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WHAT DOES A TECHNOLOGY SHOCK DO?
A VAR ANALYSIS WITH MODEL-BASED SIGN RESTRICTIONS

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Abstract

This paper estimates the effects of technology shocks in VAR models of the U.S., identified by imposing restrictions on the sign of impulse responses. These restrictions are consistent with the implications of a popular class of DSGE models, with both real and nominal frictions, and with sufficiently wide ranges for their parameters. This identification strategy thus substitutes theoretically-motivated restrictions for the atheoretical assumptions on the time-series properties of the data that are key to long-run restrictions. Stochastic technology improvements persistently increase real wages, consumption, investment and output in the data; hours worked are very likely to increase, displaying a hump-shaped pattern. Contrary to most of the related VAR evidence, results are not sensitive to a number of specification assumptions, including those on the stationarity properties of variables.

JEL classification: C3; E3.

Keywords: technology shocks; DSGE models; bayesian VAR methods; identification.

Contents

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1 Introduction[1]

An important task of macroeconomics is to develop models that account for specific, quantitative features of the business cycle. Modern business cycle theory, emanating from the seminal work of Kydland and Prescott [1982], envisions a central role of random fluctuations in technological progress in driving the bulk of aggregate fluctuations. Precisely, when technology shocks as volatile and persistent as estimated total factor productivity (TFP) are fed through a standard real business cycle (RBC) model, the simulated economy appears to be able to replicate the patterns of *unconditional* volatilities and cross-correlations of key macroeconomic time series of the postwar U.S. economy (e.g., see King and Rebelo [1999]). This is a remarkable result for, as stressed by Uhlig [2003b], alternative, "demand-driven theories need to be worked pretty hard to cough up" key business cycle facts like the strong unconditional procyclicality of both labor productivity and hours worked.

The notion that technology shocks have anything to do with business cycles, however, has been recently questioned by a growing literature that aims at testing the predictions of the theory in terms of *conditional* moments in the data, i.e. conditional on technology shocks being the source of fluctuations. Galí [1999] originally identified technology shocks with structural VAR methods as the only source of a unit root in labor productivity. His results show that a positive technology shock induces a fall in hours worked so persistent that a negative conditional correlation between output and hours worked ensues. As stressed by Galí [1999], not only does this evidence, taken at face value, reject a key prediction of standard RBC theory, but it highlights a feature of the economy's response to aggregate technology shocks whose relevance goes beyond any specific macroeconomic paradigm. Because of the procyclicality of hours worked, some other shock(s) rather than technology shocks must be driving observed aggregate fluctuations.[2] While initially the structural VAR literature reached conclusions similar to Galí [1999]

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[2] Other recent contributions that, using different methodologies, have called into question the role of technology shocks in driving business cycles include Basu, Fernald and Kimball [1998] and Shea [1998]. See Christiano,

(e.g., see Francis and Ramey [2003]), the result that hours fall after a technology improvement has been disputed by several recent contributions, either challenging its robustness or radically questioning the "credibility" of long-run restrictions for identifying technology shocks.[3]

This paper reconsiders the important VAR evidence on the dynamic effects of technology shocks by proposing an identification scheme based on model-consistent sign restrictions. When this methodology is applied to the U.S. postwar data, an unexpected improvement in technology is found to lead to a significant and persistent rise in labor productivity, real wages, output, consumption and investment, and, in line with the predictions of standard RBC models, it is much more likely to drive hours worked up, not down.[4] With a 4/5 probability, a typical shock will increase U.S. hours worked per capita after one year. In addition, these results are consistent with the view that technology shocks play an important role in accounting for output fluctuations, although the uncertainty surrounding the contributions of these shocks to the variance of the forecast errors is large. Technology shocks, however, leave unexplained most of the variation in hours worked.[5]

The paper's contribution is twofold. First, in contrast to most of the VAR literature, technology shocks are identified by means of restrictions on the sign of impulse responses, similarly to the approach proposed by Canova and De Nicoló [2002], Faust [1998] and Uhlig [2001] for monetary policy shocks.[6] Differently from those contributions, however, the degree of agnosticism inherent in this kind of restrictions explicitly reflects uncertainty over the precise parameters values of a class of widely used dynamic general equilibrium (DSGE) models, encompassing most frictions proposed in the macroeconomic literature — like habits formation in consumption, investment adjustment costs, variable capacity utilization and nominal rigidities in prices and wages (e.g., see Christiano, Eichenbaum and Evans [2003]). These models, though

Eichenbaum and Vigfusson [2004] for some contrarian evidence.

[3] For an exhaustive survey of this large literature, beyond the selected contributions mentioned below, see Galí and Rabanal [2004] and the comments by McGrattan [2004] and Ramey [2004].

[4] In the working paper version, Dedola and Neri [2004], we report broadly similar results also for Japan and (West) Germany.

[5] See Kydland [1995] for a survey of the literature addressing the well-known fact that hours worked are too volatile in the data, relative to the predictions of standard RBC models.

[6] Several recent papers question whether one can properly identify technology shocks using long-run restrictions. For instance, Fisher [2002] and Uhlig [2003a, b] have convincingly argued that a unit root in labor productivity may result from permanent shocks other than the standard RBC shock to TFP, like shocks to the efficiency of investment, affecting the rate of transformation between current consumption and productive capital in the future, and to the capital income tax, respectively. Erceg, Guerrieri and Gust [2004] and Chari, Kehoe and McGrattan [2004] assess with Monte Carlo experiments the ability of long-run restrictions to recover the true impulse responses when applied to simulated data from calibrated models.

implying that across all parameterizations the responses of several variables to a positive shock to technology be positive for a number of quarters, are inconclusive concerning the effects on hours worked. The latter can either increase or fall depending on the values of key preference and technology parameters, independently of the presence of nominal rigidities.[7] Thus, while models with different implications are conceivable, these restrictions are likely to enjoy a fairly broad support, as they are derived for a wide range of parameterizations. Moreover, the sign restrictions we impose are weak in the sense that they lead to a plurality of candidate structural impulse responses. Rather than as a shortcoming, this is a potentially important advantage of this approach, for it eschews exact restrictions, such as exclusion restrictions, that are likely not to be robust to small perturbations to model specification and parameterization. For instance, our restrictions are valid independently of the fact that technology shocks be exactly nonstationary and the only source of a stochastic trend in labor productivity. Therefore, the full specification of the stochastic structure and long-run properties of the VAR model that is an essential part of structural VARs with long-run restrictions is not needed in our analysis.

In this respect, the paper's second contribution is to the debate in the VAR literature on the robustness of the evidence on the effects of technology shocks. Several recent papers have shown that the key findings in Galí [1999] are extremely sensitive to a number of auxiliary specification assumptions, including the selection of the specific variables entering the VAR and their transformation, and the data sample considered. As argued by Cooley and Dwyer [1998], long-run restrictions critically hinge on a careful distinction between the almost observationally equivalent trend- and difference-stationarity of the variables included in the VAR. Misspecification of these auxiliary assumptions, although inconsequential for many purposes, could severely impinge on the estimated dynamics of the VAR model. Christiano, Eichenbaum and Vigfusson [2003] give empirical content to this critique, documenting that the sign of the response of labor inputs to technology shocks identified with the same long-run restrictions as in Galí [1999] and Francis and Ramey [2003] is positive when hours worked per capita are assumed to be stationary and thus enter the VAR in levels, rather than in first differences as in the latter contributions.

This result has been confirmed by Galí and Rabanal [2004], who, however, raise a further issue, showing that the response of labor inputs is always negative when per capita hours worked are included in levels but detrended by (these authors' preferred) quadratic trend, or when instead total hours worked are used without a normalization by working age popula-

[7] As discussed in Section 3 below, Francis and Ramey [2003] argue that real business cycle models, suitably modified to allow for habit formation and capital adjustment costs, may be consistent with Galí's [1999] findings. The latter contribution originally suggested nominal rigidities as the most natural explanation for the negative response of hours worked.

tion — regardless of the assumed deterministic or stochastic trend (see also Ramey [2004]). Moreover, contributions like Galí, López-Salido and Vallés [2003] have shown that the effects of technology shocks estimated with long-run restrictions change drastically between the two sample periods before and after the early 1980's, in coincidence with the beginning of Paul Volcker's tenure at the helm of the Federal Reserve System. Precisely, a positive technology shock identified as in Galí [1999] brings about a decline in hours worked in the subsample up to the early 1980's, against a rise afterwards. These authors attribute this difference to a change in systematic monetary policy.[8]

In light of these diverse and contrasting findings, it is not unjustified to conclude that results on the effects of technology shocks estimated with long-run restrictions have been shown to be a rather mixed bag. By contrast, our results are not sensitive to a number of specification assumptions. First, we show that our findings are not affected by transformation of the variables — notably both per capita and total hours worked — in levels or first differences. Second, we document the robustness of the results to different sample periods, and to the adoption of a diffuse prior on the reduced form coefficients of the VAR entertained in our Bayesian inferential approach. Third, we argue that our approach is very unlikely to mix up technology shocks with other shocks that may entail a (more) positive response of labor inputs, like monetary policy shocks, price markup shocks and investment-efficiency shocks, as results are robust to imposing further restrictions to better rule out this possibility. Finally, we conclude showing that even when we focus on those (relatively unlikely) structural impulse vectors that explain a large fraction of labor productivity in the long run, we always find that hours worked, regardless of how uncertain their response on impact might be, sharply rise after a few quarters with a hump-shaped pattern. Overall, these results lend strong support to the view that theory-based (sign) restrictions are helpful in avoiding many of the subtle specification issues that arise when using long-run restrictions.[9]

The remainder of the paper is organized as follows. Section 2 outlines the identification approach with sign restrictions, while Section 3 briefly presents the benchmark model, reporting the theoretical impulse responses of a selected vector of variables that are used to identify technology shocks. Section 4 illustrates the results of the VAR analysis in terms of impulse re-

[8] However, Christiano, Eichenbaum and Vigfusson [2003] find that also this result depends on the transformation of hours worked used to remove the assumed trend. Dotsey [1999] first argued that with sticky prices the response of labor inputs to technology shocks crucially depends on whether the systematic response of monetary policy is accommodative.

[9] In the working paper version, Dedola and Neri [2004], we also investigate whether our approach has any inherent bias toward finding an increase in hours worked. Our results show that, when applied to simulated data from a model parameterized so that hours worked fall after a technology shock, the correct negative sign is recovered.

sponses and variance decomposition. In Section 5 the differences between our results and those in the VAR literature are investigated. Finally, Section 6 offers some concluding observations.

2 The identification framework with sign restrictions

In this section, we briefly describe our strategy to estimate the dynamic effects of technology shocks by means of sign restrictions, following Canova and De Nicoló [2002], and especially Uhlig [2001]. Both approaches yield nearly identical results when applied to identifying technology shocks with our sign restrictions. It is well-known that the reduced form of a VAR of order p has the following standard representation (omitting a constant c):

$$Y_t = B\,(L)\,Y_{t-1} + U_t,$$

where the vector Y includes the variables of interest in levels and $B\,(L)$ is a lag polynomial of order p. The covariance matrix of the vector of reduced-form residuals U_t is denoted as Σ. The reduced form can be estimated consistently using ordinary least squares, which, conditional on Gaussian U_t and initial conditions, is equal to the maximum-likelihood (ML) estimator. Identification in the structural VAR literature amounts to providing enough restrictions to uniquely solve — up to an orthonormal transformation — for the following decomposition of the $n \times n$ estimated covariance matrix of the reduced-form VAR residuals Σ:

$$\Sigma = \mathcal{A}_0 \mathcal{A}_0'.$$

This defines a one-to-one mapping from the vector of orthogonal *structural* shocks V to the reduced form residuals U, $U = \mathcal{A}_0 V$. Because of the latter orthogonality assumption, and the symmetry of Σ, at least $\frac{n(n-1)}{2}$ restrictions on $\mathcal{A}_0$ need to be imposed.[10]

The j-th column of the identifying matrix $\mathcal{A}_0$, a_j, is called an *impulse vector* in $\mathcal{R}^n$, as it maps the innovation to the j-th *structural* shock V_j into the contemporaneous, impact responses of all the n variables, Ψ_0. With the structural impulse vector a_j in hand, the set of all structural impulse responses of the n variables up to the horizon k, $\Psi_1, ..., \Psi_k$ can then be computed using the estimated coefficient matrix $B\,(L)$ of the reduced form VAR, $B_1, B_2, ...B_p$:

$$\begin{aligned} \Psi_s &= \sum_{i=0}^{s} B_{s-i}\Psi_i, && s \geq 1, B_{i-s} = 0, s-i \geq p; \\ \Psi_0 &= a_j. \end{aligned}$$

Proposition 1 in Uhlig [2001] shows that, given an arbitrary decomposition A_0 of the matrix Σ, *any* structural impulse vector a_j arising from a given identifying matrix $\mathcal{A}_0$ can be

[10] E.g., see Hamilton [1994], chapter 11.

represented as $A_0 q,$ for an appropriate vector q belonging to the hypersphere of unitary radius $\mathcal{S}^n \subset \mathcal{R}^n$.[11] For instance, natural candidates for the arbitrary decomposition A_0 are either the eigenvalue-eigenvector or the Cholesky decomposition of Σ. The basic idea of sign restrictions can thus be described as attributing equal probability to all possible structural impulse vectors a_j which, for a given reduced form estimate of the VAR, yield impulses responses whose signs are consistent with the assumed ones. Operationally, it is convenient to characterize the set of all consistent impulse responses by simulation, using the following algorithm suggested by Uhlig [2001]. For a given estimate of the VAR reduced-form matrices Σ and $B(L)$, yielding an arbitrary A_0, draw (a large number of) candidate q vectors from a uniform distribution over $\mathcal{S}^n$, compute the associated impulse vector a_j and impulse response matrix Ψ, discarding those that do not satisfy the assumed sign restrictions.

As argued by Uhlig [2001], the Bayesian approach, viewing the reduced-form VAR parameters as random variables, is particularly suited to interpreting and implementing sign restrictions. From a Bayesian point of view, sign restrictions amount to attributing probability zero to reduced-form parameter realizations giving rise to impulse responses which contravene the restrictions. To the extent that these restrictions do not lead to over-identification, they impose no constraint on the reduced form of the VAR. We can thus use standard Bayesian methods for estimation and inference, obtaining measures of the statistical reliability of estimated impulse responses. As shown by Uhlig [2001], under a standard diffuse prior on the VAR reduced form parameters $B(L)$ and Σ, and assuming a Gaussian likelihood for the data sample at hand, the posterior density of the reduced-form VAR parameters with the type of restrictions we implement will be just proportional to a standard Normal-Wishart. Therefore it is possible to draw from the posterior distribution of impulse responses consistent with our sign restrictions by jointly drawing from the Normal-Wishart posterior for Σ, $B(L)$ and the uniform over $\mathcal{S}^n$, discarding the realizations that violate the restrictions.[12]

It should be kept in mind that, as stressed by Uhlig [2001], the sign restriction approach amounts to simultaneously estimating the coefficients of the reduced-form VAR and the impulse vector. Draws of the VAR parameters from their unrestricted posterior which do not permit any impulse vector to satisfy the imposed sign restrictions are discarded as they receive zero prior weight. Therefore, below we check that our empirical results are not driven by the diffuse prior on the VAR reduced form, but mainly depend on our identifying assumptions.

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[11] As stressed by Canova and De Nicoló [2002], more generally any identifying matrix $\mathcal{A}_0$ can be expressed as the product of an arbitrary A_0 time a specific orthonormal matrix Q, such that $Q'Q = I$. Thus, the q in Uhlig's [2001] proposition is effectively the j-th column of the above Q matrix.

[12] Concretely, to draw from this posterior we use the program *montevar* described in the RATS manual (see Rats User Guide, Estima [2000]).

The procedure outlined above allows one to obtain estimates of impulse responses consistent with a given set of assumed sign restrictions, under the standard assumption in the structural VAR literature that all the structural shocks are orthogonal. Without any kind of *a priori* knowledge, it would be reasonable to assume a multivariate flat prior over the support of all possible responses $\Psi_0, \Psi_1, ..., \Psi_k$, given by an hypersphere in R^{nk} centered in 0. Economic theory can then be brought to bear, as in Canova and De Nicoló [2002] and Uhlig [2001], to shift all the probability mass to the event that the responses of $m \leq n$ variables (e.g., labor productivity, investment and so on) to the specific structural shock of interest have a given (positive or negative) sign for $s \leq k$ quarters. Clearly, this must also be the only shock that satisfies the sign restrictions. For instance, Uhlig [2001], by appealing to conventional wisdom, assumes that a contractionary monetary policy shock in the U.S. uniquely brings about a hike in the Federal Fund rate, a drop in the price level and a contraction in non-borrowed reserves. Differently from the previous contributions, in the next section we instead derive those sign restrictions from a class of DSGE models that most participants in the literature would accept, and explicitly take into account possible disagreement over parameter values, e.g. on the importance of nominal rigidities and other frictions, by simulating from a distribution function over these parameters, reasonably reflecting the degree of uncertainty over them. We also use the same model to argue that technology shocks uniquely satisfy the set of sign restrictions we use in the estimation.

3 Labor inputs dynamics in a benchmark DSGE model with real and nominal frictions

In this section we describe the model that is used as a laboratory to analyze the response of a set of variables to technology shocks. The model is basically the one estimated with different methods by Christiano, Eichenbaum and Evans [2003] for the U.S. and Smets and Wouters [2003] for the euro area. It features both real rigidities, in the form of adjustment costs for investment and variable capacity utilization, and nominal rigidities, namely sticky prices and wages. To save on space, we present only the linearized equations of the model, following the convention that a hat denotes deviations of variables either from their baseline long-run growth path (e.g. real consumption) or from their steady state (e.g. inflation). We will then consider impulse responses to technology shocks.[13] Since we are interested in implications in terms of

[13] We only focus on impulse responses to these shocks because the model has implications that would allow us to disentangle other shocks considered in the literature, like labor supply shocks (akin to labor tax rate shocks), markup shocks and preference shocks, from technology shocks. To save on space we do not report these impulse responses, that are available upon request.

the signs of the responses of variables that are robust across a broad range of parameterizations of the model, with and without nominal rigidities, we find it useful to assume that all structural parameters are uniformly and independently distributed over sufficiently wide ranges. However, very similar implications in terms of the sign of impulse responses would obtain if, to represent uncertainty over parameters, we used the posterior distribution estimated with Bayesian techniques, for example as in Smets and Wouters [2003].[14] A notable advantage of our approach, given the fundamental uncertainty on the best way to model the long-run behavior of hours in the U.S., is that it leaves this behavior unspecified in the model, as preferences are not restricted so that hours be stationary along the balanced growth path. This is consistent with our level specification of the VAR, that is agnostic on the best way to model the long-run properties of the data.

3.1 A benchmark DSGE model

The explicit consideration of a balanced growth path in which per capita real variables grow at the rate $1 + g$ implies that the subjective discount factor β in the linearized economy has to satisfy the following restriction, $\beta = b\,(1 + g)^{1-\sigma_c}$, as shown by King and Rebelo [1999], where $b \in [0.985, 0.995]$ is the discount factor in the level economy, implying an interest rate between 2% and 6.5% per annum — this latter value is the one assumed in King and Rebelo [1999]. We set $g = 0.004$, equal to the trend in U.S. labor productivity per hour worked over the 1955:1-2001:4 period. This implies a 1.6% annual growth rate in per capita output, investment and consumption.

Fluctuations in the model economy around the balanced growth path are driven by the standard RBC technology shock affecting total factor productivity, ϵ^z, and by an investment-specific technology shock, ϵ^i (see Greenwood, Hercowitz, and Huffman [1988], and Greenwood, Hercowitz, and Krusell [2000]).[15] As is customary in the macro literature, both shocks are assumed to have an autoregressive representation of order one with a coefficient $\rho_j \in [0.75, 1]$, $j = z, i$. This parameterization encompasses the case of an economy with unit root shocks to productiv-

[14] This is the approach followed by Peersman and Straub [2004], who use the signs implied for some variables by the posterior distribution of impulses responses estimated by Smets and Wouters [2003] to identify shocks in a VAR of the euro area.

[15] In a previous version of the paper we also illustrated the effects of shocks to capital income taxation, which have been suggested by some authors (e.g., Uhlig [2003a]) as posing a problem in identifying technology shock with long-run restrictions. Since these shocks bring about very similar effects to those arising from investement specific shocks, we do not report results on them. However, we will return to this issue in Section 5, when dicussing the robustness of our results.

ity; however the latter behavior is basically indistinguishable, in samples of the length of the U.S. postwar period, from that induced by values close to the upper bound of the assumed range of the autoregressive coefficients. Notice that at this stage we do not need to take a stand on the standard deviation of the shocks innovations, as the sign of the impulse responses will be invariant to it.

We consider both types of technology shocks for the following reason. In contrast to the standard RBC technology shock, ϵ_t^i does not have any immediate impact on the production function. Instead, it affects the rate of transformation between current consumption and productive capital in the future. Thus, any effects on current output must be the result of the ability of that shock in eliciting a change in the quantity of input services hired by firms. As argued by Galí and Rabanal [2004], this implies that in a model with nominal rigidities ϵ^z and ϵ^i can have different effects on hours worked but similar effects on the other variables of interest, like output and, through an increase in capacity utilization, labor productivity. Therefore, it is important to investigate whether these two different kinds of technology shocks can be distinguished on the basis of their dynamic effects on a larger set of variables.[16]

Given our assumption of separability between consumption and leisure, the Euler equation for consumption $\hat{c}_t$ is given by:

$$\hat{c}_t = \frac{h}{1+h}\hat{c}_{t-1} + \frac{1}{1+h}E_t\hat{c}_{t+1} - \frac{1-h}{(1+h)\,\sigma_c}\left(\hat{R}_t - E_t\hat{\pi}_{t+1}\right) \tag{1}$$

where the parameter $h \in [0.0, 0.8]$ measures the degree of habit formation, and the parameter $\sigma_c \in [1.0, 10]$ measures the inverse of the intertemporal elasticity of substitution for consumption (i.e., the risk aversion coefficient). The assumed ranges encompass most valued used and estimated in the literature. For instance the largest point estimate of h reported by Christiano, Eichenbaum and Evans [2003] is 0.71 (with a standard error of 0.03); these authors also set $\sigma_c = 1$. The variables $\hat{R}_t$ and $\hat{\pi}_{t+1}$ denote the nominal short-term interest rate and the inflation rate, respectively, that in the RBC economy are separately determined by the monetary policy rule, with no feedback to real variables.

[16] The argument in Galí and Rabanal [2004] is made informally in the context of a sticky price version of a model like that of Greenwood, Hercowitz and Krusell [2000], assuming for simplicity that the relationship $y_t = m_t - p_t$ holds in equilibrium, and that both m_t and p_t are pre-determined relative to the shock. In that case firms will want to produce the same quantity of the good but, in contrast with the case of neutral technology shocks, in order to do so they will need to employ the same level of inputs since the efficiency of the latter has not been affected (only newly purchased capital goods will enhance that productivity in the future). Notice, however, that to increase investment and reap the benefit of the shock, consumption will have to decline, given that output is fixed.

Because of adjustment costs, households choose the level of investment and capital according to the following linearized first order condition for investment:

$$\hat{\imath}_t = \frac{\beta}{1+\beta} E_t \hat{\imath}_{t+1} + \frac{1}{1+\beta} \hat{\imath}_{t-1} + \frac{\chi^{-1}}{1+\beta} \hat{q}_t + \frac{\beta}{1+\beta} E_t \epsilon^i_{t+1} - \frac{1}{1+\beta} \epsilon^i_t \tag{2}$$

where $\hat{q}_t$ is the price of installed capital goods in terms of consumption goods (Tobin's q), $\hat{\imath}_t$ is the level of investment, $\chi \in [0.0, 5.0]$ is the inverse of the elasticity of investment to the price of capital goods. The parameter χ is inversely related to the steady state value of the second derivative of the investment adjustment cost function. The largest point estimate in Christiano, Eichenbaum and Evans [2003] for this parameter is 3.24 (with a standard error of 0.47).

The optimal choice for the stock of capital is given by:

$$\hat{q}_t = -\left(\hat{R}_t - E_t \hat{\pi}_{t+1}\right) + \beta (1-\delta) E_t \hat{q}_{t+1} + \beta \overline{r} E_t \hat{r}_{t+1} \tag{3}$$

where $\hat{r}_t$ $(\overline{r})$ is (the steady state value of) the rental price of capital (determined solely by β and δ), and δ is the depreciation rate, usually assumed to be equal to 0.025 in the RBC literature (see Cooley and Prescott [1995]). Because of variable capacity utilization, the following approximate relation exists between the rental rate of capital and capacity, $\widehat{u}_t$:

$$\psi \hat{r}_t = \widehat{u}_t, \tag{4}$$

where $\psi \in [0.0, 50]$ is the elasticity of capital utilization with respect to the rental rate of capital. Thus, a zero value of ψ corresponds to the standard case in which capacity does not adjust. This parameter is not estimated by Christiano, Eichenbaum and Evans [2003], but set to 100 *a priori*.

The aggregate resource constraint and the capital accumulation equations close the real side of the economy:

$$\hat{y}_t = \frac{\overline{c}}{\overline{y}} \widehat{c}_t + \frac{\overline{i}}{\overline{y}} \widehat{i}_t = \alpha \widehat{k}_t + (1-\alpha) \widehat{l}_t + \alpha \psi \widehat{r}_t + \widehat{\epsilon}^z_t$$

$$\widehat{k}_{t+1} = \delta \widehat{i}_t + (1-\delta) \widehat{k}_t,$$

where the variable $\hat{\epsilon}^z_t$ represents the standard technology shock shifting the production possibility frontier, $\hat{l}_t$ is hours worked per capita, $\hat{k}_t$ is the capital stock, while α is the capital share in the (Cobb-Douglas) production function, usually assumed to be around $1/3$ in the RBC literature (see Cooley and Prescott [1995]). Notice that because of variable capacity utilization aggregate output is a function of the return on capital $\hat{r}_t$.

Nominal rigidities are introduced in the form of both wage and price stickiness. Households choose the level of nominal wage for the type of labor they supply in order to maximize their intertemporal utility function. As shown by Smets and Wouters [2003], the log-linearization of the first order condition for this problem delivers the following real wage equation:

$$\begin{aligned}\hat{w}_t &= \frac{\beta}{1+\beta}E_t\hat{w}_{t+1} + \frac{1}{1+\beta}\hat{w}_{t-1} + \frac{\beta}{1+\beta}E_t\hat{\pi}_{t+1} - \frac{1+\beta\gamma_w}{1+\beta}\hat{\pi}_t + \frac{\gamma_w}{1+\beta}\hat{\pi}_{t-1} \\ &\quad - \frac{1}{1+\beta}\frac{(1-\beta\xi_w)(1-\xi_w)}{\left(1+\frac{(1+\lambda_w)\sigma_l}{\lambda_w}\right)\xi_w}\left[\hat{w}_t - \sigma_l\hat{l}_t - \frac{\sigma_c}{1-h}(\hat{c}_t - h\hat{c}_{t-1})\right] \end{aligned} \tag{5}$$

where $\hat{w}_t$ is the real wage. The parameter $\xi_w \in [0.0, 0.8]$ measures the probability that the wage is not reoptimized in every period. The higher this parameter, the more sticky wages will be. The lagged term of the real wage $\hat{w}_{t-1}$ is introduced assuming that wages that are not chosen optimally are indexed to last period inflation rate. The parameter $\gamma_w \in [0.0, 1.0]$ measures the degree of indexation of wages to last period inflation. The larger this parameter, the more nominal wages are persistent. Clearly, the standard Euler equation for the labor choice under flexible wages, appearing in the above equation in brackets, is obtained by setting $\xi_w = \gamma_w = 0$. Christiano, Eichenbaum and Evans [2003], while setting $\gamma_w = 1$, report estimates of ξ_w within the above range, with a maximum value equal to 0.8. The parameter $\sigma_l \in [0.0, 10]$ measures the inverse of the elasticity of the labor supply. Finally, $\lambda_w \in [0.0, 1.0]$ measures the wage-setter markup, ranging from 0 to 100 percent.

The inflation equation:

$$\hat{\pi}_t = \frac{\beta}{1+\beta\gamma_p}E_t\hat{\pi}_{t+1} + \frac{\gamma_p}{1+\beta\gamma_p}\hat{\pi}_{t-1} + \frac{1}{1+\beta\gamma_p}\frac{(1-\beta\xi_p)(1-\xi_p)}{\xi_p}\left[\alpha\hat{r}_t^k + (1-\alpha)\hat{w}_t - \hat{\epsilon}_t^z\right] \tag{6}$$

is derived by linearizing the first order condition of the optimization problem of monopolistic competitive firms who choose the price to be set in order to maximize the expected discounted stream of future profits (see Smets and Wouters [2003]).

Allowing firms that do not reoptimize their price to adjust it to last period inflation rate delivers an equation in which current inflation depends on last period inflation. The parameter $\xi_p \in [0.0, 0.8]$ measures the probability the price of a good is not reoptimized in the current period. The higher this parameter, the more prices will be sticky. The parameter $\gamma_p \in [0.0, 1.0]$ measures the degree of indexation of prices. The larger this parameter, the more inflation is persistent. Again, setting $\xi_p = \gamma_p = 0$ recovers the standard expression for marginal costs with flexible prices and Cobb-Douglas production function, in brackets in the above equation. Christiano, Eichenbaum and Evans [2003], while setting $\gamma_w = 1$, report a maximum estimate of ξ_p equal to 0.92, but argue that this value is way to high given the evidence on individual price changes in Bils and Klenow [2002], implying that firms change prices roughly every 5 months on average. Therefore we set the upper limit to 0.8 — implying that the average duration of prices is 5 quarters at most.

Finally, the monetary authority sets the short-term interest rate according to the following Taylor rule:

$$\hat{R}_t = (1 - \rho_r)\,\rho_y \hat{y}_t + (1 - \rho_r)\,\rho_\pi \hat{\pi}_t + \rho_r \hat{R}_{t-1}, \tag{7}$$

with parameters $\rho_r \in [0.0, 0.99]$, $\rho_y \in [-0.25, 0.25]$, $\rho_\pi \in [1.1, 2.0]$, encompassing most values considered in the literature.

3.2 Deducing sign restrictions on impulse responses

We now present and discuss the impulse responses of the model's variables to the above two types of productivity shocks, with a view to deriving identifying restrictions on their sign. This identification strategy for VARs is very much in line with the methodology outlined by Canova [2002]. We assume that all structural parameters are uniformly and independently distributed over sufficiently wide ranges. Table 1 summarizes the ranges of the uniform distributions for the parameters of the model including real and nominal frictions. As argued above, these ranges cover reasonable values for the parameters, encompassing most calibrated and also estimated values used in the literature. Clearly, the distribution for the RBC model augmented with real frictions can be viewed as a particular case in which the (degenerate) density functions over the relevant parameters (namely, $\xi_p, \gamma_p, \xi_w, \gamma_w, \lambda_w$,$\rho_r, \rho_y, \rho_\pi$) have all the probability mass concentrated at zero.

In principle, the uniform densities on structural parameters would transpire into a pattern of the distribution of impulse responses that has richer implications than the sign restrictions we use in recovering structural shocks in the data. However, two considerations lead us to focus on sign restrictions only. First, the latter are more likely to be robust to changes in the specification of the functional form of the distributions of the structural parameters of the model economy. The sign restrictions we impose are broadly similar to those that would be obtained by estimating the parameters' posterior distribution of the model with Bayesian methods, as in Smets and Wouters [2003]. In this sense the uniform distribution on parameters can be thought as a convenient device to put discipline on the derivation of sign restrictions on impulse responses, without having to carry out the estimation of a complete model, which would require to take a stand on a number of specification issues, like the number of shocks and the appropriate detrending of the data. Indeed, when we estimated the model as in Smets and Wouters [2003] using our set of variables, we found that the impulse responses to a positive technology shock have signs that are consistent with those obtained with the uniform densities, but for hours worked. The latter actually fall with more than a 95 percent probability. Second, it is computationally more viable to impose sign restrictions in the context of Bayesian VARs, rather than a whole shape of the implied distribution of impulse responses, thus allowing us to use standard methods for estimation and inference, thus facilitating comparisons with most of the VAR literature.

In order to derive robust implications for the responses to technology shocks we carried out the following Monte Carlo simulation. We drew a large number of vectors of parameters from the uniform densities reported in Table 1 for the RBC model and the model with nominal rigidities (henceforth NR). For each draw we saved the responses to a one per cent positive neutral technology and investment-efficiency shock, and computed the 2.5 and 97.5 percentiles of their distributions point-by point. This ensures that parameters combinations that bring about extreme responses in the tails are ruled out.[17]

The results are reported in Figures 1A to 1D, displaying impulse responses up to 20 quarters. From Figures 1A and 1B, presenting the dynamic effects of a 1 percent positive shock to ϵ_t^z, it is clear that neutral technology shocks have qualitatively similar effects on real variables irrespective of nominal rigidities. Labor productivity, real wages, output, investment and consumption increase for several quarters. However, these positive responses can be more or less persistent, and revert to steady state more or less slowly, reflecting our rather uninformative densities over both the parameters governing the internal propagation mechanism and the serial correlation of the shocks. Moreover, for both the RBC and NR model, hours worked can either fall or rise depending on the parameterization, not only on impact but up to 20 quarters after the shock, with a median response that is negative for most quarters. Finally, Figure 1B also shows that, for the parameters range considered, the sign of the response of inflation and the short-term interest rate in the nominal rigidities model is *a priori* indeterminate as well. However, both variables always move in the same direction on impact, implying a positive correlation in the first quarter at least. The intuition for these results is straightforward. In the case of the RBC model, e.g. as argued by Francis and Ramey [2003], whether the response of hours worked is positive to a TFP technology shock depends on the strength of the investment adjustment costs and of the consumption habit formation. Precisely, the more relevant the latter two frictions, the less the representative agent will find convenient to raise investment and consumption in response to the increased productivity, and thus the only way to benefit from the shock will be to work less and enjoy more leisure. Conversely, as first stressed by Dotsey [1999], in the case of the model with nominal rigidities, the systematic response of monetary policy is a further determinant of the response of hours. Namely, the more technology shock are accommodated with a monetary

[17] For instance this can occur because of parameter values implying singularity of some of the matrices of the model's state space representation. In addition, since several parameterizations of the monetary policy rule in the nominal rigidities economy may transpire into local indeterminacy of the steady state, we discard draws that imply local indeterminacy. Clearly, in the presence of sunspots any exercise in identification of impulse responses to orthogonal shocks would be rather meaningless. See Lubik and Schorfheide [2004] for an estimated DSGE model that allows for indeterminacy arising from the monetary policy rule.

easing and a drop in interest rates, the more positive the response of labor inputs and output, other things equal.

What about investment efficiency shocks? In Figures 1C and 1D we report the 2.5 and 97.5 percentiles of the impulse responses to a 1 percent positive shock to ϵ_t^i. Figure 1C shows that, in a model without nominal rigidities, these shocks have radically different implications for many variables, relative to the neutral technology shock. In particular, in the face of an investment and output increase triggered by a rise in hours worked, they bring about a decline of consumption and labor productivity in the first few quarters. Interestingly, this occurs notwithstanding the fact that the model features variable capacity utilization, so that all these variables could in principle increase when hours increase.

Conversely, Figure 1D shows a less clear-cut picture for the (NR) model with nominal rigidities. An expansionary response of systematic monetary policy may bring about an increase in both investment and consumption on impact, by appropriately inducing a magnified increase in hours worked. However, since the benefit of forgoing current consumption for a higher investment level in the presence of such a shock is generally quite high, the median response of consumption remains always negative, and its maximum response (the 97.5 percentile), though marginally positive for the first couple of quarters, subsequently becomes negative — in contrast with the dynamic effects of a neutral technology shock displayed in Figure 1B.

Therefore, given the uniform distribution of parameter values in Table 1, there is a unique set of restrictions that allow to disentangle these two shocks in the data, independent of the presence of nominal rigidities, as the TFP technology shock entails a more persistent increase in both consumption and investment. Although such a positive comovement could be intuitively brought about by a very expansionary monetary stance in the face of an investment-specific shock, this kind of systematic policy response is quite unlikely when a standard monetary reaction function like (7) is assumed. Since, however, it is conceivable that modifications in the form of the Taylor rule we consider may better capture the reality of the operation of monetary policy in the face of this kind of shocks, in Section 5 we will thoroughly assess the implications of this issue for our empirical findings.[18]

From the above analysis of the theoretical impulse responses clearly emerges that, given the class of DSGE models that we consider under the assumed densities of structural parameters, the set of sign restrictions that we impose allow us to unambiguously disentangle (neutral) technology shocks from other shocks *a priori*. Precisely, the effects of a neutral technology shock could be separated not only from those of an investment efficiency shock, but also from those of an expansionary monetary policy shock. The latter would bring about a persistent

[18] As a similar result may hold for shocks to a capital income tax, we will address this possibility in Section 5 as well.

decline in the interest rate and an increase in inflation, inducing a negative comovement between these two variables. In contrast, they always move in the same direction on impact following a technology shock, as shown in Figure 1B.[19]

Given these results, we broadly interpret uncertainty over structural parameters as being consistent with the requirement that a positive technology shock increases labor productivity for the first 20 quarters, investment and output for the first 10 quarters, real wages for 17 quarters from the 3rd to the 20th, and consumption for the first 5 quarters, as summarized in Table 2. The response of hours, inflation and the short-term interest rate are left unrestricted. This is the set of restrictions on the signs of the impulse responses that are imposed in the VAR analysis below.[20] Since we include inflation and the short-term interest rate in our empirical analysis, it is natural to focus on the implications of the model with nominal rigidities. In addition, these implications are also less restrictive and thus more general than those implied by the model with only real frictions. For instance, the latter implies a very tight link between the responses of the real wage and labor productivity, that instead is not borne out in the model with nominal rigidities, as clearly emerges comparing Figure 1A with 1B.[21]

4 VAR evidence on technology shocks with sign restrictions

In this section, we begin by specifying the variables that enter in the VAR and the number of lags. We then proceed in illustrating the results on impulse responses and variance decompositions to technology shocks obtained for the U.S. economy, as well as conducting some sensitivity analysis. The variables that we include in the VAR are the logarithm of hourly labor productivity, real wages, per capita hours worked, per capita real investment, per capita real consumption, the quarterly gross inflation rate (based on the GDP deflator), all seasonally adjusted, and the quarterly gross short-term interest rate over the sample period going from 1955:1

[19] Importantly, we can also rule out confusion with price markup shocks, which, according to Galí and Rabanal [2004] estimates, play an important role in driving the procyclicality of hours worked. Under the assumed uniform densities, this kind of shocks implies that the difference between labor productivity and the real wage be persistently negative, whereas it is positive, at least on impact, after a positive technology shock. When we add the latter requirement to the restrictions in Table 2, the estimated impulse responses are virtually indistinguishable from those reported in Figure 2.

[20] We also investigated whether in the model the imposition of this set of sign restrictions, implying positive comovements among several variables, would constrain the behavior of hours worked, finding that the latter's response is broadly similar to that reported in Figures 1A-1B.

[21] In a previous version of the paper we also considered the more restrictive, RBC-consistent sign restrictions. The results were broadly similar. See the working paper version.

to 2003:4. Based on likelihood methods, we choose 3 lags, although results would be virtually unchanged with 4 lags.[22]

4.1 The dynamic effects of technology shocks in the U.S. economy

The estimated impulse responses to a positive technology shock obtained under the restrictions in Table 2 for the United States, and the associated variance decomposition, are presented in Figures 2 and 3. In each case the Figures show the median (the thick, solid line) and the 5th, 16th, 84th and 95th percentiles (the dashed lines) of the pointwise distribution of the variables responses, obtained from 500 draws from the unitary hypersphere $\mathcal{S}^7$ for each of 1000 draws from the posterior distribution of the reduced form of the VAR. Output per capita is constructed by adding up the responses of labor productivity and total hours worked, per capita.

The results presented in Figures 2 and 3 are based on around 35000 different impulse vectors a_j identified out of the total of 500000 draws.[23] Figure 2 shows that a positive technology shock determines a sizable increase in labor productivity, the real wage, consumption, investment and output that is also quite persistent: the 16th percentile of the responses of these variables is generally above zero even after 4 full years. The increase in investment is between 2 and 4 times larger than that of output. The response of consumption, generally less strong than that of output, is much more persistent than the assumed 5 quarters, displaying a pattern broadly similar to that of output. While the median response of labor productivity remains pretty much around the level on impact, that of the other variables displays more of a hump shape, reaching a maximum a few quarters after the impact and then declining, more fast in the case of investment. The maximum median response of consumption and output occurs after around 2 years.

Concerning the variables whose responses are left unconstrained by our identifying assumptions, a clear-cut result is obtained for hours worked. The median response of this variable is also positive and hump shaped, reaching a peak between 5 and 8 quarters after impact, and approaching zero by the 5-year horizon. Around its peak, this response is positive with over 0.8

[22] A detailed description of the data and its sources can be found in the Data Appendix. The first five variables are the same as those used by Francis and Ramey [2003] in their U.S. study. Conversely, in its largest, five variable system Galí [1999] includes, beside the ratio of GDP to total hours worked and total hours worked, our last two nominal variables and a monetary aggregate.

[23] The shapes of the distributions of the impulse responses are extremly robust to increasing the number of draws from both the posterior distribution of the reduced form VAR and the vector q from the unitary hypersphere. On average, we find that around 7% of the candidate impulse vectors satisfy the sign restrictions (37 out of 500 draws). Moreover, we never reject any of the 1000 draws from the VAR reduced-form posterior for lack of finding impulse vectors consistent with the sign restrictions.

probability in each single period for 6 quarters. This finding stands out against the fall estimated in the VAR literature studying technology shocks — with the notable exception of Christiano, Eichenbaum and Vigfusson [2003]. This is most apparent from Table 3, reporting the probability of hours worked being at a given horizon. Clearly, it shows that this probability is 0.8 for the first year after the shock, increasing to over 0.9 for the first 2 years.

Conversely, the effects of technology shocks on the short-term nominal interest rate and inflation in the U.S. appear largely inconclusive. While the response of inflation is more likely to be slightly negative in the first few quarters, with more than a 4/5 probability one quarter after impact, that of the interest rate is basically zero, with equal probability of being either positive or negative. This finding seems to be consistent with our view that both systematic and unsystematic monetary policy are not playing a big role in shaping our results. However, we will explore further this issue in the next subsection and in Section 5.

What are the implications of our estimates in terms of the contribution of the technology shocks to aggregate fluctuations? We address this issue by computing the percent of the variance of the k-step ahead forecast error that is accounted for by technology shocks. We find that (i) technology shocks cannot be ruled out as an important driving force of business cycles, and (ii) yet, to account for the bulk of cyclical fluctuations in hours worked (and inflation, interest rates), would require considering other sources of economic disturbances. In this latter respect, our results do not seem dissimilar from those obtained with long-run restrictions.

Figure 3 presents the variance decomposition results at horizons up to 40 quarters, also reporting the median and the pointwise 68 and 90 percent error bands. We see that technology shocks can explain up to over 50 percent of the variability in labor productivity, output, consumption, investment and real wages up to 5 years, although it must be noted that there is a large degree of uncertainty around these estimates. For longer horizons this fraction remains around 50 percent for most variables, with the notable exception of investment. The median fraction, however, is always lower and generally included between 20 and 30 percent. These shocks are very unlikely to come close to explaining 100 percent of the variability of labor productivity at any horizon, thus casting some doubts on identification strategies that exclusively rely on this kind of assumptions.[24]

The explained fraction of variability in hours is generally below 40 percent with 95 percent probability, with a median of around 5-10 percent only. Strikingly, this finding is pretty much in line with the results reported in Galí [1999] and Francis and Ramey [2003]. In this respect, it appears likely that the bulk of movements in hours should reflect shocks different from those affecting technology. However, this fact, i.e. that other shocks than technology shocks would

[24] As shown by Fisher [2002] investment specific shocks may play an important role in accounting for some of the unexplained variation in labour productivity.

be needed to account for important features of labor markets at business cycle frequency, has been well known to represent a major challenge to RBC theory, since the early contributions of Kydland [1984] and Christiano and Eichenbaum [1992].

Finally, turning to nominal variables, Figure 3 shows that the contribution of technology shocks to the forecast error variance of inflation and especially the short-term interest rate is generally quite limited, with a somehow higher ceiling in the short-run, at most up to 40 percent, falling to below 30 percent in the long run.[25]

4.2 Sensitivity analysis

In this subsection we investigate to what extent the above results are robust to the following three features of our procedure: (i) the inclusion of the variables in the VAR in levels; (ii) the adoption of a Bayesian approach with a joint diffuse prior on the VAR reduced form coefficients and the covariance matrix of the residuals; (iii) the assumption of no structural change in the sample. We think the first two checks are important in light of the controversy on the appropriate modelling of the time-series properties of the variables that has surrounded the identification of technology shocks with long-run restrictions. This may raise the legitimate concern that the two assumptions above be a source of bias of our results toward finding a positive response of hours worked.[26] Moreover, authors like Galí, López-Salido and Vallés [2003], among others, have argued that systematic monetary policy may have changed after 1979, and that resulted in a structural change in VARs parameters and in the effects of technology shocks, especially on hours worked. Therefore we also examine the subsample stability of our results to changes in the U.S. monetary policy regime.[27]

As shown below, our findings turn out to be quite robust to all these checks. Running our estimation with all real variables in first differences not only does not change our results, but actually leads to an even higher probability of a positive, persistent response of hours to a technology shock. Likewise, our results are not driven by the form of the prior on the VAR reduced form parameters and are robust across the two subsamples considered.

[25] In contrast, Christiano, Eichenbaum and Vigfusson [2003] find that technology shocks identified with long-run restrictions account for over 60 percent of the one step ahead forecast error variance of inflation, and almost 40 percent at even the 20 quarter horizon.

[26] See Phillips [1991] on how diffuse, “uninformative” priors can effectively turn out to imply strong restrictions on posterior estimates in the case of nonstationary time series.

[27] Clarida, Galí and Gertler [2000] show that monetary policy became more responsive to changes in expected inflation in the Volcker-Greenspan period; a similar result is obtained by Cogley and Sargent [2003]. On the other hand Sims and Zha [2004] find that changes in the variances of structural shocks are the major source of instability in a VAR including the main U.S. macroeconomic variables.

4.2.1 Level vs difference specification

Christiano, Eichenbaum and Vigfusson [2003] show that the findings in Galí [1999] are turned on their head when per capita hours worked are treated as a stationary process rather than as a difference stationary process. This result has been confirmed by Francis and Ramey [2003] and Galí and Rabanal [2004] with VARs specifications including variables different from those originally used by Christiano, Eichenbaum and Vigfusson [2003]. Since our VAR in levels can be viewed as extending that estimated in first differences by Francis and Ramey [2003], for it appends to their five-variable specification inflation and nominal interest rates, it is natural to ask whether our results are also sensitive to our assumption that all variables enter the VAR in levels.

We therefore applied our methodology to a VAR for the U.S. in which labor productivity, hours worked, consumption, investment and the real wage are in first differences, as in Francis and Ramey [2003]. The impulses responses presented in Figure 4, obtained with the same procedure as in Figure 2, show not only that our previous findings are broadly independent of the way variables in the VAR are modelled, but actually they come out stronger with this first difference specification. Technology shocks have a more persistent effect on labor productivity, real wages, consumption and investment, quite likely to be permanent. The 5th percentiles of the distribution of all these variables is now positive for the whole horizon. Remarkably, a similar behavior is displayed by hours worked, which are now even more likely than in the level specification to increase after a positive technology shock. The median response is always positive, gradually increasing from impact to around 0.4 percent. The probability of an increase in hours worked exceeds 4/5 from the 4th quarter on. Finally, the responses of inflation and the short-term interest rate is indistinguishable from that illustrated in Figure 2.[28]

Galí and Rabanal [2004] raise a further concern on the robustness to the VAR specification, arguing that the reversal in the response of labor inputs to a technology shock documented by Christiano, Eichenbaum and Vigfusson [2003] between the level and difference specification is due to a distortion in their estimated short-run responses, as a consequence of the presence of a spurious low frequency correlation between labor productivity growth and total per capita hours. Galí and Rabanal [2004] show that the response of labor inputs is always negative regardless of the transformation when total hours worked are used without a normalization by working age population. Therefore, to make sure that our framework is unaffected by this criticism, we carried out another experiment replacing per capita hours with total hours worked in both the

[28] Another experiment run with hours, inflation and the interest rate in levels, and all other variables in first differences, yielded very similar results. Namely, the response of hours worked was more persistent and more likely to be positive for a longer period than with the benchmark level specification.

level- and difference- specification VARs. As results are very similar to those presented above, for the sake of brevity we do not report them here. The key finding, however, is that the response of hours worked is more likely to be positive for even a *longer* period than with the benchmark specification.

Overall, these results show that, in stark contrast to the VAR-based literature on technology shocks, our findings are robust to the level or first difference specification of the VAR. This lends strong support to our view that theory-based sign restrictions are helpful in avoiding a great deal of the subtle specification issues that arise when long run restrictions are used.

4.2.2 Results from the maximum-likelihood estimates

Given that the level specification does not introduce any bias in our procedure, as a further check in this subsection we report results abstracting from the (diffuse) prior on the reduced form parameters of the VAR, and just consider only the uncertainty on the identification of the technology shocks. Precisely, we keep the values of the VAR parameters fixed at their OLS-Maximum Likelihood estimates and draw a large number of candidate impulse response vectors, discarding those that do not satisfy the sign restrictions in Table 2.

This exercise should highlight any bias in favor of estimating a positive response of labor inputs potentially introduced in the VAR posterior distribution by our diffuse prior, in case the latter was dominating the data likelihood. Given our interest in impulse responses, it is not immediately clear whether a prior that is diffuse over the VAR reduced form coefficients could be actually giving more weights to particular impulse response coefficients.

Figure 5 displays the estimated impulse responses of the variables in our system to a technology shock obtained from the OLS estimates of the VAR in levels and from drawing 50000 candidate impulse vectors from $\mathcal{S}^7$. As before, we report the median (the thick, solid line) and the 5th, 16th, 84th and 95th percentiles (the dashed lines) of the pointwise distribution of the accepted impulse responses.

The key result is as follows. All impulse responses are quite similar to those displayed in Figure 2. Fixing the VAR parameters and abstracting from the uncertainty on their estimation only makes the band between the 5th and the 95th percentiles slightly narrower. This shows that the posterior distribution from which we draw the realizations of the VAR reduced form coefficients is actually quite concentrated around the OLS-ML estimates, mainly reflecting the likelihood shape. Therefore, the prevailing source of dispersion in our estimated impulse responses clearly reflects the multiplicity of impulse vectors that satisfy our sign restrictions, qualifying as technology shocks. It is thus remarkable that they lead to quite definite conclusions on the response of hours worked to technology shocks.

4.2.3 Subsample stability

In this subsection we briefly discuss subsample stability of our specification. Galí, López-Salido and Vallés [2003] have found that the effects of technology shocks estimated with long-restrictions differ drastically between the two periods before and after Volcker's tenure at the helm of the Federal Reserve System. Precisely, a positive technology shock identified as in Galí [1999] brings about a decline in hours worked in the subsample up to the early 1980's, and a rise afterwards, because of the kind of systematic monetary policy adopted by the Federal Reserve System in the two subperiods. Due to the inclusion of inflation and the short-term interest rate in our VAR, our sample is actually different from those originally used by Galí [1999] and Francis and Ramey [2003], giving relatively more weight to the second sample used by Galí, López-Salido and Vallés [2003]. Here therefore we assess the robustness of our conclusions to the possibility of subsample instability.

Figures 6A and 6B present the estimated impulse responses of the variables in our system to a technology shock for the pre-1979 and post-1983 sample periods respectively, obtained using the same algorithm as in Figure 2.[29] As before, each figure shows the median (the thick, solid line) and the usual percentiles (the dashed lines) of the pointwise distribution in the indicated subsample.

The following results stand out. First, the qualitative patterns of all variables responses are broadly similar across both periods and to those estimated in the full sample. In particular, hours worked rise in a hump-shaped pattern in both subsamples. Interestingly, this increase appears to be slightly more likely in the early period, in which the 5th percentile is now positive from the 5th to the 12th quarter after the shock. Second, in the late period, the estimated effects of technology appear somehow smaller relative to the earlier period. For instance, the median response of labor productivity is always below 0.4 percent, whereas it close to 0.5 percent in the first few quarters in the earlier subsample. This is consistent with the well-documented drop in aggregate volatility in the last two decades.[30]

This evidence, similar to that obtained by Christiano, Eichenbaum and Vigfusson [2003] with long-run restrictions, is consistent with the view that the responses in the subperiods are the same as they are for the full sample and there is no break in the response of the interest rate and inflation to technology shocks. In particular, although in the first subsample a drop in the interest rate is slightly more likely as the median response is definitely negative for a couple of

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[29] The sample period ranging from 1979:3 to 1982:4 is avoided because of the nonborrowed targeting regime adopted by the Federal Reserve, which induced a significant increases in the volatility of the Federal funds rate (see Bernanke and Mihov [1998]).

[30] See Stock and Watson [2002], among others. Intriguingly, these authors argue that the decrease in volatility is mainly due to smaller, less volatile shocks.

quarters, this does not appear sufficient to reject the hypothesis of no sample break in the VAR in more formal a way. Nevertheless, the crucial finding from our perspective is that inference about the response of hours worked to a technology shock is not affected by subsample stability issues.

5 Interpreting the results with sign restrictions

In this section we first ask whether our results may be due to the fact that our approach mixes up technology shocks with other shocks that may bring about a more positive response of hours worked, like monetary policy shocks, investment efficiency shocks and capital-income tax shocks. It should already be clear from our analysis of the theoretical impulse responses in Section 2 that, given the class of DSGE models we consider under the assumed distribution of structural parameters, the set of sign restrictions that we impose should allow us to uniquely disentangle (neutral) technology shocks from other shocks. However, here we take a more empirical view, showing that our results are quite robust to controlling for interest rate cuts, changes in capital-income taxes, and to assuming a positive lower bound on the response of consumption — highly unlikely to be exceeded after an investment efficiency shock.

After having shown that our findings are not likely to be due to mistaken inference caused by the above sources of misspecification, we briefly investigate the reasons why our findings about the behavior of hours worked are different relative to most of the VAR literature using long-run restrictions. Even when we focus on those structural impulse vectors that explain a large fraction of labor productivity in the long run, we always find that hours worked, regardless of how uncertain their response on impact might be, sharply rise with a hump-shaped pattern. Moreover, this kind of impulse responses that yield dynamic effects similar to those estimated with long-run restrictions are also relatively unlikely. Most structural impulse vectors uncover technology shocks whose long run effects are somehow smaller and less persistent, and bring about an increase in hours worked within the first few quarters.

5.1 Are sign restrictions confusing different shocks?

In this subsection we turn to the task of investigating whether our results may be due to the fact that our procedure is retrieving not only technology shocks but also other shocks that may bring about a more positive response of hours worked, mixing up their effects. Obvious candidates are monetary policy shocks, investment efficiency shocks and capital-income tax shocks, given the discussion in Section 2.3. From the analysis of the theoretical impulse responses clearly emerges that the set of sign restrictions that we impose should allow us to unambiguously disentangle (neutral) technology shocks from other potentially important shocks *a priori*. For instance, while both investment and consumption rise following a shock to total factor productivity that boosts current output, they will tend to comove negatively in response to an investment-specific shock that does not shift the current production function, with consumption declining. The same reasoning applies to a negative shock to — a fall in — the capital-income tax, which also increases the cost of current consumption relative to future consumption (investment), leaving current production possibilities unaffected.

Nevertheless, our goal in this section is to go beyond these theoretical results, and assess the robustness of our findings more broadly. For instance, it is possible to write models in which a monetary expansion brings about a temporary increase in inflation because of cost channel effects, as argued by Barth and Ramey [2002]. Likewise, interactions between real and nominal frictions, and particularly systematic monetary policy, different from those we assumed in Section 2, may trigger an increase in consumption in response to an investment-specific shock.

Therefore, in order to address these concerns we carried out the following two experiments. First, in order to unequivocally rule out confusion with monetary policy shocks we redid our empirical analysis imposing the further restriction that after the shock the nominal interest be positive. Second, we checked the sensitivity of our inference on the behavior of hours to the requirement of a large and positive bound on the response of consumption, and to controlling for capital-income tax changes. Again, across all these experiments our results turn out to be broadly unaffected, thus confirming their robustness beyond the narrower validity of the assumptions underlying our analysis.

Monetary policy. In our first experiment, we complemented the restrictions in Table 2 with the requirement that the interest rate be positive for the first 2 quarters following the shock. This way it should be very unlikely that our identification strategy mistakenly picks expansionary monetary policy shocks for technology shocks.

Figure 7 presents the relevant results, again reporting the usual percentiles of the point by point distribution of the impulse responses. It is clear that the impulse responses are very similar to those depicted in Figure 2 with no restrictions on the short-term interest rate — obviously barring the latter's response. In particular, hours worked, if anything, are slightly more likely to increase immediately under this specification, as the 16th percentile hovers very close to zero for the first 10 quarters.

This evidence has at least two noteworthy implications. Not only does it strongly support the contention that our identification strategy does not mix up technology and monetary policy shocks, but it also suggests that our findings are difficult to rationalize in terms of other kinds of shocks — like an investement-specific technology shock — accompanied by an expansionary monetary policy stance that makes their effects look similar to those of a technology shock.[31]

Shocks to the consumption-investment transformation rate. Notwithstanding the above observation, we carried out an experiment aimed at uncovering a possible influence on our results of disturbances affecting the rate of transformation between consumption and investment. In particular we added to the benchmark list of restrictions the requirement that consumption be not only positive for 5 quarters, but also larger than the 16th percentile of its estimated response reported in Figure 2. The idea is that a relatively large response of consumption is very unlikely to be consistent with an investment-efficiency shock.

The results, reported in Figure 8, are again in broad support of our overall findings. Effectively, requiring a more pronounced positive response of consumption makes the increase in hours worked even larger. The 5th percentile is now positive from the 5th to the 10th quarter. Moreover, the responses of the other variables are barely affected, especially those of investment and the nominal short-term interest rate. In light also of the above evidence on the quite limited role that systematic monetary policy plays in shaping our results, if our approach was confusing different kinds of technology shocks with opposite effects on hours worked, the stronger response of consumption would have to be associated with a more negative response of hours worked, rather than a more positive one.[32] We also verified that including consumption durables in investment and leaving only nondurables in consumption proper did not change our findings. Results are available upon request.

Our last exercise was to examine the robustness of our results to capital-income tax shocks. Following Francis and Ramey [2003], we tackled this problem by constructing a series for the capital tax rate shock (as in Jones [2002]), and included it as an exogenous variable when estimating the reduced form VAR, before imposing the sign restrictions in Table 2. Since our results are again unaffected, to save on space we do not report them here. We also computed the correlation between our estimates of technology shocks in the U.S., across all identifications, and the AR(1) innovations to the series of the capital tax rate, interpreted by Jones [2002] as tax shocks. This correlation is not significantly different from zero.

[31] We also run an experiment requiring, in addition to the restrictions in Table 2, that inflation and the interest rate have the same sign for 2 quarters after the shock, as prescribed by the model with nominal rigidities. The results of this experiment, available upon request, confirm and actually even strengthen our original findings.

[32] 0.1cm

5.2 Exploring the long-run effects of technology shocks

As mentioned above, in contrast to standard structural VAR analyses, relying on just- or over-identifying restrictions to estimate a unique impulse vector that maps reduced form residuals into structural shocks one-to-one, our procedure yields a number of impulse vectors that have a structural interpretation. Thus, a useful starting point to understand the differences between our findings and those in the literature using long-run restrictions is to ask whether among those structural impulse vectors there is any subset that is associated with large and permanent effects. Obviously, there should be in principle just one, if any, impulse vector that accounts for all variation in labor productivity in the — however defined — "long run". Nevertheless, an advantage of our approach is that it allows to assess if quantitative changes in the amount of variation in labor productivity explained at a given distant horizon are reflected in qualitative changes in impulse responses. Therefore, among the set of structural impulse vectors that satisfy our sign restrictions, we selected those that account for over 70 percent of the forecast error variance of labor productivity after 10 years— i.e., the "long run" in this exercise is meant to be 40 quarters — rather than focusing on just one of them.[33] Given the results in Section 4.2.2, for simplicity the candidate impulse vectors were computed with the parameters of the reduced form VAR for the U.S. held constant at their OLS-ML estimates.

Figure 9 presents the usual five percentiles. The key results are as follows. First, the distributions are generally much less dispersed than those reported in Figure 5, for less than 5 percent of the impulse vectors exceed the 70 percent threshold, as should also be clear from the findings on variance decomposition in Section 4.1. However, for labor productivity, real wages, investment, inflation and the interest rate, the dispersion regarding the effects on the first couple of quarters is still substantial. The short-run effects on consumption and hours are instead rather tightly estimated. Therefore, the dispersion of the responses does not seem to depend only on whether a sign restriction is imposed on the specific variable.

Second, the dynamic effects of the shock on labor productivity, real wages, output, consumption and investment appear indisputably permanent, similarly to those estimated with long-run restrictions. Interestingly, however, the maximum fraction explained by the candidate shocks never exceeds 85 percent at the 10-year horizon. All these variables responds positively on impact and then rise reaching a new long run level. Output, consumption and investment display a marked hump-shaped pattern, peaking around 10 quarters after the shock, before converging from above to the new level.

Third, the impact response of the unrestricted variables, i.e. hours worked, inflation and the

[33] Results below are reasonably robust to changes in the 40 quarters horizon or the 70 percent variance decomposition threshold.

interest is now clear-cut: they all fall. Inflation and the short-term interest rate remain negative for 3 and 10 years, respectively. By contrast, hours worked strongly rise with a hump-shaped pattern, becoming positive after 5 quarters and peaking around 3 years at roughly 0.4 percent, to return to the baseline value only very slowly. Most importantly, this increase is such that the correlation at business cycle frequencies between the technology component of hours worked and output in the data — extracted using the band pass filter suggested by Baxter and King [1999] — is positive and significant, on average equal to 0.60. This result is clearly in contrast with the findings of Galí and Rabanal [2004] in a similar exercise based on long-run restrictions (see Figure 3 in their paper).

Our procedure thus recovers a subset of impulse responses implying dynamic effects that are very similar to those obtained by means of long-run restrictions. Nevertheless, in line with the results in Christiano, Eichenbaum and Vigfusson [2003], these permanent shocks still lead to a significant increase in hours, though with a few quarters delay. The level specification of hours, however, cannot be all the story, as should be clear from Section 4.2.1, unquestionably showing that our results are independent of whether the VAR is estimated in levels or first differences. There is another important message that stems from the impulse vectors identified by our procedure and *not included* in Figure 9. Across all the identification schemes satisfying the sign restrictions, those associated with the large, permanent effects reported in the figure account only for a fraction, though important, of all possible ones. The vast majority of structural impulse vectors imply that a positive technology shock has somehow smaller and less persistent effects in the long run, but brings about an increase in hours worked in the first few quarters. This finding is more in line with the RBC tradition, in which technology shocks are usually assumed to be very persistent but trend stationary. Interestingly, the different initial effect on hours worked may be easily rationalized with the different size and persistence of the two types of shocks and the implied different wealth and substitution effects on labor supply.[34]

6 Concluding remarks

This paper identifies technology shocks in VAR models of the United States by means of restrictions on the sign of impulse responses, derived from an explicit modelling of the uncertainty over the parameters of a popular class of dynamic general equilibrium models, encompassing both nominal and real rigidities. Technology shocks are found to bring about a significant and persistent increase in real wages, consumption, investment and output; hours worked increase

[34] See Rotemberg [2003] for an explanation of Galí’s [1999] finding in terms of slow diffusion of permanent technology shocks.

with a humped-shape pattern. In addition, the view that technology shocks may play a substantial role in accounting for business cycle fluctuations cannot be rejected, although these shocks leave unexplained most of the variation in hours worked.

This paper has focused on the estimation of impulse responses and variance decompositions to technology shocks. However, a natural question to ask is whether it would possible to draw implications on the parameterizations that are more likely to be associated with relevant features of the density of the estimated impulse responses. This is important as it could shed light on key aspects of the internal propagation mechanism of DSGE models, e.g., whether the fact that consequences of a technology shock resemble those in an RBC model might in reality reflect that the actual economy has various nominal frictions, and monetary policy has successfully mitigated those frictions, as for instance recently argued by Altig *et al.* [2003]. In this respect, we obtain two contrasting results. On the one hand, it is clear that, in stark contrast with the impulses responses of the RBC model, the estimated response of real wages is in general lower than that of labor productivity — the probability that the former is lower than the latter exceeds 0.85 on impact. On the other hand, we were unable to find any evidence that the well-documented changes in the systematic conduct of U.S. monetary policy in the last two decades have had any significant effect on the economy's response to technology shocks.

As the exercise was started out by motivating identifying restrictions on impulse responses with a set of model economies, a clear advantage of its approach is the clear link between structural impulse responses and theoretical properties of the models. Therefore, if the (highly nonlinear) mapping from the model's parameter space into impulse responses could be inverted, it would be possible to map the posterior density of impulse responses back into posterior densities of structural parameters, thus providing a precise answer to the above questions. There are, however, several nontrivial aspects of this task, due to the fact that we would be trying to form our inference from a vector-valued function of a vector of parameters, with the dimensionality of both vectors quite high. Hence, an interesting issue for future research would be to compute the likelihood of a vector of impulse responses and estimate the posterior distribution of the parameters of the underlying DSGE model, perhaps suitably adapting methods such as the Sequential Monte Carlo algorithm recently applied to DSGE models by Fernández-Villaverde and Rubio-Ramírez [2004].

Appendix: Description of the data

Labor productivity: index of output per hour, non-farm business sector (Bureau of Labor Statistics, BLS)

Hours worked: index of total hours worked, non-farm business sector (BLS)

Real wage: real hourly compensation, non-farm business sector (BLS)

Consumption: personal consumption expenditures, billions of chained (1996) dollars (Bureau of Economic Analysis, BEA)

Investment: gross private capital formation, billions of chained (1996) dollars (BEA)

Short-term interest rate: Federal funds rate (Federal Reserve Bank of St. Louis)

Inflation: quarterly changes in the implicit GDP deflator (BEA)

Tables and figures

Table 1

Parameters ranges

parameter	low	up	mean
b	0.985	0.995	0.99
σ_c	1.0	10.0	5.50
σ_l	0.0	10.0	5.0
h	0.0	0.8	0.4
χ	0.0	5.0	2.5
ξ_p	0.0	0.8	0.405
γ_p	0.0	1.0	0.5
ξ_w	0.0	0.8	0.405
γ_w	0.0	1.0	0.5
λ_w	0.0	1.0	0.5
ψ	0.0	50.0	25.0
ρ_r	0.0	0.99	0.495
ρ_y	-0.25	0.25	0.0
ρ_π	1.1	2.0	1.55
ρ_z	0.75	1.0	0.85
ρ_i	0.75	1.0	0.85

Table 2

Sign restrictions on VAR variables[a]

Variable	Horizon in quarters
$lp_k \geq 0$	$k = 0, ..., 19$
$w_k \geq 0$	$k = 2, ..., 19$
$i_k \geq 0$	$k = 0, ..., 9$
$y_k \geq 0$	$k = 0, ..., 9$
$c_k \geq 0$	$k = 0, ..., 4$

Table 3

Probability of a positive response of hours worked[a]

horizon	1	3	5	9	11	13	15	17	19
probability	0.73	0.88	0.88	0.85	0.77	0.68	0.60	0.53	0.48

[a]The impact response is denoted as the response at horizon 0.

Figure 1A

Impulse responses to positive technology shock: RBC model

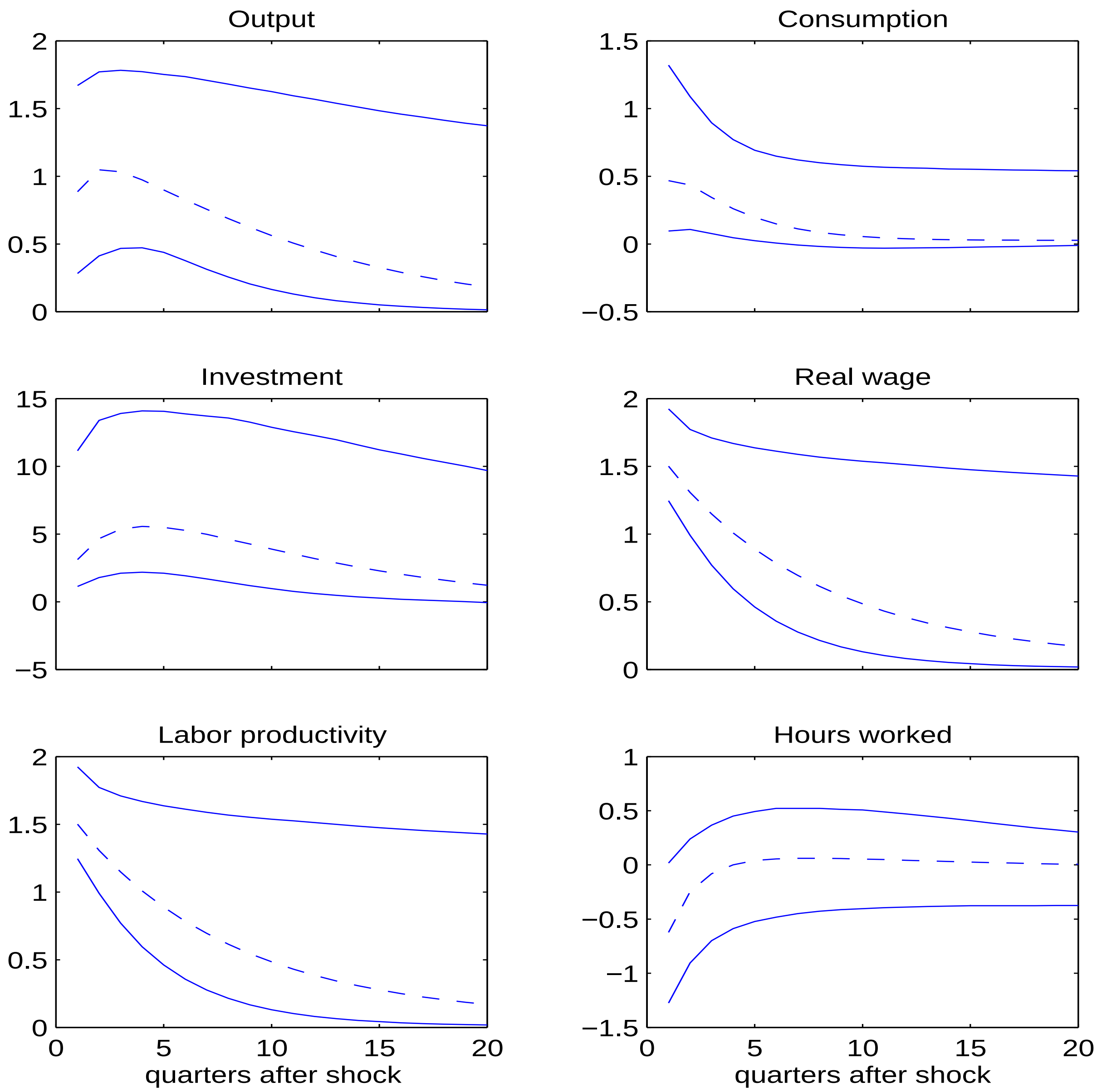


Figure 1B

Impulse responses to positive technology shock: NR model

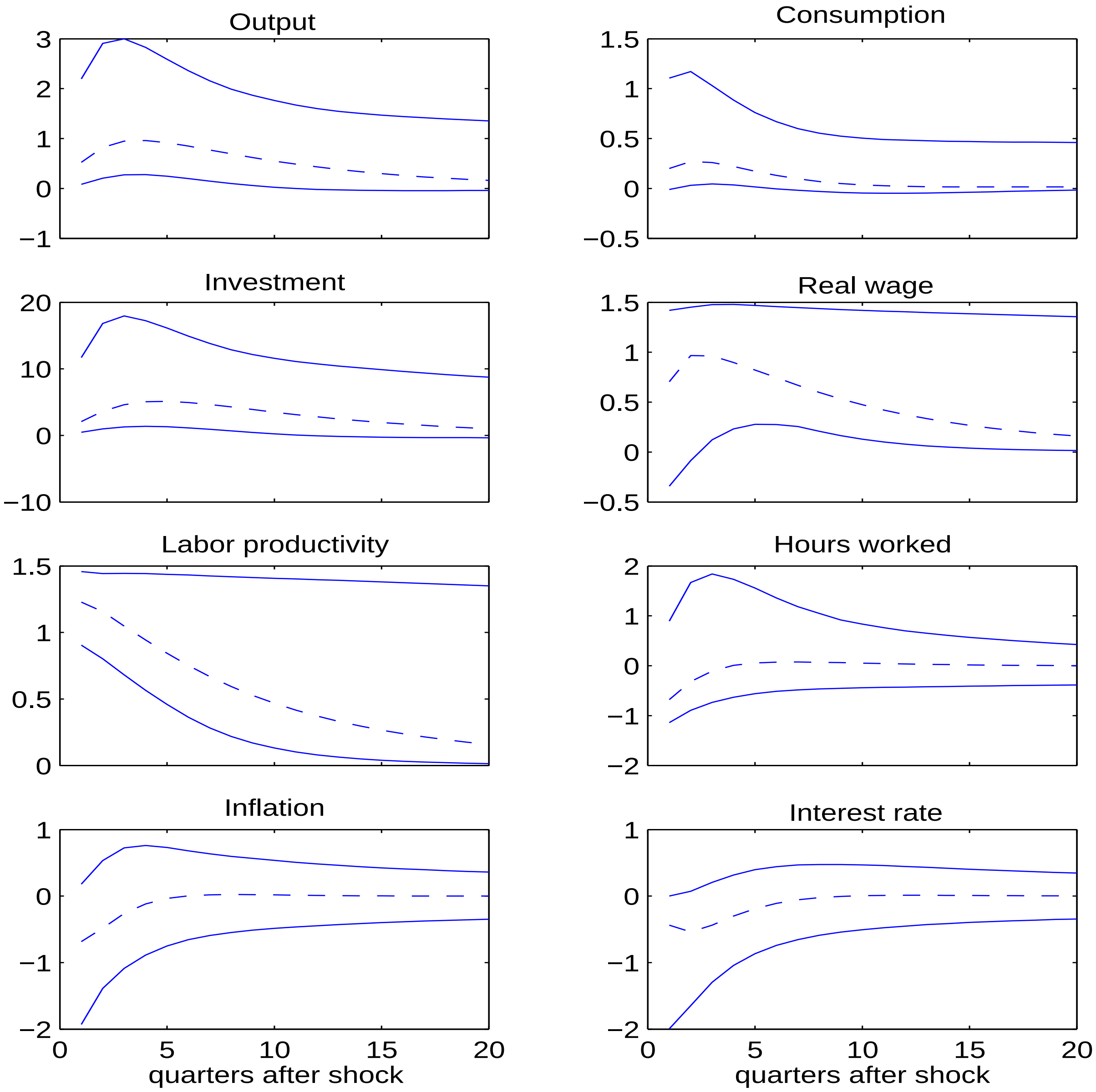


Figure 1C

Impulse responses to a positive investment efficiency shock: RBC model

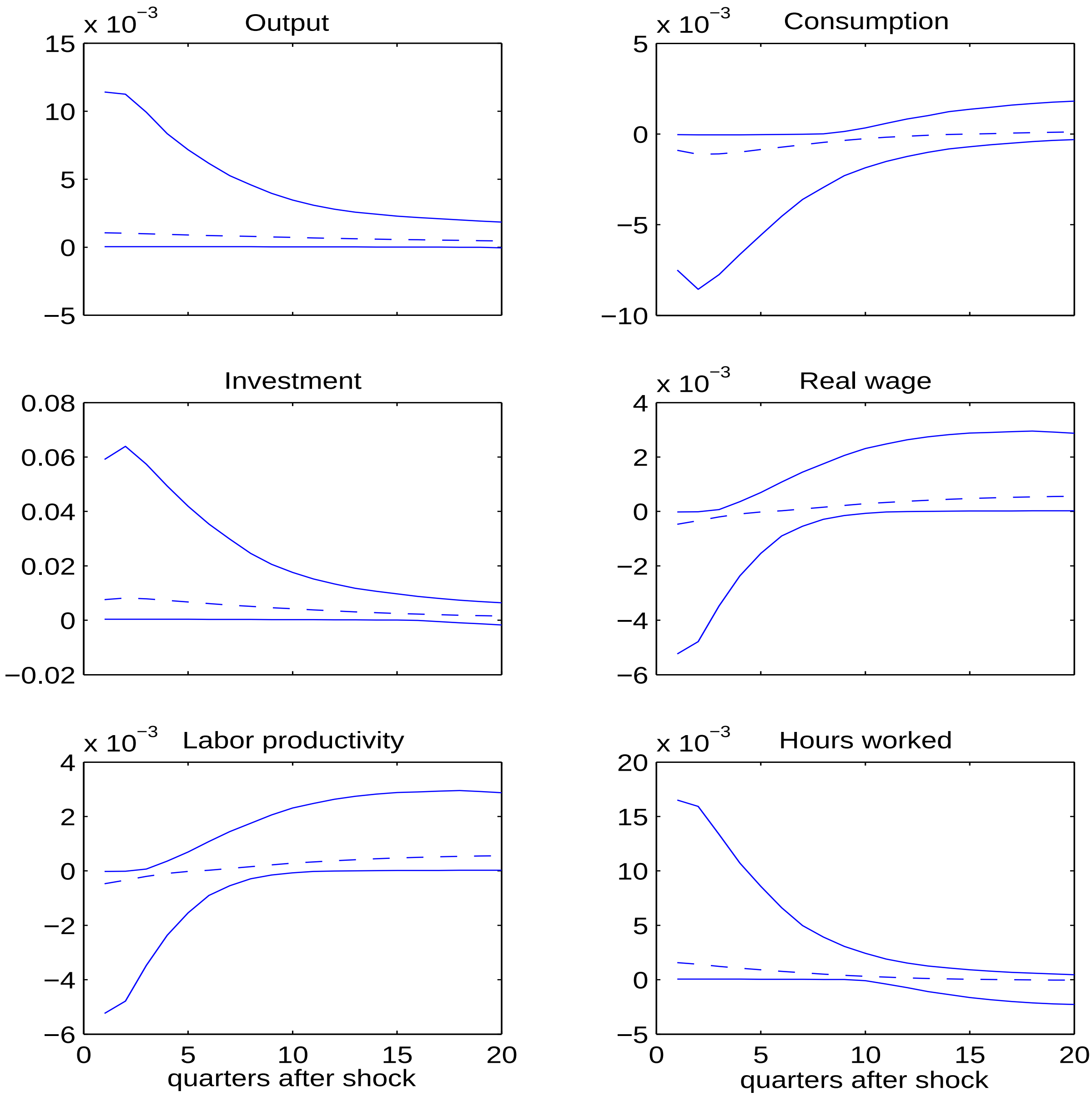


Figure 1D

Impulse responses to a positive investment efficiency shock: NR model

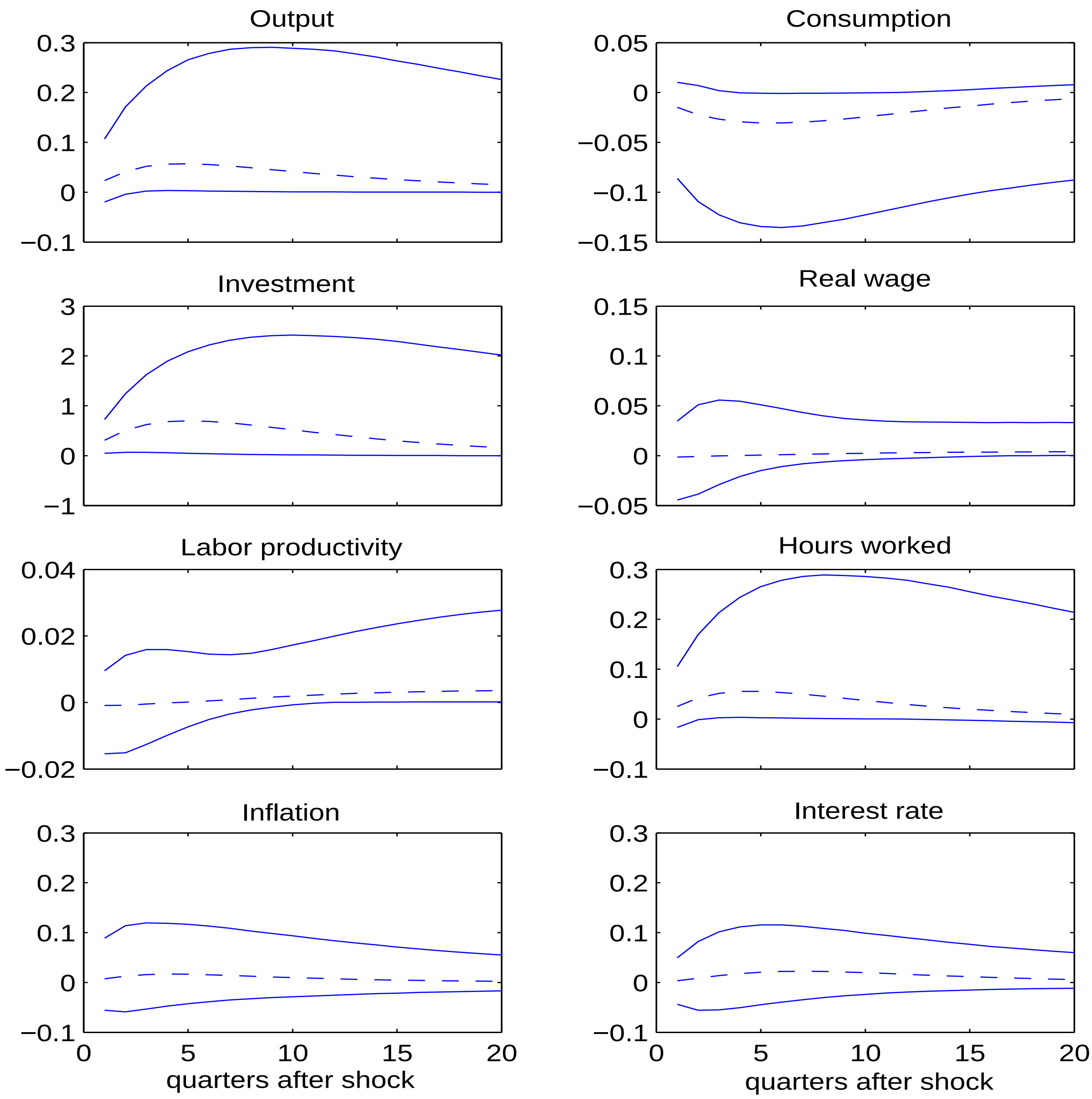
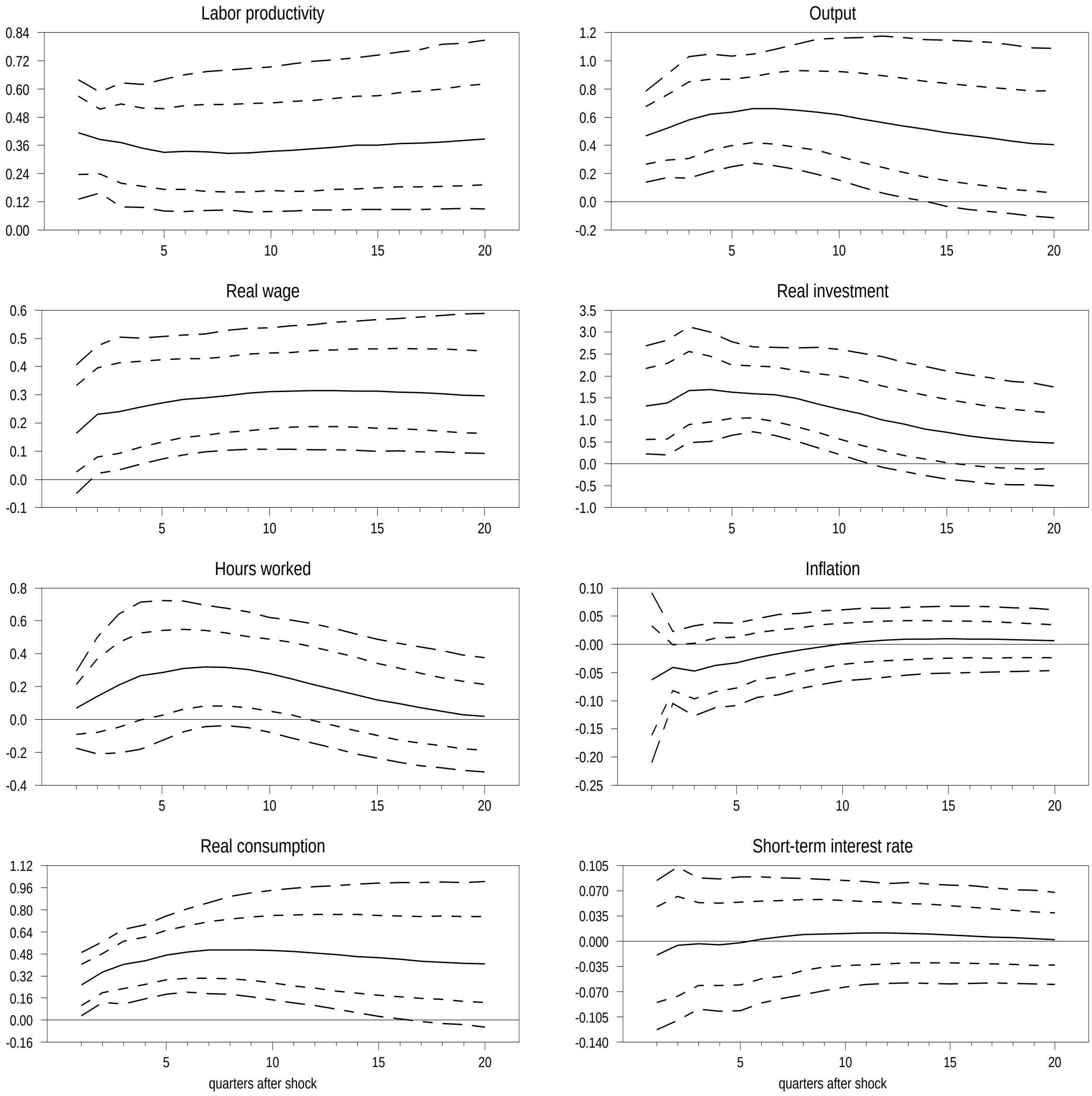


Figure 2

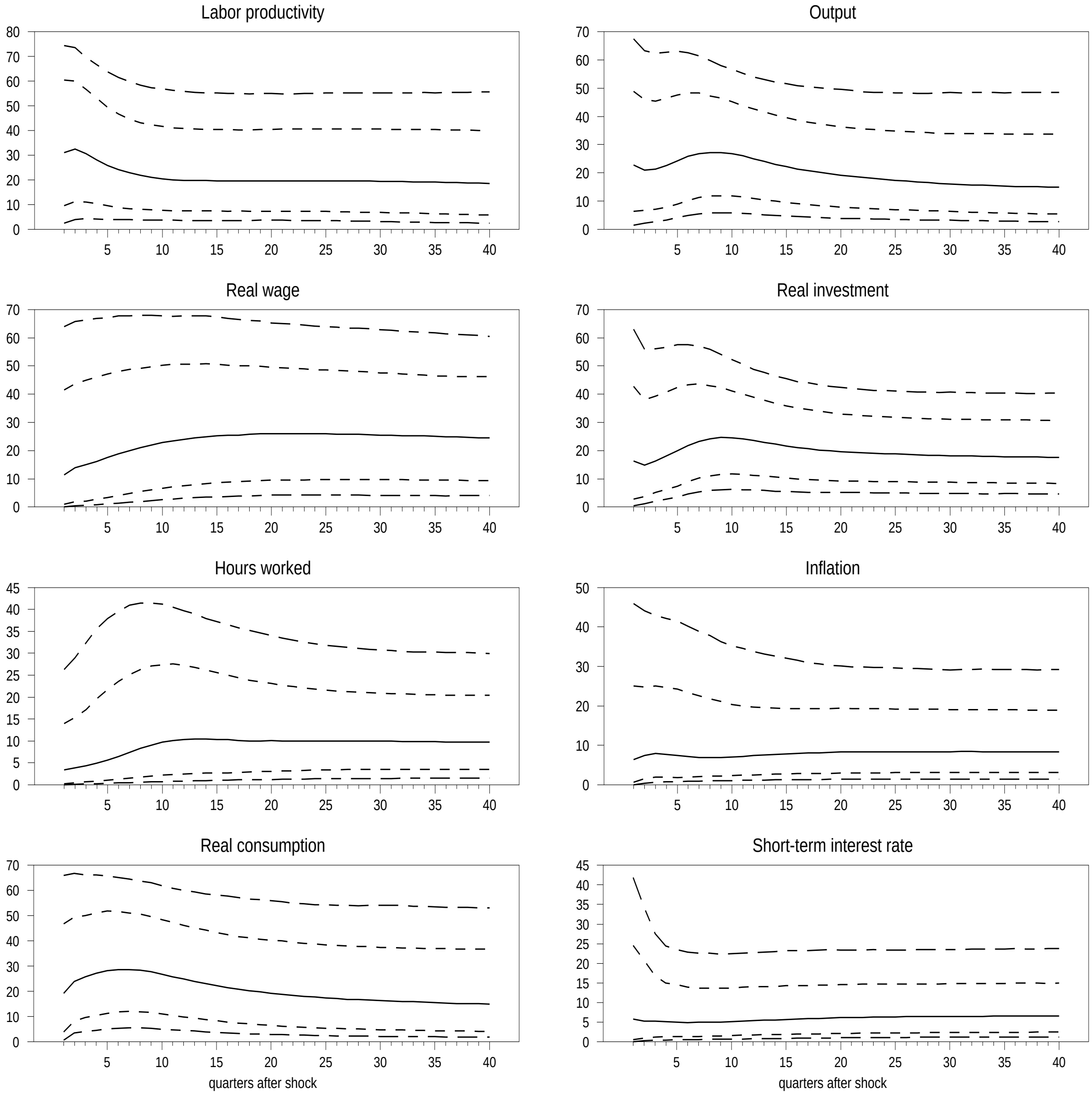
Impulse responses to positive technology shock: Benchmark specification[a]



[a]All variables in levels. Assumed sign restrictions are reported in Table 2. Sample period is 1953:1-2003:4.

Figure 3

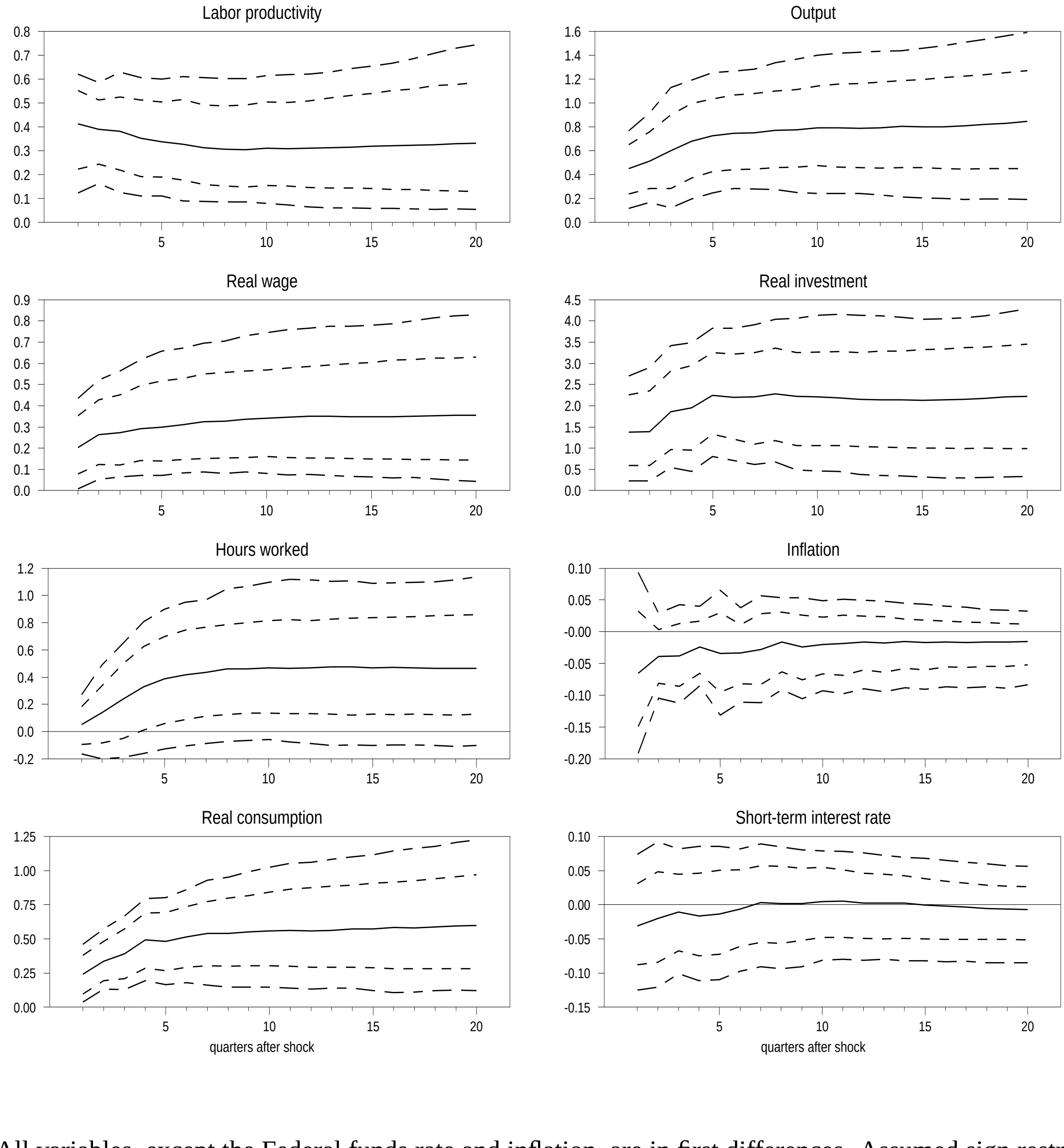
Contribution of technology shocks to the variance of the forecast error: Benchmark specification[a]



[a]All variables in levels. Assumed sign restrictions are reported in Table 2. Sample period is 1953:1-2003:4.

Figure 4

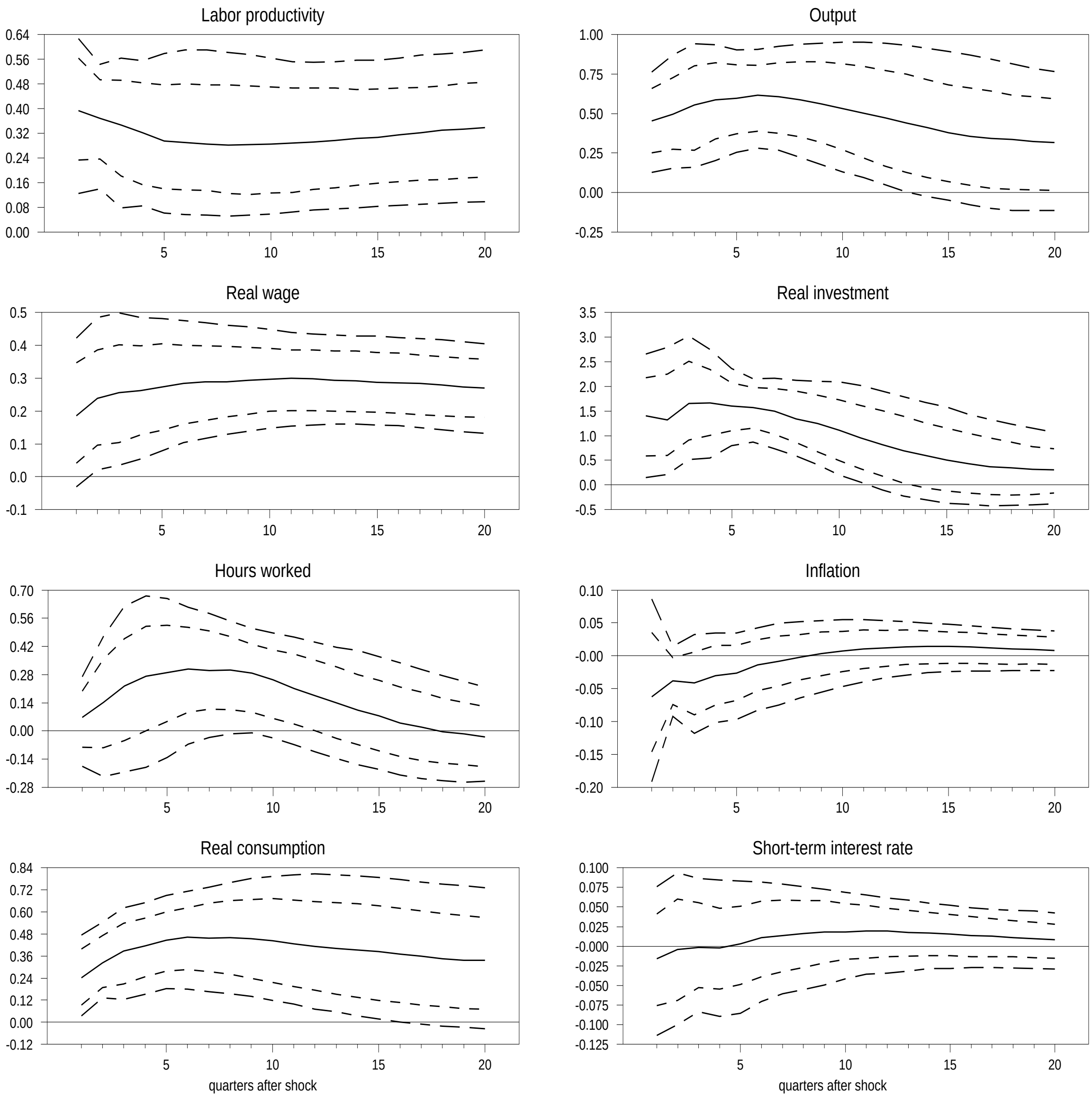
Impulse responses to positive technology shock: difference specification[a]



[a]All variables, except the Federal funds rate and inflation, are in first differences. Assumed sign restrictions are reported in Table 2. Sample period is 1953:1-2003:4.

Figure 5

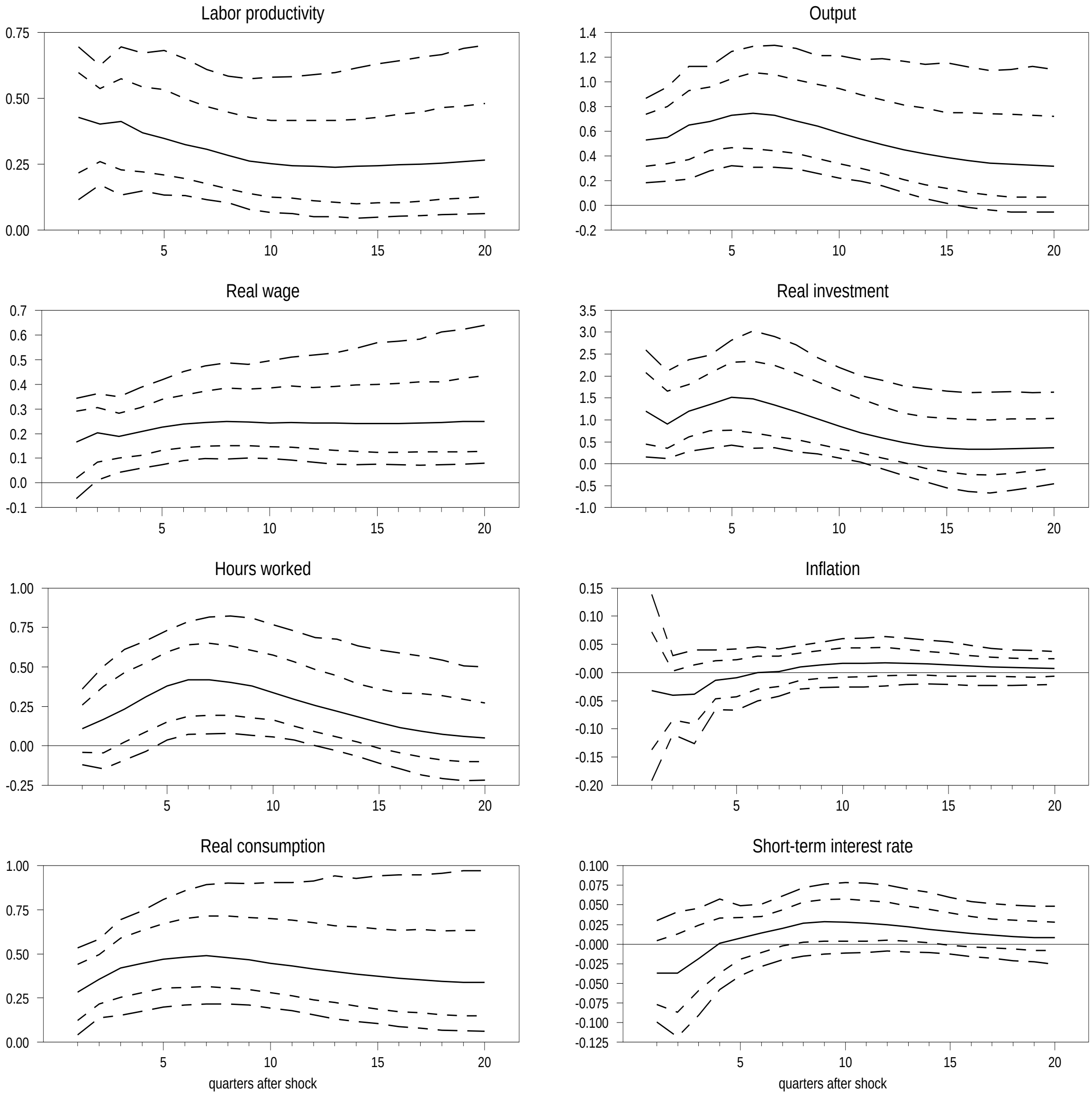
Impulse responses to positive technology shock: The effect of identification uncertainty [a]



[a]All variables in levels. Assumed sign restrictions are reported in Table 2. Sample period is 1953:1-2003:4. The reduced form of the VAR and the covariance matrix are fixed at their OLS-ML estimates.

Figure 6A

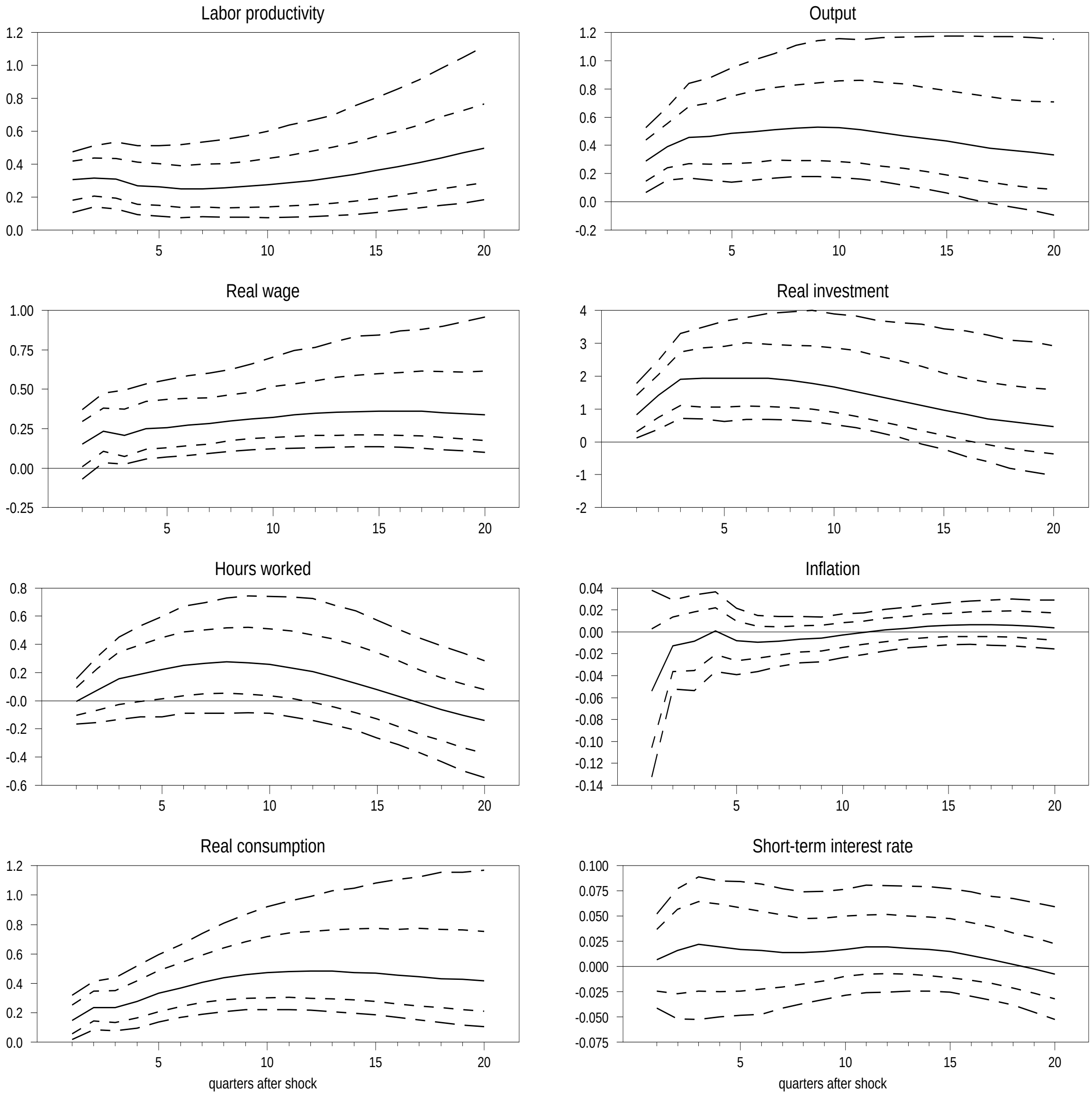
Impulse responses to positive technology shock: Pre-1979:2 period[a]



[a]All variables in levels. Assumed sign restrictions are reported in Table 2. Sample period is 1953:1-1979:2.

Figure 6B

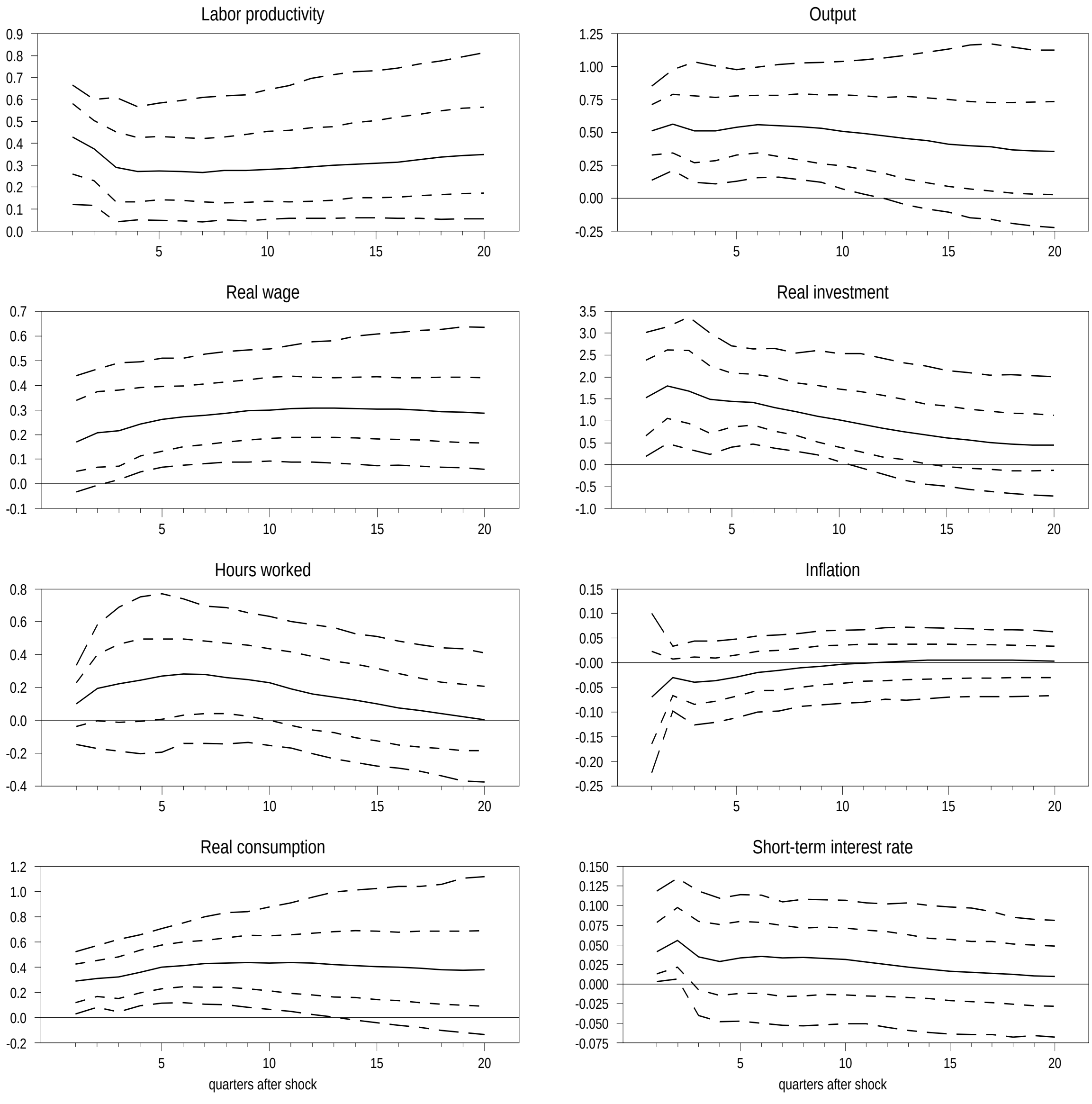
Impulse responses to positive technology shock: Post-1983:1 period[a]



[a]All variables in levels. Assumed sign restrictions are reported in Table 2. Sample period is 1983:1-2003:4.

Figure 7

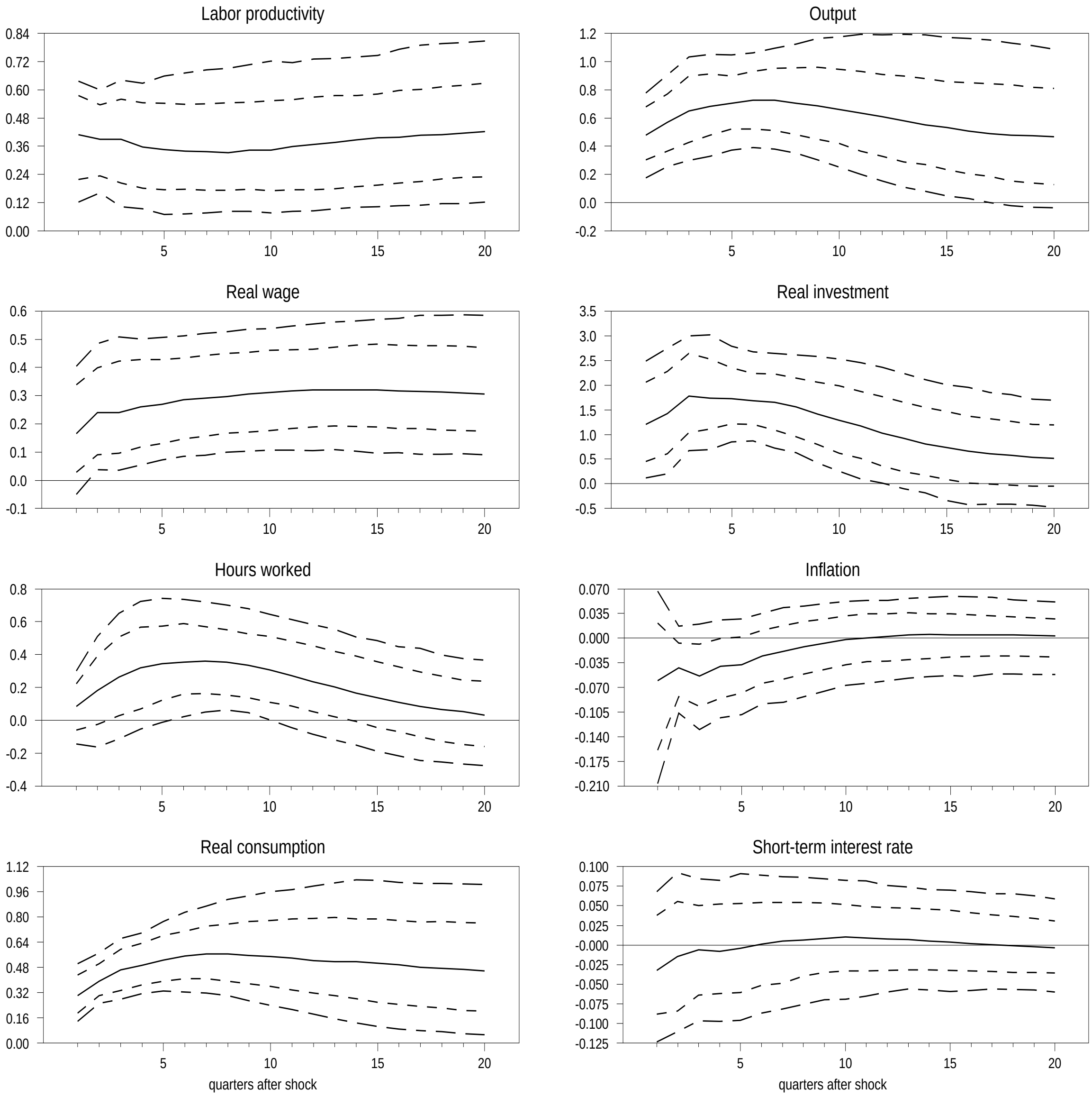
Impulse responses to positive technology shock: The effect of assuming a positive interest-rate response[a]



[a]All variables in levels. Sample period is 19531:1-2003:4. Assumed sign restrictions are reported in Table 2 with the additional requirement of a positive response of the Federal funds rate in the first 2 quarters.

Figure 8

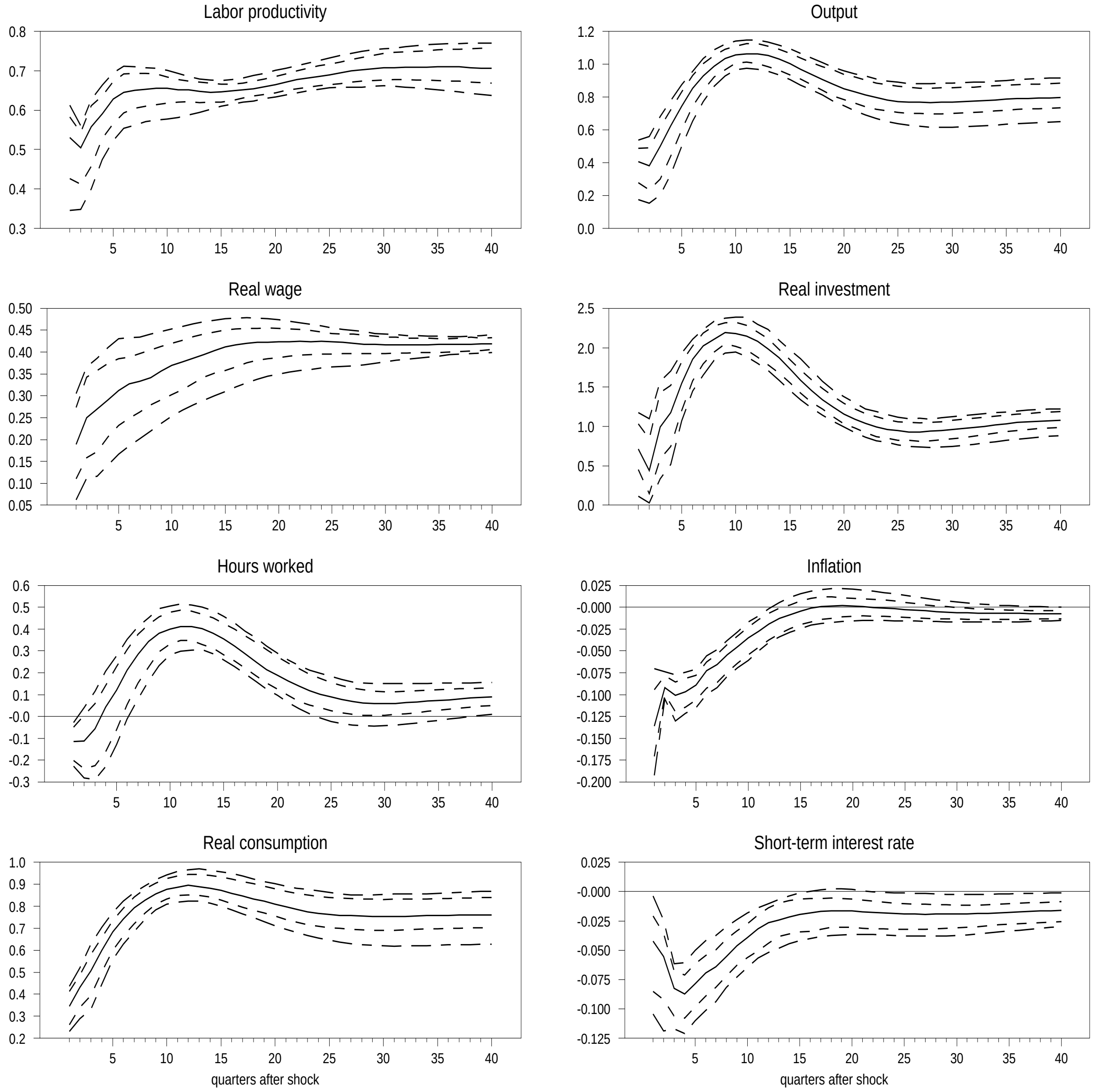
Impulse responses to positive technology shock: The effect of assuming a large response of consumption[a]



[a] All variables in levels. Sample period is 1953:1-2003:4. Assumed sign restrictions are reported in Table 2 with the additional requirement that for the first five quarters the response of real consumption is larger than the 16 percentile under the benchmark specification (see Figure 2).

Figure 9

Impulse responses to positive technology shock: The effect of requiring a large contribution to labor productivity long-run changes[a]



[a] All variables in levels. Sample period is 1953:1-2003:4. Assumed sign restrictions are reported in Table 2 with the additional requirement that technology shocks account for at least 70 percent of the variance of the forecast error of labor productivity at 40 quarters. The reduced form of the VAR and the covariance matrix of the residuals are fixed at their OLS-ML estimates.

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