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2 European Central Bank and Université de la Méditerranée; e-mail: benoit.mojon@ecb.int

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Address
Kaiserstrasse 29
60311 Frankfurt am Main, Germany

Postal address
Postfach 16 03 19
60066 Frankfurt am Main, Germany

Telephone
+49 69 1344 0

Internet
http://www.ecb.int

Fax
+49 69 1344 6000

Telex
411 144 ecb d

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CONTENTS

Abstract

An important stylized fact to emerge from the 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 corrects for the effects of large shifts (breaks) in average inflation or examines periods without such shifts (such as the 1984-2004 period). An important consequence is that the hump shaped VAR estimated response of inflation is not appropriate to fit stylised models of the response of inflation around a stable steady state inflation level.

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Keywords: monetary policy transmission, VARs,
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Non technical summary

One of the most widely accepted stylised facts of monetary economics is that US inflation has a hump shape response to exogenous monetary policy shocks. This stylised fact is supported by a series of robustness checks along several dimensions. It is robust to a series of alteration of the identification of monetary policy shocks, i.e.

whether one assumes that the Federal Reserve fix the short term interest rate as a function of only inflation and economic activity or that the level of interest rate that is relevant to measure the stance of monetary policy depends also on monetary aggregates, long term interest rates, commodity prices and indicators of demand for liquidity in the inter-bank market.

However, this characterisation of the effects of unsystematic monetary policy on inflation is very sensitive to the choice of the sample period on which the econometric model of monetary policy decisions is estimated. In particular, if one considers the last twenty years, estimated monetary policy shocks have no effect of 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 sample over which the model 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 until 2004. 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.

1 Introduction

One of the most widely accepted stylised facts of monetary economics is that US inflation has a hump shape response to exogenous monetary policy shocks. For instance, Christiano, Eichenbaum and Evans (2005) state (page 5-8) "*after an expansionary monetary policy shock [. . .] inflation responds in a hump shape 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*".

This stylised 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 Handbood of Macroeconomics: "Monetary Policy Shocks, What have we learned and to what hand?" (also by Christiano, Eichenbaum and Evans, 1999a; thereafter CEE-99a) and in many contributions reviewed therein.

However, this VAR based characterisation 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 of 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 sample 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 till 2004. 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 shows that VAR identifications of monetary policy shocks, estimates of the inflation and the price level impulse responses based on the 1984-2004 are radically different from the ones obtained for the full last 45 years. Section 3 shows that the standard result is largely due to a few discrete changes in the mean of inflation. Section 4 concludes.

[1] Through out the text I refer indifferently to either monetary policy shocks, exogenous monetary policy shocks or unsystematic monetary policy.

2 VAR based identification of US monetary policy

2.1 The identification of US monetary policy shocks

Two decades 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 a earlier paper that limited participation models and sticky price models predict different path for money and the interest rate following a monetary policy shock (CEE, 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 policy 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, *à la Sims* non-recursive models, as well as across models that differ in terms of the number of variables entering the VAR, and therefore the information set on the basis of which the central banks sets its instrument (usually assimilated to the interest rate on Federal Funds).

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 dlog(GDP), dlog(CPI), dlog(Commodity prices), the interest rate on Federal Funds, dlog(Total R), dlog(NBR) and dlog(M1).[3]

[2] The other two most cited surveys are Leeper, Sims and Zha (1996) and Bernanke and Mihov (1998).

[3] dlog stands for fist difference of the variables logarythm.

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, we orthogonalise the vector of estimated residuals u_t using a Choleski decomposition.

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

Among these orthogonalized residuals ε_t, we consider that the one associated to the i_t equation captures the monetary policy structural shocks, i.e. we call the fourth equation orthogonalized residual $\varepsilon_t^{MonetaryPolicy}$. In substance, $\varepsilon_t^{MonetaryPolicy}$ 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, on average, not react to current quarter developments in $dTR_t, dNBR_t, dM1_t$.[4]

Third: we can then invert the estimated autoregressive model to obtain the MA representation of Y_t :

$$Y_t = A(L)Y_{t-1} + A_0\varepsilon_t = (I - A(L))^{-1}A_0\varepsilon_t$$

$$= MA(\infty)\varepsilon_t$$

from which we compute the impulse responses reported in the Figures of the paper.

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

2.2 The effects of monetary policy as estimated for the 1960-2005 sample

Figure 2 reports the impulse responses of the main variables of interest in Y_t as well as the price level and the GDP level responses[5]. The picture matches very well 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 3 years after the initial shock;
- given that the initial responses of reserves and M1 is negative, $\varepsilon_t^{MonetaryPolicy}$ 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 also humped and significantly different from zero. It picks noticeably after the pick of the GDP growth response.

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

[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 the interest rate shock corresponds to a money supply shock.

[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, 2001; Giordani, 2004; and CEE-99 using monthly data) with which we check for the robustness of our results. The description of these models is done in the appendix.

Eichenbaum and Evans) and central bankers (Angeloni et al. 2003, Papademos, 2004; 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; 2004; Mojon and Peersman, 2003, Kim, 1997; 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. CEE, 2005; Altig et al. 2004). As stressed in the introduction of CEE-99a, these models ought to reproduce this "well measured in the data and well accepted" effects of US monetary policy shocks.

2.3 The effects of monetary policy as estimated for the post 1984 sample

Are the results presented in Figure 2 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 standard proceedure to check for robustness. Moreover, the ongoing transformation of the economy (e.g. the spreading of information technology) suggests that policy maker should give more prominence to the more recent evidence on macroeconomic adjustments. Second, several economists have provided evicence of structural breaks in the mid-1980s. McConnel 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) 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 Gali and Gertler (2000) and

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

Boivin and Giannoni (2002) 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 (2005) and Primiceri (2005) showed that the implicit reaction function of the Federal reserve appeared rather stable if changes in the variance of monetary policy shocks is accounted for.[9] However, both of these contributions obtain that this variance was higher in the 1970's than over the last twenty years. Hence, estimation of constant parameters model is arguably more legitimate in the recent "homoschedastic" 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.

Figure 3 reports the impulse responses for exactly the same model as described in the previous section but restricting the sample to the last two decades. There are many similarites with the estimates based on 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 smaller than for the full sample period.

2. Neither the inflation nor the price level responses are significant.

These results are actually robust across identifications schemes representative of the most popular identifications in the literature, including[10]:

[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 are described in the appendix.

- 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 reseres from the set of variables entering the VAR.

- the Gordon and Leeper (1994) non recursive identification model;

- the CEE-01 model that on the contrary 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 the capacity utilization, inflation and the FFR.

The impulse responses of the variable of interest are reported in the appendix Figures A1 to A4.

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 abritrary chosen size of the initial shock.

However, the other difference point to a markedly different set of stylised facts for the effects of exogenous monetary policy shocks on inflation than the one described in by Mankiw (2001), Christianno et al (1999a, 2005) and textbooks (e.g. Woodford, 2003). The response of inflation ocsillates around zero. Its cummulated effect on the price level is essentially flat. This difference 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 on the determinants of the inflation response over the full sample period.

3 What is "wrong" with estimating VARs over the 1960-2005 sample?

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 imposes constant parameters for the dynamics of the variables through out 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 Gianonni (2002), argue that the reaction function of the central bank has changed in the late 1970's. Others, Sims and Zha (2005) 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, nor the price level.

3.1 Has the mean of US inflation been stable?

Can we assume, as is implicit in VARs, that inflation had a stable mean between 1960 to 2005? Not really. The US has experienced long term changes in the level of inflation. Inflation has risen 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. This reading of Figure 1 can actually be tested using formal statistical tests on breaks in the mean of US inflation. These tests, which results are summarized in Table 1, usually obtain three or four breaks in the mean of US inflation since 1947.

I estimate the break dates by implementing the Altissimo-Corradi (2003) procedure for both CPI inflation and GDP deflator inflation. This break test has the advantage that it allows for multiple breaks. In addition, results of alternative tests (e.g. the Bai and Perron test) are already available in from other papers (Benati 2003, 2004; Rapach and Wohar 2002; see also as well as with the results of Levin and Piger 2004; and Gadzinski and Orlandi 2004, although for a much shorter period than mine). My results therefore usefully complement these already available tests.

Three break dates appear relatively robust[11]:

1. 1967 Q3 when the mean of CPI (GDP deflator) inflation increases 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 increases 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 decreases to 2.5 % (3.0 %).

These results are compared to the evidence reported in previous papers in Table 1.

Two observations are in order. First, the timing of the breaks is probably at odd with the prior of most economists who associate the 1970 great inflation to oil. The 1967 break in the mean of US inflation occurs much earlier than the first oil shock. And the last one, in 1982, precedes the 1986 counter oil shock by four years. We will nevertheless also investigate the effects of controlling for the 1991 break on measures of the effects of monetary policy.

[11] We don't find a break in the early 1990's as Levin and Piger and Gadzinski and Orlandi and one of Benati's tests do. We also note that the last observations of the sample tend to correct the downward shift that appears in the inflation time series between the early 1990's and the late 1990's (see Figure 1).

Second, the break dates that we observe 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". Cogley and Sargent (2004) show that the US low frequency swings in inflation are consistent with the Federal Reserve gradually upgrading its view on the (im)possibility of exploiting an output-inflation trade off. Clarida et al. (2000) conclude that the major change in US inflation, i.e. the Volcker disinflation, coincide with a change in the central bank's reaction function from setting pro-cyclical real interest rates to counter-cyclical ones.

Turning to the 1967 and the 1973 breaks, it appears that the persistent mis-perception of the productivity trend in real time may have led the Federal Reserve to consistently under-estimated the neutral level of interest rates (Orphanides, 1999).

I don't think though that the results of these tests should be taken at face value. Indeed, break tests tend, when applied to highly persistent processes, to be subject to type II error (O'Reilly and Whelan, 2004), the estimated breaks may be spurious. However, the presence of breaks in the inflation time series indicates that the adjustment of inflation may be discontinuous. "Normal times", when inflation adjusts around a stable mean sharply contrast with times of shifts in the mean. This contrast raises the concern that models which impose homogeneity in the adjustment of inflation through out time could be mis-specified.

Knowing that adjustment around the breaks is of a different nature than adjustments in normal times is very useful insofar as it avoids overrating or underestimating the response of inflation. For instance, a few recent studies of univariate models of inflation stress that allowing for breaks in the mean of inflation reduces the estimates of inflation persistence (Levin and Piger 2004, Benati, 2004, Gadzinski and Orlandi, 2004). In addition, Corvoiser and Mojon (2004) show that breaks in the mean of inflation generally (across 23 OECD countries including the US) affect the speed of the response of inflation to a variety of nominal and real shocks. To the extent that breaks in the mean of inflation reflect changes in policy regimes, one interpretation is that inflation is less persistent within monetary policy regimes than across policy regimes, possibly because inflation expectations adjust sluggishly to the new policy regime (Ball, 1994; Nicolae and Nolan, 2004; Corvoisier and Mojon, 2004). Therefore, the next section will extend the evaluation of the effects of breaks in the mean of inflation from univariate models that encompass all types of shocks to the analysis of monetary policy shocks alone.

3.2 Controlling for breaks in the VAR

3.2.1 Methodology

In addition to sub-sample estimates, which allow for changes in all the parameters of the VARs, two other strategies for controlling the effects of breaks in the mean of inflation are possible[12]. The first strategy consists of estimating the exact same VARs as discussed in section 2 while substituting the raw inflation series by "demeaned inflation" (which fluctuates around the zero line in Figure 1), i.e. the deviations of inflation from its *breaking* mean. The second one is to allow for changes in the intercepts of the VAR equations at the dates of the breaks in the mean of inflation.

I opt for the latter strategy because it nests the former, as it allows breaks in the mean of the other variables of the VAR when the mean of inflation breaks. I find that breaks in the intercepts of the VAR equations are significant only for the CPI inflation, the Commodity price index inflation and the growth rate of total reserve, which is, perhaps surprisingly, smaller in the period between 1973 and 1982. Changes in the intercept of the inflation equation are significantly different from zero. 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.

[12]Given the short span of data prior to 1967 and the proximity of breaks, this approch, which we followed in section 2 for the post 1984 period, can unfortunately not be extended to other sub-samples.

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 respectively from 2.0, 2,7 and 3.8 in the OLS estimate to respectively 1.6, 2.1 and 3.6 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 %.

Altogether, this result contradicts somewhat the point of Primiceri (2005) and Sims and Zha (2005) who both conclude that outside the changes in the variance of the shocks, the VAR parameters are stable. My results indicate instead that I cannot reject that the changes in the mean of US inflation, as dated thanks to the break tests presented above, are associated to statistically significant changes in the intercept of the inflation equation. This result also confirms the graphical impression that the time series of inflation has indeed seen its mean changing in the 1970's.

3.2.2 Results

Figure 4 compares the impulse response of the Federal funds rate, GDP growth, inflation and the price level in the CEE-99a with and without breaks in the intercept of the model equations. The point estimate of the impulse responses with breaks is the dashed line, while the full lines are similar to the standard model showed in Figure 2. Figures A1 to A5 focus on the response of inflation and the price level for the four other VARs[13]. There again, the point estimate of the model with breaks is compared to the 10th, 50th and 90th percentile of 1000 Monte Carlo simulations of the model without breaks.

The breaks in the mean of inflation affect mostly the impulse response of inflation. The latter is close or above the 10th percentile of the IRF of inflation

[13] For this exercise, I also report estimates of a VAR where instead of the interest rate on Federal Funds, I use the Boschen and Mills index of monetary policy tightness. This model cannot however be estimated for the post 1984 sample because the Boschen and Mills index has not been updated after 1995. See the presentation of this model in the appendix.

in the model without breaks. Looking at the price level response reveals an even much sharper contrast between the VAR model ignoring the breaks and the model controlling for them.[14] [15]

What do these results imply?

First, the impact of including breaks corroborates the view that there may be indeed discontinuous adjustments of inflation including in its response to unsystematic monetary policy. It is remarkable that only 3 degrees of freedom have such a large impact on the shape and magnitude of the responses of inflation and the price level. Second, that the omission of breaks result in a larger response of inflation is at least suggestive that breaks in the mean of US inflation are monetary policy phenomena. It seems that the response of inflation as it is measured in the literature is capturing these very persistent changes in the mean of inflation. In other words, our result is consistent with unsystematic monetary policy having contributed to the shifts toward higher and then lower levels of inflation that took place around the break dates. This view is actually reinforced by plotting together the breaks in the mean of inflation and the moving average of order 5 of the exogenous monatery policy shocks (see Figure 5)[16]. It is clear that the 1982 drop in the mean of inflation was preceded by a tightening of monetary policy between a sample low in 1975 to the sample pick in 1981. Likewise, the 1967 and 1973 breaks occur at a period when the tightness of monetary policy is trending down.

Third, the response of inflation to unsystematic monetary policy is not significant outside the period of the breaks in the mean of inflation. This does not, however, imply that systematic monetary policy has no effect on inflation. Ac-

[14] Although the point estimates of the inflation and the price level responses may actually increase (Figure 4 and A1 to A5), these increases are typically not significant.

[15] Using a similar approach on respectively the 1974-2005 and the 1960-1981 sub-samples showed that controling respectively for the the 1982 break alone or both the 1967 and 1973 breaks had the same consequence of reducing dramatically the response of inflation and the price level to monetary policy shocks. Results, which are not reported for the sake of space, are available upon request.

[16] As estimated in CEE-1999 in the model that does not include controls for breaks in the mean of inflation.

tually, the VAR can only picks up deviations of the monetary instrument from the estimated monetary policy rule. My estimates of this policy rule/reaction function are such that the policy interest rate has a positive response to inflation and GDP growth, and so in both models (with and without breaks). It may still be the case that systematic monetary policy stabilizes inflation. However, on average, the deviations of the interest rate from the policy rule, which are "orthogonal" to other shocks and therefore identifiable, had no impact on inflation once we control for a few shifts in the mean of inflation.

To summarize, unsystematic monetary policy, which is defined in the VAR as deviations from the monetary policy rule, has a much smaller effect on inflation than is usually thought, except at the three occasions when the mean of inflation changed. It therefore appears that much of the response of inflation in the model that ignores such breaks comes from the rare and large adjustments of US inflation around the dates identified by the break tests. Even if we do not take these break tests literally, the estimates that I reported here point to instability of the response of inflation to monetary policy shocks.

4 Conclusion

This paper shows that the hump shape response of US inflation which is obtained with VAR based identification and considered a stepping stone of the development stuctural 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.

Second, the paper suggests that the typical hump shape response is actually picking up the large and persistent adjustments that took place around the Great Inflation of the 1970's. This is why, controlling for shifts in the mean of inflation radically reduces the VAR-based measures of the response of US inflation to unsystematic monetary policy. Once these shifts are filtered out, the response of inflation and the price level appear much weaker than in the consensual picture.

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

Appendix: An overview of the five alternative identifications used in the paper

Models and identification

a) Christiano, Eichenbaum, and Evans (1999)

See section 2 of the main text.

Our vector of variables is Yt with $Yt' = (dy_t, \pi_t, \pi_t^{CP}, i_t, dTR_t, dNBR_t, dM1_t)$. See the next section for the definition and source of the variables.

In the paper we refer to estimates of this model estimated both with quarterly and monthly data. In the latter case, dy_t is measured by industrial production.

b) Gordon and Leeper (1994)

The second model is based on an identification procedure proposed by Gordon and Leeper (1994). This model adds a long-term (ten-year) interest rate and substitute M2 to M1, while measure of reserves are excluded. Gordon and Leeper opt for an alternative set of identifying restrictions that focus on the information that the central bank could be expected to have at the time of setting the short-term interest rate. Accordingly, contemporaneous observations on inflation and GDP cannot influence this decision—leaving only contemporaneous commodity prices, the long-term interest rate, and M2 as potentially affecting the contemporaneous Federal funds rate. In contrast, contemporaneous prices and GDP components enter the "private sector driven" money demand equation. Here I slightly modify this identification by allowing the long-term interest rate to react to contemporaneous innovation in the federal funds rate.

The vector of variables is given by: $Yt' = (dy_t, \pi_t, \pi_t^{CP}, l_t, i_t, dM2_t)$

We estimate the money supply and the money demand shocks by imposing zero contemporaneous restrictions as follows:

$$
\begin{pmatrix} dy_t \\ \pi_t \\ \pi_t^{CP} \\ l_t \\ i_t \\ dM2_t \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ X & 1 & 0 & 0 & 0 & 0 \\ X & X & 1 & 0 & 0 & 0 \\ X & X & X & 1 & X & 0 \\ 0 & 0 & X & X & 1 & X \\ X & X & 0 & 0 & X & 1 \end{pmatrix} \begin{pmatrix} \varepsilon_t^1 \\ \varepsilon_t^2 \\ \varepsilon_t^3 \\ \varepsilon_t^4 \\ \varepsilon_t^{Money\ Supply} \\ \varepsilon_t^{Money\ Demand} \end{pmatrix}
$$

c) Christiano, Eichenbaum, and Evans (2001)

The third model is taken from Christiano, Eichenbaum, and Evans (2001). This model includes consumption, investment, GDP, the CPI, a real wage variable, a labor productivity measure, real corporate profits, the federal funds rate, M2 growth, and the Standard and Poor's 500 stock price index deflated by the CPI.

The model's vector of variables is

$Yt' = (dy_t, \pi_t, d\left(\frac{W}{P}\right)_t, d\left(\frac{Y}{H}\right)_t, i_t, d(PY - WH)_t, dM2_t, dS\&P_t^{500})$

and the identification is recursive.

d) Giordani (2004)

In the fourth model, I follow Giordani (2004). This model uses capacity utilization instead of GDP because it is a better approximation of the output gap that enters the reaction function of the Federal Reserve. Giordani (2004) shows both in a simple theoretical model and in a simple 3 variables estimated VAR model (capacity utilization, inflation and the Federal funds rate) that the price puzzle is very much attenuated with respect to models using GDP as the measure of real activity.

The model includes $(CAP\_U._t, \pi_t, i_t)$, which respectively stand for Capacity Utilization, dlog(CPI) and the FFR.

The identification is recursive.

e) Boschen and Mills (1995) inspired VAR

We simply replace the 3 monetary policy variables of the CEE-99a model, i.e. FFR, total reserves and non-borrowed reserves with the Boschen and Mills index of monetary tightness.

The vector of variables is given by: $Yt' = (dy_t, \pi_t, \pi_t^{CP}, l_t, BSI_t, dM2_t)$
The identification is recursive.

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Data definition and notation

Definition	Notation
Federal funds rate	i_t
Log(GDP$_t$/GDP$_{t-1}$)	dy_t
Log(CPI$_t$/CPI$_{t-1}$)	π_t
Log(Commodity PI$_t$/Commodity PI$_{t-1}$)	π_t^{CP}
Log(TR$_t$/TR$_{t-1}$)	dTR_t
Log(NBR$_t$/NBR$_{t-1}$)	$dNBR_t$
Log(M1$_t$/M1$_{t-1}$)	$dM1_t$
Log(M2$_t$/M2$_{t-1}$)	$dM2_t$
Bond interest rate	l_t
Real wage inflation	$d\left(\frac{W}{P}\right)_t$
Log(Hours' productivity$_t$/H' prod.$_{t-1}$)	$d\left(\frac{Y}{H}\right)_t$
Log(Profits$_t$/Profits$_{t-1}$)	$d(PY - WH)_t$
Log(Stock prices$_t$/Stock prices$_{t-1}$)	$dS\&P_t^{500}$
Capacity Utilistation	$CAP\_U._t$
Boschen and Mill Index	BSI_t

Sources

Interest Rates: Federal Funds Rate, (% P.A.); Fed H15 and 10-Year Constant Maturity Securities, (% P.A.); Fed H15

Monetary aggregates: M1 and M2, (SA Billions $);Fed H.6 Money Stock and Liquid Assets, and Debt Measures

Reserves: Nonborrowed 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; Capacity Utilization: Manufacturing - SIC, (% Capacity, SA);Federal Reserve Board, G.17; Industrial Production: Total index, (Index 1997=100, SA), Federal Reserve Board, G.17

Prices: CPI: Urban Consumer - All items, (1982-84=100, SA), BLS; KR-CRB Futures Price Index, (1967=100);Knight-Ridder, Commodity Index Report

Wages: Avg. Hrly Earnings: Total Private, ($ Per Hrs., SA);Bureau of Labor Statistics: Form 790, P is the CPI above.

Productivity: Productivity & Costs: Nonfarm Business - Output Per Hour All persons, (Index 1992 = 100);BLS: Productivity & Costs

Profits: Corporate profits with inventory valuation and capital consumption adjustments; Non financial (row 4); Survey of Current Business: Table 6.16 b,c and d; BEA

S&P **Stock Price** Index: 500 Composite, (Index 1941-43=10, Monthly Average);Standard & Poors: Security Price Index Record

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.

Table 1: Dates of Breaks in the mean of US inflation

Paper	Test	Sample period	Series used	Dates of break			
				1960s	1970s	1980s	1990s
Benati (2003)	Bai and Perron	1960-2002	CPI/P_GDP	59-63		76-82	
	Bai-Andrews-Ploberger 1	1960-2002	CPI/P_GDP	56-60	72-73	79-84	85-96
Benati (2004)	Bai-Andrews-Ploberger 2	1960-2002	CPI/P_GDP		77-84		
Corvoisier&Mojon	Altissimo and Corradi	1960-2003	CPI/P_GDP	67Q3	73Q1	82Q2	
Rapach&Wohar	Bai and Perron	1960-1998	CPI/P_GDP	66-68	71-73	80-82	
Gadzinski&Orlandi	Single break test.	1984-2003		n.a.	n.a.	n.a.	91
Levin&Piger	Single break test.	1984-2002	CPI/P_GDP	n.a.	n.a.	n.a.	91

Source: Corvoisier and Mojon Appendix Table A2

Figure 1a: Time series of selected US macroeconomic variables

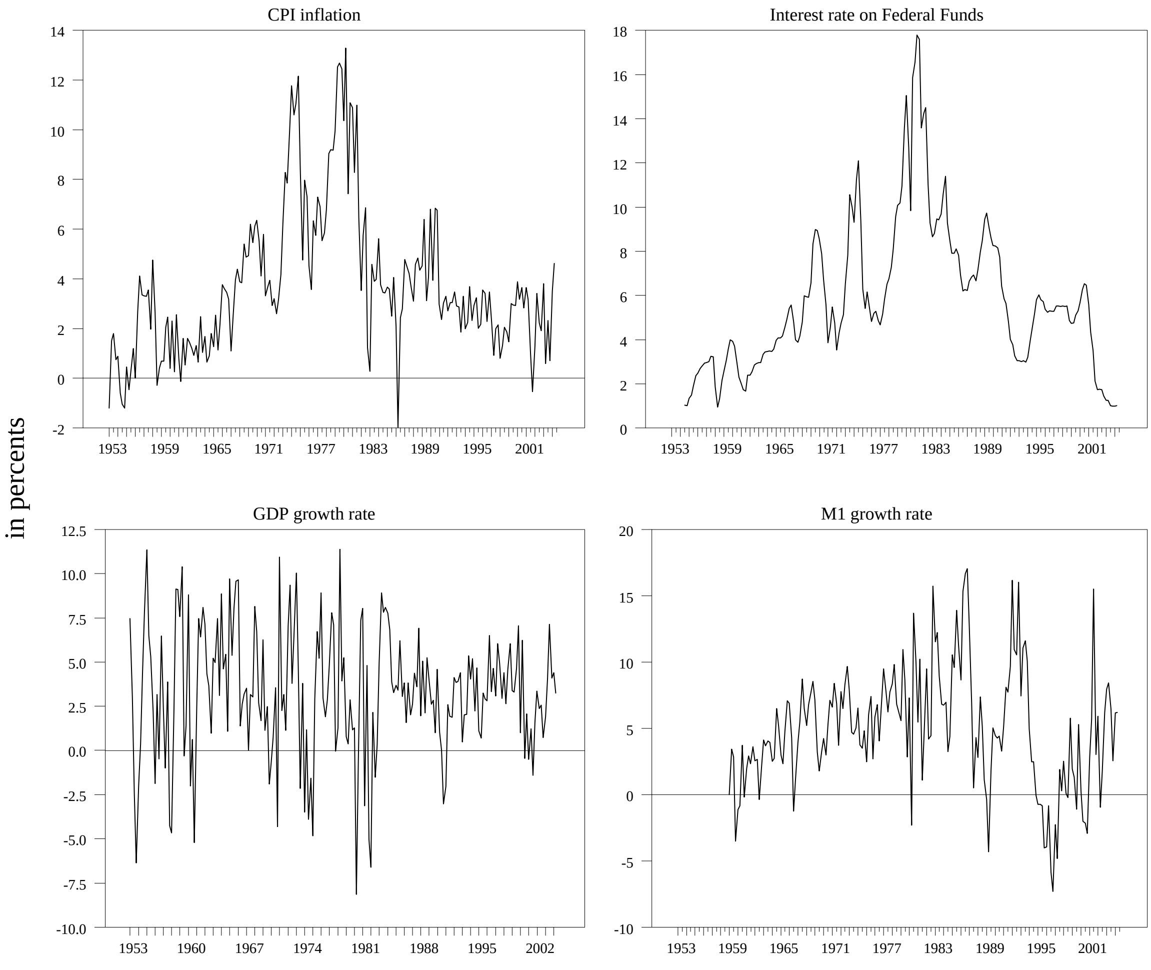


Figure 1b: Time series of selected US macroeconomic variables

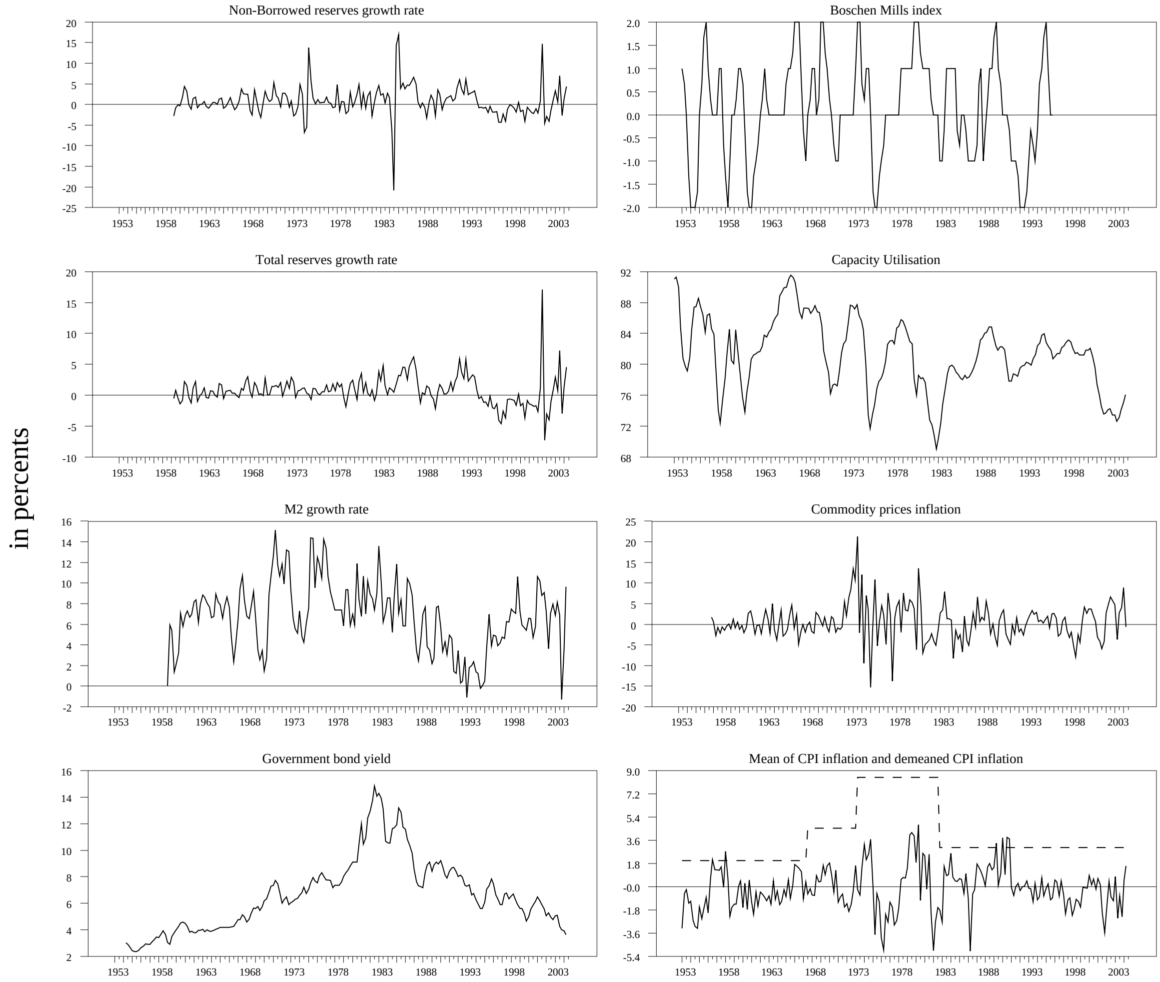


Figure 2: Effects of monetary policy when estimated for the 1960-2005 sample period

CEE-99a model, 4 lags; 10 th 50th and 90th percentiles of the impulse responses

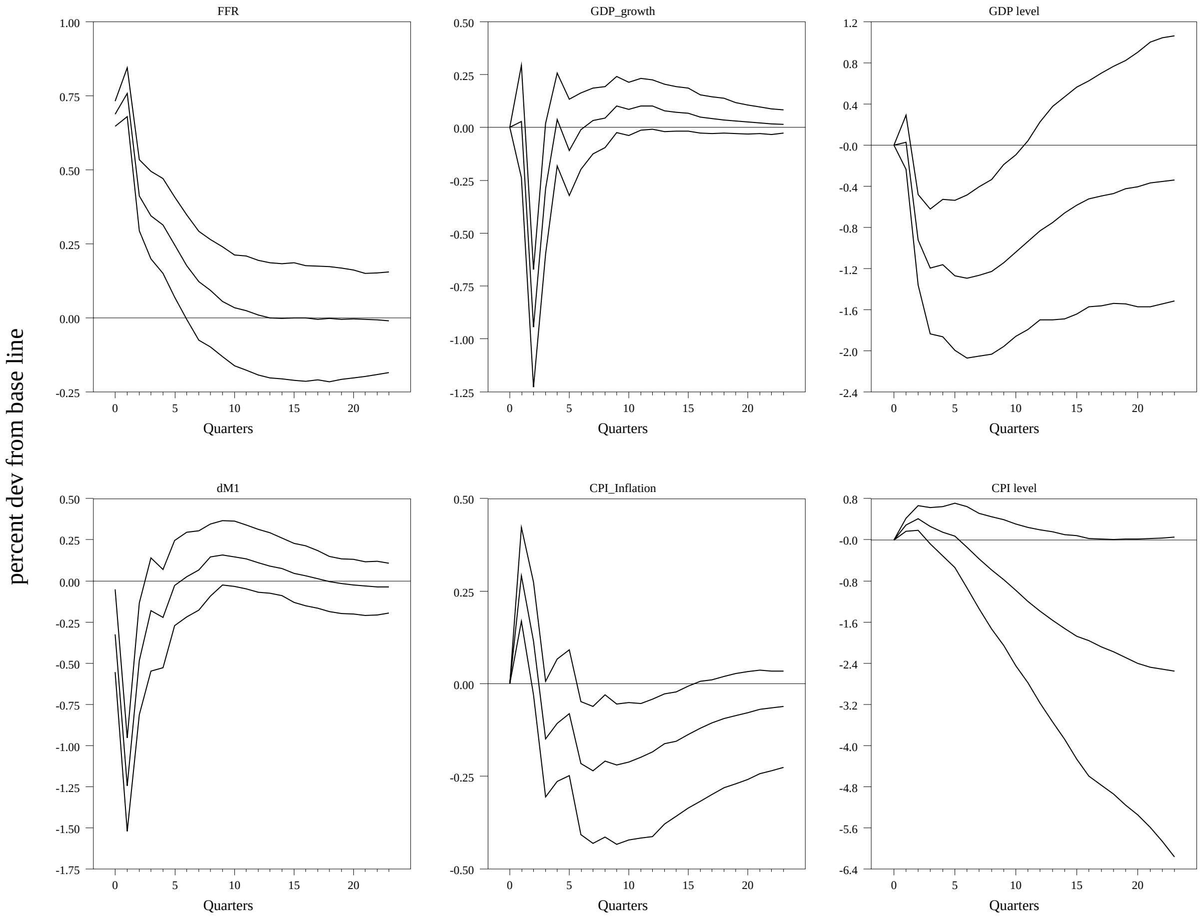


Figure 3: Effects of monetary policy, estimates for the 1984-2005 period

CEE-99a model, 4 lags; 10 th 50th and 90th percentiles of the impulse responses

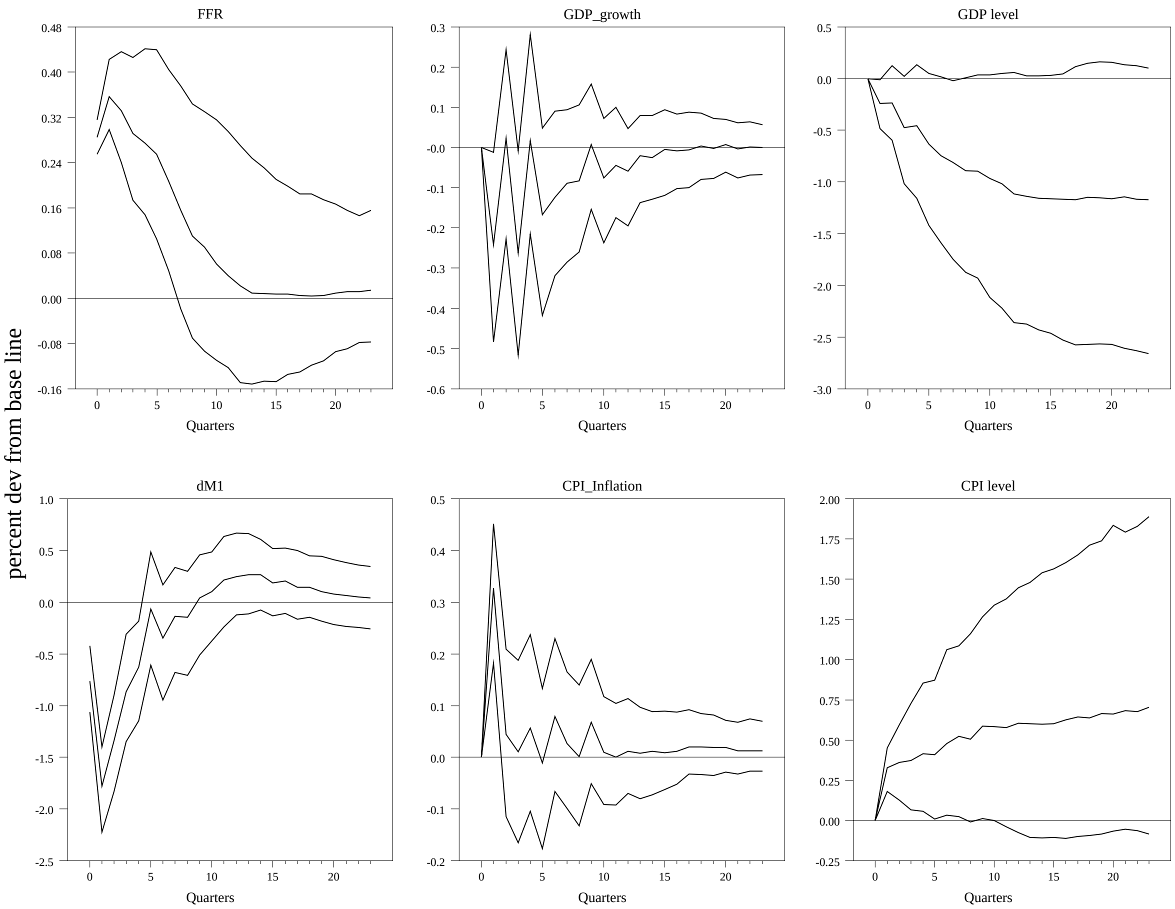


Figure 4: Controlling for changes in the intercepts at break dates

1960-2005; dotted lines, which correspond to constant intercepts, are similar to Figure 2

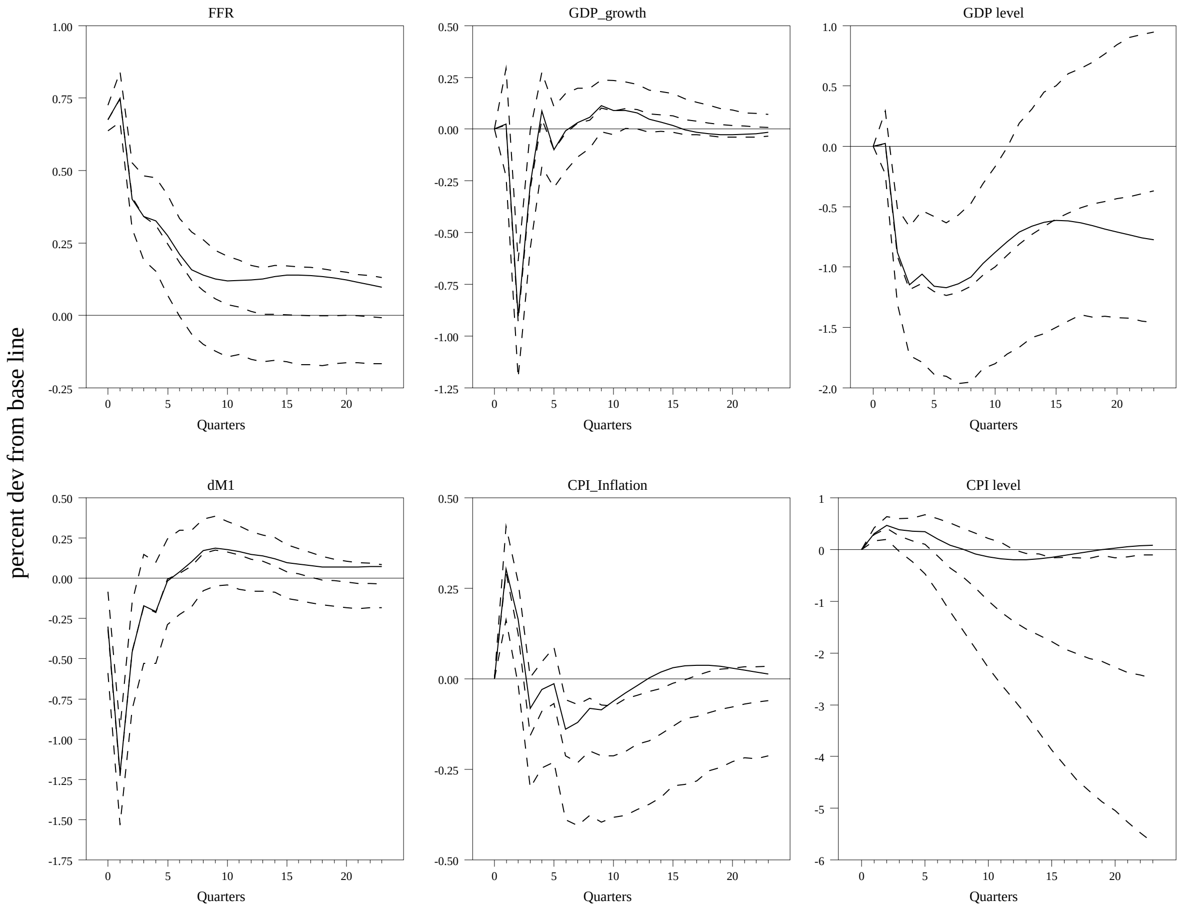


Figure 5: MA(5) of the CEE-99 monetary policy shocks and mean of inflation

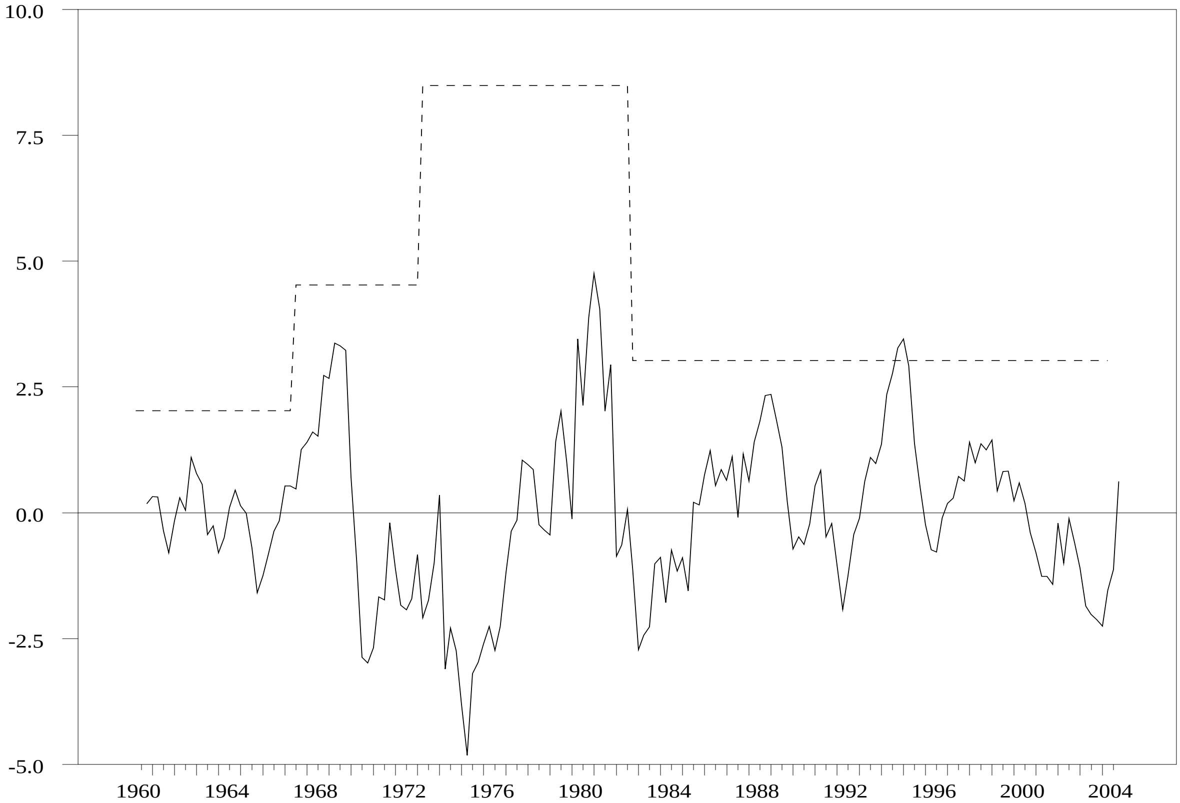


Figure A1: Effects of MP shocks in three versions of the monthly model

1984:1 to 2005:8 (full line); 1960:1 to 2005:8 without (broken line) and with changes in intercepts (dotted lines)

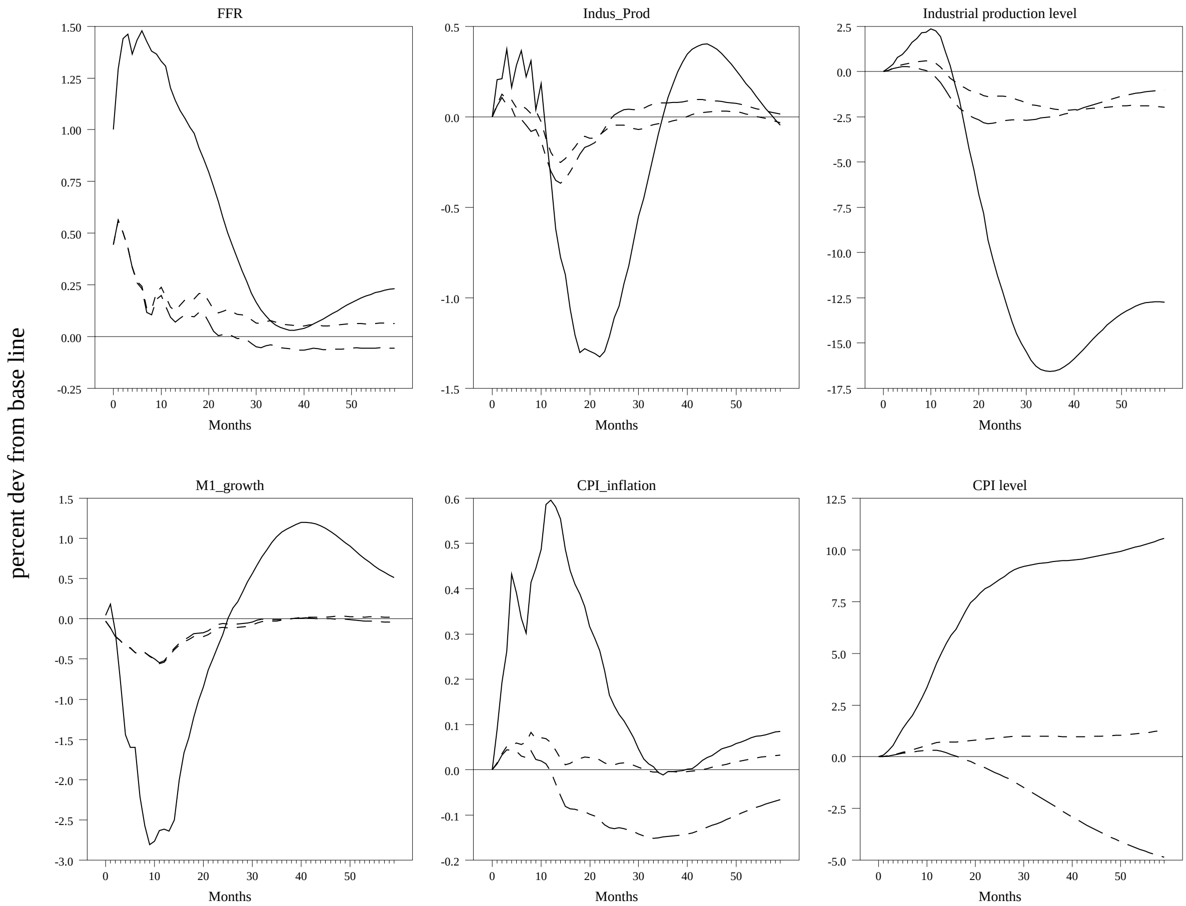


Figure A2: Effects of MP shocks in three versions of the GL model

1984:1 to 2005:8 (full line); 1960:1 to 2005:8 without (broken line) and with changes in intercepts (dotted lines)

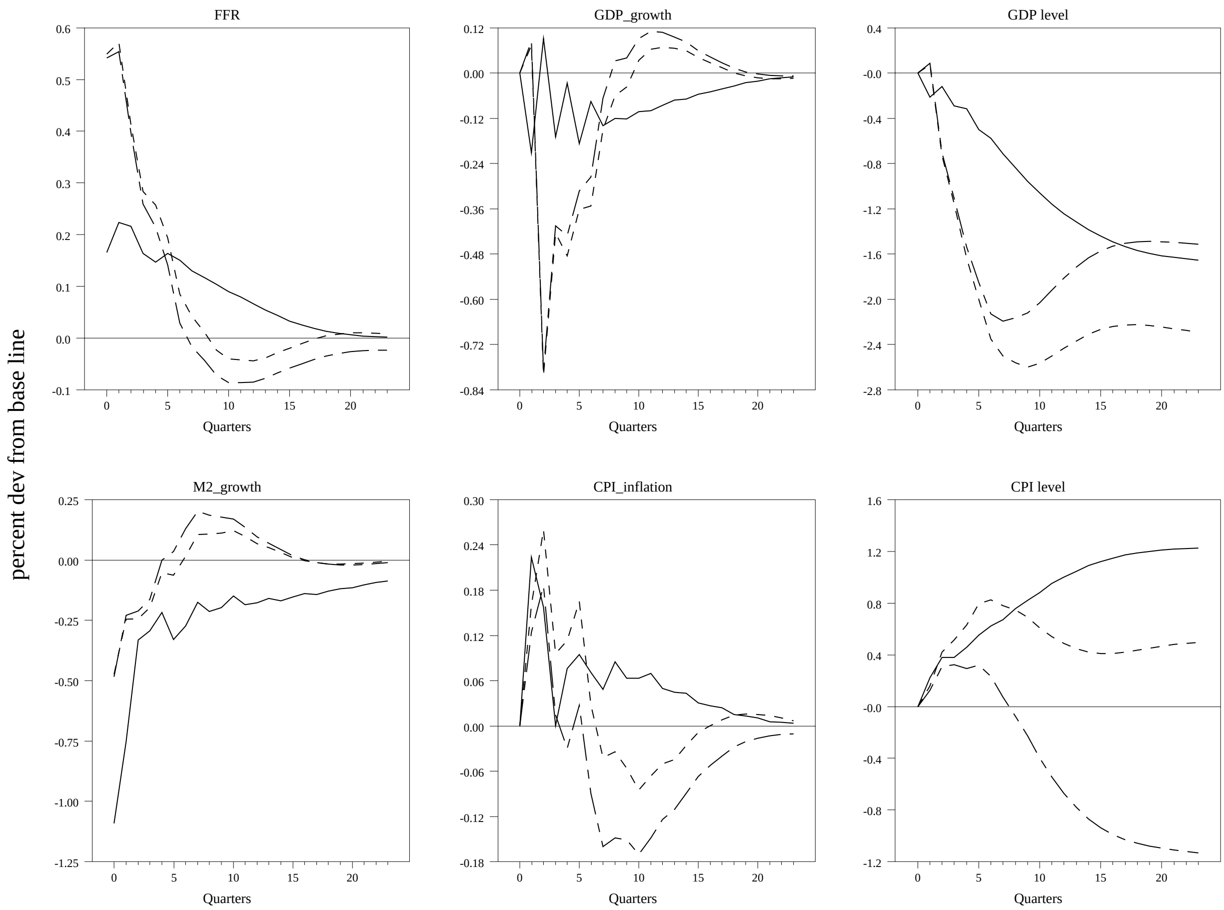


Figure A3: Effects of MP shocks in three versions of the CEE-01 model

1984:1 to 2005:8 (full line); 1960:1 to 2005:8 without (broken line) and with changes in intercepts (dotted lines)

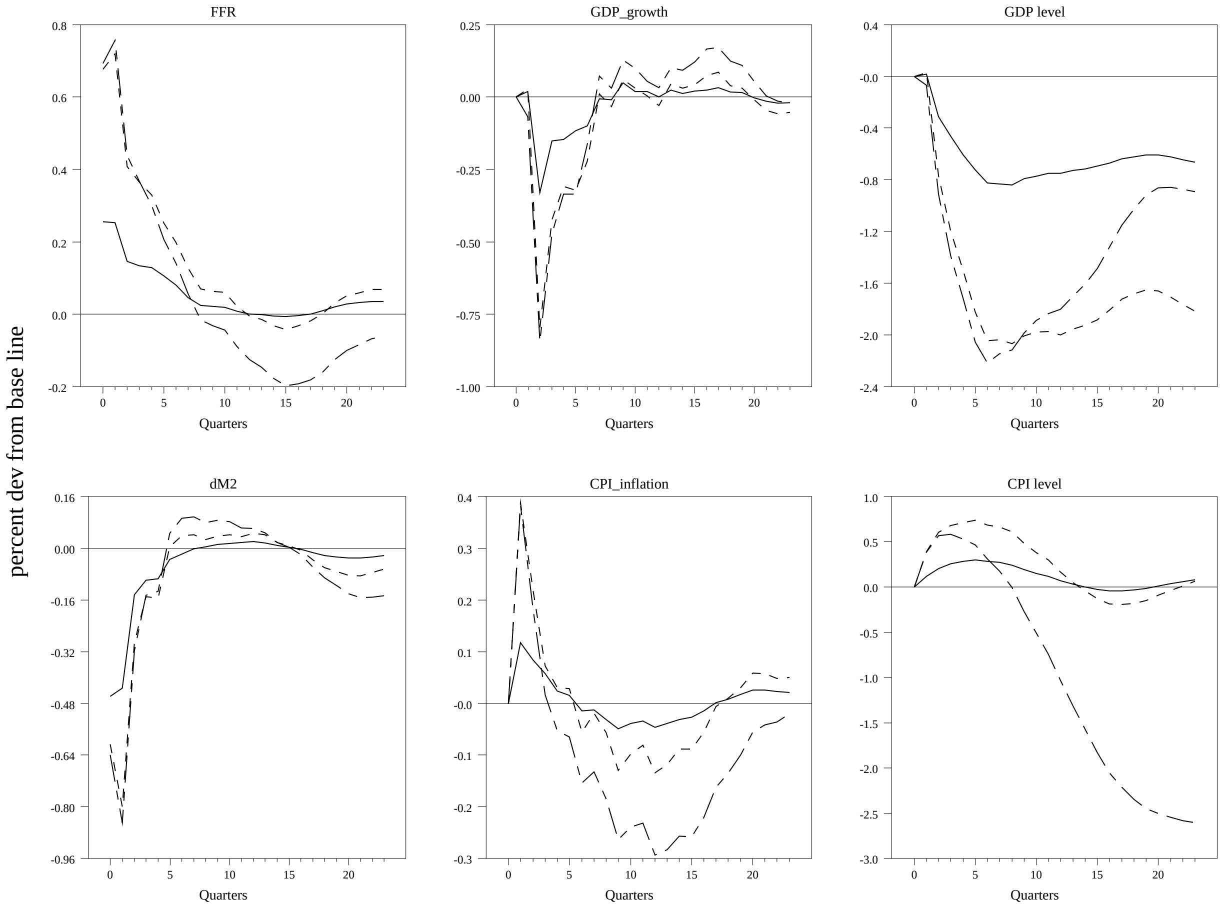


Figure A4: Effects of MP shocks in three versions of the Giordani model

1984:1 to 2005:8 (full line); 1960:1 to 2005:8 without (broken line) and with changes in intercepts (dotted lines)

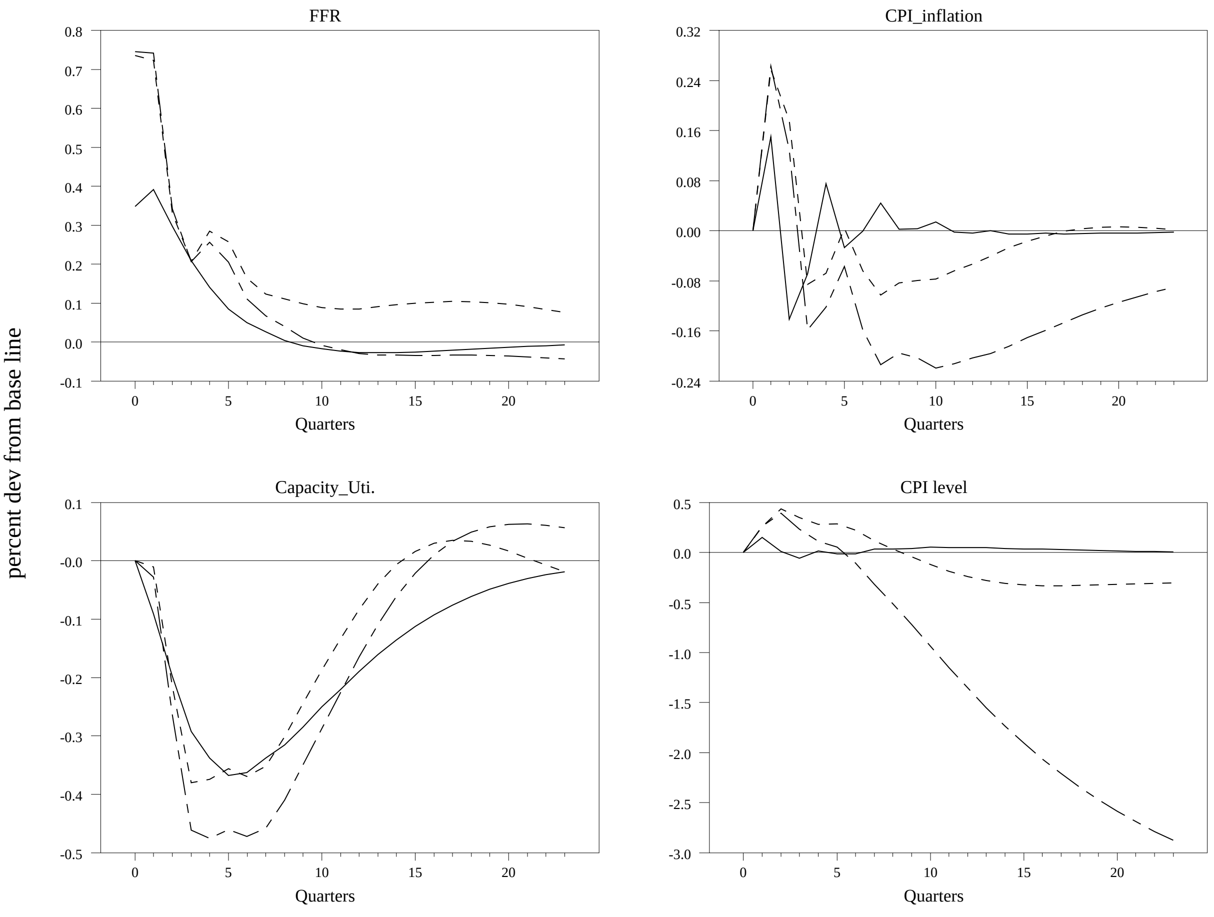
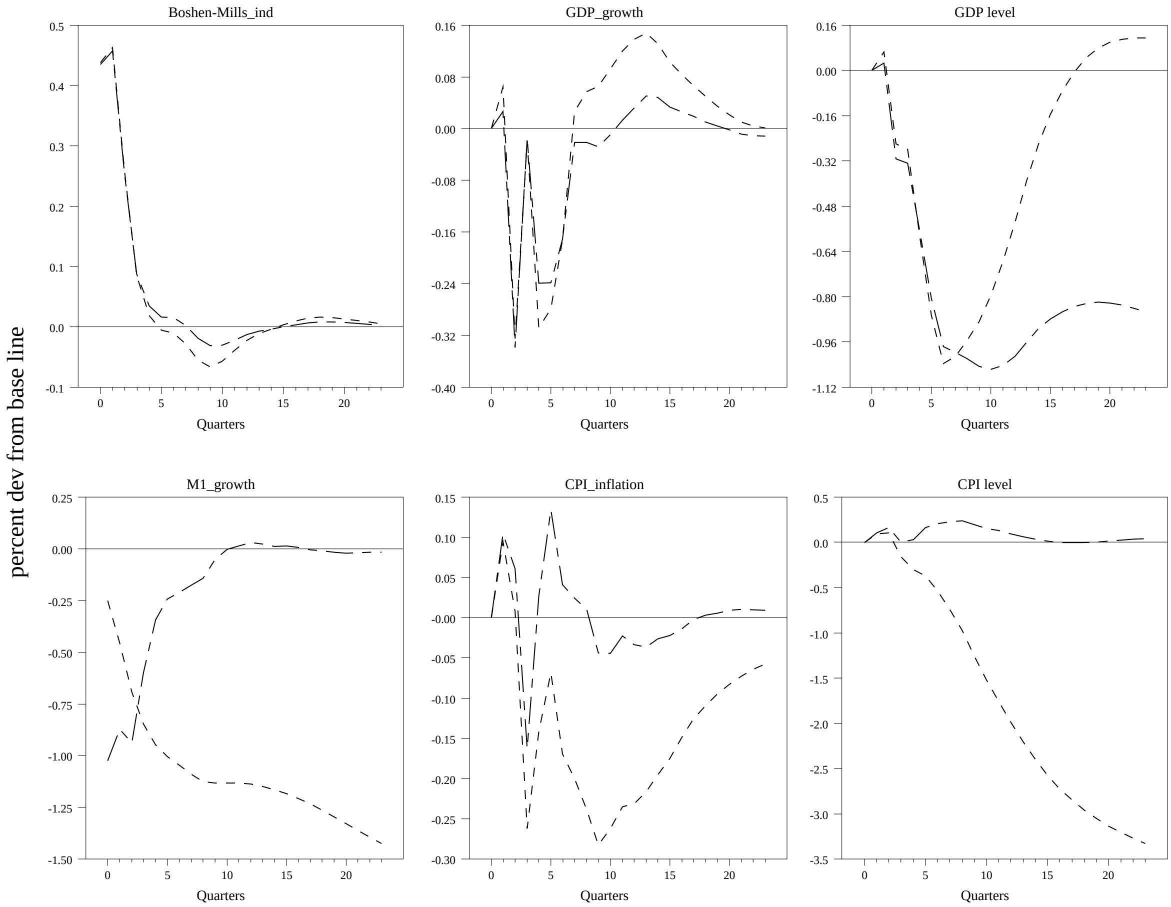


Figure A5: Effects of MP shocks in two versions of the BM95 model

1960:1 to 1995:4 without (broken line) and with changes in intercepts (dotted lines)



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