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European Economic Review 52 (2008) 487–497
EUROPEAN ECONOMIC REVIEW
www.elsevier.com/locate/eer

When did unsystematic monetary policy have an effect on inflation?

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Received 7 April 2006; accepted 14 February 2007
Available online 7 May 2007

Abstract

An important stylized fact to emerge from VAR estimates is that exogenous monetary policy shocks (also labelled unsystematic monetary policy) have a delayed, persistent, hump-shaped effect on inflation. I argue that this empirical pattern is fragile. In particular, it disappears when one examines periods without large shifts in the level of inflation (such as 1984–2005). An important consequence is that the hump-shaped VAR estimated response of inflation is not appropriate to fit stylized models of the response of inflation around a stable steady state inflation level.

JEL classification: E52

Keywords: Monetary policy transmission; VARs; Inflation dynamics

1. Introduction

One of the most widely accepted stylized facts of monetary economics is that US inflation has a hump-shaped response to exogenous monetary policy shocks. For instance, Christiano et al. (2005, pp. 5–8) state, "after an expansionary monetary policy shock [...] inflation responds in a hump-shaped fashion peaking after about two years." Likewise Mankiw (2001) writes, "According to the consensus view among central bankers and monetary economists, a contractionary monetary shock raises unemployment, at least temporarily, and leads to a delayed and gradual fall in inflation."

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0014-2921/$ - see front matter
doi:10.1016/j.euroecorev.2007.02.007

This stylized fact is supported by a series of robustness checks along several dimensions, though all within the VAR framework, reported in the second chapter of the 1999 Handbook of Macroeconomics: "Monetary Policy Shocks, What have we learned and to what end?" (also by Christiano et al., 1999a; thereafter CEE-99a) and in many contributions reviewed therein.

However, this VAR based characterization of the effects of unsystematic monetary policy[1] on inflation is very sensitive to the choice of the sample period. In particular, if one considers the last twenty years, VAR estimated monetary policy shocks have no effect on inflation nor the price level. The hump-shaped response of inflation is obtained only if either the building up or the collapse of the 1970 Great Inflation is included in the samples over which the VAR is estimated. One important implication is that models that are consistent with the evidence estimated over long sample periods may be mixing up the response of inflation to monetary policy shocks in periods of large adjustments of inflation, such as the so-called Volker disinflation, and periods when the mean of inflation is stable, e.g. from 1984 to 2005. There is therefore a risk that these models provide a poor approximation of inflation dynamics for both periods of large adjustments and periods when the mean of inflation is stable.

The paper proceeds as follows. Section 2 reviews the literature on VAR identifications of US monetary policy shocks. Section 3 focuses on the changes in the inflation impulse responses estimated on the 1984–2005 sample period relative to the ones obtained for the 1960–2005 period. Section 4 concludes.

2. VAR based identification of US monetary policy shocks

Twenty six years after the seminal contribution of Sims (1980), vector autoregressions (VARs) have become the most widely used econometric apparatus to describe stylized facts on the effects of structural, i.e., economically meaningful, shocks. In particular, the study of US monetary policy with VAR models has developed as a literature of its own. CEE-99a, which is the second chapter of the latest Handbook of Macroeconomics, is entirely dedicated to VAR based identification of US monetary policy shocks and estimates of their effects on US macroeconomic variables.[2] In their introduction, CEE-99a argue that US monetary policy shocks are "good candidates" to evaluate the ability of models to mimic actual economies. For instance, they showed in an earlier paper that limited participation models and sticky price models predict different paths for money and the interest rate following a monetary policy shock (Christiano et al., 1999b).

One remarkable result of CEE-99a is that most competing identification schemes of US monetary policy shocks deliver quite similar results in terms of their effects on output and prices. A monetary tightening triggers a hump-shaped response of the GDP log-level and a negative response of the price log-level that is gradual. This is true for both recursive identification and non-recursive models as initially put forth by Sims, as well as across models that differ in terms of the number of variables entering the VAR, and therefore the information set the central banks uses to set its monetary policy instrument (usually the interest rate on federal funds).

[1]Throughout the text I consider either monetary policy shocks, exogenous monetary policy shocks or unsystematic monetary policy as interchangeable terms.

[2]The other two most cited surveys are Leeper et al. (1996) and Bernanke and Mihov (1998).

The CEE-99a VAR model consists of the following vector of variables:

$$Y'_t = (dy_t, \pi_t, \pi_t^{CP}, i_t, dTR_t, dNBR_t, dM1_t),$$

which stand for d log(GDP), d log(CPI), d log(commodity prices), the interest rate on federal funds, d log(total reserves), d log(non-borrowed reserve) and d log($M1$).[3]

The monetary policy shocks and their effects on the variables of the model are estimated in three steps. First, the variables are regressed on their lag values to estimate the parameters of $A(L)$, the autoregressive form of the model:

$$Y_t = A(L)Y_{t-1} + u_t.$$

Second, the vector of estimated residuals u_t is orthogonalized using a Choleski decomposition

$$\varepsilon_t = A_0^{-1}u_t \quad \text{with } u_t u'_t = \Omega \text{ and } \varepsilon_t\varepsilon'_t = D, \text{ where } d_{ij} = 0 \text{ for } i \neq j.$$

Among these orthogonalized residuals, ε_t, the one associated with the i_t equation captures the structural monetary policy shocks, called $\varepsilon_t^{\text{MonetaryPolicy}}$. In substance, $\varepsilon_t^{\text{MonetaryPolicy}}$ are the deviations of the monetary policy instrument from the linear average reaction function to the variables of the VAR during the sample period. The recursiveness of the CEE-99a identification amounts to assuming that the central bank may react to current quarter observations of dy_t, π_t, and π_t^{CP} while it would not react to current quarter developments in $dTR_t, dNBR_t, dM1_t$.[4]

Third, the estimated autoregressive model can be inverted to obtain the MA representation of Y_t:

$$\begin{aligned} Y_t &= A(L)Y_{t-1} + A_0\varepsilon_t = (I - A(L))^{-1}A_0\varepsilon_t \\ &= MA(\infty)\varepsilon_t, \end{aligned}$$

from which the impulse responses reported in the Figs. 1–3 are derived.

3. The effects of monetary policy

3.1. Constant parameters VAR model for the 1960–2005 sample

Fig. 1 describes the effects of a monetary tightening shock. It reports the impulse responses of the main variables of interest as well as the price level and the GDP level responses.[5]

[3] d log stands for first difference of the variables logarithm.

[4] See Leeper et al. (1996) and CEE-99 for a discussion of recursiveness in the identification of US monetary policy shocks.

[5] The observations are quarterly and each equation of the model is estimated with a constant term and four lags. In all impulse response figures, which are estimated with Rats 5, the confidence bands correspond to the 10th and the 90th percentiles of 1000 Monte Carlo replications of the model.

Through out the text we systemically report the response of the largest monetary aggregate included in the model in order to check that the interest rate shock corresponds to a money supply shock.

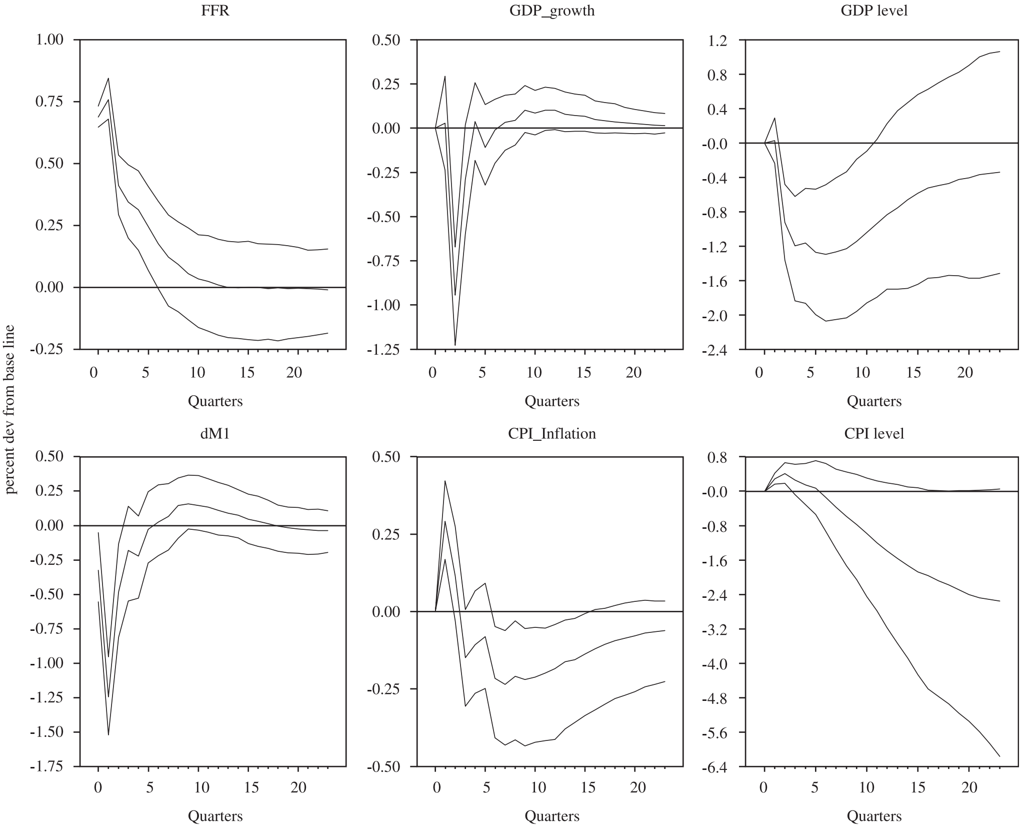

Fig. 1. Effects of monetary policy when estimated for the 1960–2005 sample period CEE-99a model, 4 lags; 10th, 50th and 90th percentiles of the impulse responses.

The picture matches very well with the one obtained with the same model estimated by CEE-99a on levels directly for the 1960–2005 sample period[6]:

- The interest rate returns to baseline within three years after the initial shock;
- given that the initial responses of reserves and M1 is negative, $\varepsilon_t^{\text{MonetaryPolicy}}$ can be interpreted as a money supply shock;
- the hump-shaped response of GDP is significant before GDP returns to baseline;
- the response of inflation is initially positive, i.e., we observe a price puzzle;
- the inflation response eventually shows a negative hump that is significantly different from zero. It reaches a minimum noticeably after the trough of the GDP growth response.

A long list of alternative identifications of US monetary policy shocks confirm these results.[7] What is even more remarkable is that these results are widely agreed upon in the

[6]We simply use the largest set of available data for the estimation. We therefore start the sample in 1960, which is the first observation of monetary aggregates that are consistent to date.

[7]This is for instance the case for the alternative identification (Gordon and Leeper, 1994; CEE, 2005; Giordani, 2004; and CEE-99a using monthly data) with which I check for the robustness of my results. The description of these models is available in Mojon (2005).

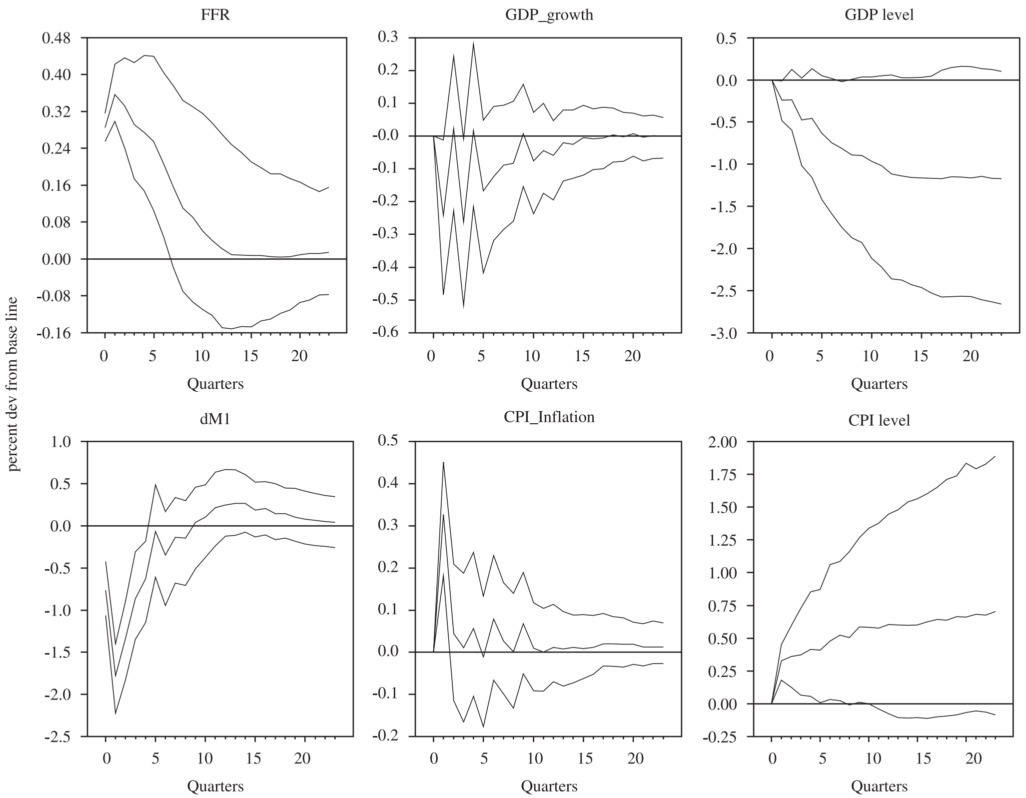

Fig. 2. Effects of monetary policy, estimates for the 1984–2005 period CEE-99a model, 4 lags; 10th, 50th and 90th percentiles of the impulse responses.

profession as describing the effects of unsystematic (and in some case also systematic) US monetary policy. From leading neo-Keynesian academics (e.g., Mankiw, 2001) to RBC developers (e.g., Christianno, Eichenbaum and Evans) and central bankers (Angeloni et al., 2003; Papademos, 2003; the Federal Reserve Board web site introduction on the transmission mechanism), ''we economists'' consider that monetary policy affects inflation first by affecting demand, which eventually puts pressure on prices and wages. This consensus view has also became a benchmark to model the transmission mechanism of other OECD countries (Sims, 1992; Peersman and Smets, 2003; Mojon and Peersman, 2003; Kim, 1999; Angeloni et al., 2003).

As a result, this pattern of impulse responses has become one of the targets for calibrating or estimating structural models (e.g., Christiano et al., 2005; Altig et al., 2005). As stressed in the introduction of CEE-99a, these models ought to reproduce these ''well measured in the data and well accepted'' effects of US monetary policy shocks.

3.2. Estimates for the post 1984 sample

Are the results presented in Fig. 1 stable over time? In particular, there are several good reasons for focusing on the post 1984 period. First, estimating models on sub-samples is a

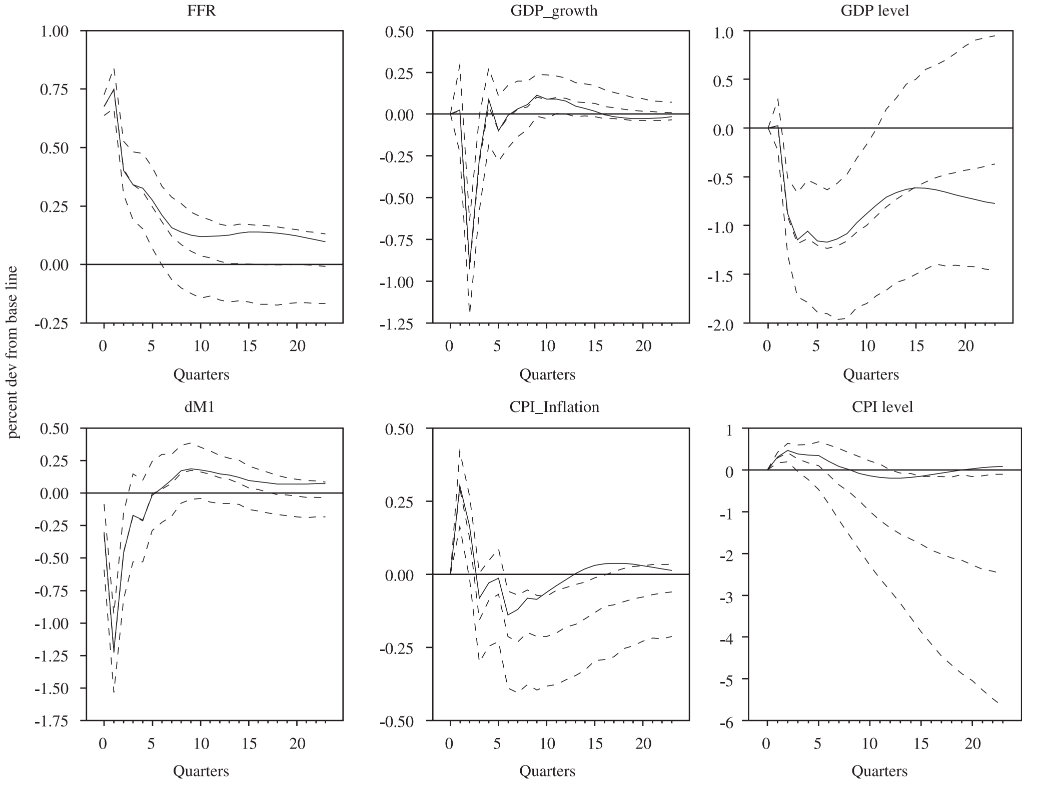

Fig. 3. Controlling for changes in the intercepts at break dates 1960–2005; dotted lines, which correspond to constant intercepts, are similar to Fig. 1.

standard procedure to check for robustness. Moreover, the ongoing transformation of the economy (e.g., the spreading of information technology) suggests that policy makers should give more prominence to the more recent evidence on macroeconomic adjustments. Second, several economists have provided evidence of structural breaks in the mid-1980s. McConnell and Perez-Quiros (2000) showed that the volatility of US output growth admits a significant break in 1984. Turning to inflation, the amplitude of its fluctuations have become an order of magnitude smaller than the one observed in the 1970's. Levin and Piger (2004) and Gadzinski and Orlandi (2004) also argue that the mean and the persistence of inflation may have dropped after the Volker disinflation was completed.

Third, there is also an active debate on the existence of the monetary policy regime shifts since the early 1960's.[8] Clarida et al. (2000) and Boivin and Giannoni (2006) have argued that inflation expectations were not anchored before the Volker–Greenspan era because the implicit reaction function of the Federal reserve was, up to the late 1970's accomodative of inflationary shocks. Sims and Zha (2006) and Primiceri (2005) showed

[8]Time series based analyses of monetary policy usually start in 1960 due to limited availability of monetary aggregate data prior to that date.

that the implicit reaction function of the Federal Reserve appeared rather stable if changes in the variance of monetary policy shocks is taken into account.[9] However, both of these contributions find that this variance was higher in the 1970's than over the last twenty years. Hence, estimation of the constant parameters model is arguably more legitimate in the recent ''homoscedastic'' sub-sample than for the full sample. Finally, Goodfriend and King (2005) argue that it took Paul Volcker several years to anchor inflation expectations at lower levels than in the 1970's. The first years of the Federal Reserve under Volcker therefore correspond to an ''incredible disinflation'', where the dynamics of inflation may not be similar to the one that prevailed once the disinflation was completed. For all these reasons, looking at the effects of monetary policy as it can be estimated in the period starting in 1984, i.e., after the Volker disinflation is completed, seems warranted.

Fig. 2 reports the impulse responses for exactly the same model as described in the previous section but restricting the sample to the last two decades. Many impulse responses are similar to ones estimated for the 1960–2005 sample. In particular, the response of the interest rate, reserves and M1 validate that we are describing a money supply shock. However, there are also two striking differences:

1. The size of the monetary policy shocks is half the size estimated for the full sample period.
2. Neither the inflation nor the price level responses are significant.

These results are actually robust across identification schemes representative of the most popular identifications in the literature, including:

- A monthly version of a simplified CEE-99a model, i.e., using industrial production instead of GDP as an indicator or real activity and excluding total reserves and non-borrowed reserves from the set of variables entering the VAR.
- The Gordon and Leeper (1994) non-recursive identification model.
- The CEE-01 model that controls for many additional variables with respect to CEE-99a (real wage, productivity, profits and the stock price).
- The Giordani (2004) model that contains only capacity utilization, inflation and the FFR.[10]

The smaller scale of interest rate shocks with respect to the full sample result is well documented (e.g., Primiceri, 2005). The volatility of exogenous monetary policy shocks was higher in the 1970's and especially so in the 1979–1982 period of strict monetary targeting under Paul Volker. Besides, the size of the shock is irrelevant for VAR based impulse responses which are proportional to the arbitrary chosen size of the initial shock.

However, the other difference points to a markedly different set of stylized facts for the effects of exogenous monetary policy shocks on inflation than the one described in Mankiw (2001), Christiano et al. (1999a, 2005) and textbooks (e.g., Woodford, 2003). The response of inflation oscillates around zero. Its cummulated effect on the price level is

[9] See also Hanson (2003).

[10] The exact definition and source of the variables used in these models and the identification procedure implemented for each of them as well as the estimated effects of monetary policy shocks for 1960–2005 and 1984–2005 are described in Mojon (2005).

essentially flat. This contrast with the 1960–2005 estimates should not be neglected. It makes a big difference, both for the development of structural models and for the implementation of monetary policy whether in a stable monetary policy regime such as the post 1984 sample, deviations from the reaction function of the central bank have an effect on inflation or not.

The next section is therefore an attempt to gain insights into the determinants of the inflation response over the full sample period.

3.3. *Estimates for the 1960–2005 sample allowing for changes in the mean of inflation*

This section proposes a simple explanation of the contrast between the results after 1984 and the ones reported in CEE-99a. By construction, the impulse responses based on standard VARs impose constant parameters for the dynamics of the variables throughout the sample. This point has been criticized when applied to the modelling of US monetary policy. Some, including Clarida et al. (2000) and Boivin and Giannoni (2006), argue that the reaction function of the central bank changed in the late 1970's. Others, Sims and Zha (2006) and Primiceri (2005) consider that these changes are not significant once controlling for the time variation in the variance of the VAR shocks.

We show in this section that one parameter of the VAR has actually changed significantly in addition to the change in the shocks' variance. This parameter is the intercept of the inflation equation, i.e., the mean of inflation. Once controlling for these changes in the mean of inflation, monetary policy shocks do not have a significant effect on inflation or the price level.

3.3.1. *Changes in the mean of US inflation*

The US has experienced persistent changes in the level of inflation. Inflation rose in the late 1960s and again in the early 1970s before it declined sharply in the early 1980's. Except for these three episodes, inflation seems to have fluctuated around a stable mean (about 2.5%) both before 1965 and after 1982 and clearly above that level in between. Statistical tests on breaks in the mean of US inflation corroborate this description of the data. In particular, the multiple break test developed by Altissimo and Corradi (2003) points to three breaks in the mean of the US inflation process:

1. 1967 Q3 when the mean of CPI (GDP deflator) inflation increased by 2.5% from 2.1 to 4.6 (from 2.0 to 4.5),
2. 1973 Q1 when the mean of CPI (GDP deflator) inflation increased by 2.8% from 4.6 to 7.3 (by 4.0 from 4.5 to 8.5),
3. 1982 Q2 when the mean of CPI (GDP deflator) inflation decreased to 2.5% (3.0%). [11,12]

[11]These dates are consistent with the results of alternative break tests that have been implemented by others. For a synthesis of these results, see Mojon (2005, Table 1).

[12]It should be stressed that these estimated break dates correspond to some well-known events in the US monetary policy history. To start with, the 1982 break marks the success of the Volcker "Conquest of the Post War US inflation." See the discussion in Cogley and Sargent (2005) and Clarida et al. (2000).

Turning to the 1967 and the 1973 breaks, it appears that the persistent misperception of the productivity trend in real time may have led the Federal Reserve to consistently under-estimate the neutral level of interest rates (Orphanides, 2003).

3.3.2. VAR estimates with changes in the mean of inflation

One simple experiment therefore consists of estimating the effects of monetary policy over the full 1960–2005 sample in a specification which allows for changes in the mean of inflation. The VAR is estimated with dummies that permit changes in the intercept of the VAR equations in 1967 Q3, 1973 Q1 and 1982 Q2. Two results of this experiment are particularly striking.

First, this form of time variation in the VAR parameters is not rejected by the data. The changes in the intercept of the VAR equation are significant. Hence, breaks that are estimated with the Altissimo and Corradi procedure (and alternative break tests), which are based on the information contained in the inflation series only, are not due to the omission of some variable to which the mean of inflation would endogenously respond.[13]

Second, this slight modification of the VAR specification, i.e., dropping 3 degrees of freedom to allow for steps in the intercept, dramatically reduces the response of inflation and prices to monetary policy shocks (Fig. 3). Effectively, estimating the same VAR as in Section 3.1, except for letting the intercept of the inflation equation take a higher value during the 1970's, obtains the same results as the ones estimated over the post 1984 sample: Inflation wiggles around zero and the price level response is flat.[14]

One possible interpretation of the difference in the impulse responses from the models with and without changes in the intercept is that the widely accepted hump-shaped response of inflation in the CEE-99a results come from a few changes in the mean of inflation which took place around the 1970's.

This result contradicts somewhat the point made by Primiceri (2005) and Sims and Zha (2006), who conclude that outside the changes in the variance of the shocks, the VAR parameters are stable. Perhaps this is because these authors check simultaneously for changes in all parameters of the VARs, with, arguably, little statistical power to spot breaks in any single coefficients of the VAR, while this paper tests only whether the constant term changed at given dates. In the end, the impulse responses of prices in Sims and Zha (2006) are not significantly different from zero and hence consistent with the ones reported in Fig. 3.

4. Conclusion

This paper shows that the hump-shaped response of US inflation, which is obtained with VAR based identification and widely considered a stepping stone for the development of structural models, is not robust. In particular, the post 1984 experience is one where VAR estimated money supply shocks have no significant effect on inflation and the response of the price level is flat. This change in the response of inflation points to time variation in the effects of monetary policy. Outside periods of large and persistent adjustments, the VAR based measures of non-systematic monetary policy do not affect inflation. While this result may simply reflect some limitations of the VAR methodology, it nevertheless establishes a

[13]These changes in the intercept of the inflation equation are robust when controlling for the heteroskedasticity of the errors. The T-statistics of the breaks in the intercept of the inflation equations in 1967 Q3, 1973 Q1 and 1982 Q2 drop from 2.0, 2,7 and 3.8, respectively, in the OLS estimate to 1.6, 2.1 and 3.6, respectively, when estimated with GLS controlling for heteroschedasticity of the errors as in White (1980). Hence two of the three breaks remain significant at the standard 5% threshold while the first 1967 Q3 remains so only at the 12% threshold.

[14]Figs. A1–A5 in Mojon (2005) show that this result holds for the four alternative identification schemes listed in Section 3.2.

new benchmark stylized fact for the calibration of macroeconometric models of the US business cycle.

Definition	Notation
Federal funds rate	i_t
$\text{Log}(\text{GDP}_t/\text{GDP}_{t-1})$	dy_t
$\text{Log}(\text{CPI}_t/\text{CPI}_{t-1})$	π_t
$\text{Log}(\text{Commodity PI}_t/\text{Commodity PI}_{t-1})$	π_t^{CP}
$\text{Log}(\text{TR}_t/\text{TR}_{t-1})$	dTR_t
$\text{Log}(\text{NBR}_t/\text{NBR}_{t-1})$	$dNBR_t$
$\text{Log}(\text{M1}_t/\text{M1}_{t-1})$	$dM1_t$

Sources

Interest rates: Federal funds rate, (% P.A.); Fed H15.
Monetary aggregates: M1, (SA Billions $); Fed H.6 Money stock and liquid assets, and debt measures.
Reserves: Non-borrowed reserves adjusted for changes in reserve requirements, (Mil. $, SA) and total reserves adjusted for changes in reserve requirements, (Mil. $, SA); FRB: Aggregate reserves of depository institutions—H.3.
Economic activity: GDP in billions of chained 2000 dollars, BEA.
Prices: CPI: Urban consumer—all items, (1982–1984 = 100, SA), BLS; KR-CRB futures price index, (1967 = 100); Knight–Ridder, commodity index report.

All data are downloadable from www.bea.gov and www.freelunch.com and the web page of Boschen and Mills for their index of monetary policy tightness.

Acknowledgements

This paper was written while I was working at the European Central Bank and an Associate Professor at CEDERS, Université Aix-Marseille 2. I would like to thank Jürgen von Hagen, two anonymous referees, Harris Dellas, Anil Kashyap, Giorgio Primiceri, Sergio Rebello, Tao Zha and participants in seminars at the ECB and the Eurosystem IPN meeting held in Vienna, for useful comments and suggestions on a previous draft. The views expressed in the paper are not necessarily those of the Federal Reserve Bank of Chicago or the Federal Reserve Bank System. The author is the only person responsible for any remaining errors.

References

Altig, D., Christiano, L., Eichenbaum, M., Linde, J., (2005). Firm-specific capital, nominal rigidities, and the business cycle. NBER Working Papers No. 11034.

Altissimo, F., Corradi, V., 2003. Strong rules for detecting the number of breaks in a time series. Journal of Econometrics 117, 207–244.

Angeloni, I., Kashyap, A., Mojon, B., Terlizzese, D., 2003. The output composition puzzle: A difference in the monetary transmission mechanism in the Euro Area and the United States. Journal of Money, Credit and Banking 35 (6), 1265–1306.

Bernanke, B., Mihov, I., 1998. Measuring Monetary Policy. Quarterly Journal of Economics 113 (3), 869–902.

Boivin, J., Giannoni, M.P., 2006. Has monetary policy become more effective? The Review of Economics and Statistics 88 (3), 445–462.

Christiano, L., Eichenbaum, M., Evans, C., 2005. Nominal rigidities and the dynamic effects of a shock to monetary policy. Journal of Political Economy 113 (1), 1–45.

Christiano, L.J., Eichenbaum, M., Evans, C., 1999a. Monetary policy shocks: What have we learned and to what end? In: Taylor, J.B., Woodford, M. (Eds.), Handbook of Macroeconomics, vol. 1A. pp. 65–148.

Christiano, L.J., Eichenbaum, M., Evans, C., 1999b. Sticky price and limited participation models: A comparison. European Economic Review 41 (6), 1173–1200.

Clarida, R., Gali, J., Gertler, M., 2000. Monetary policy rules and macroeconomic stability: Evidence and some theory. Quarterly Journal of Economics 115 (1), 147–180.

Cogley, T., Sargent, T., 2005. Then conquest of US inflation: Learning and robustness to model uncertainty. Review of Economic Dynamics 8, 528–563.

Gadzinski, G., Orlandi, F., 2004. Inflation persistence for the EU countries, the Euro Area and the US. ECB WP 414.

Giordani, P., 2004. An alternative explanation of the price puzzle. Journal of Monetary Economics 51 (6), 1271–1296.

Goodfriend, M., King, R., 2005. The incredible Volker disinflation. NBER Working Paper 11562.

Gordon, D., Leeper, E., 1994. The dynamic impacts of monetary policy: An exercise in tentative identification. Journal of Political Economy 102 (6), 1228–1247.

Hanson, M., 2003. Varying monetary policy regimes: A vector autoregressive investigation. Wesleyan University Mimeo.

Kim, S., 1999. Do monetary policy shocks matter in the G-7 countries? Using common identifying assumptions about monetary policy across countries. Journal of International Economics 48 (2), 387–412.

Leeper, E., Sims, C., Zha, T., 1996. What does monetary policy do? Brookings Papers on Economic Activity 2, 1–78.

Levin, A.T., Piger, J.M., 2004. Is inflation persistence intrinsic in industrial economies? ECB Working Paper No. 334.

Mankiw, G., 2001. The inexorable and mysterious trade off between inflation and unemployment. The Economic Journal 111 (471), C45–C61.

McConnell, M., Perez-Quiros, G., 2000. Output fluctuations in the United States: What has changed since the early 1980s? American Economic Review 90, 1464–1476.

Mojon, B., 2005. When did unsystematic monetary policy have an effect on US inflation? ECB Working Paper No. 559.

Mojon, B., Peersman, G., 2003. A VAR description of the effects of monetary policy in the individual countries of the Euro Area. In: Angeloni, I., Kashyap, A., Mojon, B. (Eds.), Monetary Policy Transmission in the Euro Area. Cambridge University Press, Cambridge.

Orphanides, A., 2003. The quest of prosperity without inflation. Journal of Monetary Economics 50 (3), 633–663.

Papademos, L., 2003. Business Cycles and Monetary Policy. In: Speech given at the Banque de France International Conference on: Politique monétaire, cycle économique et dynamique financière, 7 March 2003, Paris.

Peersman, G., Smets, F., 2003. The monetary transmission mechanism in the Euro Area: More evidence from VAR analysis. In: Angeloni, I., Kashyap, A., Mojon, B. (Eds.), Monetary Policy Transmission in the Euro Area. Cambridge University Press, Cambridge.

Primiceri, G., 2005. Time varying structural vector autoregressions and monetary policy. Review of Economic Studies 72 (3), 821–852.

Sims, C., 1980. Macroeconomics and reality. Econometrica 47 (1), 1–40.

Sims, C., 1992. Interpreting the macro-economic time series facts: The effects of monetary policy. European Economic Review 36 (4), 975–1011.

Sims, C., Zha, T., 2006. Were there regime switches in US monetary policy? American Economic Review 96 (1), 54–81.

White, H., 1980. A heteroskedasticity-consistent covariance matrix estimator and direct test for heteroskedasticity. Econometrica 48, 817–838.

Woodford, M., 2003. Interest and Prices. Princeton University Press, Princeton, NJ.