inflation, the price markup, and real wages into components attributable to each of the five shocks. In doing so, we distinguish between the two sample periods considered, namely, 1970:1-1989:4 and 1990:1-2006:1.

The variance decomposition results show that technology and monetary policy disturbances have driven real output in the first sample period, accounting for at least some one third of the latter variable's variability each.

While technology shocks continue to be an important driving force behind real output in the post-1990 period, the role of monetary disturbances in explaining real activity's fluctuations has diminished markedly. Moreover, in the second period movements in real output appear to be dominated by preference shocks. Concerning oil price disturbances, their contribution has declined from somewhat over 10% of total real output's variability at the end of the first year in the pre-1990 period to some 6% in the post-1990 period. After the first year, the contribution of oil shocks has roughly halved to reach some 5% of the total.

With regard to inflation, markup shocks are found to be the driving force in the pre-1990 period, but their importance is considerably reduced in the second period. Technology shocks play a relatively important role in aggregate price developments in both sample periods, with their share in inflation variability rising however to reach almost half of the total in the post-1990 period. Finally, the role of preference shocks appears to have risen over time, while the relative importance of monetary policy and oil disturbances has declined.

Turning to the remaining two variables in Table 4, price markup fluctuations appear to be dominated by its own shock in both sample periods, with oil shocks in particular playing a minor role. This result stands in contrast with Rotemberg and Woodford's (1996) emphasis on the role played by oil shocks in inducing higher markups in an environment combining monopolistic competition and produced intermediate inputs. Regarding real wages, they have been driven in the pre-1990 period by markup shocks in the first three years or so after the shock, and by technology shocks over the longer run. In the more recent sample period, preference disturbances play a dominant role within the first two years, while technology shocks remain the key driving force behind real wage developments beyond that time horizon. The share of real wage movements explained by oil price disturbances is reasonably large in the short run (at 8% and 7% after four quarters in the pre- and post-1990 sample periods, respectively), thereafter declining in the case of both estimation periods.

By way of summary, our variance decomposition analysis shows that technology disturbances have played an important role in explaining movements in real output and inflation over time, as well as real wages in the medium- to long-term. Preference shocks have increased their share in overall macroeconomic variability in the post-1990 period, becoming of special relevance in the cases of real output, inflation and real wages. Markup disturbances have lost some of the importance in explaining inflation and real wages seen in the pre-1990 period, while still remaining the key force behind price markup developments themselves. The decline in the role of monetary policy shocks in driving macroeconomic variables can be traced to the higher predictability of policy as captured by means of a simple Taylor rule. Finally, the empirical relevance of oil shocks for macroeconomic developments has declined in the post-1990 period, in line with the reduction observed in these shocks' persistence and - to a lesser degree - volatility. Yet, these disturbances still explain no less than 5% of the forecast error variance of real output, inflation and (within the first year) real wages.

3.7 Forecast accuracy

This subsection uses so-called bias-corrected S-tests (see Harvey *et al.*, 1997) to evaluate the null hypothesis that there is no difference between the k-step-ahead forecast accuracy of our DSGE model and two benchmark unrestricted vector autoregressions (VAR). The latter are given by VAR(1) and VAR(2) models estimated using 1 and 2 lags of observable variables (the same variables also used in estimating the DSGE model), respectively. The k-step-ahead forecasts involved in these tests start using the competing models estimated over the sample period 1990:1-2003:1 and end using the models estimated over the sample period 1990:1-2006:1.

Table 5 reports bias-corrected S-tests alongside root mean squared errors (RMSE) for the three alternative models. For both types of statistics, we report results obtained varying the number of step-ahead quarters from 1 to 4. In Table 5, the S statistics adopt negative values, which indicates that the DSGE model's forecasts are outperformed by the VAR models. However, in the vast majority of cases the tests fail to reject the null of equal forecast accuracy at the 5% significance level. The only cases in which the forecast errors are significantly larger in the DSGE model correspond to the 1-quarter-ahead forecasts for real output and the interest rate. Overall, therefore, the DSGE model's forecasting performance appears to be reasonably good.

3.8 Counterfactuals

Table 6 reports counterfactual exercises that show how changes in parameter values may lead to a different macroeconomic outcome from oil price shocks. The first counterfactual experiment assumes that the euro area economy enjoys a 20% efficiency gain in oil use. This is found to considerably reduce the effect of a sudden oil price hike on both inflation and real output. The second counterfactual exercise reported in Table 6 consists of a fall in λ_w, which amounts to reducing the wage contract length by one quarter. This counterfactual exercise indicates that more flexible wages on balance induce a milder contractionary impact following the disturbance. The lower degree of labour market rigidity is also found to induce somewhat larger inflationary pressures.

In addition, we have performed a third counterfactual exercise (not shown here), in which the price markup is assumed not to respond to the oil shock over the initial three years following the disturbance.[21] We find that this does not generate any noticeable relief to inflation and output. This is in line with the variance decomposition results reported earlier showing that

[21] This experiment is technically similar to the absence of interest rate reaction for 3 years implemented in Smets and Wouters (2004). Our results are available to the interested reader upon request.

the markup does not appear to be much affected by oil price shocks. This third counterfactual exercise suggests that markup behaviour does not play an important role in amplifying the macroeconomic consequences of oil price disturbances, as instead argued by Rotemberg and Woodford (1996) for the US economy.

4 Conclusions

This paper uses a maximum likelihood approach to estimate a DSGE model for the euro area. The model combines staggered nominal contracts in both product and labour markets, as well as intertemporal mechanism operating via consumption and investment demand. Moreover, the model is extended to incorporate oil usage in production and endogenous price markups.

The results reported in this paper look reasonable, in light of the euro area economy's structural characteristics as well as the existing empirical and theoretical literature. The DSGE model estimated here has been found to exhibit reasonably good forecasting properties. With regard to structural parameters, we estimate an intermediate degree of wage rigidity, with the average length of wage contracts slightly exceeding 3 quarters. The price markup is found to react positively to the ratio of expected discounted profits to current output, in line with a number of existing studies for the US. Moreover, the postulated Taylor rule displays higher predictability in the post-1990 period, while there has been a decline in the persistence and - to a lesser degree - volatility of oil prices. The present study has uncovered evidence of significant structural change, as captured by parameter instability between pre- and post-1990 sample periods concerning behavioural parameters and exogenous processes driving structural disturbances.

The finding that the price markup reacts positively to the ratio of expected discounted profits to current output deserves further discussion. As stressed by Rotemberg and Woodford (1996), this finding can be rationalised in terms of the "implicit collusion" model of Rotemberg and Saloner (1986), which predicts that implicit collusion can be sustained in a context of higher markups given that any given firm's deviation to a lower markup would induce punishment from its competitors. Given that the ratio of expected discounted profits to current output is normally dominated by short-run developments in the latter, our finding would normally be labelled as a situation of "countercyclical" markups. It is however worth saying that the expression "countercyclical markup" may be misleading in some contexts. To see this, consider a real-world situation in which there is an output expansion, but the price markup rises due to a favourable reassessment of expected profitability prospects. In this case, a model positing a positive relation between the afore-mentioned ratio and markups would rightly predict the increase in the latter. This is the case even if, taken literally, the markup would in the present example behave in a procyclical (as opposed to countercyclical) fashion.

With respect to the propagation mechanisms captured by the empirical DSGE model, the qualitative features of the responses reported here are broadly in line with those obtained in the literature. In the case of oil price shocks, they are found to generate inflationary and recessionary pressures as well as a reduction in real wages and an increase in the price markup. Overall, there is evidence that impulse responses to shocks have become milder in the post-1990 sample period compared with the pre-1990 period. Variance decomposition analysis allows us to detect the smaller importance of shocks affecting monetary policy and oil prices in the post-1990 period, which is in line with the higher predictability of policy and the fall in the persistence and - to a lesser extent - variability of oil disturbances. Among other shocks, technology disturbances have played an important role in explaining movements in real output and inflation, as well as real wages in the medium- to long-term. Preference shocks have increased their share in overall macroeconomic variability in the post-1990 period, becoming of special relevance in the cases of real output, inflation and real wages. Markup disturbances have lost some of the importance in explaining inflation and real wages seen in the pre-1990 period, while still remaining the key force behind price markup developments themselves.

Counterfactual exercises show that oil efficiency gains would alleviate the inflationary and contractionary consequences of an oil shock, while higher wage flexibility would help constrain the contractionary effects of the disturbance at the expense of wider fluctuations in inflation. Finally, the rise in price markups induced by an oil disturbance is not found to considerably amplify the inflationary and contractionary effects of the shock. This finding stands in contrast with the prediction of Rotemberg and Woodford's (1996) model which was calibrated for the US economy.

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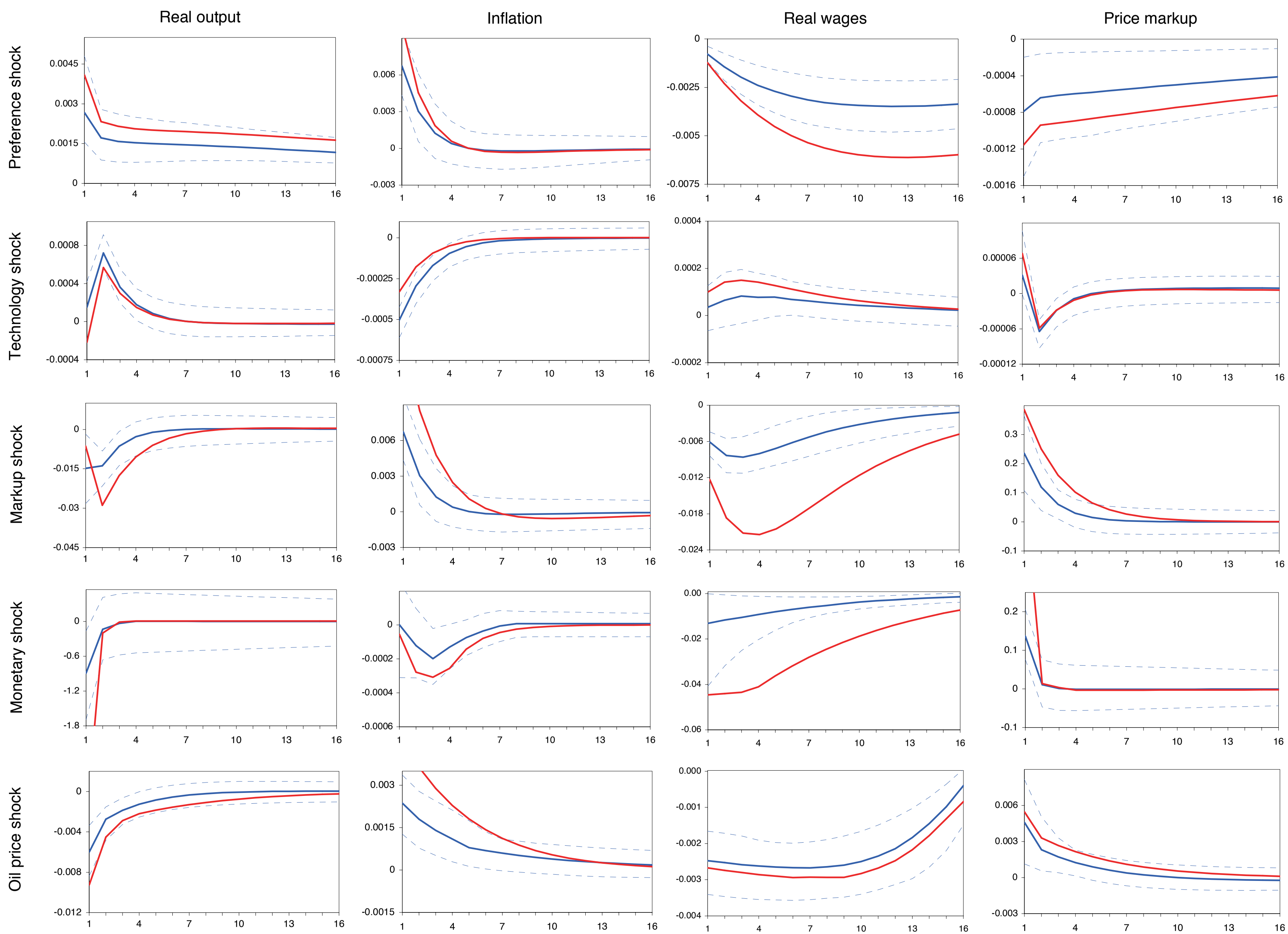

Fig. 1. Impulse responses for selected variables

Table 1. Parameter estimates and stability tests

A) Parameter estimates

Parameter	Pre-1990 estimate	Standard deviation	Post-1990 estimate	Standard deviation
γ	1.10	0.13	1.15	0.41
λ_w	0.73	0.08	0.68	0.15
μ	0.14	0.06	0.15	0.07
ρ_r	0.47	0.09	0.35	0.10
τ_π	0.30	0.07	0.25	0.04
τ_y	1.17	0.06	1.21	0.36
ρ_a	0.50	0.08	0.36	0.07
ρ_z	0.55	0.07	0.60	0.25
ρ_θ	0.64	0.05	0.50	0.02
σ_a (%)	45.92	4.10	34.94	1.76
σ_z (%)	1.33	0.66	1.81	0.17
σ_θ (%)	38.37	11.72	23.38	1.19
σ_r (%)	56.23	18.71	14.45	1.61

B) Tests for parameter stability

Stability of all 13 estimated parameters	W = 363.46***
Stability of the 6 structural parameters	W = 53.13***
Stability of the 3 policy parameters	W = 0.67
Stability of the 4 shocks' standard errors	W = 711.56***

The asterisks */**/*** denote rejection of the null hypothesis of parameter constancy at 10%/5%/1% level.

Table 2. Impulse responses for selected model specifications

accumulated responses to a 100% oil price shock at the end of a given year

	DSGE models				SVAR model		SV model	
	Pre-1990 period		Post-1990 period					
	Real output	inflation	Real output	inflation	Real output	inflation	Real output	inflation
1st year	-1.5	0.9	-0.9	0.5	-1.3	1.4	-0.9	0.7
2nd year	-2.3	1.7	-1.3	0.8	-2.0	1.5	-1.2	0.9
3rd year	-2.7	1.9	-1.4	1.0	-1.9	1.6	-1.3	1.1

Structural VAR (SVAR) and stochastic volatility (SV) models have been estimated over the period 1970:1-2006:1.

Table 3. Real output and inflation effects from alternative models

accumulated responses to a 100% oil price shock

	Real output			Inflation		
	1st year	2nd year	3rd year	1st year	2nd year	3rd year
AWM[1]	-0.2	-0.6	-0.8	1.0	1.8	2.0
Interlink[2]	-0.8	-0.4	0.2	1.2	1.6	1.8
Multimod[3]	-0.2	-0.8	-0.6	3.2	5.4	6.4
NiGEM[4]	-1.6	-1.6	-1.4	0.6	1.0	1.0
QUEST						
- EU Commission (2004)	-1.0	-1.2	-1.4	0.6	1.0	1.2
- Roeger (2005)	-0.7	-1.2	n.a.	0.4	0.9	n.a.

Source: Author's calculations. Adapted from Roeger (2005).

1) Dieppe et al. (2004)

2) Dalsgaard et al. (2001)

3) Hunt et al. (2002)

4) Barrell and Pomeranz (2004)

Table 4. Forecast error variance decompositions

Quarters	Preference shock		Technology shock		Markup shock		Monetary shock		Oil price shock	
ahead	Pre-1990	Post-1990	Pre-1990	Post-1990	Pre-1990	Post-1990	Pre-1990	Post-1990	Pre-1990	Post-1990
					A) Real output					
1	16.3	56.1	31.8	23.1	0.7	5.2	36.4	4.3	14.8	11.3
4	15.1	55.9	38.1	30.3	2.8	2.0	33.1	5.6	11.0	6.3
8	14.2	53.0	41.8	34.0	3.4	1.5	30.7	6.1	9.9	5.4
12	13.9	52.5	43.4	34.7	3.5	1.4	29.9	6.4	9.3	5.0
20	13.8	52.4	44.2	34.7	3.5	1.3	29.5	6.6	9.0	4.9
40	13.8	52.3	43.8	34.8	3.6	1.3	29.5	6.6	9.3	5.0
∞	13.7	52.0	43.6	34.9	3.7	1.3	29.4	6.6	9.5	5.1
					B) Inflation					
1	9.1	38.8	57.2	41.9	5.9	8.5	10.0	4.9	17.8	5.9
4	6.0	38.6	39.1	44.0	38.8	7.0	5.4	4.6	10.6	5.8
8	4.7	39.2	29.8	45.2	52.3	6.0	5.3	4.0	7.9	5.6
12	4.4	39.5	27.1	45.5	56.1	6.0	5.3	3.6	7.1	5.4
20	4.3	39.8	25.8	45.7	58.1	5.8	5.2	3.4	6.8	5.3
40	4.2	39.9	25.6	45.7	58.3	5.8	5.2	3.4	6.7	5.3
∞	4.2	39.9	25.6	45.8	58.3	5.6	5.1	3.4	6.7	5.3
					C) Price markup					
1	1.2	9.2	0.3	0.3	92.9	90.4	4.7	0.0	0.9	0.1
4	0.8	9.4	0.6	1.1	95.1	89.2	2.9	0.2	0.6	0.2
8	0.8	9.8	0.8	1.6	94.9	88.2	2.9	0.3	0.6	0.2
12	0.8	9.9	0.8	1.7	94.9	87.9	2.9	0.3	0.6	0.2
20	0.9	10.0	0.8	1.8	94.8	87.8	2.9	0.3	0.6	0.2
40	0.9	10.0	0.9	1.8	94.7	87.7	2.9	0.3	0.6	0.2
∞	0.9	10.0	0.9	1.8	94.7	87.6	2.9	0.3	0.7	0.2
					D) Real wages					
1	13.0	57.7	2.1	2.8	50.2	11.0	18.5	14.5	16.1	13.9
4	13.2	63.6	6.8	12.4	57.7	10.9	14.1	5.7	8.1	7.4
8	12.8	50.0	11.4	30.9	60.5	10.9	10.3	3.6	5.0	4.6
12	7.8	32.9	22.9	50.5	58.0	10.9	7.5	2.7	3.8	3.0
20	2.2	19.2	48.7	66.5	42.3	10.9	4.2	1.9	2.6	1.6
40	1.2	7.6	73.1	80.0	22.6	10.7	1.7	1.0	1.4	0.7
∞	1.2	7.6	73.1	80.0	22.6	10.7	1.7	1.0	1.4	0.7

The values reported here decompose the forecast error variance in each variable at each quarterly forecast horizon into percentages attributable to each disturbance.

Table 5. Forecast accuracy (2003:2-2006:1)

Quarters ahead	1	2	3	4
A) Real output				
RMSE: DSGE	0.39	0.87	1.08	1.46
RMSE: VAR(1)	0.28	0.52	0.72	0.90
RMSE: VAR(2)	0.29	0.60	0.91	1.18
S: DSGE vs. VAR(2)	-3.22**	-2.12	-1.79	-1.68
S: DSGE vs. VAR(1)	-3.22**	-2.12	-1.79	-1.68
B) Real private consumption				
RMSE: DSGE	0.31	0.67	0.08	1.06
RMSE: VAR(1)	0.23	0.04	0.05	0.06
RMSE: VAR(2)	0.26	0.46	0.67	0.86
S: DSGE vs. VAR(2)	-3.01*	-2.10	-1.83	-1.72
S: DSGE vs. VAR(1)	-3.00*	-2.10	-1.83	-1.71
C) Inflation				
RMSE: DSGE	0.43	0.78	0.39	0.54
RMSE: VAR(1)	0.34	0.36	0.26	0.17
RMSE: VAR(2)	0.31	0.53	0.48	0.44
S: DSGE vs. VAR(2)	-3.09*	-2.12	-1.80	-1.68
S: DSGE vs. VAR(1)	-3.11*	-2.12	-1.81	-1.71
D) Short-term interest rate				
RMSE: DSGE	0.77	1.17	2.50	3.01
RMSE: VAR(1)	0.56	1.12	1.67	2.22
RMSE: VAR(2)	0.33	0.83	1.43	2.07
S: DSGE vs. VAR(2)	-3.19**	-2.24*	-1.80	-1.71
S: DSGE vs. VAR(1)	-3.20**	-2.22	-1.81	-1.71

The asterisks */**/*** denote rejection of the null hypothesis of parameter constancy at 10%/5%/1% level.

Table 6. Counterfactual exercises from the DSGE model

accumulated responses to a 100% oil price shock at the end of a given year

	Baseline model		20% oil efficiency gain		1 quarter shorter wage contract duration	
	Real output	inflation	Real output	inflation	Real output	inflation
1st year	-0.93	0.47	-0.73	0.35	-0.83	0.48
2nd year	-1.34	0.84	-1.05	0.62	-1.23	0.92
3rd year	-1.40	1.02	-1.11	0.74	-1.29	1.16

The responses correspond to the model estimated over the period 1990:1-2006:1.

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