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Francesco Giuli e Massimiliano Tancioni

Contractionary Effects of Supply Shocks:
Evidence and Theoretical Interpretation

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Francesco Giuli and Massimiliano Tancioni*

Sapienza University of Rome

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Abstract

The debate on the response of hours worked after productivity improvements is still an open issue in the theoretical and empirical literature. In this work we show that, once conditional correlations are taken into account, both hours and investment decline temporarily following a positive technology shock. We first provide evidence about this apparent puzzle employing weakly identified SVECs. We then set-up and estimate a sticky price/wage DSGE model in which the presence of strategic complementarities in price setting lowers the slope of the New Keynesian Phillips curve, and show that the posterior impulse responses are consistent with the SVEC-based evidence.

JEL CLASSIFICATION: E32, E22, C11.

KEYWORDS: technology shocks, investment dynamics, vector error correction model, Bayesian inference.

*Corresponding author: massimiliano.tancioni@uniroma1.it

1 Introduction

The observed strong and positive covariation between GDP, investment and hours worked has been, for a long time, one of the most uncontroversial stylized facts of macroeconomic dynamics. This empirical regularity is consistent with the RBC view on business cycle determination and, in principle, with any theoretical setting in which the level of factor inputs is adjusted proportionately to their expected marginal revenue. Price flexibility in RBC models and a rationale use of the demand channel by the monetary authority in New-Keynesian models are the key elements which, following a productivity improvement, prevents the aggregate demand response to constrain the increase in factor inputs.

In this paper we question the standard view on the determinants of short/medium-term investment and hours worked dynamics from both an empirical and a theoretical perspective. Our line of argument is motivated by the finding that, even if the *unconditional* correlation between productivity and inputs is positive, their *conditional* correlation on identified technology shocks is negative.

The idea that positive technology shocks could have some contractionary effects dates back to Hansen and Wright (1992) and has been recently supported by abundant empirical emphasis, mostly focusing on the negative employment effects of productivity improvements, the so-called "productivity-employment puzzle"¹.

Differently from the employment puzzle, the evidence of a negative short-term investment response is not yet established in the empirical literature on business cycles. One major exception is the analysis in Basu *et al.* (2006) who show that, once a "purified" measure of TFP is considered, technology improvements turn out contractionary on impact for both factor inputs. However, the literature is far from reaching a consensus and the short-run effect of technology shocks is still an open issue in the debate on macroeconomic dynamics.

A crucial and controversial point is the identification of technology shocks. Two approaches can be singled out. The first relies on the long-run identification of vector autoregressions (SVARs), based on the assumption that only technology shocks can have permanent effects on labor productivity (Galí, 1999). The second, proposed by Basu *et al.* (2006), uses a direct

¹See, among others, Galí (1999), Francis and Ramey (2002, 2005), Basu *et al.* (2006).

measure of technology obtained with a bottom-up growth accounting approach.

The empirical literature based on SVARs, identified with long-run restrictions, reaches conflicting conclusions about whether hours worked rise or fall after a productivity improvement. Cooley and Dwyer (1998) and Chari *et al.* (2005) show that SVAR-based results may be seriously affected by the use of long-run restrictions in small samples, while the use of low-dimensional VARs may lead to truncation biases (Fry and Pagan, 2005). Moreover, Christiano *et al.* (2003, 2004) criticize the results obtained by Galí (1999) and by the following literature on the employment puzzle sustaining that it emerges because of the wrong consideration of nonstationary hours in SVARs.

The first part of our analysis aims at establishing an explicit empirical premise that corroborates the results obtained by Basu *et al.* (2006) from a more standard methodological perspective, relying on unprocessed data for productivity. To do this, we use a Structural Vector Error Correction (SVEC) model, specified such that it can render an approximation of a fairly general NK-DSGE model with capital accumulation subject to permanent supply shocks. The main outcome from the SVEC analysis is that the short-term response of both investment and hours to a positive technology shock is negative. This result confirms that obtained by Basu *et al.* (2006) and provides further evidence questioning the very relevance of supply shocks in explaining the observed pro-cyclicality of productivity, investment and hours.

A major advantage of our approach is that results are potentially more challenging with respect to those provided by the literature on the negative short-term response of hours, since: *i*) we also focus on the dynamics of the most pro-cyclical macro-variable, investment; *ii*) from robustness checks, we show that the critical arguments advanced by the opponents to the "productivity-employment puzzle", do not apply to our analysis. We in fact show that neither the introduction of a measure of capital in the SVEC to control for the truncation bias, nor the consideration of stationary hours to take into account Christiano *et al.*'s (2004) critical arguments, alter the main results of our analysis.

The second part of the analysis focuses on providing a theoretical explanation for the SVEC evidence. We set-up and estimate a monetary DSGE model in which the slope coefficient of the new Keynesian Phillips curve (NKPC) depends, aside from the frequency at which firms

are allowed to reset their prices, on the presence of strategic complementarities in price setting. These arise as a result of two important modifications to the baseline NK-DSGE model: *i*) the hypothesis that capital is firm-specific (Altig *et al.*, 2005; Woodford, 2005; Sveen and Weinke, 2005-2007), and *ii*) the hypothesis that the elasticity of demand is endogenous (Eichenbaum and Fisher, 2007; Smets and Wouters, 2007). We show that, when considering these model extensions, the negative investment and hours response to supply shocks emerges even with reasonable degrees of nominal price/wage rigidity and of monetary policy accommodation.

The strategy adopted for the parameterization of theoretical models is a key issue, as it mirrors the conflicting SVAR-based evidence. Some of the most recent and influential contributions supporting the procyclicality of investment and hours worked (Christiano *et al.*, 2005; Altig *et al.*, 2005) are in fact parameterized by using estimators that minimize a weighted distance between the theoretical and the SVAR-based impulse responses. In our opinion, the evidence produced with a matching estimator applied to a model with particularly flexible dynamic properties cannot be considered conclusive, since it does not provide additional evidence to that implied by the SVAR-based impulse responses. This strategy basically forces the model calibration to a parameters space for which the implied model dynamics is consistent with that obtained with the SVAR.

These considerations lead us to use a more transparent empirical strategy for the calibration of our model. We perform a fully structural Bayesian estimate, where we counterfactually adopt a prior parameterization for which our extended model does not replicate the SVEC-based evidence on the negative conditional correlation between productivity and factor inputs. By employing the same data set, we show that both the weakly identified SVEC results and the structural Bayesian estimate of our model predict a very similar behavior of the relevant macroeconomic variables.

The economic rationale for our results is related to the flatter New Keynesian Phillips curve (NKPC) characterizing the firm-specific capital and/or endogenous demand elasticity specifications: when marginal costs are firm-specific, the incentive to cut prices following a positive productivity shock is countervailed by the increase in marginal costs due to an increased demand at the firm level. Such an effect is even stronger when considering the endogenous demand elasticity in addition to the firm-specific capital specification. This, other

things equal, leads to a weak demand response that constrains the use of inputs in production.

The fact that a firm specific capital monetary model can address the negative short-run response of investment to supply shocks has been acknowledged in Altig *et al.* (2005). However, from the empirical evaluation of their model, which employs an impulse response matching strategy, they rule out such a result, since from the simulation of their benchmark near-SVAR they obtain a positive short term response of factor inputs following an unexpected productivity improvement. This is the very reason why Altig *et al.*'s (2005) model fails to replicate the observed decrease in inflation after a technology shock. In order to capture the increase in the general economic activity, their estimation criterion forces the monetary policy response to be accommodative enough so as to prevent a decrease in inflation.

The relevance of the aggregate demand response highlights the role played by monetary authorities. It is well known that, by adopting a Taylor rule that responds to the theory-consistent output gap, the monetary authority can mimic the optimal monetary policy that minimizes the volatility of the target variables around their natural levels. However, the implementation of such a rule in real life operations requires the knowledge of the natural rate of interest or of the level of potential output, that are hardly in the monetary authority information set. An abundant literature has shown that the conduct of monetary policy is not fully accommodative and can be approximated by an empirical Taylor rule targeting trend output instead of the theory-consistent potential output.

In order to rule out that our estimation results may depend on the preferred policy rule, we assume a monetary rule that adjusts the interest rate with respect to a weighted average considering trend output deviations and the theory-based output gap. Our estimates tend to favor the weight assigned to the former, rendering the policy response not fully accommodative and thus leading to a weak aggregate demand response to the supply shock.

The paper proceeds as follows. Section 2 provides the results of the SVEC analysis, showing that the short-term hours and investment responses to a positive supply shock are negative. Based on this evidence, section 3 introduces an extended sticky price/wage model with firm-specific capital and endogenous demand elasticity. Section 4 provides the details of the Bayesian MCMC estimates of the model and discusses the results in the light of the SVEC outcomes. Section 5 concludes.

2 The empirical premise

In this section we provide empirical evidence challenging the standard theoretical prediction that the short-term hours and investment responses to technology shocks are non-negative. To do this, differently from Basu *et al.* (2006), who use growth accounting measures of technology shocks in a SVAR-based experiment in the spirit of Blanchard and Quah (1989) and Galì (1999), we employ unprocessed data for productivity and adopt a SVEC representation. The use of the SVEC is justified by the fact that, when in the presence of nonstationary variables, the stationary SVAR representation is potentially mis-specified, due to the fact that co-integration (CI) may emerge among the variables in the VAR.

2.1 The vector error correction model

The VEC model is estimated considering hourly productivity (y_t), labor supply (h_t), inflation (π_t), the short-term nominal interest rate (r_t), consumption per hour (c_t) and non-residential investment per hour (i_t). The vector of endogenous variables is thus $\mathbf{x}'_t = [y_t \ h_t \ \pi_t \ r_t \ c_t \ i_t]$. A convenient structural formulation of the m -dimensional VEC assumes no contemporaneous correlations among variables and is given by:

$$\mathbf{\Gamma}(L) \Delta \mathbf{x}_t = \mathbf{\Pi} \mathbf{x}_{t-1} + \mathbf{B} \boldsymbol{\varepsilon}_t \quad (1)$$

where $\mathbf{\Gamma}(L) = \mathbf{\Gamma}_0 - \mathbf{\Gamma}_1 L - \dots - \mathbf{\Gamma}_{p-1} L^{p-1}$ are structural coefficients matrices and $\mathbf{\Gamma}_0 = \mathbf{I}_m$. Under this hypothesis, $\mathbf{B}$ contains the contemporaneous structure of the system, i.e. the contemporaneous correlations among variables and among errors. The long-run relation matrix $\mathbf{\Pi} = \boldsymbol{\alpha} \boldsymbol{\beta}'$, in the presence of cointegration, is a reduced rank matrix and $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $m \times r$ full-column rank matrices containing the loading coefficients and the long-run equilibrium relations, i.e. the r cointegrating vectors, respectively. The vector of disturbances $\boldsymbol{\varepsilon}_t \sim (\mathbf{0}, \mathbf{I}_m)$ contains the orthonormal structural innovations. The system of linear equations relating the estimated reduced-form errors $\mathbf{u}_t$ to the structural shocks is thus $\mathbf{u}_t = \mathbf{B} \boldsymbol{\varepsilon}_t$, which implies $\boldsymbol{\Omega} = \mathbf{u} \mathbf{u}' = \mathbf{B} \mathbf{B}'$.

VARs unavoidably consider only a subset of the key variables that describe the dynamics of an economic system. When omitted variables are Granger-causal, or related to the maintained set by dynamic identities in which autoregressive terms are present, infinite-order VARMA

representations may emerge, leading to relevant biases in the finite-order VAR estimates (Fry and Pagan, 2005)². In order to control for such approximation problems, we choose a lag-order of the VEC which is higher than that suggested by the standard information criteria, and check the robustness of results with respect to an alternative specification including the fixed assets quarterly series (k_t) ³.

2.1.1 Data

We employ quarterly time-series for the sample period 1954:3 - 2007:2. The labor supply is calculated as the log ratio between non farm total hours worked and population aged over 16⁴. The capital stock time-series is obtained by applying the permanent inventory method on US quarterly data on gross fixed capital formation and depreciation and annual data for the capital stock. Data on real variables and the GDP price deflator are obtained from the BEA database, while data employed to construct the labor supply measure are obtained from the BLS database. The nominal rate is the Federal Funds Rate and is obtained from the FRED database. All variables enter the VAR in logs. A more detailed description of data and data manipulations is provided in the Appendix.

2.1.2 Integration and cointegration tests

From DF-ADF tests we obtain that real and monetary variables are all $I(1)$ in levels, irrespective of how the deterministic components are specified⁵. The tests are instead inconclusive with respect to the labor supply measure. The hours to population ratio is $I(0)$ at the standard 5% significance level when a constant term is included, and $I(1)$ under the alternative

²For this reason, the VAR representation of models with capital accumulation cannot be compressed by omitting the capital stock, since capital is dynamically related to investment via the permanent inventory identity. The relevance of the truncation bias and of more general approximation errors in VAR analyses has been largely addressed by the literature (Lippi and Reichlin, 1994; Cooley and Dwyer, 1998; Chari *et al.*, 2005). However, in most cases these problems can be controlled by augmenting the lag-order of the VAR (Erceg *et al.*, 2005; Ravenna, 2007) or avoided by appropriately selecting the variables to be included.

³In this case the vector of endogenous variables of the level VAR representation is $\mathbf{x}'_t = [y_t \quad h_t \quad \pi_t \quad r_t \quad c_t \quad i_t \quad k_t]$.

⁴The use of hours to population ratios as a labor supply measure is standard in the literature (see Christiano *et al.*, 2003; Gali and Rabanal, 2004; Chari *et al.*, 2005; Del Negro *et al.*, 2005). We adopt a correction to the denominator in order to take into account the calendar effects on total available time.

⁵The appropriate lag order of the DF-ADF tests is identified employing a two-step procedure. The order is at first specified according to the Schwartz Bayesian Criterion (SBC); then, on the basis of the results of the Portmanteau statistic, it is augmented such that residuals are not autocorrelated.

specifications⁶. Considering the weak power of standard unit root tests (Maddala and Kim, 1998) and the very high degree of persistence resulting from the estimated autoregressive representation of hours, in the benchmark specification we assume that the labor input measure is nonstationary. However, since the $I(1)$ assumption for hours may result crucial for the outcomes of the analyses (Christiano *et al.*, 2004), we also re-estimate the six variables VEC by imposing stationary hours. This alternative specification provides a second robustness check of results.

Johansen's trace test indicates the presence of three cointegrating vectors at the 95% significance level. When the capital stock series is included in the VEC a fourth cointegrating vector emerges at the same significance level⁷. According to the estimated CI ranks, there are thus three permanent components.

2.1.3 Long-run components and CI space

We assume that the permanent component observed in the three real variables y_t , c_t , i_t is due to the technological stochastic trend (King *et al.*, 1991), while the permanent component observed in the monetary variables r_t and π_t is due to how the central bank adjusts its policy target (Vlaar, 2004)⁸. The third permanent component is attributed to labor supply shocks. Given variables' ordering, the first CI relation in β' defines the Fisher interest parity $r_t - \beta_{13}\pi_t$, i.e. the stationary real interest rate, while the two remaining CI relations in β' define the stationary "great ratios" of the economy, i.e. $c_t - \beta_{21}y_t$ and $i_t - \beta_{31}y_t$. These stationary ratios are consistent with the balanced growth hypothesis (King *et al.*, 1991; Pesaran and Smith, 1995, 1998; Garratt *et al.*, 2003). The additional CI relation emerging when capital is included in the first alternative VEC specification is attributed to the theoretical steady-state relation

⁶Detailed results are reported in the technical appendix of the paper, available upon request from the authors.

⁷Johansen's trace tests are referred to a VEC structure considering three lags, (thus four lags in the levels VAR), unrestricted constants and two dummies, the first controlling for the post-Volker period, the second capturing the increase in the consumption to output ratio since the early 90's. The chosen lag-order is higher than that indicated by the Schwartz and Akaike information criteria (suggesting one and three lags, respectively). Standard autocorrelation tests indicate that such a choice ensures non autocorrelated errors. The peculiar specification of the constant term in the VEC, which is consistent with the presence of trending variables, has been chosen according to the results of standard LM exclusion tests.

⁸This hypothesis is justified by the fact that, in the absence of inflation biases, the long-run output gap is zero, while long-run inflation is determined by the monetary authority policy target. This view is consistent with the standard results of the NK-DSGE literature.

between capital and investment $k_t - \beta_{46}i_t$. The case of stationary hours in the six variables VEC (the second alternative VEC) is evaluated by imposing a further CI relation in which only labor h_t enters the corresponding cointegrating vector⁹.

2.2 Identification of the SVMA representation

The forecast error impulse response functions (IRFs) are obtained from the structural vector moving average (SVMA) representation, which for the SVEC model takes the form:

$$\mathbf{x}_t = \mathbf{C}(1) \sum_{i=1}^t \mathbf{B}\boldsymbol{\varepsilon}_i + \mathbf{C}^0(L) \mathbf{B}\boldsymbol{\varepsilon}_t + \tilde{\mathbf{x}}_0 \quad (2)$$

where $\mathbf{C}(1) = \boldsymbol{\beta}\perp(\boldsymbol{\alpha}'\perp\boldsymbol{\Gamma}\boldsymbol{\beta}\perp)^{-1}\boldsymbol{\alpha}'\perp$ is the long-run effects matrix, in which $\boldsymbol{\Gamma} = \mathbf{I}_m - \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i$ and $\boldsymbol{\alpha}\perp, \boldsymbol{\beta}\perp$ are the orthogonal complements of the loading coefficients and the long-run equilibrium matrices, respectively; $\mathbf{C}^0(L) = \sum_{j=0}^{\infty} \mathbf{C}_j^0 L^j$ is a convergent infinite order polynomial in the lag operator and $\tilde{\mathbf{x}}_0 = [\boldsymbol{\beta}\perp(\boldsymbol{\alpha}'\perp\boldsymbol{\Gamma}\boldsymbol{\beta}\perp)^{-1}\boldsymbol{\alpha}'\perp] \mathbf{x}_0$ depends on initial conditions (Johansen, 1995).

Leaving out the coefficients attached to the lagged variables, the reduced-form VEC provides $m(m+1)/2 = 21$ nonredundant coefficients in the dispersion matrix $\boldsymbol{\Omega}$, while the SVEC has $m^2 = 36$ unknown structural coefficients in $\mathbf{B}$. The identification of the structural errors thus requires imposing $m(m-1)/2 = 15$ restrictions.

The rank of the $m \times m$ total impact matrix $\mathbf{C}(1)$ is given by the number of permanent components in the system, which in the benchmark specification is $m-r = 3$. This implies that the last three columns of $\mathbf{C}(1)$, which correspond to the r transitory components of the model (the CI vectors), are zero. However, given the reduced rank of $\mathbf{C}(1)$, CI provides only $(m-r)r = 9$ constraints for the long run response matrix $\mathbf{C}(1)$, leaving $(m-r)(m-r-1)/2 = 3$ additional restrictions to exactly identify the permanent shocks and $r(r-1)/2 = 3$ restrictions to exactly identify the transitory shocks (Breitung *et al.*, 2004). The structural identification of the two alternative SVMA representations (one considering the capital stock, the other stationary hours) requires the imposition of three additional short-run constraints, since $r = 4$.

⁹The estimation of the theoretical long-run structure implies the imposition of 15 restrictions (24 in the VEC including the capital stock and 21 in the VEC with stationary hours), while exact identification requires nine restrictions (16 in the two specifications with four transitory components). This implies that the CI space is over-identified. The over-identifying restrictions are always rejected by the data when asymptotic chi-sq. critical values are employed.

The identification of the permanent shocks in the SVMA representation considering stationary hours requires only one restriction, since in this case $m - r = 2$.

2.2.1 Permanent components

The permanent components are identified by adopting the standard hypothesis that only supply shocks have permanent effects on productivity (Shapiro and Watson, 1989; Blanchard and Quah, 1989; Galí, 1999; Francis and Ramey, 2002, 2005) and by imposing monetary (inflation) long-run neutrality for hours. The first hypothesis provides two exclusion restrictions on the long-run effects of labor supply and inflation shocks on productivity (the elements c_{12} and c_{13} of the $\mathbf{C}(1)\mathbf{B}$ matrix are zero), while the latter imposes long-run inflation neutrality for hours ($c_{23} = 0$). The hypothesis that productivity and labor supply shocks have zero long-run effects on inflation provides two over-identifying constraints, that we impose to enhance the structural identification of the model ($c_{31} = c_{32} = 0$). These hypotheses are consistent with the predictions of a broad class of theoretical models¹⁰. The same long-run identification scheme is adopted for the two alternative SVEC specifications, with the peculiarity that in the stationary hours version of the VEC the elements of the second column of the $\mathbf{C}(1)\mathbf{B}$ matrix are zero by construction.

2.2.2 Transitory components

Contrary to the long-run, theory is less informative for the definition of the short-run constraints. Forward-looking behavior in DSGE models implies that the structural shocks affect all variables simultaneously, thus the use of exclusion restrictions on the impact responses is not legitimate¹¹. For this reason, instead of assuming pre-set contracts, nominal rigidities or informational lags to justify a sparse $\mathbf{B}$ -matrix, we employ cross-coefficient restrictions that are directly derived from the calibrated policy functions of a standard monetary DSGE

¹⁰The theoretical literature suggests a further exclusion restriction for the long-run effects of productivity shocks on hours (Francis and Ramey, 2002). We do not take into account such restriction in our identification scheme on the grounds that, under permanent supply shocks, its validity is limited to a particularly restrictive specification of utility, i.e. one in which income and substitution effects exactly cancel-out in the long-run.

¹¹The recourse to nominal rigidities and to decision or informational lags to justify the contemporaneous exclusion restrictions has been widely used in the literature (Blanchard, 1989; Galí, 1992; Christiano *et al.*, 2005). Vlaar (2004) criticizes this practice by noting that, if decision/information lags are at work, they affect the contemporaneous response to all shocks, thus hampering their structural identification.

model¹². In order to exactly identify the transitory components, we explicit the **B**-matrix cross-elements restrictions implied by the following theoretical relations:

i) the log-linear aggregate resources constraint $y_t = (1 - \psi) c_t + \psi i_t$, where ψ represents the steady-state share of investment in output, which depends on deep parameters, and that we fix to 0.25. This relation provides a proportionality restriction among consumption, investment and output;

ii) the log-linear Cobb-Douglas production function $y_t = \alpha k_{t-1} + (1 - \alpha) h_t$, where $1 - \alpha$ represents the labor share in production, fixed at 0.36. This relation provides a theory-based proportionality condition between the impact responses of log-output and hours (provided that the amount of installed/operated capital is predetermined);

iii) the monetary policy Taylor rule $r_t = \rho + \phi_\pi \pi_t + \phi_y y_t$, $\phi_\pi + \phi_y > 1$, which provides an exact relation among the impact responses of the nominal interest rate, log-output and inflation. According to standard calibration values, we fix $\phi_\pi = 1.5$ and $\phi_y = 0.125$.

Three restrictions are needed to exactly identify the transitory components. Considering the transitory component affecting consumption, two cross-coefficients restrictions for the contemporaneous effects **B**-matrix are derived from the theoretical relations above: the first imposes the proportionality condition among output, consumption and investment (implied by the aggregate resource constraint) to the impact effects of a transitory shock to consumption (preferences), i.e. $b_{15} - (1 - \psi) b_{55} - \psi b_{65} = 0$; the second restriction imposes the proportionality condition between output and labor implied by the production function, i.e. $b_{15} - (1 - \alpha) b_{25} = 0$. Considering the transitory component affecting investment, we impose the cross-coefficient proportionality condition between the interest rate, inflation and output implied by the policy rule, i.e. $b_{46} - \phi_\pi b_{36} - \phi_y b_{16}$.

The identification of the SVMAs of the two alternative SVEC specifications requires the imposition of three additional restrictions. For both the stationary hours specification and the one including the capital stock, these are provided by the following cross-coefficients

¹²Our identification strategy is close in spirit to that adopted by Vlaar (2004), who uses a proportionality condition between the impact response of inflation and real money balances. The main difference is that we extend the cross-coefficients restriction approach by focusing on the theoretical relations that hold exactly in each period. These relations are those in which only state variables are present and the shock being considered is not present. The equations including jump variables do not provide exact relations, given the presence of expectational terms.

proportionality conditions: $b_{45} - \phi_{\pi} b_{35} - \phi_y b_{15}$, $b_{16} - (1 - \psi) b_{56} - \psi b_{66} = 0$ and $b_{14} - (1 - \alpha) b_{24} = 0$. Compared to the exclusion restrictions strategy, our identification scheme has the advantage of satisfying the order and rank conditions for *empirical* identification without violating the theoretical implications for the structure of the contemporaneous response matrix¹³.

To check the robustness of results, we also implement a more standard identification scheme in which we impose the orthogonality among permanent and transitory components. Results are basically unchanged under this alternative identification scheme.

2.3 SVEC impulse response analysis and its theoretical implications

Figure 1 shows the impulse responses to a one standard deviation productivity shock of the variables entering the SVEC model. Solid lines denote the impulse responses point estimates, while dotted lines define 95% asymptotic confidence intervals.

Three features of Figure 1 are worth noting. First, real variables (output, wage, consumption and investment) denote a positive and statistically significant long-run impulse response, confirming the standard theoretical prediction that supply shocks are expansionary in the long-run. This feature is consistent with the nonstationary properties of the data and with the long-run identification strategy imposing a technology-driven common stochastic trend among real variables. Secondly, the impulse responses of monetary variables are consistent with the theoretical predictions of standard monetary models. The inflation and interest rate responses are significantly negative for nearly two and four quarters, respectively. Interestingly, the response of the nominal rate denotes gradual accommodation, confirming the well-documented inertial behavior of monetary authorities. Thirdly, and most importantly, while the short-term responses of output and consumption are all consistent with the predictions of standard business cycle models, those of hours and investment are not. Hours decline immediately following a supply shock, peaking at nearly three quarters. The 95% error bands indicate that the negative response of hours is significant for approximately 7 quarters. Similarly, investment denotes a hump-shaped short-term negative response; the impulse response point estimate crosses the zero line after 10 quarters and then converges to its positive long-run

¹³This strategy does not guarantee the *theoretical* identification of the transitory shocks, since a given cross-equation restriction is employed for different shocks, and the theoretical constraints depend on the chosen calibration of the underlying models. This is not problematic for our analysis, since we are only interested in the structural identification of productivity shocks.

response value. According to the 95% confidence intervals, the negative response is significant for approximately five quarters.

Figure 1 - Impulse responses to a supply unit shock and 95% confidence interval region
Benchmark specification

The finding of a negative short-term response of investment is particularly relevant since it provides further evidence questioning the very relevance of supply shocks in explaining the observed pro-cyclicality of productivity, investment and hours. Moreover, compared to the well documented negative short-term response of hours, the conditional investment dynamics entails a more robust confirmation of the evidence on the contractionary effects of supply shocks. In fact, differently from the productivity-employment puzzle, the short-term contractionary effects on investment cannot be criticized on the grounds of the statistical properties of the investment time series: on the empirical side, the nonstationarity hypothesis is largely accepted for the investment series; on the theoretical side, since technology is the unique source of nonstationarity, investment must be nonstationary to guarantee the validation of the long-run balanced growth hypothesis.

The two alternative specifications of the SVEC confirm these results. The short-term response of hours and investment to an identified supply shock are in fact negative and significant in both the seven variables SVEC (the one including the fixed assets series) and in the six variables SVEC with stationary hours (see Figures A1-A2 in the Appendix).

3 The model

In this section we provide the linearized version of a cash-in-advance sticky price/wage NK-DSGE model characterized by different sources of strategic complementarities in price setting. Our theoretical apparatus is developed along the lines of Altig *et al.* (2005), Sveen and Weinke (2005), and Woodford (2005) for the modelization of the firm-specific capital accumulation, Erceg *et al.* (2000) for the introduction of nominal wage staggering and Smets and Wouters (2007) for the endogenous demand elasticity specification.

The model economy is populated by maximizing households and firms, while monetary and fiscal authorities follow exogenous policy rules. Final sector firms operate in a perfectly competitive environment as simple aggregators of the differentiated goods produced by intermediate sector firms. These combine labor and capital services employing a Cobb-Douglas production technology, which is subject to permanent productivity shocks that give rise to a common stochastic trend and to long-run stationary ratios among real variables. This feature makes the model consistent with the nonstationary and co-trending behavior of the data addressed by the SVEC analysis.

Each intermediate firm rents differentiated labor services from the households and makes an investment decision to adjust its capital stock to the desired level considering a capital adjustment cost. The firm-specific capital hypothesis implies that a predetermined stock of capital is employed in production. Intermediate sector firms can change price only infrequently according to a random duration Calvo-lottery and those that are not allowed to change their price adopt a partial indexation scheme.

Households maximize a separable utility function defined over consumption and leisure. Each household purchases consumption goods by means of labor income and cash holdings net of deposits. Their preferences exhibit external habit persistence in consumption and are assumed to be log-linear in consumption and CRRA in leisure in order to guarantee balanced growth. The presence of differentiated labor services implies some monopoly power in labor supply, and wages are set in staggered contracts according to a Calvo-scheme. In analogy with the pricing behavior of firms, wages that are not reoptimized are partially indexed to past inflation.

With respect to other models explicitly addressing the analysis of the business cycle in a structural perspective (Smets and Wouters, 2003, 2007), the peculiar theoretical features of our DSGE model are basically two: *i*) capital accumulation is decided by firms; *iii*) since productivity shocks are permanent, the model economy evolves stochastically around a deterministic growth path γ_z ¹⁴.

3.1 The (stationary) log-linear representation

In this section we describe the linearized model expressed in stationary form¹⁵. The stationary representation is needed since we deal with the hypothesis of nonstationary technology shocks, that induce a common stochastic trend in real variables (Juillard *et al.*, 2008). To obtain model stationarity, we scale the real variables with respect to the stochastic technology level Z_t by imposing the transformation $X_t = \hat{X}_t Z_t$, where the "hat" superscript indicates that level variables are expressed in terms of stationary ratios. The model is then log-linearized around the steady state of the scaled (thus stationary) variables. Lower case letters with a "hat" thus denote log-deviations of the corresponding detrended variables.

3.1.1 Supply-side

On the supply-side, the linearized aggregate production function is the following:

$$\hat{y}_t = \alpha \hat{k}_{t-1} + (1 - \alpha) h_t - \alpha \log g_t^z \quad (3)$$

where we have assumed that firms produce their output $\hat{y}_t$ combining in a Cobb-Douglas production function their accumulated (thus firm-specific) capital endowment $\hat{k}_t$ with hired labor services h_t . The parameter α and its complement $1 - \alpha$ denote the capital and labor shares in production, respectively. The term $\log g_t^z$ is the growth rate of the labor-augmenting technology, which is assumed to follow a first-order autoregressive process $\log g_t^z = (1 - \rho_z) \log \gamma_z + \rho_z \log g_{t-1}^z + \varepsilon_t^z$, where γ_z is the deterministic long-run growth rate¹⁶. Under

¹⁴The main difference with respect to the model developed in Giuli and Tancioni (2009) is that utility is separable, log linear in consumption and CRRA in leisure. Moreover, households' preferences exhibit external habit persistence in consumption and there are no cost channels in the monetary transmission mechanism. Finally, we investigate more deeply the role of monetary policy by exploring different kinds of interest rate rules.

¹⁵More details are provided in the technical appendix of the paper, available upon request from the authors.

¹⁶The appearance of technology growth in the log-linear production function stems from the stationary representation of the model. It appears in each equation where dynamic relations among real variables are present. The formal derivation of the model

this specification of productivity growth, the evolution of the technology level has a non-stationary second-order autoregressive representation, since $\log Z_t = \log Z_{t-1} + \log g_t^z$ can be rewritten as $\log Z_t = (1 - \rho_z) \log \gamma_z + (1 + \rho_z) \log Z_{t-1} - \rho_z \log Z_{t-2} + \varepsilon_t^z$. The choice for such a flexible specification of the technology process is motivated by the need to separate the model-specific dynamics from that potentially emerging from a fairly general specification of the stochastic components. In fact, for a particular specification of the stochastic process of productivity, even RBC models can be made consistent with the contractionary effects of positive technology shocks, due to the operation of intertemporal substitution effects (Lindé, 2004).

We assume that firms face convex adjustment costs of changing their fixed assets holdings, that become productive with one quarter lag¹⁷. By log-linearizing the capital adjustment cost function we obtain the following capital law of motion:

$$\hat{k}_t = \frac{(\delta + \gamma_z - 1)}{\gamma_z} \hat{i}_t + \frac{(1 - \delta)}{\gamma_z} \hat{k}_{t-1} + \frac{(\delta - 1)}{\gamma_z} \log g_t^z \quad (4)$$

where $\hat{i}_t$ is the stationary log-deviation of gross investment and the parameter δ denotes capital depreciation. The steady-state growth parameter γ_z enters the convolutions of parameters of the dynamic equations since it represents the deterministic (thus expected) component in the long-run model solution.

3.1.2 Pricing behavior of firms

The aggregate price dynamics π_t , i.e. the pricing behavior of the monopolistically competitive firms obtained from the log-linearized first-order condition for profit maximization and the price index, is described by the following specification of the New Keynesian Phillips Curve (NKPC):

$$\pi_t = \iota_p \pi_{t-1} + \beta E_t (\pi_{t+1} - \iota_p \pi_t) + \kappa \widehat{mc}_t + \log u_t^\pi \quad (5)$$

where β is the discount factor, $\widehat{mc}_t = \hat{w}_t^r - \hat{y}_t + h_t$ is the log-linearized real marginal cost ($\hat{w}_t^r$ is the real wage), κ is the reduced-form NKPC slope coefficient and the stochastic term $\log u_t^\pi$ denotes a cost-push disturbance which is assumed to follow the stationary first-order

is provided in the Technical Appendix, available upon request from the authors.

¹⁷This assumption justifies why the capital stock enters the production function with one period lag.

autoregressive process $\log u_t^\pi = \rho_\pi \log u_{t-1}^\pi + \varepsilon_t^\pi$. The backward looking component in (5) is due to the hypothesis of partial indexation, whose degree is given by ι_p .

Equation (5) is compatible with a large class of models. The adoption of the firm-specific or rental capital specification (FSK or RK, respectively) and of the constant or endogenous demand elasticity of substitution among differentiated goods (CDE or EDE, respectively) only affects the convolution of parameters defining the reduced-form slope coefficient κ (Eichenbaum and Fisher, 2007)

Under RK and CDE, the reduced form NKPC slope coefficient (κ_{RK}) is given by $\kappa_{RK} = \frac{(1-\beta\theta_p)(1-\theta_p)}{\theta_p}$, where θ_p defines the fraction of firms that are not allowed to reset their price. Under FSK, the reduced form NKPC slope coefficient can be written as $\kappa_{FSK} = \kappa_{RK}\Lambda$, where Λ is a function of model's parameters, whose computation is not straightforward and can only be obtained with the method of undetermined coefficients¹⁸. However, Sveen and Weinke (2005) provide a useful approximation of Λ in terms of the capital share in production and of the elasticity of substitution among differentiated goods. Formally, $\Lambda = \frac{1-\alpha}{1-\alpha+\alpha\epsilon}$, where ϵ denotes the constant elasticity of substitution among intermediate goods.

The value of the NKPC slope coefficient under FSK thus decreases in ϵ . By reducing the responsiveness of prices to marginal costs, the firm-specific capital hypothesis implies a slower operation of the price mechanism as a reequilibrating device, and a bigger role for demand constraints. This provides a key insight on the role of strategic complementarities among intermediate sector firms for their price setting behavior. The economic rationale for the flatter slope of the NKPC under FSK is that, since firms operate with a predetermined (firm-specific) stock of capital, their marginal cost increases in the level of output. Compared to a situation in which capital services can be chosen period by period in a rental capital market, this implies that reoptimizing firms facing a positive productivity shock are induced to cut their prices by a smaller amount, since they anticipate that price reductions eventually lead to higher marginal costs due to increased demand and output at the firm level.

Aside from the firm-specificity of capital, there are other theoretical hypotheses that can induce strategic complementarity in price setting. Among these hypotheses, the EDE assump-

¹⁸A description of the procedure to derive κ with the method of undetermined coefficients can be found in Christiano (2004) and Eichenbaum and Fisher (2007).

tion has been employed in the literature (Eichenbaum and Fisher, 2007; Smets and Wouters, 2007). In such a case, the only difference is that the coefficient relating inflation to the marginal cost in the EDE model (k^{EDE}), irrespective of the RK or FSK specification, is further reduced by the factor $\frac{1}{\frac{1}{\epsilon-1}\phi_k+1}$, where ϕ_k is the percent change in the elasticity of demand due to a one percent change in the relative price of the good, evaluated in the steady state, and $\frac{1}{\epsilon-1}$ now denotes the firm's steady state markup (Kimball, 1998). The relation between the CDE and EDE specification of the reduced-form NKPC slope coefficient is thus approximated by the following equation:

$$k^{EDE} = k^{CDE} \frac{1}{\frac{1}{\epsilon-1}\phi_k + 1}$$

The fact that both the FSK and the EDE models differ from the baseline rental capital model in the size of the NKPC slope coefficient makes the different model specifications observationally equivalent when considering a specification in which the reduced-form coefficient κ enters the NKPC. This allows us to estimate the model in a form that is consistent with that adopted by Altig *et al.* (2005), i.e without specifying whether capital is firm-specific or rental and whether the elasticity of substitution among differentiated goods is constant or endogenous. From the estimated value of the reduced-form coefficient κ , given some hypotheses on the demand elasticity parameter and on the parameter defining its degree of endogeneity, we can then infer the degree of price duration in each model alternatives¹⁹.

3.2 Pricing behavior of wage setters

Turning to the pricing behavior of the monopolistically competitive wage setters, by log-linearizing household's first-order condition, the following real wage equation is obtained:

$$\begin{aligned} \hat{\pi}_t^w &= \beta E_t [\hat{\pi}_{t+1}^w + (\pi_{t+1} - \iota_w \pi_t) + \log g_{t+1}^z] - \log g_t^z - (\pi_t - \iota_w \pi_{t-1}) + \\ &\quad + \kappa^w (\widehat{mrs}_t - \hat{w}_t^r - \log \chi_t) + \log u_t^{\pi^w} \end{aligned} \quad (6)$$

where $\hat{\pi}_t^w = \hat{w}_t^r - \hat{w}_{t-1}^r$ denotes real wage growth and $\widehat{mrs}_t = \frac{\gamma_z}{\gamma_z - \lambda} \left[\hat{c}_t - \frac{\lambda_c}{\gamma_z} (\hat{c}_{t-1} - \log g_t^z) \right] + \eta h_t$ denotes the (real) marginal rate of substitution between consumption $\hat{c}_t$ and labor, where λ_c denotes the degree of habit persistence in consumption. The convolution of parameters

¹⁹To do this, we will employ a modified version of Christiano's matlab routines, in which we have added the EDE hypothesis and omitted the firm-specific labor specification. Details of these modifications are provided in the technical appendix, provided upon request from the authors.

$\kappa^w = \frac{(1-\beta\theta_w)(1-\theta_w)}{\theta_w} \frac{1}{1+\epsilon_H}$ defines the reduced-form slope coefficient, where the parameter θ_w defines the degree of nominal wage stickiness, while parameters η and ϵ_H denote the inverse Frish labor elasticity and the elasticity of substitution among differentiated labor services, respectively. The stochastic term $\log u_t^{\pi^w}$ denotes the wage-push disturbance, and is assumed to follow the first-order autoregressive process $\log u_t^{\pi^w} = \rho_{\pi^w} \log u_t^{\pi^w} + \varepsilon_t^{\pi^w}$. Even in this case, the backward looking component in (6) is due to the hypothesis of partial indexation to past inflation, whose degree is given by ι_w .

3.2.1 Demand-side

Concerning the demand-side of the economy, the dynamics of consumption $\hat{c}_t$ resulting from the corresponding Euler equation is described by:

$$\begin{aligned} \hat{c}_t = & \frac{\lambda_c/\gamma_z}{1 + \lambda_c/\gamma_z} (\hat{c}_{t-1} - \log g_t^z) + \left(1 - \frac{\lambda_c/\gamma_z}{1 + \lambda_c/\gamma_z}\right) E_t (\hat{c}_{t+1} + \log g_{t+1}^z) + \\ & - \frac{1 - \lambda_c/\gamma_z}{1 + \lambda_c/\gamma_z} (r_t - E_t \pi_{t+1} - \rho) - E_t (\Delta \log \chi_{t+1}) \end{aligned} \quad (7)$$

where π_{t+1} is the expected price inflation rate while r_t and ρ are the current and the steady state nominal interest rates, respectively. Current consumption thus depends on the expected real interest rate and on a weighted average of past and future consumption, with weights depending on the degree of external habit persistence λ_c . The term $\log \chi_t$ denotes a consumption preference shock and is assumed to follow the stationary first-order autoregressive process $\log \chi_t = \rho_\chi \log \chi_{t-1} + \varepsilon_t^\chi$.

The stationary investment dynamics depends on firms' choices on capital accumulation, defined by the log-linear capital Euler equation:

$$\begin{aligned} \hat{k}_t = & \frac{1}{(1 + \beta)} \hat{k}_{t-1} + \frac{\beta}{(1 + \beta)} E_t \hat{k}_{t+1} + \frac{1 - \beta \gamma_z^{-1} (1 - \delta)}{\epsilon_\psi \gamma_z (1 + \beta)} E_t \widehat{ms}_{t+1} + \\ & - \frac{1}{\epsilon_\psi \gamma_z (1 + \beta)} (r_t - E_t \pi_{t+1} - \rho) - \frac{1}{(1 + \beta)} \log g_t^z + \\ & + \frac{\beta}{(1 + \beta)} \log g_{t+1}^z + \frac{1}{\epsilon_\psi \gamma_z (1 + \beta)} [\beta \gamma_z^{-1} (1 - \delta) E_t \log \zeta_{t+1} - \log \zeta_t] \end{aligned} \quad (8)$$

where $\beta = (1 + \rho)^{-1}$ and $\widehat{ms}_{t+1} = \hat{w}_{t+1}^r + h_{t+1} - \hat{k}_t + \log g_{t+1}^z$ is the expected stationary log-deviation of return on capital, which under firm-specific capital is expressed in terms of firm's marginal savings in labor costs. The current installed capital thus depends on its past and

expected future values, on the expected marginal saving and on the expected real interest rate. Moreover, the investment/capital dynamics is assumed to be affected by a stationary first-order autoregressive disturbance $\zeta_t = \rho_\zeta \log \zeta_{t-1} + \varepsilon_t^\zeta$ to the convex capital adjustment cost function, whose steady-state elasticity is provided by the parameter $\epsilon_\psi > 0$.

3.2.2 Model closure

The model is closed considering the aggregate resources constraint and the policy reaction function. The log-linearized aggregate resources constraint is given by:

$$\hat{y}_t = (1 - \psi - g^y) \hat{c}_t + \psi \hat{i}_t + g^y \hat{g}_t \quad (9)$$

where $\psi = \frac{\alpha(\delta+\gamma_z-1)}{\frac{\epsilon}{\epsilon-1}\gamma_z(\rho+\delta)}$ is the steady-state investment to output ratio and g^y is the steady state public expenditure to output ratio. The term $\log u_t^{me}$ is an AR(1) measurement error capturing the evolution of public expenditure and of other exogenous components affecting the aggregate resource constraint. To account for the relevant changes in US net exports, we follow Smets and Wouters (2007) and assume that exogenous expenditure is also affected by the technology shock, i.e. $\hat{g}_t = \rho_g \hat{g}_{t-1} + \varepsilon_t^g + \rho_{z,g} \varepsilon_t^z$. The coefficient g^y captures the long-run value of the public expenditure to GDP ratio.

The empirical monetary policy reaction rule considers the inflation deviation from a nonzero policy target π^* and a weighted average considering the output growth deviation from the deterministic long-run rate of growth $\Delta y_t - \log \gamma_z = \Delta \hat{y}_t + \log g_t^z - \log \gamma_z$ and the expected theory-based output gap growth $\Delta \hat{y}_t - \Delta \hat{y}_t^p$, where $\hat{y}_t^p$ is the level of output that would have prevailed in the absence of nominal rigidities. The policy instrument is adjusted gradually, giving rise to interest rate smoothing:

$$\begin{aligned} r_t = & \rho + \rho_r r_{t-1} + (1 - \rho_r) [\phi_\pi (\pi_t - \pi^*)] + \\ & + \phi_{\Delta y} [\tau (\Delta \hat{y}_t + \log g_t^z - \log \gamma_z) + (1 - \tau) E_{t-1} (\Delta \hat{y}_t - \Delta \hat{y}_t^p)] + \log u_t^r \end{aligned} \quad (10)$$

The parameter ρ_r denotes the degree of interest rate smoothness, while ϕ_π and $\phi_{\Delta y}$ define the strength of the policy reaction to inflation and the weighted average (τ) of output growth deviations from the respective targets, respectively. The stochastic term $\log u_t^r = \rho_{ur} \log u_{t-1}^r + \varepsilon_t^r$ denotes a stationary first-order autoregressive monetary policy error.

We are aware of the fact that a policy rule targeting potential output only (i.e. $\tau = 0$) would make the negative hours and investment response less likely, while a policy rule targeting the long-term growth rate (i.e. $\tau = 1$) implies that authorities underestimate the actual growth rate of natural output, resulting in a not fully accommodative interest rate response. However, there are at least three reasons justifying our choice for an empirical rule: first, since our interest is mainly empirical, the chosen policy rule should be consistent with both the standard and the inverted hours and investment response for a reasonable parameterization of the model. Secondly, a rule targeting the theory-based output gap requires perfect knowledge of the natural level of output, which is - by definition - unobservable. In other terms, we have to consider how policy makers - in real life operations - form their opinions about the natural level of output, which is hardly in monetary authorities' information set. This is why, as suggested by McCallum (2001), we make a step towards greater realism by introducing the expectations of the output gap growth formed one period in advance. Thirdly, given that real-time data on potential output are subject to relevant imperfections, under model uncertainty and when technology evolves according to a random walk with drift process, the estimated long-term deterministic growth component γ_z might represent the best prediction for output growth. For this reason, policy effectiveness resulting from targeting a misperceived output gap can result inferior to a policy rule responding to output deviations from trend (Orphanides, 2003a, 2003b, 2007; Del Negro *et al.*, 2005). Nevertheless, in the estimation process, we assume a weakly informative prior for τ and let the data speak on which measure of the output gap is preferred.

The linearized system is thus composed of four behavioral equations (7), (8), (5) and (6), the production function (3), the permanent inventory equation (4), the aggregate resources constraint (9) and the empirical Taylor rule (10). Four definitory equations for $\hat{w}_t^r$, $\widehat{ms}_t$, $\widehat{mrs}_t$, $\widehat{mc}_t$ complete the economic system.

The stochastic component is described by seven autoregressive processes. As previously stressed, in order to match the stochastic properties of the data, we have assumed that the stochastic process describing the evolution of the labor-augmenting technology level Z_t is permanent and potentially follows a second-order autoregressive process. For the remaining stochastic components (the aggregate utility shock χ_t , the labor supply preference shock ξ_t ,

the wage and cost-push shocks $u_t^{\pi^w}$ and u_t^{π} , the capital adjustment cost shock ζ_t , the monetary policy shock u_t^r and the government expenditure measurement error $\hat{g}_t$) we assume a stationary first-order autoregressive specification.

4 Bayesian estimation, model comparison and simulation

This section provides some details of the strategy adopted to estimate the structural parameters of the monetary DSGE model presented in the previous section. Our specification, in which a reduced form slope coefficient of the NKPC is estimated, does not imply any prior idea about the CDE/EDE or the RK/FSK specifications of the model.

4.1 The posterior distribution

The scope of Bayesian estimators is to obtain the posterior distribution for model parameters conditioning on prior beliefs on models M_j ($j = 1, 2, \dots$), structural parameters θ_j , and sample information. The methodology thus nests the formalized prior distribution $P(\theta_j, M_j)$ for the j -th model's parameters vector $\theta_j \in \Theta$, and the conditional distribution (pseudo-likelihood) $P(\mathbf{Y}_T | \theta_j, M_j)$, where $\mathbf{Y}_T = \{\mathbf{y}_t\}_{t=1}^T$ contains sample information, to get the posterior density $P(\theta_j | \mathbf{Y}_T, M_j)$. This is obtained by employing the Bayes rule:

$$P(\theta_j | \mathbf{Y}_T, M_j) = \frac{P(\mathbf{Y}_T | \theta_j, M_j) P(\theta_j, M_j)}{P(\mathbf{Y}_T, M_j)} \quad (11)$$

where $P(\mathbf{Y}_T, M_j)$ is the marginal data density, which can be normalized since it does not depend on θ_j .

The posterior distribution is basically the result of a weighted average of prior information and the conditional distribution (i.e. the empirical information). Weights are inversely related to the variance of the prior distributions and the variance of the sample information ("precisions"), respectively. Thus, formalizing a tight prior will result in a highly constrained estimation, while a diffuse prior will result in a weakly constrained estimation²⁰.

²⁰Asymptotically, the conditional distribution (objective information) dominates the prior distribution (subjective information) and the posterior distribution of the parameters collapses to their pseudo-true values. This property guarantees that the relevance of priors in posterior estimates vanishes as the sample size increases. A further feature of the Bayesian estimator that is particularly important in standard applications is that its small sample performances outperform those of the FIML estimator (Geweke *et al.*, 1997; Fernandez-Villaverde and Rubio-Ramirez, 2004).

The posterior density of interest is a complex nonlinear function of the deep parameters θ_j , thus its analytical calculation is not generally feasible. For this reason, we calculate the posterior distribution via numerical integration. Operationally, the Bayesian MCMC posterior estimates are obtained adopting a two steps procedure, first employing the Kalman smoother to approximate the conditional distribution and then the Hastings-Metropolis (M-H) algorithm to perform Monte Carlo integration²¹.

4.2 Data and prior distributions

4.2.1 Data

For the Bayesian estimation of our model we use the same US data sources employed for the SVECM analysis of Model 2, briefly described in section 2.3. The sample is thus composed of quarterly data observed over the sample period 1954:1 - 2007:2²². Following Del Negro *et al.* (2005) and Smets and Wouters (2007), seven quarterly time series are considered: the log differences of real GDP $\Delta \log y_t$, real consumption $\Delta \log c_t$, real investment $\Delta \log i_t$ and the real wage $\Delta \log w_t$, the log-levels of hours $\log h_t$, of GDP price inflation $\log \pi_t$ and of the federal funds rate $\log r_t$. The vector of observables employed in the estimations is thus:

$$\mathbf{x}'_t = \left[\Delta \log y_t \quad \Delta \log c_t \quad \Delta \log i_t \quad \Delta \log w_t \quad \log h_t \quad \log \pi_t \quad \log r_t \right]$$

Considering that we express the models in log deviations around the stochastic growth path ($\log Z_t$), the resulting measurement equations linking the model variables to observables

²¹The posterior mode is estimated employing the Sims' optimizer, while numerical integration is performed employing three parallel chains of 200000 M-H replications each. The fraction of drops of the initial parameters vector estimates is set at 30%. The scale parameter for the variance of the jumping distribution is calibrated such that we obtain an acceptance rate of nearly 25% in the three blocks. This value guarantees that the M-H algorithm considers the entire support of the posterior distributions. For the application of the Bayesian method we employ the open-source software Dynare-Matlab.

²²A detailed description of data sources and manipulations is provided in the Appendix.

are the following:

$$\begin{aligned}
\Delta \log y_t &= \hat{y}_t - \hat{y}_{t-1} + \log g_t^z \\
\Delta \log c_t &= \hat{c}_t - \hat{c}_{t-1} + \log g_t^z \\
\Delta \log i_t &= \hat{i}_t - \hat{i}_{t-1} + \log g_t^z \\
\Delta \log w_t &= \hat{w}_t - \hat{w}_{t-1} + \log g_t^z \\
\log h_t &= h_t + h \\
\log \pi_t &= \pi_t + \log g^p \\
\log r_t &= r_t - \log(\beta) + \log \gamma_z + \log g^p
\end{aligned} \tag{12}$$

where $\log g_t^z = (1 - \rho_z) \log \gamma_z + \rho_z \log g_{t-1}^z + \varepsilon_t^z$.

4.2.2 Prior distributions

Priors are basically defined according to the results obtained in previous analyses and to economic reasoning. However, we impose a prior parameterization for some key parameters such that the estimate is initialized over a parameter space for which the model does not replicate the SVEC-based evidence on the negative conditional correlation between productivity and investment.

To enhance the estimation of the key structural parameters, we impose one dogmatic priors to the 34-dimensional parameters vector θ' . We fix the discount factor $\beta = 0.998$ and the steady state value for the elasticity of substitution among differentiated goods $\epsilon = 41$. The high value of β is chosen to facilitate a data-consistent estimate of the long-run growth parameter γ_z and of the steady-state inflation parameter g^{p23} . Concerning the elasticity parameter ϵ , the chosen value implies a price mark-up $\epsilon(\epsilon - 1)^{-1}$ of 2.5%²⁴.

All the remaining parameters are estimated. The shape of the prior distributions is chosen according to the following standard practice: the reference distribution for the structural shocks is the inverted gamma which is consistently defined over the $\mathbb{R}^+$ range; for parameters theoretically defined in a $[0 - 1]$ range a beta distribution is assumed, while the normal

²³Our sample data indicate an average quarterly output growth rate of nearly 0.5%, while the average quarterly inflation and nominal rate are 0.88% and 1.42%, respectively.

²⁴The literature has employed very different values of the demand elasticity parameter. These range from a value of 3, as in Smets and Wouters (2007) to a value of 101, as in Altig *et al.* (2005). Giuli and Tancioni (2009) provide an estimate of the elasticity parameter employing a NK-DSGE model with firm-specific capital, obtaining a value of nearly 42, which implies a mark-up of nearly 2.5%.

distribution is adopted for priors on parameters theoretically defined over the $\mathbb{R}$ range.

For the parameter defining the steady state value of productivity growth (γ_z) we adopt a tight prior defined by a normal distribution centered on sample means (nearly 1.005), with standard deviation equal to 0.001. This value is basically the same employed by Smets and Wouters (2007). An informative beta-distributed prior is also adopted for α and δ , whose prior means (standard deviations) are set to 0.36 and 0.025 (0.02 and 0.002), respectively. The chosen mean values are standard in the literature and consistent with sample data on the labor share and capital depreciation. For the price/wage indexation parameters we assume a beta distribution with prior mean value $\iota_p = \iota_w = 0.75$ and prior standard deviation (0.1). The prior distribution for the inverse labor supply Frish elasticity parameter is assumed to follow a normal distribution centered around $\eta = 1$ with standard deviation equal to 0.1. The prior for the parameter defining the convex capital adjustment cost ϵ_ψ follows a normal distribution with prior mean 3 and standard deviation 0.2. For the habits persistence parameter we assume a Beta distribution with prior mean 0.5 and prior standard deviation 0.1²⁵.

For the reduced-form NKPC slope coefficient κ , we adopt a beta-distributed prior with mean 0.01 and prior standard deviation 0.001. Under the standard RK specification, this prior implies an average price duration of nearly ten quarters, i.e. a Calvo parameter value ($\theta_p = 0.9$) which is slightly above the available macroeconomic evidence obtained with the standard specification (Galí and Gertler, 1999; Smets and Wouters, 2003; Eichenbaum and Fisher, 2007; Del Negro *et al.*, 2005). Under FSK hypothesis, and given the chosen value for the elasticity of substitution among differentiated goods ($\epsilon = 41$ in the present application), the prior implies a lower Calvo parameter of about 0.6, which is slightly above the microeconomic evidence produced by Bils and Klenow (2004), suggesting an average price duration of nearly two quarters. The adoption of the EDE hypotheses, and for standard values of the curvature parameter of the Kimball aggregator ($\phi_\kappa = 10$), does not change significantly the implied Calvo parameter²⁶.

For the parameter defining the degree of nominal wage rigidity θ_w we adopt a beta-

²⁵The prior mean is in this case slightly below the values generally adopted in the literature. Our choice is justified by preliminary checks in which we obtained a lower posterior estimate of the habit persistence parameter.

²⁶In fact, the sensitivity of the NKPC slope to the curvature parameter ϕ_κ is increasing in the steady state mark-up value, thus decreasing in the elasticity parameter ϵ .

distributed prior with mean 0.5 and prior standard deviation 0.1.

The coefficients of the monetary policy reaction rule are assumed to follow a normal distribution centered on the standard prior mean values $\phi_\pi = 1.5$ and $\phi_{\Delta y} = 0.2$, with standard deviations equal to 0.1 and 0.05, respectively. For the coefficient τ , defining the weight assigned to trend output deviations and to the theory-based output gap, we assume a Beta distribution with prior mean 0.5 and prior standard deviation 0.1. The autoregressive coefficient ρ_r , defining the degree of interest rate smoothness, is beta-distributed with prior mean 0.25 and prior standard deviation 0.1²⁷.

For the standard errors of innovations we assume a prior mean of 0.01 with two degrees of freedom. These diffuse priors on perturbations reflect a very imprecise opinion about the dimensionality of shocks. Even if we assume that all shocks are serially correlated, we adopt differentiated priors for the autoregressive coefficients. For that defining the autocorrelation in differences, i.e. for the AR(2) non stationary technology process, we adopt a prior zero-mean normal distribution with a diffuse prior standard deviation of 0.2. For the autoregressive coefficients of the stationary stochastic components affecting capital adjustment costs and the aggregate resource constraint we adopt a beta distribution with a prior mean of 0.75 and standard deviation of 0.10. The coefficient capturing the effects of productivity improvements on net exports is assumed to be beta-distributed with prior mean 0.5 and standard deviation 0.1. For the remaining autoregressive coefficients we assume a beta distribution with prior mean 0.25 and standard deviation 0.1. The choice for a low degree of autocorrelation of the stationary disturbances favors the separation between stationary and nonstationary components (Smets and Wouters, 2003). Moreover, since the estimated Calvo parameters are particularly sensitive to the degree of persistence of the stochastic components, our choice should enhance the identification of the economic sources of persistence²⁸.

²⁷The prior mean for the interest rate smoothing parameter is decisively lower than that adopted in other applications. Our choice is justified by the need to initialize the estimates over a parameterization for which the monetary policy is sufficiently accommodative, such as ruling out the emergence of the contractionary effects of supply shocks on factor inputs under the prior parameterization. However, in preliminary checks of the estimation process, we found that our model specification does not tend to select a high degree of inertia in monetary policy behavior.

²⁸This is particularly relevant for the estimation of the sources of persistence of price and wage dynamics. Smets and Wouters (2007), by assuming a higher prior value for the autoregressive coefficients (0.5), estimate a very high degree of stochastic persistence for price and wage dynamics (nearly 0.9). On the basis of our preliminary checks, this reduces the estimated Calvo parameters, i.e. the degree of persistence explained by well-defined economic relations.

These prior opinions on structural parameters are summarized in the first three columns of Tables 1 and 2.

4.3 Results

Tables 1 and 2 report the posterior mode and mean estimates of the estimated 32-dimensional parameters vector θ . Table 1 contains the estimates of the 17 parameters defining the model structure, while Table 2 the estimates of the 15 parameters defining the persistence and size of the 7 exogenous stochastic components.

Table 1 - Priors and posterior distribution of structural parameters

	Prior distribution			Posterior distribution				
	Distr	Mean	St.Dev	Mode	St.Dev	Mean	5%	95%
γ_z	$\mathcal{N}$	1.004	0.001	1.005	0.001	1.005	1.004	1.006
γ^p	$\mathcal{N}$	1.009	0.001	1.009	0.001	1.009	1.008	1.010
g^y	$\mathcal{B}$	0.200	0.025	0.246	0.024	0.248	0.209	0.286
α	$\mathcal{B}$	0.360	0.015	0.286	0.011	0.290	0.272	0.305
δ	$\mathcal{B}$	0.025	0.002	0.032	0.002	0.032	0.029	0.036
ϵ_H	$\mathcal{N}$	21.00	5.000	25.06	4.603	25.68	18.27	32.93
η	$\mathcal{N}$	1.000	0.100	1.017	0.099	1.021	0.857	1.186
λ_c	$\mathcal{B}$	0.500	0.100	0.359	0.047	0.384	0.302	0.468
ϵ_ψ	$\mathcal{N}$	3.000	0.200	3.680	0.181	3.667	3.376	3.973
κ	$\mathcal{B}$	0.010	0.001	0.009	0.001	0.009	0.007	0.010
θ_w	$\mathcal{B}$	0.500	0.100	0.858	0.026	0.853	0.810	0.897
ι_p	$\mathcal{B}$	0.750	0.100	0.860	0.059	0.833	0.734	0.942
ι_w	$\mathcal{B}$	0.750	0.100	0.407	0.076	0.424	0.296	0.550
ρ_r	$\mathcal{B}$	0.300	0.100	0.456	0.045	0.456	0.384	0.528
τ	$\mathcal{B}$	0.500	0.200	0.9575	0.031	0.943	0.896	0.992
ϕ_π	$\mathcal{N}$	1.500	0.100	1.471	0.068	1.500	1.382	1.607
$\phi_{\Delta y}$	$\mathcal{N}$	0.250	0.050	0.496	0.032	0.500	0.452	0.551

Table 2 - Priors and posterior distribution of shock processes

	Prior distribution			Posterior distribution				
	Distr	Mean	St.Dev	Mode	St.Dev	Mean	5%	95%
ρ_z	$\mathcal{N}$	0.000	0.200	0.213	0.052	0.218	0.135	0.301
ρ_ζ	$\mathcal{B}$	0.750	0.100	0.936	0.010	0.934	0.918	0.951
ρ_g	$\mathcal{B}$	0.750	0.100	0.975	0.009	0.970	0.954	0.987
$\rho_{z,g}$	$\mathcal{B}$	0.500	0.100	0.460	0.091	0.458	0.306	0.597
ρ_χ	$\mathcal{B}$	0.250	0.100	0.801	0.035	0.785	0.726	0.849
ρ_π	$\mathcal{B}$	0.250	0.100	0.080	0.042	0.106	0.028	0.181
ρ_{π^w}	$\mathcal{B}$	0.250	0.100	0.246	0.057	0.251	0.157	0.344
ρ_{u^r}	$\mathcal{B}$	0.250	0.100	0.060	0.025	0.068	0.026	0.107
σ_z	$\mathcal{IG}$	0.010	∞	0.010	0.001	0.010	0.009	0.011
σ_ζ	$\mathcal{IG}$	0.010	2	0.013	0.001	0.013	0.011	0.015
σ_g	$\mathcal{IG}$	0.010	2	0.020	0.003	0.020	0.016	0.024
σ_χ	$\mathcal{IG}$	0.010	2	0.012	0.001	0.013	0.010	0.015
σ_π	$\mathcal{IG}$	0.010	2	0.004	0.000	0.004	0.003	0.004
σ_{π^w}	$\mathcal{IG}$	0.010	2	0.005	0.000	0.005	0.004	0.006
σ_{u^r}	$\mathcal{IG}$	0.010	2	0.005	0.000	0.005	0.004	0.006

According to the estimated posterior standard deviations and the implied pseudo t-values, all parameter estimates appear significant at the standard significance level. The estimated autoregressive coefficient of the technology growth process is $\hat{\rho}_z = 0.22$ and significant, indicating an integrated second order autoregressive process for the stochastic process driving technology. The preferred AR(2) specification implies that the impact and the long-run responses of productivity are not the same; however, the low posterior estimate for ρ_z implies that the relevance of intertemporal substitution effects due to expected increases in productivity is negligible. This basically provides evidence in favor of a New Keynesian explanation of the contractionary effects of supply shocks as opposed to explanations based on technological diffusion delays (Rotemberg, 2003).

Concerning the stationary disturbances, we obtain a high degree of autocorrelation for the process for the measurement error entering the aggregate resource constraint equation and for the stochastic process affecting capital adjustment costs ($\hat{\rho}_g = 0.97$ and $\hat{\rho}_\zeta = 0.93$). A very moderate degree of autocorrelation, with the exception of the shock affecting preferences ($\hat{\rho}_\chi = 0.78$), is obtained for all the remaining shocks. The exogenous innovations are all significant according to their standard errors and there are signals that the most important

shocks are those affecting productivity, the capital adjustment cost and the aggregate resources constraint.

The steady state value for productivity growth is estimated to be around 0.49% on a quarterly basis, and 95% of its posterior distribution lies over the intervals $[0.38\%, 0.59\%]$. These values are strictly in line with priors and sample moments for the output growth. The estimated mean values of the Cobb-Douglas coefficient $\hat{\alpha} = 0.29$ and capital depreciation $\hat{\delta} = 0.03$ are instead relatively distant from the corresponding priors and from sample information on the labor share and depreciation.

The posterior mean estimates for the behavioral parameters are close to the respective modal values and denote reasonable estimates when compared to our prior opinions and to the results of the related literature. There are however some exceptions that are worth mentioning. The estimate of the habit persistence parameter is lower than the prior and that addressed by the literature ($\hat{\lambda}_c = 0.38$), while the estimated degrees of indexation show that there is more inflation indexation on the firm side than on the worker side ($\hat{\iota}_p = 0.83, \hat{\iota}_w = 0.42$). Moreover, the estimated Calvo parameter on wage contracts is nearly 0.85, consistent with an average wage duration of nearly seven periods.

The estimated reduced form NKPC slope coefficient is slightly lower than the prior mean value ($\hat{\kappa} = 0.009$), indicating a particularly weak transmission mechanism from marginal costs to price inflation. This value lies in the wide range of the estimated NKPC coefficients provided by the vast literature on the subject (Schorfheide, 2008). Considering the standard RK-CDE model specification, the estimated slope coefficient implies a price Calvo parameter θ_p of nearly 0.9. This value is reduced to 0.6 when considering the FSK-CDE and the FSK-EDE specifications. Given the preferred calibration of the demand elasticity parameters ϵ and ϕ_κ , the EDE hypothesis is thus not relevant. However, for different calibrations of ϵ and ϕ_κ , and in particular for calibrations in which a high steady state mark-up is assumed (low values of ϵ), the EDE hypothesis would be crucial for obtaining a low value of the Calvo parameter. Unfortunately the elasticity parameters ϵ and ϕ_κ are unobserved and unidentifiable in the present setting, thus we are left to conjectures over their reasonable values.

Concerning the demand elasticity parameter, previous analyses provide little help, as very different calibrations have been employed in the literature. The literature provides little

information even for the calibration of ϕ_κ , as the different values employed are basically chosen in order to make the different estimates of the reduced form NKPC coefficient consistent with the different calibrations of the demand elasticity and Calvo parameters (Eichenbaum and Fisher, 2007). Smets and Wouters (2007) adopt a very low value for the demand elasticity parameter ($\epsilon = 3$) and a curvature coefficient $\phi_\kappa = 10$, and obtain an estimated Calvo parameter of 0.66. Given their model specification, the implied reduced form NKPC slope coefficient is nearly 0.02, a value which is above our estimate. The reason for this result is that they obtain (and allow for) a very high persistence of the shock affecting the NKPC ($\hat{\rho}_\pi = 0.89$). In other terms, the weak correlation between price inflation and marginal costs present in the data tends to be explained by an exogenous component determining persistent (and parallel) translations of the NKPC. Our estimates are instead obtained with an almost negligible degree of persistence of the shock ($\hat{\rho}_\pi = 0.08$), which forces the explanation of the weak correlation between inflation and marginal costs in the slope of the NKPC.

The parameters defining the monetary policy reaction function are equally important in determining the demand response to a positive supply shock. In particular, our estimates suggest that the monetary authority attributes more weight to the output deviations from long-run growth than to the theory-based output gap, since $\hat{\tau} = 0.94$.

These results clearly drive the model over a parameter space for which the contractionary effects of positive supply shocks on inputs are relevant and persistent.

4.4 Posterior impulse responses

In this section we provide an evaluation of the dynamic properties of the model employing stochastic simulations based on posterior estimates. In discussing our results we will focus on the economic mechanism determining the sign and the persistence of the response of factor inputs to a positive technology shock.

Figure 2 shows the posterior impulse responses to a positive supply shock and compares them with those obtainable with the prior parametrization. Note that, under the prior, the technology shock is expansionary even in the short run. This signals that the posterior IRFs are not affected (qualitatively) by the peculiar prior parameterization employed in the estimation process.

There are at least three indications from posterior IRFs that merit to be highlighted. First, the estimates confirm the results obtained with the SVEC analysis presented in section 2, with the investment and hours responses qualitatively mirroring those obtained with the SVEC-based impulse response analysis. The investment response turns out negative in the short to medium run, and becomes expansionary after nearly seven periods, when the demand constraint becomes less binding as more firms cut their prices, letting the standard expansionary mechanisms display their effects. The posterior hours response is negative for nearly eight periods, indicating that both specifications are consistent with the well known productivity-employment puzzle. Secondly, consumption, output and real wage responses are quite standard: consumption increases in response to the expected permanent increase in productivity and output, driving the expansionary aggregate demand response (output), while the real wage increases following the increase in productivity. Thirdly, the interest rate reduction signals that the estimated monetary policy rule is quite accomodative but not enough to prevent a decrease in inflation. This result is particularly important for the empirical performances of our model, as it is consistent with our SVEC results and with the standard SVAR-based evidence. Our estimated model is thus able to reproduce the short/medium term puzzling evidence on labor and investment without compromising the empirical consistency of the estimated dynamics of the other real and monetary variables²⁹.

²⁹Note that Altig *et al.* (2005) obtain a counterfactual increase in inflation after a neutral technology shock, which they address as originating in a particularly accomodative monetary policy response, needed in order to account for the general rise in economic activity obtained with the benchmark SVAR.

Figure 2 - Prior and posterior impulse responses to a supply shock

When a technology shock hits the economy, the degree to which real activity follows its natural level depends on the resulting price cut, which is constrained by the Calvo scheme. Under FSK and EDE, the incentive to cut prices is countervailed by the increasing marginal cost at the firm level. In fact, with these assumptions at work, firms' marginal costs depend on their own pricing decisions. This implies that reoptimizing firms facing a positive supply shock are induced to cut their prices by a smaller amount than under the standard RK-CDE specification, since they expect price reductions to lead to higher marginal costs, due to increased demand and output. The increased inertial behavior of inflation implies a weaker demand response that may reduce the use of inputs in production.

Note that the mechanics of firms' investment behavior in the FSK and RK specifications is the same operating in standard frictionless models, which ultimately depends on Tobin's q dynamics stimulated by the shock. The introduction of nominal frictions leads to temporary deviations from natural levels. However, in the RK model and for reasonable degrees of stickiness, the emergence of gaps in Tobin's q does not overturn model predictions: since gaps are positive, a positive response in investment is obtained³⁰.

³⁰For a credible parameterization and a standard specification of the policy rule, the NK rental capital hypothesis may imply that Tobin's q gap is bigger than that emerging in the flexible price economy, because the central bank's reactions operate in

The demand constraint implied by the presence of nominal rigidities may reverse the effects of the supply shock on Tobin's q gap, but only in the presence of unrealistically high degrees of price stickiness. By contrast, under the FSK and EDE specification, the degree of nominal stickiness needed to reverse Tobin's q dynamics is lower and basically consistent with microeconomic evidence. This is why, in our opinion, Basu *et al.*'s (2006) conclusion that standard NK monetary models can account for the contractionary effects of supply shocks is not fully legitimate, since it requires an excessively high degree of nominal frictions.

5 Conclusions

This paper addresses the issue of the contractionary effects of positive supply shocks. By employing a Structural Vector Error Correction model, we first show that the short-term response of hours and investment to an identified productivity shock is negative. We then provide a theoretical explanation of the SVECM-based results by showing that they are consistent with a reasonably parameterized monetary sticky price DSGE model in which capital is firm-specific and/or demand elasticity is endogenous. These hypotheses ensure that the negative investment response is observed even for reasonable degrees of nominal stickiness, basically because of the flatter slope of the NKPC.

Our results are consistent with those obtained by Basu *et al.* (2006), who employ a purified measure of the Solow residual in VAR estimates, but contrast those obtained by the abundant literature employing unprocessed data in standard stationary SVAR representations and with those obtained by the empirical NK-DSGE literature imposing optimal or nearly optimal policy rules.

Considering the latter, our posterior estimates only partly confirm the results obtained by Smets and Wouters (2007), who estimate a NK-DSGE model with nominal and real frictions characterized by a particularly flexible dynamics, obtaining good adaptability to sample data and out of sample performances. However, even if they obtain a drop in hours following a positive technology shock, the impulse response of investment implied by their estimates is positive even on impact. The basic reason for this result is that Smets and Wouters (2007)

the same direction as increased expected capital returns. This is the basic reason why Galí and Gertler (2007) argue that, other things equal, in a NK environment firms may invest more than they would under flexible prices.

assume that the monetary authority targets flexible price output, resulting in a particularly accommodative rule which prevents the short-term contraction of investment after the productivity improvement.

Our results are instead distant from those provided by Altig *et al.* (2005), who develop a firm-specific capital monetary model that can account for the negative short-run response of investment to supply shocks even under a plausible parameterization. In fact, from the empirical evaluation of their model, based on an impulse response matching strategy, they are forced to rule out such a result, since their benchmark SVAR estimates produce a positive short-term response of both factor inputs. In our opinion, the evidence produced with a matching estimator applied to a model with a particularly flexible dynamics cannot be considered conclusive, since it does not provide additional evidence to that implied by the SVAR-based impulse responses. The matching estimator simply forces the model calibration to a parameter space for which the implied model dynamics is consistent with the SVAR estimates.

The reasons for our different outcomes are thus basically two: with respect to the SVAR-based evidence, we adopt an alternative and - in our opinion - more adequate SVEC representation of the joint data density. The cointegrated VAR is in fact consistent with the nonstationary properties of the data and allows a long-run identification approach providing explicit representation of the theory-based common stochastic trend among real variables; with respect to the NK-DSGE model-based evidence, the contractionary effects of productivity improvements on inputs are the result of the presence of relevant demand constraints emerging, on the one hand, because of the nearly flat NKPC, denoting a limited operation of the price adjustment mechanism, and on the other hand, because of the presence of a not fully accommodative monetary policy rule.

Our results thus provide additional evidence against the empirical relevance of standard flexible price models, which basically address supply shocks as main cause of the observed procyclicality of productivity, investment and hours. Our contribution to the debate is twofold: first, it provides support to the literature addressing the contractionary effects of positive technology shocks showing that not only hours, but also investment respond negatively to productivity improvements; secondly, it shows that the negative responses to productivity

improvements of hours and investment are theoretically consistent with monetary models taking into account the firm-specificity of capital and/or endogenous demand elasticity in addition to a not fully accommodative policy rule.

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6 Appendix

6.1 Sources of data and data transformations

Non farm business sector GDP (Y_t), personal consumption (C_t), private non residential investment (I_t), private wages (W_t), and scrapping (δI_t) are taken from the US Department of Commerce - Bureau of Economic Analysis (BEA) database. Employment (N_t), hours worked by the non farm civilian employees (H_t) and population aged 16 and older ($P16+$) are taken from the Bureau of Labour Statistics (BLS). Inflation is approximated with the quarterly log-differences of the GDP price deflator (PY_t). The nominal rate is the Federal Funds Rate (FFR_t) and is taken from the FRB's database.

Real GDP is expressed in chained 2000 dollars. Nominal consumption, investment, wages and scrapping are deflated with the chained price GDP deflator. This choice makes our dataset fully consistent with that employed by Altig *et al.* (2005), Del Negro *et al.* (2005) and Smets and Wouters (2007). Moreover, it eliminates the positive trend in the investment share resulting from the almost flat dynamics of the investment relative price deflator.

Hourly productivity, consumption, investment and wages are obtained by dividing the real variables by total quarterly hours worked.

Labor supply is obtained by dividing total quarterly hours worked ($N \times H$) by total available time. The latter is given by the product between population 16 and older and a quarterly-specific constant value approximating quarterly available time, obtained by considering 16 hours per day. This approach basically imposes a correction to the denominator in the hours to population ratio, which is needed to take into account the calendar effects on total available time. The use of hours to population ratios as labor supply measure is standard in the literature (see Christiano *et al.*, 2004, Galí and Rabanal, 2004, Chari *et al.*, 2005 and Del Negro *et al.*, 2005).

The capital stock employed in the SVEC analyses (K_t) is obtained applying the permanent inventory method (PIM) recursively from the first available observation for the capital stock. This procedure is necessary, since the fixed assets time series is available on a yearly basis only.

All the series are seasonally adjusted and entered in logs. The log quarterly nominal rate is obtained employing the transformation $r = \log(1 + FFR/400)$ to the original variable. The table below summarizes data sources and the manipulations from which the dataset has been obtained.

TableA1 - Sources of data and data transformations

Variable	Source	Definition	Database Table/Code	Transformation
Y	BEA	Business sector GDP GDP - government expenditure	NIPA Table 1.1.5	$y_t = \log\left(\frac{Y_t}{N_t \times H_t}\right)$
C	BEA	Personal consumption expenditures	NIPA Table 1.1.5	$c_t = \log\left(\frac{C_t}{N_t \times H_t}\right)$
I	BEA	Gross domestic investment Private sector	NIPA Table 1.1.5	$i_t = \log\left(\frac{I_t}{N_t \times H_t}\right)$
K_t	BEA	Fixed assets Private sector	NIPA Table 5.9	$k_t = \log\left(\frac{K_{t-1} - \delta I_t + I_t}{N_t \times H_t}\right)$
δI	BEA	Consumption of fixed capital Private sector	NIPA Table 1.7.6	-
W	BEA	Wage and salary accruals Private sector	NIPA Table 1.1.2	$w_t = \log\left(\frac{W_t}{N_t \times H_t}\right)$
PY	BEA	Implicit price deflator for GDP 2000 = 100	NIPA Table 1.1.9	$\pi_t = \log\left(\frac{PY_t}{PY_{t-1}}\right)$
FFR	FRB	Effective Federal Funds Rate	FF	$r = \log\left(1 + \frac{FFR}{400}\right)$
H	BLS	Weekly hours of production workers Private sector	CES0500000007	$h_t = \log\left(\frac{N_t \times H_t}{const \times P16+}\right)$
N	BLS	Non farm employees Private = total - government	CES0000000000 CES9000000000	-
$P16+$	BLS	Civilian non institutional population 16 years and older	LNU000000000	-

6.2 Impulse responses in the alternative SVEC specifications

Figure A1 - Impulse responses to a supply unit shock and 95% confidence interval region:
seven variables SVEC (k_t included)

Note: the IRFs of the benchmark specification are represented by solid lines

Figure A2 - Impulse responses to a supply unit shock and 95% confidence interval region:
six variables SVEC with stationary hours

Note: the IRFs of the benchmark specification are represented by solid lines

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