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ABSTRACT

Recent research provides evidence of important changes in the U.S. economic environment over the last 40 years. This appears to be associated with an alteration of the monetary transmission mechanism. In this paper we investigate the implications for the evolution of monetary policy effectiveness. Using an identified VAR over the pre- and post-1980 periods we first provide evidence of a reduction in the effect of monetary policy shocks in the latter period. We then 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 determine the source of the observed change in the monetary transmission mechanism, as well as in the economy's response to supply and demand shocks. The main finding is that monetary policy has been more stabilizing in the recent past, as a result of both the way it has responded to shocks, but also by ruling out non-fundamental fluctuations.

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1 Introduction

A growing body of evidence, both anecdotal and from formal statistical investigations, suggests that the behavior of the economy has changed in substantial and fundamental ways over the last decades. The important decline in the volatility of U.S. real activity and inflation since the early 1980's is a striking illustration. This evolving economic environment appears to be associated with an alteration of the propagation mechanism of monetary policy. In fact, evidence already exists that points to a change in the impact of monetary policy on output and inflation.[1] Recent studies using vector autoregressions (VAR) find that the impact of monetary policy "shocks" — defined as unexpected exogenous changes in the Federal funds rate — have had a much smaller impact on output and inflation since the beginning of the 1980's.[2] This is illustrated in Figure 1, which shows the response of a measure of detrended output and inflation to a monetary shock of the same size, separately for the pre- and post-1979:4 periods.[3]

This evidence raises the possibility that the effect of monetary policy on the economy has changed in important ways. One possible interpretation is that monetary policy has lost

[1] 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 these papers is that monetary policy's effects appear somewhat weaker recently than in previous decades (see e.g., Kuttner and Mosser (2002)).

[2] 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.

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

some of its influence on the economy. Indeed, various innovations in firms and consumers behavior, perhaps induced by technological progress or financial innovations, might have allowed consumers to better cushion themselves from the impact of interest rate fluctuations.[4] But this is not the only possible interpretation. In fact, not only does the response to monetary policy shocks depend on the behavior of households and firms — in short, the private sector — but also on the way monetary policy is conducted. An alternative interpretation could thus be that monetary policy itself has come to systematically respond more decisively to economic conditions,[5] thereby moderating the real effects of demand fluctuations. In this case, the change in the responses to monetary shocks would reflect an improvement in the effectiveness of monetary policy.

The effectiveness of monetary policy might have changed along other dimensions as well. Not only might monetary policy stabilize more effectively the economy in response to its own shock, but also to other shocks such as real demand and supply disturbances. Another possibility, as suggested by Clarida, Gali and Gertler (2000), is that monetary policy is now more successful at ruling out undesired non-fundamental fluctuations. Furthermore, the size of the policy shocks itself – which could represent random policy mistakes – might have changed over time.

The goal of this paper is to investigate the evolution of monetary policy effectiveness along these various dimensions. We do so by following a two-step strategy. First, using a VAR estimated over the 1959:1-1979:3 and 1979:4-2002:4 periods, we identify a reduced form

[4] 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.

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

policy reaction function and the implied policy shocks. This allows us to identify, with a minimum amount of structure, the monetary transmission mechanism, and to characterize its evolution over the past four decades.

However, this VAR evidence alone does not allow us to properly interpret the source of changes, as the private sector behavior and expectations are not identified separately from the monetary policy behavior. This motivates the second step of our 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 inflation inertia. Since our 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 then perform a series of counterfactual experiments to determine the causes of the observed changes in the monetary transmission mechanism, and the implications for monetary effectiveness. In particular, we can determine how the response to the different shocks of the model has changed, and what part of these changes is due to 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 is that monetary policy has been more stabilizing since the early 1980's. In particular, we find that the reduced effect of monetary policy shocks in the post 1980 period can be almost entirely explained by a increase of the Fed responsiveness to inflation and output. We also find that the current conduct of monetary policy more effectively stabilizes the economy in response to supply and demand shocks. Finally, as Clarida *et al.* (2000) have suggested, we find that the current policy prevents the existence of non-fundamental – sunspot – fluctuations, which was not the case in the pre-1980 period.

The rest of the paper is organized as follows. Section 2 describes our VAR model of the monetary transmission mechanism, and documents the reduction in the effect of monetary policy shocks since the early 1980's. Section 3 constructs and estimates a fully-specified general equilibrium model of the U.S. economy. Section 4 uses this model to interpret the nature of the changes in the monetary transmission mechanism through various counterfactual analyses. Section 5 concludes.

2 Investigating changes in the monetary transmission mechanism

2.1 Empirical model

Our baseline 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 3, as well as a commodity price measure.[6] The commodity price inflation, although not formally justified by the theoretical model, is added to limit the extent of a "price-puzzle" in this VAR.[7] 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 instrument in the U.S. over most of that period (see e.g., Bernanke and Blinder (1992), Bernanke and Mihov (1998)).[8]

In order to identify the policy reaction function and the policy shocks from this VAR, we assume that the economy (Z_t) responds only with a lag to changes in the Fed funds

[6] All series are taken from the Standard and Poor's –DRI database. Detrended output is measured as the deviation of the natural logarithm of quarterly real GDP (mnemonic GDPQ) from a linear deterministic trend. The results are robust to the use of alternative detrending methods, including a quadratic trend and a band pass filter (see our NY Fed Staff Report # 144). $\hat{Y}_t$ 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 2002:2. Four lags are included in the VAR, as determined by the Schwarz information criterion.

[7] This practice is fairly standard in this literature. An alternative proposed by Bernanke, Boivin and Eliasz (2002) is to incorporate more information using a factor-augmented VAR (FAVAR). As shown in the appendix, our VAR results are robust to this alternative.

[8] 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.

rate. Although debatable, this identifying assumption is consistent with many recent VAR analyses.[9] Under this recursive structure, the identified VAR can be expressed as:

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

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

Equation (2) constitutes an unrestricted specification of the policy reaction function, which can be estimated directly by OLS. As we discuss below, the policy reaction function so identified can be seen as a reduced-form expression for the structural policy rule used in the estimation of the structural model.

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 different specifications 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. Furthermore, the robustness analysis discussed in the Appendix shows that the inclusion of more information in our VAR does not affect this conclusion.

[9] See for instance Bernanke and Blinder (1992), Rotemberg and Woodford (1997), Bernanke and Mihov (1998) and Christiano, Eichenbaum and Evans (1999).

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

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 in the bivariate relationships among 76 macroeconomic variables. In the VAR context, mixed results have been obtained.[11] 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.

To investigate the stability of the parameters in the VAR described above, we use an heteroskedasticity robust version of the Bai, Lumsdaine and Stock (1998) multivariate stability test. Under the alternative of this test, the VAR parameters are experiencing a discrete shift at some unknown date. The test allows to test jointly for instability of all the parameters of the VAR and if instability is detected, confidence intervals for the break date can be constructed[12]. Moreover, this class of tests 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-value of the test applied to our VAR is 0.01, suggesting that its parameters and the

[11] Bernanke, Gertler and Watson (1997) and Stock and Watson (2002) find evidence of instability in a monetary VAR, while Bernanke and Mihov (1998) and Christiano, Eichenbaum and Evans (1999) conclude the opposite.

[12] The implementation of the test follows Stock and Watson (2002).

implied propagation mechanism have changed at some point in the last four decades. The economic significance of this change is further emphasized by Boivin and Giannoni (2002) who show, using a similar VAR, that the observed reduction in the volatility of inflation and output is explained roughly equally by a reduction in the variance of the shocks and a smaller propagation.[13]

If we impose on the VAR the structure of the policy rule used in the structural model of Section 3 and apply the test to this equation only, the p-value is 0.00. We can thus already conclude that part of the instability observed in the reduced form VAR arises from changes in the conduct of monetary policy *per se*.

The 90% confidence interval for the VAR parameters break date ranges from the fourth quarter of 1977 to the second quarter of 1986. The break dates is thus quite imprecisely estimated. However, this confidence interval is consistent with a structural change in the economy occurring in the early 1980's, as many of the studies mentioned in the introduction have suggested.

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 for the effect of monetary policy. 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 two different sub-samples. Based on compelling

[13] Boivin and Giannoni (2002) reports stability tests applied only to subsets of these parameters and also suggest the presence of instability. Note also that allowing for a break in the constant at the estimated break date does not change the conclusions reached here.

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 comparison on the following sub-samples: Sample 1 corresponding to the 1959:1-1979:3 period and Sample 2 corresponding to the 1979:4–2002.2 period. The assumed break date corresponds to the one at which Fed chairman Paul Volcker announced a shift in policy. This date is within the confidence interval of the break date estimate and is the one chosen by Rotemberg and Woodford (1997), which we use as a benchmark for our structural analysis in Section 3. Of course, the changes could be argued to have occurred at other points within the confidence interval of the break date estimates. An alternative would be to start the second sample in 1984:1, a date consistent with some estimates of the date of change in the volatility of the U.S. economy.[14] However, as we argue in the Appendix, with a proper account of the relevant information, the differences between the pre- and post-break samples we emphasize are robust to this alternative choice of the break date.

Figure 1 displays — for both 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.[15] 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 previous period; the trough of the response of output is about four times larger in Sample 1 than in Sample 2. This result suggests that the effect of a monetary policy shock of a *given* size was stronger before the 1980's.

Given the imprecision of the estimated impulse response functions, it is difficult to assess

[14] See McConnell and Perez-Quiros (2000).

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

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 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 difference in the point estimates of the impulse response function in the two samples are statistically significant. We see these results, together with the existing evidence[16], as providing compelling evidence of changes in the propagation of monetary policy shocks.

Finally, a by product of this estimation is a measure of the standard deviation of the monetary policy shocks in the two samples. In the first sample the standard deviation is 0.48 compared to 0.60 in the second sample.[17] Taking these numbers literally, this would suggest that monetary policy has not become more successful in reducing random variations in its instrument, perhaps stemming from policy mistakes. However, we know from the

[16] 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).

[17] This result might be surprising at first. However, this is clearly due to the inclusion of the early 1980's in the second sample, where the Fed funds rate was very volatile. While as argued above, the starting date of the sample does not affect the estimated impulse response functions, it does affect size of the policy shocks. For instance, if we start the second sample in 1984:1 instead, the standard deviation of the policy shocks is 0.23.

existing literature that these policy shocks are small and while they are useful to help us identify the monetary transmission mechanism, they have a marginal contribution in business cycle fluctuations. As a result, one would not expect the loss, or the potential gain for that matter, in monetary policy effectiveness to be important along this dimension.

3 Structural analysis of the monetary transmission mechanism

The main goal of this investigation is to determine the implications of these changes for the effectiveness of monetary policy. This requires identifying separately the parameters describing the private sector behavior from those describing the policy behavior. 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. In the next section, we perform counterfactual experiments with this model, to determine the origin of the changes in the impulse response functions observed for the two samples.

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

The model that we consider extends the Rotemberg and Woodford (1997) model by allowing for two additional key elements: habit formation in consumption, and inflation inertia. 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

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

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 $\tilde{\eta} = 0$, we allow the parameter $\tilde{\eta}$ to lie between 0 and 1, so that the households' utility depends on the deviation of consumption C_t^j from some habit stock $\tilde{\eta} C_{t-1}^j$.[18] 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 aggregate consumption index is of the form

$$C_t^j = \left(\int_0^1 c_t^j\left(z\right)^{\frac{\theta-1}{\theta}} \right)^{\frac{\theta}{\theta-1}} \tag{3}$$

with a constant elasticity of substitution $\theta > 1$. It follows that optimal consumption of the

[18] One specification of the utility function u could be for instance $u = \left(C_t - \tilde{\eta} C_{t-1} + M\right)^{1-\rho} / \left(1 - \rho\right)$, 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 $u = \log(C_t - \tilde{\eta} C_{t-1})$. In contrast, Amato and Laubach (2000) and Fuhrer (2000) consider monetary models with "multiplicative" habit formation introduced by Abel (1990) and Galí (1994).

good z is given by the usual expression $c_t^j(z) = C_t^j \left(p_t(z)/P_t\right)^{-\theta}$, where $p_t(z)$ is the price of good z at date t, and P_t is the corresponding Dixit-Stiglitz 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 $C_{t+2} = E_t C_{t+2}$.[19] 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.

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,[20] but we hope that this effect is not too large on the business cycle frequency movements that we consider.

The household's optimal choice of consumption satisfies

$$E_{t-2}\left\{\Lambda_t P_t\right\} = E_{t-2}\left\{u_c\left(C_t - \tilde{\eta} C_{t-1}; \xi_t\right) - \beta\tilde{\eta} u_c\left(C_{t+1} - \tilde{\eta} C_t; \xi_{t+1}\right)\right\}, \tag{4}$$

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

[20] 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).

where Λ_t represents the household's marginal utility of additional nominal income at date t. This equation indicates that at date $t-2$, the household chooses a consumption level C_t for period t 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 (4) represents the expected effect of a change in consumption at date t on instantaneous utility at that date, the second term represents the effect of a change in C_t on instantaneous utility in the following period, through its effect on the stock of habit. The marginal utilities of income furthermore satisfy

$$\Lambda_t = \beta \left(1 + R_t\right) E_t \Lambda_{t+1}, \tag{5}$$

where R_t is the rate of return on a riskless nominal one-period asset.

In addition, we assume that the government purchases an aggregate G_t of all goods in the economy of the form (3). This implies that the demand for good z is given by

$$y_t\left(z\right) = Y_t \left(\frac{p_t\left(z\right)}{P_t}\right)^{-\theta} \tag{6}$$

where the aggregate demand for the composite good, Y_t, satisfies $Y_t = C_t + G_t$.[21] For consistency with the assumption made in our VAR for the identification of monetary policy shocks, we assume that G_t is determined before the interest rate is set in period t, so that G_t is determined on the basis of information available at date $t-1$.

The equations (4) and (5) together with the goods market equilibrium condition just mentioned characterize the link between the interest rate and aggregate demand. We consider log-linear approximations of these equations around a steady state in which there are no

[21] While G_t is associated here with government expenditures, it can more generally represent variations in autonomous (i.e., not interest-rate sensitive) spending.

exogenous disturbances and prices are stable. Approximations of these equations yield

$$
\begin{aligned}
E_{t-2}\left\{\hat{\lambda}_t\right\} &= -\frac{\sigma}{1-\beta\tilde{\eta}} E_{t-2}\left[\left(1+\beta\tilde{\eta}^2\right)\hat{C}_t - \tilde{\eta}\hat{C}_{t-1} - \beta\tilde{\eta}\hat{C}_{t+1} - \bar{C}_t + \beta\tilde{\eta}\bar{C}_{t+1}\right], && (7)\\
\hat{\lambda}_t &= E_t\left(\hat{\lambda}_{t+1} + \hat{R}_t - \pi_{t+1}\right), && (8)\\
\hat{Y}_t &= \left(\bar{C}/\bar{Y}\right)\hat{C}_t + \hat{G}_t && (9)
\end{aligned}
$$

where $\hat{\lambda}_t$, $\hat{C}_t$, $\hat{Y}_t$, and $\hat{R}_t$ represent respectively percent deviations of $(\Lambda_t P_t)$, C_t, Y_t, and $1+R_t$ from their steady-state level, $\hat{G}_t \equiv \left(G_t - \bar{G}\right)/\bar{Y}$, $\pi_t \equiv \log\left(P_t/P_{t-1}\right)$, and $\bar{C}_t \equiv \frac{u_{c\xi}}{u_c\sigma}\xi_t$ represents exogenous shifts in marginal utility of consumption.[22] In the absence of habit-formation, the coefficient $\sigma \equiv -u_{cc}\bar{C}/u_c > 0$ would represent the inverse of the intertemporal elasticity of substitution (EIS) of consumption evaluated at the steady-state.

Equations (7) – (9) form our "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

$$\hat{\lambda}_t = \hat{r}_t^L \equiv \sum_{T=t}^{\infty} E_t\left(\hat{R}_T - \pi_{T+1}\right), \tag{10}$$

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 (9), and recalling that $E_{t-2}\hat{C}_t = \hat{C}_t$, we obtain

$$\hat{Y}_t = \eta\hat{Y}_{t-1} + \beta\eta E_{t-2}\hat{Y}_{t+1} - \psi E_{t-2}\hat{r}_t^L + g_t, \tag{11}$$

where

$$
\begin{aligned}
\psi &\equiv \frac{(1-\beta\tilde{\eta})}{\sigma\left(1+\beta\tilde{\eta}^2\right)}\frac{\bar{Y}}{\bar{C}} > 0\\
0 &\leq \eta \equiv \frac{\tilde{\eta}}{\left(1+\beta\tilde{\eta}^2\right)} \leq (1+\beta)^{-1},
\end{aligned}
$$

[22] 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 $\hat{R}_t$.

and where g_t is an exogenous real demand disturbance that depends on past, present, and expected future values of $\hat{G}_t$ and $\bar{C}_t$. Note that because $\hat{G}_t$ is determined at $t-1$, both g_t and $\hat{Y}_t$ are determined at date $t-1$. Since it is difficult to interpret the parameter σ in the presence of habit formation, we prefer to focus on a pseudo-EIS, ψ, 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.[23] In the absence of habit formation, (11) reduces to the familiar equation $\hat{Y}_t = E_{t-2}\left(-\sigma^{-1}\bar{Y}/\bar{C}\hat{r}_t^L + g_t\right)$.

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 disturbances. 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. 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 again consistent with the assumption made in the structural VAR to identify monetary policy shocks. Following Christiano, Eichenbaum and Evans (2001), and Woodford (2002, ch. 3), we assume that if a price is not re-optimized, it is indexed to lagged inflation according to the rule

$$\log p_t(z) = \log p_{t-1}(z) + \gamma\pi_{t-1},$$

for some $0 \leq \gamma \leq 1$. Since every supplier faces the same demand function given by (6), all suppliers allowed to change their price in period t will choose the same price p_t^* that

[23] This can be seen by taking first differences of equation (11), and combining with (8).

maximizes

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

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, 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 this problem, solving for $\hat{p}_t^* \equiv \log(p_t^*/P_t)$, and quasi-differentiating the resulting expression yields the optimal pricing decision

$$\hat{p}_t^* = \alpha\beta E_{t-1}\hat{p}_{t+1}^* + \frac{1-\alpha\beta}{1+\omega\theta}E_{t-1}\left(\omega\hat{Y}_t - \hat{\lambda}_t + \frac{v_{y\xi}}{v_y}\xi_t\right) + \alpha\beta E_{t-1}\left(\pi_{t+1}-\gamma\pi_t\right), \qquad (12)$$

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. It can be shown that $\omega\hat{Y}_t - \hat{\lambda}_t + \frac{v_{y\xi}}{v_y}\xi_t$ is proportional to a measure of the output gap corresponding to the deviation of equilibrium output from its natural rate, i.e., the level that would obtain in the case of perfectly flexible prices.

Assuming furthermore that the price-setters who are allowed to change their price are chosen independently of their history of price changes implies that the price index satisfies

$$P_t = \left\{(1-\alpha)\,p_t^{*1-\theta} + \alpha\left[P_{t-1}\left(P_{t-1}/P_{t-2}\right)^{\gamma}\right]^{1-\theta}\right\}^{1/(1-\theta)}.$$

Log-linearizing this law of motion for P_t, and combining the resulting expression with (12) yields the following variant of the new-Keynesian aggregate supply equation

$$\pi_t - \gamma\pi_{t-1} = \kappa E_{t-1}\left(\omega\hat{Y}_t - \hat{\lambda}_t\right) + \beta E_{t-1}\left(\pi_{t+1}-\gamma\pi_t\right) + s_t, \qquad (13)$$

where $s_t \equiv \frac{v_{y\xi}}{v_y \kappa} E_{t-1}\xi_t$ is an aggregate supply shock — determined at date $t-1$ — that measures exogenous shifts in the disutility of producing output, and where the slope of the aggregate supply equation, $\kappa \equiv \frac{(1-\alpha)(1-\alpha\beta)}{(1+\omega\theta)\alpha}$, depends on the degree of price stickiness. As in the basic 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 $t-1$.

While the aggregate supply equation (13) is very stylized, it is important to realize that it nests as special cases some popular models that have very different implications, in particular regarding the degree of persistence in inflation. Except for the fact that pricing decisions at date t are taken here on the basis information available at date $t-1$, (13) reduces to the basic New Keynesian aggregate supply equation with Calvo pricing when $\gamma = 0$ and $\eta = 0$. Alternatively, when $\gamma = 1$, $\eta = 0$, and $\beta = 1$, (13) corresponds to the aggregate supply equation of Fuhrer and Moore (1995). While we assume that prices which are not re-optimized are adjusted to lagged inflation, we could have derived an equation almost identical to (13) — but with different restrictions on the model parameters — by assuming instead that some sellers are not rational and that they set their prices according to a simple rule of thumb, as in Galí and Gertler (1999).

Finally, the model is closed by a description of the central bank behavior. To the extent that the central bank is forward-looking, the coefficients of the VAR policy equation will subsume policy parameters — i.e., the parameters characterizing the Fed's systematic behavior

— as well as the remaining parameters needed to form the expectations, conditional on the time-t information set. To distinguish between changes in the private sector and policy behavior we need to specify a structural form of the reaction function. The forward-looking Taylor rule is one such policy reaction that is consistent with the reduced form policy of the VAR. It takes the form:

$$\hat{R}_t = \phi^{\pi} E_t \pi_{t+h_\pi} + \phi^{y} E_t \hat{Y}_{t+h_y} + \rho_1 \hat{R}_{t-1} + \rho_2 \hat{R}_{t-2} + \varepsilon_t, \tag{14}$$

where ε_t is an unforecastable random variable that represents monetary policy shocks. For the horizons $h_\pi = 0$ and $h_y = 0$, equation (14) corresponds to the popular rule proposed by Taylor (1993), augmented by lags of the Fed funds rate.[24] As another special case, the baseline case considered by Clarida, Galí and Gertler (2000) obtains when $h_\pi = 1$ and $h_y = 1$.[25]

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" (10) – (11), the aggregate supply equation (13), and an interest-rate feedback rule of the form (14). The resulting system of linear difference equations can then be solved using standard methods (e.g., King

[24] 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)). The specification that we consider involves two lags of the interest rate. This turns out to be the most parsimonious specification that is not rejected by the test of overidentifying restrictions.

[25] These authors estimate such a rule by GMM, in the single equation framework, assuming rational expectations on the part of the central bank. In contrast, we estimate this equation together with the rest of our structural model as described in the next sub-section.

and Watson (1998), McCallum (1998)).

For some parameter configurations, the model may result in an indeterminate equilibrium.[26] This may arise when the policy reaction function involves too little a response to changes in economic conditions, as Clarida *et al.* (2000) argue might have been the case for the pre-Volcker period.[27] This is in fact one dimension along which monetary policy could have become more effective and that we wish to investigate. But allowing parameters configuration to lie in the indeterminate region raises some difficulty: whenever this is the case, one equilibrium must be selected and there is unfortunately no natural criterion to select a particular one. Our strategy is to select the minimum-state-variable solution advocated by McCallum (1983), i.e., the single bubble-free solution. This solution corresponds to a situation where the economy could be subject to sunspots fluctuations, but there happens to be no such shock. While we recognize that the criterion that we adopt to select a solution may not be the only one, we find it advantageous, in particular when compared to the alternative of ruling out *a priori* the possibility of indeterminacy. For instance, it allows us

[26] 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 $\{\varepsilon_t, g_t, s_t\}$. 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.

[27] See, e.g., Woodford (2002, ch. 2 and 4) for a discussion of the problem of indeterminacy of the equilibrium in monetary models of the kind analyzed here.

to implement the estimation strategy described in the next section, even when parameter configurations yield an indeterminate equilibrium.[28]

3.2 Estimation of the structural model

3.2.1 Minimum Distance estimation of the structural parameters

We now turn to the estimation of the structural model just described.[29] In section 2, we estimated a structural VAR, that allowed us to generate impulse response functions to monetary policy innovations. In the previous 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)).[30]

More formally, we consider the vector of structural parameters for Sample s, Δ_s, the

[28] Lubik and Schorfheide (2002) propose an alternative approach that allows for the possibility of multiple equilibria in the estimation of DSGE models. However, unlike the estimation strategy discussed below, their Bayesian approach requires a complete specification of the shock processes and prior distributions.

[29] A similar estimation procedure can be found in Rotemberg and Woodford (1997), Amato and Laubach (2003), Gilchrist and Williams (2000) and Christiano, Eichenbaum, and Evans (2001).

[30] Our estimation method is also similar in spirit to the specification test used by Cogley and Nason (1995), although it was based on matching autocorrelation functions and they were not concerned with the estimation of the structural parameters.

vector Ω_s containing the identified VAR coefficients and $G_V(\Omega_s)$, 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 $G_M(\Delta_s)$ the corresponding vector-valued function that collects the *model-based* impulse response functions yielded by its rational expectations solution. Let $G(\Omega_s, \Delta_s) \equiv G_M(\Delta_s) - G_V(\Omega_s)$. Having estimated Ω_s, using the technique described in Section 2, we minimize

$$L(\Delta_s) = G\left(\hat{\Omega}_s, \Delta_s\right)' W_s G\left(\hat{\Omega}_s, \Delta_s\right) \tag{15}$$

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.

Given the goal of our investigation, we find this estimation strategy appealing 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 samples considered, it is 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 potential efficiency gain has to be weighted against the cost of additional identifying assumptions that would be required. 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 identification of other shocks and to the specification of parts of the model that are not related to the impulse response functions we are interested in.[31] Specifically,

[31] The robustness of this estimation approach to a misspecification of the theoretical model is discussed

while the endogenous variables are affected by demand and supply shocks g_t and s_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 $\{s_t\}$. Finally, as Hall (2001) pointed out, estimation through impulse response functions has an important advantage over the application of GMM to Euler equations: it indirectly imposes the model's boundary conditions.[32]

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), (13), and the policy reaction function (14). We need to quantify a total of ten structural parameters: $\{\phi^{\pi}, \phi^{y}, \rho_1, \rho_2, \beta, \psi, \kappa, \omega, \eta, \gamma\}$. All of these parameters could in principle be separately identified from the impulse response functions to a monetary policy shock. However, in order to reduce the dimension of the estimation, 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. We thus attempt to estimate the remaining nine parameters $\Delta = \{\phi^{\pi}, \phi^{y}, \rho_1, \rho_2, \psi, \kappa, \omega, \eta, \gamma\}$ by matching the model-based impulse response functions with those of the VAR, subject to the

more generally in Dridi and Renault (2001).

[32] 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.

model constraints on the sign and magnitude of the parameters[33]. We consider the responses of the variables over the first twelve 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.

To estimate the structural parameters, we also need to determine an asymptotically non-stochastic weighting matrix W_s indicated in (15). To account for the fact that some points of the impulse response functions are less precisely estimated than others, we use a diagonal weighting matrix that involves the inverse of each impulse response's variance on the main diagonal.[34]

3.2.2 Estimation of the forecasting horizon

The estimation of the policy reaction function 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

[33] In turns out however that for the second sample, ω is not well identified which renders the minimization problem ill-behaved. As a result, we decided to calibrate to $\omega = .47$, a value close to the one estimated in the first sample and the one used by Rotemberg and Woodford (1997).

[34] Ideally, one might want to use the complete variance-covariance matrix of the impulse response functions. But this approach relies on many more estimated elements, which seems to adversely affect the numerical stability of our minimization problem.

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.

As the forward-looking rule (14) 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 minimizing 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 (14).

Table 1 reports the p-values of the J-test for all combinations of h_π and h_y up to 4 quarters, and for the two samples. The best horizon, i.e., the one with the highest $p-$value, is $h_\pi = 3$ and $h_y = 0$, for the 1959:1-1979:3 period and $h_\pi = 2$, $h_y = 0$, respectively for the 1979:4-2002:2 period. For these horizons, the forward-looking Taylor rule specification is not rejected at the 5% level. Noting that a 95% confidence interval for these horizons would include any combination (h_π, h_y) with a p-value larger than 5%, the table suggest that within the set of horizons less than or equal to four, the horizons are fairly precisely estimated.

3.3 Estimation results

Table 2 reports the structural parameters' estimates, along with the associated standard deviations for both samples. Looking first at the parameters describing the behavior of the private sector, the main differences between the two samples are in terms of the sensitivity of output to the long-run interest rate in the IS curve, ψ, and the slope of the Phillips curve,

κ: from Sample 1 to Sample 2, ψ increases from 0.89 to 1.07 and κ from 0.008 to 0.024. On the other hand, the degree of habit formation and inflation inertia are similar in both samples, and remain close, or on the theoretical upper limits imposed in the estimation. This evidence suggests: 1) that detrended output has become more sensitive to the real rate of return and 2) that inflation responds more quickly to change in detrended output, consistent with a possible increase in price flexibility (i.e., a decline in the probability α). This implies, everything else equal, that changes in the instrument of monetary policy should have had a *stronger* effect on output after 1980.

It is difficult to provide justification for changes in certain "deep" parameters, such as those of the utility function embedded in ψ. Although we doubt that the intertemporal elasticity of substitution has changed, we view instead these estimates as capturing the fact that the private sector of the economy has reacted more strongly to changes in economic conditions in the post-1980 sample than in the pre-1980 sample. Moreover, rather than ruling out changes in the private sector parameters *a priori*, in the next section we account for these changes and determine whether or not they affect our conclusions about the effectiveness of monetary policy. We find that they do not.

Turning to the policy parameters, the main result is that the response to inflation and output is considerably higher in the second sample. The response to inflation is 60% larger in the second period and that of output, although still quite small and insignificant, is more than six times larger. Monetary policy is thus more responsive since the early 1980's to the economic environment, a result consistent with the evidence of Clarida *et al.* (2000), Boivin (2001) and Cogley and Sargent (2002), among others. We find it interesting that we obtain qualitative results similar to these studies, using a very different estimation strategy that relies solely on impulse response functions.

An important implication is that in the first sample, the configuration of estimated parameters yields an indeterminate rational expectation equilibrium, as in Clarida *et al.* (2000). This is not the case in the second sample. There is a strong presumption that this is due to the weakness of the central bank response in the first sample, thus implying that monetary policy became more effective along that dimension.[35] However, this cannot be asserted definitely without taking into account the changes in the other structural parameters. We investigate this issue in the counterfactual experiments of the next section.

Figure 2 plots 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 estimated structural model (solid lines), for both samples. 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.[36] 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.

[35] In fact, by plotting the number of unstable eigenvalues of the dynamic system characterizing our estimated model in the (ϕ^{π}, ϕ^{y}) space (not reported here), we observe that the equilibrium would be determinate for values of ϕ^{π} above 0.33 and values of ϕ^{y} above 0.04.

[36] The only exception is the response of inflation one period after the shock in Sample 1.

4 Model-based counterfactual analysis

Having argued that our model replicates reasonably well the effects of monetary shocks on output, inflation, and the interest rate in both samples, we can now investigate whether monetary policy has become more effective. Through counterfactual experiments, we first determine whether monetary policy or the behavior of the private sector has been instrumental in removing indeterminacies of the rational expectations equilibrium in the post-1980 period. We then seek to determine to what extent the reduced effect of monetary policy shocks in the post-80 sample is due to an improvement in monetary policy or a change in the private sector's behavior. Next, we will assess the relevance of monetary policy and the private sector to account for changes in the response to real demand and supply disturbances. In these counterfactual experiments, we denote by $\Delta_s^{MP} = \{\phi^\pi, \phi^y, \rho_1, \rho_2\}$ the set of monetary policy parameters, while $\Delta_s^{PS} = \{\psi, \kappa, \omega, \eta, \gamma\}$ contains the private sector parameters.

4.1 Indeterminacy

In the previous section we documented the presence of an indeterminate equilibrium in the pre-1979:4 period. To determine what is the cause of indeterminacy, Table 3 reports for various combinations of the estimated private sector (Δ_s^{PS}) and monetary policy parameters (Δ_s^{MP}) whether or not they involve a unique rational expectations equilibrium. The only results we want to emphasize from this Table is that monetary policy is indeed the source of indeterminacy in the first sample. In fact, if the pre-1979:4 conduct of monetary policy had been maintained over the second sample, an indeterminate equilibrium would have resulted.

In contrast, if the recent conduct of monetary policy would have ruled out non-fundamental fluctuations in the first sample. This is thus clearly a dimension along which monetary policy has become more effective.

4.2 Equilibrium responses to exogenous disturbances

Figure 3 displays the impulse response functions generated by the model to an innovation — a unit increase — in each of the three exogenous disturbances ε_t, g_t, and s_t. Each panel of Figure 3 contains impulse responses for the four possible combinations of monetary policy (Δ_s^{MP}) and private sector (Δ_s^{PS}) parameters, where $s = 1$ for the 1959:1–1979:3 period and $s = 2$ for the 1979:4–2002:2 period.

The response to the monetary shocks (in the first column) clearly shows that the observed change in the monetary transmission mechanism is due to a change in the conduct of monetary policy itself. A comparison of the two sets of responses generated by $(\Delta_1^{MP}, \Delta_1^{PS})$ and $(\Delta_1^{MP}, \Delta_2^{PS})$ — i.e. maintaining monetary policy as estimated in the first sample — reveals that, if anything, the change in the private sector implies a *larger* response of output, not smaller one. While inflation reacts initially more strongly with the post-80 parameters of the private sector, its response is also less persistent. This is consistent with the fact that we have estimated larger values of ψ and κ the second sample. By comparing the corresponding impulse responses $(\Delta_2^{MP}, \Delta_1^{PS})$ and $(\Delta_2^{MP}, \Delta_2^{PS})$ — i.e. maintaining monetary policy as estimated in the second sample — we note that the change in the structural parameters has almost no effect on the impulse response functions.

The striking result, however is that the observed reduction in the *magnitude* of the impulse responses is almost entirely attributable to monetary policy. In fact, by changing monetary policy and maintaining the structural parameters fixed — i.e., by comparing the lines $(\Delta_1^{MP}, \Delta_1^{PS})$ to $(\Delta_2^{MP}, \Delta_1^{PS})$, and $(\Delta_1^{MP}, \Delta_2^{PS})$ to $(\Delta_2^{MP}, \Delta_2^{PS})$ — we observe that the responses of output and inflation associated with the policy estimated for the second sample involve considerably less variation than those associated with the policy of Sample 1. By maintaining the structural parameters constant at Δ_1^{PS}, we observe that a change in policy from Δ_1^{MP} to Δ_2^{MP} almost entirely explains the impulse responses $(\Delta_2^{MP}, \Delta_2^{PS})$ obtained in the second period. This counterfactual experiment thus suggests that the change in the estimated impulse responses to a monetary shock, first reported in Figure 1, is attributable almost entirely to a change in the systematic conduct of monetary policy. The fact that the response of output and inflation has become considerably smaller in the post-80 period thus does not appear to reflect a diminished effect of monetary policy on these variables. Our analysis suggests that it is rather the fact that monetary policy has become more aggressive — by reacting more strongly to fluctuations in expected inflation and to a lesser degree to fluctuations in expected output — that has helped stabilize the economy in response to monetary shocks.

The second column of Figure 3, which plots the counterfactual responses to a real demand disturbance — i.e., an innovations in g_{t+1} — conveys a similar message.[37] While the output response is slightly smaller with the private sector parameters of Sample 2 — comparing $(\Delta_1^{MP}, \Delta_1^{PS})$ and $(\Delta_1^{MP}, \Delta_2^{PS})$ — most of the change in the impulse response of output

[37] Note that because g_{t+1} is predetermined, as discussed in Section 3, an unexpected demand disturbance at date t does not affect output before date $t+1$.

between Sample 1 and Sample 2 is explained by a change in monetary policy from Δ_1^{MP} to Δ_2^{MP}. In addition, although the change in the private sector parameters from Δ_1^{PS} to Δ_2^{PS} appears to have rendered the inflation response less persistent, the line $(\Delta_2^{MP}, \Delta_1^{PS})$ indicates that the reduction in the magnitude of the inflation response is largely due to the change in monetary policy.

The last column of Figure 3 displays impulse response functions to an adverse unanticipated supply shock s_t that raises inflation by one percent in period 1.[38] In contrast to the counterfactual experiments with the two previous shocks, the responses to a supply shock yield a less clear picture regarding the effectiveness of monetary policy. The inflation responses suggest that both a change in the private sector and a change in the monetary policy can account for the smaller and less persistent response of inflation in the second sample. Although the shock's effect on output is relatively small, the change in monetary policy is responsible for a change in the sign of the output response. The expected decline in output following the adverse shock (with monetary policy Δ_2^{MP}) is intuitive. In contrast, the increase in output under the policy Δ_1^{MP} may be surprising: it relates to the fact that the real rate of interest (not displayed) can decrease substantially following the shock, as the policy of Sample 1 responds relatively little to the higher expected inflation.

Because the evidence about monetary effectiveness is less clear in the face of aggregate supply shocks, we finally turn to a simulation of the model under alternative policies, using estimated processes for the real shock processes $\{g_t\}$ and $\{s_t\}$.

[38] Note that s_t is by construction predetermined, so that an unexpected supply disturbance at date t does not affect inflation before date $t+1$.

4.3 Counterfactual simulations

In order to perform counterfactual simulations of the model, we need to estimate the stochastic processes of the three disturbances $\{\varepsilon_t, g_t, s_t\}$. Combining again our identified VAR of Section 2 with our structural model, we can extract a time series for the vector of exogenous variables $x_t \equiv [\varepsilon_t, g_{t+1}, s_{t+1}]'$, along the lines of Rotemberg and Woodford (1997). First, we rewrite our structural VAR in companion form as

$$\bar{Z}_t = B\bar{Z}_{t-1} + \bar{u}_t \tag{16}$$

where $\bar{Z}_t$ is a vector containing all the variables of the VAR and their lags, and $\bar{u}_t$ is an unforecastable vector. Second, using the structural equations (10) – (11), (13), and (14), and the estimated parameters, we can express the shocks as a function of past, present and expected future values of output, inflation and the interest rate. It follows that the vector of exogenous variables x_t can be expressed as a function of present and expected future values of a vector $\tilde{Z}_t$ which contains the theoretical variables corresponding to those in $\bar{Z}_t$. Third, assuming that expectations of future variables in the model correspond to the VAR forecasts, so that $E_t\tilde{Z}_{t+j} = E_t\bar{Z}_{t+j} = B^j\bar{Z}_t$ for all $j > 0$,[39] we can express x_t as

$$x_t = C\bar{Z}_{t-1} + D\bar{u}_t \tag{17}$$

for some matrices C and D. This can then be used to generate a historical time series for x_t. Using (16) – (17) together with the structural equations of the model reproduces exactly the historical time series of all variables.

[39] It is important that both vectors $\bar{Z}_t$ and $\tilde{Z}_t$ contain all relevant variables which are part of the information set at date t (such as $\hat{Y}_{t+1}$ and π_{t+1}), for the previous equality to hold.

We model the dynamics of the constructed time series of g_t and s_t by fitting an AR(p) process on each shock. We set the policy shock equal to 0 at all dates in order to isolate the effect of systematic monetary policy in the counterfactual simulations. Our simulation model is then comprised of the estimated structural equations (10) – (11), (13), (14), the law of motion for the disturbances given by the fitted AR(p) processes for g_t and s_t, as well as a distribution for the innovations.

Table 4 contains the results of the counterfactual simulations for alternative policy rules and structural parameters, using a given shock process. Specifically, the upper part of the table reports counterfactual variances of output, inflation and the interest rate for alternative combinations of the monetary policy parameters and the private sector parameters, using the shock processes for g_t and s_t estimated in Sample 1, and assuming a zero monetary policy shock at all dates. The variances reported refer to sample variances and are expressed relative to the respective variances simulated in Sample 1.[40] The lower part of Table 4 reports the results of the same calculations, in the case that the shock processes are the ones estimated in Sample 2.

One interesting fact revealed by Table 4 is that for any variable considered, the variances, V, satisfy except for one case:

$$\begin{aligned} V\left(\Delta_2^{MP},\Delta_2^{PS}\right) &< V\left(\Delta_2^{MP},\Delta_1^{PS}\right) < V\left(\Delta_1^{MP},\Delta_1^{PS}\right) \\ V\left(\Delta_2^{MP},\Delta_2^{PS}\right) &< V\left(\Delta_2^{MP},\Delta_1^{PS}\right) < V\left(\Delta_1^{MP},\Delta_2^{PS}\right). \end{aligned}$$

Taking together the inequalities $V\left(\Delta_2^{MP},\Delta_1^{PS}\right) < V\left(\Delta_1^{MP},\Delta_1^{PS}\right)$ and $V\left(\Delta_2^{MP},\Delta_2^{PS}\right) < V\left(\Delta_1^{MP},\Delta_2^{PS}\right)$ confirm that the more responsive monetary policy of Sample 2 results in a

[40] Asymptotic variances yield similar qualitative results.

smaller variability of output, inflation and the interest rate, regardless of the set of structural parameters Δ_1^{PS} or Δ_2^{PS}, and the shock process considered. The only case that reverses the inequality $V\left(\Delta_2^{MP}, \Delta_1^{PS}\right) < V\left(\Delta_1^{MP}, \Delta_1^{PS}\right)$ refers to the variance of output in Sample 2.[41]

Combining the inequalities $V\left(\Delta_2^{MP}, \Delta_2^{PS}\right) < V\left(\Delta_2^{MP}, \Delta_1^{PS}\right)$ reveals that for given policy of the post-80 period, a change in the structural parameters from Δ_1^{PS} to Δ_2^{PS} further decreases the economy's volatility. However, since the variances are always lower when we change only policy from Δ_1^{MP} to Δ_2^{MP} than we change only the structural parameters from Δ_1^{PS} to Δ_2^{PS}, we conclude that the reduced variance is mainly due to the more responsive monetary policy. In addition, in the case of the pre-80 policy, Δ_1^{MP}, a similar change in the structural parameters is in most cases associated with an increase in the economy's volatility.

Overall, these experiments suggest that the change in monetary policy has been instrumental in reducing the economy's variability in the post-80 period. In particular, we find that the change in monetary policy has contributed more to a reduction in output, inflation, and interest rate variability than a change in the behavior of the private sector. Note finally that for the policy Δ_2^{MP}, which is more responsive to fluctuations in expected inflation and output than Δ_1^{MP}, changes in the structural parameters have relatively little effect on the impulse response functions and on the variances reported in Table 4. In contrast, the changes in the parameters of the private sector exert a larger effect on the impulse responses and the

[41] This reveals that in the case of the shock processes estimated in Sample 2, and the structural parameters are maintained at their pre-80 level, a change in monetary policy from Δ_1^{MP} to Δ_2^{MP} would increase modestly the variance of output. This slight increase in output volatility is however accompanied by a much larger decline in inflation variability — the absolute level of the variance on inflation being larger than the variance of output.

variances when the less responsive policy Δ_1^{MP} 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^{MP}, a more aggressive policy makes it more likely for the variances of output, inflation and the interest rate to be contained.

5 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 inflation since the beginning of the 1980's. In this paper, we attempt to determine the causes of this phenomenon and the implications for the effectiveness of monetary policy. We consider three dimensions of monetary policy effectiveness: 1) its ability to stabilize the effects of shocks on the economy; 2) its success in eliminating non-fundamental sources of fluctuations, and 3) the extent to which it manages to reduce the amount of randomness in the setting of its policy.

To investigate these questions, we adopt a two-step strategy. First, using a VAR, we identify changes in the monetary transmission mechanism based on a limited amount of identifying assumptions. Second, we develop a fully specified model of the U.S. economy that can replicate well the responses obtained from the VAR. Based on the estimated values of the structural parameters, estimated separately over the pre- and post-1980 periods, we perform counterfactual experiments.

We find that the dominant cause behind the alteration of the monetary transmission mechanism is a change in the conduct of monetary policy, characterized by a stronger response to inflation and output since the early 1980's. The main finding is that monetary policy has become more effective at stabilizing the economy. First, the current conduct of monetary policy prevents potential non-fundamental forces to affect the economy. Moreover, our estimated structural model implies that monetary policy response to monetary policy and demand shocks has become more successful. When all shocks, including the supply shock, are considered simultaneously, we find through counterfactual simulations, that changes in monetary policy have been instrumental in stabilizing the economy. Finally, we argue that changes in the variance of policy shocks have played a negligible role. Taken together, our results suggest that monetary has indeed become more effective.

A Robustness analysis of VAR results

In this Appendix we investigate the robustness of the VAR findings reported in Section 2, to the inclusion of more information in the VAR and to the choice of an alternative starting date for the second sample, namely 1984:1.

The specification used in this paper was favored on the ground that it contained the minimum set of variables necessary for our investigation, while delivering sensible responses of the economy consistent with existing results. As is commonly done in this literature, the commodity price index was included to alleviate the so-called price-puzzle. Yet, our VAR does contain a limited amount information and this potential misspecification could contaminate our empirical results. Bernanke and Boivin (2003) and Bernanke, Boivin and Eliasz (2002) propose a way of incorporating more information in low dimensional VARs. More specifically, building on recent development on the estimation dynamic factor models with large panels[42], their strategy is to expand VAR systems with a few factors estimated from a large panel of macroeconomic series. We follow the exact same factor-augmented VAR (FAVAR) strategy here, expanding a VAR in $\{\hat{Y}_t, \pi_t, R_t\}$ with the first factor estimated from the panel of macroeconomic series used in Bernanke, Boivin and Eliasz (2002).[43] Note that if this factor properly accounts for the existing information, there is no justification to have the commodity price index in the VAR, which we thus exclude.[44] It is important to note

[42] See Stock and Watson (2002), Forni, Lippi and Reichlin (1999) among others.

[43] See Bernanke, Boivin and Eliasz (2002) for details on the implementation, the data set and the identification strategy – consistent with our VAR – in this framework.

[44] In fact, adding the commodity price index on top of the factor does not affect the result. These results are not reported.

that if the baseline VAR is properly specified, the inclusion of additional information should not affect the results.

We estimate the impulse response functions for the FAVAR model for the two samples considered in the text, as well as a third one corresponding to the alternative break date, that is the post-84:1 period. The results are reported in Figure 4 together with those obtained from the VAR specification used in the paper. The confidence intervals displayed are those obtained from the baseline VAR.

Looking first at the results for the baseline VAR model – the -+- line in the figure – we find that the finding of a reduced effect of monetary policy shock in the post 1980 period is robust to the alternative break date considered. But, there are still notable differences between the post-79:4 and post-84:1 samples. In particular, the response of inflation appears somewhat stronger when the VAR is estimated on the latter, and the response of output 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 post-84:1 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. This might suggest misspecification of our baseline VAR for the post-84:1, but given the much larger confidence interval, this could also be due to the imprecision of the estimation on this shorter sample.

Turning now to the results obtained from the FAVAR model, two key conclusions emerge. First, for the pre- and post 79:4 periods – the first two columns in the figure – the results are essentially the same as those obtained with the baseline specification. There is a somewhat stronger response of output in the pre-79 period that goes outside the confidence intervals of the baseline VAR; but this would reinforce our finding that there has been an important reduction in the effect of monetary policy shocks on output. The second conclusion is that the results of the baseline VAR for the post-84:1 period are not robust to the inclusion of the additional information. More strikingly, the inclusion of the factor has the effect of reversing the sign of the response of output, thus becoming consistent with conventional wisdom and with the results from the other samples. This suggests that more information was used in the conduct of monetary policy in the post-84 period, which needs to be accounted for to properly identify the impulse response functions.

Overall, this robustness analysis suggests that the pre- and post 79:4 comparison undertaken in the paper is justified, as the inclusion of more information does not affect the VAR conclusions for these two samples and makes the conclusions obtained from the post-84 period broadly consistent with those of the post-79:4 period.

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

	h_y				
h_π	0	1	2	3	4
			1959:1–1979:2		
0	0.001	0.007	0.013	0.009	0.003
1	0.000	0.000	0.000	0.000	0.000
2	0.012	0.029	0.021	0.023	0.027
3	0.053	0.015	0.003	0.005	0.003
4	0.011	0.007	0.001	0.001	0.000
			1979:3–2002:2		
0	0.009	0.017	0.022	0.024	0.030
1	0.013	0.003	0.003	0.002	0.005
2	0.307	0.278	0.250	0.307	0.258
3	0.005	0.002	0.003	0.009	0.009
4	0.000	0.000	0.000	0.000	0.000

Note: The table report 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 percent level. See text for details.

Table 2: **Estimates of structural parameters**

Parameters	Sample 1	Sample 2
ψ	0.894	1.073
	(0.104)	(0.184)
η	0.502	0.487
	(—)	(.004)
κ	0.008	0.023
	(0.000)	(0.005)
ω	0.474	.470
	(0.419)	(—)
γ	1	1
	(—)	(—)
ϕ^{π}	0.307	0.503
	(0.009)	(0.019)
ϕ^{y}	0.006	0.001
	(0.004)	(0.017)
ρ_1	0.946	0.589
	(0.021)	(0.015)
ρ_2	-0.250	-0.070
	(0.019)	(0.019)

Note: Results based on the minimum distance estimation described in the text. Standard errors are in parentheses. (—) denotes that the standard error is not available, either because the parameter is calibrated or is hitting the parameter space boundary.

Table 3: **Indeterminacy in counterfactual experiments**

	Δ_1^{PS}	Δ_2^{PS}
Δ_1^{MP}	I	I
Δ_2^{MP}	D	D

Note: This table reports whether the structural model results in a determinate (D) equilibrium or an indeterminate (I) equilibrium for various combinations of policy rule coefficients $\left(\Delta_i^{MP}\right)$ and parameters of the private sector $\left(\Delta_i^{PS}\right)$, in samples $i = 1, 2$.

Table 4: **Normalized variances in counterfactual experiments**

Sample 1 shock process	$\text{var}\left(\hat{Y}\right)$	$\text{var}(\pi)$	$\text{var}\left(\hat{R}\right)$
$\left(\Delta_1^{MP}, \Delta_1^{PS}\right)$	1	1	1
$\left(\Delta_1^{MP}, \Delta_2^{PS}\right)$	2.595	12.388	15.907
$\left(\Delta_2^{MP}, \Delta_1^{PS}\right)$	0.765	0.032	0.026
$\left(\Delta_2^{MP}, \Delta_2^{PS}\right)$	0.514	0.320	0.024

Sample 2 shock process	$\text{var}\left(\hat{Y}\right)$	$\text{var}(\pi)$	$\text{var}\left(\hat{R}\right)$
$\left(\Delta_1^{MP}, \Delta_1^{PS}\right)$	1.313	5.624	4.307
$\left(\Delta_1^{MP}, \Delta_2^{PS}\right)$	5.447	3.209	3.990
$\left(\Delta_2^{MP}, \Delta_1^{PS}\right)$	1.433	1.890	1.879
$\left(\Delta_2^{MP}, \Delta_2^{PS}\right)$	1	1	1

Note: The table reports counterfactual sample variances for alternative combinations of the monetary policy parameters and the private sector parameters, relative to the variance simulated in the respective sample.

Figure 1: Impulse Response Functions to a Monetary Shock Over Different Samples

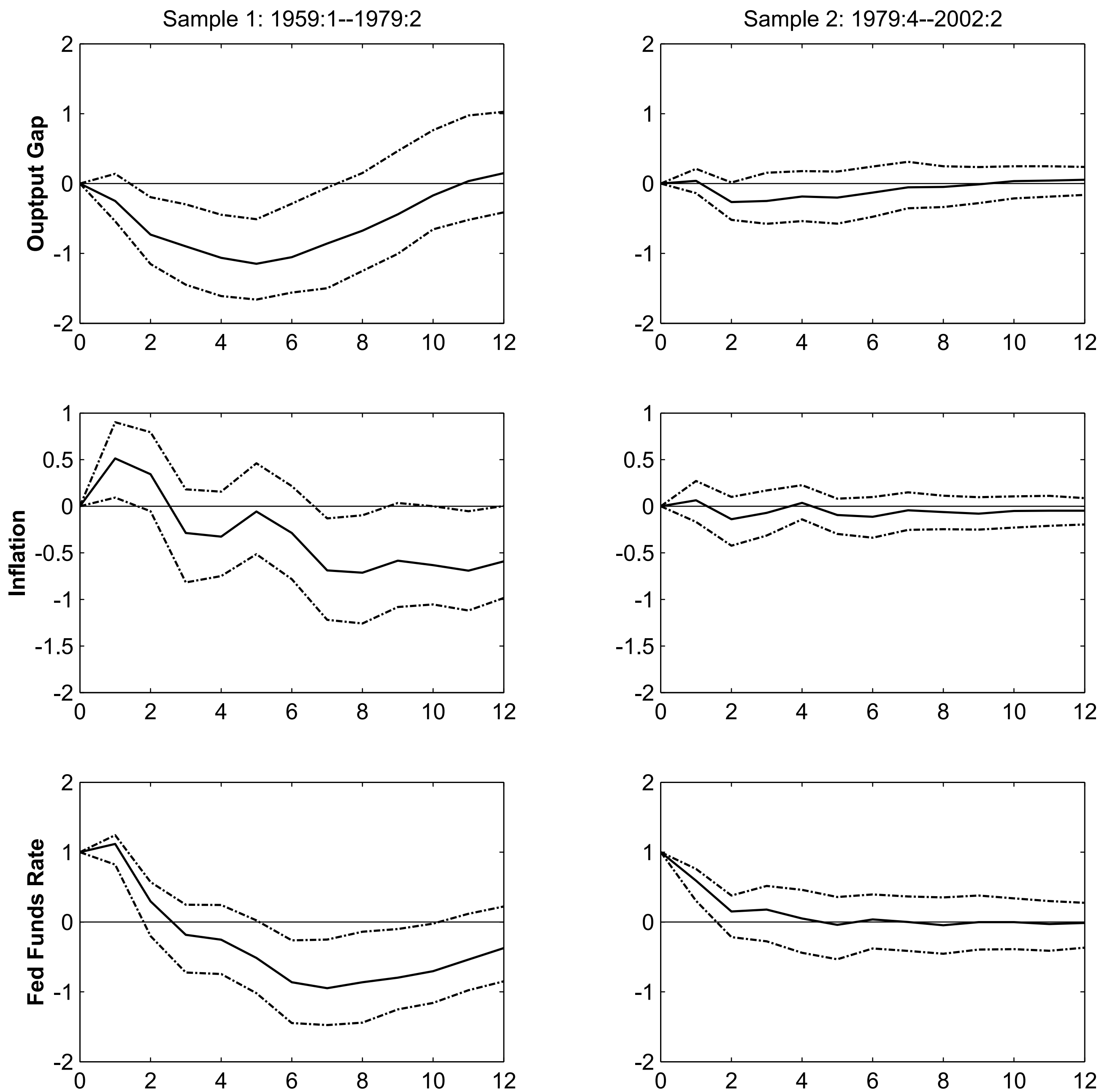

Figure 2: VAR and Model-Based Impulse Responses to a Monetary Shock

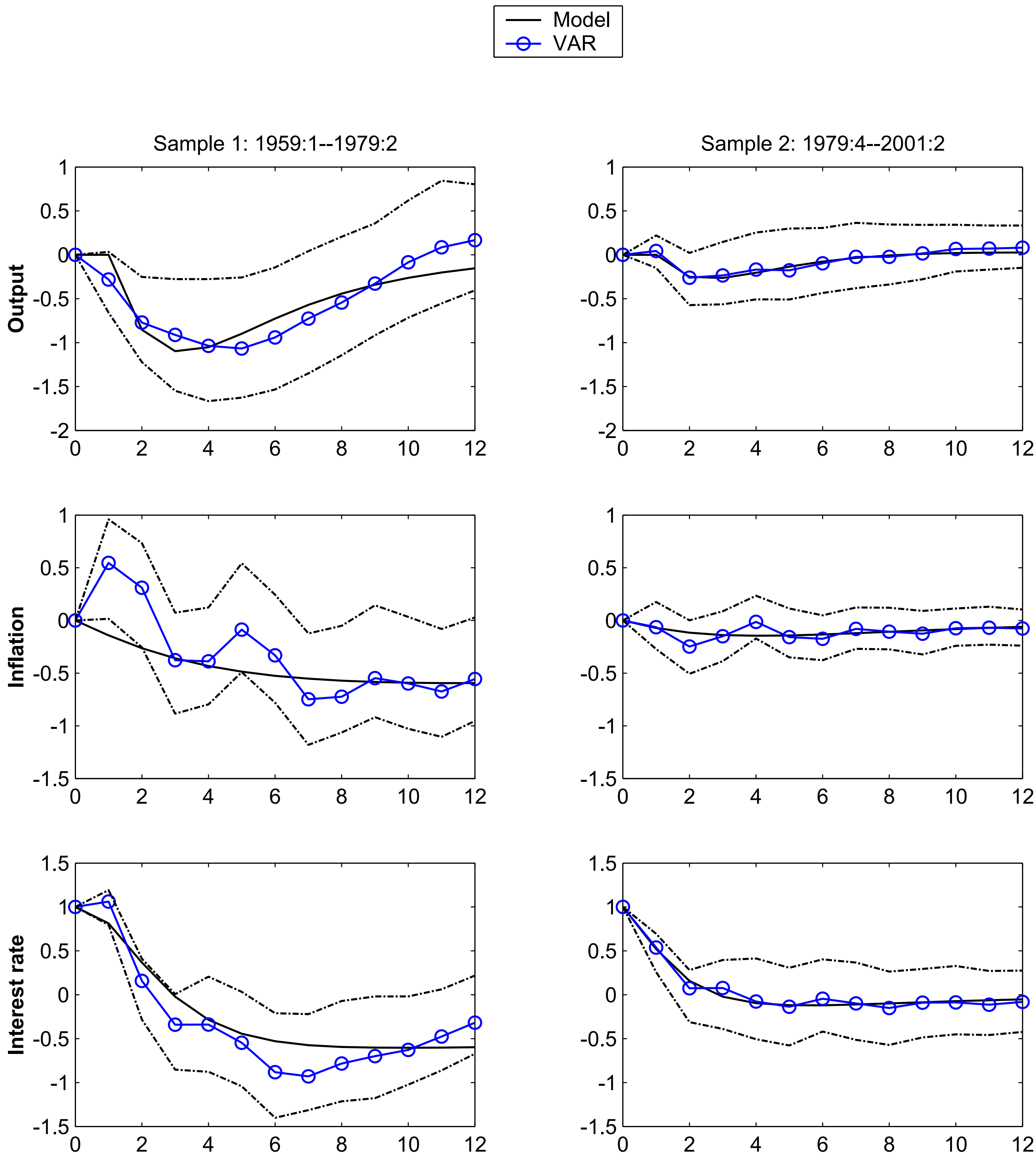

Figure 3. Model–Based Counterfactual Analysis

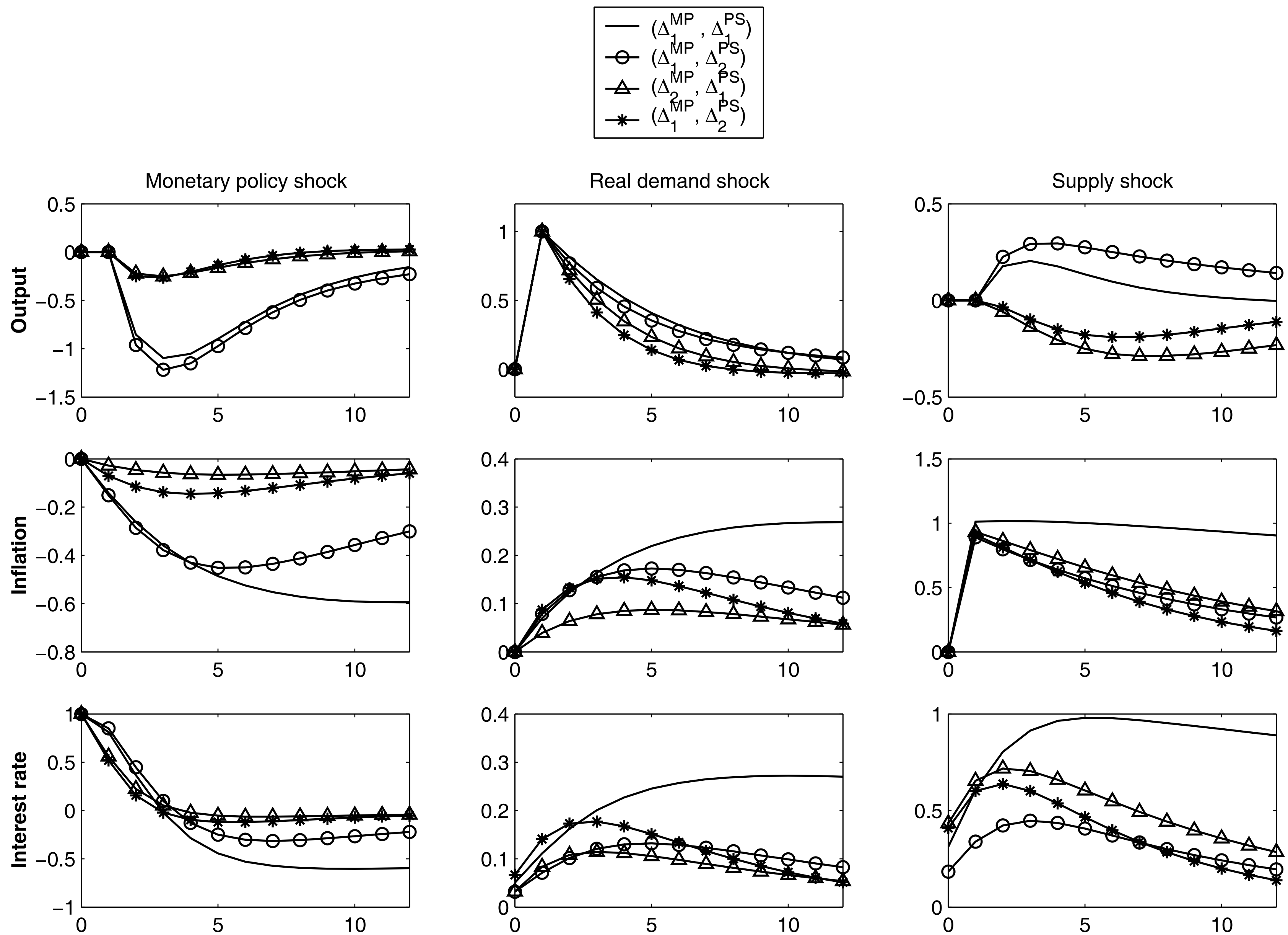

Figure 4: Impulse response functions over different samples

PCOM and Factor Augmented VAR