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TECHNOLOGY SHOCKS AND ROBUST SIGN RESTRICTIONS IN A EURO AREA SVAR*

BY GERT PEERSMAN AND ROLAND STRAUB[1]

Ghent University, Belgium; European Central Bank, Germany

We use a model-based identification strategy to estimate the impact of technology shocks on hours worked and employment in the euro area. The sign restrictions applied in the vector autoregression (VAR) analysis are consistent with a large class of dynamic stochastic general equilibrium (DSGE) models and are robust to parameter uncertainty. The results are in line with the conventional Real Business Cycle (RBC) interpretation that hours worked rise as a result of a positive technology shock. By comparing the sign restrictions method to the long-run restriction approach of Galí (*Quaterly Journal of Economics* (1992) 709–38), we show that the results do not depend on the stochastic specification of the hours worked series or the data sample but only on the identification scheme.

1. INTRODUCTION

The direction and the magnitude of the response of hours worked and employment following a technology shock are subject to an active controversy in the academic literature. The debate has its origin in the Real Business Cycle (RBC) research program. The workhorse of this program, as introduced in the seminal paper by Kydland and Prescott (1982), has been a flexible-price, dynamic general equilibrium model with optimizing agents. The motivation behind this approach was to explain aggregate fluctuations in actual economies using the RBC model subject to stochastic technology shocks. In the RBC framework, technology shocks act as labor demand shifters and have a positive impact on both per capita hours worked and output.

This prediction of the RBC model has been challenged by Galí (1999). Using a structural vector autoregression (VAR) with long-run restrictions, Galí (1999) provides empirical evidence that hours worked fall as a result of a positive technology shock in the United States. Furthermore, Galí (1999) demonstrates that the latter result is in line with the prediction of a standard New Keynesian (NK)

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model with sticky prices. Price rigidities in combination with an active monetary policy rule imply that aggregate demand cannot change immediately following a positive technology shock, forcing firms to contract employment.[2] Similar results are presented in Shea (1998), Basu et al. (2006), Francis and Ramey (2005), and Francis et al. (2003).

Recent studies, however, argue that the results provided in Galí (1999) are not robust to certain modifications. First, in Galí's setup, only technology shocks have a long-run impact on labor productivity. Uhlig (2004) shows, however, that capital income taxation shocks or long-run shifts in the social attitudes to the workplace can also be a source of changes in long-run labor productivity. Second, Faust and Leeper (1997) demonstrate that using long-run restrictions substantial distortions are possible because of small sample biases and measurement errors. Furthermore, in a similar framework as Galí (1999), Christiano et al. (2003) test the sensitivity of the results to the stochastic specification of the hours worked series. Interestingly, if per capita hours worked is modeled as a difference stationary process, the results confirm that hours worked will fall following a positive technology shock. But in case the system is estimated using the level of the hours worked series, the impulse responses are in line with the predictions of the RBC model. Other papers casting doubts on the robustness of Galí's results are Bils (1998) and Chang and Hong (2006). A comprehensive overview of the empirical and theoretical debate on the effects of technology shocks is presented in Galí and Rabanal (2004).

In this article, we propose an alternative, model-based identification strategy to estimate the effects of technology shocks on hours worked and employment in the euro area. In particular, we utilize conditional moments of dynamic stochastic general equilibrium (DSGE) models, which hold under both flexible and sticky prices/wages and a given sensible range of structural parameter values, as sign restrictions in a structural VAR (SVAR). In order to identify the shocks, however, we use only a minimum set of sign restrictions that are robust to model and parameter uncertainty. Since we are mainly interested in the response of hours worked following a technology shock, we do not apply any restriction on its response. Hence, the estimated reaction of hours worked in our VAR allows us to discriminate between the NK and the RBC models. As an identification approach, sign restrictions were first used by Faust (1998), Canova and De Nicoló (2002), and Uhlig (2005) to identify monetary policy shocks. Peersman (2005) showed how they can be used to also identify aggregate supply, demand, and oil price shocks.

In line with the RBC hypothesis, we find a significant increase in hours worked following a positive technology shock. The results are robust whether we estimate the model in levels or first differences or when we use total employment instead of hours. Interestingly, when using long-run restrictions, as defined in Galí (1999), to identify technology shocks in our data set, hours worked fall on impact. This

[2] The ability to hold inventory might change the nature of firms' response to technology shocks even under sticky prices. Chang et al. (2004) demonstrate that even when the prices are fixed, firms may want to produce more, hire workers, and build up inventories for future sales in response to a favorable technology shock.

indicates that our results do not depend on the chosen data set or the stochastic specification of hours worked but rather on the identification scheme[3]. However, although our results assign a more important role for technology shocks in explaining variations in output and hours worked as compared with the results by Galí (1999), they do still question the original RBC hypothesis that technology shocks are the main source of business cycle fluctuations. In particular, technology shocks can only explain less than 25% of variations in output and hours worked in a five-year horizon.

In related work, Dedola and Neri (2007) also find a positive impact of technology shocks on hours worked by applying sign restrictions in a VAR for the United States and Germany. Their identification scheme relies, however, on restrictions on labor productivity, output, investment, consumption, and real wages and is therefore more restrictive than our approach. Furthermore, the restrictions in Dedola and Neri (2007) are not sufficient to disentangle productivity shocks from government spending shocks in the NK models with limited asset market participation (see, e.g., Galí et al., 2007). Note that these types of NK models generate positive effects on consumption following a government spending shock in line with the empirical evidence, as discussed, for instance, in Fatás and Mihov (2001) and Coenen and Straub (2005).

The rest of this article is organized as follows. In Section 2, we describe our model-based identification strategy. First, we set up a baseline DSGE model that nests both an NK sticky price/wage model and an RBC model as a special case and utilize the impulse responses of the models to derive a minimum set of sign restrictions that are robust to model and parameter uncertainty. In Section 3, we present the results of the SVAR, provide robustness analysis, and identify the underlying cause for our contrasting results. We also discuss the importance of technology shocks for the euro area business cycle. Finally, Section 4 concludes the analysis.

2. IDENTIFICATION

In this section, we present the equilibrium conditions of a standard DSGE model that are utilized to derive the sign restrictions imposed in the empirical exercise. As we will discuss in the next section, the model presented below nests both an NK sticky price/wage model and an RBC model as a special case.

2.1. *Households.* In the first step, we present the optimization problem of a representative household denoted by h. The household maximizes lifetime utility by choosing consumption $C_{h,t}$, financial wealth in form of bonds $B_{h,t+1}$, and the next period's capital stock $K_{h,t+1}$:

$$\max E_0 \sum_{t=0}^{\infty} \beta^t \left\{ \frac{1}{1-\sigma} C_{h,t}^{1-\sigma} - \frac{\varepsilon_t^n}{1+\zeta} N_{h,t}^{1+\zeta} \right\}, \tag{1}$$

[3] Canova et al. (2005) and Peersman and Straub (2006) also confirm a rise of hours using sign restrictions in more recent work.

where β is the discount factor, σ denotes the coefficient of relative risk aversion, and ζ is the inverse of the elasticity of work effort with respect to the real wage. The household's utility depends positively on the level of consumption, $C_{h,t}$, and negatively on hours worked, $N_{h,t}$. We denote a serially correlated shock to labor supply with ε_t^n.[4] The intertemporal budget constraint of the representative household is given by

$$\begin{aligned} C_{h,t} + I_{h,t} + R_t^{-1}\frac{B_{h,t+1}}{P_t} \\ = \frac{W_{h,t}}{P_t}N_{h,t} + R_t^K K_{h,t} + D_{h,t} + T_{h,t} + \frac{B_{h,t}}{P_t}, \end{aligned} \tag{2}$$

and the capital accumulation process by

$$K_{h,t+1} = (1-\delta)K_{h,t} + I_{h,t}. \tag{3}$$

Here, R_t is the nominal interest rate, R_t^K is the rate of return to capital, $W_{h,t}$ is the nominal wage, $K_{h,t}$ is the capital stock, $T_{h,t}$ are lump-sum taxes paid to the fiscal authority, P_t is the price level, and $D_{h,t}$ is the dividend income. In the following sections, we will assume the existence of state-contingent securities that are traded among households in order to insure households against variations in household-specific wage income. As a result, where possible, we neglect the indexation of individual households.

The maximization of the objective function with respect to consumption, bond holding, and next period's capital stock can be summarized by the following two standard Euler equations:

$$\beta R_t \mathrm{E}_t\left[\frac{{C_t}^{\sigma}}{{C_{t+1}}^{\sigma}}\frac{P_t}{P_{t+1}}\right] = 1 \tag{4}$$

and

$$1 = \beta \mathrm{E}_t\left[\frac{{C_t}^{\sigma}}{{C_{t+1}}^{\sigma}}\left(1-\delta+R_{t+1}^K\right)\right]. \tag{5}$$

2.2. *Firms.* There are two types of firms, a continuum of monopolistically competitive firms indexed by $f \in [0, 1]$, each of which produces a single-differentiated intermediate good, $Y_{f,t}$, and a distinct set of perfectly competitive firms, which combine all the intermediate goods into a single final good, Y_t.

[4] Recent literature presents empirical evidence for the importance of labor supply shifts in explaining business cycle fluctuation. Chang and Schorfheide (2003), for example, show that labor supply shocks account for about 30% of the cyclical fluctuation in the U.S. hours worked series. Smets and Wouters (2003) report that after two-and-a-half years, about 33% of the variation of euro area output is caused by labor supply shocks.

2.2.1. *Final-good firms.* The final-good-producing firms combine the differentiated intermediate goods $Y_{f,t}$ using a standard Dixit–Stiglitz aggregator

$$Y_t = \left(\int_0^1 Y_{f,t}^{\frac{1}{1+\lambda_p}} \, df \right)^{1+\lambda_p}, \tag{6}$$

where λ_p is a parameter determining the degree of imperfect competition in the goods market. Minimizing the cost of production subject to the aggregation constraint (Equation (6)) results in demand for the differentiated intermediate goods as a function of their price $P_{f,t}$ relative to the price of the final good P_t,

$$Y_{f,t} = \left(\frac{P_{f,t}}{P_t} \right)^{-\frac{1+\lambda_p}{\lambda_p}} Y_t, \tag{7}$$

where the price of the final good P_t is determined by the following index:

$$P_t = \left(\int_0^1 P_{f,t}^{-\frac{1}{\lambda_p}} \, df \right)^{-\lambda_p}.$$

2.2.2. *Intermediate-goods firms.* Each intermediate-goods firm f produces its differentiated output using a production function of a standard Cobb–Douglas form:

$$Y_{f,t} = A_t N_{f,t}{}^{1-\alpha} K_{f,t}^{\alpha}, \tag{8}$$

where A_t is a technology shock and α the capital share of output in the steady state. Taking the rental cost of capital, R_t^k, and the aggregate wage index, W_t, as given, cost minimization subject to the production technology (Equation (8)) yields first-order conditions for the inputs that can be expressed as relative factor demands and nominal marginal cost MC_t:

$$\frac{K_{f,t}}{N_{f,t}} = \left(\frac{\alpha}{1-\alpha} \right) \frac{W_t}{R_t^k}$$

and

$$MC_t = \frac{1}{A_t \, \alpha^{\alpha} (1-\alpha)^{(1-\alpha)}} \, W_t^{(1-\alpha)} \big(R_t^k \big)^{\alpha}.$$

2.2.3. *Price setting.* Following Calvo (1983), intermediate-goods-producing firms receive permission to optimally reset their price in a given period t with probability $1 - \theta_p$. All firms that receive permission to reset their price choose the same price $P_{f,t}^*$. Each firm f receiving permission to optimally reset its price in period t maximizes the discounted sum of expected nominal profits,

$$\mathrm{E}_t \left[\sum_{k=0}^{\infty} \theta_p^k \, \chi_{t,t+k} \, D_{f,t+k} \right],$$

subject to the demand for its output (Equation (7)), where $\chi_{t,t+k}$ is the stochastic discount factor of the households owing the firm and

$$D_{f,t} = P_{f,t}\, Y_{f,t} - MC_t Y_{f,t}$$

are period-t nominal profits that are distributed as dividends to the households.

Hence, we obtain the following first-order condition for the firm's optimal price-setting decision in period t:

(9)

$$P^*_{f,t}\, Y_{f,t} - (1+\lambda_p)\, MC_t Y_{f,t} + \mathrm{E}_t \left[\sum_{k=1}^{\infty} \theta_p^k\, \chi_{t,t+k}\, Y_{f,t+k} \big(P^*_{f,t} - (1+\lambda_p)\, MC_{t+k} \big) \right] = 0.$$

With the intermediate-goods prices $P_{f,t}$ set according to Equation (9), the evolution of the aggregate price index is then determined by the following expression:

$$P_t = \left((1-\theta_p) \big(P^*_{f,t} \big)^{-\frac{1}{\lambda_p}} + \theta_p \, (P_{f,t-1})^{-\frac{1}{\lambda_p}} \right)^{-\lambda_p}.$$

2.3. *Wage Setting.* There is a continuum of monopolistically competitive unions indexed over the same range as the households, $h \in [0, 1]$, which act as wage setters for the differentiated labor services supplied by the households, taking the aggregate nominal wage rate W_t and aggregate labor demand N_t as given. Following Calvo (1983), unions receive permission to optimally reset their nominal wage rate in a given period t with probability $1 - \theta_w$. All unions that receive permission to reset their wage rate choose the same wage rate $W^*_{h,t}$. Each union h that receives permission to optimally reset its wage rate in period t maximizes the household's lifetime utility function (Equation (1)), subject to its intertemporal budget constraint (Equation (2)) and the demand for labor services of variety h, the latter being given by

$$N_{h,t} = \left(\frac{W_{h,t}}{W_t} \right)^{-\frac{1+\lambda_w}{\lambda_w}} N_t,$$

where λ_w is a parameter determining the degree of imperfect competition in the labor market. As a result, we obtain the following first-order condition for the union's optimal wage-setting decision in period t:

(10)

$$\frac{W^*_{h,t}}{P_t} - (1+\lambda_w)\, \varepsilon^n_t\, MRS_t + \mathrm{E}_t \sum_{k=1}^{\infty} \theta_w^k\, \beta^k \left[\frac{W^*_{h,t}}{P_{t+k}} - (1+\lambda_w)\, \varepsilon^n_{t+k}\, MRS_{t+k} \right] = 0,$$

where MRS_{t+k} stands for the marginal rate of substitution

$$MRS_t = {N_{h,t}}^{\zeta}\, {C_{h,t}}^{\sigma}.$$

The aggregate labor demand, N_t, and the aggregate nominal wage rate, W_t, are determined by the following Dixit–Stiglitz indices:

$$N_t = \left(\int_0^1 (N_{h,t})^{\frac{1}{1+\lambda_w}} \, dh \right)^{1+\lambda_w}$$

and

$$W_t = \left(\int_0^1 (W_{h,t})^{-\frac{1}{\lambda_w}} \, dh \right)^{-\lambda_w}.$$

With the labor-specific wage rates $W_{h,t}$ set according to Equation (10), the evolution of the aggregate nominal wage rate is then determined by the following expression:

$$W_t = \left((1-\theta_w)\left(W_{h,t}^*\right)^{-\frac{1}{\lambda_w}} + \theta_w (W_{h,t-1})^{-\frac{1}{\lambda_w}} \right)^{-\lambda_w}.$$

2.4. *Market Clearing and Shock Processes.* The labor market is in equilibrium when the demand for the index of labor services by the intermediate-goods firms equals the differentiated labor services supplied by households at the wage rates set by unions. Similarly, the market for physical capital is in equilibrium when the demand for capital services by the intermediate-goods firms equals the capital services supplied by households at the market rental rate. Finally, the final-good market is in equilibrium when the supply by the final-good firms equals the demand by households

$$Y_t = C_t + I_t + G_t,$$

where G_t is an aggregate demand shock, for example, a shock to government spending. The model is simulated in its log-linearized form, that is, small letters will characterize in the following percentage deviations from the steady state. The exogenous technology, labor supply, and aggregate demand shocks follow an AR(1) process described by the following equations:

$$a_t = \rho^a a_{t-1} + \eta_t^a, \qquad (11)$$

$$\varepsilon_t^n = \rho^n \varepsilon_{t-1}^n + \eta_t^n,$$

and

$$g_t = \rho^g g_{t-1} + \eta_t^g.$$

Finally, monetary policy follows a standard log-linearized Taylor rule:

$$r_t = \rho^r r_{t-1} + (1-\rho^r)\left(\phi^y y_t + \phi^\pi \pi_t\right) + \eta_t^r, \qquad (12)$$

where ρ^r is a parameter determining the degree of interest rate smoothing, η_t^r is a white noise monetary policy shock, and ϕ^y and ϕ^π represent the elasticity of the interest rate to output and inflation, respectively.

2.5. *Sign Restrictions and Robustness Analysis.* For testing the models, one could estimate the RBC and the NK model using, for example, Bayesian methods and thereby computing the odds of the respective models. Such a strategy is chosen, for example, in Galí and Rabanal (2004). However, the prototype RBC and NK model are likely to be too stylized to be taken directly to the data. Hence, we will utilize the models only to derive a set of robust sign restrictions and will base the empirical analysis on a more flexible VAR specification. In this section, we will discuss the derivation of the sign restrictions from the impulse response functions of the theoretical model. First, note that when prices and wages are perfectly flexible, that is, $\theta_p = 0$ and $\theta_w = 0$, the equilibrium conditions in the goods and the labor market converge to

$$P_t = (1 + \lambda_p)\, MC_t$$

and

$$\frac{W_t}{P_t} = (1 + \lambda_w)\, \varepsilon_t^n\, MRS_t.$$

We will simulate the model under both scenarios, assuming that the economy is subject to nominal and real rigidities, as in the NK case, as well as to flexible prices and wages and perfect competition in goods and labor market (i.e., $\lambda_p = 0$ and $\lambda_w = 0$), as in the standard RBC model.

In order to test whether a flexible-price RBC model or an NK model with nominal and real rigidities is better in matching the dynamics present in the data, we use the methodology discussed, for example, in Canova (2002), Pappa (2004), and Peersman (2005). In the first step, we identify robust implications in each of the two models that are not sensitive to variations of structural parameters. In order to do so, we define a range for each of the structural parameters by conducting a brief survey of the related empirical literature. Papers such as Smets and Wouters (2003) use the Bayesian methods to estimate medium-scale DSGE models providing the corresponding posterior distribution of the structural parameters. Similar models using alternative estimation techniques have been analyzed by Christiano et al. (2005), Altig et al. (2005), Onatski and Williams (2004), Rabanal and Rubio-Ramirez (2005), and Coenen and Straub (2005).

We use the estimated posterior distribution of structural parameters in these models as a benchmark but have extended the ranges beyond the 90% interval. For example, the preference parameter driving the labor supply utility ζ is allowed to vary in the interval [0, 10], the risk-averse coefficient $\sigma \in [1, 10]$, and the Calvo parameters determining the degree of nominal wage and price rigidities θ_p and θ_w are both allowed to vary in the interval [0.01, 0.95].

For the monetary policy rule, we delimit the range of parameters to cover the values generally discussed in the Taylor rule literature. In order to ensure determinacy of the model, we restrict the inflation response to the range between [1, 3], whereas the output response and the degree of interest rate smoothing are allowed to vary in the interval [0, 1].

We set the range for the subjective discount rate β between [0.985, 0.995], implying an annual steady-state real interest rate between 2% and 6%. The interval

TABLE 1

PARAMETER VALUES AND RANGES

Parameter	Description	Range
β	Discount factor	[0.985, 0.995]
σ	Risk-aversion coefficient	[1, 10]
ζ	Preference parameter	[0, 10]
λ_p	Degree of monopolistic competition in the goods market	[0, 0.5]
λ_w	Degree of monopolistic competition in the labor market	[0, 0.5]
θ_p	Degree of nominal rigidities in the goods market	[0.01, 0.95]
θ_w	Degree of nominal rigidities in the labor market	[0.01, 0.95]
α	Capital share	[0.2, 0.5]
δ	Depreciation rate	[0.01, 0.05]
ϕ^y	Coefficient on output in the monetary policy rule	[0, 1]
ϕ^π	Coefficient on inflation in the monetary policy rule	[1, 3]
ρ^r	Degree of interest rate smoothing	[0, 1]
ρ^a	Persistence of technology shocks	[0.5, 0.99]
ρ_n	Persistence of labor supply shocks	[0.5, 0.99]
ρ_g	Persistence of aggregate demand shocks	[0.5, 0.99]

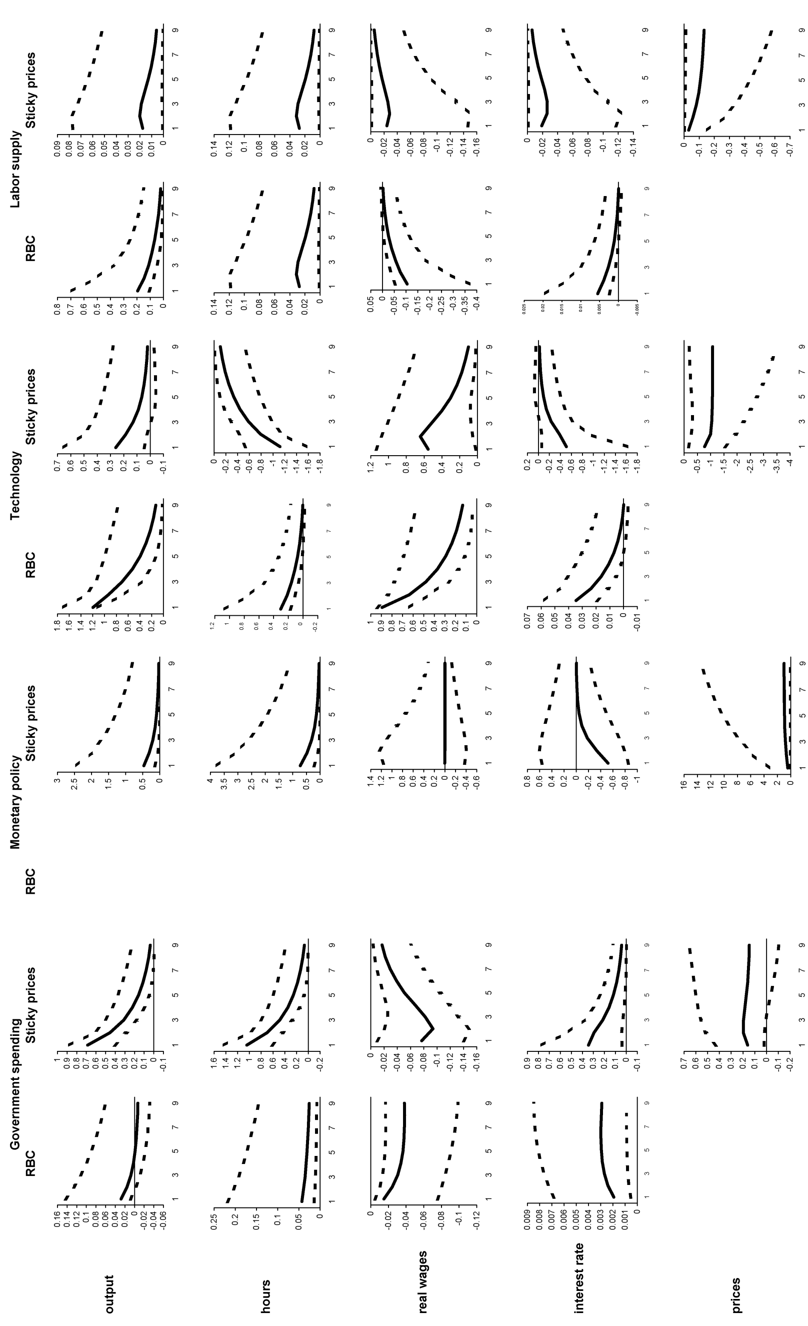
NOTE: In the RBC model, $\theta_p = 0$, $\theta_w = 0$, $\lambda_p = 0$, and $\lambda_w = 0$.

determining the capital share in the Cobb–Douglas production function α is set between [0.2, 0.5], including a steady-state share of capital income of 30% usually assumed in the literature. We also allow for variation in the depreciation rate $\delta \in [0.01, 0.05]$, price mark-up $\lambda_p \in [0, 0.5]$, and the wage mark-up $\lambda_w \in [0, 0.5]$. Finally, and in line with the empirical literature, we restrict the persistence of the shocks to the interval [0.5, 0.99]. The intervals for all parameter values are reported in Table 1.

After defining a sensible range for the parameter values, we proceed with the simulation exercise. First, we assume that the parameters are uniformly distributed over the selected parameter range. Second, we draw a random value for each parameter from the presented intervals and calculate the corresponding impulse response functions of the model. This exercise is repeated for 500,000 simulations. The median and the 10th and 90th percentiles of all the conditional responses are shown in Figure 1.

The impulse responses of both models are in line with the well-known results in the literature. In the RBC model, technology shocks act as labor demand shifters and result in an increase of the equilibrium real wage, output, hours worked, and interest rate. In contrast, an exogenous shock to labor supply has a negative impact on real wages but a positive impact on output and interest rate. Government spending shocks generate a negative wealth effect, leading to an increase in hours worked and a corresponding fall of the real wage.

In the NK model, positive technology shocks have a negative impact on hours worked but, similarly to the RBC model, a positive impact on real wages. On the other hand, the sign of the impulse response functions to labor supply shocks appears to be insensitive to the existence of nominal rigidities, implying a positive response of output and a negative response of prices and real wages. As labor supply shocks and technology shocks have asymmetric effects on real wages, a



NOTE: Median of simulations with 90th and 10th percentiles.

FIGURE 1

THEORETICAL IMPULSE RESPONSE FUNCTIONS

TABLE 2
SIGN RESTRICTIONS

	Output	Prices	Interest Rate	Hours	Wages
Monetary policy	↑	↑			
Aggregate demand	↑	↑			
Technology	↑	↓			↑
Labor supply	↑	↓			↓

feature common in both RBC and NK models, we are able to use the latter as a sign restriction in our VAR. In contrast to monetary and government spending shocks, expansionary technology and labor supply shocks have a negative impact on the price level. Therefore, we apply the latter restriction on prices to distinguish technology shocks from expansionary monetary policy and aggregate demand shocks in our empirical exercise.[5] All sign restrictions that will be used to identify a technology shock in our empirical VAR are summarized in Table 2.

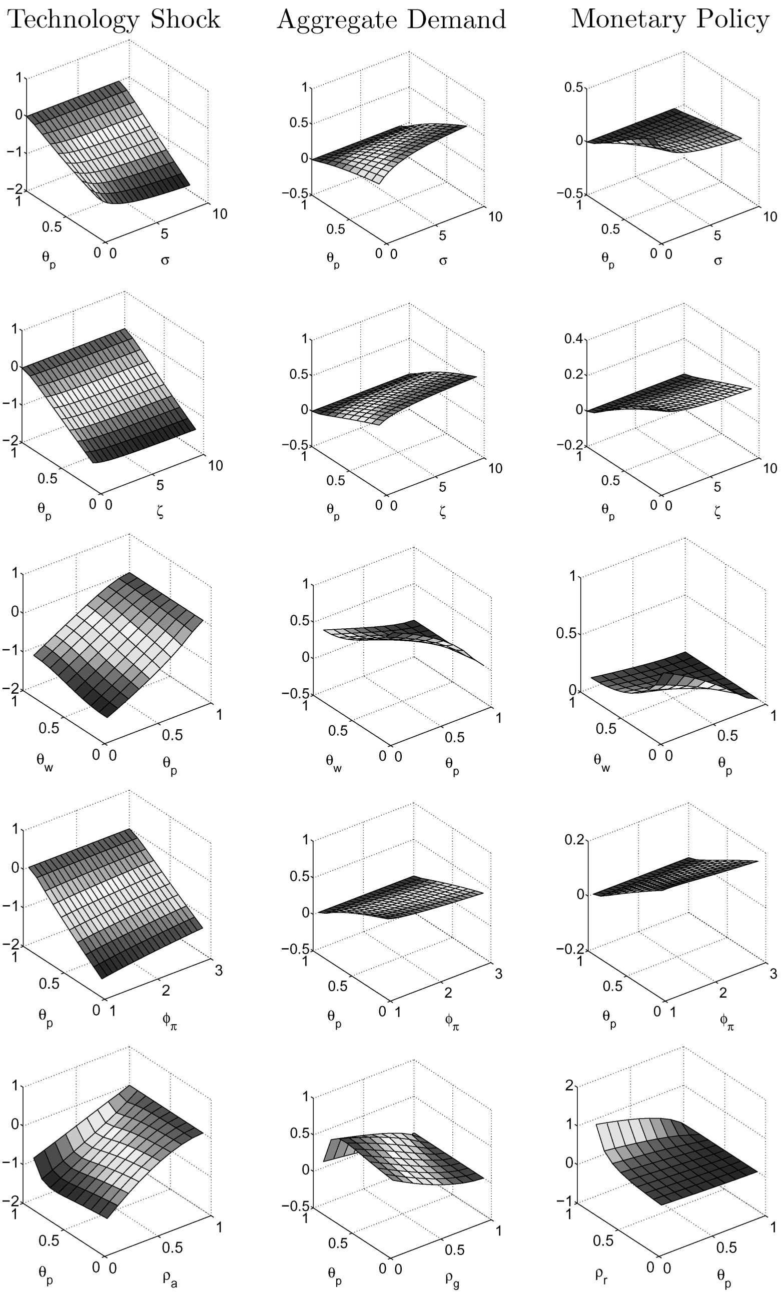
In order to provide some further robustness checks of our selected sign restrictions, we trace out boundaries in the parameter space, beyond the predefined ranges, across which the signs of the impact response of relevant variables, such as the price level and real wages, are switching. The latter procedure gives us a better hint than the multidimensional Monte Carlo analysis about which parameter constellations are crucial for the chosen identification scheme. The outcome of this exercise is shown in Figures 2(a) and (b). We present the results in three-dimensional figures by varying two crucial parameters at the same time, while leaving the remaining set of parameters at the mean value of the parameter range presented in Table 1. For reasons of legibility, the analysis only focuses on the sign of the impact responses of the NK model.[6]

First, Figure 2(a) plots the impact response of the price level to a technology, aggregate demand, and monetary policy shock. The response of the price level is used to separate monetary policy and aggregate demand shocks from technology shocks. For all possible combinations of parameter values, the response of prices is always negative following a technology shock and always positive following a monetary policy and aggregate demand shock. In the upper panel, for example, we show that the sign of the price level response is insensitive for different combinations of nominal rigidities θ_p and intertemporal elasticity of substitution σ. Similar results hold for other combinations of structural parameters.[7]

[5] This is demonstrated in Figure 1 only for the NK case. In models with flexible prices and imperfect competition, the aggregate price level is indeterminate and only relative prices are pinned down in equilibrium. We generally allow, however, for a possible zero impact of shocks in our empirical approach because restrictions are imposed as $\geq$ or $\leq$.

[6] Note that in the RBC model, the sign of the response of real wages is not sensitive to variations in structural parameters, as technology shocks act as labor demand and labor supply shocks act as labor supply shifters.

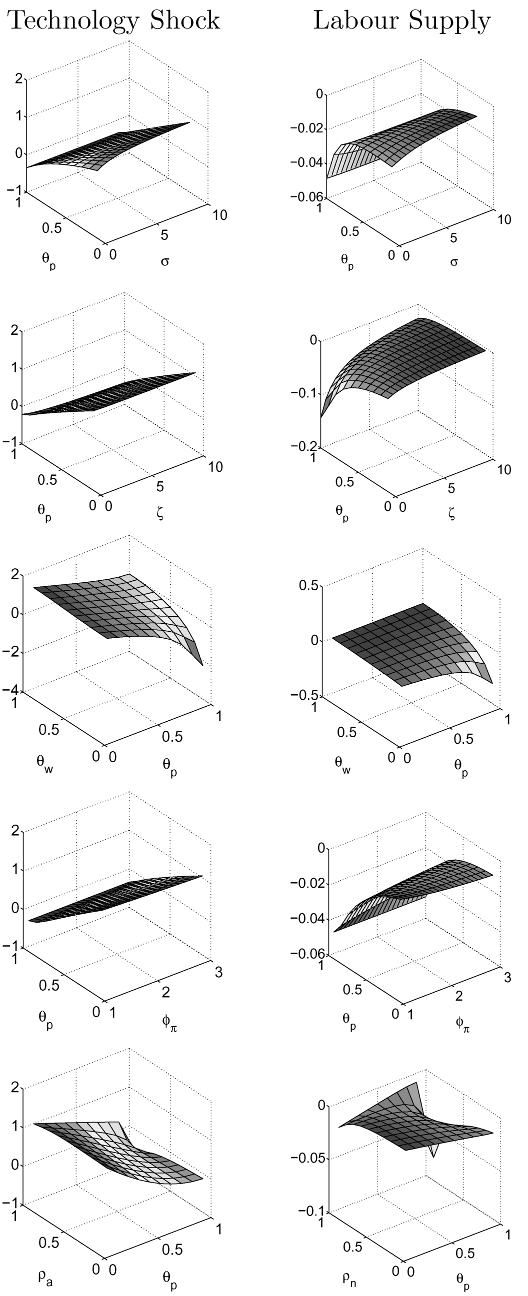
[7] We focus only on a subset of variables, but similar results hold for combinations of other structural parameters. The latter results are available upon request.



NOTES: The figure shows the impact of the price level by varying two structural parameter at the same time, while leaving the remaining set of parameters at the mean value of the parameter range presented in Table 1.

FIGURE 2A

FURTHER SENSITIVITY ANALYSIS/PRICE-LEVEL RESPONSE ON IMPACT



NOTES: The figure shows the impact response of the real wage by varying two structural parameter the same time, while leaving the remaining set of parameters at the mean value of the parameter range presented in Table 1.

FIGURE 2B

FURTHER SENSITIVITY ANALYSIS/REAL WAGE RESPONSE ON IMPACT

In Figure 2(b), we focus on the impact response of real wages to exogenous increases in labor supply and technology. The real wage response is crucial to differentiate labor supply shocks from technology shocks. As indicated in the right column, the negative impact of labor supply shocks on real wages is always robust to variations in key parameters. The response of real wages to a technology shock, however, switches signs under certain parameter combinations. In particular, if price rigidity θ_p is high, the real wage response can become negative.[8] The latter is reinforced by high intertemporal elasticity of substitution (denoted by the inverse of σ), high labor supply elasticity (denoted by the inverse of ζ), and/or low wage rigidity (denoted by θ_w). The rationale behind the results is as follows: In standard NK models, technology shocks trigger generally an inward shift of both labor supply and labor demand curve, resulting in a fall in equilibrium hours worked. As a result, real wages increase only if the shift in the labor supply curve (induced by the positive wealth effect) is more pronounced than the shift in the labor demand curve (as a result of the dampened response of real aggregate demand). As indicated in the previous section, this is the case for a wide range of model parameters. However, high labor supply elasticity and/or high intertemporal elasticity of substitution will dampen the relative importance of the wealth effect induced by the technology shock. Also, higher price rigidity implies, ceteris paribus, a weaker response of real aggregate demand to an exogenous increase in technology, resulting in a more pronounced downward shift of the labor demand curve. In the two lower panels of Figure 2(b), we also illustrate the sensitivity of the impact response of real wages to variations in the persistence of the technology shock ρ_a and to the interest rates elasticity with respect to inflation ϕ^{π}. As indicated, low shock persistence implies a weaker wealth effect on households, increasing in combination with high price rigidity the probability of a negative real wage response. Similarly, a weak response of policy rates to inflation is more conducive with a negative reaction of real wages following an expansionary technology shock.

In equilibrium, a combination of the discussed factors can induce a negative response of real wages. The probability that this happens is, however, low in the presented model, as is also indicated in the Monte Carlo exercise in the previous section.

In order to summarize, the RBC and the NK models differ with regards to the sign of the impulse response function of hours worked following a technology shock but resemble in a number of other conditional moments, allowing us to derive a sufficient set of sign restrictions that are fairly robust to model and parameter uncertainty. The corresponding sign restrictions for our empirical analysis can be found in Table 2.

3. EMPIRICAL EVIDENCE

In this section, we present the results of the SVAR using euro area data for the sample period 1982–1 to 2002–4. All data are taken from the area-wide model

[8] The boundary is around $\theta_p = 0.9$ and varies depending on the value of other key structural parameters presented in Figure 2(b).

TABLE 3
SUMMARY OF VAR SPECIFICATIONS

Variables	Identification	Impact on Hours	Results
y_t, p_t, i_t, h_t, w_t	Sign restrictions	Positive	Figure 3, col. 1
y_t, p_t, i_t, e_t, w_t	Sign restrictions	Positive	Figure 3, col. 2
$d(y_t - h_t), dp_t, i_t, dh_t, dw_t$	Long-run restrictions	Negative	Figure 4, col. 1
$d(y_t - h_t), dp_t, i_t, dh_t, dw_t$	Sign restrictions	Positive	Figure 4, col. 2
$y_t, p_t, i_t, h_t, w_t, p_t^{oil}$	Sign restrictions	Positive	Not reported

NOTE: $y_t =$ output; $p_t =$ prices; $i_t =$ interest rate; $h_t =$ hours; $e_t =$ employment; $p_t^{oil} =$ oil price.

(Fagan et al., 2001). Hours worked is a series constructed by the European Central Bank (ECB) Euro Area Department. The latter is only available from 1981 onward, which determines our sample period. In Section 3.1, we present the baseline results. The robustness of these results and the source of our findings are discussed in Sections 3.2 and 3.3. Section 3.4 analyzes the exogeneity of the identified shocks and, finally, the importance of technology shocks for aggregate fluctuations is investigated in Section 3.5. A summary of all the VAR specifications that are considered in this section is reported in Table 3.

3.1. *Baseline Results.* Consider the following specification for a vector of endogenous variables Y_t:

$$Y_t = c + \sum_{i=1}^{n} A_i Y_{t-i} + B\varepsilon_t, \tag{13}$$

where c is an $(n \times 2)$ matrix of constants and linear trends, A_i is an $(n \times n)$ matrix of autoregressive coefficients, and ε_t is a vector of structural disturbances. The endogenous variables, Y_t, that we include in the VAR are real gross domestic product (GDP) (Y_t), the GDP deflator (p_t), short-term nominal interest rate (i_t), hours worked (h_t), and real wages (w_t). We estimate this VAR model in log levels, except the interest rate, which is included in percent. Lag length is determined by standard likelihood ratio tests and Akaike's information criterion (AIC), which turns out to be three.[9]

Within this VAR, we only identify technology shocks. In order to identify these shocks, we use the restrictions reported in Table 2.[10] Specifically, a positive

[9] By doing the analysis in levels, we allow for implicit cointegration relationships in the data and still have consistent estimates of the parameters (Sims et al., 1990). In Section 3.3, we check the robustness of our results when we use a first-difference specification of the VAR. We can, however, not reject the hypothesis of the existence of a cointegration relation in the level specification when we perform the tests on the reduced-form point estimates using the procedure of Johansen and Juselius in CATS. All results reported in this article are, however, Bayesian. Including or excluding the time trend has no qualitative impact on the results. Given our sample period (downward trend in inflation and interest rate), a linear trend is appropriate for the first-difference specification. The results are also not sensitive with respect to the number of lags.

[10] For the estimation results of, respectively, the monetary policy, aggregate demand, and labor supply shock, we refer to the ECB Working Paper no. 373 version of this article.

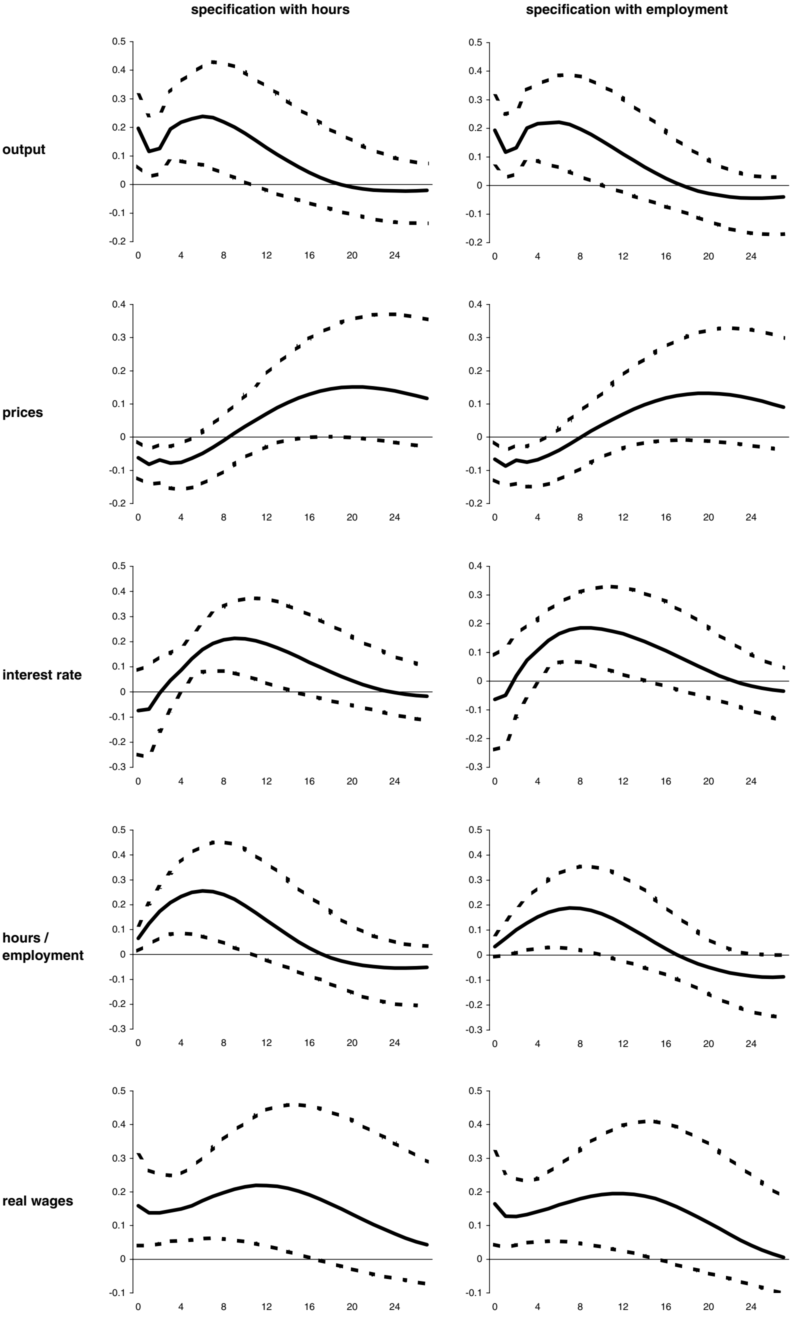
technology shock is a shock with a nonnegative effect on output, prices do not rise, and there is no decrease in real wages. These restrictions are sufficient to uniquely disentangle them from, respectively, monetary policy, aggregate demand, and labor supply shocks, as shown in the previous section. No restrictions are imposed for the response of hours, which allows us to compare the theoretical responses with the data and discriminate between an RBC and an NK model. Following Uhlig (2005) and Peersman (2005), we set the time period over which the sign restriction is binding equal to four quarters ($k = 4$). For the implementation of the restrictions, we refer to Peersman (2005) or the Appendix of this article. All restrictions are imposed as $\leq$ or $\geq$. Impulse responses are computed based on 50,000 draws from the posterior simulator. In all figures, we report the median and the 90th and 10th percentiles of the posterior distribution.

The first column of Figure 3 shows the baseline results. By construction, there is a rise of output and real wages after a positive technology shock and a fall in prices. We also find a positive effect on the nominal interest rate, a response that was unrestricted. Most important, there is a positive and significant reaction on the impact of hours worked. This effect even lasts for more than three years after the initial shock. This striking finding stands in contrast to the results of Galí (1999) and others and is in line with the prediction of the RBC model. In the next sections, we discuss the exact source of our conflicting results with Galí (1999) and the role of the stochastic specification of the hours worked series.

3.2. *Robustness of Results.* As a first robustness check, we re-estimate the basic model, including employment instead of hours worked. The latter sensitivity analysis was also conducted in Galí (1999). The results are reported in the right column of Figure 3. The magnitude of the effect is slightly smaller for employment, but there are no significant differences between the estimated impulse response functions of the employment and the hours worked specification. Also, for other alternative specifications (e.g., consumer price index instead of GDP deflator), we always find a significant rise in hours worked following a positive technology shock. Furthermore, our conclusions do not depend on the inclusion of a time trend in the VAR, the number of lags, and the number of quarters for which the sign restrictions are imposed.[11]

3.3. *A Comparison with the Existing Literature.* In the next step, we aim at identifying whether our results are robust to certain modifications in the empirical model, which has been intensively discussed in the related literature. There are several reasons why our results can differ, for example, from Galí (1999). First, the results might be data-driven; that is, they are dependent on the chosen sample period and data set. Second, the results can be sensitive to the stochastic specification of the hours worked series, as discussed in Christiano et al. (2003). Third, the results might be simply driven by the choice of our new identification scheme.

[11] All these results are available upon request. The results for a VAR with restrictions only imposed for one lag after the shock ($k = 1$) are also shown in the working paper version of this article.



NOTE: Median and 90th and 10th percentiles impulse responses based on the output of the posterior simulator. Technology shock is identified using sign restrictions. Horizon is quarterly.

FIGURE 3

IMPULSE RESPONSES TO A TECHNOLOGY SHOCK

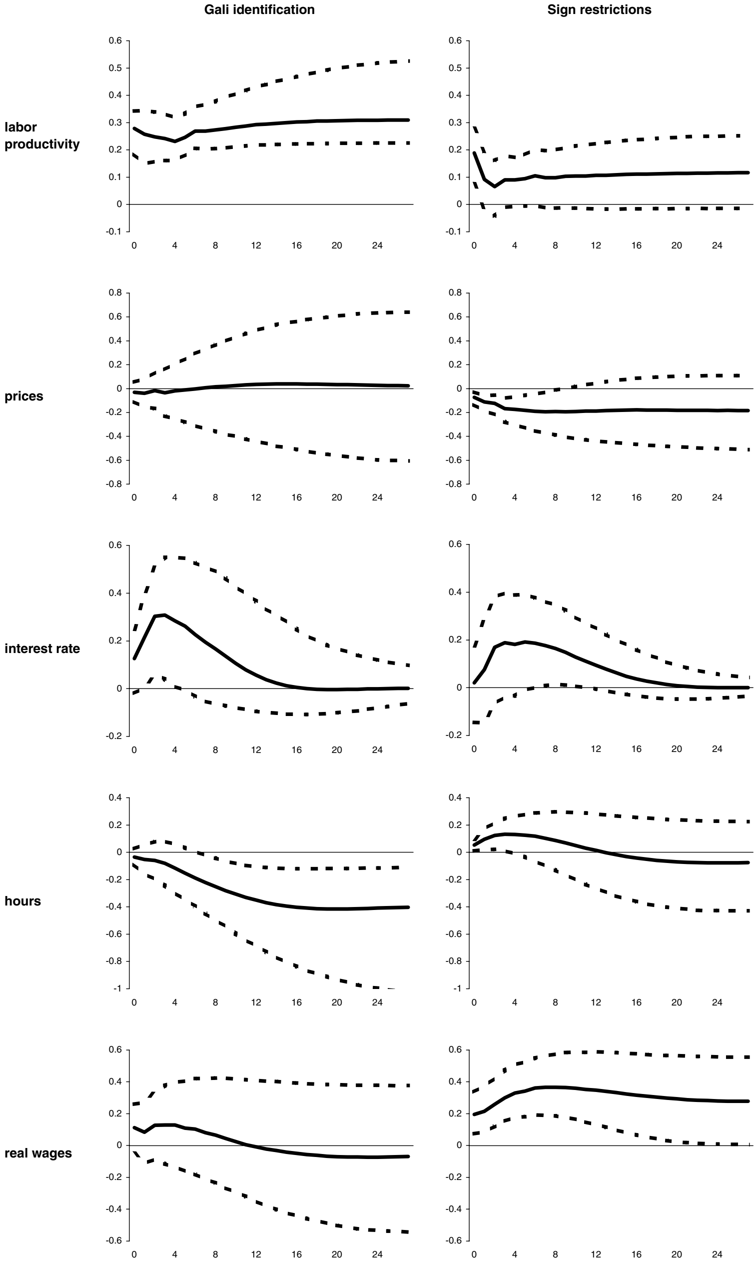
In order to conduct the exercise, we first impose Galí's identification scheme on our euro area VAR. Galí (1999) provides evidence for G7 countries using data starting between 1948 (the United States) and 1970 (Italy) and ending in 1994. The latter results are confirmed in Galí (2004) using data from 1970 to 2002 for the euro area. Given the availability of the hours worked series, our sample starts only in 1982. Moreover, the identification of a technology shock with sign restrictions makes it necessary to add real wages into the VAR. Therefore, for comparison, and following Galí, we estimate a VAR including the first differences of labor productivity $d(y_t - h_t)$, prices (dp_t), real wages (dw_t), hours worked (dh_t), and the level of the interest rate (i_t).[12] A technology shock is identified as the only shock that has a permanent effect on labor productivity. The results are shown in the left column of Figure 4. Consistent with Galí (1999, 2004), we also find a significant fall of hours worked, concluding that the choice of the data sample is not the main driver of our contrasting results.

In order to evaluate the sensitivity of our results to the stochastic specification of the VAR, as discussed in Christiano et al. (2003), we re-estimate the above defined VAR including the first differences of labor productivity $d(y_t - h_t)$, prices (dp_t), real wages (dw_t), hours worked (dh_t), and the level of the interest rate (i_t) but identify the impact of technology shocks on hours worked using our preferred sign restrictions methodology. If, in the given setup, hours worked rise following a technology shock, our contrasting results must be driven by the choice of the identification strategy. On the other hand, if hours worked fall, we could argue that our results are driven by the stochastic specification of the VAR, since the baseline results were produced using a VAR in levels. Estimation results are presented in the right column of Figure 4.[13] In contrast to Christiano et al. (2003), it turns out that the treatment of variables has no consequences for our baseline findings. We still find a significant positive impact of technology shocks on hours worked. The response of all other variables are in line with Galí's results. The impact on labor productivity is somewhat lower and the reaction of real wages is stronger. In sum, we find that the difference in the results is driven by the identification strategy and not by the choice of data set or stochastic specification of hours worked.

3.4. *Exogeneity of the Identified Technology Shocks.* Francis and Ramey (2005) argue that technology shocks identified in VARs are potentially correlated with other shocks that are in fact not related to technology. They therefore present a procedure for testing the exogeneity of the estimated technology shocks. In particular, they regress the identified technology shock on three sets of dummy variables: (i) monetary policy indicators, (ii) oil shock dummies, and (iii) war dates. Given that we are able to disentangle technology from monetary policy shocks with our identification strategy, there is, by construction, no correlation of the latter with technology shocks. Furthermore, our sample period does not include

[12] The results are similar also when the first difference of the interest rate is used instead of its level in this specification.

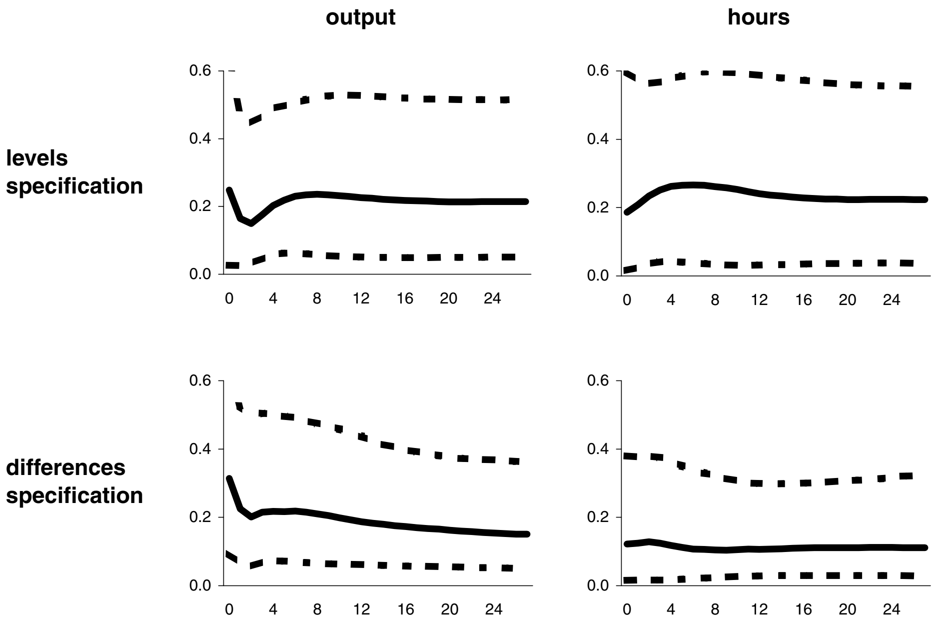
[13] An alternative exercise could be to estimate a Vector Error Correction Model (VECM) specification of the DSGE model and implement the long-run restrictions, in the spirit of Chang and Schorfheide (2003), but this is out of the scope of this article.



NOTE: Median and 90th and 10th percentiles impulse responses based on the output of the posterior stimulator. Variables included are first differences of labor productivity, prices, hours, real wages, and the level of the interest rate. Technology shock in the left panel is identified with Gali's long-run restrictions and with sign restrictions in the right panel. Horizon is quarterly.

FIGURE 4

IMPULSE RESPONSES TO A TECHNOLOGY SHOCK



NOTE: Median and 90th and 10th percentiles values based on the output of the posterior simulator. Horizon is quarterly.

FIGURE 5

CONTRIBUTION OF TECHNOLOGY SHOCKS TO FORECAST VARIANCE

important war dates for the euro area. In order to check the potential correlation with oil price shocks, we perform two robustness checks. First, we calculate a simple correlation between the identified technology shocks and the pure oil price shocks obtained from the study of Peersman (2005). This correlation varies between −0.17 and −0.20, depending on the specification, and is always insignificant. Second, we re-estimate all VAR models, with oil prices (or commodity prices) as an additional exogenous variable. For all specifications, we still find a significant positive impact of technology shocks on hours worked.[14]

3.5. *How Important Are Technology Shocks for Aggregate Fluctuations?* In Figure 5, we report the contribution of technology shocks to the forecast error variance of output and hours worked series for the level and first-difference specification, respectively. In contrast to Galí (1999), who finds almost no role for technology shocks in explaining business cycle fluctuations, we find an important role for technology shocks in explaining variations in output and hours worked. The error bands are, however, very wide, which is typical for this type of exercise. On the basis of the median estimates, we find that between 20% and 25% of variations in hours worked and output can be explained by technology shock. The

[14] These results are not reported but are available upon request.

latter is, however, still significantly lower than the 40% reported by Christiano et al. (2003). Hence, our results stand somewhat in contrast to the original RBC hypothesis that technology shocks are the main drivers of business cycle fluctuations, as they assign an important role for other structural shocks for business cycle fluctuations.

4. CONCLUSIONS

In this article, we have provided empirical evidence on the impact of technology shocks on hours worked using a VAR for the euro area. The structural shocks are identified using sign restrictions derived from the DSGE models. The suggested procedure utilizes, however, only a minimum set of restrictions that are robust to model and parameter uncertainty. The results presented in this article are in favor of the RBC hypothesis that hours worked increase following a positive shock to technology. Although our results assign a more prominent role for technology shocks in explaining variations in output and hours worked than Galí (1999), they do stand, however, somewhat in contrast to the original RBC hypothesis that technology shocks are the main source of business cycle fluctuations.

Furthermore, we show that our results are not sensitive (i) to different stochastic specifications or (ii) if employment instead of hours worked is used in the VAR. In addition, we argue by comparing our results to a VAR with long-run restrictions, as in Galí (1999), that our results depend only on the chosen identification scheme and not on the stochastic specification of hours worked and/or the data sample.

However, our findings do not necessarily imply that the NK models are generally not a good representation of the reality. But, the results indicate that the NK models stand in contrast to the empirical evidence at least in one particular aspect, namely, the transmission of technology shocks to the labor market. Hence, reconsidering the transmission mechanism of technology shocks within the NK framework might be a worthwhile exercise.[15] Also, the structural shocks in our empirical analysis are identified at a fairly aggregated level. Identifying additional shocks, such as price and wage mark-up shocks, could potentially provide further information. This is left for future research.

APPENDIX

A.1. *Implementation of the sign restrictions.* In this Appendix, we explain how to implement the sign restrictions in our SVAR. For a detailed explanation, we refer to Peersman (2005). Consider Equation (13) in Section 3. Since the shocks are mutually orthogonal, $E(\varepsilon_t \varepsilon_t') = I$, the variance–covariance matrix of Equation (13) is equal to $\Omega = BB'$. For any possible orthogonal decomposition B, we can find an infinite number of admissible decompositions of Ω, $\Omega = BQQ'B'$, where Q is any orthonormal matrix, that is, $QQ' = I$. Possible candidates for B are the Choleski

[15] See also, on a similar issue, the discussion about the effects of government spending shocks on private consumption in Galí et al. (2007) and Bilbiie and Straub (2004). As discussed in Chang et al. (2004), introducing inventories into the standard NK model might be a promising first step.

factor of Ω or the eigenvalue-eigenvector decomposition, $\Omega = PDP' = BB'$, where P is a matrix of eigenvectors, D is a diagonal matrix with eigenvalues on the main diagonal, and $B = PD^{\frac{1}{2}}$. Following Canova and De Nicoló (2002) and Peersman (2005), we start from the latter in our analysis. More specifically, $P = \prod_{m,n} Q_{m,n}(\theta)$, with $Q_{m,n}(\theta)$ being rotation matrices of the form

$$\text{(A.1)} \qquad Q_{m,n}(\theta) = \begin{bmatrix} 1 & \cdots & 0 & \cdots & 0 & \cdots & 0 \\ \cdots & \ddots & \cdots & \cdots & \cdots & \cdots & \cdots \\ 0 & \cdots & \cos(\theta) & \cdots & -\sin(\theta) & \cdots & 0 \\ \vdots & \vdots & \vdots & 1 & \vdots & \vdots & \vdots \\ 0 & \cdots & \sin(\theta) & \cdots & \cos(\theta) & \cdots & 0 \\ \cdots & \cdots & \cdots & \cdots & \cdots & \ddots & \cdots \\ 0 & \cdots & 0 & \cdots & 0 & \cdots & 1 \end{bmatrix}.$$

Since we have five variables in our model, there are 10 bivariate rotations of different elements of the VAR, $\theta = \theta_1, \ldots, \theta_{10}$, and rows m and n are rotated by the angle θ_i in Equation (A.1). All possible rotations can be produced by varying the 10 parameters θ_i in the range $[0, \pi\ ]$. For the contemporaneous impact matrix determined by each point in the grid, B_j, we generate the corresponding impulse responses:

$$\text{(A.2)} \qquad R_{j,t+k} = A(L)^{-1} B_j \varepsilon_t.$$

A sign restriction on the impulse response of variable p at lag k to a shock in q at time t is of the form

$$\text{(A.3)} \qquad R^{pq}_{j,t+k} \gtrless 0.$$

Following Uhlig (2005), Peersman (2005), and Farrant and Peersman (2006), we use a Bayesian approach for estimation and inference. Our prior and posterior belong to the normal-Wishart family used in the RATS manual for drawing error bands. Because there are an infinite number of admissible decompositions for each draw from the posterior when using sign restrictions, we use the following procedure. In order to draw the "candidate truths" from the posterior, we take a joint drawing from the posterior for the usual unrestricted normal-Wishart posterior for the VAR parameters as well as a uniform distribution for the rotation matrices. We then construct impulse response functions. If all the imposed conditions of the impulse responses are satisfied, we keep the draw. Decompositions that do not match the restrictions are rejected. This means that these drawings receive zero prior weight. On the basis of the drawings kept, we calculate statistics and report the median of the posterior distribution, together with the 90th and 10th percentiles.

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