The Fed's Effect on Excess Returns and Inflation is Much Bigger Than You Think

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Abstract

We find that between 20 and 25 percent of the negative covariance between excess returns and inflation is explained by shocks to monetary policy variables. The finding is robust to changes in the monetary policy rule that have occured during the 1966-1998 period. The result contradicts the theory that money supply shocks induce a positive correlation between inflation and returns. Our findings also cast doubt on models that explain the negative correlation in a money-neutral environment (Boudoukh, Richardson, and Whitelaw (1994)), and on models that account for this correlation as being due solely to money demand shocks (Fama (1981), Marshall (1992)). We argue that contractionary monetary policy lowers excess stock market returns through various channels. Furthermore, if the Fed has some private information about future inflation, then a contractionary monetary shock will be followed by an increase in inflation, in the short run. The combined effect is a negative inflation/excess returns correlation. The results lend support to the argument that if asset pricing models are to capture the observed negative correlation, they must incorprate monetary policy effects.

JEL: G10, G12, E44, E51, E61, C32

1 Introduction

> "Greenspan and the powerful Open Market Committee can raise short-term interest rates to keep the economy from overheating. When they raise short-term rates, bonds and money market mutual funds look more attractive relative to stocks. And companies must pay higher rates on their borrowing, which reduces corporate earnings. Both of which, in theory, should bring stock prices down,..."
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> —USA TODAY, March 27, 2000.

People outside of academia take it for granted that the actions of the Federal Reserve Board (Fed) have a considerable impact on stock market returns, but a consensus amongst economists has yet to emerge. However, most economists will agree that if monetary policy is to have an effect on real returns, it must be either by influencing future net cash flows or by affecting the discount factor, at which the cash flows are capitalized. In either case, monetary policy must be "non-neutral", i.e. it must have an influence on real quantities, such as real dividend growth, in order to affect real returns. The focus of this paper is on the effect of monetary policy on excess stock returns, and on the correlation between those returns and inflation. Ours is a very natural direction of research, since the Federal Reserve System was founded by Congress precisely to promote price stability and to stimulate long-run economic growth[1]. To the extent that Fed policy is successful in fulfilling those goals, it must have an effect on excess returns and on the correlation between excess returns and inflation.

The effect of monetary policy on stock returns is of clear interest, as demonstrated by the constant analysis of Alan Greenspan's comments and the actions of the Federal Open Market Committee (FOMC) by economists, the media and Wall Street[2]. After the influential papers by Lucas (1976) and Sims (1980, 1986), the impact of monetary policy on financial and real variables has most often been evaluated using "weakly identified" vector autoregressions[3] (VARs) – large dynamic linear systems, allowing us to explore the statistical features of the data. If appropriate restrictions are placed on a VAR, its results might be given a structural interpretation, providing us with guidelines on what features asset pricing models should possess. Under the paradigm that only unsystematic monetary policy should have an effect on real variables (Lucas (1972, 1996), Sargent and Wallace (1975)), the emphasis has often been placed on analyzing the dynamic effects

of shocks on the estimated system. The analysis in Cornell (1983a,b) is in the same spirit, but uses different assumptions and econometric techniques to identify monetary policy shocks and their effect on stock prices. The VAR literature has recently produced results that conform to prior belief and economic theory, but monetary policy shocks have only been able to explain no more than 20% of the variation in real activity (Cochrane (1994), Bernanke et al.(1997)). Moreover, since real variables account for no more than half of the variation in stock returns (Geske and Roll (1983), Fama (1990)), it is not surprising that only a small percentage of the total variation of equity returns are explained by unanticipated policy shocks, as shown by Patelis (1997), Thorbecke (1997), and corroborated by our own results.

The negative correlation between excess returns and inflation has received independent interest, since this finding was published by Fama and Schwert (1977). The result is surprising in a money-neutral world, because if stocks are claims against real assets, the correlation must be *zero*. But the result is even more surprising if we allow monetary policy to have an effect on real activities. A successful contractionary monetary policy, implemented for instance by raising the short-term interest rate, must be followed by decreasing inflation and excess returns, resulting in a *positive* correlation between those two variables. Starting with Fama's (1981) "proxy" hypothesis, many people have tried to explain the observed negative correlation. Surprisingly, most explanations focus either on money demand effects (Fama (1981), Marshall (1992)), or on money-neutral environments (Boudoukh et al.(1994)). The effect of monetary policy on the relationship between returns and inflation has been given much less attention. The papers by Geske and Roll (1983) and Kaul (1987, 1990) consider the effect of systematic money supply shifts through various channels, and we discuss those papers below. The focus here is more in line with the current macroeconomics literature, that only unanticipated monetary policy will have an effect on excess stock returns and inflation. Moreover, the identification of monetary policy and its systematic effects on returns and inflation is also different from those in Geske and Roll, and Kaul.

We make the following contributions. Using traditional VAR analysis, we show that unexpected contractionary monetary policy results in a negative correlation between excess returns and inflation. In fact, during the 1966-1998 period, between 20 and 25% of the covariance of inflation and excess returns can be explained by monetary supply shocks. This finding suggests that monetary policy variables are an important factor and must be included in any asset pricing model that

aims to capture the negative returns/inflation covariance. To reach this conclusion, we assume that monetary policy shocks are identified in a recursive system, where the Fed follows a simple interest rate rule. More specifically, the policy instrument of the central bank is the federal funds rate, which responds systematically to inflation, output growth, and is subject to exogenous policy shocks. Then, using covariance decomposition, we find the percentage of the inflation/excess returns covariance that is explained by those shocks.

To understand why a monetary shock causes a negative returns/inflation correlation, we look at the separate effect of such a shock on excess returns and inflation. Unexpected contractionary policy leads to a decrease in excess returns, as is to be expected if monetary shocks have real effect and as stocks are claims against real assets. We discuss several channels, commonly thought to provide a transmission mechanism of monetary shocks onto real variates. We also observe that an unexpected contractionary monetary policy leads to a temporary increase in inflation, a fact that has been labeled the "price puzzle" by Sims (1992). Despite the usual precautions and variable selection in our VARs (such as including the price of commodities), the price puzzle is always present. In sum, the negative correlation caused by monetary supply shocks is inextricably linked to the price puzzle. There might be several reasons for the price puzzle, mostly associated with the asymmetric information between the Fed and the public. Under the assumption that the Fed has more information about future inflation than other players in the economy, we discuss a simple model that can account for the seemingly anomalous increase in prices following a contractionary monetary policy. Our results are consistent with the findings of Romer and Romer (2000) who provide a direct test of the Fed's superior information in forecasting future inflation.

The above arguments are based on the assumptions that only unsystematic monetary policy shocks can have real effects, that those shocks are correctly identified, and that the monetary policy rule has remained unchanged during the entire sample period. All of those assumptions have received some attention in recent years. However, some assumptions are more untenable than others. For example, it is highly unlikely that monetary policy has remained unchanged for more than 30 years. Changes in Fed chairmen, other Federal Open Market Committee (FOMC) members, and knowledge of the economy must have resulted in different monetary policy rules over the years (Clarida et al. (1999), Taylor (1998), Boivin (1999), and Friedman and Schwartz (1963a,b) for a historical perspective). In order to take the changing monetary policy into consideration, we

split the sample into two periods, 1966-1979 and 1983-1998, corresponding to tenures of FOMC chairmen, known to have conducted monetary policy in different fashions[4]. Indeed, the estimated monetary policy rule is very different in the two sub-samples. We re-estimate the VAR and analyze the effects of monetary policy on excess returns and on the correlation between excess returns and inflation during those two periods. This is a crude way of allowing the VAR policy parameters to change, without taking a stance on the provenance of those changes. A monetary policy does have an effect on the inflation/excess returns correlation in both periods, although the effects are less pronounced during 1983-1998. The price puzzle is remarkably stable during both periods. It must be noted that, given the changing monetary policy rule, it is highly unlikely that long-horizon VAR identifying restrictions, proposed by Blanchard and Quah (1989) and recently used by Hess and Lee (1999), can capture true monetary policy shocks[5]. The fact that excess returns do not respond to monetary shocks with the same magnitude during both periods and the fact that monetary policy influence on the covariance between returns and inflation changes in those periods prompts us to look deeper into how changes in the systematic part of the Fed policy will have effects on returns and inflation.

Changes in the Fed policy function may come from various sources. For instance, the central bank might be learning how to optimally respond to evolving economic conditions (Clarida et al. (1999)). Alternatively, the varying quality of Fed forecasts, due to changing economic conditions, might induce time variation in the policy function. This line of reasoning has led authors to consider monetary policy rules under uncertainly or imperfectly observed economic conditions. (Aoki (1999), Giannoni (1999)). Ideally, one would want to estimate a VAR with time-varying parameters, since as pointed out by Lucas (1976), a change in the monetary policy rule will result in changes of the parameters in the rest of the reduced form equations. However, a general time-varying parameter model is not easy to accommodate in the VAR framework, due to the sheer number of parameters to be estimated.

We specify the VAR so that time variation comes only from the monetary policy rule, whereas the rest of the system is time invariant. The coefficients in the VAR can easily be estimated by maximum likelihood, implemented with Kalman filtering. Time variation even in one equation will percolate throughout the entire system from the dynamics of the regressions. We estimate the VAR allowing for a changing monetary policy rule and examine the effects of monetary shocks on

the economy using time-varying response analysis. Comparing the responses of the system under time-varying and time-invariant policies provides a measure of how significant, statistically and economically, the variations in the Fed policy rule are. The impact of monetary disturbances on the excess returns/inflation covariance varies dramatically over time and the fluctuations are strongest during periods of economic instability, consistent with our account of the price puzzle. However, despite such drastic monetary policy changes, the effect of a contractionary monetary policy on excess returns is always negative, conforming to theory, and the effect on the returns/inflation covariance is also negative, supporting our previous findings. Surprisingly, we find that the impact of monetary shocks on the negative excess returns/inflation correlation has been steadily declining.

The paper is structured as follows. In section 2, we present the Fed policy function, discuss its changing character, and elaborate on the effect of monetary disturbances on stock returns and inflation. Section 3 lays out the methodology used in the empirical investigation. First, we discuss the usual VAR toolbox and a useful extension; covariance decomposition. Then, we present the estimation and dynamic response analysis of a VAR with time-varying interest rate equation. The results are analyzed in section 4. Section 5 offers concluding remarks.

2 Fed policy, excess returns, and inflation

Monetary policy may affect both the cash flows of firms and the discount factor at which the cash flows are discounted. A considerable fraction of the variation of annual stock returns is due to real variables such as industrial production and investment, which are important determinants of the cash flows of firms (Fama (1981), Geske and Roll (1983), Kaul (1987), and Fama (1990)). To the extent that monetary policy affects real variables, it will also have an effect on firm cash flows and returns. Monetary policy may also affect the discount factor at which cash flows are discounted, by influencing the risk structure of the economy. For instance, Schwert (1989) documents that stock market volatility is generally higher during recessions than during expansions and shows that the spread between lower- and higher-grade corporate bond yields is directly related to subsequently observed stock volatility. As discussed by Schwert, this result might be due to the increasing operating and financial leverage of firms during recessions. Therefore, to the extent that monetary policy shocks affect a firm's financial health, they may also affect the expected risk premium demanded by investors.

Using the framework of Campbell and Shiller (1988), Campbell (1991), and Patelis (1997), the excess stock returns – the spread of nominal stock returns over nominal risk-free interest rate – may well be decomposed into expected future sums of dividend growth, real interest rate, and excess returns. In a money-neutral setting, these real variables are solely determined by real factors such as productivity of capital, time preference, and risk aversion. However, if money is assumed to have real effects, then monetary shocks will affect excess returns through the three expectations, as demonstrated by Patelis (1997). Since the main focus of this paper is on demonstrating that part of the negative covariance between excess returns and inflations is due to monetary shocks, we are only interested in establishing that those shocks do have an effect on excess returns and in providing an explanation for the results. The more daunting task of identifying the exact channels through which unanticipated monetary policy affects the covariance of interest is left for further research[6].

The positive correlation between short interest rates and inflation[7] is a well documented empirical fact. This robust empiricism might be interpreted as evidence supporting the Fisher relation. However, we think of the federal funds rate *not* as a state variable, but rather as being controlled by the Fed. The FOMC uses the federal funds rate as its instrument of monetary policy[8] in order to promote price stability and economic growth. In this respect, an exogenous shock to the federal funds rate (tighter monetary policy) must result in *lower* future, expected and realized, inflation. However, as mentioned above, increases in the funds rate have historically been followed by increases in the price level at the short horizon, followed by a decrease at longer horizons. Following Romer and Romer (2000), we attribute this fact to the Fed's superior information about future economic conditions. The Fed possesses private information about future inflation that is not captured by a VAR. To the extend that the Fed is unable to offset the inflationary impetus that led it to predict higher (than the public) inflation, the unanticipated (by the public) increase in the federal funds rate will be followed by (will forecast) a rise in inflation. The persistence in the inflation series, due to the rigidity in wages and contracted prices, as well as the Fed's desire to smooth interest rates, results in a short-horizon positive correlation between monetary policy shocks and inflation.

Summing the effects of unanticipated monetary policy on returns and inflation, we have the desired result. In the short run, a tighter monetary policy, implemented by a positive shock to the federal funds rate, induces a decrease in excess rates of return and is followed by an increase in

inflation. Hence, such a policy will deliver the observed negative correlation between excess returns and inflation.

One might argue that our mechanism is one of the many that produce the negative correlation between inflation and returns. However, the reasoning above implies that a big part of the negative correlation between inflation and excess returns must be explained by unexpected shocks to monetary policy variables. This dimension of the correlation is not captured by any other model. As discussed below, we find that between 20% and 25% of the covariance between excess returns and inflation is due to monetary policy shocks. Therefore, models that imply that the negative correlation is entirely due to money demand shocks (Fama (1981), Marshall (1992)) are not supported by the data, since we find that money supply shocks account for a significant fraction of that variation. Moreover, our findings suggest that models that treat money as being completely neutral in the short run also do not offer the complete explanation (Boudoukh et al.(1994)). Lastly, under the assumption that time-variation in the VAR comes only from the policy function, we find that contractionary monetary disturbances are followed by lower excess returns and a negative returns/inflation covariance during any time period, which is consistent with the previous discussion.

The rest of the section is structured as follows. First, we discuss the Fed policy function (Taylor rule) and lay out several channels through which monetary shocks can be transmitted to real variables and stock returns. Next, we describe the asymmetric information problem between the Fed and the public, within the framework of the Taylor rule. Finally, we discuss instabilities in the policy function.

2.1 The Fed's policy function

In recent years, a lot of attention has been devoted to the issue of how to conduct monetary policy. This increased interest can be credited to the success of recent empirical papers showing that monetary shocks do have an impact on the course of the real economy[9]. A virtual consensus in this literature is that monetary policy can be characterized by looking at the federal funds rate rather than at monetary aggregates, such as M0, M1, or M2. In other words, VAR's that include federal funds rate (or nonborrowed reserves) exhibit impulse response functions that conform to our previous beliefs and economic intuition. Therefore, in this study, we take the federal funds rate to be the main policy instrument of the central bank. In this respect, our analysis differs from

the ones in Geske and Roll (1983) and Kaul (1987, 1990) who look at the growth of monetary aggregates.

On the theoretical side, the monetary policy literature has enriched the dynamic general equilibrium models, developed in the real business cycles field, by introducing frictions, such as nominal price rigidities. A significant product of this literature is the development of specific rules of Fed policy that are justified by general equilibrium considerations, and supported by the data. A popular class of rules are the simple linear interest rate functions, also known as "Taylor rules", after the article by John Taylor (1993). Taylor suggests that the Fed should try to set the short rate as:

$$r_t^{ff} = \alpha + \gamma\left(\pi_t - \pi^*\right) + \eta x_t + v_t$$

where r_t^{ff} is the federal funds rate, π_t is the level of inflation, π^* is the target inflation, x_t is the log deviations of GNP from its trend, and v_t is a policy shock. In the same spirit, and using insights from theoretical work by Rotemberg and Woodford (1997) and Woodford (1999), Clarida et al.(1999) propose a "forward looking" version of the Taylor rule that also exhibits interest rate smoothing, often observed in the Fed's behavior. The feedback rule, proposed by Clarida et al.(1999) is (setting $\pi^* = 0$):

$$r_t^{ff} = \alpha + \phi r_{t-1}^{ff} + (1-\phi)\left[\gamma E\left(\pi_{t+1}|I_t\right) + \eta x_t\right] + v_t \tag{1}$$

The parameter ϕ captures the degree of interest rate smoothing. Expectations are taken with respect to the information set I_t available to everyone in the economy: the public and the Fed. In what follows, we assume that the Fed's systematic policy is captured by a Taylor rule, as in (1). A contractionary policy shock is captured by a positive innovation, v_t.

2.2 The Fed's impact on real variables and stock returns

We focus on two channels, through which the Fed's actions can be transmitted onto real economic variables and excess returns. The first one is a traditional IS-LM or a "money" channel, which relies on price rigidities. The second channel, often called a lending or a credit channel, relies on capital market imperfections. We do not try to differentiate between the two effects, since it has proven to be a difficult task with even more disaggregated data. The presented evidence is consistent with both channels.

Traditional IS-LM channel: Brainard and Tobin (1963), and later Fama (1980), convincingly argue that the impact of monetary policy on the economy can be analyzed through its effect on investor portfolios. Suppose that we are in a Fama (1980) banking world, where banks are merely a medium of rebalancing portfolios. The analysis can be conducted with only two financial assets: "outside" money, provided monopolistically by a central bank, that serves as the numeraire good and as a medium of exchange, and "bonds". We assume that money and bonds are not perfectly substitutable, and that there is a non-zero demand for money. In an unexpected contractionary move, the central bank decreases the quantity of money. Households must then hold more bonds and less money in their portfolios. If there is price-rigidity in some sectors of the economy, prices do not fully (or instantaneously) adjust to changes in money supply, and the fall in money holdings represents a decline in real money balances. To restore equilibrium, the real return on bonds must increase. Thus, fewer projects are available at higher required rates of returns–investment and industrial production decrease.

The mechanism described above hinges on the central banks's ability to control the supply of outside money, and on prices being somewhat rigid. Since, both of those assumptions are uncontroversial for the US, most researchers would agree that monetary policy is not neutral, at least in the short run. However, the VAR literature has produced "protracted, hump-shaped and large[10]" responses of real variables to unexpected monetary shocks that are difficult to believe. Many recent papers have approached this problem in different fashions[11]. In the next section, we discuss an already large literature that analyzes frictions in capital markets as possible channels that amplify and propagate the effects of monetary policy.

Capital market-imperfections channels: The capital market-imperfection channels provide additional mechanisms of translating monetary policy actions into variations in real variables. However, it is a mistake to think that the two views are anything else but complementing and reinforcing each other. In this extensive literature[12], monetary policy impacts the difference in cost between external funds (issuing equity or debt) and internal funds (retained earnings), known as the external finance premium. In order to analyze market imperfections, we need to introduce three financial assets: money, bonds, and bank loans.

The banking literature has proposed two linkages between monetary policy actions and the

external finance premium. The first one, known as the *balance-sheet* or net-worth channel, emphasizes that an increase in interest rates weakens the financial conditions of consumers and firms, by directly impacting their cash flows, net worth, and assets. For example, higher rates of interest will result in lower present discounted value of a firm's assets (equipment, buildings, etc.), which are often used as a collateral for loans. Therefore, such a monetary action will exacerbate any existing agency and information costs of issuing credit and will result in the firm having a reduced access to bank loans (Bernanke et al.(1994)). Balance sheet constraints also lead to lower investment and, ultimately, decreasing rates of return. The second channel, suggested by Bernanke and Blinder (1988), and further analyzed by Kashyap et al. (1993), is the *bank lending* channel. The main idea is that a reduction in bank reserves by the Fed also reduces bank deposits and, hence, banks' loanable funds. If bank loans are imperfect substitutes for other forms of financing (such as commercial paper), a reduced supply of bank loans will lower economic activity by bank-dependent borrowers. As an example, consider a company, whose primary source of short-term debt financing are bank loans. A contractionary monetary policy, resulting in a reduction of loanable funds, will compel this firm to look for credit in the commercial paper market. To the extent that bank loans and commercial paper are not perfectly substitutable, and if the commercial paper market is not sufficiently "deep", then a contractionary monetary policy will result in an increase of commercial paper rates[13], and in the total cost of short-term debt financing of the firm. (Kashyap et al. (1993)). Therefore, the cash flow of indebted firms will decrease, resulting in lower stock returns.

In sum, a tighter monetary policy, working through the two channels, has the effect of reducing the excess rate of return in financial markets. It is also likely that the combination of those two channels induces monetary shocks to have a more "protracted" effect on economic variables. As a result of their similar effects on real variables, the channels and their nuances are difficult to identify in aggregate data (Friedman and Kuttner (1993), Oliner and Rudebusch (1993), Cecchetti (1995), Bernanke (1995)). However, in our study, we are not focusing on identifying those channels: our aim is to establish that a tighter monetary policy ultimately results in a decrease of excess returns. We include the default premium, the spread between yields on Aaa and Baa rated bonds, as a simple proxy for capital market imperfections.

2.3 The "Price Puzzle" and the Fed's information set

In a VAR framework, if the weak restrictions that identify the structural shocks are inappropriate or if there is crucial information missing from the vector of variables, then the system will exhibit dynamics that are counter-intuitive to established beliefs. A well-known example is the puzzling increase in the price level to a contractionary shock of monetary policy in a recursive system. This result, first observed in U.S. and international data by Sims (1992), has been dubbed the "price-puzzle". Sims (1992), Sims and Zha (1995), and Christiano et al. (1996,1998) all suggest that this anomalous result obtains because the Fed's reaction function contains information about inflation that is missing from the consumer price index (CPI). To remedy this problem, the authors include commodity prices or the producer price index (PPI). Those measures alleviate the problem, but the price-puzzle is still present in the short-horizon, for up to 6-12 months[14]. Romer and Romer (2000) argue that the Fed has a considerable amount of private information beyond what is available to commercial forecasters (and everybody else). This viewpoint is stronger than the one in the above cited papers, because the information that needs to be included in the VARs is not only omitted but also unavailable to forecasters (or the public).

These arguments can be incorporated in our discussion about the Taylor rule in the following way. Suppose we denote by I_t^c the information set available to forecasters and the public, and $I_t^c \subset I_t$, where I_t is the information set of the Fed. Then, using the fact that $E(\pi_{t+1}|I_t) = E(\pi_{t+1}|I_t^c) + E(\pi_{t+1}|I_t \backslash I_t^c)$, we can write the Taylor rule as:

$$r_t^{ff} = \alpha + \phi r_{t-1}^{ff} + (1-\phi)\left[\gamma E(\pi_{t+1}|I_t^c) + \gamma E(\pi_{t+1}|I_t \backslash I_t^c) + \eta x_t\right] + v_t \tag{2}$$

where v_t is a monetary policy shock. On the other side, the forecast of the public, denoted by r_t^p is:

$$r_t^p = \alpha + \phi r_{t-1}^{ff} + (1-\phi)\left[\gamma E(\pi_{t+1}|I_t^c) + \eta x_t\right] + \varepsilon_t \tag{3}$$

where $\varepsilon_t = \gamma E(\pi_{t+1}|I_t \backslash I_t^c) + v_t$ is the monetary policy shock, identified in the literature. In our information asymmetry hypothesis, ε_t is comprised on the following two components: (i) the Fed's policy endogenous adjustment to $E(\pi_{t+1}|I_t \backslash I_t^c)$, i.e. the Fed's assessment of future inflation based on its private information, and (ii) exogenous policy disturbances. For the public and econometricians, both (i) and (ii) come as unanticipated policy shocks that have been found to affect real and

financial variables. However, the unanticipated (to the public) shocks in fact contain (i), which is merely a systematic adjustment of the Fed to its private information. Thus to the extent that ε_t contains $\gamma E\left(\pi_{t+h}|I_t\backslash I_t^c\right)$, the unanticipated monetary policy shocks predict future inflation, and thus the price puzzle obtains. To put this differently, the price puzzle can be attributed to the Fed's systematic reaction to private assessment of future inflation, and to the inability of the public to disentangle between "true" policy shocks, v_t, and systematic policy reaction.

Our story is a middle ground between the arguments in Sims (1992) and Romer and Romer (2000). We agree with Sims (1992) that the price puzzle is due to the Fed's information about future inflation that is missing from the econometrician's information set. However, we espouse the views in Romer and Romer (2000) that part of the Fed's information is not available to the public and cannot be proxied by available data. In other words, we argue that variables, such as commodity prices, are part of I_t^c and cannot be used to proxy for $E\left(\pi_{t+1}|I_t\backslash I_t^c\right)$. There are several reasons why the Fed might have information that is unavailable to other forecasters or to the public. Romer and Romer (2000) show that if one were to form an expectation of future inflation using a weighted average of internal Fed forecasts and private forecasters, all the weight should be placed on the former. Moreover, the Fed has a large and competent staff of researchers in charge of forecasting inflation, who often have an early access to relevant economic data.

Under the assumption that the available data cannot proxy for $E\left(\pi_{t+1}|I_t\backslash I_t^c\right)$, it is difficult to formulate a direct test of our asymmetric information hypothesis[15]. The VAR results, discussed below, are consistent with our hypothesis, but cannot be taken as anything more than indirect support. As a final note, we tried to keep the asymmetric information story as simple as possible. There are several directions for future work. For example, instead of assuming that the public has no information about variables in $I_t\backslash I_t^c$, one might think of a setup, where some of the information is known, and the public is attempting to forecast the forecast of the Fed. For such a model in a different context, see Townsend (1983).

2.4 Changing policy function

It would be unreasonable to assume that the conduct of monetary policy has remained unchanged during the 1966-1998 period, which encompasses 5 chairmen of the Board of Governors[16], 6 business cycles[17], and countless academic articles on monetary policy. There are several excellent papers

documenting the various monetary policy periods. Before 1966, the Fed relied on monetary aggregates, such as M1 and M2 as strategic targets for monetary policy. The first three years of Volcker are known as the "Volcker experiment"; the central bank adopted a pure non-borrowed reserve targeting strategy. During the last 5 years of Volcker and during Greenspan, the Fed switched to federal funds rate targeting as an operating procedure. An empirical study of the effect of monetary policy on stock returns and inflation must take into account such changes in the monetary policy rule. Such changes might induce rational agents to behave differently, thereby also changing the rest of the parameters in the system, as argued in the Lucas (1976) critique. An easy way of taking into account the time-varying Fed policy is to estimate the VAR for two different periods, 1966:1-1979:6 and 1983:1-1998:12. The "Volcker experiment" period is too short to be analyzed with a VAR. Table 1 presents the results from estimating the Taylor rule (1) for the entire sample, and for various sub-samples. The policy function during Volcker-Greenspan is very different from the one in previous years, as documented by Clarida et al (1999). Moreover, the 1979:7–1982:12 period of the "Volcker experiment" seems to be particularly peculiar.

[Table 1 about here]

The above analysis would have been adequate if we were ready to argue that monetary policy changes occur only because of changes in FOMC members. However, this is not the case and changes in the reaction function need not be regime shifts. As discussed above, it might be that the Fed is slowly learning how to optimally respond to changing economic conditions (Clarida et al.(1999)). Alternatively, it might be the case that the Fed's policy varies because, in different states of the world, it has better (or worse) estimates of present and future economic conditions (Aoki (1999), Giannoni (1999)). Changes in the monetary policy rule might also have an impact on the correlation between returns and inflation. For instance, a less effective monetary policy, i.e. a monetary policy that does not have real effects, will result in a zero correlation between excess returns and inflation. Also, a more transparent monetary policy, where there is no information asymmetry, should lead to the resolution of the price puzzle and in a positive correlation between excess returns and inflation, if the price puzzle were really caused by such asymmetric information.

A time varying monetary policy function is captured by:

$$r_t^p = \alpha_t + \phi_t r_{t-1}^{ff} + (1-\phi_t)\left[\gamma_t E\left(\pi_{t+1}|I_t^c\right) + \eta_t x_t\right] + \varepsilon_t$$

Figure 1 plots the estimated coefficients γ_t and η_t and compares them with their time invariant analogues and with the case when the function is estimated for the sub-periods 1966:1–1979:6, 1979:7–1982:12, and 1983:1–1998:12. It is interesting to notice that the variation in the policy function is substantial and not entirely captured by regime shifts. For simplicity, the estimation is carried out by using $\pi_{t,}$ instead of π_{t+1} or $E\left(\pi_{t+1}\right)$. Since inflation is a persistent process, π_t is a good proxy for expected inflation. Boivin and Watson (1999) use a more sophisticated method, accounting for possible inconsistency, arising from using forecasts of future variables, which might be correlated with the residual in the regression. Their results (figure 2) are similar to ours, despite the fact that they use different variables at different frequency. To capture the forward looking behavior in (1), we also include DEFP and TERM in the VAR, as discussed below.

[Figure 1 about here]

3 Data and Methodology

3.1 Data

We work with a system of eight variables, available at monthly frequency: industrial production growth (IPG), consumer price inflation (INF), commodity price inflation (DPCOM), federal funds rate (FF), growth of non-borrowed reserves (DNBRD), default premium (DEFP), term spread (TERM), and excess market return (EP). The exact data description and sources can be found in Appendix A. The use of IPG as a proxy for theoretical dividend growth has been motivated by Fama (1981,1990), Geske and Roll (1983), and Boudoukh et al. (1994)[18]. As forward-looking variables, DEFP, TERM, and EP enable the VAR to span larger information set than usually considered in this literature. The power of these spreads to predict future economic activity and inflation is well documented in the literature (e.g., Chen et al. (1986), Stock and Watson (1989), Bernanke and Blinder (1992)).

The sample period runs from 1966:01 to 1998:12. The federal funds rate was not the main instrument of monetary policy before 1966. We also divide the sample into two periods: (i) 1966:01

to 1979:06 (pre-Volcker), and (ii) 1983:01 to 1998:12 (post-Volcker). The first three and a half years of the chairmanship of Volcker (1979:07 – 1982:12) are omitted, because of substantial changes in the Fed's operating procedure (see Table 1 and Figure 1)[19].

3.2 VAR Analysis

3.2.1 Isolating Monetary Policy Shocks using the VAR

The aim of this paper is not to write down a fully specified asset pricing model; doing so would require making some restrictive assumptions. Given the apparent lack of fit of available general equilibrium models with the time-series facts, our goal is to impose as little a priori assumptions on the data as possible. In the spirit of recursive weakly-identified VARs, we only assume a contemporaneous recursive relationship between the variables in order to isolate monetary policy shocks. First, we fit an unrestricted VAR system to the data:

$$y_t = \phi + \sum_{l=1}^{m} \Phi_l y_{t-l} + u_t, \quad E\left(u_t u_t'\right) = V \tag{4}$$

where the residuals u_t are assumed to be serially uncorrelated with a covariance matrix V, ϕ is a vector of constants, and m denotes the VAR lag order[20]. To give some structural interpretation to the fitted VAR system, we decompose the VAR residuals u_t into unobserved economic shocks, w_t, and assume that the economic shocks are mutually and serially uncorrelated. The relationship between u_t and w_t can be written as $u_t = Gw_t$, where G is a square matrix that imposes the identifying restrictions. In this paper, we assume a recursive contemporaneous (or Wold) ordering (e.g. Sims (1986) and Christiano et al. (1996,1998), among others), or G is a unit diagonal lower triangular matrix. Those restrictions imply that economic shocks have a contemporaneous effect only on variables placed at the same level, or lower, in the system.

The ordering of the variables in the VAR is: IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP. In other words, we treat IPG, INF, and DPCOM as predetermined for monetary policy shocks, as in Christiano et al. (1996, 1998) and Thorbecke (1997)[21]. The financial variables DEFP, TERM, and EP are placed below the policy variables, implying that they respond to monetary policy shocks within the same month[22]. In the VAR literature, monetary policy shocks have been identified as innovations in federal funds rate or non-borrowed reserves[23]. We use both federal funds rate and the growth of non-borrowed reserves as our policy variables, since they both may

contain information about the stance of the Fed (Leeper and Gordon (1992), Bernanke and Mihov (1998)).

The above assumptions are just enough to let us recover (or identify) the unobservable economic shocks w_t from the VAR residuals u_t. Let $E(w_t w_t') = D$, where D is a diagonal matrix with the variances of the economic shocks on the principal diagonal. Then, we have the set of restrictions

$$V = GDG'$$

It is often convenient to let $P = GD^{1/2}$ or $V = PP'$, where P is the lower triangular Choleski factor of V.

The results presented below are fairly robust to the choice and ordering of variables in the VAR. We have experimented with several alternative identifying restrictions. In fact, the identification imposed by P, although not very restrictive, can be relaxed even further. The variables in our analysis can naturally be grouped into three sets: macro variables (IPG, INF, DPCOM), policy variables (FF, DNBRD), and financial variables (DEFP, TERM, and EP). If there are only three structural, uncorrelated shocks, then we can write $V = \widetilde{G}\widetilde{D}\widetilde{G}'$, where $\widetilde{G}$ and $\widetilde{D}$ are block lower triangular, and block diagonal, respectively. Such restrictions provide a robustness check of the results, since the ordering of variables within a given group would not have an impact on the analysis. In the next section, we show that V, $\widetilde{D}$, and D are very similar, thus confirming our robustness claims. We also tried several other specifications by omitting the variables DEFP, TERM, DPCOM, in different combinations. Placing the financial variables DEFP, TERM, and EP prior to the policy variables did not quantitatively change the results. In a VAR without DPCOM, the price puzzle is much more persistent. Results from the alternative specifications are available upon request.

Lee (1992) uses a VAR in a similar context and finds no causal relationship between real stock returns and inflation. However, his analysis is limited to the lead-lag (causal) relationships among variables, and hence does not explain what shocks account for the negative correlation. This point has been taken up in Hess and Lee (1999) who use long-run identification assumptions (Blanchard and Quah (1989)) to decompose stock returns and inflation into transitory demand components and permanent supply components and argue that the money supply shocks lead to positive correlation between real stock returns and inflation. Although Hess and Lee (1999)

interpret the aggregate demand disturbances as money supply shocks, monetary shocks can affect both the aggregate demand and the aggregate supply as argued by Bernanke (1995), and Bernanke and Gertler (1995). Moreover, given the instability of the policy function, it is questionable whether long-run restrictions can really identify monetary policy shocks.

3.2.2 Dynamic Response to Monetary Policy Shocks

In the VAR literature, impulse response functions are typically used to trace out the effects of innovations in one of the variables on the system. Specifically, the impulse response of the j-th variable, $y_{t+h}^{(j)}$ $(h > 0)$ to an innovation in k-th variable, $w_{t+1}^{(k)}$ is defined as

$$\frac{\partial y_{t+h}^{(j)}}{\partial w_{t+1}^{(k)}}. \tag{5}$$

Suppose we cast the VAR in (4) as $y_{t+h} = \sum_{s=0}^{h-1} \Psi_s u_{t+h-s} + y_t(h)$, where $y_t(h) = \sum_{q=1}^{h} \Phi_q y_{t+h-q}$ ($\Phi_q = 0$ for $q > m$) is the h-period VAR forecast of y_t. Starting with $\Psi_0 = I$, Ψ_s may be obtained from the Φ_s using the recursion[24]

$$\Psi_s = \sum_{r=1}^{s} \Psi_{s-r} \Phi_r, \qquad s = 1, 2, \ldots \tag{6}$$

The impulse response matrices to recursively identified one-standard-deviation innovations, Θ_s $(s = 1, 2, ...)$, are obtained by postmultiplying Ψ_s by the lower triangular Choleski factor, P, i.e., $\Theta_s \equiv \Psi_s P$. Then the h-period impulse response of j-th variable to a one-standard deviation shock in k-th innovation is given by the (j, k)-th element of Θ_{h-1} $(h = 1, 2, ...)$.

The impulse responses provide a measure of correlation between economic variables and unsystematic economic shocks. Following a rational expectations argument (e.g., Lucas (1972)) that only unsystematic monetary policy shocks have real effects on the economy[25], the VAR literature has focused on the responses of the system to unsystematic policy shocks. In particular, the impulse response analysis has been used to trace out the net effects of unanticipated policy shocks onto other variables.

3.2.3 Dynamic Response of the Covariance between Inflation and Excess Returns

Finding the portion of the total variance of an observed variable that is due to the various structural shocks is called variance decomposition.

In order to analyze the effects of monetary policy shocks on the covariance between excess stock returns and inflation, we examine the dynamic response of the covariance to orthogonalized economic shocks. We focus on the causal relationship between the return/inflation correlation and economic shocks identified in our system. Using the fact that w_t are serially and contemporaneously uncorrelated with the identity covariance matrix, the h-period ahead forecast error covariance matrix[26] is simply given by $\Sigma_y(h) = \sum_{s=0}^{h-1} \Theta_s \Theta_s'$. Then the h-period forecast error covariance between i-th and j-th variables, $\Sigma_y^{(i,j)}(h)$ is calculated as $\Sigma_y^{(i,j)}(h) = \sum_{k=1}^{K} \sum_{s=0}^{h-1} \theta_s(i,k)\,\theta_s(j,k)$, where K denotes the number of endogenous variables in the system (8 in our model), and $\theta_s(j,k)$ denotes the (j,k)-th element of Θ_s. The covariance decomposition, i.e., the percentage that h-period forecast error covariance between i-th variable and j-th variable accounted for by shocks in k-th variable, is given by

$$\frac{\sum_{s=0}^{h-1} \theta_s(i,k)\,\theta_s(j,k)}{\sqrt{\left(\sum_{k=1}^{K} \sum_{s=0}^{h-1} \theta_s(i,k)^2\right)\left(\sum_{k=1}^{K} \sum_{s=0}^{h-1} \theta_s(j,k)^2\right)}}. \tag{7}$$

The covariance decomposition, which nests the variance decomposition for $i = j$, is a rather obvious extension to the existing VAR toolbox. However, to the best of our knowledge, it has never been used in the literature[27].

We can also trace out the net effect of monetary shocks on the return-inflation covariance by examining the conditional moment profile[28]. The conditional forecast given a unit impulse in the k-th economic shock is given by $y_t^{(k)}(h) \equiv y_t(h) + \Theta_{h-1}\epsilon_{t+1}^{(k)}$ for $h = 1, 2, \ldots$, where $\epsilon_{t+1}^{(k)}$ is an impulse vector with 1 in the k-th position and 0 in the others. The conditional moment profile for the forecast error covariance matrix, due to unit impulse in the k-th innovation, is

$$E\left[\left(y_{t+h} - y_t^{(k)}(h)\right)\left(y_{t+h} - y_t^{(k)}(h)\right)'\right] - \Sigma_y(h) = \theta_{h-1}(\cdot,k)\,\theta_{h-1}(\cdot,k)', \quad h = 1, 2, \ldots,$$

where $\theta_{h-1}(\cdot,k)$ is the k-th column of Θ_{h-1}. Therefore, we obtain the dynamic responses of the h-period forecast error covariance between i-th and j-th variables to the k-th identified innovation by

$$\theta_{h-1}(i,k)\,\theta_{h-1}(j,k), \quad h = 1, 2, ... \tag{8}$$

The covariance decomposition is nothing but a normalized cumulative sum of (8).

3.2.4 Error Bands

The estimators for the impulse responses and the covariance decompositions are computed from the estimated coefficients of the reduced-form VAR model (4) using the recursion (6). However, the mapping from the VAR coefficients (Φ, V) to the impulse response functions (Ψ, Θ) becomes increasingly nonlinear as the forecast horizon increases. It has been shown that large bias and skewness in the small-sample distribution of these estimators render traditional confidence intervals based on asymptotic normal approximation or standard bootstrap methods extremely inaccurate (Kilian (1998), Sims and Zha(1999)).

We adopt a Bayesian approach since it allows for a computationally and conceptually simple way of constructing error bands for impulse responses and covariance decompositions[29]. In order to construct Bayesian error bands, we simulate from the posterior distribution of coefficients of the unrestricted VAR parameters (Φ's) and the covariance matrix of the VAR residuals (V). Assuming Gaussian innovations, the posterior distribution can easily be simulated[30]. Specifically, assuming a diffuse prior on the elements of Φ's and a Jefferys prior on V, we run the Gibbs sampler for $1,200$ iterations and discard the first 200 draws, leaving $1,000$ posterior samples of each coefficient for analysis. We calculate impulse responses and covariance decompositions for each posterior draw, and extract the probability bands (see Sims and Zha (1999) for details[31]). Throughout this paper, we report point estimates and intervals with coverage probability 0.68 (one standard error in the Gaussian case) and 0.90.

3.3 Time-Varying VAR

3.3.1 Estimation

The traditional VAR analysis implicitly assumes that the system is invariant over the entire sample. However, there are many reasons to believe that the Fed's conduct of monetary policy has undergone systematic changes due to changes in chairmen and FOMC members, and in socio-economic conditions, as well as due to the Fed's evolutionary learning toward more optimal monetary policy. To isolate monetary policy shocks, while allowing for a changing monetary policy function, we estimate a time-varying reaction function within the VAR framework. Similar approaches have been pursued by Fuhrer (1996) and Boivin (1999).

In general, we can write our reduced-form VAR model (4) in a structural form as

$$G^{-1}y_t = c + \sum_{l=1}^{m} A_l y_{t-l} + \delta_t, \quad E\left(\delta_t \delta_t'\right) = D \tag{9}$$

where G is the unit diagonal unique lower triangular Choleski factor G such that $V = GDG'$ (or $P = GD^{\frac{1}{2}}$) with a diagonal matrix $D = diag\left(\sigma_1^2, \ldots, \sigma_K^2\right)$, $A_l \equiv G^{-1}\Phi_l$, and $c \equiv G^{-1}\phi$. This structural form (9) can be extended to a case where some or all coefficients are time-varying. Here we consider a simple case

$$G_t^{-1}y_t = c_t + A_{1,t}y_{t-1} + \delta_t, \quad E\left(\delta_t \delta_t'\right) = D \tag{10}$$

where $m = 1$ and $A_{1,t}$, c_t, and G_t are time-varying though the same methodology can be applied to more general setting. Equation (9) can also be rewritten as

$$y_t = c_t + A_{0,t}y_t + A_{1,t}y_{t-1} + \delta_t, \tag{11}$$

with $A_{0,t} \equiv I - G_t^{-1}$. Denoting (i,j)-th element of $A_{l,t}$ by $a_{l,t}(i,j)$ for $l = 0$ and 1, and since δ_t is mutually uncorrelated, we can re-express (11) as the following K scalar equations:

$$y_t^{(i)} = c_t^{(i)} + \sum_{l=0}^{1}\sum_{j=1}^{K} a_{l,t}(i,j)\, y_{t-l}^{(i)} + \delta_t^{(i)}, \quad i = 1, ..., K \tag{12}$$

where $a_{0,t}(i,j) = 0$ for $i \leq j$ (Kitagawa and Gersch (1996)). Due to the triangular nature of the system, we do not have to consider the possibility of the regression being correlated with any elements of δ_t, and thus we can estimate each scalar equation successively without resorting to an IV estimation[32]. Assuming that the time-varying coefficients follow random walk processes[33], we cast each equation into a state-space form for which the Kalman filtering and smoothing algorithm is applied. The hyperparameters – state (time-varying parameters) and measurement error covariances – are chosen to maximize a likelihood function. In order to avoid the numerical problems associated with estimating the MLE's for the hyperparameters[34], we apply the EM algorithm of Shumway and Stoffer (1982) and Watson and Engle (1983) at the cost of somewhat slow convergence.

Once we estimate the time-varying regressions (12) successively, we can construct the reduced-form time-varying VAR(1) system

$$y_t = \phi_t + \Phi_t y_{t-1} + u_t, \quad E\left(u_t u_t'\right) = V_t \tag{13}$$

by the linear algebraic transformations $\phi_t = G_t c_t$, $\Phi_t = G_t A_{1,t}$, and $V_t = G_t D G_t'$. The one-to-one correspondences between (11) and (13) are shown by Kitagawa and Akaike (1978). Although the structural shocks δ_t are homoskedastic, the VAR residuals u_t may exhibit heteroskedasticity when the $a_{0,t}(i,j)$'s are time varying.

In this paper, we focus on the case where only policy reaction functions, i.e., equations for FF and DNBRD in (11) are time-varying. The number of time-varying parameters to be estimated is 12 in the FF equation and 13 for the DNBRD equation. In addition, the number of hyperparameters to be estimated (coefficients in the system and error covariances) is 79 for the FF equation and 92 for the DNBRD equation.

3.3.2 Time-varying Impulse Responses

In the context of time-varying VARs, the impulse response functions should be carefully defined and analyzed. For example, given the reduced-form VAR (13), the *historical* h-period responses, realized at t, given a shock at time $t-h$ can be traced out as $\left[\prod_{l=0}^{h-1} \Phi_{t-l}\right] P_{t-h}$ (Lutkepohl (1993, Ch.12)), where $P_{t-h} = G_{t-h} D^{\frac{1}{2}}$ is the lower triangular Choleski factor of V_{t-h}. However, the realized impulse response obtained in this way does not tell us about the potential net effects that a given shock would have in the presence of parameter uncertainty. A more appropriate measure of the dynamic effect of a shock would be the conditional moment profile (Gallant et al.(1993)). The conditional mean profile is given by $E\left(y_t^{(k)}(h) - y_t(h)\right)$ where $y_t^{(k)}(h)$ is the h-period VAR forecast of y_t given a unit impulse in k-th innovation. However, both $y_t(h)$ and $y_t^{(k)}(h)$ are very difficult to calculate since they are nonlinear functions of unknown parameters Φ_{t+l} and y_{t+l} for $l = 1, 2, ..., h-1$ (Hamilton (1994), Ch.13). The accurate computation of the conditional mean profile would require heavy simulation studies for different time period and for different forecasting horizon.

In this paper, we take a somewhat simpler approach by assuming that the structural relationship, given by (11), is contemporaneously invariant at each period. Fuhrer (1996) uses similar approach by assuming that agents use current period's reaction function parameters to forecast short-term interest rates. Boivin (1999) computes the effect of a structural shock assuming no further change in parameters. Given the random walk specifications of the parameters in the structural equation, using the current parameter estimates to forecast the dynamic effects of policy shocks to other

variables does not introduce large systematic errors in our analysis (Fuhrer (1996)). Thus we compute the impulse response functions $\Theta_{s,t}$, $s = 1, 2, ...$, based on the estimated (smoothed) time-varying coefficients Φ_t and the time-varying lower Choleski factor P_t by

$$\Theta_{s,t} = \Phi_t^{s-1} P_t.$$

We do not report the error bands for the time-varying impulse responses since they are computationally involved. Note, however, that our Bayesian simulation procedure can be extended to the computation of error bands for time-varying impulse responses by combining the simulation smoothing algorithm of de Jong and Shephard (1995) with the method of Sims and Zha (1999).

4 Results

4.1 Monetary Policy Shocks to Return–Inflation Correlation

The main focus of this article is on the dynamic properties of the excess returns/inflation covariance as a result of a Fed policy shock. Figure 2a presents the response of the covariance to all economic shocks, as identified by a Choleski decomposition. In the full sample, a shock to the funds rate has a significantly negative impact during the first 4 to 6 months. Shocks to non-borrowed reserves contribute to the negative covariance mildly, but the effect seems to be more persistent. Interestingly, most other shocks also have a negative impact on the covariance. However, the response to the funds rate is the most significant, economically and statistically. A covariance decomposition, Figure 3a, shows that approximately 20–25% of the negative correlation is explained by the shocks in both federal funds rate and non-borrowed reserves. Those results are surprisingly robust to different orderings in the VAR and also hold in the sub-samples of interest.

Figures 2b,c and 3b,c present the same analysis for the sub-samples 1966:01-1979:06 and 1979:07-1998:12 as do figures 2a and 3a for the entire sample, and the results are very similar. However, it must be noted that the effect of monetary policy shocks is more pronounced in the pre-Volcker period: shocks to the funds rate explain about 25% of the negative covariance at horizons up to 1 year. Monetary policy shocks explain 15-20% of the covariance in the post-Volcker period. During this period, only inflation innovations and shocks in federal funds rate explain the return/inflation covariance.

[Figures 2a,b,c, and 3a,b,c about here]

The presented results are robust to alternative system orderings. As a more formal analysis of robustness, we compare the (unrestricted) covariance matrix V of the VAR residuals with the covariance from the block diagonal restictions, $\tilde{D}$. Table 2 presents the estimate of V for different periods. The elements on the principal diagonal are the standard deviations of the VAR residuals, whereas the off-diagonal elements are the correlations. Interestingly, most correlations are very small, indicating that the unrestricted residuals are close to being uncorrelated. Therefore, imposing any orthogonality restrictions on the residuals will not significantly alter the results. Indeed, the estimate of $\tilde{D}$ (imposing block-triangular restrictions), shown in Table 3, is only slightly different from V. The bold off-diagonal terms in the diagonal blocks of Tables 2 and 3 are very similar and close to zero[35], indicating that the ordering of variables within the macro, monetary policy, and financial groups would not alter the results substantively. The results of such specifications are available upon request.

[Tables 2 and 3 about here]

The role of monetary policy shocks in explaining the negative return/inflation covariance can be attributed to the their opposite effects on excess return and inflation. In the full-sample (Figure 4a), a 25 basis-points shock in federal funds rate leads to a 0.25% decline in the growth of industrial production with 3 months lag. The negative effect lasts for more than a year. A decline in expected output growth should lead to an immediate decline in excess returns. In fact, we find that a 25 basis-points contractionary shock in the federal funds rate reduces excess returns by about 2% in the same month of the shock. Meanwhile, a 25 basis-points shock in the funds rate is followed by a short-run increase in inflation by about the same amount. The commodity prices are more responsive to contractionary policy shocks; after an initial increase, they decline a few months after the shock.

[Figure 4a about here]

Combining these effects, monetary policy shocks have a significant effect on the short-run negative return/inflation relationship. Our empirical evidence is consistent with the arguments laid out in Section 2. We find that a contractionary policy shock brings about a short-run decline in industrial production growth and short-run increase in inflation rate. Furthermore, we notice that a contractionary shock leads to a significant increase in the default premium with a few months' lag. This result suggests that the credit channel is an important mechanism of Fed policy transmission. Our finding contradicts Hess and Lee (1999), who argue that monetary shocks are associated with positive correlation between inflation and real stock returns.

4.2 Dynamic Changes in the VAR System

4.2.1 Dynamic Effects of Monetary Policy Shocks

While we find that monetary policy shocks explain 20-25% of the negative return/inflation relationship in the full sample, we also find that the role of monetary policy has diminished in the latter period. Indeed, there are significant differences in the private sector responses to monetary policy shocks (Figures 4b, 4c). For example, while shocks in federal funds rate explain about 30% of inflation variation and 5% of excess return variation in the pre-Volcker period, they explain less than 10% of inflation variation and do not explain excess return variation in the post-Volcker period (Figures 5a, 5b). Consequently, monetary policy shocks explain a smaller fraction of the negative return/inflation covariance in the latter period than in the former. This result is consistent with Patelis (1997), who finds that monetary policy shocks explain only 3% of excess return variation between 1962 and 1994.

[Figures 4b, 4c, 5a, and 5b about here]

4.2.2 Inflation Variation

The variance decomposition of inflation, displayed in Figures 6a-b, exhibits striking differences between the two sub-periods. In hindsight, this result had to be expected, given the documented changes in the Fed policy function in Table 1. During the pre-Volcker period, significant fraction of inflation variation can be explained by innovations in commodity prices and monetary shocks. Inflation innovations explain only 50% of their own variation in the long run. However, both

commodity price innovations and monetary shocks have lost their explanatory power in the post-Volcker period. 80–90% of inflation variation is explained by its own innovations. Consequently, more of the negative return/inflation relationship is attributable to inflation innovations in the latter period.

[Figures 6a and 6b about here]

The decline in the indicative role of the commodity prices in explaining inflation variation conforms with the finding of Blomberg and Harris (1995), who provide a few explanations for this phenomenon: (i) diminished use of commodities as inflation hedges, (ii) sharp decline in the commodity composition of final output reflecting the shift in of the U.S. economy from commodity-intensive production, and (iii) recent exchange rate fluctuations account for a lot of the recent commodity price fluctuations that may be unrelated to domestic inflation. Moreover, they point out that (iv) the indicative role of the commodity prices may have been offset by more effective countervailing monetary policy movements.

The fourth point in Blomberg and Harris (1995) deserves further attention. Usually commodity prices give early signals of an inflationary surge in aggregate demand. This is because any inflationary impetus is first observed in commodity prices which are continuously updated in thick markets, while consumer price inflation is reported with a lag of a few weeks. If the Fed systematically reacts to the inflationary surge observed in the commodity prices, in order to offset the early signal of inflation in commodity prices, and to the extent that consumer prices do not adjust to the decline in commodity prices quickly, we may not observe the indicative signal of the commodity prices while the Fed in fact does. If this is correct, the Fed employs more effective policy rule in the latter period to offset inflationary movements in the aggregate demand by reacting to commodity prices. This may imply the importance of systematic policy rules in controlling inflation over the unanticipated policy surprises (Cochrane (1998)). As long as the Fed offsets the inflationary impetus systematically and quickly, the monetary policy shocks may not convey inflationary signal as they did in the pre-Volcker period. This hypothesis, discussed below, gives another interpretation for the diminishing role of monetary policy shocks in explaining the variation in inflation. Note that this argument is different from the one made in section 2.3, where we argue that the Fed has private

information about future inflation that is not available even in commodity prices.

4.2.3 Time-Varying Impulse Responses

Figures 7a,7b, and 7c display the effect of a one-standard-deviation shock of the federal funds rate on the inflation/excess return covariance, based on the time-varying VAR model. Note that, in a time-varying VAR, the impulse responses are also time varying, as discussed above. The results are consistent with the findings from the two sub-samples. Monetary policy shocks account for more of the negative return/inflation covariance in the earlier period of our sample, i.e. before the early 1980's. Furthermore, the effect of monetary shocks on the return/inflation covariance exhibits significant time variation, which is a direct result of the significant fluctuations in the Taylor rule parameters (Figure 1).

[Figures 7a-7c about here]

Interestingly, the explanatory power of monetary policy shocks on the negative return/inflation correlation has diminished steadily in recent years. This finding can be traced to: (i) the decreasing price puzzle (top panels of Figures 8a-8c) and (ii) the dwindling negative effect on excess stock returns (bottom panels of Figures 8a-8c). The diminishing price puzzle may be due to several reasons. First, the public has become more capable to forecast inflation, and less prone to surprises (through learning), resulting in the weaker forecasting relationship between identified policy shocks and future inflation. Second, the Fed has learned to control inflation using more effective systematic feedback policy rules. Third, Fed policy has become more transparent, especially in the last few years, thus mitigating some of the information asymmetry effects. Supporting the above arguments is the differing response of inflation and excess returns to monetary shocks during the pre- and post-Volcker periods (Figures 7 and 8). In the former period, a shock to the funds rate produces a large positive increase in prices and a large negative increase in excess returns, whereas in the latter, the effects are not as pronounced.

[Figures 8a-8c about here]

In figures 8a, 8b, and 8c we observe a striking correspondence between the rise in inflation and the fall in excess returns following monetary policy shocks. We argue that this evidence is consistent with the information asymmetry hypothesis of section 2.3. It is reasonable to suspect that the private information of the Fed will be particularly valuable during periods of higher economic instability[36], coinciding with times when monetary policy shocks will have a great effect on future cash flows (especially through the credit channel). Therefore, during periods when the price puzzle is particularly pronounced, we will also see a big impact of monetary shocks on excess returns, resulting in the pattern, observed in Figure 8.

The evidence presented above is not a direct test of information asymmetry; alternative interpretations are also possible. One such interpretation would be that a higher expected inflation, induced by contractionary monetary policy shocks, leads to a decline in excess returns. In this case, the negative relation between excess return and expected inflation, predicted by the money demand theory (Fama(1981), Marshall(1992)), would follow a contractionary shock when the shock in fact conveys information about future inflation. A second interpretation is that an increase in economic volatility would lead to an increase in monetary policy uncertainty. In the presence of information asymmetry between the Fed and the public, the increased policy uncertainty raises expected real interest rates, as in Stulz (1986), which in turn results in lower excess returns. Naturally, those explanations are not mutually exclusive. Finding the most appropriate one would involve writing down and testing a fully specified asset pricing model.

The empirical results, presented above, strongly suggest that monetary policy shocks must play a role in a complete account of the negative returns/inflation covariance. The considerable variation in the response of the covariance might suggest that there are several channels at play. The importance of some of those channels might be difficult to disentangle, but this issue is left for further research. The results clearly contradicts the argument that money supply shocks lead to positive return/inflation correlation (Hess and Lee (1999)).

4.3 Beyond Unanticipated Shocks: Money Demand or Systematic Policy?

In the VAR analysis, we focused on the dynamic effects of unsystematic monetary shocks. Surprisingly, we found that about a quarter of the negative excess return/inflation covariance can be traced to unanticipated Fed policy. But, what about the remaining three quarters? Are they due

to money demand effects, as suggested by Fama (1981) and modelled by Marshall (1992)? Or, is it the case that systematic monetary policy might also have an effect on the covariance, as suggested by Geske and Roll (1983), and Kaul (1987)? Unfortunately, the VAR analysis is uninformative about the role played by systematic or endogenous response of monetary policy to the economy. For example, if the Fed's policy is completely characterized as a feedback rule, then the VAR analysis would conclude that monetary policy has no real effect. However, the response to non-policy shocks may depend importantly on the way monetary policy adjusts endogenously (Walsh (1998)). For example, as the opening quote suggests, higher-than-expected inflation might lead one to anticipate contractionary monetary policy response which in turn would affect real output and real stock returns. In recent articles, Bernanke et al. (1997) and Cochrane (1998) analyze the effects of systematic policy effects as well as unsystematic policy shocks. Both studies conclude that more attention must be devoted to systematic policy effects, though distinguishing systematic changes from unsystematic shocks remain a difficult task.

According to the money demand theory put forth by Fama (1981) and Marshall (1992), positive innovations to inflation (unanticipatedly high inflation) implies lower demand for real balances, which may be associated with lower output growth and lower excess return. In essence, the money demand theory relies on the negative association between inflation and real activity. However, we find less evidence of the negative relationship in post-Volcker sub-samples (Figures 9a, 9b). While we observe that inflation innovations lead to a lower industrial production growth in the pre-Volcker period, we do not find any negative association between inflation innovation and industrial production growth in the post-Volcker period[37]. Nevertheless, negative return/inflation relationship exists in both sub-samples. These findings cast doubt on the validity of the explanation of negative return/inflation relationship solely based on money demand theory.

[Figures 9a-9b about here]

Here we provide an alternative interpretation, based on systematic effects of monetary policy. If innovations in inflation lead us to anticipate endogenous response of the Fed toward contractionary feedback to the economy, and if *anticipated* monetary policy in fact has real effects, we would observe the contemporaneous negative correlation between innovations in excess stock returns and inflation.

In other words, to the extent that the inflation innovations induces anticipated policy actions, and to the extent that the anticipated policy actions are believed to affect the economy, a positive inflation innovation may result in lower excess stock returns. If this interpretation is correct, the negative return/inflation correlation may be proxying for a negative relation between excess returns and anticipated contractionary policy, and may be induced by the positive relationship between unanticipated inflation innovations and anticipated contractionary policy actions. To the extent that the private sector cannot adjust their activity quickly to the anticipated money supply change, the change in the anticipated money supply may have short-run effects on real economic activity, and hence affect excess return. This interpretation emphasizes the role of the systematic monetary policy on real economic activities and financial markets[38].

However, using time-series evidence to uncover the effects of systematic monetary policy rules remain a difficult task, as stressed by Bernanke et al. (1997) and Cochrane (1998). It is not possible to infer the responses to anticipated policy actions from a VAR, since dynamic effects can be cleanly traced out only in the case of unanticipated random shocks to the VAR system. Nevertheless, our empirical findings are consistent with the hypothesis that the negative return/inflation correlation proxies for the negative relation between systematic monetary policy movements and excess stock returns. First, in the pre-Volcker period, a positive inflation innovation leads to an increase in federal funds rate with a few months' lag. In the post-Volcker period, a positive inflation innovation leads to a concurrent rise in federal funds rate. Thus an unexpected increase in inflation in fact leads to contractionary movements of the Fed. Second, the diminished relationship between inflation and industrial production growth can also be explained by the systematic action of the Fed to offset the effect of inflation innovations on industrial production.

5 Conclusion

The correlation between excess returns and inflation must be zero, if monetary policy is neutral, or positive, if monetary policy has real effects. However, we find that about a quarter of the negative correlation between excess returns and inflation is explained by shocks to the monetary policy function. The results are robust to alternative VAR ordering schemes and also hold for any time period in the 1966-1998 sample. We show that our finding, which contradicts simple economic intuition, is linked to the price puzzle. In the short run, a contractionary Fed policy

shock, implemented by an increase in the federal funds rate, induces lower excess rates of return through its effect on real variables. Such a policy has also been followed by a seemingly anomalous increase in consumer prices, thus producing the observed negative correlation between excess returns and inflation.

A cynical view is that we have replaced one puzzle with another one. However, we present a simple story, based on the Fed's superior information, that might account for the "price puzzle," and which is consistent with the previous findings in Romer and Romer (2000). As further support of our information asymmetry hypothesis, we find that in periods of high economic instability, when Fed shocks have a particularly large effect on excess returns, we also observe a more pronounced "price puzzle." During those periods, the private information of the central bank has a big impact on future consumer prices. Despite the instability of the policy function and the variability of the Fed impact on excess returns and inflation, monetary policy shocks have always been followed by an increasingly negative correlation between those two variables.

Our paper leaves some unanswered questions: First, how can we account for the remaining 75% of the negative covariance? Some of it is surely due to money demand shocks, some of it might be caused by cross-sectional fluctuations in industry output (Boudoukh et al. (1994)), and some of it might even be traceable to effects of fiscal policy. Second, is the VAR framework able to test the propositions of Geske and Roll (1983) and Kaul (1986) that some of this negative covariance might also be caused by *systematic* and endogenous monetary policy actions? We are sympathetic to the view, expressed by Bernanke et al. (1997) and Cochrane (1998), that the VAR literature needs to find ways to incorporate and measure the systematic effects of monetary policy onto financial and real variables. However, no matter what the answers to these questions are, the fact remains that a significant fraction of the negative excess returns/inflation covariance is explained by Fed policy shocks. Therefore, if an asset pricing model is to capture the entire negative correlation, it must find ways to account for the policy of the Fed.

Appendix A

The eight variables are obtained in monthly frequency from 1966 to 1998. The variables are listed in the order of the Wold causal ordering employed in our VAR analysis.

Name	Description	Source	Mnemonic
IPG	The log difference of the Industrial production, seasonally adjusted, 1987=100.	CITIBASE	IP
INF	The inflation rate, defined by the log difference in the Consumer Price Index, all items, seasonally adjusted.	CITIBASE	PUNEW
DPCOM	The log difference of spot market index for all commodities.	CITIBASE	PSCCOM
FF	The federal funds rate, average of business day figures.	CITIBASE	FYFF
DNBRD	Minus the log difference in non-borrowed reserves. "Minus" is taken to facilitate comparisons with FF.	CITIBASE	FMRNBC
DEFP	The spread of Baa-rated over Aaa-rated corporate bond yields.	CITIBASE	FYAAAC, FYBAAC
TERM	The spread between 1-year and 3-month Treasury bill rates, converted to continuously compounding basis.	CITIBASE	FYGM3, FYGMYR
EP	The excess return (the equity premium), obtained by subtracting the CRSP one-month T-bill rate from the NYSE value-weighted stock return including dividends..	CRSP	VWRETD

Tables and Figures

Table 1: Taylor Rule

	Period				
	1966:01-1998:12	1966:01-1979:06	1979:07-1998:12	1979:07-1982:12	1983:1-1998:12
γ	0.7657	0.1896	1.2255	1.0517	0.7438
	0.0456	0.0516	0.0719	0.2793	0.0456
η	0.1557	0.0488	0.2635	1.4061	0.0612
	0.0186	0.0183	0.0314	0.1841	0.0121
T	395	161	233	41	191

Notes: The estimated regression is: $r_t^{ff} = \alpha + \phi r_{t-1}^{ff} + (1-\phi)(\gamma\pi_t + \eta x_t) + \varepsilon_t$, where r_t^{ff} is the federal funds rate, π_t is the inflation rate, and x_t is the growth rate of industrial production. This is the Fed's policy function, as suggested by Taylor (1993) and modified for interest smoothing, as in Clarida et al.(1999). Some studies use the GDP gap for x_t. Unfortunately, GDP is only available at a quarterly frequency, whereas in our case, high frequency (monthly) data is important. However, the obtained results are similar to those obtained from using the GDP gap and quarterly data. The coefficient ϕ is set equal to 0.96, the estimate obtained from the entire sample. The standard errors are given below the estimates. Using π_{t-1} or π_{t+1} in the regression does not change the results qualitatively, because inflation is a fairly persistent variable.

Table 2: Variance/Correlation Matrices of VAR Residuals

(a) VAR(3) from 1966.01 to 1998.12 (Full sample period)

	IPG	INF	DPCOM	FF	DNBRD	DEFP	TERM	EP
IPG	*8.4567*							
INF	**0.0959**	*2.4380*						
DPCOM	**-0.0473**	**0.1290**	*23.6925*					
FF	0.0187	0.0455	0.0796	*0.5538*				
DNBRD	0.0370	0.1000	0.1454	**0.3295**	*17.4317*			
DEFP	-0.0882	-0.0450	-0.2011	-0.0122	-0.0412	*0.1015*		
TERM	-0.0219	-0.0166	0.1295	-0.2553	-0.0176	**-0.0741**	*0.2059*	
EP	-0.0511	-0.1252	-0.0483	-0.1313	-0.1561	**0.1205**	**-0.1437**	*52.9881*

(b) VAR(2) from 1966.01 to 1979.06 (Pre-Volcker period).

	IPG	INF	DPCOM	FF	DNBRD	DEFP	TERM	EP
IPG	*9.6075*							
INF	**0.0800**	*2.4475*						
DPCOM	**-0.0664**	**0.1289**	*27.3741*					
FF	0.0414	-0.0898	0.0695	*0.3563*				
DNBRD	0.0705	0.1527	0.1695	**0.2043**	*17.8208*			
DEFP	-0.1587	-0.0136	-0.1302	-0.1341	0.1171	*0.0834*		
TERM	-0.0324	0.0197	-0.0273	0.2892	-0.0047	**-0.0616**	*0.1801*	
EP	-0.0680	-0.1734	0.0178	-0.1677	-0.1627	**0.2752**	**-0.1373**	*54.3924*

(c) VAR(2) from 1983.01 to 1998.12 (Post-Volcker period).

	IPG	INF	DPCOM	FF	DNBRD	DEFP	TERM	EP
IPG	*7.5484*							
INF	**0.0921**	*1.8560*						
DPCOM	**-0.0844**	**0.1313**	*19.4920*					
FF	0.0762	0.1415	0.2388	*0.2362*				
DNBRD	-0.0563	-0.0830	0.1537	**-0.0411**	*13.4409*			
DEFP	-0.0245	-0.0145	-0.1956	0.0330	-0.2522	*0.0726*		
TERM	-0.0031	-0.0049	0.1759	0.1302	0.1904	**-0.1511**	*0.1352*	
EP	-0.0915	-0.1384	-0.0464	-0.0687	-0.0078	**0.0062**	**-0.2998**	*50.6572*

Note: In each matrix, diagonal elements (italicized) are standard deviations of the unrestricted VAR residuals, in annual percentage points. Off-diagonal elements are their correlations.

Table 3: Variance/Correlation Matrices of Block-Orthogonalized VAR Residuals

(a) VAR(3) from 1966.01 to 1998.12 (Full sample period)

	IPG	INF	DPCOM	FF	DNBRD	DEFP	TERM	EP
IPG	*8.4567*							
INF	**0.0959**	*2.4380*						
DPCOM	**-0.0473**	**0.1290**	*23.6925*					
FF	0.0000	0.0000	0.0000	*0.5516*				
DNBRD	0.0000	0.0000	0.0000	**0.3203**	*17.1758*			
DEFP	0.0000	0.0000	0.0000	0.0000	0.0000	*0.0989*		
TERM	0.0000	0.0000	0.0000	0.0000	0.0000	**-0.0514**	*0.1963*	
EP	0.0000	0.0000	0.0000	0.0000	0.0000	**0.1094**	**-0.1812**	*51.7876*

(b) VAR(2) from 1966.01 to 1979.06 (Pre-Volcker period).

	IPG	INF	DPCOM	FF	DNBRD	DEFP	TERM	EP
IPG	*9.6075*							
INF	**0.0800**	*2.4475*						
DPCOM	**-0.0664**	**0.1289**	*27.3741*					
FF	0.0000	0.0000	0.0000	*0.3531*				
DNBRD	0.0000	0.0000	0.0000	**0.2090**	*17.3596*			
DEFP	0.0000	0.0000	0.0000	0.0000	0.0000	*0.0794*		
TERM	0.0000	0.0000	0.0000	0.0000	0.0000	**-0.0275**	*0.1712*	
EP	0.0000	0.0000	0.0000	0.0000	0.0000	**0.2924**	**-0.0944**	*52.1736*

(c) VAR(2) from 1983.01 to 1998.12 (Post-Volcker period).

	IPG	INF	DPCOM	FF	DNBRD	DEFP	TERM	EP
IPG	*7.5484*							
INF	**0.0921**	*1.8560*						
DPCOM	**-0.0844**	**0.1313**	*19.4920*					
FF	0.0000	0.0000	0.0000	*0.2270*				
DNBRD	0.0000	0.0000	0.0000	**-0.0672**	*13.2000*			
DEFP	0.0000	0.0000	0.0000	0.0000	0.0000	*0.0690*		
TERM	0.0000	0.0000	0.0000	0.0000	0.0000	**-0.0939**	*0.1304*	
EP	0.0000	0.0000	0.0000	0.0000	0.0000	**-0.0065**	**-0.3022**	*49.9297*

Note: In each matrix, diagonal elements (italicized) are standard deviations of block-orthogonal residuals, in annual percentage points. The blocks are: (IPG, INF, DPCOM), (FF, DNBRD), and (DEFP, TERM, EP). Off-diagonal elements are their correlations.

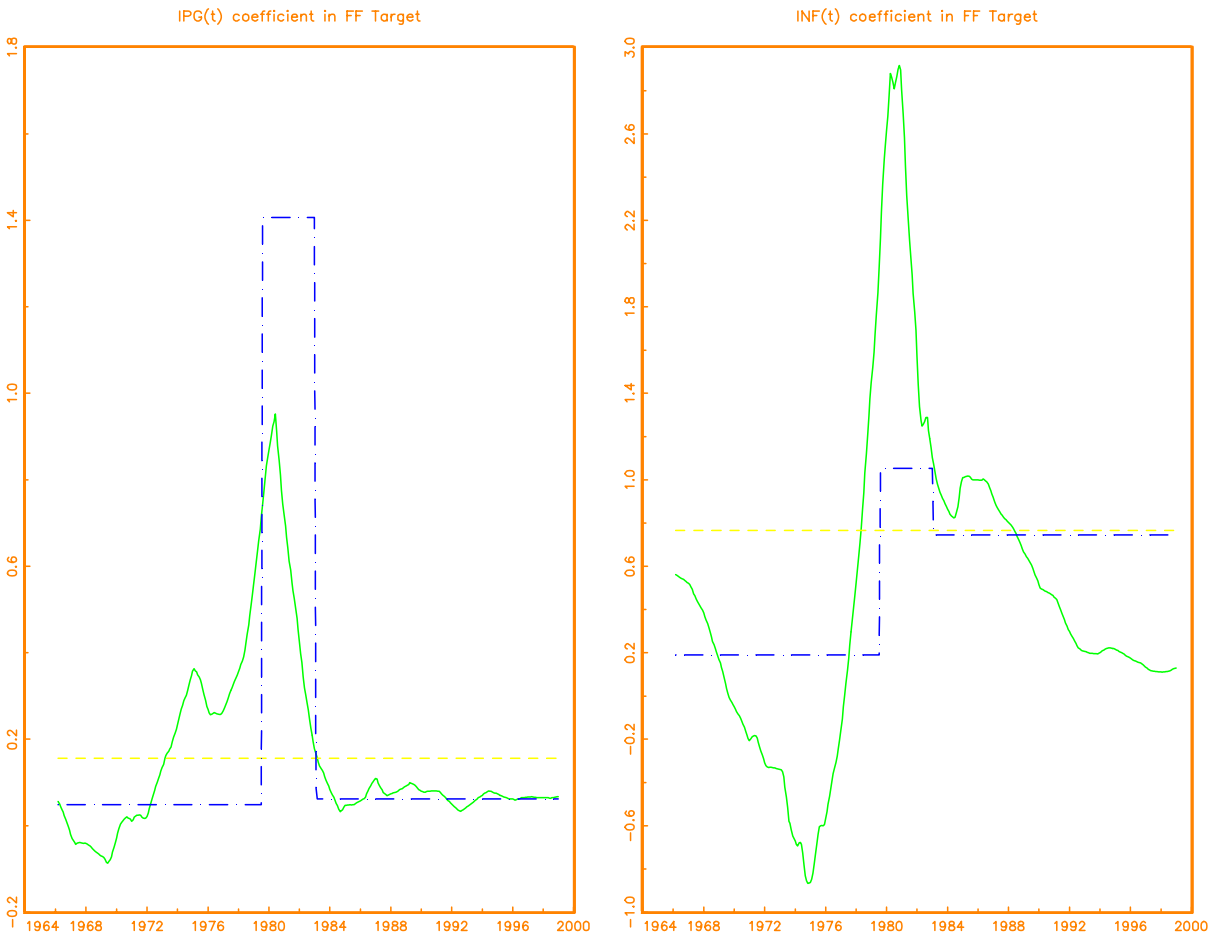

Figure 1

Each graph plots an estimated coefficient of a time-varying-parameter regression: $r_t^{FF} = \alpha_t + \phi_t r_{t-1}^{FF} + (1-\phi_t)(\eta_t x_t + \gamma_t \pi_t) + \varepsilon_t$ where r_{t-1}^{FF} is the federal funds rate, x_t is the growth rate of industrial production, and π_t is the inflation rate. In this regression, coefficients $\Xi_t = [\alpha_t, \phi_t, \eta_t, \gamma_t]'$ are assumed to follow a random walk process $\Xi_t = I_4 \times \Xi_{t-1} + \varsigma_t$ where $\varepsilon_t, \varsigma_t$ are white noise sequences with $E(\varepsilon_t \varepsilon_t') = R$, $E(\varsigma_t \varsigma_t') = Q$, and $E(\varepsilon_t \varsigma_t') = 0$. The ML estimators of R, Q are used to obtain smoothed estimates of Ξ_t. Each graph plots smoothed estimates of η_t and γ_t, respectively. Both estimates are divided by $(1-0.96)$ to facilitate comparison with the constant-coefficient counterpart, given in Table 1. The dotted and dashed lines correspond to the full-sample and subsample estimates from Table 1.

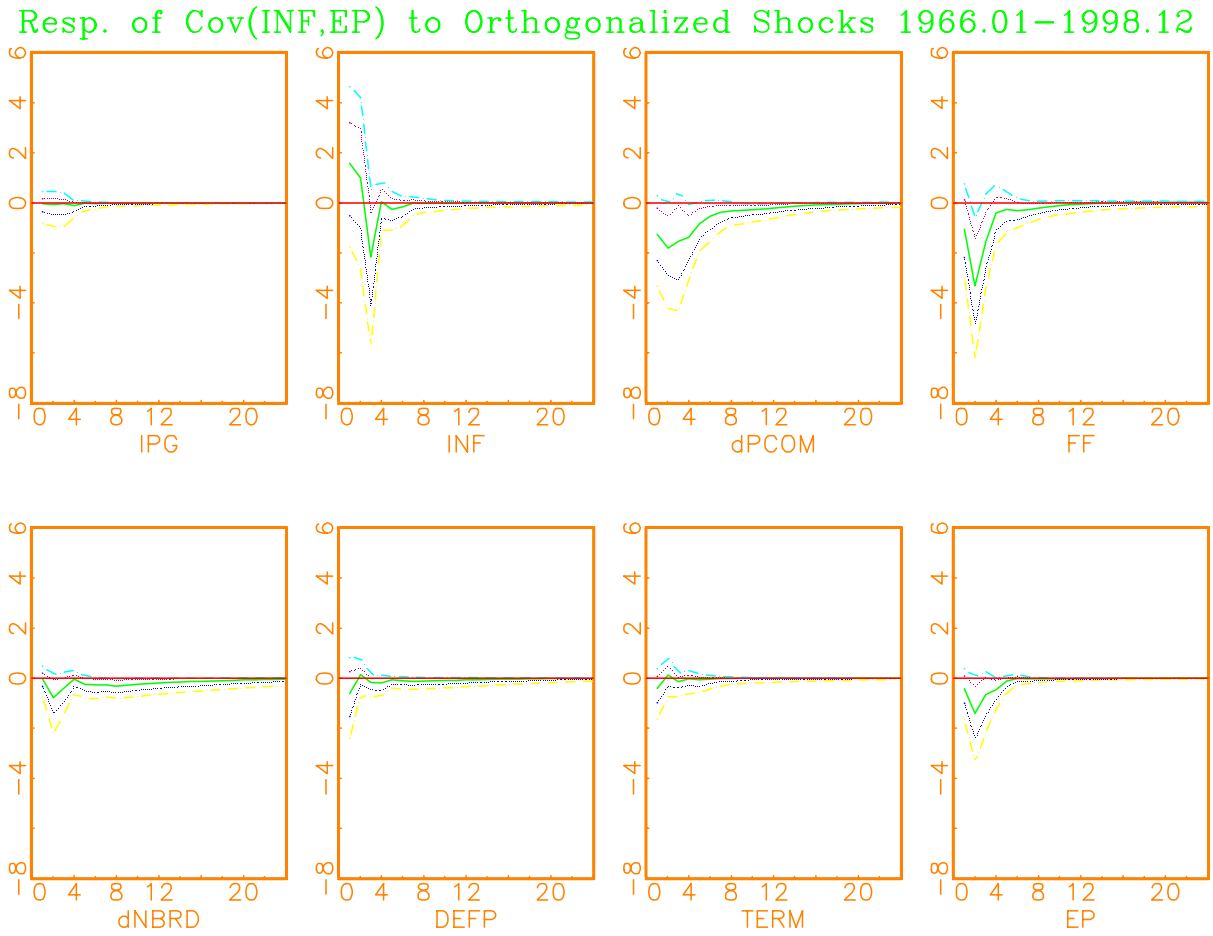

Figure 2a

Response of the covariance between inflation and excess return for full sample period (1966:01–1998:12). Each graph is the response of covariance between INF and EP to a one-standard-deviation impulse in triangularly orthogonalized innovations in IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP. All variables are in annual percentage points. Each response is plotted over a 24-month horizon following the month of the shock. The vertical axis scales represent deviation in covariance ($\times 10^4$) due to shocks in each variable. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point. Responses of the h-period forecast error covariance between i-th and j-th variables to the k-th identified innovation by $\theta_h(i,k)\,\theta_h(j,k)$, $h = 1, 2, ...$, where $\theta_h(i,k)$ is the (i,k)-th element of the impulse response matrix Θ_h after the recursive (triangular) identification assumptions have been imposed.

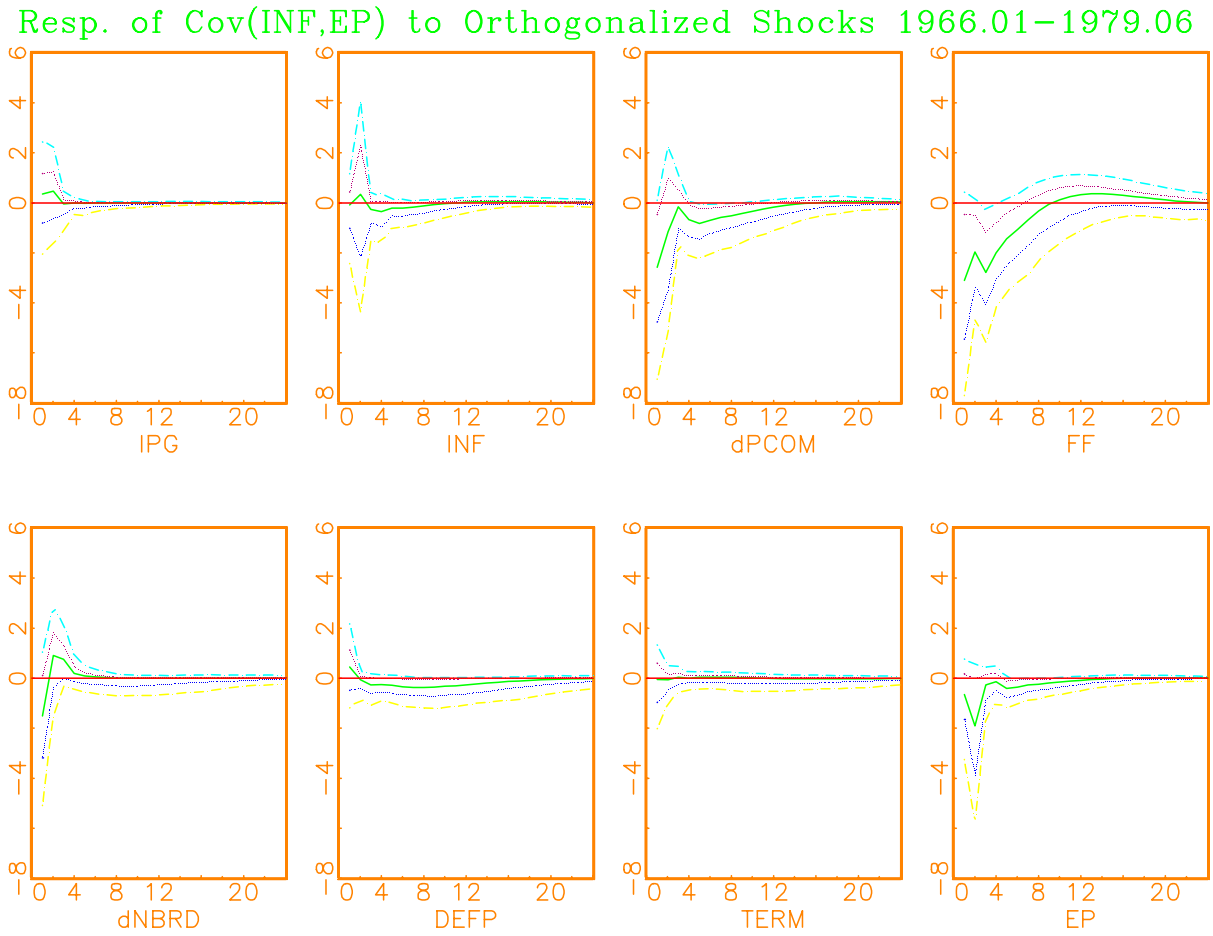

Figure 2b

Response of the covariance between inflation and excess return for pre-Volcker period (1966:01–1979:06). Each graph is the response of covariance between INF and EP to a one-standard-deviation impulse in triangularly orthogonalized innovations in IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP. All variables are in annual percentage points. Each response is plotted over a 24-month horizon following the month of the shock. The vertical axis scales represent deviation in covariance ($\times 10^4$) due to shocks in each variable. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point. Responses of the h-period forecast error covariance between i-th and j-th variables to the k-th identified innovation by $\theta_h(i,k)\,\theta_h(j,k)$, $h = 1, 2, ...$, where $\theta_h(i,k)$ is the (i,k)-th element of the impulse response matrix Θ_h after the recursive (triangular) identification assumptions have been imposed.

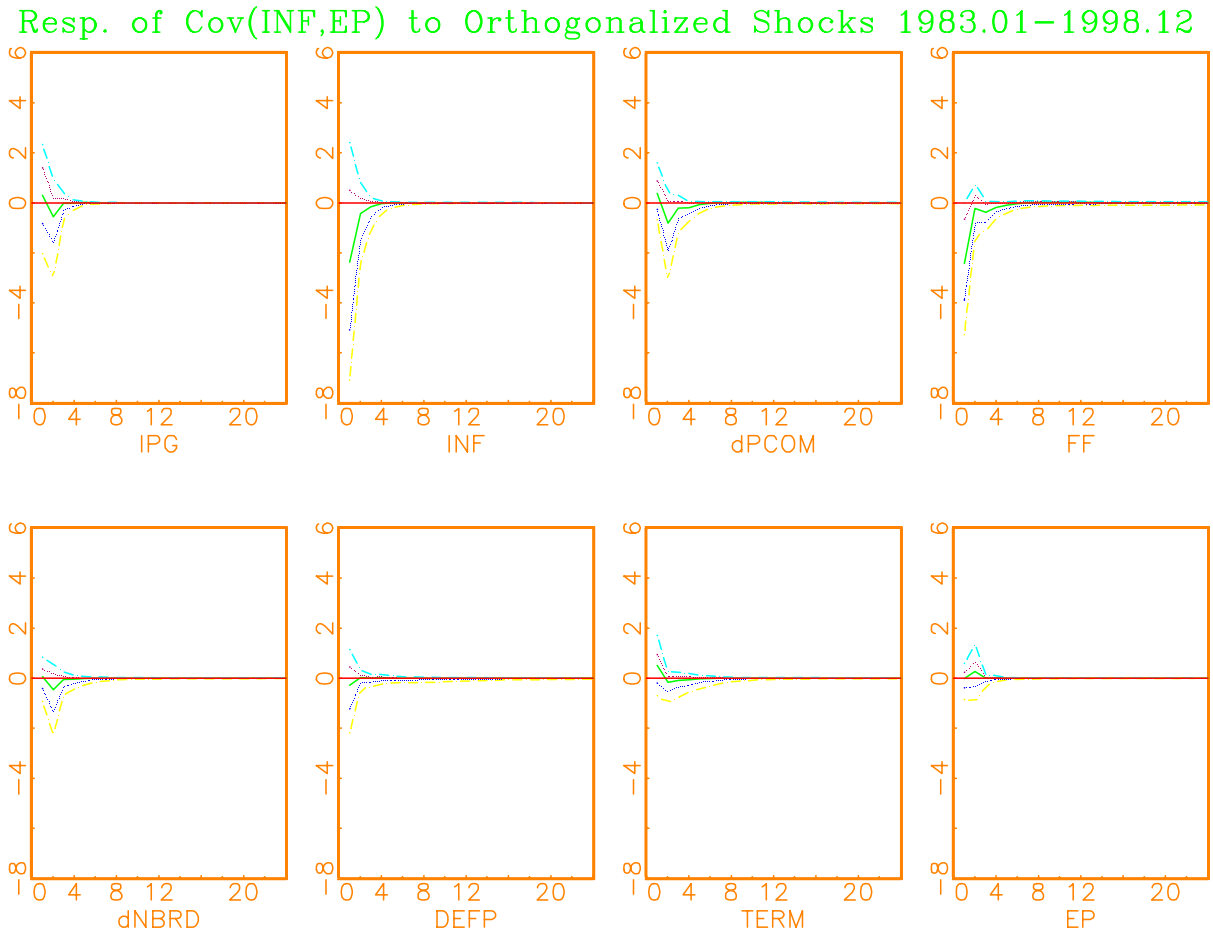

Figure 2c

Response of the covariance between inflation and excess return for post-Volcker period (1983:01–1998:12). Each graph is the response of covariance between INF and EP to one-standard-deviation impulse in triangularly orthogonalized innovations in IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP. All variables are in annual percentage points. Each response is plotted over a 24-month horizon following the month of the shock. The vertical axis scales represent deviation in covariance ($\times 10^4$) due to shocks in each variable. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point. Responses of the h-period forecast error covariance between i-th and j-th variables to the k-th identified innovation by $\theta_h(i,k)\,\theta_h(j,k)$, $h = 1, 2, ...$, where $\theta_h(i,k)$ is the (i,k)-th element of the impulse response matrix Θ_h after the recursive (triangular) identification assumptions have been imposed.

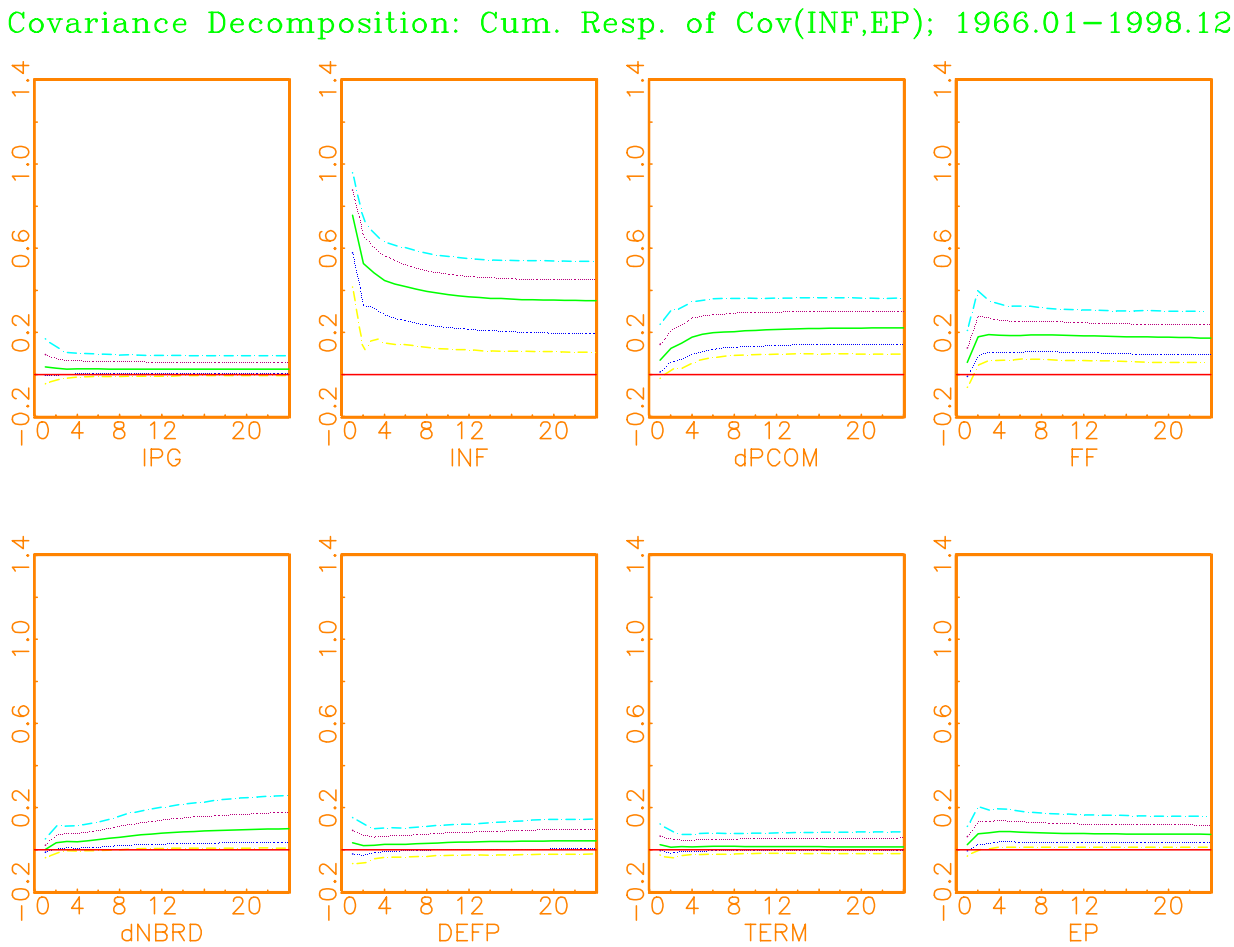

Figure 3a

Decomposition of covariance between inflation and excess return for full sample period (1966:01–1998:12). Each graph is the cumulative response of covariance between INF and EP to a one-standard-deviation impulse in triangularly orthogonalized innovations in IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP. All variables are in annual percentage points. Each response is plotted over a 24-month horizon following the month of shocks. The vertical axis scales represent deviation in covariance ($\times 10^4$) due to shocks in each variable. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point. The covariance decomposition, i.e., the percentage that h-period forecast error covariance between i-th variable and j-th variable accounted for by shocks in k-th variable, is given by $\frac{\sum_{s=1}^{h}\theta_s(i,k)\theta_s(j,k)}{\sqrt{\left(\sum_{k=1}^{K}\sum_{s=1}^{h}\theta_s(i,k)^2\right)\left(\sum_{k=1}^{K}\sum_{s=1}^{h}\theta_s(j,k)^2\right)}}$ for $h = 1, 2, ..., 24$, .where $\theta_h(i,k)$ is the (i,k)-th element of the impulse response matrix Θ_h after the recursive (triangular) identification assumptions have been imposed.

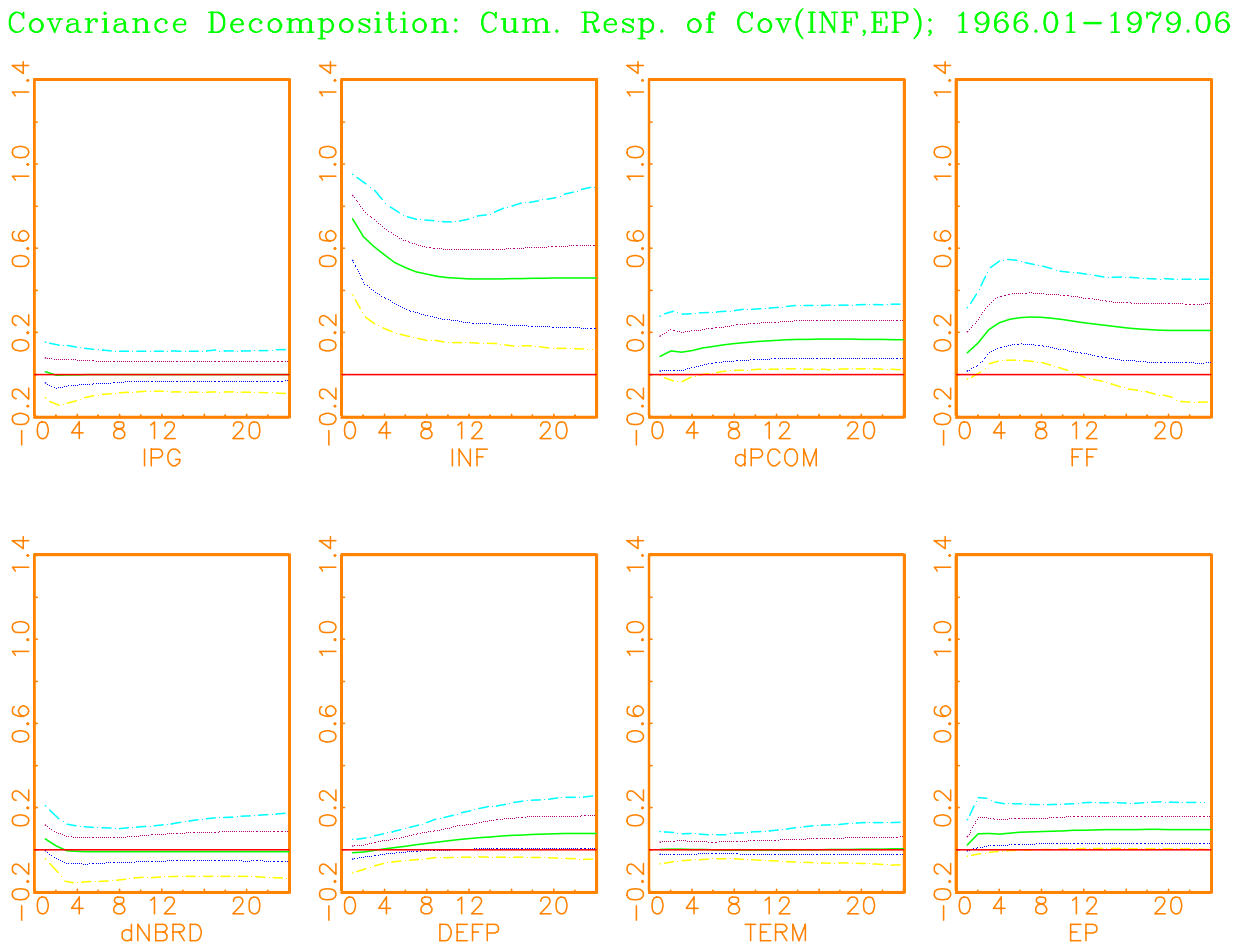

Figure 3b

Decomposition of covariance between inflation and excess return for pre-Volcker period (1966:01–1979:06). Each graph is the cumulative response of covariance between INF and EP to a one-standard-deviation impulse in triangularly orthogonalized innovations in IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP. All variables are in annual percentage points. Each response is plotted over a 24-month horizon following the month of shocks. The vertical axis scales represent deviation in covariance ($\times 10^4$) due to shocks in each variable. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point. The covariance decomposition, i.e., the percentage that h-period forecast error covariance between i-th variable and j-th variable accounted for by shocks in k-th variable, is given by $\frac{\sum_{s=1}^{h}\theta_s(i,k)\theta_s(j,k)}{\sqrt{\left(\sum_{k=1}^{K}\sum_{s=1}^{h}\theta_s(i,k)^2\right)\left(\sum_{k=1}^{K}\sum_{s=1}^{h}\theta_s(j,k)^2\right)}}$ for $h = 1, 2, ..., 24$, .where $\theta_h(i,k)$ is the (i,k)-th element of the impulse response matrix Θ_h after the recursive (triangular) identification assumptions have been imposed.

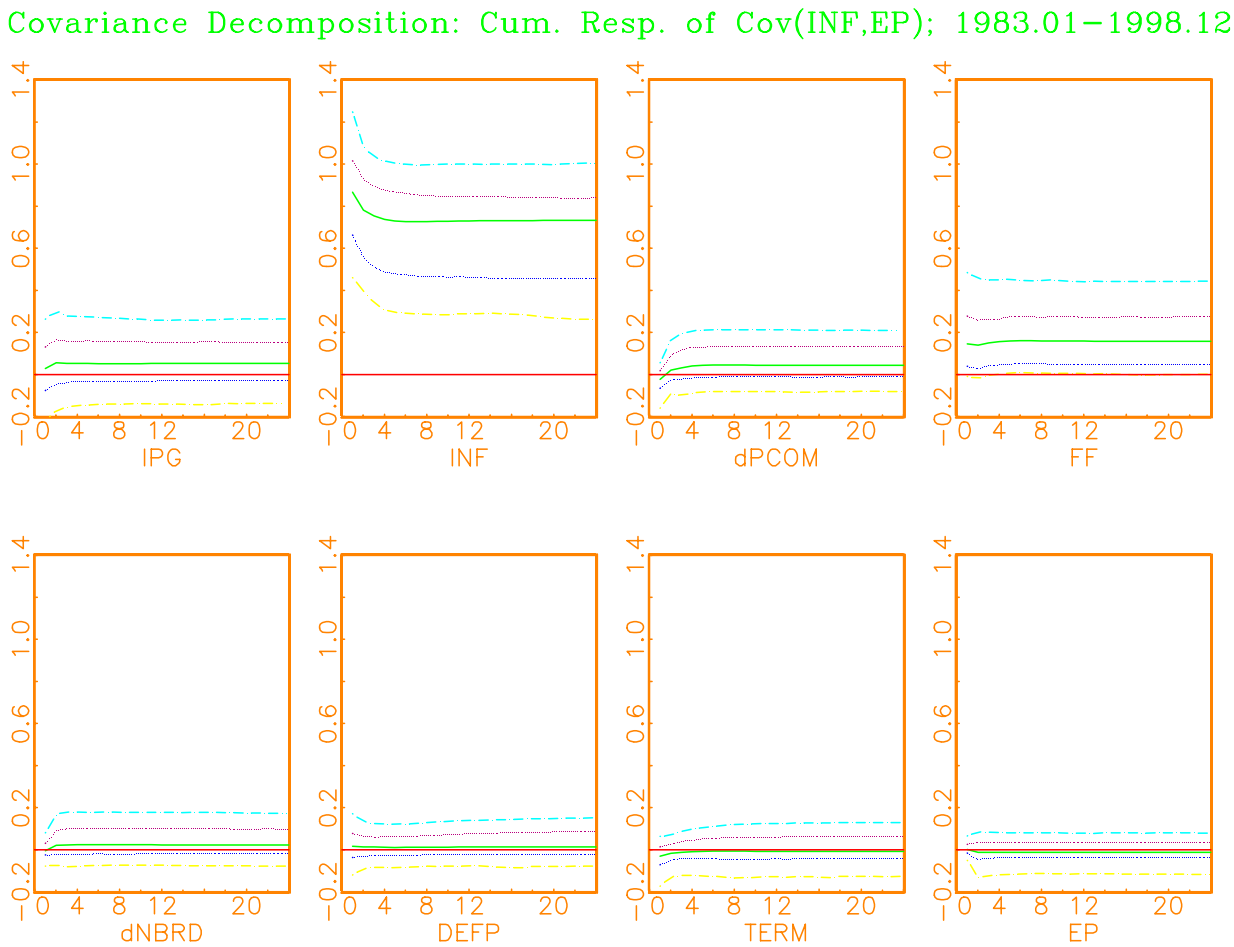

Figure 3c

Decomposition of covariance between inflation and excess return for post-Volcker period (1983:01–1998:12). Each graph is the cumulative response of covariance between INF and EP to a one-standard-deviation impulse in triangularly orthogonalized innovations in IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP. All variables are in annual percentage points. Each response is plotted over a 24-month horizon following the month of shocks. The vertical axis scales represent deviation in covariance ($\times 10^4$) due to shocks in each variable. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point. The covariance decomposition, i.e., the percentage that h-period forecast error covariance between i-th variable and j-th variable accounted for by shocks in k-th variable, is given by $\frac{\sum_{s=1}^{h}\theta_s(i,k)\theta_s(j,k)}{\sqrt{\left(\sum_{k=1}^{K}\sum_{s=1}^{h}\theta_s(i,k)^2\right)\left(\sum_{k=1}^{K}\sum_{s=1}^{h}\theta_s(j,k)^2\right)}}$ for $h = 1, 2, ..., 24$, .where $\theta_h(i,k)$ is the (i,k)-th element of the impulse response matrix Θ_h after the recursive (triangular) identification assumptions have been imposed.

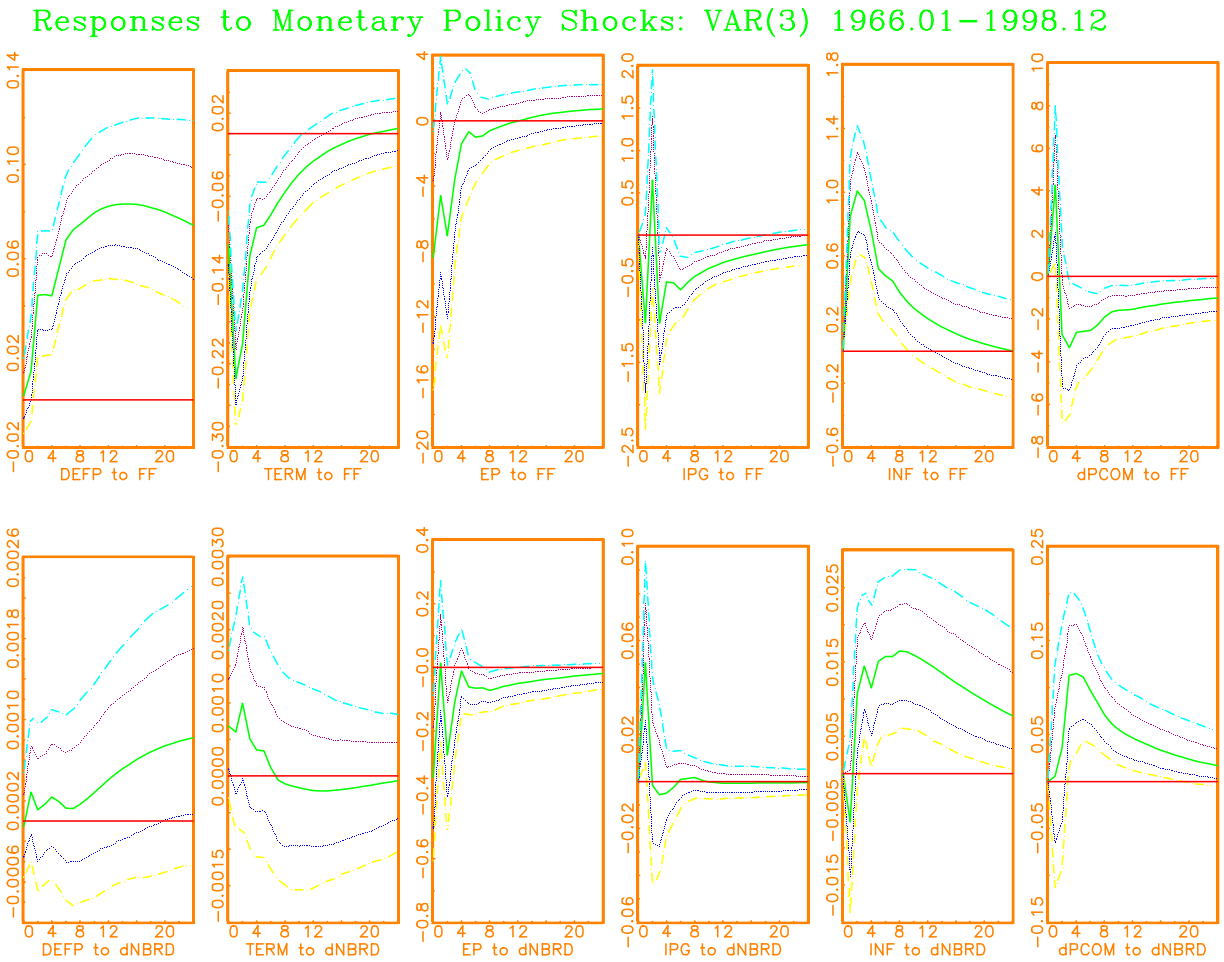

Figure 4a

Effects of monetary policy shocks on non-policy variables for full sample period (1966:01–1998:12). Each graph shows 24-month response of variables to a one percent shock in monetary policy variables. Top panel: effects of shocks in federal funds rate. Bottom panel: effects of shocks in minus log difference nonborrowed reserves. Vertical axis scales represent percent deviation of variables. Responses of DEFP, TERM, and EP are depicted first since these variables are assumed to respond to policy shocks within the same month. IPG, INF, and DPCOM are assumed to respond to policy shocks with one-month lag. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point.

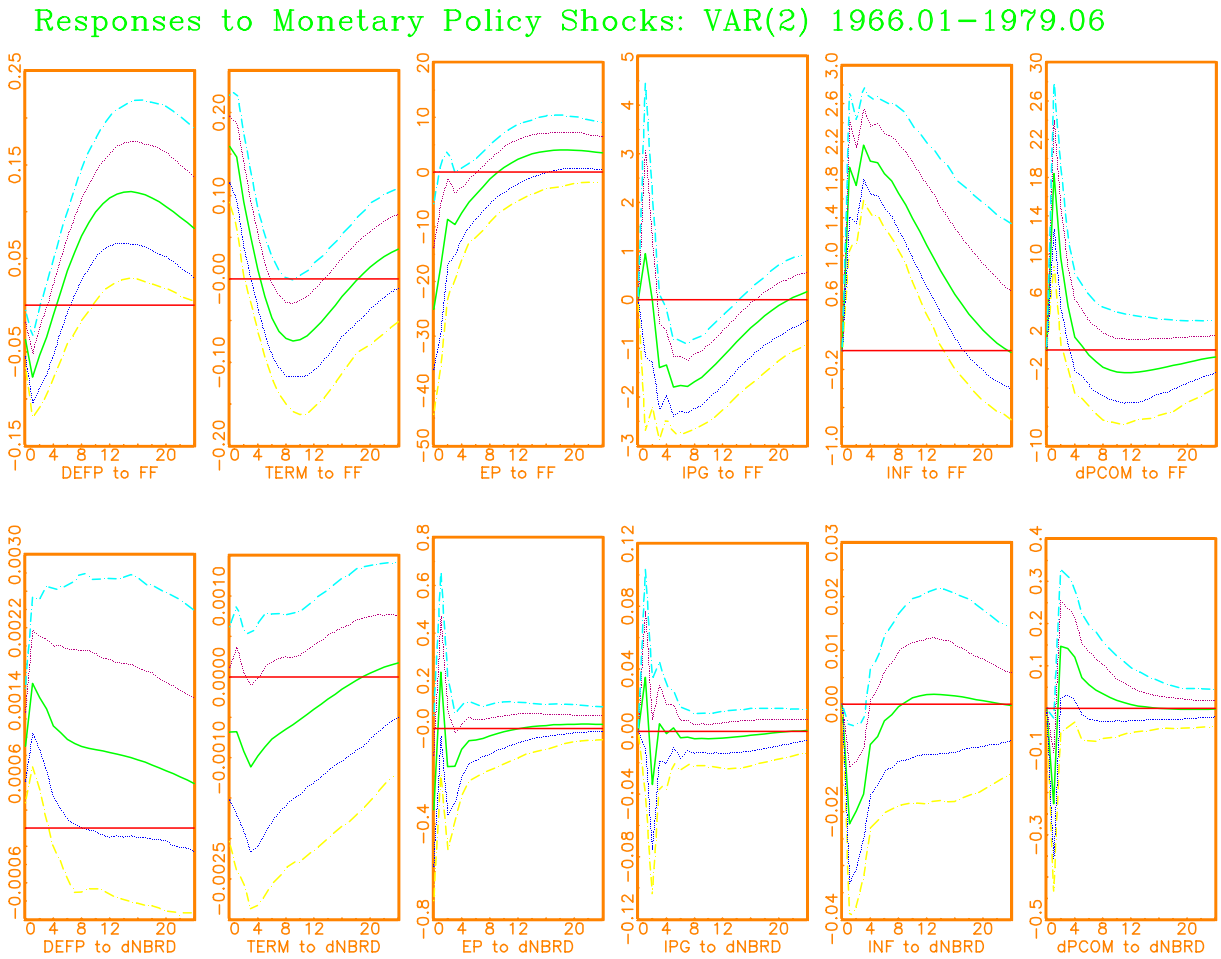

Figure 4b

Effects of monetary policy shocks on non-policy variables for pre-Volcker period (1966:01–1979:06). Each graph shows 24-month response of variables to a 1 percent shock in monetary policy variables. Top panel: effects of shocks in federal funds rate. Bottom panel: effects of shocks in minus log difference nonborrowed reserves. Vertical axis scales represent percent deviation of variables. Responses of DEFP, TERM, and EP are depicted first since these variables are assumed to respond to policy shocks within the same month. IPG, INF, and DPCOM are assumed to respond to policy shocks with one-month lag. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point.

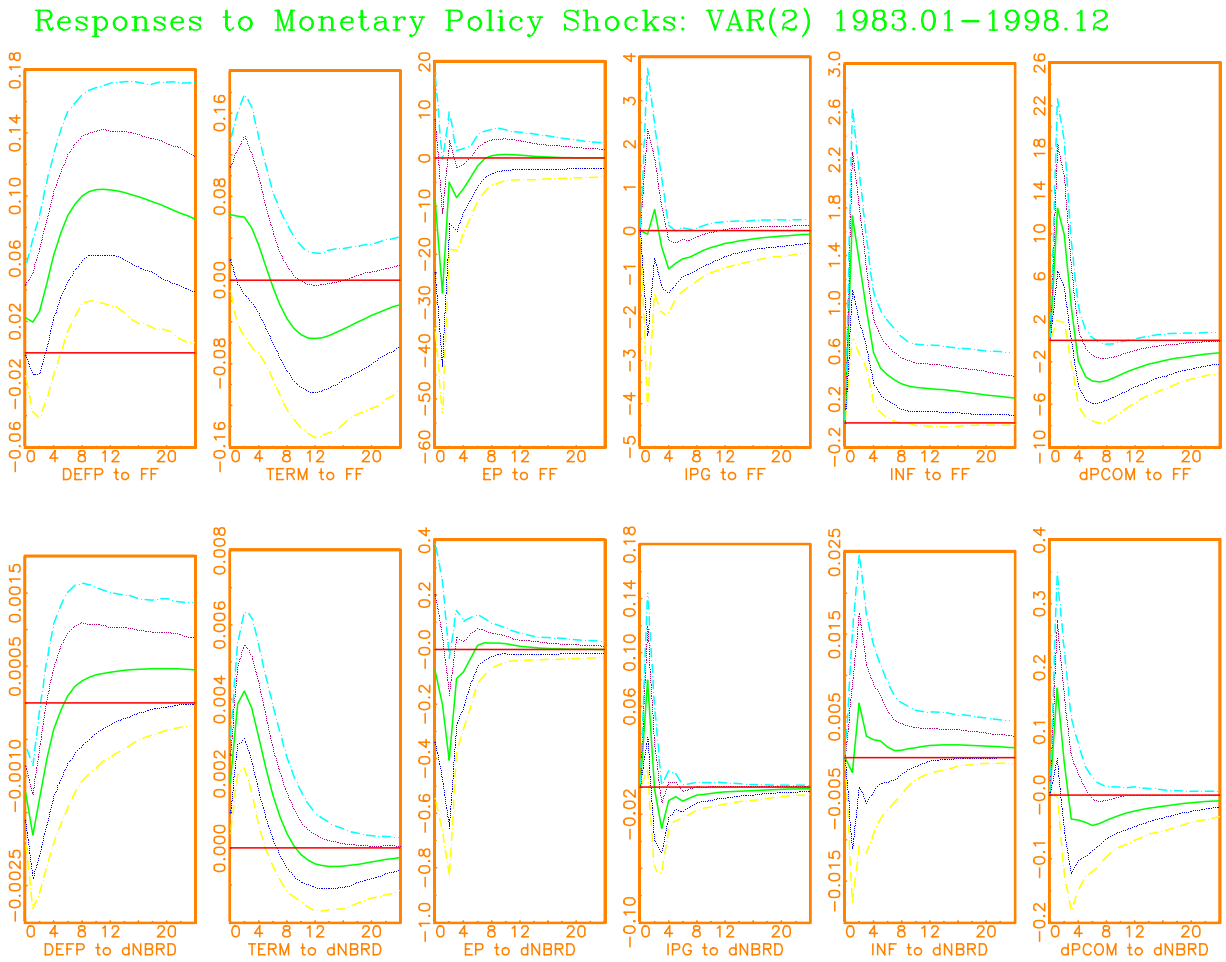

Figure 4c

Effects of monetary policy shocks on non-policy variables for post-Volcker period (1983:01–1998:12). Each graph shows 24-month response of variables to a 1 percent shock in monetary policy variables. Top panel: effects of shocks in federal funds rate. Bottom panel: effects of shocks in minus log difference nonborrowed reserves. Vertical axis scales represent percent deviation of variables. Responses of DEFP, TERM, and EP are depicted first since these variables are assumed to respond to policy shocks within the same month. IPG, INF, and DPCOM are assumed to respond to policy shocks with one-month lag. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point.

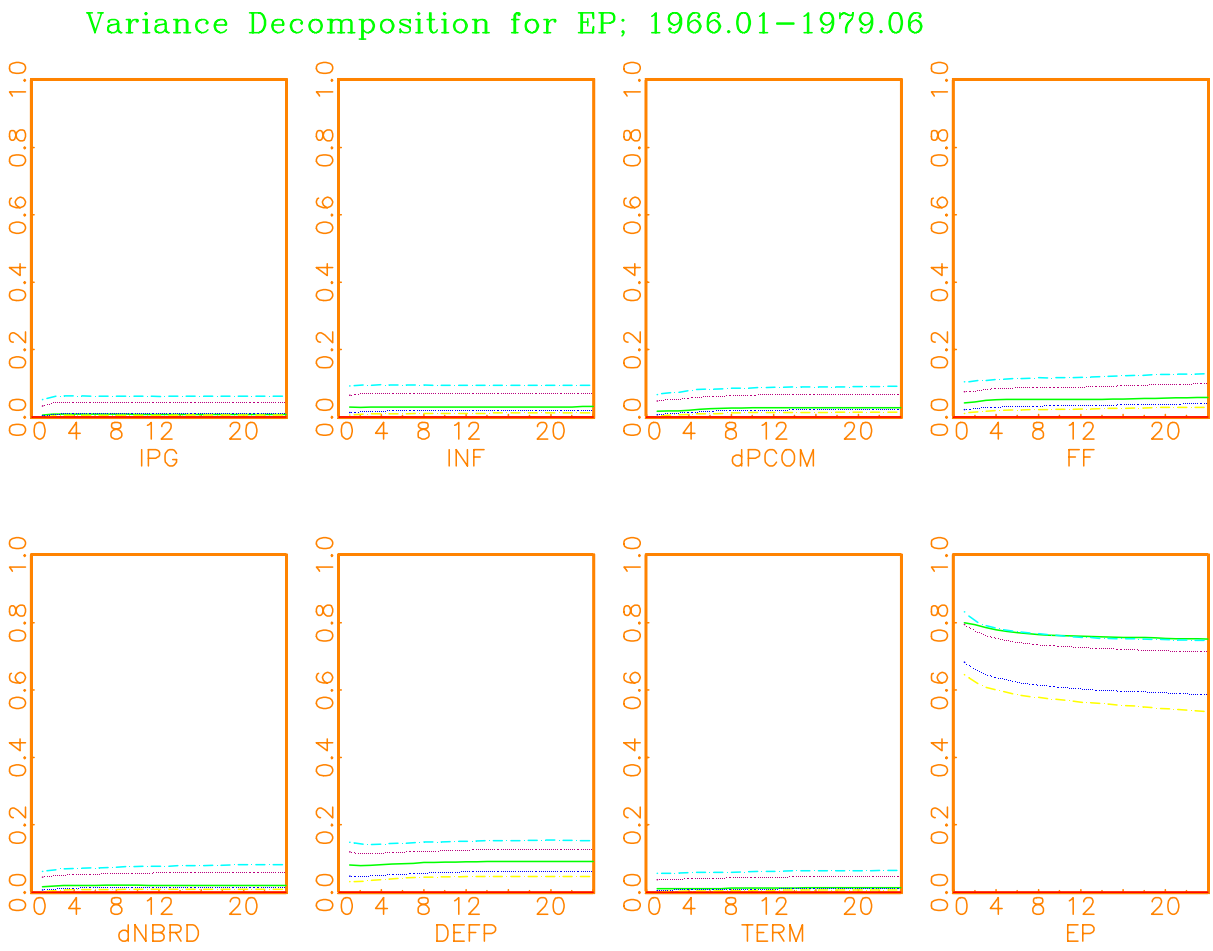

Figure 5a

Variance decomposition of excess return for pre-Volcker period (1966:01–1979:06). Each graph is the cumulative response of variance of EP to a one-standard-deviation impulse in triangularly orthogonalized innovations in IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP. All variables are in annual percentage points. Each response is plotted over a 24-month horizon following the month of shocks. The vertical axis scales represent deviation in variance ($\times 10^4$) due to shocks in each variable. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point.

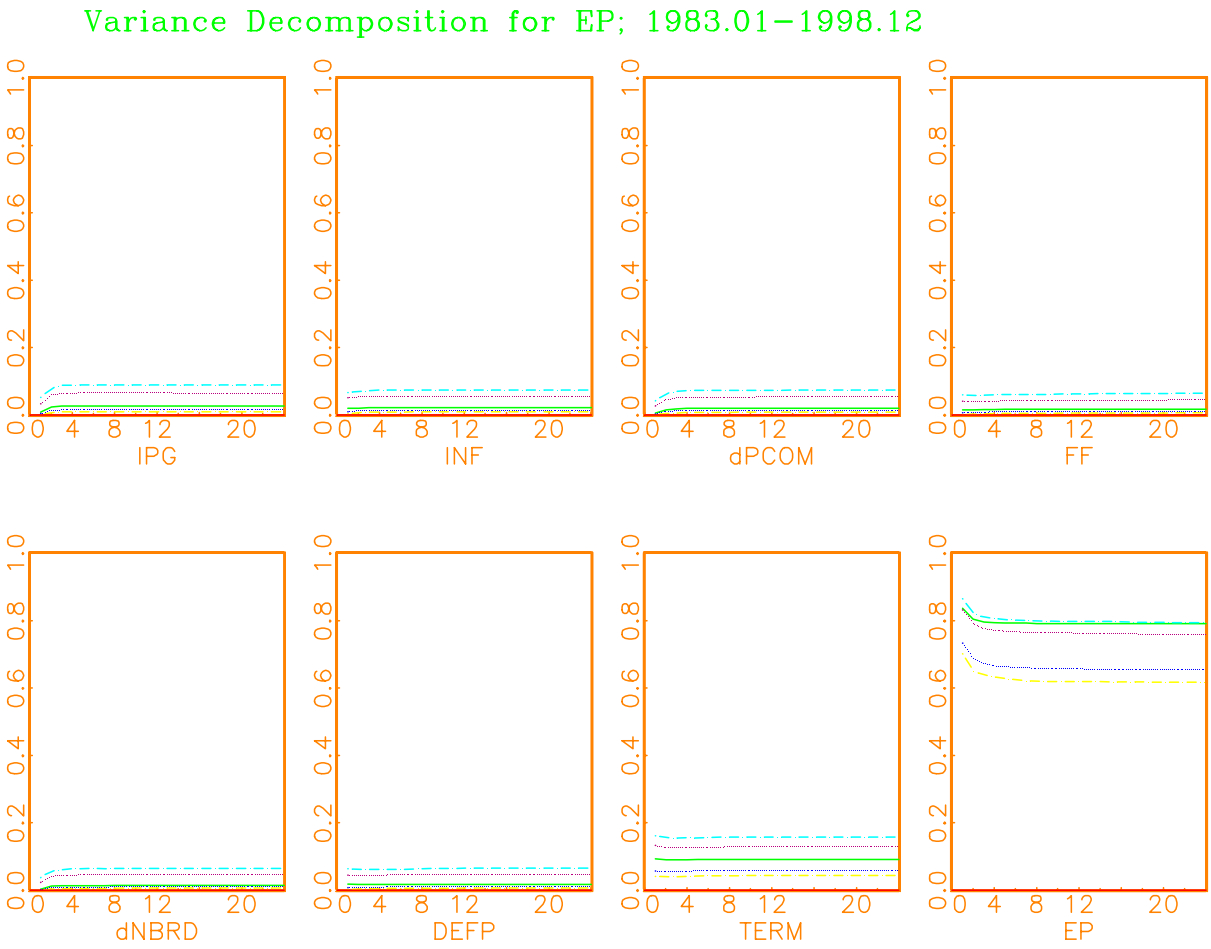

Figure 5b

Variance decomposition of excess return for post-Volcker period (1983:01–1998:12). Each graph is the cumulative response of variance of EP to a one-standard-deviation impulse in triangularly orthogonalized innovations in IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP. All variables are in annual percentage points. Each response is plotted over a 24-month horizon following the month of shocks. The vertical axis scales represent deviation in variance ($\times 10^4$) due to shocks in each variable. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point.

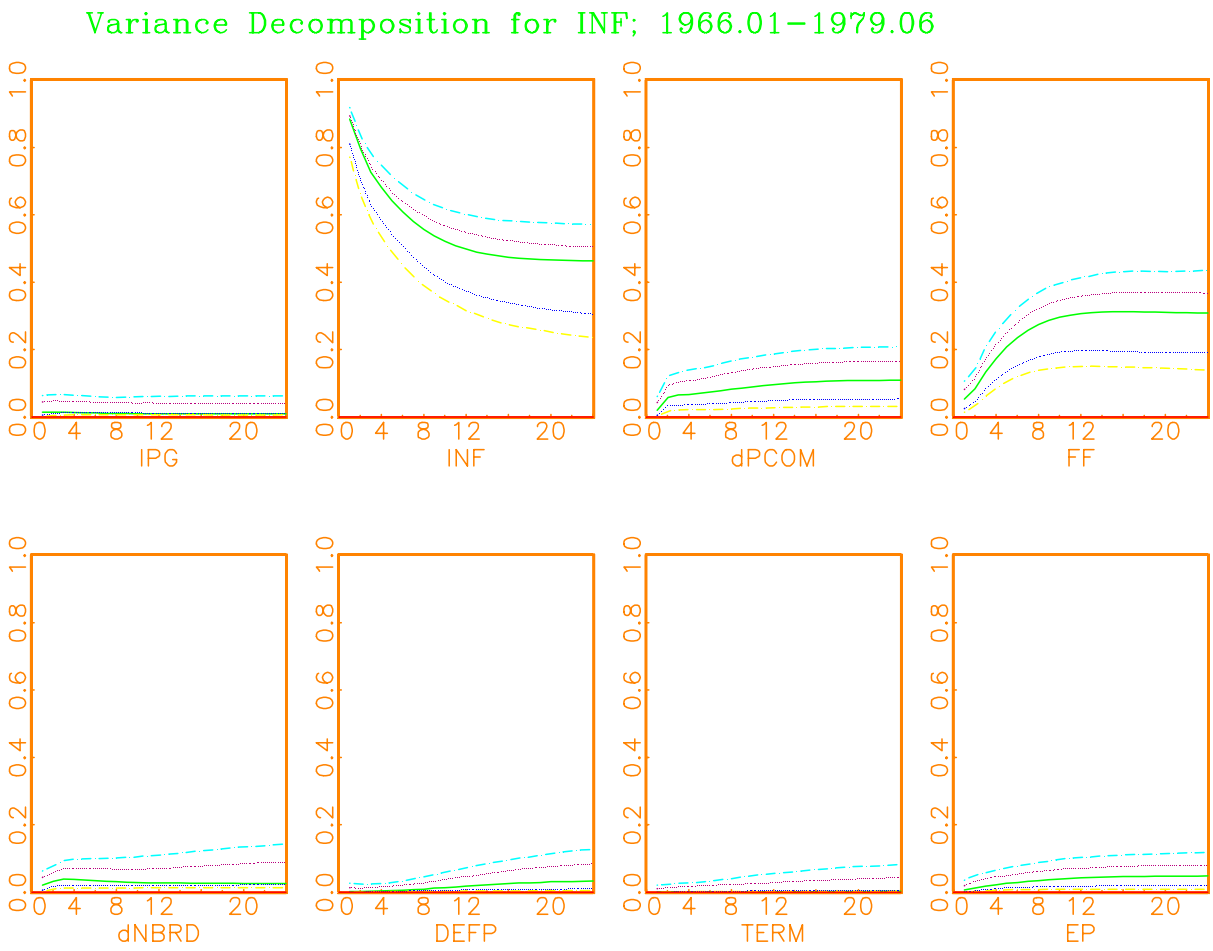

Figure 6a

Variance decomposition of inflation for pre-Volcker period (1966:01–1979:06). Each graph is the cumulative response of variance of INF to a one-standard-deviation impulse in triangularly orthogonalized innovations in IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP. All variables are in annual percentage points. Each response is plotted over a 24-month horizon following the month of shocks. The vertical axis scales represent deviation in variance ($\times 10^4$) due to shocks in each variable. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point.

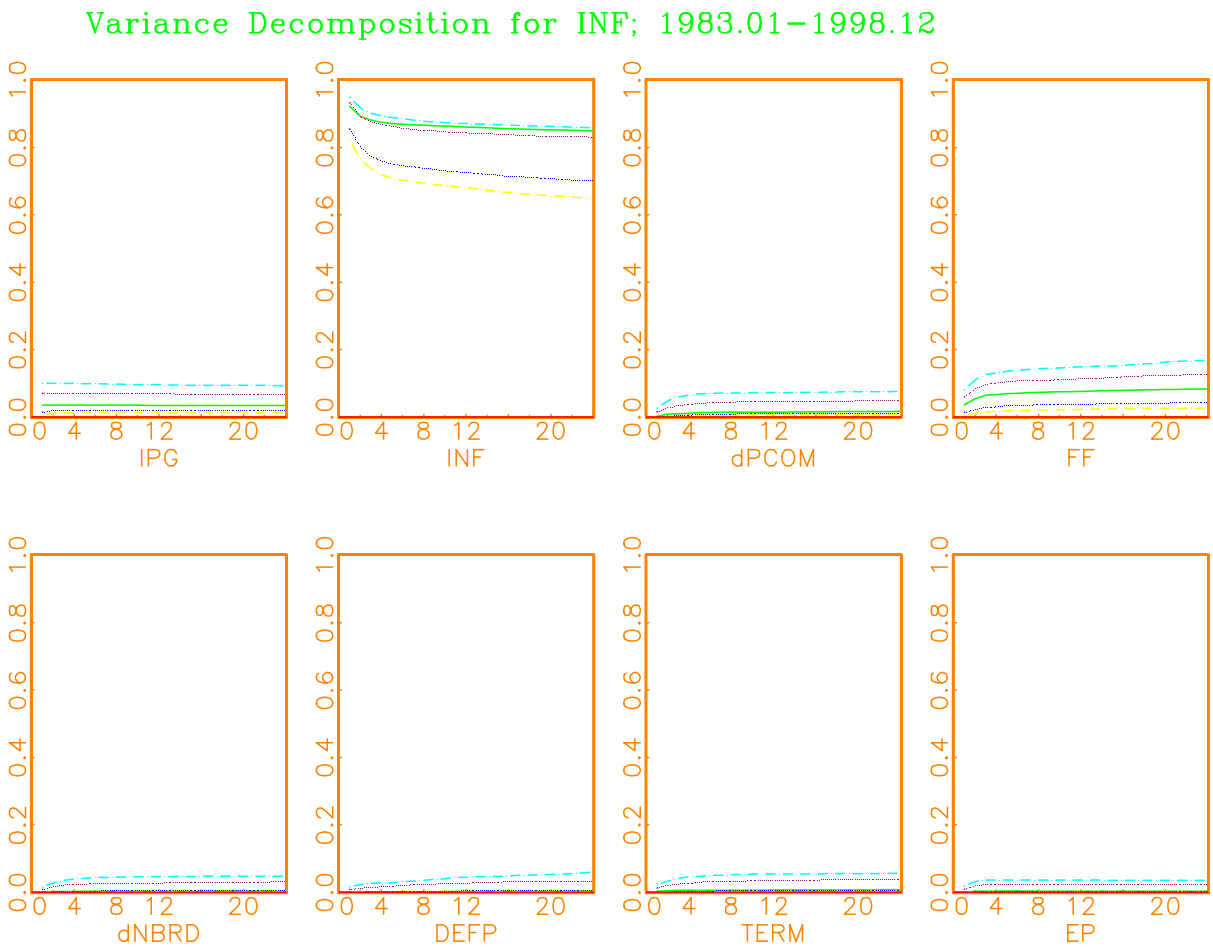

Figure 6b

Variance decomposition of inflation for post-Volcker period (1983:01–1998:12). Each graph is the cumulative response of variance of INF to a one-standard-deviation impulse in triangularly orthogonalized innovations in IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP. All variables are in annual percentage points. Each response is plotted over a 24-month horizon following the month of shocks. The vertical axis scales represent deviation in variance ($\times 10^4$) due to shocks in each variable. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point.

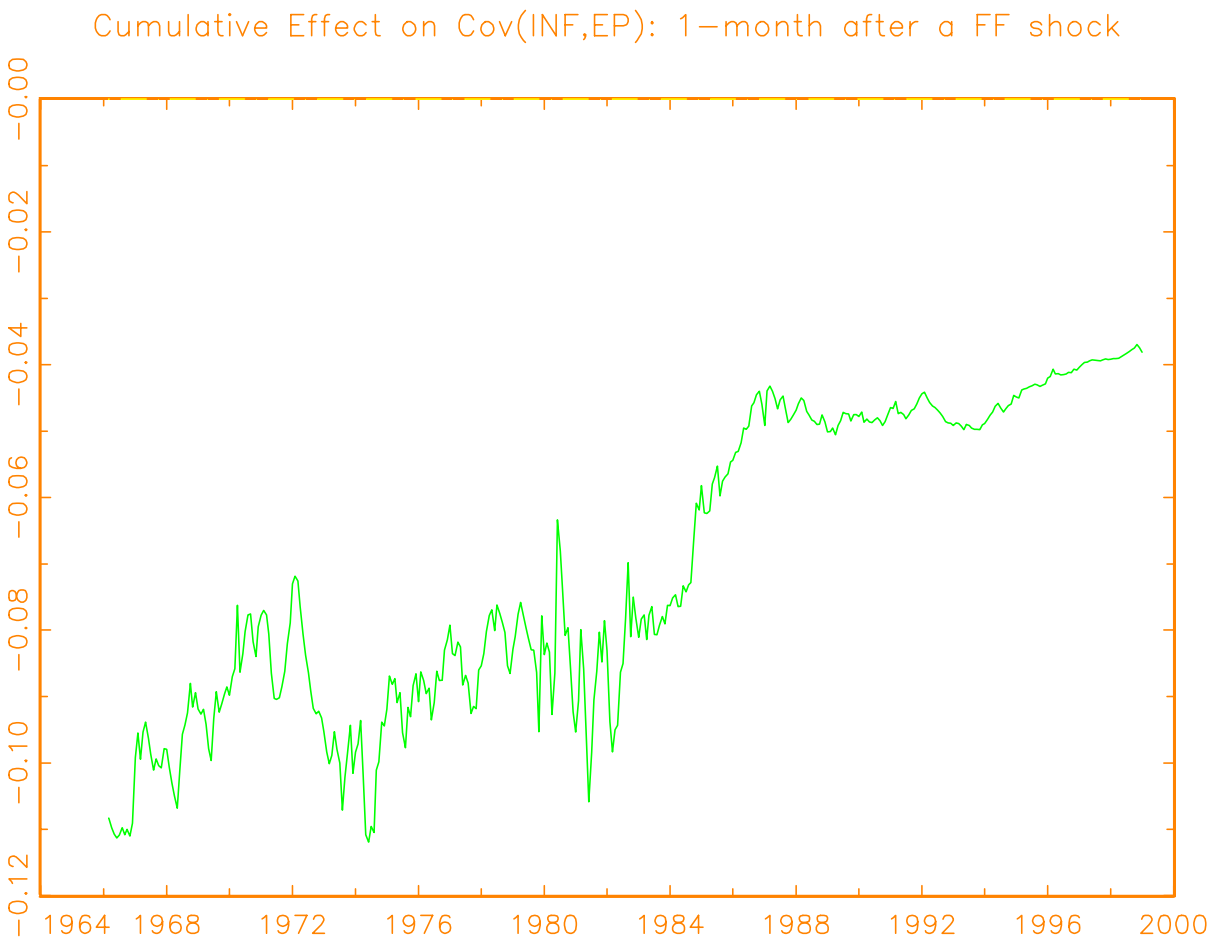

Figure 7a

Cumulative effects of a one-standard-deviation impulse in the federal funds rate on the covariance between inflation and excess return. This graph depicts the time-varying cumulative effects for a 1-month horizon following the month of the shock, across the sample period (1966-1998). Vertical axis scales represent deviation in covariance $\left(\times 10^4\right)$ due to shocks in FF. See section 3.3 for the estimation of time-varying VAR and time-varying impulse responses. In our time-varying VAR, all coefficients of the two policy equations – reaction functions for FF and DNBRD – are assumed to be time-varying, while the coefficients of other equations are kept constant. Point estimates are based on smoothed estimates in the time-varying VAR model. The cumulative effects on the covariance between inflation and excess return are computed as $\sum_{h=0}^{1} \theta_{h,t}\left(INF, FF\right) \times \theta_{h,t}\left(EP, FF\right)$, where $\theta_{h,t}\left(INF, FF\right)$ and $\theta_{h,t}\left(INF, FF\right)$ denote the h-month responses of INF and EP to a one-standard-deviation impulse in FF at period t.

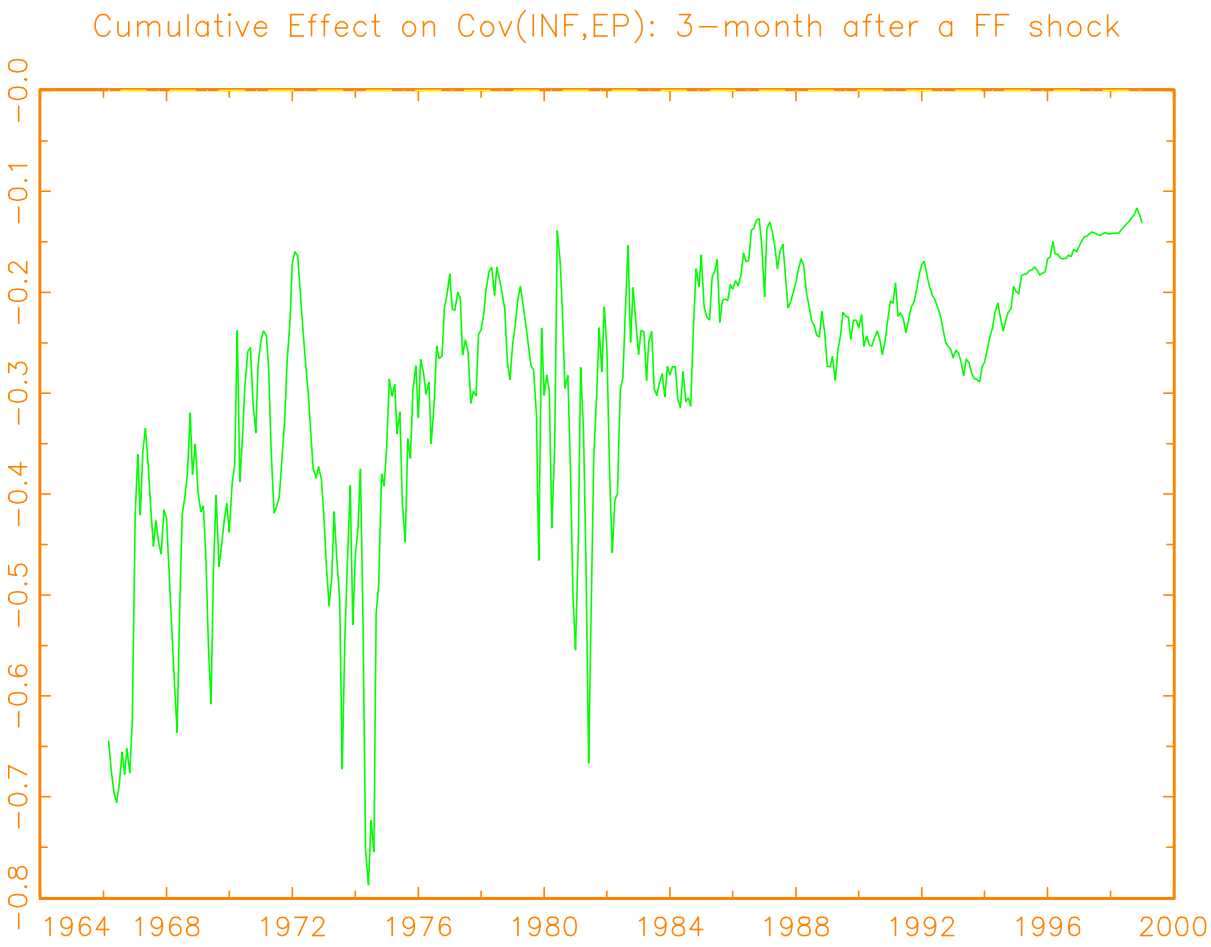

Figure 7b

Cumulative effects of a one-standard-deviation impulse in the federal funds rate on the covariance between inflation and excess return. This graph depicts the time-varying cumulative effects for a 3-month horizon following the month of the shock, across the sample period (1966-1998). Vertical axis scales represent deviation in covariance $\left(\times 10^4\right)$ due to shocks in FF. See section 3.3 for the estimation of time-varying VAR and time-varying impulse responses. In our time-varying VAR, all coefficients of the two policy equations – reaction functions for FF and DNBRD – are assumed to be time-varying, while the coefficients of other equations are kept constant. Point estimates are based on smoothed estimates in the time-varying VAR model. The cumulative effects on the covariance between inflation and excess return are computed as $\sum_{h=0}^{3} \theta_{h,t}\left(INF, FF\right) \times \theta_{h,t}\left(EP, FF\right)$, where $\theta_{h,t}\left(INF, FF\right)$ and $\theta_{h,t}\left(INF, FF\right)$ denote the h-month responses of INF and EP to a one-standard-deviation impulse in FF at period t.

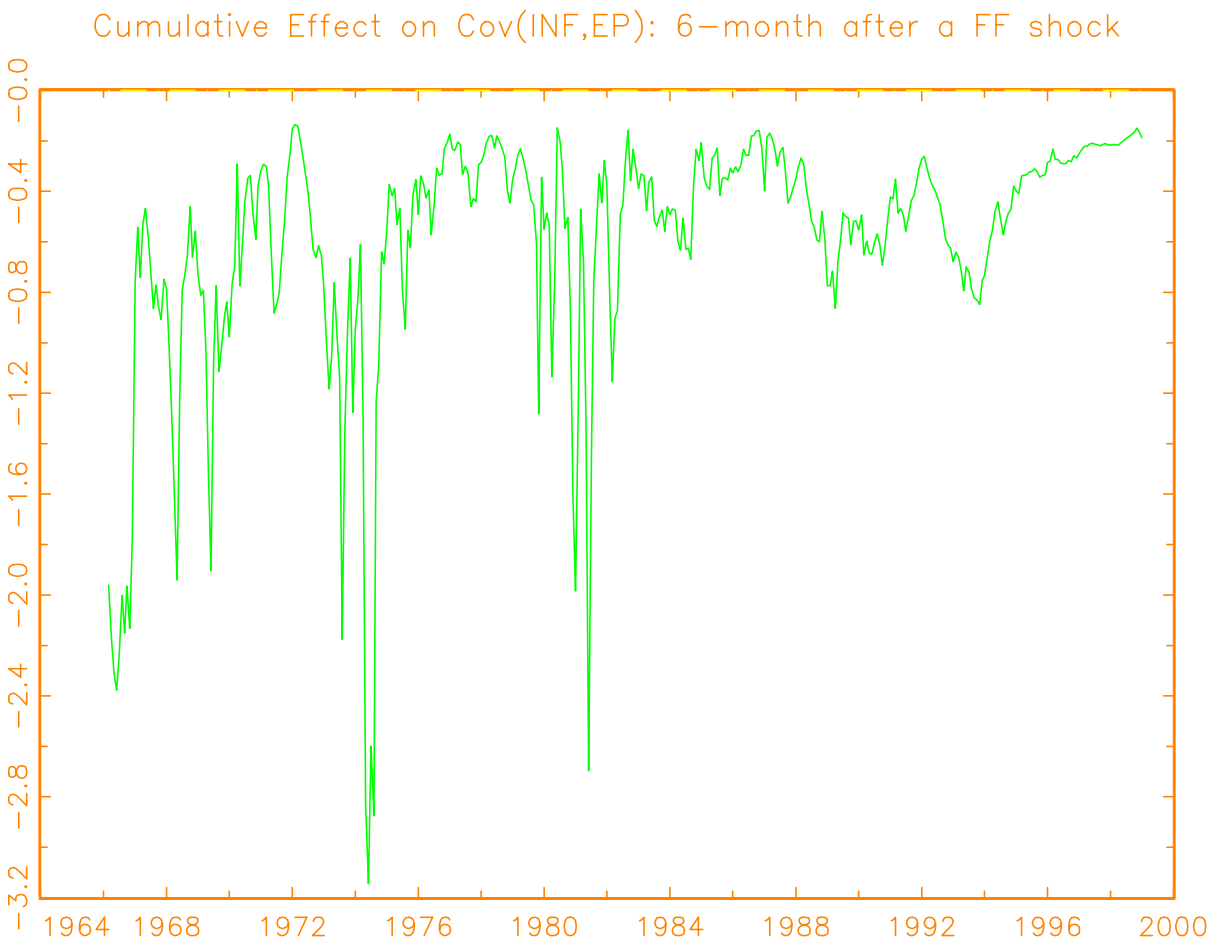

Figure 7c

Cumulative effects of a one-standard-deviation impulse in the federal funds rate on the covariance between inflation and excess return. This graph depicts the time-varying cumulative effects for a 6-month horizon following the month of the shock, across the sample period (1966-1998). Vertical axis scales represent deviation in covariance $\left(\times 10^4\right)$ due to shocks in FF. See section 3.3 for the estimation of time-varying VAR and time-varying impulse responses. In our time-varying VAR, all coefficients of the two policy equations – reaction functions for FF and DNBRD – are assumed to be time-varying, while the coefficients of other equations are kept constant. Point estimates are based on smoothed estimates in the time-varying VAR model. The cumulative effects on the covariance between inflation and excess return are computed as $\sum_{h=0}^{6} \theta_{h,t}\left(INF, FF\right) \times \theta_{h,t}\left(EP, FF\right)$, where $\theta_{h,t}\left(INF, FF\right)$ and $\theta_{h,t}\left(INF, FF\right)$ denote the h-month responses of INF and EP to a one-standard-deviation impulse in FF at period t.

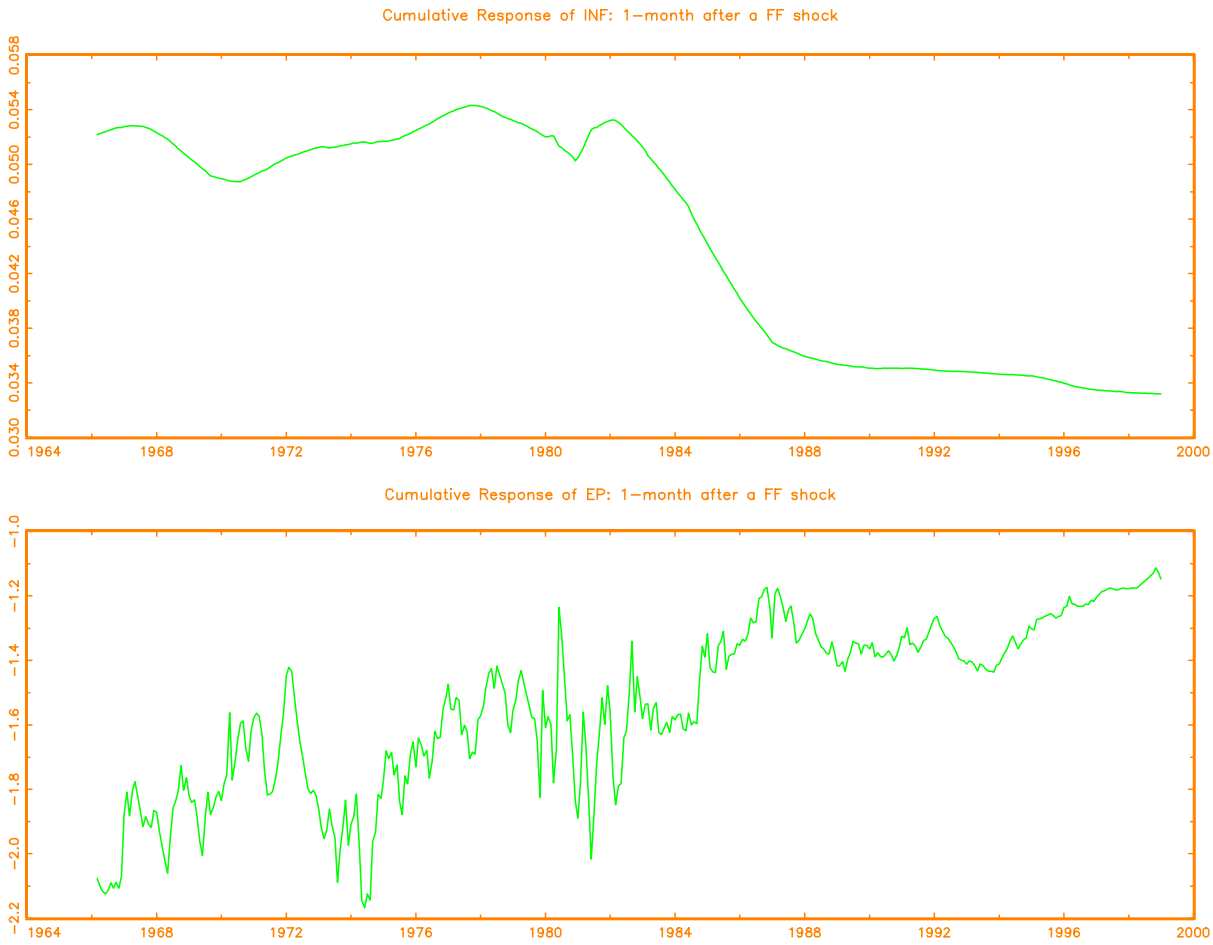

Figure 8a

Cumulative effects of a one-standard-deviation impulse in the federal funds rate on inflation and excess return. Each graph depicts the cumulative response of inflation and excess returns for a 1-month horizon following the month of the shock. Cumulative time-varying impulse responses are plotted across the sample period (1966-1998). Vertical axis scales represent percent deviation of variables due to the shock in FF. See section 3.3 for the estimation of time-varying VAR and time-varying impulse responses. In our time-varying VAR, all coefficients of the two policy equations – reaction functions for FF and DNBRD – are assumed to be time-varying, while the coefficients of other equations are kept constant. Point estimates are based on smoothed estimates in the time-varying VAR model. The cumulative responses are computed as $\sum_{h=0}^{1} \theta_{h,t}(INF, FF)$, and $\sum_{h=0}^{1} \theta_{h,t}(EP, FF)$,where $\theta_{h,t}(INF, FF)$ and $\theta_{h,t}(INF, FF)$ denote the h-month responses of INF and EP to a one-standard-deviation impulse in FF at period t.

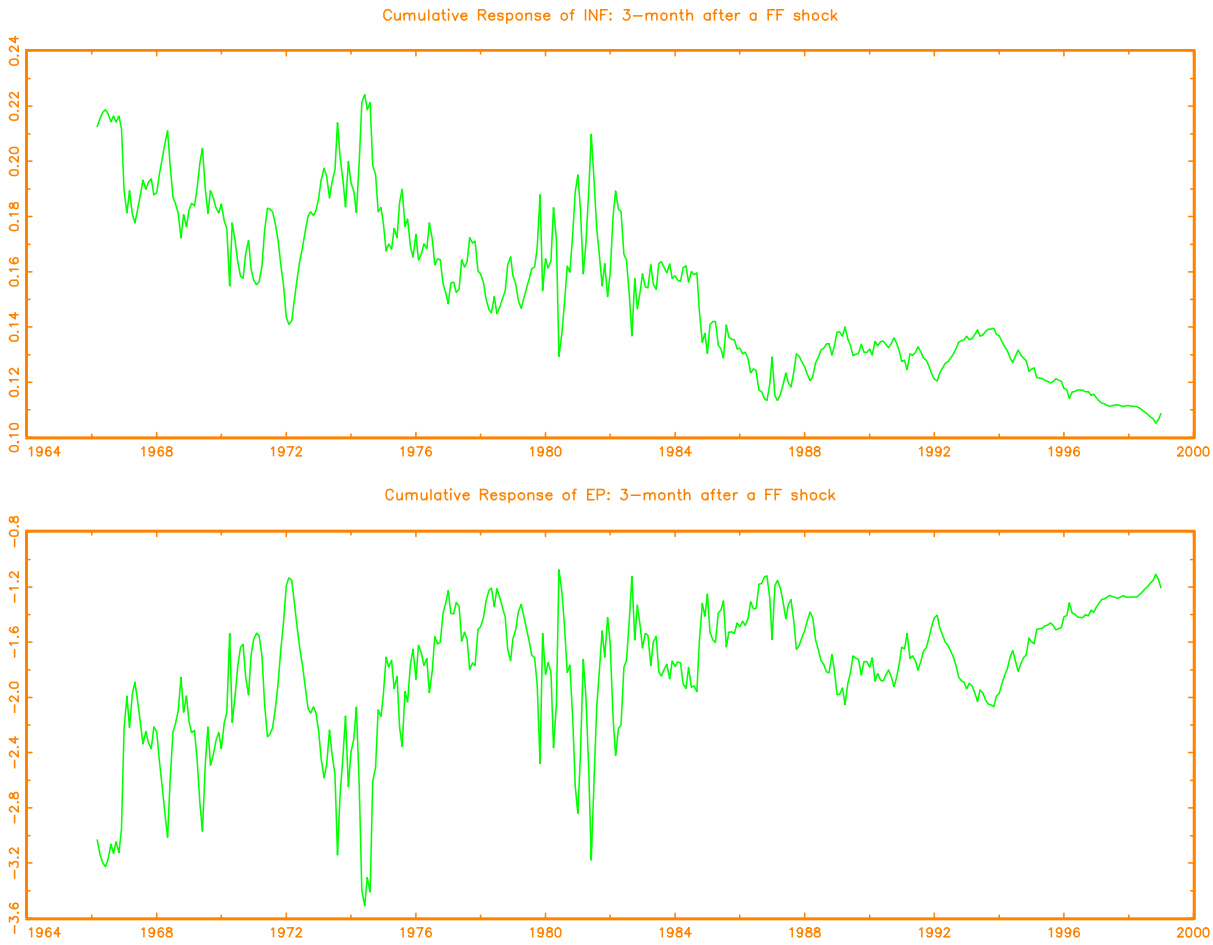

Figure 8b

Cumulative effects of a one-standard-deviation impulse in the federal funds rate on inflation and excess return. Each graph depicts the cumulative response of inflation and excess returns for a 3-month horizon following the month of the shock. Cumulative time-varying impulse responses are plotted across the sample period (1966-1998). Vertical axis scales represent percent deviation of variables due to the shock in FF. See section 3.3 for the estimation of time-varying VAR and time-varying impulse responses. In our time-varying VAR, all coefficients of the two policy equations – reaction functions for FF and DNBRD – are assumed to be time-varying, while the coefficients of other equations are kept constant. Point estimates are based on smoothed estimates in the time-varying VAR model. The cumulative responses are computed as $\sum_{h=0}^{3} \theta_{h,t}(INF, FF)$, and $\sum_{h=0}^{3} \theta_{h,t}(EP, FF)$,where $\theta_{h,t}(INF, FF)$ and $\theta_{h,t}(INF, FF)$ denote the h-month responses of INF and EP to a one-standard-deviation impulse in FF at period t.

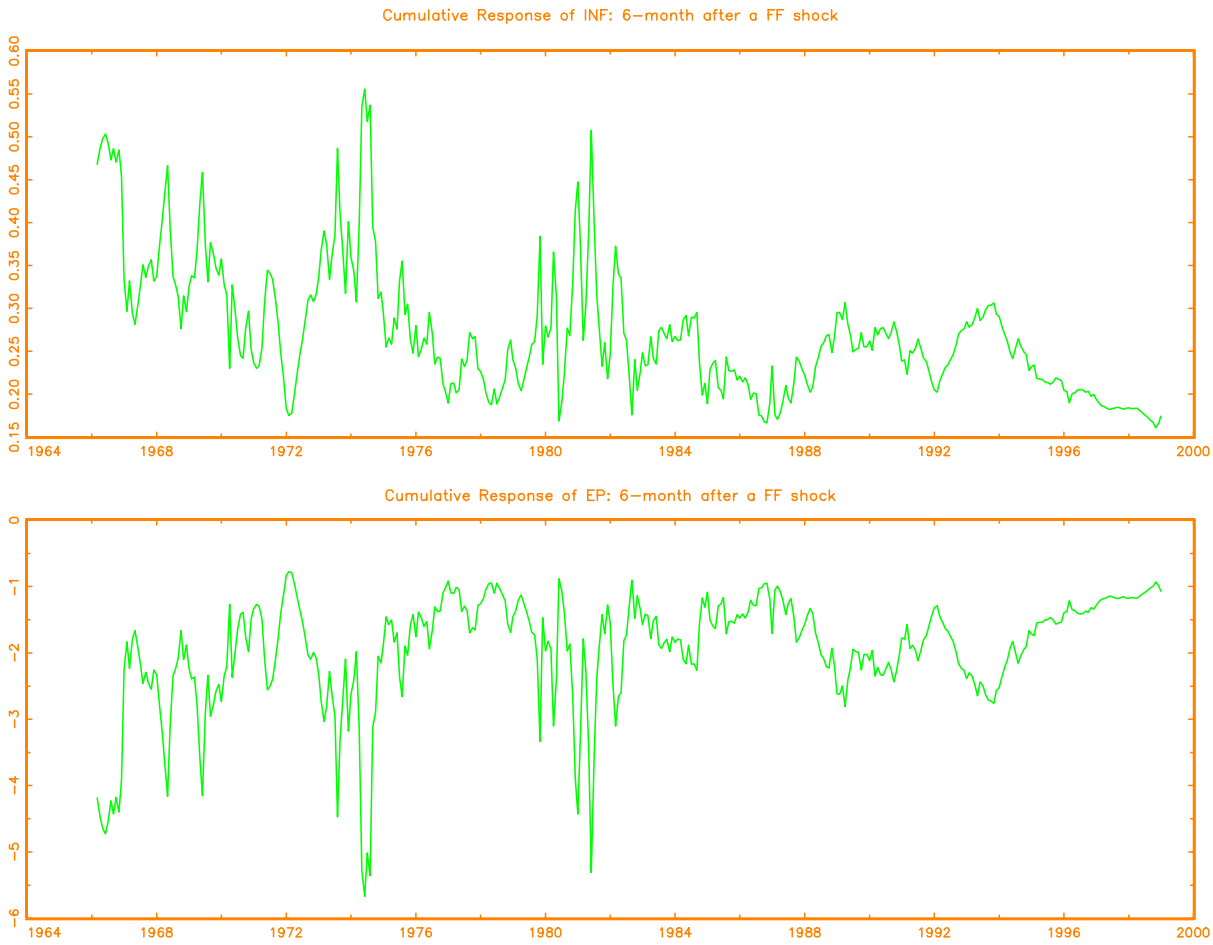

Figure 8c

Cumulative effects of a one-standard-deviation impulse in the federal funds rate on inflation and excess return. Each graph depicts the cumulative response of inflation and excess returns for a 6-month horizon following the month of the shock. Cumulative time-varying impulse responses are plotted across the sample period (1966-1998). Vertical axis scales represent percent deviation of variables due to the shock in FF. See section 3.3 for the estimation of time-varying VAR and time-varying impulse responses. In our time-varying VAR, all coefficients of the two policy equations – reaction functions for FF and DNBRD – are assumed to be time-varying, while the coefficients of other equations are kept constant. Point estimates are based on smoothed estimates in the time-varying VAR model. The cumulative responses are computed as $\sum_{h=0}^{6} \theta_{h,t}(INF, FF)$, and $\sum_{h=0}^{6} \theta_{h,t}(EP, FF)$,where $\theta_{h,t}(INF, FF)$ and $\theta_{h,t}(INF, FF)$ denote the h-month responses of INF and EP to a one-standard-deviation impulse in FF at period t.

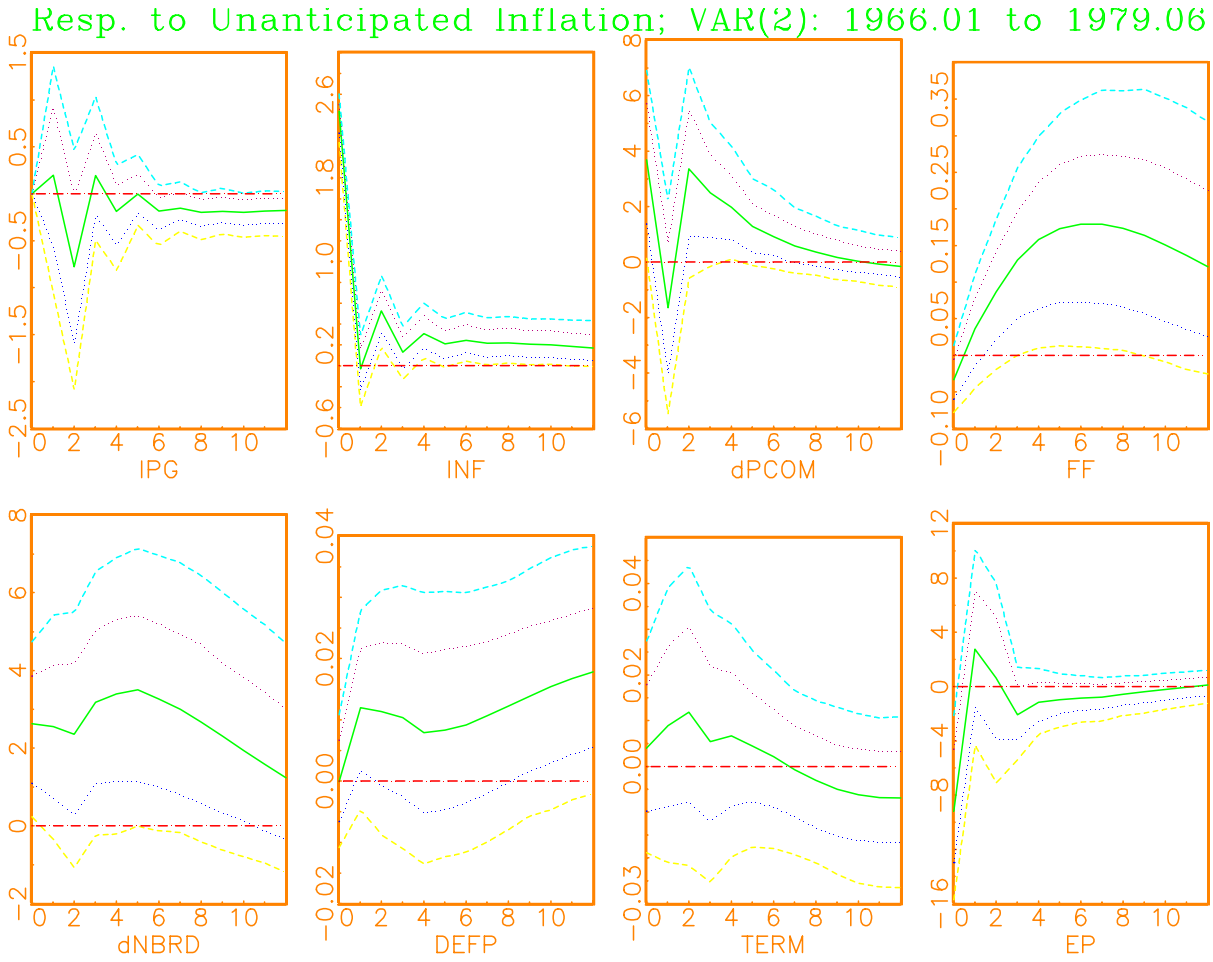

Figure 9a

Effects of unanticipated inflation on economic variables: pre-Volcker period (1966:01–1979:06). Each graph shows 12-month response of variables to a one-standard-deviation innovation in INF. Vertical axis scales represent percent deviation of variables. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point.

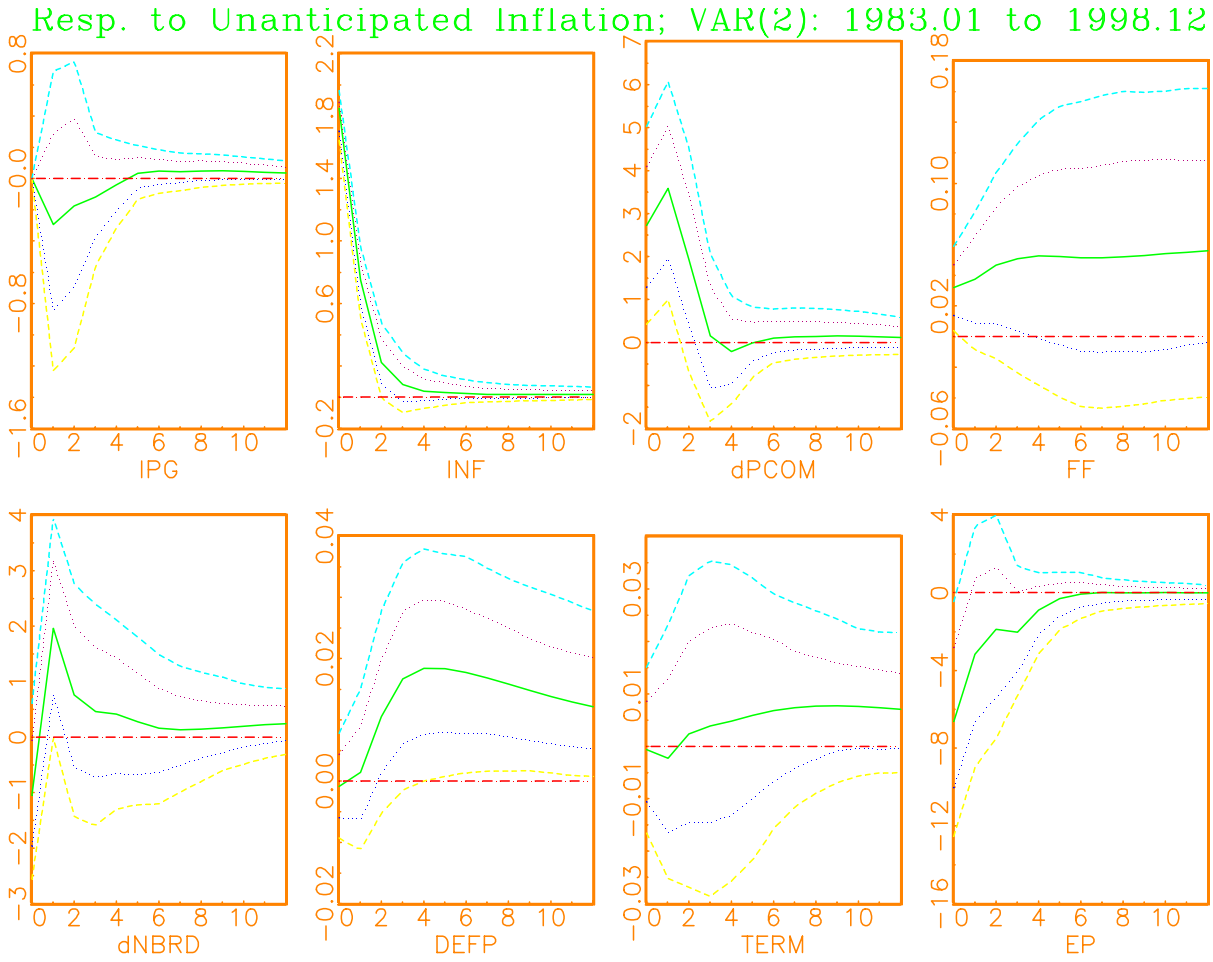

Figure 9b

Effects of unanticipated inflation on economic variables: post-Volcker period (1983:01–1998:12). Each graph shows 12-month response of variables to a one-standard-deviation innovation in INF. Vertical axis scales represent percent deviation of variables. Solid lines are point estimates, dotted lines are 68% error bands, and dashed lines are 90% error bands, estimated point by point.

Notes

[1]The original Federal Reserve Act of 1913, signed by President Woodrow Wilson, was much narrower in scope. However, the Banking Act of 1935, the Employment Act of 1946, the Full Employment and Balanced Growth Act of 1978 (a.k.a the Humphrey-Hawkings Act) established the objectives of the Federal Reserve System to be price stability, economic growth, a high level of employment, and moderate long-term interest rates.

[2]As additional evidence of how important a role the Fed plays in today's stock market, some practicioners have expressed a concern that Fed actions might be guided by the desire to stabilize or control asset prices. This spreading belief has prompted the Chairman of the Federal Reserve Board to make the following comment:

> "The persuasive evidence that the wealth effect is contributing to the risk of imbalances in our economy, however, does not imply that the most straightforward way to restore balance in financial and product markets is for monetary policy to target asset price levels, ...Leaving aside the deeper question of whether asset-price targeting is an appropriate government function, there is little if any evidence that monetary policy aimed at achieving that goal would be successful."
>
> —Alan Greenspan, as reported in the New York Times, April 06, 2000.

However, there is little, if any, evidence that the Fed might be setting its policy as a response to the stock market valuation.

[3]Sims (1986), Bernanke (1986), Blanchard and Watson (1986), Bernanke and Blinder (1992), Lee (1992), Cochrane (1994), Christiano et al.(1996,1998), Thorbecke (1997), Patelis (1997), etc.

[4]During the first three and a half years of Volcker, 1979-1983, also known as the "Volcker experiment", the Fed pursued a vastly different monetary policy. For more details, see the discussions in the paper.

[5]Moreover, as argued by Bernanke (1995) and Bernanke and Gertler (1995) monetary shocks can affect both the aggregate demand and the aggregate supply. Therefore, identifying aggregate demand and aggregate supply shocks and interpreting the former as money supply shocks might not be the best way to identify monetary policy.

[6]As explained below, we are decomposing the excess returns/inflation covariance into different unanticipated shocks, one of which is a monetary policy shock. A considerably more involved exercise would be to first decompose the covariance into excess returns, interest rate, and dividend growth components, and then decompose those components into monetary supply shocks. Such an exercise, although interesting, will divert us conciderably from the main point in this paper.

[7]Inflation is defined as the log difference in Consumer Price Index (CPI), following the convention in the literature.

[8]The VAR literature of the late 1980's and 1990's has established that it is better to focus on the federal funds rate as a tool of monetary policy, rather than the other monetary aggregates, such as M0, M1, or M2. (Sims (1980, 1986), Bernanke and Blinder (1992), Christiano et al.(1996), among others). Some papers advocate the use of non-borrowed reserves (NBR), and we include it in our analysis.

[9]For example, Romer and Romer (1989), Bernanke and Blinder (1992), Gali (1992), Bernanke and Mihov (1998), Christiano et al.(1996, 1998), and Leeper et al.(1996).

[10]Cochrane, 1998, p. 278

[11]For example, Bernanke et al. (1997), and Cochrane (1998) analyze the effect of anticipated (or endogenous) policy effects as well as unanticipated monetary shocks. Both studies conlcude that more attention must be devoted to systematic policy effects.

[12]Bernanke and Blinder (1992), Bernanke (1993), Cecchetti (1995), Friedman and Kuttner (1993), Hubbard (1995), Kayshap et al. (1993), Bernanke and Gertler (1995), and are some of the good articles in this area.

[13]This is one of the reasons why the commercial paper-Treasury bill spread is a good predictor of future economic activity.

[14]Most studies that use commodity prices to alleviate the price-puzzle, still exhibit an increase, albeit insignificant, of prices to contrationary monetary policy (Sims (1992), Christian et al.(1996), Cochrane (1998), and Bernanke et al.(1997))

[15]Romer and Romer (2000) do conduct some simple tests of the private information in Fed forecasts, because they have data of the actual Fed forecasts. Unfortunately, this data is not available to us.

[16]In chronological order, the chairmen are: McChesney, Burns, Miller, Volcker, and Greenspan.

[17]As defined by the NBER.

[18]Replacing the industrial production data with interpolated GNP/GDP data constructed as in Bernanke et al. (1997), Bernanke and Mihov (1998) did not change our qualitative results.

[19]The Fed abandoned its interest rate smoothing effort, and adopted a pure targeting of nonborrowed reserves from 1979:10 to 1982:10. Hamilton(1990), Ball and Torous(1995), and Gray(1996) show the excessive interest rate volatility during this period using regime-switching models.

[20]We estimate unconstrained VAR using a constant and 2 lags for each subperiod and using a constant and 3 lags for the full sample period, as determined by AIC(Akaike information criterion). Eight variables – IPG, INF, DPCOM, FF, DNBRD, DEFP, TERM, and EP – are included in this order.

[21]Leeper et al.(1996) cast doubt on treating commodity prices as predetermined for monetary policy shocks since they may respond to monetary policy shocks quickly in their thick auction markets. This problem may be significant in quarterly data. However, we did not find any qualitative changes in our results by placing DPCOM after monetary policy variables. See Bernanke and Mihov (1998) for similar observation.

[22]In other words, monetary policy variables respond to innovations in financial variables only with a one-month decision-lag. We do not consider this as restrictive since we use the federal funds rates are taken to be monthly averages.

[23]Strongin (1995), Bernanke and Mihov (1998), Christiano et al.(1998), and Walsh (1998) provide institutional details of the banking system to justify these approaches. For a review of identifying restrictions, see Leeper et al.(1996) and Christiano et al.(1998).

[24]When the system is stable, the Ψ_s are the coefficient matrices of the moving average representation of y_t by the Wold theorem. However, impulse responses are defined for unstable systems as well.

[25]This implies that systematic policy rules have no real effects. We turn to this point later.

[26]If the VAR system is stable, the h-period forecast error covariance matrix converges to the unconditional covariance matrix as $h \to \infty$, though we do not make any assumption on the stability of our system.

[27]The covariance decomposition is easily extended to the analysis of auto-covariances.

[28]Gallant et al.(1993) introduce conditional moment profile for computing impulse responses of nonlinear dynamic models.

[29]For recent developments in classical approach, we refer readers to the bias-corrected bootstrap-after-bootstrap method recently proposed by Kilian (1998).

[30]The RATS Bayesian Monte Carlo procedure has been widely used to compute first and second moments of the simulated distributions of impulse responses.

[31]Although we do not impose any over-identifying restrictions on the VAR, we note that this procedure is only approximately correct for over-identified models since the density of the restricted parameter space is not used. Sims and Zha (1999) propose a procedure for constructing Bayesian error bands for over-identified models using the Metropolis-Hastings algorithm.

[32]Boivin (1999), and Boivin and Watson (1999) propose an IV framework for time-varying parameter models in a more general setting.

[33]We allow for non-zero correlation among time-varying parameters.

[34]Standard numerical procedures tend to estimate the variances of time-varying parameters to be zero. See Boivin (1999), Boivin and Watson (1999) for discussion.

[35]A Monte-Carlo experiment, available upon request, indicates that none of the differences in the unrestricted terms of the two covariance matrices are significant at conventional levels.

[36]This assertion can be tested, for example, by looking at interest rate volatility as a proxy for economic instability. However, this is a project for future research.

[37]Lee (1992) finds little variation in real activity, which responds negatively to inflation innovations from his analysis of impulse responses. However, he does not construct error bands for his impulse responses. We find that the mildly negative impulse response he finds is in fact insignificant.

[38]This interpretation is commonly found in media: For example,

> Friday's plunge came after the government said prices at the consumer level showed surprising strength last month, triggering fears that the Federal Reserve may raise interest rates more aggressively. – "Unnerved investors rapidly unload stocks amid inflationary fears" by Catherine Tymkiw, April 14, 2000, http://www.cnn.com/

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