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Abstract

Recently, two stylized facts about the behavior of the U.S. economy have emerged: first, macroeconomic aggregates appear to be less volatile post-1984 than in the preceding two decades; second, monetary policy appears more responsive to inflationary pressures — and thereby more "stabilizing" — during the Volcker/Greenspan chairmanships relative to earlier regimes. Does a causal relationship exist between these two observations? In particular, has "better" policy by the Federal Reserve Board contributed significantly to the lessened volatility of the U.S. economy? This paper uses a structural vector autoregressive (VAR) specification to address these questions, examining the advantages and limitations of such an approach. In contrast with much of the existing research on these topics, I find that most of the quantitatively significant changes in volatility are attributed to breaks in the non-policy portion of the structural VAR, and not to the identified policy equation.

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1 Introduction

Has the prolonged U.S. economic expansion of the past few decades been due to good fortune or good policy? In the popular and business press, the Federal Reserve Board — and Chairman Alan Greenspan in particular — often receives accolades for the sustained growth and moderate to low inflation experienced since the mid–1980s. Recent academic work also has focused upon the Fed, crediting monetary policymakers for a share of the recent economic outcomes while blaming them for the poor performance experienced in the late 1960s and 1970s. This literature raises a pair of questions. First, is the greater stability experienced by the U.S. economy recently — as documented by McConnell and Perez-Quiros (2000), Blanchard and Simon (2001), and Stock and Watson (2002), for example — due primarily to an "improvement" in monetary policy, or have policy makers benefited from some transformation in the economy itself? Put more glibly, has Alan Greenspan been smart or lucky — or both?

Once attention is focused upon monetary policy, a second question naturally arises: what attributes of policy have changed? In assessing the role of monetary policy, it is important (indeed, from the standpoint of econometric identification, necessary) to distinguish between *exogenous* (unforecastable) innovations to policy and the *endogenous* response of policy to the state of the economy. This dichotomy determines whether changes in the policy instrument can be prescribed to factors other than actions by the monetary authority. Interestingly, the two main empirical literatures on monetary policy place very different emphasis on each of these components.

Beginning with Taylor (1993), a large literature has investigated simple specifications of a feedback rule for monetary policy. Usually no attempt is made to interpret or explain the residuals of the estimated rule. Recent work in this "Taylor rules" literature (e.g. Judd and Rudebusch, 1998; Taylor, 1999; Clarida et al., 2000) provides a single direct answer to the pair of questions posed at the opening of this paper: by failing to respond forcefully enough to inflationary pressures, "inappropriate" Federal Reserve policy introduced instability into the economy prior to 1980. In contrast, a conventional view of U.S. economic performance during these years posits an unavoidable policy dilemma in the face of adverse supply shocks (and possibly domestic political pressure).[1] While the choice between high inflation or high unemployment may have been difficult and undesirable, this more traditional view does not regard the Fed's policy response as negligent.

[1] DeLong (1997) provides an informative discussion of the policy environment in the late 1960s and 1970s.

It is important to recognize that much of the Taylor rules literature *presumes* monetary policy to be the culprit: single equation models cannot address the question of changing dynamics in the broader economy, while small, stylized macroeconomic models that fail to account for instability elsewhere could lead to improper inference about the role of the policy rule — especially if the true nature of the policy rule is sensitive to the equations that govern the broader economy.

The structural vector autoregressive (SVAR) literature represents the second main empirical approach. Until fairly recently, this literature had focused almost exclusively upon the effects of exogenous monetary policy shocks, largely ignoring the estimated monetary policy rule or reaction function. Yet the shocks account for a very small fraction of the total variation of the policy instruments; most is due to the endogenous component of policy. This lack of attention to the estimated reaction function has been a subject of criticism (e.g. Cochrane, 1994; Rudebusch, 1998; McCallum, 1999).

In a widely-cited paper, Bernanke and Mihov (1998) address this issue with a "semi-structural" VAR model and find evidence of regime changes in U.S. monetary policy in both 1979 and 1982, but report limited evidence of structural instability in the VAR coefficients. In their *Handbook of Macroeconomics* chapter, Christiano et al. (1999) summarize much of the structural VAR literature at the time, and report no evidence of structural instability across their estimates. Cogley and Sargent (2001) use Bayesian techniques in a three-variable reduced-form VAR and report to find evidence of changes in an implied monetary policy rule like that of Clarida et al. (2000); Cogley and Sargent (2005) repeat this exercise in a specification that allows for breaks in both the variances and the coefficients of the VAR, and re-affirm most of their earlier results. Sims (2001) criticizes the results based on the implied policy rules from these two studies along similar lines as the above criticisms of the Taylor rules approach. Sims and Zha (2004) use a Bayesian regime switching model of a fully-identified, non-recursive SVAR. They report some evidence of regime switching in U.S. monetary policy, but their preferred specification finds instability only in the variances of the structural shocks and not in the estimated coefficients. Overall, there does not yet appear to be a consensus in the VAR literature on the sources or significance of instability due to changes in the practice of monetary policy.

In trying to disentangle the questions posed above, it is essential to take into account the feedback between policy and the other variables of the model — particularly in a dynamic setting. Therefore, I investigate a semi-structural VAR model, in which the identifying restrictions yield an estimate of the policy rule, as outlined in section 2. My approach is similar in spirit to that of Sims and Zha (2004), but I work with existing structural VAR models estimated in a classical statistical paradigm. Section 3 examines the empirical findings, starting with the properties of the estimated reaction function across various sub-sample periods, then explores the possibility of significant instability residing in the non-policy part of the model. Sections 3.1 and 3.2 pursue methods to quantify the economic significance of some of the instability uncovered in section 2.3. To foreshadow the main result, despite some evidence of instability in the practice of monetary policy, these changes do not appear to be able to account for most of the varying economic outcomes observed in the data. Section 3.3 explores some interpretations of the results, and section 4 concludes.

2 Monetary VAR Specification and Estimation

2.1 Specification and Identifying Restrictions

A variety of identifying assumptions can be found in the vector autoregressive (VAR) literature that attempts to quantify the effects of monetary policy. In this paper, a semi-structural block-recursive framework is employed. This approach is common in the literature: see Leeper et al. (1996), Bernanke and Mihov (1998) or Christiano et al. (1999) for some prominent examples.[2] The data (described below) are separated into two groups or blocks: those that measure macroeconomic activity (such as output and the aggregate price level), and those that involve monetary policy. The latter group can be further divided into the policy instrument of the Federal Reserve (in this study, the federal funds rate) and other variables in the market for bank reserves.

Formally, let X_t be the $n \times 1$ vector of endogenous variables in the model, partitioned as $X_t = [Y_t \;\; M_t]'$. Y_t is a vector of "activity" variables, comprised of monthly measures of output (log of the industrial production index), the price level (log of the consumer price index[3]) and an index of commodity prices (also in logs). M_t is a vector of policy variables: the Federal funds rate, nonborrowed reserves and total

[2] Other authors characterize this class of models as "partially identified;" as noted below, only a subset of the shocks are imbued with a structural economic interpretation. Keating (1996) provides a detailed investigation of block-recursive identification techniques in VARs.

[3] Due to mismeasurements of the CPI prior to 1983 resulting from the method used to impute the cost of owner-occupied housing, the CPI excluding shelter is used.

reserves (both in logs).[4] A VAR model then can be written as:

$$\Phi(L)\,X_t = \varepsilon_t\,, \tag{1}$$

where ε_t is an n-vector of unobserved structural disturbances with the covariance structure $E[\varepsilon_t \varepsilon_t{}'] = \Omega$. Since the structural disturbances are mutually and serially uncorrelated by assumption, the covariance matrix Ω will be diagonal.

Pre-multiplying equation (1) by Φ_0^{-1} yields an expression for the reduced-form model,

$$A(L)\,X_t = \upsilon_t\,, \tag{2}$$

where $A(L) = I - A_1 L - A_2 L^2 - \cdots - A_p L^p$ is a lag polynomial of order p, and A_k is the $n \times n$ matrix of coefficients such that $A_k = \Phi_0^{-1}\Phi_k$ for $k = 1, \ldots, p$, and I is the $n \times n$ identity matrix.[5]

Given the division of the data into the Y_t and M_t vectors, the reduced-form model of equation (2) can be expressed in terms of the structural parameters as:

$$\begin{bmatrix} (\Phi_0^{-1})_{YY} & (\Phi_0^{-1})_{YM} \\ (\Phi_0^{-1})_{MY} & (\Phi_0^{-1})_{MM} \end{bmatrix} \Phi(L) \begin{bmatrix} Y_t \\ M_t \end{bmatrix} = \begin{bmatrix} (\Phi_0^{-1})_{YY} & (\Phi_0^{-1})_{YM} \\ (\Phi_0^{-1})_{MY} & (\Phi_0^{-1})_{MM} \end{bmatrix} \begin{bmatrix} \varepsilon_{Y,t} \\ \varepsilon_{M,t} \end{bmatrix}.$$

Identification of the monetary policy shock requires sufficient restrictions be placed on the Φ_0^{-1} matrix. In the block-recursive scheme employed in this paper, the upper right-hand block, $(\Phi_0^{-1})_{YM}$, is set to zero. This restriction is justified by assuming sluggish adjustment of the variables in the macroeconomic activity block to innovations in the policy sector: examples include lags in the ability to gather and process information about the monetary sector, or sluggish adjustment of price setters or production decisions. Specifically, the activity block is assumed to react to changes in $\varepsilon_{M,t}$ with a one period lag. Conversely, the $(\Phi_0^{-1})_{MY}$ block is left unrestricted, thus allowing each of the policy variables to react to contemporaneous movements in all of the activity variables.

[4] While unit root tests on most of the series in this specification tend not to reject a null of non-stationarity, I follow previous authors and estimate the VAR with log levels. Bernanke and Mihov (1998) report few differences between estimation in log differences and in log levels. Sims et al. (1990) report that estimation in (log) levels of a VAR model with an unknown number of unit roots is consistent for sufficient lag length.

[5] The leading identity matrix reflects the normalization of placing the variable x_{it} on the left-hand side of the i^{th} equation of the reduced-form system, which then can be estimated consistently by OLS. This notation ignores intercept terms for convenience; they are included in estimation.

To identify the monetary policy reaction function and structural monetary policy shock, additional restrictions must be placed on $(\Phi_0^{-1})_{MM}$. Over most of the sample the federal funds rate serves as the policy instrument, and the Fed accommodated changes in reserves demand. Consistent with Christiano et al. (1996a,b) and Bernanke and Mihov (1998) the funds rate thus is ordered first in the M_t block. A recursive structure for the $(\Phi_0^{-1})_{MM}$ block implies that the funds rate does not respond to contemporaneous innovations to the reserves market variables. It does respond to lagged movements in these variables, which reflect a combination of supply and demand shocks within the market for reserves. Under these identifying assumptions, the residuals of the funds rate equation capture the structural monetary policy shocks: the unforecastable component of changes in the policy instrument. The funds rate equation in the structural model is associated with the monetary policy reaction function: the component of the policy instrument that responds endogenously to the state of the economy.

The above restrictions are sufficient for the identification of the exogenous component of monetary policy. The dynamic effects of these monetary policy shocks can be found by inverting $\Phi(L)$ in equation (1) to construct the structural vector moving-average representation:

$$X_t = \Theta(L)\,\varepsilon_t\,, \tag{3}$$

where $\Theta(L) = \Theta_0 + \Theta_1 L + \Theta_2 L^2 + \cdots = \Phi(L)^{-1}$ is a (possibly infinite-order) lag polynomial. The estimated coefficients in $\Theta(L)$ determine the impulse response functions to the structural shocks, ε_t; the responses to the estimated monetary policy shock are analyzed in section 3.

2.2 Estimation

The model is estimated with data from January 1966 through August 2001.[6] Strongin (1995) and Meulendyke (1998) note that the modern form of the reserves market came into existence in the mid-1960s. As emphasized by Strongin (1995) and Bernanke and Mihov (1998), the Fed shifted to a non-borrowed reserves targeting regime during the Volcker disinflation.[7] As a result, there are *a priori* reasons to isolate this period from the rest of the sample: once the policy instrument shifted from the funds rate to non-borrowed reserves, the residual of the fed funds equation no longer would represent exogenous policy

[6] The Akaike Information Criteria suggests a lag length of 8 months. The impulse responses shown below are fairly similar with a one-year lag length (which is fairly common in the VAR literature) and with 14 lags (as in Bernanke and Mihov, 1998).

[7] See Meulendyke (1998) and Walsh (2003) for further discussion of the change in operating procedures after October 1979.

shocks.[8] Failure to account for such a change in the operation of monetary policy would result in a misspecified model, incorrect impulse response functions, and thereby false inference about the role of policy.[9] Sims and Zha (2004) also find evidence of a distinct policy regime during the Volcker disinflation.

Monthly data are used as the plausibility of the above timing restrictions for identification depends critically on the length of a "period;" much of the VAR literature uses quarterly data, which implies a full three months between the time a monetary policy action is undertaken and the time agents are allowed to respond. Moreover, these timing restrictions are often motivated as approximations of state-contingent behavior, and such approximations are less appropriate as the time that agents are not permitted to respond lengthens.[10]

2.3 Stability of Reduced-Form Estimates

Table 1 examines the Akaike and Schwarz Information Criteria for VAR models estimated across several combinations of sample period and lag length. The information criteria are log likelihood statistics, with penalties for additional parameters. (The SIC penalizes additional parameters more heavily than the AIC.) Heuristically, the lower the AIC or SIC measure, the better the "fit" of the equation. These information criteria are computed for the reduced-form models in equation (2).

Sims (1998a,b) has argued in favor of using these measures to examine the overall stability of a VAR model. By this metric, a model estimated over a full sample is preferred to one estimated separately over two sub-sample periods if the former has a lower value of the AIC or SIC than a weighted-average of the information criteria for the latter case. The weights are determined by the proportion of the full sample in each sub-sample estimation.[11]

Table 1 permits an analysis of structural instability for two variants of the reduced-form model. The first is estimated over the full 1966:01 – 2001:08 sample, and breakpoints after October 1979 and December 1982 — the onset and conclusion of the Volcker disinflation experiment, respectively — are considered. The second investigates a joint sample with the intervening 1970:10 – 1982:12 period excluded. As cited above, some authors have suggested that, outside of the "Volcker disinflation," there has not been

[8] Notice that it is not sufficient to simply impose a break at October 1979 in light of the shift in operating procedures.

[9] Because the identifying assumptions employed by the above-cited Bayesian literature do not depend upon the particular policy instrument in the reserves market, they do not face the same issues when confronting the Volcker disinflation period.

[10] If the true delay in response of any of the Y_t variables to innovations to $\varepsilon_{M,t}$ is three months (or more), the monthly VAR will recover that behavior in the estimated coefficients of the appropriate impulse response functions.

[11] This result can be shown by comparing the log likelihood functions for the two cases. I thank Phil Howrey for helping to clarify this point.

any statistically discernible change in the monetary policy regime.

The first row of table 1, labeled "Fixed Coefficients," lists the AIC and SIC values for the reduced-form model estimated over the 1966:01 –2001:08 period. The second row lists the same statistics for a model that allows for a break in all coefficient values (and the variances) after October 1979. Notice that at all reported lag lengths, the model that allows for a break is preferred (i.e. has a lower value for the information criteria) to the model with fixed coefficients. The third row allows for a break after 1982:12, with similar (if slightly less strong) results.[12]

For each lag length considered the joint sample period, which excludes the Volcker disinflation period, always has a noticeably lower value of the reported statistics than does the full sample. Thus, the statistical evidence on model fit supports the exclusion of the 1979:10 – 1982:12 period as well. Interestingly, the AIC always favors treating the joint sample period as two distinct samples, whereas the SIC suggests the best fitting model does not feature a break. These results merit further investigation of the nature of changes in the stability of the VAR even after accounting for the (temporary) disruption caused by the Volcker experiment.[13]

Table 2 shifts to examining the stability of the individual equations. Panel (*a*) reports the p-values for Wald tests of parameter stability of each equation of the reduced-form VAR model. The null hypothesis of cross-sample stability can be strongly rejected for the CPI, the funds rate, and total reserves with 8 lags in the VAR. At longer lag lengths the null for each of the variables is rejected at conventional levels of significance. A Wald test is used instead of a Chow test as the former is robust to changes in the variance across the sub-samples; panel (*b*) reports p-values for the Goldfeld-Quandt test of homoskedasticity, which indicate rejection of a constant variance for four of the six variables in the VAR (at all lag lengths). Interestingly, the CPI is one of the variables for which the variance appears constant across sub-samples.

All of the variables in the model exhibit instability in either their estimated coefficients or their variance. The rejection of stability for the fund funds equation coefficients is not surprising in light of the results reported in the Taylor rules literature. However, both the VAR equation in this section and a Taylor rule are reduced-form relationships. Stability tests on reduced-form equations cannot determine the

[12] Note that the results in table 1 are not directly comparable across different lag lengths: the initial values of the VAR are drawn from within the stated sample period, thereby making the number of observations used for estimation vary inversely with the number of lags.

[13] These results imply that including the full sample would increase the likelihood of rejecting the null hypothesis of stable coefficients, as the model fit is worse in this case. All of the subsequent analysis is conducted with a VAR system that excludes the Volcker disinflation time period.

economic sources of instability. In particular, so long as the Fed's understanding of the nature of the macroeconomy informs the policy decisions of the FOMC (as would be the case for both optimal policy and certain classes of robust policy rules), the coefficients in a Taylor rule are not themselves preference parameters, but functions of both the Fed's preferences — the relative weights on the arguments of its objective function — and the constraints imposed by the structural equations that describe the dynamics of the macroeconomy. Any changes in the equations that determine the non-policy variables of the model — which are also quite consistent with the results in table 2 — also should lead to instability in the coefficients of a Taylor rule, even if the preferences embedded in the Fed's objective function remain unchanged.

3 Sources of Regime Variation

While a change in operating procedures necessitates segmenting the sample, it does not immediately imply that the practice of monetary policy — or the dynamic behavior of the economy more generally — differs between the periods before and after the Volcker disinflation. In this section I explore the magnitude and importance of any changes over time in the identified structural blocks of the model.

Figure 1 compares the impulse response functions to a 25 basis point contractionary monetary policy shock for four of the main variables in the VAR model of section 2.[14] The VAR is estimated over three sample periods: the first column reports the impulse responses in the full sample from 1966:01 to 2001:08 (excluding, as above, the Volcker disinflation period); the second column represents a "pre-Volcker" sample (1966:01 – 1979:08) and the third is a "post-Volcker" sample (1983:01 – 2001:08). In each panel, the solid line represents the point estimates of the impulse response function, while the dark and light grey shaded regions represent the 68% and 95% bootstrapped confidence intervals, respectively.[15]

Several features are common to all three estimation periods. First, the initial response of output to a contractionary policy shock is positive, albeit not statistically significant and short-lived, followed by a protracted negative response. A studied in detail in Hanson (2004), with monthly data each period exhibits a "price puzzle:" the CPI price level increases in response to a monetary policy shock for a protracted period of time. These positive responses are generally not statistically significant, and are

[14]While the size of a "typical" monetary policy shock varies across the sample periods, the magnitude of the shock is normalized in figure 1 to facilitate comparison.

[15]See Kilian (1998, 1999) for a description of the bias-corrected bootstrap procedure used herein.

mitigated by the inclusion of commodity prices in the VAR.[16] The fed funds rate illustrates a persistent response to the initial 25 basis-point contractionary policy innovation, even rising further for a few months following the shock. The higher interest rates are generally associated with a small decline in non-borrowed reserves, consistent with a liquidity effect.[17]

Splitting the sample reduces the statistical significance of the output response. In the earlier sample (1966 – 1979) output is below its pre-shock value for about two years, and the point estimate returns to slightly (though not statistically significantly) above the initial value about 3 years following the structural policy innovation. On the other hand, the impulse responses in the latter sample period (1983 – 2001) generally do not illustrate a statistically discernible effect of the structural innovations to policy when the 95% confidence interval is the criteria for significance. The 68% confidence interval, on the other hand, suggests a protracted negative response to contractionary monetary policy, lasting about three years. Moreover, the overall decline exhibited by the point estimate is greater than in the pre-Volcker sample period. Because the average amplitude and variance of the estimated policy shocks are lower in the post-Volcker sample period than in the pre-Volcker one, the confidence intervals in the latter sample period are correspondingly larger. Without more careful analysis of the nature of these policy shocks, however, it is difficult to say why the variance changes over time. On the one hand, the change might reflect "better" exogenous policy, in that the Fed is introducing less exogenous instability into the economy. On the other hand, the changes in the distribution of the exogenous policy shocks may reflect a change in the structure of the broader economy that reduced control error by the Fed, thereby allowing for smaller on average exogenous policy shocks to achieve the same objectives that would have required larger exogenous shifts in the funds rate prior to 1979.[18] The mapping of the decline in the monetary policy shock variance into "good luck" or "good policy" is not so clear cut.

The estimated price responses are also interesting. In general, there is very limited evidence of any statistically discernible response of the price level to contractionary policy shocks. This finding is fairly

[16]Sims (1992) first proposed including commodity prices as an inflation indicator in monetary VARs to avoid the "price puzzle." Hanson (2004) demonstrates that commodity prices are not sufficient to eliminate the price puzzle, particularly in the pre-Volcker period, and that the puzzle is generally more pronounced with monthly than quarterly data. Interestingly, the above evidence of a post-Volcker price puzzle is slightly larger — at least in terms of the point estimate — than that reported by Hanson (2004), whose sample ends in 1998.

[17]These results, as well as all those reported below, are qualitatively similar if output is measured by the unemployment rate, and/or if prices are measured with the personal consumption expenditure deflator. Further, estimates based on the gap between the actual and detrended (via a quadratic-trend) unemployment rate, and the inflation rate (of either price series), return similar results as well. In the interest of brevity these results are not reported here; contact the author for details.

[18]Moreno (2004) reports evidence of a change in the behavior of price setters between the sub-samples analyzed here.

common for monthly VARs in which the policy instrument is the fed funds rate; recall that the timing assumptions used to identify the structural policy shocks are more plausible with higher-frequency data. The point estimate of the response in the pre-Volcker period is similar to that of the jointly estimated sample, with a positive response for nearly one and one-half years. In contrast, the point estimate is always positive for the post-Volcker period. Recall that these estimates include commodity prices, ostensibly to avoid the "price puzzle." One interpretation of these results is that they provide evidence for a "cost channel" of monetary policy, as in Barth and Ramey (2001) and Christiano et al. (2005) — and that the nature of the cost channel may have changed over time.

Finally, the funds rate takes on a noticeably different dynamic pattern in the sub-sample estimates relative to the joint sample. In particular, the earlier sample period provides some evidence of a reversal in monetary policy between 18 and 36 months following a contractionary policy shock. This reversal, while at best only minimally significant in a statistical sense, is consistent with an interpretation of monetary policy during this period as having a "stop-go" nature. The story, as revealed in the impulse responses, might be stated as follows: an exogenous tightening of monetary policy leads to a recession within six months of the policy action; the reduction in output lasts for at least 18 months. Prices, however, adjust much more sluggishly to the policy action and the shortfall in output, only beginning to decline after the funds rate increase has dissipated and the response of output has reached its nadir. The Fed, when faced with the adverse consequences of the recession, decides (endogenously) after about 18 months to push the funds rate below its initial pre-shock value for some time. This policy reversal helps pull output back towards its initial level and eventually stabilizes the reduction in the price level as well.

Such a tale is not inconsistent with many analyses of policy making in the pre-Volcker Fed: the canonical interpretation within the Taylor rules literature faults the Fed for reacting too weakly to incipient inflation and being too concerned with output.[19] What is striking about the sub-sample VAR evidence, however, is that a very similar pattern is apparent in the post-Volcker sample as well — indeed, both the magnitude and duration of this apparent reversal in policy is larger in the second part of the sample than the first. Notice that the response of non-borrowed reserves tells a consistent story: within 18 months of the initial contractionary shock, the point estimate for non-borrowed reserves turns positive as the funds rate response becomes negative.

[19] See, for example, Taylor (1999) and Clarida et al. (2000) for supporting arguments and evidence; Orphanides (2004) provides an intetesting counter-argument.

This latter finding is surprising in light of the Taylor rules literature, and not broadly reported elsewhere in the VAR literature. Arguably, the dynamic responses of output and prices differ more substantially between the two sub-samples plotted in figure 1 (as well as non-borrowed reserves) than do those for the Fed funds rate. This result raises the question as to whether (and which of) the estimated VAR equations exhibit statistically discernible differences over time, as well as questions regarding the source of any instability across sub-samples.

3.1 Quantifying The Effects of Regime Variation

The above statistical evidence reveals instability in several of the equations in the VAR, but does not convey the economic significance of such variation over time. Put succinctly: is there evidence of sizable differences in the effects of monetary policy over time, and if so, which parts of the empirical model appear to have contributed the most to any observed changes?

Since individual coefficient estimates in a VAR do not lend themselves to convenient interpretation, I propose a series of simulations designed to explore the contribution of each structural block of the VAR to the time variation in the dynamic responses to a monetary policy shock shown in figure 1. Recall that, due to the identifying assumptions on the activity sector, Y_t, one cannot meaningfully separate the individual innovations within that block, but all are predetermined with respect to the monetary policy instrument. To the extent that the identifying assumptions allow treating the equation for the fed funds rate as a monetary policy rule, and its residuals as the structural monetary policy shocks, I also investigate the effects of changes to the estimated policy rule. A parallel investigation of changes to the policy block as a whole returned similar results, and thus are not included here for the sake of brevity.

At an intuitive level, if reasonable variation in the group of coefficients that comprise one of the structural blocks of the VAR does not lead to noticeable differences in the implied impulse responses, then one could conclude that this block does not play a significant quantitative role in the determination of the dynamics. I use the results from the joint and sub-sample estimates above, so that "reasonable" variation is based on data rather than arbitrarily chosen values. This approach has the additional advantage of mirroring studies in the Taylor rule literature that focus on the contribution of changes in the estimated policy reaction function across sub-sample periods.[20]

At the outset it is important to recognize that the purpose of these simulations is to gauge the contri-

[20]See, for example, Judd and Rudebusch (1998).

bution of structural changes in different parts of the model: to provide a quantifiable measure of practical significance that complements the statistical tests of section 2.3 above. They are not intended to represent counterfactual policy experiments. Rather, I want to locate the sources of variation and determine the sensitivity of the above response functions to changes in the specification, in a way that is consistent with the sample data. Later in this section I compare the modest objectives of this investigation with related approaches in the literature.

The first two simulations investigate the sensitivity of the impulse responses by replacing the estimated structural reaction function from the joint sample with the one from the pre-Volcker (1966:01 – 1979:09) and post-Volcker (1983:01 – 2001:08) sub-samples. These results are shown in the second and third columns, respectively, of figure 2; the left-most column reproduces the joint-sample estimates from figure 1. The estimated standard confidence intervals from the joint-sample data also are replicated in the second and third columns of figure 2. While there are some differences across the point estimates, they are not nearly as pronounced as one might expect based upon the incongruent nature of estimated Taylor rules for these two sub-samples as reported in that literature. Arguably, the endogenous component of policy is not a significant source of the cross-sample variation as illustrated in figure 1.[21]

In a similar vein, figure 3 shows the effects of varying the specification of the activity block by replacing the structural estimates from the two sub-samples into the full sample. In this case, the common estimate of the reaction function over the joint sample is retained.[22] As in the experiments above, relatively little change is found in the dynamics of the funds rate itself. However, the dynamics of the remaining variables change quite noticeably. The most dramatic is output, which shows a larger initial decline for the 1966 – 1979 simulation, followed by an *increase* in the point estimate some 40 quarters after the policy shock. For the 1983 – 2001 simulation, the decline in output is much sharper and larger in magnitude, lying below even the 95% confidence interval on the estimated model. These simulated results also generate larger price puzzles than in the estimated model. In total, figure 3 provides evidence of a economically significant change in the underlying structure of the non-policy portion of the U.S. economy during the sample period in question. These results are not able to identify why the broader economy may have changed over time. However, it seems unlikely that the change is due to different monetary policy regimes, given that the results in figure 2 indicate minimal quantifiable consequences

[21] Nearly identical results are obtained with variation in the monetary block as a whole.

[22] As noted above, nearly identical results are achieved when the full monetary block is fixed and the sub-sample estimates of the activity block are used to generate the simulated IRFs.

of changing the reaction function while those in figure 3 are derived under the assumption of no shifts in the specification of the reaction function.

3.2 Additional Simulations

The previous results can be interpreted as suggesting that most of the cross-sample variation in the dynamics can be attributed to changes in the non-policy sector of the economy. That said, figure 3 also can be seen as showing the impact of imposing the same reaction function — one estimated for the joint sample period — on the two distinct sub-samples. Under this interpretation, the form of the reaction function can be seen to be a much more significant source of fluctuations. That is, once the broader economy is assumed to differ between the two sub-sample periods, moderate modifications to the monetary reaction function may generate substantially different dynamics.

This interpretation suggests an additional set of simulations, in which sub-sample variation of the broader economy is taken as the starting point — in contrast with much of the Taylor rules literature.[23] In this setting, replacing the estimated policy rule for each distinct sub-sample with the other constitutes an additional measure of the relative importance of variation in the two model components for the dynamic responses. Figure 4 examines this experiment for the 1966:01 – 1979:09 estimated model with the 1983:01 – 2001:08 policy rule; figure 5 shows the converse. In both of these figures, the left-hand column reports the actual responses to a 25 basis point contractionary policy shock for the sub-sample listed, as originally shown in figure 1. The right-hand column reports the impulse responses when the actual reaction function is replaced with the one estimated for the other sub-sample. The confidence intervals from the actual model again are replicated in the right-hand column.

Figure 4 reveals greater instability in the impulse responses of the simulations vis-à-vis the in-sample estimated responses. The funds rate response, in particular, appears much more volatile in the presence of the post-Volcker reaction function. While this difference is not as strong when the full 1983:01 – 2001:08 monetary block is employed (not shown for brevity), the activity variables nonetheless exhibit greater volatility in response to contractionary policy shocks of identical size. Figure 5 also illustrates some differences between the actual and simulated results for the 1983:01 –2001:08 sub-sample, but these are less extreme than those of figure 4: in general, the counterfactual responses lay inside the confidence intervals of the estimated in-sample responses. (The main exception appears to be the response

[23] These sub-sample VARs are estimated separately, which also allows for the variance estimates to vary across samples.

of the funds rate under the alternative reaction function about three years after the initial shock.)

The results in figure 4 are difficult to reconcile with the conventional wisdom in the Taylor rules literature that states that the post-Volcker reaction function is stabilizing whereas the pre-Volcker one was not — if the reaction function is viewed as the key to understanding the changing dynamics of the economy. While these simple simulations are far from definitive, they do suggest that the activity block plays the more important role in determining the dynamic responses to the monetary policy shock. Indeed, figure 5 shows that replacing the actual Volcker-Greenspan reaction function with the "sub-optimal" one from the Martin-Burns-Mitchell era does not appear to significantly worsen the nature of the responses given the activity block for the more recent 1983 – 2001 period. This apparent disparity between the VAR and the Taylor rules literature deserves further investigation.

3.3 Discussion

As noted in the introduction, VAR analysis has traditionally emphasized the role of the exogenous policy shocks at the expense of the estimated reaction function. Recent attempts to model the effects of alternative monetary policy regimes have generally taken two approaches. The first includes counterfactual simulations such as Bernanke et al. (1997) or Sims (e.g. Sims, 1998a,b; Sims and Zha, 1998), and more recently Hoover and Jordá (2001) and Leeper and Zha (2002). While my approach is similar in spirit (if not structure) to those papers, I do not claim to learn anything about the effects of alternative policy rules directly: any such assertion would have to address the Lucas Critique, as discussed below. Rather, I merely am interested in determining how sensitive the estimated dynamics are to reasonable (i.e. data-driven) variation in the blocks of the identified VAR.[24]

A variation on this approach uses Bayesian methods to model stochastic volatility in the coefficients and/or variances of a VAR model. There are some important differences in the specific assumptions — Cogley and Sargent (2001, 2005) and Canova and Gambetti (2005) use time-varying parameter models while Sims and Zha (2004) employ a regime-switching approach — and identifying assumptions — Cogley and Sargent (2001, 2005) use a small-dimension, recursive reduced-form VAR while Sims and Zha (2004) and Canova and Gambetti (2005) use structural VAR models — that are used in this literature, and consensus on these modeling choices has not yet arisen. Sims and Zha (2004) also note that the resulting

[24]Ahmed et al. (2002) and Boivin and Giannoni (2002) undertake similar investigations as I do here, albeit with differences in the data series and model specification.

posterior distributions tend to be diffuse. In much of this research, the primary focus is more on whether or not there is evidence of a direct shift (or drift) in policy, and less on the relative contribution of potential changes in monetary policy to understand sub-sample changes in the transmission mechanism, as I investigate here (although Canova and Gambetti, 2005, do report support for my main results in their framework).

A second approach attempts to imbed the question into a more structural framework, then examine the effects of varying one or more parameters. Recent examples include Boivin and Giannoni (2003), Favero and Rovelli (2003), and Stock and Watson (2002), amongst others. This approach clearly has strong appeal, but is not without some difficulties.[25] Often parameters related to the behavior of private agents are held constant while Fed preferences are varied, which presumes the answer to the issues raised in the introduction.[26] A particularly awkward example occurs when one varies the monetary policy regime while holding fixed the price setting parameter of a Calvo model — itself not a "deep structural" parameter, but one that likely would vary endogenously with changes in the mean level or persistence of inflation. More fundamentally, the solutions to many of these types of models are based upon a log-linearized approximation around a steady-state inflation rate of zero — whereas the data indicate the steady-state level of inflation is neither zero nor constant over the joint sample period studied here. While I do not wish to minimize the importance of a structural modeling approach for understanding the role of monetary policy, many of the commonly-used models have a flavor of "incredible restrictions" when confronting the data. Thus my motivation in a way is akin to that of the seminal work of Sims (1980): impose a minimal set of restrictions necessary for identification and then investigate the time series properties of the resulting estimates, particularly the impulse response functions.[27]

An objection to the above exercise is that the estimated policy rules are not truly structural. For the simple question of tracing how variation in parts of the estimated model equations impact the estimated dynamics implied by these equations, this criticism is not of primary importance. At the same time, the arguments above make it clear that the same objection can and should be applied to many of the results in the Taylor rules literature, as it does not represent a truly structural relationship either. If one does not adequately model policy rules as depending on the specification of the broader economy, any structural

[25] In addition to the theoretical caveats I outline, Heckman (2000) notes that "[t]he empirical track record of the structural approach is, at best, mixed. Economic data . . . have not yielded many stable structural parameters." (p. 49)

[26] Note that this approach must be taken in certain structural models in order to achieve identification for estimation.

[27] An additional advantage of this approach is that it is broadly consistent with multiple model structures and assumptions, given that macroeconomists still disagree about the "correct" model of the economy — and likely will for some time.

change in the behavior of private agents would have to be attributed to changes in the preferences of policy makers. Indeed, any model that relies upon a simple, fixed parametric specification of the private economy while varying the monetary policy regime may be subject to this general argument.

Moreover, the logic of the Lucas Critique should apply with equal force to models with policy makers who set policy in an optimizing framework: any change in the underlying determinants of private sector behavior should be incorporated into the rules that describe the actions of the policy makers. This bi-directional nature of structural instability changing the behavior of both policy makers and private actors is implicitly ignored in many approaches in the existing literature. The evidence above, while hardly definitive, does point towards the importance of incorporating the feedback from the private sector into a model of central bank behavior. Moreover, a view of changes in the preferences of policy makers as the primary explanation for the changes in the observed macroeconomic dynamics seems difficult to reconcile with the results shown above: to a first-order approximation, changes in the policy rule should have had much more noticeable effects on the simulated impulse responses. Perhaps the impact of such policy rule changes are reflected primarily in the behavior of the private sector, which then possibly could account for some of the importance of variation in the activity block noted above. But this explanation at best implies that one cannot identify the source of changes in the activity block — they might flow from changes in the behavior of policy makers, or they might originate elsewhere — and therefore nothing definitive can be said about the contribution of changes in policy to the overall stability of the economy.

On the other hand, the Lucas Critique only says that optimizing private agents should respond to changes in the policy environment — it does not say how large those responses will be in the data. It is possible that the optimal response to the empirical policy variation identified here is not especially large: recently Rudebusch (2003) has illustrated that autoregressive equations for the activity variables largely are insensitive to historically estimated variation in the policy rule. Alternatively, the empirical variation in the policy rule itself may not be so sizable — witness the similarities in the response of the Fed funds rate to a monetary policy shock between the two sub-samples shown in figure 1 — and that could justify minimal if any change in behavior by private agents, especially in the face of uncertainty about the true objectives of policy makers. In other words, the behavior of the private sector might be robust to certain types and magnitudes of changes in the policy rule.[28] Notice that in each of these cases

[28]One also could conceive of “modest” variation in the policy rule, which rational agents are unable to detect as distinct policy regimes, in analogy with the “modest interventions” (i.e. temporary deviations from a given policy rule) of Leeper and Zha (2003). Other costs of learning about and/or adjusting to policy changes could produce “optimal inertia” in the feedback of

the above simulations may admit a counterfactual policy interpretation, although the main conclusions of this paper do not depend in any way on such an interpretation.

4 Conclusion

The two main empirical literatures that attempt to quantify the effects of monetary policy on the U.S. economy — the Taylor rule approach and the structural VAR approach — have focused on distinct components of monetary policy and, in general, have reached rather different conclusions about the "importance" of monetary policy in the economy. While the Taylor rules literature places primary emphasis on changes in the policy rule specification leading to broader changes in macroeconomic dynamics (as in, e.g. Taylor, 1999, or Clarida et al., 2000), until recently many papers in the VAR literature did not even recognize a meaningful break in the time series processes of the variables in the model (see, e.g., the Handbook chapter of Christiano et al., 1999). This paper steers a path between these two extremes, using the multivariate framework of a structural VAR to explicitly recognize the possibility of changes in both the policy rule and the broader economy, while (ideally) minimizing the "incredible" restrictions imposed upon the data. It also straddles the small-dimension, reduced-form estimates of Cogley and Sargent (2001, 2005) and the larger, fully-identified regime-switching model of Sims and Zha (2004). Changes in the operating procedures of the Federal Reserve during the Volcker disinflation are accommodated by estimation over a joint sample of monthly data from 1966 to 2001 that explicitly excludes October 1979 through December 1982.

Simple statistical tests reveal much instability of the equations in the VAR between the two subsample periods, although the impulse response functions are fairly similar — including, somewhat surprisingly, evidence of "stop-go" monetary policy in both the 1966 – 1979 and 1983 – 2001 sub-samples. Simulations to ascertain the quantitative significance of changes in the coefficient values of the VAR point towards instability in the activity (i.e. non-policy) block of the model contributing more to instability in the estimated dynamics of the model as a whole than the estimated policy rule (or monetary block). These results do not depend upon the simulations having a counterfactual policy interpretation.

These results are novel in that much of the existing empirical literature — and related "structural" investigations of how changes in the monetary policy regime impact the economy — does not represent

policy into private sector behavior. Discrete changes in policy may translate into gradual changes in behavior that are difficult to identify in the data.

a framework in which the questions put forth in the introduction can be adequately addressed. In some instances, the underlying structure of the broader economy is often treated as fixed; in others the behavior of policy makers is assumed invariant to changes in the structure of the economy. Recent research in a Bayesian framework also had attempted to address these issues; see Sims and Zha (2004) and Canova and Gambetti (2005) for examples of similarly motivated research that uses very different estimation techniques and identifying assumptions, yet whose results broadly complement those reported herein. While not providing a definitive answer to these questions of causality and the sources of breaks, the results reported here do strongly suggest the need to investigate further the nature of non-policy related structural change in the broader economy, and how such changes have affected (or should affect) the practice of monetary policy.

The gap between Taylor rule models and structural VAR models is still wide. Reconciling the approaches — and ultimately the conclusions — of these two literatures would be an important step towards improving our understanding of the role of monetary policy in economic fluctuations. In the interim, this exploration of the potential sources of variability both policy and the private economy may suggest important avenues to pursue.

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Table 1: Comparison of Akaike and Schwarz Information Criteria

	8 lags		*12 lags*		*14 lags*	
Sample Period	AIC	SIC	AIC	SIC	AIC	SIC
1966:01 – 2001:08:						
Fixed Coefficients	−50.1418	−49.4203	−50.0227	−49.0634	−50.0575	−48.9781
Break after 1979:09	−51.2054	−50.0044	−51.3013	−49.6948	−51.4678	−49.6544
Break after 1982:12	−51.0622	−49.8511	−51.0227	−49.4019	−50.9992	−49.1694
1966:01 – 1979:09, 1983:01 – 2001:08:						
Fixed Coefficients	−51.4692	−50.6823	−51.3471	−50.2915	−51.2940	−50.1007
Separate Pre- and Post-Volcker periods	−51.9632	−50.6735	−51.9376	−50.2099	−51.9593	−50.0077

Notes: Akaike Information Criteria: $\text{AIC} = \log(|\Sigma|) + \frac{2}{T}\left(n(p+1) + \frac{n(n+1)}{2}\right)$, Schwarz Information Criteria: $\text{SIC} = \log(|\Sigma|) + \frac{\log(T)}{T} \cdot \left(n(p+1) + \frac{n(n+1)}{2}\right)$, where $|\Sigma|$ is the determinant of the reduced-form covariance matrix, n is the number of variables in the VAR, p is the lag length, and T is the sample size. Within a column, a more negative value of an information criteria suggests a better fit. Values for information criteria are not directly comparable across different lag lengths shown, as initial values for estimation are drawn from within the stated sample period.

Table 2: Tests for Instability of Reduced-Form Model

a. Wald Tests of Parameter Stability

Equation	*8 lags*	*12 lags*	*14 lags*
IP	0.1510	0.0133	0.0165
CPI	0.0055	0.0503	0.0034
PCOM	0.2126	0.0113	0.0236
FF	0.0004	0.0011	0.0001
NBR	0.2233	0.0343	0.0119
TR	0.0000	0.0000	0.0000

Notes: Part (*a*) reports p-values for the null hypothesis of equal coefficent vectors across the two sub-sample periods, 1966:01 – 1979:09 and 1983:01 – 2001:08. Test statistic has a χ^2 distribution.

b. Goldfeld–Quandt Tests of Homoskedasticity in Innovation Variance

Equation	*8 lags*	*12 lags*	*14 lags*
IP	0.0000	0.0006	0.0009
CPI	0.3652	0.2389	0.4134
PCOM	0.0153	0.0424	0.0137
FF	0.0000	0.0000	0.0000
NBR	0.0003	0.0034	0.0143
TR	0.8330	0.8249	0.7990

Notes: Part (*b*) reports p-values for the null hypothesis of equal residual variances across the two sub-sample periods, 1966:01 – 1979:09 and 1983:01 – 2001:08. Test statistic has an F distribution.

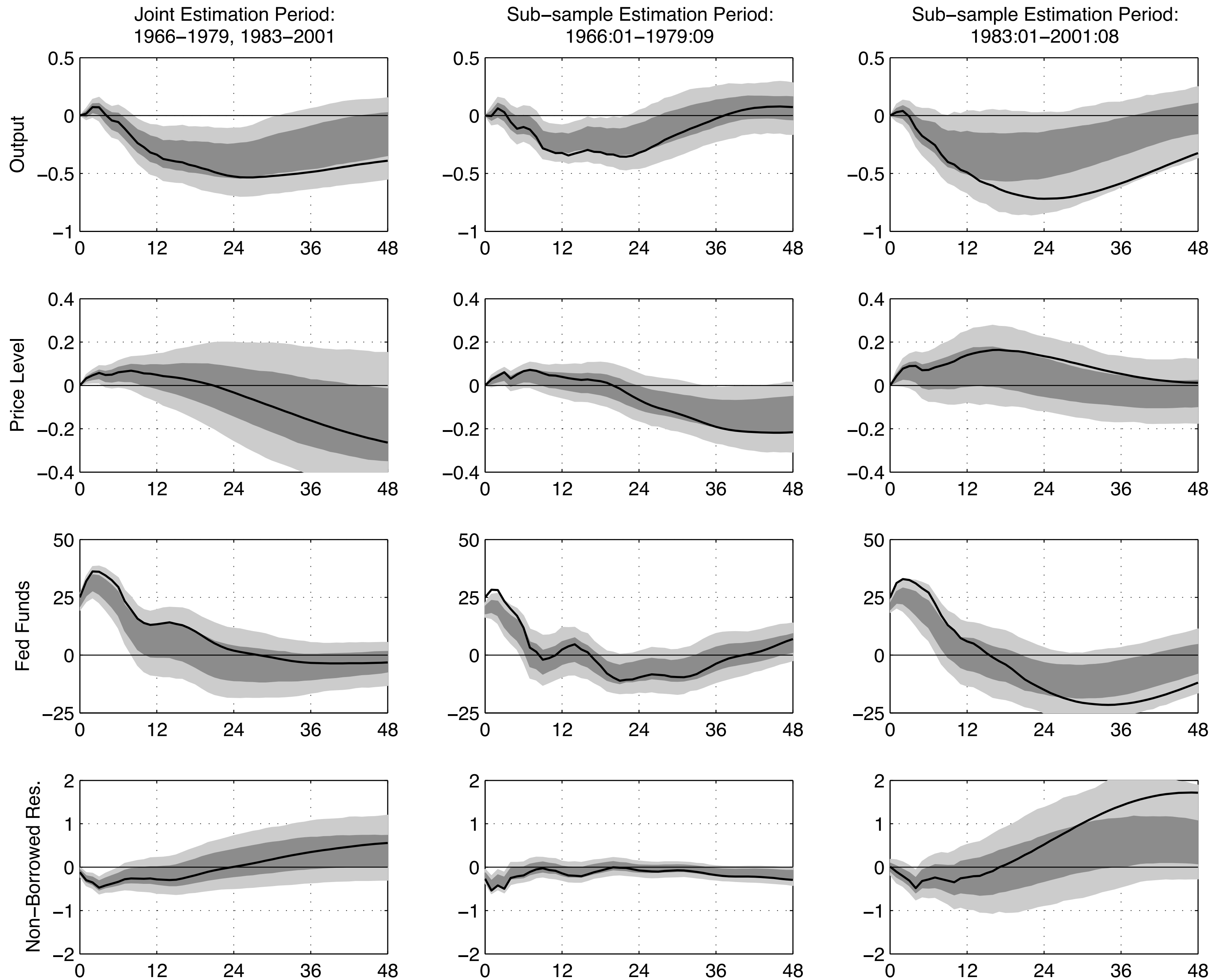

Figure 1: Comparison of Estimated Impulse Response Functions:
Responses to a 25-basis point contractionary monetary shock

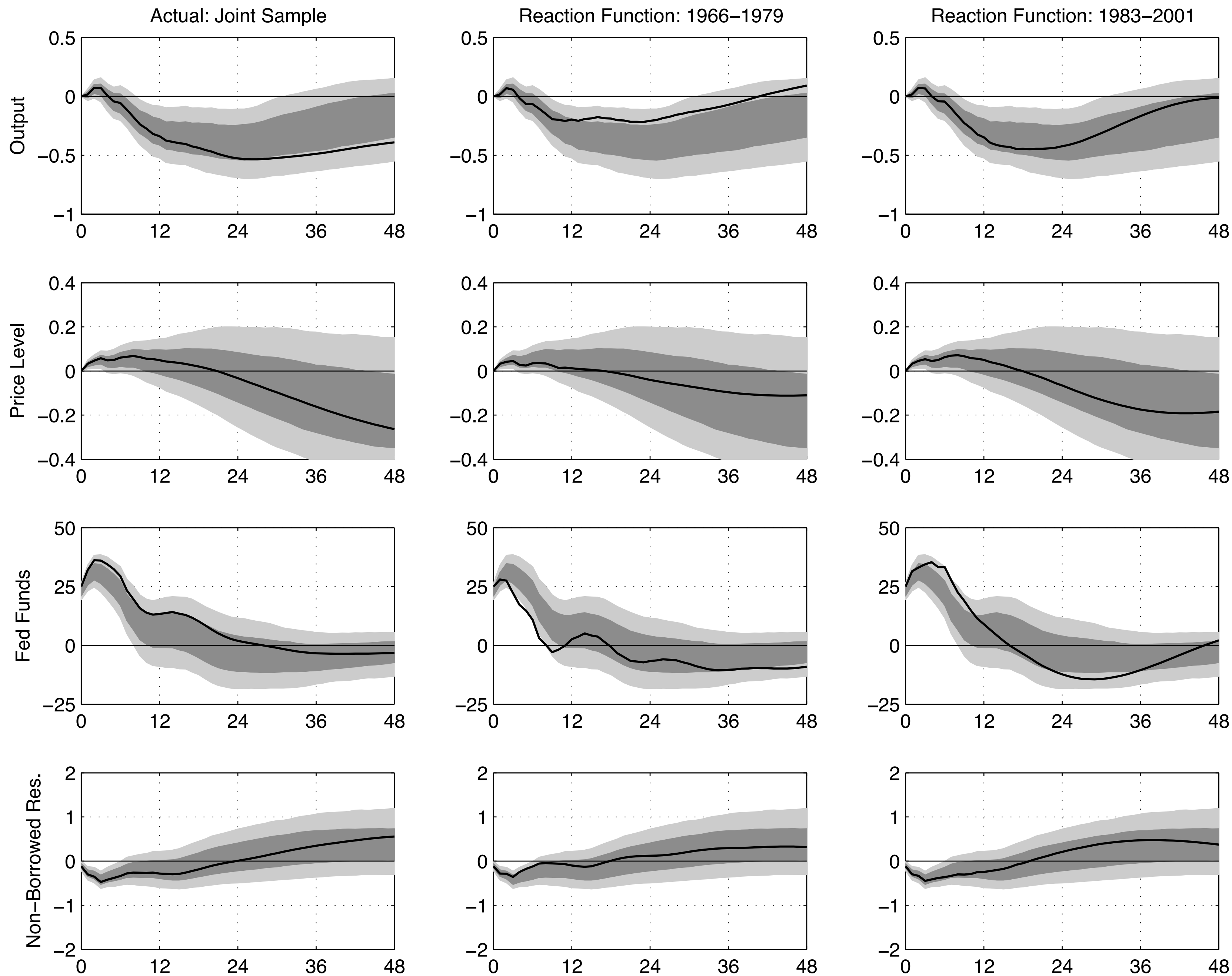

Figure 2: Impulse Response Functions with Varied Reaction Function: Responses to a 25-basis point contractionary monetary shock

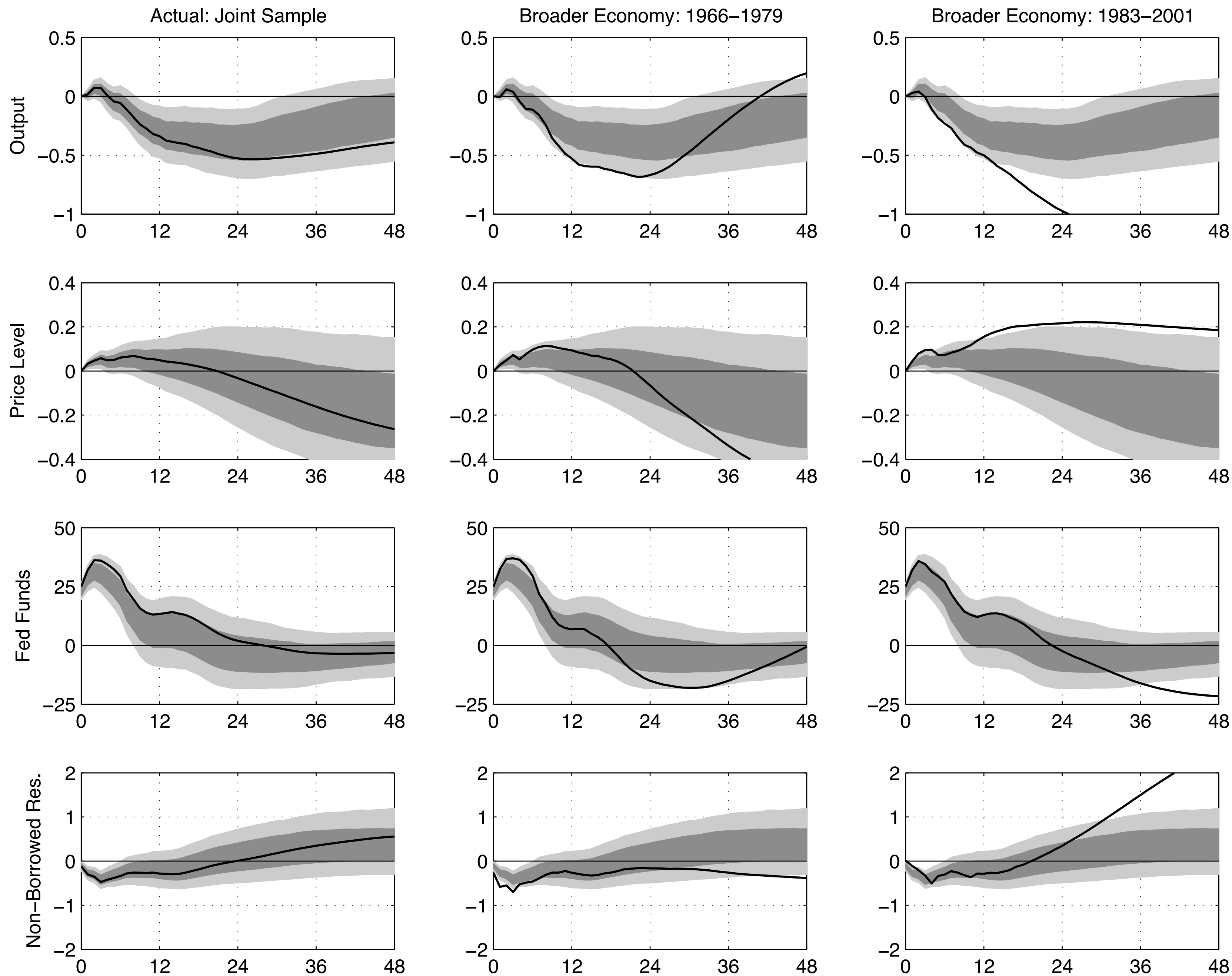

Figure 3: Impulse Response Functions with Fixed Reaction Function: Responses to a 25-basis point contractionary monetary shock

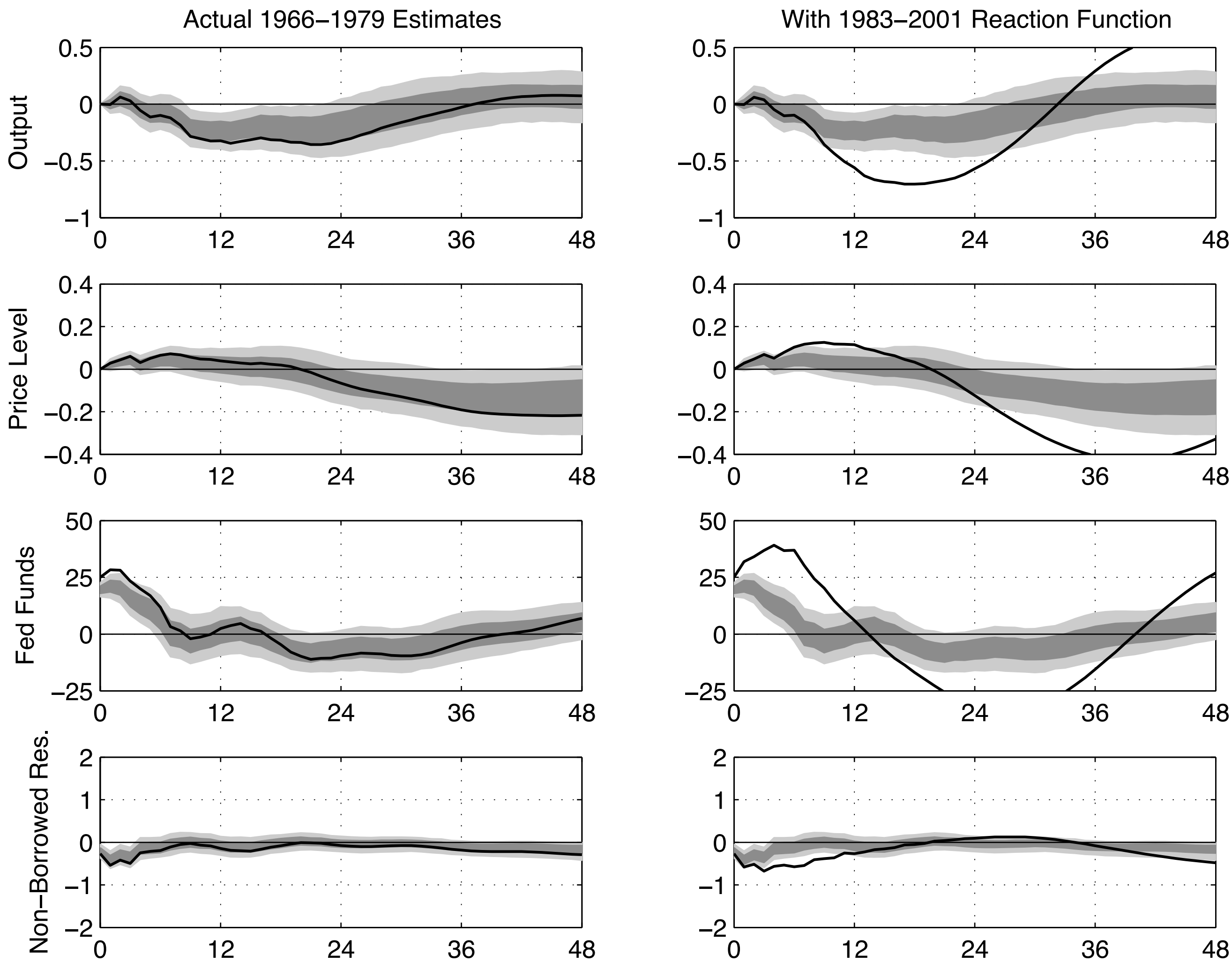

Figure 4: 1966-1979 Impulse Response Comparison:
Actual vs. simulated responses to a contractionary 25-basis point shock

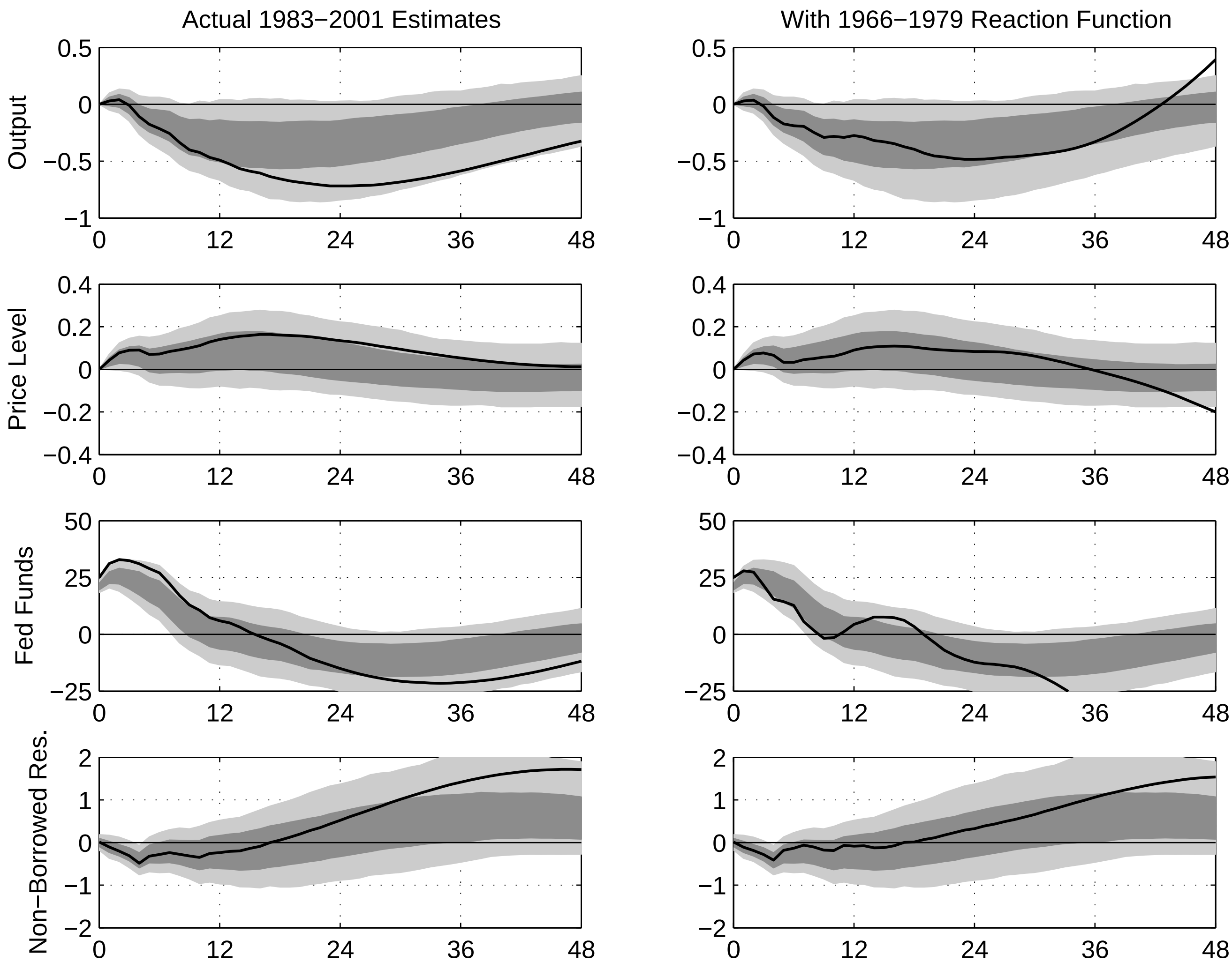

Figure 5: 1983-1997 Impulse Response Comparison:
Actual vs. simulated responses to a contractionary 25-basis point shock