Has Monetary Policy Become Less Powerful?*

Jean Boivin†

Columbia University

Marc Giannoni‡

Federal Reserve Bank of New York

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†Columbia Business School, 719 Uris Hall, 3022 Broadway, New York, NY 10027; e-mail: jb903@columbia.edu; www.columbia.edu/~jb903.

‡Federal Reserve Bank of New York, Domestic Research, 33 Liberty Street, New York, NY 10045; e-mail: marc.giannoni@ny.frb.org; www.newyorkfed.org/rmaghome/economist/giannoni/giannoni.html.

Abstract

Recent vector autoregression (VAR) studies have shown that monetary policy shocks have had a reduced effect on the economy since the beginning of the 1980s. This paper investigates the causes of this change. First, we estimate an identified VAR over the pre- and post-1980 periods, and corroborate the existing results suggesting a stronger systematic response of monetary policy to the economy in the later period. Second, we present and estimate a fully specified model that replicates well the dynamic response of output, inflation, and the federal funds rate to monetary policy shocks in both periods. Using the estimated structural model, we perform counterfactual experiments to quantify the relative importance of changes in monetary policy and changes in the private sector in explaining the reduced effect of monetary policy shocks. The main finding is that changes in the systematic elements of monetary policy are consistent with a more stabilizing monetary policy in the post-1980 period and largely account for the reduced effect of unexpected exogenous interest rate shocks. Consequently, there is little evidence that monetary policy has become less powerful.

1 Introduction

A growing body of evidence, both anecdotal and from formal statistical investigations, suggests that the economy has changed in substantial and fundamental ways over the last decades. Perhaps the most striking characteristic of the post-1980 U.S. experience in terms of real activity and inflation is the important decline in their volatility. Various phenomena might be at the root of these changes. The conduct of monetary policy has recently received a fair amount of attention as a potential source,[1] but various innovations in firms and consumers behavior, induced by technological progress and financial innovations for instance, are also likely to have occurred.[2]

This raises important questions concerning the role of monetary policy in this new environment. In particular, do these changes imply a different effect — perhaps smaller — of monetary policy on the economy? Evidence already exists that points to a change in the impact of monetary policy.[3] Recent studies using monetary vector autoregressions (VAR) have demonstrated that the impact of monetary policy "shocks" — defined as unexpected exogenous changes in the Federal funds rate — have had a smaller impact on output and inflation since the beginning of the 1980's.[4] This is illustrated in Figure 1, which shows the response of a measure of detrended output and inflation to the same size monetary shock, separately for the pre- and post-1980 periods.[5] In so far as these

[1] See Clarida, Galí and Gertler (2000), Boivin (2001), and Cogley and Sargent (2001) among others.

[2] McConnell and Perez-Quiros (2000) and Kahn, McConnell and Perez-Quiros (2001) argue that progress in inventory management could explain the lower volatility of GDP after 1984.

[3] A special issue of the Federal Reserve Bank of New York's *Economic Policy Review* is dedicated to this question, following a Conference on *Financial innovation and monetary transmission.* One broad conclusion from the papers included in the volume — which approach the question from a variety of different angles — is that monetary policy's effects appear somewhat weaker recently than in previous decades (see e.g., Kuttner and Mosser (2002)).

[4] See the NBER working paper version (no. 5145, June 1995) of Bernanke and Mihov (1998), Gertler and Lown (2000), Barth and Ramey (2001) and Boivin and Giannoni (2002) among others.

[5] The exact definitions of these variables and how the responses were computed is described in Section 2 below.

impulse response functions trace out the effect of monetary policy, a casual look at this evidence might suggest that monetary policy has become less powerful. However, this can only be determined once we understand the causes of these changes.

In fact, the effect of monetary policy depends both on the way it is conducted, and on the response of households and firms — the private sector for short — to variations in the monetary policy instrument. One potential reason for the observed change in the responses to monetary policy shocks is that the private sector's response has changed; perhaps financial innovations, or other structural changes have allowed consumers and firms to better cushion themselves from the impact of interest rate fluctuations. In such a case, the reduced response to monetary shocks is likely to reflect less powerful monetary policy. Another possibility however is that monetary policy itself has come to respond more decisively to economic conditions, thereby moderating the real effects of demand fluctuations. In this case, the change in the responses to monetary shocks does not result from a less powerful monetary policy. These are not the only possible explanations, of course, but these examples illustrate that determining the causes of the shocks' smaller impact is crucial for understanding the consequences of this finding.

The goal of this paper is to investigate the driving forces behind the reduced effect of unexpected exogenous interest-rate shocks on the economy, in order to determine if monetary policy has become less powerful.[6] We are not interested in the policy shocks — which are known to be small — *per se,* but rather in the mechanism that propagates them throughout the economy. These shocks allow us to identify this propagation mechanism, which includes the systematic part of monetary policy — i.e., the policy reaction function. We view the reaction function as the most important component

[6] It is important to note that, although related, our goal differs significantly from the one of other papers — such as Khan, McConnell and Perez-Quiros (2001) — which attempt to explain the reduced volatility of macroeconomic variables.

of monetary policy. We thus seek to determine to what extent changes in the reaction function affect the economy.

We follow a two-step strategy to uncover the causes behind the smaller effect of exogenous changes in the interest rate. First, using a VAR estimated over the 1963:1-1979:3 and 1980:1-1997:4 periods, we identify a forward-looking policy reaction function and investigate the extent to which its estimated changes can account for the differences in the impulse response functions across sub-samples. We do this through a counterfactual experiment where we combine the policy rule estimated over the second sub-sample with the other equations of the VAR estimated over the first sample, and vice-versa. From this fairly unstructured approach, we find that while the changes in monetary policy are important to explain the reduced responses of output and inflation to policy shocks, changes in the other parameters of the VAR also play a role.

It is likely, however, that this approach does not allow to properly interpret the source of changes. In fact, it does not take into account the modifications in the agents' behavior resulting from changing policy. As a result, it would only be valid if firms and consumers were entirely backward looking, or if the forward-looking agents did not account for the change in policy while forming their expectations. This is more likely to happen if the policy changes are very small and of temporary nature. However, when the changes are important and more persistent — such as the 1979 regime shift may have been — the Lucas (1976) critique becomes more relevant.

This observation motivates our second and main strategy, which is to use a general equilibrium macroeconomic model to interpret the changes in the VAR impulse response functions. We consider a model similar to that of Rotemberg and Woodford (1997), but that allows for additional frictions such as habit formation and some degree of backward-looking behavior. Since the ultimate goal is to use the structural model to interpret the evolution of the impulse response functions, a natural approach is to estimate the model by minimizing the distance between the theoretical and empirical (i.e., VAR-based) impulse response functions. Although akin to a calibration exercise, this is a well-defined estimation problem and thus statistical inference on the structural parameters can be performed. Using the estimated structural model we can interpret the causes of the observed changes in the estimated VAR. In particular, we can isolate the part that stems from changes in the systematic elements of monetary policy. An important by-product of our investigation is to provide a set of structural parameter estimates for the New Keynesian model that we consider, and for different sub-samples.

The main finding of this paper is that changes in the policy reaction function largely account for the reduced effect of exogenous interest rate shocks. The results also suggest that by responding more strongly to changes in economic conditions, the recent conduct of monetary policy is more robust to potential changes in the private sector behavior. Furthermore, an interpretation of the discrepancy between the VAR-based and model-based counterfactual analyses is that changes in systematic policy are taken into account by the private sector when forming expectations; this suggests that the systematic component of monetary policy affects the economy, and also that the Lucas (1976) critique is relevant in the present context.

The rest of the paper is organized as follows. Section 2 describes the identification of the policy reaction function in a VAR, and the results from stability tests and split-sample estimation of the empirical model. Section 3 investigates the source of the observed changes through a counterfactual analysis based on the identified VAR. Section 4 describes and estimates a fully-specified general equilibrium model of the U.S. economy. Section 5 uses this model to interpret the nature of the changes in the monetary transmission mechanism. Section 6 concludes.

2 Empirical identification of the monetary transmission mechanism

2.1 Identification of the policy reaction function

2.1.1 VAR and forward-looking policy reaction function

Our empirical model of the economy is a VAR in variables describing the economy (Z_t) as well as monetary policy (R_t)

$$\begin{pmatrix} Z_t \\ R_t \end{pmatrix} = a + A(L) \begin{pmatrix} Z_{t-1} \\ R_{t-1} \end{pmatrix} + u_t.$$

Three variables are included in the non-policy block Z_t: detrended output ($\hat{Y}_t$) and the inflation rate (π_t), as suggested by the theoretical model developed in Section 4, as well as a commodity price measure.[7] The commodity price inflation (πc_t), although not formally justified by the theoretical model, is added to limit the extent of a "price-puzzle" in this VAR.[8] The policy instrument, R_t, is assumed to be the Federal funds rate. While the Fed's operating procedure has varied in the last four decades, many authors have argued that the Federal funds rate has been the key policy

[7] All series are taken from the Standard and Poor's –DRI database. Detrended output, which is measured as the percent deviation of quarterly real GDP (mnemonic GDPQ) from a stochastic trend, is obtained from a high-pass filter that isolates frequencies associated to periods less than 32 quarters. The results are robust to the use of a linear or quadratic deterministic trend. $\hat{Y}_t$ is often referred to as the "output gap" in the literature. The inflation rate is the annualized rate of change in the GDP deflator (mnemonic GDPD) between two consecutive quarters. The commodity price measure is the quarterly average of the monthly spot market commodity price index (mnemonic PSCCOM). The original data set runs from 1959:1 to 2000:4. Because of the data transformation (first-differences and high-pass filter), the analysis is performed on the 1963:1–1997:4 period, except for the estimation of the forecasting horizon (see below) where we use the 1963:1–1995:5 period to accomodate the greatest horizon considered. Four lags are included in the VAR.

[8] This practice is fairly standard in this literature.

instrument in the U.S. over most of that period (see e.g., Bernanke and Blinder (1992), Bernanke and Mihov (1998)).[9]

Results from VAR models are known to be quite sensitive to their specification. Our simple but standard specification has the virtue of containing the minimum set of variables necessary for our investigation, and yet delivering sensible impulse response functions, broadly consistent with existing results in the literature. Importantly, the key empirical feature that we are trying to explain, namely the reduced effect of monetary shocks on output and inflation, is corroborated by quite different specification and identifying assumptions. For instance, Bernanke and Mihov (1998) report a similar reduction in the effect of a policy shock using a much more sophisticated model of the Fed's operating procedure.[10] Barth and Ramey (2001) reach similar conclusions using instead long-run restrictions.

In order to identify the policy reaction function from this VAR, we assume that the economy (Z_t) responds only with a lag to changes in the Fed funds rate. Although debatable, this identifying assumption is consistent with many recent VAR analyses,[11] including Rotemberg and Woodford (1997) which serves as a benchmark to our structural investigation in Section 4. Under this recursive structure, the identified VAR can be expressed as:

$$Z_t = b + \sum_{i=1}^{P} B_i^Z Z_{t-i} + \sum_{i=1}^{P} B_i^R R_{t-i} + u_t^Z \tag{1}$$

$$R_t = \phi^0 + \sum_{i=0}^{P} C_i^Z Z_{t-i} + \sum_{i=1}^{P} C_i^R R_{t-i} + u_t^R. \tag{2}$$

[9] The Federal funds rate provides probably a less adequate measure of monetary policy stance for the period running from 1979 to 1982, as non-borrowed reserves were set to achieve a level of interest rates consistent with money growth targets, but Cook (1989) argues that the Fed funds rate may still provide a satisfactory indicator during this episode.

[10] See NBER working paper version (no. 5145, June 1995) of Bernanke and Mihov (1998).

[11] See for instance Bernanke and Blinder (1992), and Bernanke and Mihov (1998).

Equation (2) constitutes an unrestricted specification of the policy reaction function, which, under this recursiveness assumption, can be estimated directly by OLS.

The policy reaction function so identified can be seen as a reduced-form expression of a forward-looking policy rule according to which the Fed responds to expected inflation and detrended output at given horizons h_π and h_y respectively. The coefficients of this reduced-form equation subsume policy parameters — i.e., the parameters characterizing the Fed's systematic behavior — as well as the remaining parameters needed to form the expectations of π_{t+h_π} and $\hat{Y}_{t+h_y}$ conditional on the time-t information set. To uncover the policy response parameters, we need to impose more structure on the policy rule. We assume, in particular, that it takes the form:

$$R_t = \phi^0 + \phi^\pi \pi^e_{t+h_\pi|t} + \phi^y \hat{Y}^e_{t+h_y|t} + \sum_{i=1}^{P} \rho_i R_{t-i} + \varepsilon_t, \tag{3}$$

where $\pi^e_{t+h_\pi|t}$ and $\hat{Y}^e_{t+h_y|t}$ represent the projection of π_{t+h_π} and $\hat{Y}_{t+h_y}$ on the time-t information set, and the unforecastable random variable ε_t represents the monetary policy shock. For $h_\pi = 0$ and $h_y = 0$, equation (3) corresponds to the popular rule proposed by Taylor (1993), augmented by the lags of the Fed funds rate.[12] Another special case is when $h_\pi = 1$ and $h_y = 1$, which corresponds to the forward-looking rule estimated by Clarida, Galí and Gertler (2000). These authors estimate such a rule by GMM, in the single equation framework, assuming rational expectations on the part of the central bank.

Efficient estimation of the identified VAR with the unrestricted policy reaction function can be performed simply by estimating each equation of the system (1) and (2) by OLS. However, once the structure of the forward-looking rule (3) is explicitly imposed, the equation-by-equation estimation of the resulting system does not provide an efficient estimate. Instead, all the moments of the

[12] The Fed funds rate is known to display a lot of persistence. This phenomenon might arise from a Fed's interest-rate smoothing concern or could reflect optimal policy under commitment (cf. Woodford (1999b)).

system (1) and (3) should be exploited simultaneously. In what follows, however, we will ignore this efficiency issue. More specifically, we estimate each equation of the system (1) separately by OLS, and we estimate (3) by GMM, as in Clarida *et al.* (2000), using the valid instruments implied by the multivariate system.

2.1.2 The forecasting horizon

The estimation of equation (3) requires the specification of the horizons h_π and h_y. Such horizons are usually specified on *a priori* grounds, based on what is thought to be reasonable lags for the effect of monetary policy on the economy. But the horizon that a central bank should be considering is not clear in theory. While forward-looking rules can be motivated from the existence of lags in the effect of monetary policy, there is also a case to be made for backward-looking rules, which might provide more stability. Given the absence of a clear criterion to select the horizon a priori and, importantly, given the sensitivity of the results to this choice, it appears desirable to infer the horizon from the data.

Intuitively, one could infer the horizon of the policy rule by determining which one is the most consistent with the unrestricted version of the policy reaction function. In fact, estimation of equation (2) consistently identifies the systematic part of the forward-looking rule, independently of the true underlying horizon. Different specifications of the horizon imply different sets of restrictions on the parameters C_i^Z, and an estimate of the horizon can be obtained by determining which of these sets of restrictions is best supported by the data.

More formally, as the forward-looking rule (3) is just an over-identified version of equation (2), one can select the horizon that minimizes the distance of the over-identified model from the unrestricted model. A measure of this distance is provided by the Hansen J-test. We thus select the horizon that minimizes this test statistic. As a by-product, this statistic provides a measure of the accuracy of the specification — other than the horizon — embedded in equation (3). Table 1 reports the test statistics and p-values of the J-test for all combinations of h_π and h_y up to 8 quarters, and for various samples.

The best horizon, i.e., the one with the highest $p-$value, is $h_\pi = 2$ and $h_y = 1$, for the full sample.[13] For the 1980:1-1995:4 and 1984:1-1995:4 periods we obtain $h_\pi = 1$, $h_y = 1$ and $h_\pi = 3$, $h_y = 1$, respectively. For the 1963:1-1979:3 sample, the horizon is very imprecisely estimated.[14] In fact, almost all horizon combinations cannot be rejected by the data. Moreover, if we were to take these numbers literally, we would obtain a surprisingly long horizon for $\hat{Y}^e_{t+h_y|t}$: in fact the p-value of the J-test is maximized at $h_\pi = 1$ and $h_y = 7$. This is at odds with conventional wisdom, and in fact leads to an estimate of the policy reaction function that has a negative coefficient on the output gap. Based on this, and given the imprecision of the horizon's estimation, we proceed with the horizon estimated for the full sample, i.e. $h_\pi = 2$ and $h_y = 1$, which is not rejected by the data in Sample 1.[15]

2.2 Documenting changes in the effect of monetary policy

2.2.1 Stability tests on the reduced form VAR

The stability of macroeconomic relationships has been investigated in a number of recent papers. The most general evidence is provided by Stock and Watson (1996) who find widespread instability

[13] Since we are considering up to eight-quarter ahead forecasting horizons, we perform the forecasting horizon estimation based on data up to 1995:4.

[14] Note that all combinaison of h_y and h_π with a p-value greater than 0.05 in Table 1 are part of the 95% confidence set.

[15] Note that the impulse response functions obtained for Sample 1 using this horizon are very similar to the ones obtained using the unrestricted equation (2). This choice of horizon has thus no significant impact in the results reported below.

in the bivariate relationships among 76 macroeconomic variables. In the VAR context, mixed results have been obtained.[16] Boivin (1999) argues that the difference are due mainly to the small sample properties of the stability tests, and to the effect of the number of parameters tested on the power of these tests. He concludes that there is compelling evidence of instability in monetary VARs.

We now perform a similar stability investigation on the VAR described in the previous section. For each equation of the reduced form VAR, we test jointly for the stability of all the coefficients on the lags of a given variable, using the Wald version of the Quandt (1960) likelihood-ratio test (i.e., Andrews (1993) sup-Wald test). We use an heteroskedasticity-robust version of the test. This test, unlike the well-known Chow test, does not assume knowledge of the date at which the break in the parameters occurs. This test is also known to have power against other alternatives, such as one in which the coefficients follow a random walk (see Stock and Watson (1998)).

The p-values of the stability tests are presented in Table 2. Overall the results suggest that instability is important in this VAR. Of the 16 tests performed, 50% reject the null of stability at the 5% level. We thus interpret these results as strong evidence of changes in the propagation mechanism. It is important to note that these changes are economically significant. In fact, in Boivin and Giannoni (2002) we showed using a similar VAR – which left the policy reaction function unrestricted – that the observed reduction in the volatility of inflation and output was explained roughly equally by a reduction in the variance of the shocks and a smaller propagation.

Given the identified policy rule, it is also possible to test directly for its stability. The p-value of the test applied jointly to ϕ^0, ϕ^π and ϕ^y is 0.000. We can thus conclude that at least part of the instability observed in the reduced form VAR arises from changes in the conduct of monetary

[16] Bernanke, Gertler and Watson (1997) find evidence of instability in a monetary VAR, while Bernanke and Mihov (1998) and Christiano, Eichenbaum and Evans (1999) conclude the opposite.

policy *per se*.

2.2.2 Split-sample estimates of the impulse response functions

Given this evidence of changes in the economy, we now turn to the implications of these changes for the effect of monetary policy. As argued in the introduction, we assess the changes in the effects of monetary policy by comparing impulse response functions of the output gap, inflation, and the Fed funds rate to a monetary policy shock, using the VAR estimated over different sub-samples. Based on anecdotal evidence regarding the conduct of monetary policy, and on previous empirical studies, while making sure that the samples are not too small, we decided to base our benchmark comparison on the sub-samples on each side of 1979:4, the date at which Fed chairman Paul Volcker announced a shift in policy.[17] An alternative would be to start the second sample in 1984:1, to eliminate the alleged non-borrowed reserves targeting experiment from the second sample. As a result, we compare three sub-samples: Sample 1 corresponds to 1963:1-1979:3, Sample 2 corresponds to 1980:1–1997:4 and Sample 3 to 1984:1–1997:4.[18]

Table 3 reports the estimates of the policy rule for the different samples. Similarly to the results obtained by Clarida *et al.* (2000), we find that the long-run response of policy to inflation — i.e., $\phi^{\pi}/(1-\rho_1-\rho_2-\rho_3-\rho_4)$ — is smaller than 1 and insignificant in Sample 1, and much larger and significant in the post-80 period, i.e. for both Sample 2 and 3.[19] The policy response to output on

[17] We do not include 1979:4 in the second sample to be consistent with the one used by Rotemberg and Woodford (1997). See also Bernanke and Mihov (1998), and Clarida *et al.* (2000) among others.

[18] In principle, one could estimate the break date as a by-product of the Quandt likelihood-ratio tests of the previous sub-section. Not surprisingly, however, our VAR's estimated break dates — for each combination of a dependent variable and lags of a regressor — do not provide of a consistent picture of the timing of the observed instability.

[19] Clarida *et al.* (2000) emphasize that a response to inflation smaller than one generates indeterminacy or the possibility of persistent self-fulfilling equilibria. While this response needs to be large enough to avoid such a situation, it can be smaller than one in the model we describe in Section 4.

the other hand is always significant, and is smallest for Sample 3.

Figure 1 displays — for all three samples — the impulse response functions to an unexpected unit increase in the Fed funds rate, and the associated 95-percent confidence interval from the unrestricted VAR.[20] The key result from this comparison is that the response of detrended output and inflation is much less pronounced and persistent since the beginning of the 1980's than in the pre-1979 period; the trough of the response of output is at least two and a half times larger in Sample 1 than in Sample 2 or 3. This result, which has already been documented in the literature, suggests that the effect of monetary policy *shocks* was stronger before the 1980's.

While this last conclusion is robust to the use of Sample 2 or 3 in the comparison, there are still notable differences between these two samples. In particular, the response of inflation appears somewhat stronger when the VAR is estimated on Sample 3, and the response of output, while overall of similar shape, is positive for most of the periods in the first two years following a positive innovation to the Fed funds rate. We feel that this latter feature of the Sample 3 impulse response functions is problematic. In fact, it implies that over the first two years, a tightening of monetary policy results mainly in an expansionary effect on the economy, which is inconsistent with the implications of any standard macroeconomic model. Since this positive response of output is likely to be due simply to the imprecision of the estimation — the confidence interval are indeed quite large — we leave this issue for future investigation, and focus in the rest of the paper on the Sample 1 – Sample 2 comparison.

Given the imprecision of the estimated impulse response functions, it is difficult to assess directly from the confidence intervals reported in Figure 1 whether the changes in impulse response functions are significant or not. However, we have provided statistical evidence of changes in the parameters of the VAR, and we have shown that these changes imply point estimates of the impulse

[20]The 95% confidence intervals were obtained using Kilian's (1998) bootstrap procedure.

response functions that are quite different. Moreover, the structural analysis that we perform below establishes that the changes in the impulse response functions are driven almost entirely by changes in the policy reaction function, no matter whether the other structural parameters have changed or not. Since the changes in the estimated policy reaction function are found to be statistically significant, we can thus conclude that the point estimates of the impulse response function in the two samples are statistically significant. Taking these results together with the existing evidence[21], there is strong evidence of changes in the effect of monetary policy shocks on output and inflation.

3 VAR-based counterfactual analysis

The previous section has established that the economy's response to interest-rate fluctuations has changed substantially over time, but the evidence does not identify the reasons why. In fact, while the stability tests on the policy reaction function suggests that monetary policy is one potential source of this varying response of the economy, the evidence from Table 2 is also consistent with the presence of changes in the private sector's response.

In this subsection, we investigate the source of change in the effect of monetary policy by performing a counterfactual analysis on the reduced form VAR. In particular, we use two counterfactual experiments to answer the following questions: (1) Are the observed changes in the policy rule sufficient, by themselves, to explain the evolution of the impulse response functions? (2) Alternatively, assuming that monetary policy did not change, can we reproduce the evolution of the impulse response functions through the observed changes in the private sector's response?

To be more precise, let Φ_s be the set of estimates of the parameters of the policy rule (3) for

[21] As noted in the introduction, see Barth and Ramey (2001), Gertler and Lown (2000), Boivin and Giannoni (2002), and other papers collected in the special issue of *Economic Policy Review* (2002) on *Financial innovation and monetary transmission.*

Sample s. Similarly, let Ω_s be the set of estimates of the other VAR parameters, i.e., the parameters of (1). A combination (Φ_s, Ω_p) completely characterizes a set of impulse response functions. For instance, (Φ_1, Ω_1) corresponds to the impulse response functions obtained in the previous section for Sample 1. The two counterfactual experiments we undertake can then be expressed as (Φ_2, Ω_1) and (Φ_1, Ω_2). Figure 3 displays the resulting impulse response functions, together with the ones obtained for Sample 1 and 2.

Based on this counterfactual exercise we conclude that both sets of parameters are important to explain the reduced effect of exogenous interest rate fluctuations. One way to see this, is that if we are starting from Sample 2, (Φ_2, Ω_2), changing only one of the two sets of parameters would not be sufficient to reproduce the impulse response functions of Sample 1, i.e. (Φ_1, Ω_1), especially in terms of the magnitude.[22]

This analysis thus implies that both changes in monetary policy and in the private sector's response are important. Yet, the underlying cause of these changes is not clear. In fact, there are two potential interpretations consistent with these results. One is that the changes in Ω, which embeds the private sector's response, occurred — at least in part — for reasons unrelated to monetary policy. A second is that the changes in Ω are linked to the change in policy, Φ. For instance, the observed changes in Ω might be entirely due to an adjustment of the way firms and consumers form their expectations to a new policy regime. Under this second scenario, monetary policy would be the only fundamental source of changes. Since this VAR-based analysis does not account for the Lucas critique, the two scenarios cannot be distinguished. This motivates a more structural investigation, to which we now turn.

[22] From the response of inflation, it is interesting to note that the presence of the price-puzzle for Sample 1 appears to be due to the non-policy parameters.

4 Structural analysis of the monetary transmission mechanism

To account for the linkages between the policy and non-policy parameters, we need to identify the structure of the non-policy block. To do so, we use a stylized, but fully specified general equilibrium model that is consistent with the identifying assumption made in the VAR. We estimate this model so that it replicates as well as possible the response of the economy to monetary policy shocks. We then attempt to determine the origin of the changes in the impulse response functions observed for the two samples by using our structural model to perform counterfactual experiments.

4.1 A stylized structural model of the U.S. economy

The model that we consider builds upon the model developed in Rotemberg and Woodford (1997) by allowing for two additional key elements: habit formation in consumption, and backward-looking price setters. These additional features allow the model to better replicate the response of real output, inflation and the interest rate to an unexpected monetary policy shock, in particular in the pre-1980 sample. The model is furthermore set up to be consistent with the structure of the VAR considered in previous sections.

We assume that there is a continuum of households indexed by j, each of which seeks to maximize its utility given by

$$E_t \left\{ \sum_{T=t}^{\infty} \beta^{T-t} \left[u\left(C_T^j - \gamma C_{T-1}^j ; \xi_T \right) - v\left(y_T\left(j\right); \xi_T \right) \right] \right\},$$

where $\beta \in (0,1)$ is the household's discount factor, C_t^j is an index of the household's consumption of each of the differentiated goods at date t, $y_t(j)$ is the amount of the specialized good that household j supplies at date t. The vector ξ_t represents disturbances to preferences. While Rotemberg and Woodford (1997) assume that utility is time-separable, corresponding to the case $\gamma = 0$, we allow the parameter γ to lie between 0 and 1, so that the households' utility depends on the deviation of consumption C_t^j from some habit stock γC_{t-1}^j.[23] As we show below, the presence of habit formation allows us to replicate the hump-shaped response of output to a monetary policy shock.

Following Dixit and Stiglitz (1977), we assume that each household's consumption index aggregates consumption of each good with a constant elasticity of substitution, θ, between goods. It follows that the demand for good z is given by

$$y_t(z) = Y_t \left(\frac{p_t(z)}{P_t} \right)^{-\theta}, \tag{4}$$

where Y_t represents aggregate demand for the composite good, $p_t(z)$ is the price of good z at date t, and P_t is the corresponding price index.

We assume that financial markets are complete, so that risks are efficiently shared. As a result, all households face an identical intertemporal budget constraint, and choose to consume the same amount at any date. We may therefore drop the superscript j in C_t^j. Furthermore, we assume, as in Rotemberg and Woodford (1997), that households must choose their consumption index C_t at date $t-2$, so that $C_{t+2} = E_t C_{t+2}$.[24] This assumption is consistent with the identifying restriction imposed in the VAR considered above, according to which both output and inflation are prevented from responding to a contemporaneous monetary shock. Moreover, an assumption of this kind is needed to account for the fact that monetary policy shocks in the U.S. start exerting a significant effect on GDP after two quarters.

[23] One specification of the utility function u could be for instance $u = (C_t - \gamma C_{t-1} + M)^{1-\rho} / (1-\rho)$, where $M \geq 0$ is large enough for the whole term in parenthesis to be positive (for all dates and all states). Boldrin, Christiano and Fisher (1999) assume a simplified version of this utility function of the form $u = \log(C_t - \gamma C_{t-1})$. In contrast, Amato and Laubach (2000b) and Fuhrer (2000) consider monetary models with "multiplicative" habit formation introduced by Abel (1990) and Galí (1994).

[24] Another interpretation of this assumption is that households choose their consumption using information regarding the state of the economy two periods earlier.

While our setup does not explicitly model the demand for capital goods, we view C_t more broadly as representing the interest-sensitive part of GDP, that is, roughly the amount of consumption and investment goods, assuming crudely that all goods purchases are made to derive utility. Certainly, our model does not take into account the effects of investment on future productive capacities, but we hope that this effect is not too large on the business cycle frequency movements that we consider.[25]

The household's optimal choice of consumption satisfies

$$E_t \left\{ \lambda_{t+2} P_{t+2} \right\} = E_t \left\{ u_c \left(C_{t+2} - \gamma C_{t+1}; \xi_{t+2} \right) - \beta \gamma u_c \left(C_{t+3} - \gamma C_{t+2}; \xi_{t+3} \right) \right\}, \tag{5}$$

where λ_t represents the household's marginal utility of additional nominal income at date t. This equation indicates that at date t, the household chooses a consumption level C_{t+2} for period $t+2$ that equates the expected utility of additional consumption with the expected marginal utility of additional nominal income. While the first term on the right-hand side of (5) represents the expected effect of a change in consumption at date $t+2$ on instantaneous utility at that date, the second term represents the effect of a change in C_{t+2} on instantaneous utility in the following period, through its effect on the stock of habit. The marginal utilities of income furthermore satisfy

$$\lambda_t = \beta R_t E_t \lambda_{t+1}, \tag{6}$$

where R_t is the gross return on a riskless nominal one-period asset. Finally, we use the goods market clearing condition $C_t = Y_t$ to substitute for consumption in (5).[26] The resulting equation, together with (6), characterize the link between the interest rate and aggregate demand.

[25] To the extent that C_t also represents investment spending, the assumption that it is planned two periods in advance also relates to the time-to-build assumption introduced by Kydland and Prescott (1982).

[26] We could easily generalize the goods market equilibrium condition to $C_t + G_t = Y_t$, where G_t represents non-interest sensitive expenditures such as government spending. None of our results would be affected by this however, as we only use the model to analyze impulse responses to monetary policy shocks.

We will consider log-linear approximations of equations (5) and (6) around a steady state in which there are no exogenous disturbances, output growth is constant, and prices are stable. The approximations of these equations yield

$$
\begin{aligned}
E_t\left\{\hat{\lambda}_{t+2}\right\} &= -\frac{\sigma}{1-\beta\gamma}E_t\left\{\left(1+\beta\gamma^2\right)\hat{Y}_{t+2}-\gamma\hat{Y}_{t+1}-\beta\gamma\hat{Y}_{t+3}-g_{t+2}+\beta\gamma g_{t+3}\right\}, &(7)\\
\hat{\lambda}_t &= E_t\left\{\hat{\lambda}_{t+1}+\hat{R}_t-\pi_{t+1}\right\}, &(8)
\end{aligned}
$$

where $\hat{\lambda}_t$, $\hat{Y}_t$, and $\hat{R}_t$ represent respectively percent deviations of $(\lambda_t P_t)$, Y_t, and R_t from their steady-state level, $\pi_t \equiv \log\left(P_t/P_{t-1}\right)$, and $g_t \equiv \frac{u_{c\xi}}{u_c\sigma}\xi_t$ represents exogenous shifts in marginal utility of consumption.[27] The coefficient $\sigma \equiv -u_{cc}\bar{C}/u_c > 0$ represents the inverse of the intertemporal elasticity of substitution (EIS) of consumption evaluated at the steady-state consumption level, in the absence of habit-formation. Since it is difficult to interpret σ in the presence of habit formation, we focus on a pseudo-EIS, which is the elasticity of expected output growth with respect to changes in the real return, conditional on output growth remaining constant in other periods. Taking first differences of equation (7), and combining with (8), we observe that the pseudo-EIS is given by $\frac{(1-\beta\gamma)}{\sigma(1+\beta\gamma^2)}$.

Equations (7) and (8) form what is sometimes called the "IS block" as they result in a negative relationship between the real interest rate and aggregate demand. To see this, we solve (8) forward, to obtain

$$\hat{\lambda}_t = \hat{r}_t^L \equiv \sum_{T=t}^{\infty} E_t\left(\hat{R}_T - \pi_{T+1}\right),$$

where $\hat{r}_t^L$ represents the percentage deviations of a long-run real rate of return from steady state. Combining this with (7), and recalling that $E_t\hat{Y}_{t+2} = \hat{Y}_{t+2}$, we obtain finally

$$\hat{Y}_t = \frac{1}{1+\beta\gamma^2}E_{t-2}\left(-\frac{1-\beta\gamma}{\sigma}\hat{r}_t^L + \gamma\hat{Y}_{t-1} + \beta\gamma\hat{Y}_{t+1} + g_t - \beta\gamma g_{t+1}\right).$$

[27] We view the variables used in the VAR – output gap, inflation and the Fed funds rate – as the empirical counterparts of $\hat{Y}_t$, π_t and $\hat{R}_t$.

Note that in the absence of habit formation, this expression reduces to the familiar equation $\hat{Y}_t = E_{t-2}\left(-\sigma^{-1}\hat{r}_t^L + g_t\right)$ derived, e.g., in Rotemberg and Woodford (1997).

Monetary policy has real effects in this model, because it is assumed that not all suppliers are able to adjust their prices in response to perturbations. Specifically, we assume as in Calvo (1983) that a fraction $(1-\alpha)$ of suppliers can choose a new price at the end of any given period, while the remaining sellers have to maintain their old prices. The timing that we assume implies that the sellers who get to change their prices at date t must decide on the basis of information available at date $t-1$, which is consistent with the assumption made in the structural VAR to identify monetary policy shocks. Following Galí and Gertler (1999), Amato and Laubach (2000a), and Steinsson (2000), we assume furthermore that there is a fraction $(1-\eta)$ of forward-looking suppliers, who seek to maximize their utility, while the remaining fraction η of suppliers is backward-looking, as sellers choose their prices by using a simple rule of thumb. While we do not attempt to model precisely why some sellers might act according to a rule of thumb, such an assumption allows the model to replicate better the sluggish response of inflation to monetary shocks.

Since every supplier faces the same demand function given by (4), all forward-looking suppliers allowed to change their price in period t will choose the same price, p_t^f, that maximizes

$$E_{t-1}\left\{\sum_{T=t}^{\infty}(\alpha\beta)^{T-t}\left[\lambda_T p_t^f Y_T\left(\frac{p_t^f}{P_T}\right)^{-\theta} - v\left(Y_T\left(\frac{p_t^f}{P_T}\right)^{-\theta};\xi_T\right)\right]\right\}.$$

While the first term inside the brackets represents the contribution to expected utility from sales revenues at date T, given that the seller chooses a price p_t^f, the second term represents disutility resulting from the supply of goods demanded at date T. The household discounts the stream of utilities by a factor $\alpha\beta$, rather than β, to account for the fact that the price chosen at date t will apply in period T with probability α^{T-t}. Log-linearizing the first-order condition to the previous problem, solving for $\hat{p}_t^f \equiv \log\left(p_t^f/P_t\right)$, and quasi-differentiating the resulting expression yields the optimal pricing decision for the forward-looking suppliers

$$\hat{p}_t^f = \alpha\beta E_{t-1}\hat{p}_{t+1}^f + \frac{1-\alpha\beta}{1+\omega\theta}E_{t-1}\left(\omega\hat{Y}_t - \hat{\lambda}_t - q_t\right) + \alpha\beta E_{t-1}\pi_{t+1}, \tag{9}$$

where $\omega \equiv v_{yy}\bar{Y}/v_y$ is the elasticity of the marginal disutility of producing output with respect to an increase in output, and $q_t \equiv -v_{y\xi}/v_y\xi_t$ measures exogenous shifts in the disutility of producing output.

Turning to the backward-looking suppliers, we assume as in Galí and Gertler (1999), that they set their prices p_t^b according to the simple rule of thumb

$$p_t^b = P_{t-1}^* \frac{P_{t-1}}{P_{t-2}},$$

where P_t^* is the aggregate of the prices newly chosen at date t by both forward- and backward-looking price setters. Even though the model does not provide a rational explanation for the choice of prices p_t^b, we note that the latter eventually converge to the prices p_t^f in the absence of further perturbations, since they depend on the prices chosen by the forward-looking price setters in the previous period.

Assuming furthermore that the price-setters who are allowed to change their price are chosen independently of their history of price changes implies that $P_t^* = \left[(1-\eta)\left(p_t^f\right)^{1-\theta} + \eta\left(p_t^b\right)^{1-\theta}\right]^{1/(1-\theta)}$ and $P_t = \left[(1-\alpha)\left(P_t^*\right)^{1-\theta} + \alpha P_{t-1}^{1-\theta}\right]^{1/(1-\theta)}$. Log-linearizing (9) and the laws of motion for P_t^* and P_t, and combining the resulting expressions with (9) yields the following variant of the new-Keynesian aggregate supply equation

$$\pi_t = \kappa E_{t-1}\left(\omega\hat{Y}_t - \hat{\lambda}_t - q_t\right) + \chi^b\pi_{t-1} + \chi^f\beta E_{t-1}\pi_{t+1}, \tag{10}$$

where the term in parenthesis is a measure of the gap between equilibrium output and its natural rate, and where $\chi^b \equiv \frac{\eta}{\alpha+\eta(1-\alpha+\alpha\beta)}$, $\chi^f \equiv \frac{\alpha}{\alpha+\eta(1-\alpha+\alpha\beta)}$, and $\kappa \equiv \frac{(1-\alpha)(1-\alpha\beta)(1-\eta)\chi^f}{(1+\omega\theta)\alpha}$. In the special case in which all suppliers are forward-looking, we have $\eta = 0$, which implies $\chi^f = 1$ and $\chi^b = 0$.

In this case, as in the familiar New Keynesian supply equation, inflation depends positively on the expectation of the gap between output and its natural rate, as well as on the expectation of future inflation. Here it is the expectation formed at date $t-1$ that is relevant for the determination of period-t inflation, as sellers are assumed to set their prices on the basis of information available at date $t-1$. More generally, when $\eta > 0$, inflation also depends on its lagged value, as some sellers set their prices according to the simple rule of thumb. While we here allow for backward-looking price setters who are not fully rational, we would obtain an aggregate supply equation identical to (10), but with slightly different restrictions on the coefficients χ^b, χ^f, if we assumed instead that all suppliers are rational and forward looking, but that prices are indexed by past inflation, as in Christiano, Eichenbaum and Evans (2001), and Woodford (2001).

The model that we use for the joint determination of the evolution of inflation, real output and the short-run and long-run interest rates (all expressed in terms of deviations from their steady state), can be summarized by the "IS block" composed of equations (7) and (8), the aggregate supply equation (10), and an interest-rate feedback rule of the form (3). The resulting system of linear difference equations can then be solved using standard methods (e.g., King and Watson (1998)), to obtain a unique bounded rational expectations equilibrium, provided that such an equilibrium exists, and that it is unique.

4.2 Estimation of the structural model

We now turn to the estimation of the structural model just described. Before discussing the results, we describe our econometric methodology.[28] In section 2, we estimated a structural VAR, that allowed us to generate impulse response functions to monetary policy innovations. In the previous

[28] A similar estimation procedure can be found in Rotemberg and Woodford (1997), Amato and Laubach (1999), Gilchrist and Williams (2000) and Christiano, Eichenbaum, and Evans (2001).

subsection, we presented a model that is consistent with the identifying assumption imposed in the VAR, and that delivers impulse responses of the variables of interest for a given set of structural parameters. Our econometric methodology involves selecting the structural parameters that minimize the distance between the estimated VAR responses and the model-based responses. In a way, this can be seen as a calibration exercise. As we now discuss, however, it is a well-defined econometric exercise that can be seen as an application of "semi-parametric indirect inference" (Dridi and Renault (2001)).[29]

More formally, we consider as before the vector of policy coefficients Φ_s for Sample s, the vector Ω_s containing the remaining coefficients of the structural VAR, and $G_V(\Phi_s, \Omega_s)$, the vector-valued function that collects the *VAR-based* impulse response functions of output, inflation and the interest rate to a monetary policy innovation. In addition, we denote by Δ_s the vector of structural parameters of our model, and by $G_M(\Phi_s, \Delta_s)$ the corresponding vector-valued function that collects the *model-based* impulse response functions, provided that the model delivers a unique bounded rational expectations equilibrium. Let $G(\Phi_s, \Omega_s, \Delta_s) \equiv G_M(\Phi_s, \Delta_s) - G_V(\Phi_s, \Omega_s)$. Having estimated Φ_s and Ω_s, using the technique described in section 2, we minimize

$$L(\Delta_s) = G\left(\hat{\Phi}_s, \hat{\Omega}_s, \Delta_s\right)' W_s G\left(\hat{\Phi}_s, \hat{\Omega}_s, \Delta_s\right) \tag{11}$$

with respect to Δ_s to obtain the minimum distance estimator $\hat{\Delta}_s$, where W_s is a positive definite weighting matrix which we discuss below. Note that since the policy coefficients Φ_s are estimated directly from the VAR, we do not need to estimate them again when we estimate Δ_s.

This estimation strategy is advantageous to us for several reasons. First, since we are interested in explaining the observed changes in the impulse response function to a monetary shock in the two

[29] Our estimation method is also similar in spirit to the specification test used by Cogley and Nason (1995), although they were not concerned with the estimation of the structural parameters.

periods considered, it is very natural to estimate the structural parameters directly on the basis of the impulse responses functions. Certainly, more efficient estimates of the structural parameters could be obtained by exploiting the response of the economy to other shocks, but this would require plausible identification of these shocks. Moreover, to the extent that the model is unable to explain all the features of the data, the estimation on the basis of responses to monetary shocks allows us to focus the estimation on the relevant empirical features of the data that we seek to explain. In this sense, the estimation approach is robust to the specification of parts of the model that are not related to the impulse response functions we are interested in.[30] Specifically, while the endogenous variables are affected by the demand and supply shocks g_t and q_t in the theoretical model, our econometric strategy allows us to estimate the structural parameters of interests without estimating the parameters that characterize the stochastic processes $\{g_t\}$ and $\{q_t\}$. Finally, as Hall (2001) pointed out, estimation through impluse response functions has an important advantage over the application of GMM to Euler equations: it indirectly imposes the model's boundary conditions.[31]

The model that we seek to estimate — in order to determine the evolution of inflation, real output and the nominal interest rate — can be summarized by the structural equations (7), (8),

[30]The robustness of this estimation approach to a misspecification of the theoretical model is discussed more generally in Dridi and Renault (2001).

[31]According to Hall (2001, p. 9): "The Euler equation holds for wildly non-optimal behavior as well as for optimal behavior that satisfies the terminal condition. Consequently, an estimator that incorporates the terminal condition pins down parameter values more effectively than one that considers only the Euler equation." The above description of our model does not formally specify terminal conditions, because these conditions are automatically satisfied once we restrict ourselves to bounded fluctuations of the endogenous variables around the steady state. Nevertheless our estimation method retains the advantage mentioned by Hall (2001) as it incorporates the assumption that endogenous variables are bounded.

(10), and the policy reaction function (3). Leaving aside the coefficients of the policy rule which we have estimated earlier, we need to quantify the seven structural parameters β, σ, κ, ω, γ, χ^b, and χ^f. All of these parameters can be separately identified from the impulse response functions to a monetary policy shock. However, in order to reduce the set of parameters to estimate, we calibrate β to 0.99, because it can be identified directly from first moments of the data. In fact β^{-1} corresponds to the steady-state gross real rate of return, which is approximately 1.01 on average. Moreover, instead of estimating separately χ^b and χ^f, we choose χ^f to equate $1-\chi^b$. While this constraint is generally not exactly satisfied in the theoretical model, the approximation error on χ^f, $\frac{\alpha\eta(1-\beta)}{\alpha+\eta(1-\alpha+\alpha\beta)}$, is very small, i.e., less than 0.006, given the value of β. We thus estimate the remaining five parameters σ, κ, ω, γ, and χ^b by matching the model-based impulse response functions with those of the VAR. We consider the responses of the variables over the first eight periods following the monetary shock. This choice is motivated by the fact that most of the difference in the output response in two samples occurs within this horizon. Moreover this corresponds approximately to the time that it takes for output to return to its initial level, following a monetary policy shock.

For some parameter configurations, the model may result in an indeterminate equilibrium.[32] This may arise when the policy reaction function involves too little a response to changes in economic conditions.[33] Clarida *et al.* (2000) argue that the policy reaction function estimated for the pre-

[32] This means that for any bounded solution $\{z_t\}$, where z_t is the vector of variables of interest $\left[\hat{Y}_t, \pi_t, \hat{R}_t\right]'$, there exists another bounded solution of the form

$$z_t' = z_t + v\epsilon_t,$$

where v is an appropriately chosen (nonzero) vector, and the stochastic process $\{\epsilon_t\}$ may involve arbitrarily large fluctuations, that may or may not be correlated with the fundamental disturbances $\{\varepsilon_t, g_t, q_t\}$. It follows that for such a parameter configuration, the model may involve arbitrarily large fluctuations of real output, inflation and the interest rate, independently of the size of the fundamental shocks.

[33] See, e.g., Woodford (1999a) for a complete discussion of the problem of indeterminacy of the equilibrium in

Volcker years is consistent with such a situation. We also find that for a range of structural parameter values, the model results in an indeterminate equilibrium, when the policy reaction function is the one estimated for Sample 1. However, the equilibrium is determinate for another range of structural parameters. In the estimation of the structural parameters, we consider only the combinations that result in a unique bounded equilibrium.[34] Moreover, as the structural model imposes restrictions on the sign and magnitude of the structural parameters, we also impose these restrictions in the numerical minimization of (11).

To estimate the structural parameters, we also need to determine an asymptotically non-stochastic weighting matrix W_s indicated in (11). We consider three weighting matrices. First, we perform the estimation with an identity weighting matrix, so that we are in fact minimizing the sum-of-squared deviations between the model-based and estimated impulse response functions. The advantage of such a matrix is that it yields estimates of the structural parameters that provide the best fit of the VAR-based impulse response functions. This weighting scheme does however not take into account the fact that some impulse responses are less precisely estimated than others. To remedy this problem, we use an alternative diagonal weighting matrix that involves the inverse of each impulse response's variance on the main diagonal. Finally, we consider the efficient weighting matrix, i.e., the inverse of the impulse response functions' variance-covariance matrix.

monetary models of the kind analyzed here.

[34]While this restriction may prevent us from obtaining the best possible fit of the impulse response functions to a monetary shock in Sample 1, it does not affect our final conclusion that the change in the policy rule is the most important source of changes in the impulse response functions. In fact, it is precisely the change in the policy rule that makes it impossible, in our model, for the equilibrium to be indeterminate in the second sample.

4.3 Estimation results

Table 4 reports the structural parameters' estimates, along with the associated standard deviations, for Sample 1 and Sample 2, and using either the identity weighting matrix or the alternative diagonal matrix discussed above. The first thing to note is that both weighting matrices yield very similar results. Given this, and since we ultimately seek to explain the overall change in the point estimates of the impulse response functions, we focus our discussion on the results obtained with the identity weighting matrix. In the first sample, the estimate of γ indicates a high and significant degree of habit formation in consumption. The implied pseudo-EIS amounts to 0.32 in the first sample. In contrast, in the second sample, which is the same as the one used by Rotemberg and Woodford (1997), our estimate of γ indicates no habit formation. It follows that the estimated pseudo-EIS, which in this case corresponds to the estimated EIS, is 2.16. While this number is smaller than the one found in Rotemberg and Woodford (1997) for the same sample, it is in the range of numbers reported in numerous studies. Note that our estimate refers to the elasticity of expected *output* growth with respect to changes in the real return, which is likely to be higher than the corresponding elasticity for nondurable consumption. In any case, the comparison between the two samples indicates that output growth has become *more* sensitive to changes in the real rate of return in the post-1980 sample than in the pre-1980 sample. This suggests that, if anything, changes in the instrument of monetary policy have had a *stronger* effect on output after 1980.

While the estimated slope of the aggregate supply equation, κ, is close to zero in the first sample, and in fact not significantly different from zero, it is higher and significant in the second sample. This is consistent with an increased price flexibility (i.e., a decline in the probability α) in the post-1980 period. In contrast, the estimate of ω, which measures the elasticity of marginal disutility of producing output with respect to an increase in output, is considerably lower in the second sample, suggesting that the disutility of supplying goods is almost linear in the amount of goods. Finally, the estimated χ^b measuring the degree of backward-looking behavior or inflation inertia in the aggregate supply equation is almost reduced by half from the first to the second sample, indicating that the price-setters are substantially more forward-looking in the post-1980 sample. Note however that the standard deviations are fairly large in the first sample — except for γ —, suggesting that most structural parameters are imprecisely estimated in this sample.

It is difficult to provide justification for changes in certain "deep" parameters, such as those of the utility function. Although we doubt that the private sector has changed in such a way that households do not care about their habit stock any more since the beginning of the 1980's, that their marginal disutility of producing output has become insensitive to changes in output, that price setters have become much more forward-looking, and that prices have become more flexible, we view instead these estimates as indications that the private sector of the economy has reacted more promptly to changes in economic conditions in the post-1980 sample than in the pre-1980 sample. Moreover, as we show in the next section, our conclusions concerning the effet of monetary policy do not depend on the interpretation of these changes in the parameters describing the private sector behavior.

Figures 3a and 3b plot both the impulse response functions estimated from the VAR (circles), along with their 95 percent confidence intervals, and the corresponding impulse response functions generated by the theoretical model (solid lines), for both samples, using the estimates obtained with the identity weighting matrix. Notice that the model is able to replicate quite precisely both the magnitude and the persistence of the impulse responses generated by the VAR, and the model-based impulse responses remain consistently within the confidence interval. For the first sample, the model reproduces reasonably well the hump-shaped response of output, the progressive decline in inflation, and the response of the interest rate. For the second sample, the fit is even better. The model captures the rapid decline followed by a return to steady state, both in inflation and output, and it tracks the response of the interest rate.

We finally estimated the model using the efficient weighting matrix, i.e., the inverse of the impulse response functions' variance-covariance matrix, to weigh each of the impulse responses. While the estimated coeffcients for Sample 1 are similar to those reported in Table 4, the estimated coefficients for Sample 2 are sensibly different from those reported. We chose not to report these estimated parameters, as the implied impulse reponse functions do not fit the VAR-based responses as well as with the alternative weighting matrices, especially in Sample 2. The efficient weighting matrix has however the advantage of providing us with Hansen's J-test of the model's specification. The J-statistics are respectively 15.54 and 9.15 for Samples 1 and 2. Given that the vectors $G\left(\hat{\Phi}_s, \hat{\Omega}_s, \hat{\Delta}_s\right)$ in (11) have 24 rows, the p-values associated to the J-statistics, and read from a χ^2_{24} distribution, are respectively 0.904, and 0.997 for Samples 1 and 2. Thus there is no evidence that our structural model is rejected on the basis of its ability to fit the VAR-based impulse response functions.

5 Model-based counterfactual analysis

Now that we have argued that our model captures reasonably well the effects of monetary shocks on output, inflation, and the interest rate in both samples, we finally investigate whether the reduced effect of monetary policy shocks in the second sample is due to an improvement in monetary policy or to a change in the private sector's response.

Figure 4 summarizes our results. It displays the impulse response functions generated by the model for four combinations of structural parameters, Δ, and policy coefficients, Φ. A comparison of the two sets of responses generated by (Φ_1, Δ_1) and (Φ_1, Δ_2), using monetary policy estimated for Sample 1, reveals that the change in the private sector's response — i.e., in the structure of the economy — has generated a more rapid response of the endogenous variables to monetary policy shocks. In particular, output and inflation first decrease more rapidly, and then return to their steady state faster, with the parameters of the second sample. Similarly, by comparing the corresponding impulse responses (Φ_2, Δ_1) and (Φ_2, Δ_2) constructed using the monetary policy of Sample 2, we note again that the change in the structural parameters is responsible for a *faster* reaction of the economy to a monetary policy shock. This is consistent with the above discussion of the estimated structural parameters in both samples.

It is important to note, however, that the change in structural parameters, for given policy, is not associated with a reduction in the *magnitude* of the impulse responses. In fact, the responses of output and inflation to an innovation in the interest rate are *larger* with the parameters Δ_2 than with the parameters Δ_1, in the first few quarters, and the maximal effect on output is slightly larger with the parameters Δ_2 than with Δ_1, for given monetary policy. After three or four quarters, however, the impulse responses of output and inflation are smaller with the parameters Δ_2, as they converge faster to the initial steady-state.

Most of the observed reduction in the *magnitude* of impulse responses appears to be attributable to monetary policy. In fact, by changing monetary policy and maintaining the structural parameters fixed — i.e., by comparing the lines (Φ_1, Δ_1) to (Φ_2, Δ_1), and (Φ_1, Δ_2) to (Φ_2, Δ_2) — we note that the responses of output and inflation associated with the policy estimated for Sample 2 involve considerably less variation than those associated with the policy of Sample 1. In addition, by maintaining the structural parameters constant at Δ_1, we observe that a change in policy from Φ_1 to Φ_2 almost entirely explains the impulse responses (Φ_2, Δ_2) obtained in the second period.

Table 5 provides further evidence of the importance of monetary policy for the reduction in the variability of output and inflation attributable to interest rate shocks. This table reports the variance of expected changes in the variables of interest between periods t and $t+k$, due to exogenous monetary policy shocks, for various horizons k, and the four combinations (Φ, Δ). For instance, the first line below var $\left(E_t\hat{Y}_{t+k} - \hat{Y}_t\right)$ corresponds to the variance of the change in output predicted by the model at a two-quarter horizon, when both the policy reaction function and the structural parameters are the ones of Sample 1. Note that since the variables $\hat{Y}, \pi, \hat{R}$ are assumed stationary, the last four lines of Table 5 contain simply the variances of the respective variables, *conditional* on monetary policy shocks. In fact, for any stationary variable x, the statistic var $(E_t x_{t+k} - x_t)$ is equal to var (x_t) when $k = \infty$. For all these experiments, we assume that there are no other shocks besides an exogenous monetary policy shock, and we set the variance of the innovations ε_t to 1.0.[35]

One interesting fact revealed by Table 5 is that for any variable x and almost any horizon k considered, the conditional variances $V \equiv \text{var}\,(E_t x_{t+k} - x_t)$ are ranked as follows:[36]

$$V_{(\Phi_1,\Delta_1)} > V_{(\Phi_1,\Delta_2)} > V_{(\Phi_2,\Delta_1)} > V_{(\Phi_2,\Delta_2)}.$$

Taking together the inequalities $V_{(\Phi_1,\Delta_1)} > V_{(\Phi_2,\Delta_1)}$ and $V_{(\Phi_1,\Delta_2)} > V_{(\Phi_2,\Delta_2)}$ confirm that the more responsive monetary policy of Sample 2 results in a smaller variability of output, inflation and the interest rate, regardless of the set of structural parameters Δ_1 or Δ_2, and independently of the horizon k considered. Taking the inequalities $V_{(\Phi_1,\Delta_1)} > V_{(\Phi_1,\Delta_2)}$ and $V_{(\Phi_2,\Delta_1)} > V_{(\Phi_2,\Delta_2)}$ indicates that for given policy, the change in the private sector's response, reflected in a change from Δ_1 to Δ_2, results also in a reduced variability of the variables of interest (except for output at horizon $k = 2$). However, since the conditional variances decrease more by changing only policy from

[35] Alternative values maintain exactly the same qualitative results.

[36] The only exception is obtained for output and $k = 2$, in which case the first and third inequalities are reversed. This reinforces our conclusion even more.

Φ_1 to Φ_2 than by changing only the structural parameters from Δ_1 to Δ_2, we conclude that the reduced variance is mainly due to the more responsive monetary policy.[37] For instance, in the case $k = \infty$, we note that the change in the private sector's response mechanism alone is responsible for a decrease in the variance of output from 2.07 to 1.45, while the change in monetary policy alone brings the variance of output down to 0.34.

Overall, these experiments suggest that the change in monetary policy has been the main cause underlying the reduced effect of exogenous interest rate fluctuations on output and inflation. Note finally that for the policy Φ_2, which is much more responsive to fluctuations in expected inflation and output than Φ_1, changes in the structural parameters have almost no effect on the impulse response functions, and relatively little effect on the variances. In contrast, the parameter changes exert a significant effect on the impulse responses and the variances when the less responsive policy Φ_1 is followed. This is consistent with the finding by Giannoni (2002) that an aggressive monetary policy rule of the kind estimated in the second sample tends to be more *robust* to uncertainty about the structural parameters, than less aggressive policies such as the one estimated in the first sample. In fact, to the extent that the central bank faces uncertainty about the exact values of the structural parameters Δ, a more aggressive policy makes it more likely for the variances of output, inflation and the interest rate to be contained.

6 Conclusion

Empirical evidence from VAR analyses, including the one presented here, suggests that unexpected exogenous changes in the Fed funds rate have been followed by a smaller response of output and

[37] In fact, for any variable and any horizon considered in Table 5, we have: $\left(V_{(\Phi_1,\Delta_1)} - V_{(\Phi_2,\Delta_1)}\right) > \left(V_{(\Phi_1,\Delta_1)} - V_{(\Phi_1,\Delta_2)}\right)$ and $\left(V_{(\Phi_1,\Delta_2)} - V_{(\Phi_2,\Delta_2)}\right) > \left(V_{(\Phi_2,\Delta_1)} - V_{(\Phi_2,\Delta_2)}\right)$.

inflation since the beginning of the 1980's. In this paper we attempt to determine the causes of this phenomenon. While some authors have pointed to a change in the conduct of monetary policy, others have argued that they are rooted in changes in the private sector's response, i.e., of the way the economy responds to interest rate fluctuations.

The main finding of this paper is that monetary policy is the dominant cause of this change. More precisely, our empirical investigation, based on an identified VAR, confirms the finding by Clarida *et al.* (2000) and Boivin (2001), among others, that monetary policy in the U.S. has become significantly more responsive to expected inflation and expected output after 1980. Moreover, our estimation of a small structural model of the U.S. economy indicates that a change in the behavior of households and firms is responsible for a more rapid response of the economy to monetary policy shocks. However, our model-based counterfactual investigation reveals that the reduced effect of monetary policy shocks on output and inflation is predominantly due to the change in the systematic part of monetary policy. Furthermore, by being more aggressive, monetary policy has also made the economy less sensitive to changes in the parameters characterizing the private sector's behavior, in the face of monetary policy shocks.

Does this evidence imply that monetary policy has become less powerful? Our analysis suggests that this has not been the case. Instead, monetary policy appears to have been conducted in a more stabilizing manner, smoothing out the effect of exogenous variations in the interest rate. Moreover, the mere fact that a large fraction of the changes between the pre- and post-1980 periods can be explained by monetary policy suggests that the policy reaction function has a powerful impact on the economy.

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Table 1: **Hansen J-test for the different horizons**

	h_y								
h_π	0	1	2	3	4	5	6	7	8
					Full sample: 1963:1–1995:4				
0	0.011	0.033	0.005	0.000	0.000	0.000	0.000	0.000	0.000
1	0.032	0.087	0.021	0.002	0.000	0.000	0.000	0.001	0.001
2	0.044	0.117	0.030	0.004	0.000	0.000	0.000	0.001	0.001
3	0.044	0.096	0.025	0.003	0.000	0.000	0.000	0.001	0.001
4	0.038	0.067	0.014	0.002	0.000	0.000	0.000	0.001	0.001
5	0.044	0.085	0.018	0.002	0.000	0.000	0.000	0.001	0.002
6	0.044	0.078	0.019	0.002	0.000	0.000	0.000	0.002	0.002
7	0.014	0.028	0.005	0.000	0.000	0.000	0.000	0.001	0.001
8	0.013	0.026	0.004	0.000	0.000	0.000	0.000	0.001	0.001
					Sample 1: 1963:1–1979:3				
0	0.107	0.229	0.075	0.025	0.006	0.013	0.134	0.514	0.116
1	0.376	0.399	0.171	0.119	0.233	0.408	0.813	0.872	0.581
2	0.258	0.323	0.132	0.121	0.409	0.497	0.769	0.617	0.327
3	0.227	0.302	0.115	0.096	0.283	0.392	0.678	0.468	0.242
4	0.182	0.257	0.087	0.048	0.131	0.176	0.400	0.326	0.182
5	0.288	0.317	0.112	0.071	0.258	0.320	0.473	0.436	0.257
6	0.238	0.283	0.103	0.099	0.462	0.584	0.674	0.417	0.257
7	0.098	0.221	0.079	0.025	0.051	0.110	0.318	0.195	0.110
8	0.096	0.221	0.079	0.027	0.017	0.058	0.283	0.197	0.119

Note: The table reports for each forecasting horizon combination (h_π, h_y), the p-value of the Hansen J-test. A p-value smaller than 0.05 signifies that the model specification is rejected at the 5% level. See text for details.

Table 1: **Hansen J-test for the different horizons (Continued)**

	h_y								
h_π	0	1	2	3	4	5	6	7	8
					Sample 2: 1980:1–1995:4				
0	0.003	0.012	0.002	0.002	0.001	0.001	0.001	0.001	0.001
1	0.013	0.053	0.012	0.010	0.004	0.003	0.002	0.002	0.002
2	0.005	0.023	0.004	0.004	0.002	0.001	0.001	0.001	0.001
3	0.004	0.025	0.004	0.003	0.002	0.001	0.001	0.001	0.001
4	0.007	0.024	0.005	0.003	0.002	0.001	0.001	0.001	0.001
5	0.006	0.021	0.004	0.003	0.001	0.001	0.001	0.001	0.001
6	0.002	0.010	0.002	0.001	0.001	0.000	0.000	0.001	0.001
7	0.001	0.007	0.002	0.001	0.001	0.000	0.001	0.001	0.001
8	0.001	0.003	0.001	0.001	0.000	0.000	0.000	0.000	0.001
					Sample 3: 1984:1–1995:4				
0	0.129	0.178	0.159	0.123	0.090	0.065	0.066	0.047	0.046
1	0.105	0.145	0.139	0.168	0.144	0.093	0.089	0.055	0.054
2	0.149	0.225	0.183	0.233	0.208	0.124	0.107	0.069	0.064
3	0.297	0.424	0.361	0.313	0.284	0.270	0.290	0.210	0.188
4	0.103	0.116	0.106	0.090	0.080	0.061	0.066	0.044	0.042
5	0.097	0.123	0.110	0.101	0.080	0.055	0.060	0.042	0.038
6	0.260	0.309	0.263	0.202	0.161	0.100	0.100	0.083	0.079
7	0.132	0.186	0.158	0.117	0.079	0.058	0.055	0.043	0.042
8	0.089	0.124	0.130	0.106	0.071	0.048	0.048	0.033	0.030

Note: The table reports for each forecasting horizon combination (h_π, h_y), the p-value of the Hansen J-test. A p-value smaller than 0.05 signifies that the model specification is rejected at the 5% level. See text for details.

Table 2: **Stability of the reduced-form VAR**

	Regressors			
Dep. Var	πc	π	$\hat{Y}$	R
πc	0.003	0.224	0.0190	0.079
π	0.062	0.005	0.447	0.000
$\hat{Y}$	0.176	0.002	0.700	0.000
R	0.050	0.067	0.005	0.247

Note: The numbers reported in this table are the p-values for the Andrews (1993) sup-Wald test. Under the null of the test, the coefficients are time-invariant. The test is applied jointly to the constant and coefficients on the lags of the variable corresponding to the given column. The p-values were computed using the simulation approach of Hansen (1997).

Table 3: **Estimates of the policy reaction function**

	Sample 1	Sample 2	Sample 3
ϕ^0	0.138	-0.050	0.663
	(0.285)	(0.326)	(0.340)
ϕ^π	0.144	0.415	0.135
	(0.087)	(0.107)	(0.135)
ϕ^y	0.253	0.461	0.329
	(0.096)	(0.112)	(0.108)
ρ_1	1.285	0.784	1.196
	(0.114)	(0.113)	(0.155)
ρ_2	-0.900	-0.287	-0.480
	(0.182)	(0.150)	(0.251)
ρ_3	0.633	0.218	0.162
	(0.192)	(0.150)	(0.252)
ρ_4	-0.162	0.106	-0.049
	(0.110)	(0.108)	(0.145)

Note: Standard errors are reported in parantheses. Sample 1 corresponds to 1963:1 - 1979:3, Sample 2 to 1980:1 - 1997:4 and Sample 3 to 1984:1 - 1997:4.

Table 4: **Estimates of structural parameters**

Parameters	Sample 1		Sample 2	
	$W_s = I$	$W_s =$ diag.	$W_s = I$	$W_s =$ diag.
σ	0.133	0.134	0.463	0.435
	(0.527)	(0.543)	(0.054)	(0.071)
γ	0.931	0.930	0.000	0.000
	(0.296)	(0.305)	(0.006)	(0.049)
κ	0.006	0.006	0.024	0.023
	(0.046)	(0.047)	(0.006)	(0.011)
ω	9.087	8.980	0.000	0.000
	(56.17)	(55.81)	(0.173)	(0.241)
χ^b	0.977	0.973	0.523	0.517
	(0.803)	(0.803)	(0.014)	(0.022)
pseudo-EIS	0.318	0.319	2.162	2.297

Note: Results based on the minimum distance estimation described in the text, for different weighting matrices. Standard deviations are in parentheses.

Table 5: **Variance of forecastable components**

k		$\text{var}\left(E_t\hat{Y}_{t+k} - \hat{Y}_t\right)$	$\text{var}(E_t\pi_{t+k} - \pi_t)$	$\text{var}\left(E_t\hat{R}_{t+k} - \hat{R}_t\right)$
2	(Φ_1, Δ_1)	0.84	0.46	3.07
2	(Φ_1, Δ_2)	1.60	0.25	2.27
2	(Φ_2, Δ_1)	0.19	0.07	1.54
2	(Φ_2, Δ_2)	0.23	0.04	1.47
4	(Φ_1, Δ_1)	1.59	1.58	6.20
4	(Φ_1, Δ_2)	1.53	0.48	2.55
4	(Φ_2, Δ_1)	0.41	0.24	1.89
4	(Φ_2, Δ_2)	0.28	0.07	1.21
8	(Φ_1, Δ_1)	1.89	3.48	11.33
8	(Φ_1, Δ_2)	1.70	0.79	3.20
8	(Φ_2, Δ_1)	0.66	0.47	2.69
8	(Φ_2, Δ_2)	0.28	0.12	1.46
12	(Φ_1, Δ_1)	2.11	3.96	12.12
12	(Φ_1, Δ_2)	1.67	1.14	2.93
12	(Φ_2, Δ_1)	0.63	0.52	2.35
12	(Φ_2, Δ_2)	0.25	0.16	1.38
∞	(Φ_1, Δ_1)	2.07	12.68	20.47
∞	(Φ_1, Δ_2)	1.45	1.81	2.43
∞	(Φ_2, Δ_1)	0.34	0.52	1.82
∞	(Φ_2, Δ_2)	0.21	0.18	1.27

Figure 1: Impulse Responses to a Monetary Shock over Different Samples

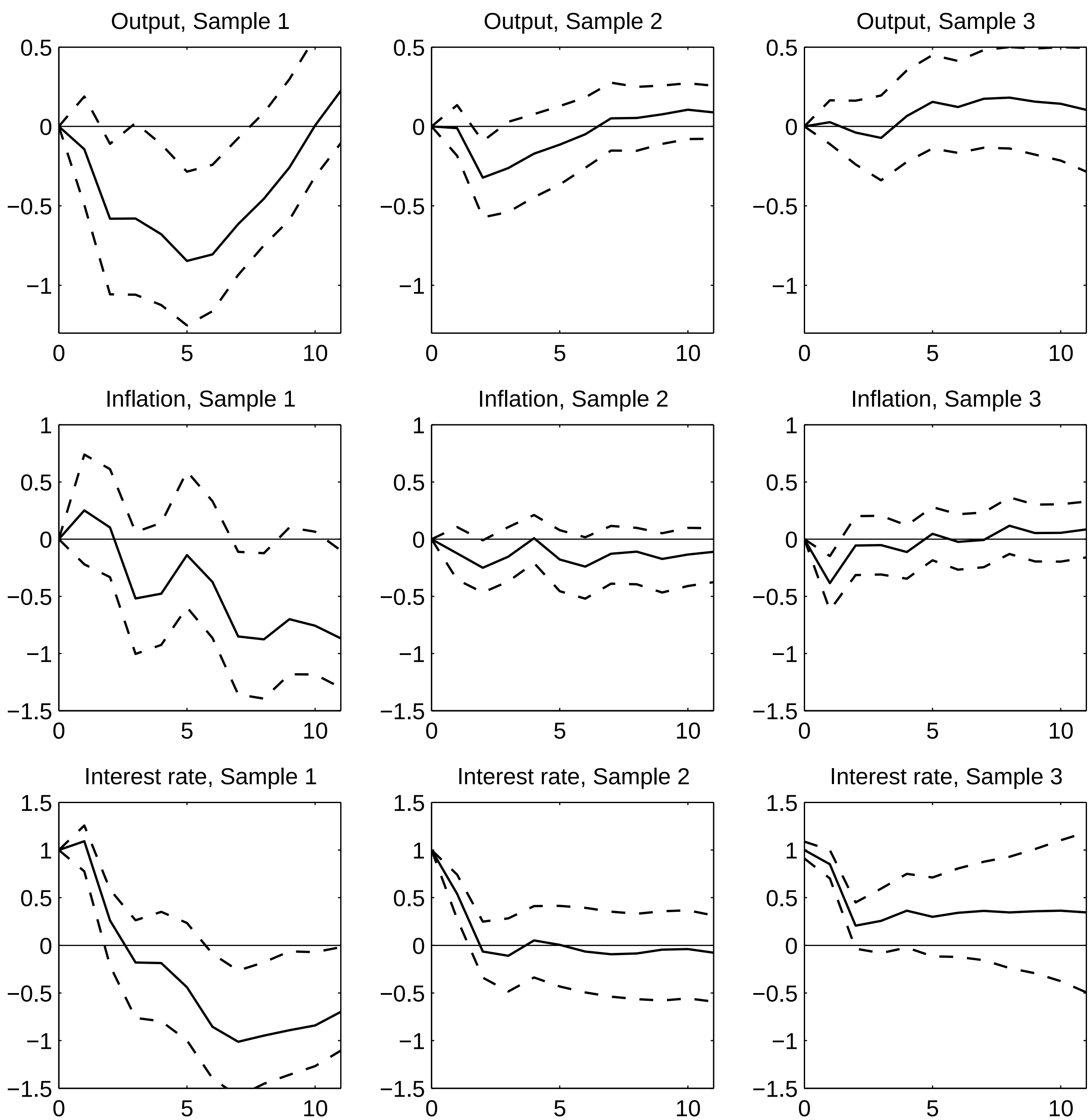

Figure 2: VAR−Based Counterfactual Analysis

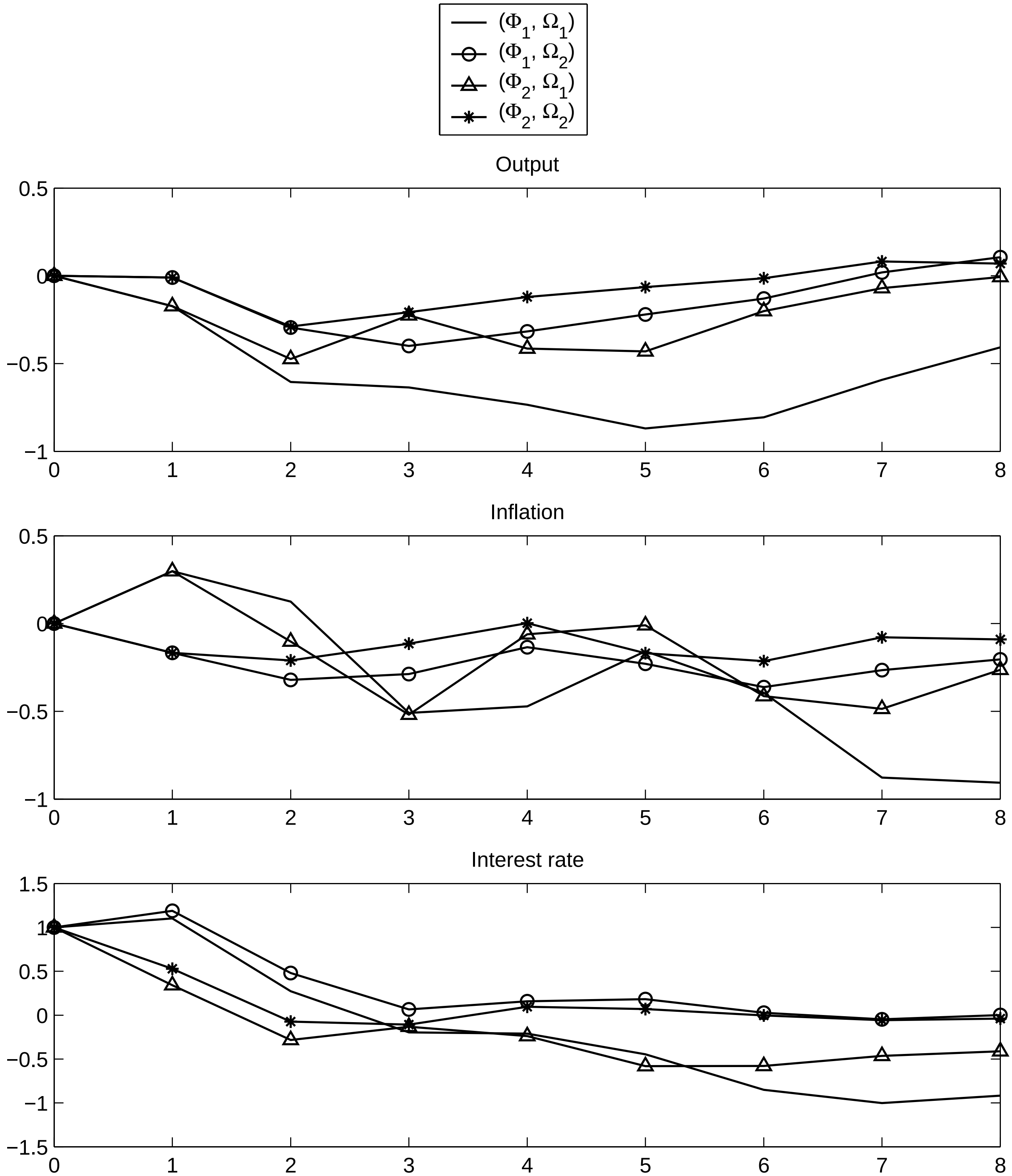

Figure 3a: VAR and Model−Based Impulse Responses to a Monetary Shock (Sample 1)

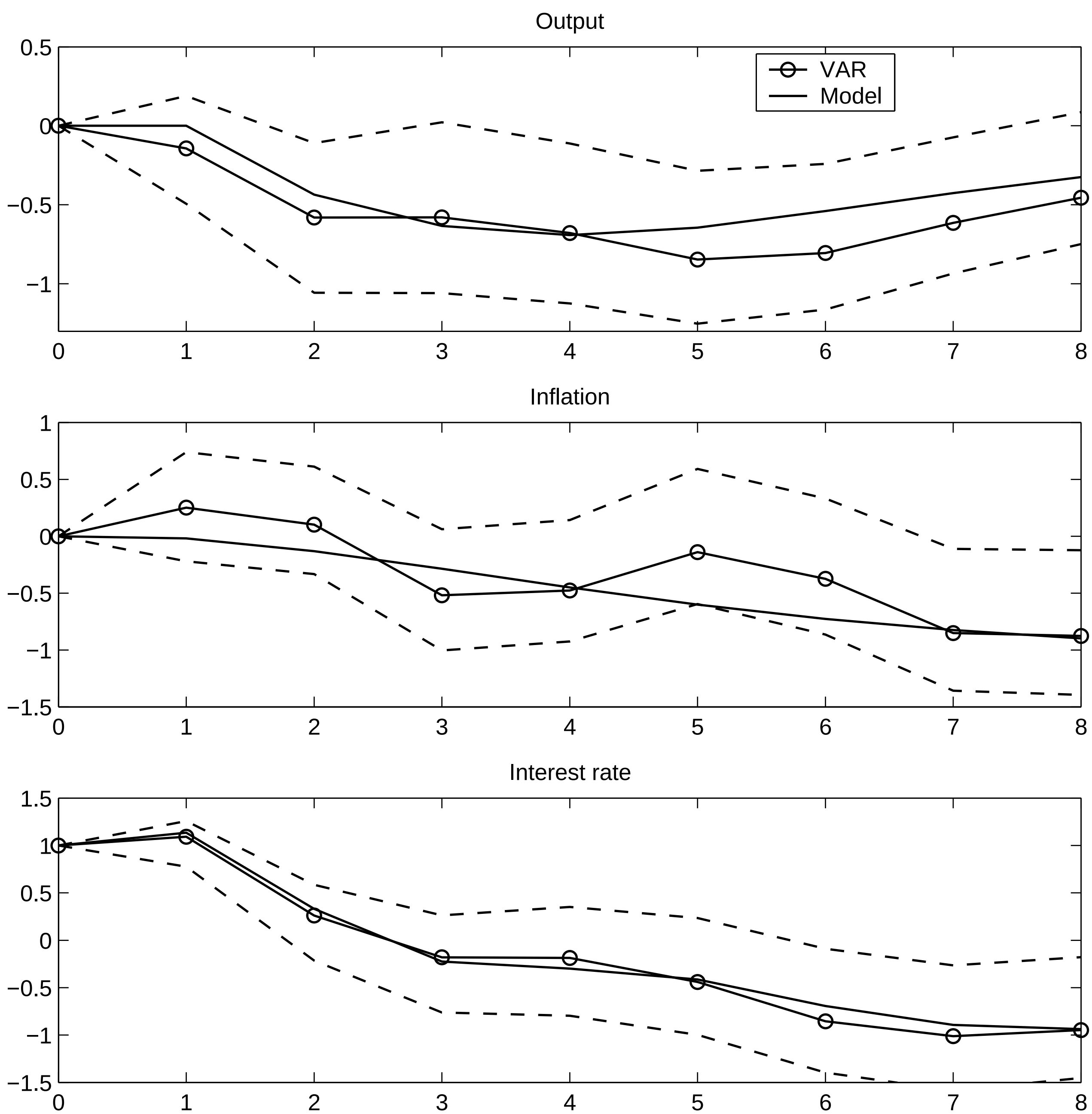

Figure 3b: VAR and Model−Based Impulse Responses to a Monetary Shock (Sample 2)

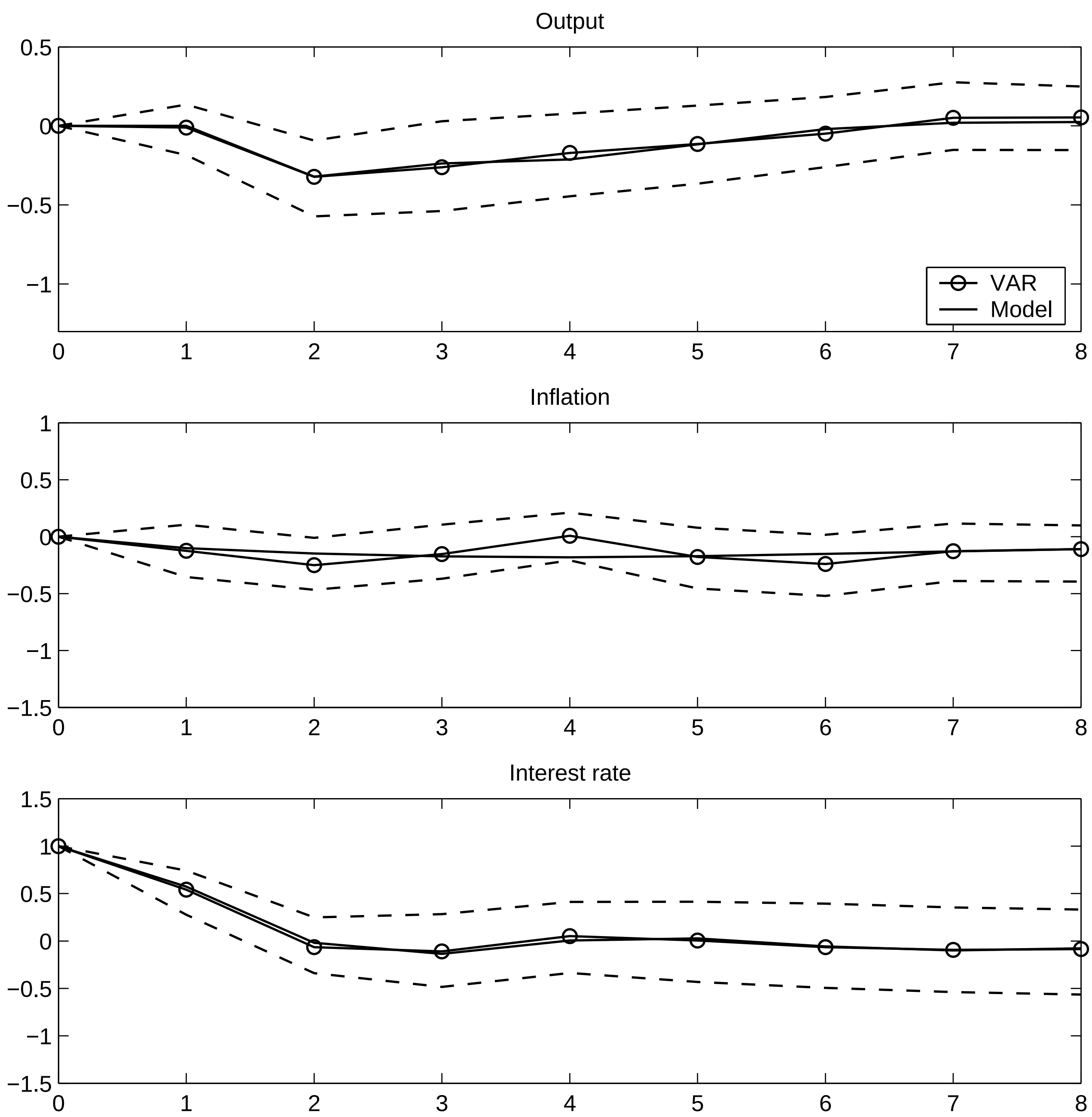

Figure 4: Model−Based Counterfactual Analysis