WORKING PAPER SERIES

STLS/US-VECM 6.1: A Vector Error-Correction Forecasting Model of the US Economy

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Working Paper 1997-008A
http://research.stlouisfed.org/wp/1997/97-008.pdf

FEDERAL RESERVE BANK OF ST. LOUIS
Research Division
411 Locust Street
St. Louis, MO 63102

STLS/US-VECM6.1:
A Vector Error-Correction Forecasting Model of the U.S. Economy

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January 1997

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Abstract

Any research or policy analysis exercise in economics must be consistent with the time-series properties of observed macroeconomic data. This paper discusses in detail the specification of a six-variable vector error-correction forecasting model. We test for cointegration among those variables: the CPI, the implicit price deflator for GDP, real money balances (M1), the federal funds rate, the yield on long-term (10-year) government bonds, and real GDP. We also examine the estimated dynamic parameters of the vector error correction structure, and analyze the properties of the model residuals in detail; discuss the forecasting performance of the model with particular reference to the 1990-91 recession and the 1994-95 expansion; compare alternative permanent/transitory decompositions of the data series that are implied by the estimated parameters of the model; discuss the role of weak exogeneity in our estimated structure, and the identifying restrictions that are sufficient to determine a "historical policy rule" within the sample; discuss the conditions required for identification of "dynamic economic models" from the reduced form VECM structure and apply one set of exactly identifying restrictions to derive impulse response functions for a permanent nominal shock and a permanent real shock; and, report some ex-ante forecasts from recent history.

1. Why a Vector Error Correction Model?

Any exercise in empirical macroeconomics must recognize the conclusions drawn from time series analyses of macroeconomic data, and utilize specifications that are consistent with those results. Such analyses, starting with the classic study of Nelson and Plosser (1982), consistently have demonstrated that macroeconomic time series data likely include a component generated by permanent or nearly permanent shocks. Such data series are said to be integrated, difference stationary, or to contain unit roots.[1] On the other hand, economic theories suggest that some economic variables will not drift independently of each other forever, but ultimately the difference or ratio of such variables will revert to a mean or a time trend.[2]

Granger defined variables that are individually driven by permanent shocks (integrated), but for which there are weighted sums (linear combinations) that are mean reverting (driven only by transitory shocks), as cointegrated variables.[3] He then demonstrated in the Granger Representation Theorem (Engle and Granger, 1987; Johansen, 1991) that variables, individually driven by permanent shocks, are cointegrated if and only if there exists a Vector Error Correction representation of the data series. Let X_t be a (px1) vector of economic time series that each contain a permanent shock component. Then the Vector Error Correction Representation (VECM) of X_t is:

(1) $$\Delta X_t = \mu + \alpha\beta' X_{t-1} + \sum_{j=1}^{k} \Gamma_j \Delta X_{t-j} + \varepsilon_t$$

where Γ_j are (pxp) coefficient matrices (j = 1, ...k), μ is a (px1) vector of constants including any deterministic components in the system, and α, β are (pxr) matrices). $0 < r < p$, and r is the number of linear combinations of the elements of X_t that are affected only by transitory shocks. The error correction terms, $\beta' X_{t-1}$, are the mean reverting

[1] Statistical tests to differentiate series with unit roots (permanent shocks) from ones with "near unit roots" (extremely persistent transitory shocks) are known to have very low power to discriminate the two alternatives. Hence it is impossible to be certain of the existence of permanent shocks, and model specification becomes a choice problem on how best to minimize the dangers of specification error (Eichenbaum and Christiano (1990)).

[2] For early discussions of "Great Ratios" in macroeconomics, see Klein and Kosobud (1961) and Ando and Modigliani (1963)

weighted sums of cointegrating vectors and data dated t-1. The matrix α is the matrix of error correction coefficients. Note that the Vector Error Correction model (1) is just a standard VAR (Sims (1980)) in the first differences of X_t augmented by the "error correction" terms, $\alpha\beta' X_{t-1}$. In the absence of cointegration, the VECM is a VAR in first differences and the number of independent permanent shocks is equal to the number of variables (p). An important characteristic of cointegrated systems is that the nonstationary behavior of a p-dimensional vector of time series data may be explained by only k=p-r independent permanent innovations.

Cointegration in a vector time series has a number of implications for work in empirical macroeconomics. One of the purported advantages of recognizing cointegration rank in an integrated vector process is that it will result in improved forecast performance. Engle and Yoo (1987) illustrate that forecasts taken from cointegrated systems are "tied together" because the cointegrating relations must "hold exactly in the long-run." They demonstrate in a series of Monte Carlo experiments that incorporating cointegration into the forecasting model can reduce mean squared forecast errors by up to 40% at medium to long forecast horizons. As recently as Stock (1995), the apparent value of incorporating cointegration into the forecasting exercise was noted. In a recent applications Clements and Hendry (1995) and Hoffman and Rasche (1996b) re-examine this issue, concluding that it is difficult to verify the predictions of Engle and Yoo in practice. Clements and Hendry (1995) find that incorporating knowledge of cointegration rank results in significant mean squared forecast error (MSFE) reduction only in models formed from relatively small samples. Both papers illustrate that the relative gains (or losses) can depend upon the particular representation of data (e.g. levels, differences, linear combinations, etc.) about which one is concerned. The value of incorporating cointegration into the forecasting exercise is examined further by Christofferson and Diebold (1996). They conclude that the value of incorporating cointegration rank may actually appear only at short-run horizons when the conventional trace MSE forecast error criterion is used. No long-run advantage is obtained because the "long-horizon forecast of

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[3] Nonstationary variables can be integrated and cointegrated of different degrees. Our focus here is on variables that are integrated of order one (I(1)).

the error-correction term is always zero". Alternative measures of forecast performance that recognize the potential value of cointegration are suggested.

Our exercise in forecasting is designed to reveal how a simple autoregressive structure anchored by cointegration in the form of standard "textbook" long-run relations performs over several recent out-sample periods. We examine representations of data in both "levels" and "differenced" forms and measure performance of out-sample periods of varying size. The system is comprised of six relevant macro aggregates. We choose a set of cointegrating vectors that have been well documented in previous applied work, test for evidence of cointegration in the context of a relevant postwar sample and examine forecast performance using forecasts calculated dynamically from the corresponding VECM representation of the system. Performance is measured using standard MSFE criterion over the recent recession years 1990 and 1991 and the expansion years 1994 and 1995 in an effort to compare performance under contrasting economic conditions. Our system is comprised of the following: two measures of inflation; real money balances; the federal funds rate; the long term (10 year) government bond rate; and real GDP.

In evaluating forecasts, we concentrate on the rate of inflation, growth in real GDP and the level of the federal funds rate. One benchmark for the performance of our model is the performance of the projections published in the Federal Reserve Board's Green Book. These projections are available to us only with a five year lag. A second benchmark is the published central tendency forecasts of the FOMC members, as published with the Humphrey-Hawkins testimony.

In the following section, we discuss the specification of a particular six variable model and test for cointegrating vectors among those variables. In section 3 we examine the estimated dynamic parameters of the Vector Error Correction Structure and analyze the properties of the model residuals in detail. In section 4 we discuss the forecasting properties and performance of the estimated model with particular reference to the 90-1 recession and the 94-5 expansion. In section 5, alternative permanent/transitory decompositions of the data series that are implied by the estimated parameters of the model are constructed and discussed. In Section 6 we discuss the role of weak exogeneity in our estimated structure and identifying restrictions that are sufficient to determine a "historical policy rule" within the sample. In Section 7 we discuss the conditions required for identification of "dynamic economic models" from the reduced form VECM structure and apply one set of exactly identifying restrictions to derive impulse response functions for a permanent nominal shock and a permanent real shock. In Section 8 we report some ex-ante forecasts from recent history. Finally, in Section 9 we point out a number of directions that we think are interesting in which this research can be extended.

2. Testing for Cointegration

In recent years, tests of cointegration have revealed that various linear combinations of individually integrated processes—such as real money balances, real income, inflation, and nominal interest rate series—are in fact linked by stationary linear combinations. Hoffman and Rasche (1991), Johansen and Juselius (1990), Baba, Hendry, and Starr (1992), Stock and Watson (1993), and Lucas (1994) among others present evidence on the stationarity of money demand relations. This is sometimes estimated as a velocity relation that links income velocity to movements in a measure of nominal interest rates as in Hoffman, Rasche and Tieslau (1995). Mishkin (1992) and Crowder and Hoffman (1996) present evidence of a Fisher equation, while Campbell and Shiller (1987, 1988) have examined cointegration between interest rates of assets with different terms to maturity. The final relation we examine is the relation between two definitions of inflation, the GDP deflator and the inflation rate for the CPI. Evidence in support of most of these cointegration relations is compiled in experiments presented in Hoffman and Rasche (1996a).

We examine a six-dimension vector process that allows us to test whether there is evidence that distinct money demand, Fisher, term structure, and inflation rate relations exist in the data. The variables used in the analysis include a measure of real M1 money balances (m1p), two measures of inflation (infgdp) and (infcpi), the long term rate on government securities, (lrate) and the federal funds rate (funds). The primary data used in this paper span 56:1 to 96:1 and are taken from the Federal Reserve Bank of St. Louis database, FRED. A consistent series for M1 over the full sample is obtained from Rasche (1987).[4]

Real money balances are obtained by deflating the nominal series by the GDP deflator. Inflation is measured as the percent changes (log differences) in the GDP deflator and CPI respectively at an annual rate and both real balances and real GDP are expressed as natural logarithms. The degree of integration maintained by these series has been widely discussed throughout the literature. We are operating under the assumption that each series either maintains a single unit root or is well approximated by the assumption that it follows an I(1) process.[5]

To illustrate the cointegration space that spans the steady-state relations which presumably underlie our VAR representation, order the variables as $z_t' = \{\mathrm{m1p}_t, \mathrm{Infdef}_t, \mathrm{lrate}_t, \mathrm{Infcpi}_t, \mathrm{gdp}_t, \mathrm{funds}\}$ to correspond with the natural log of real balances, deflator inflation, long-term government rates, CPI inflation, the natural log of real GDP and the Federal funds rate. A strict interpretation of the cointegration space that reflects all four hypothesized long-run relations is:

$$\beta' = \begin{bmatrix} 1 & \underline{\underline{0}} & \underline{\underline{0}} & \underline{\underline{0}} & -1 & \beta_1 \\ \underline{\underline{0}} & 1 & \underline{\underline{0}} & \underline{\underline{0}} & 0 & -1 \\ \underline{\underline{0}} & \underline{\underline{0}} & 1 & \underline{\underline{0}} & 0 & -1 \\ \underline{\underline{0}} & \underline{\underline{0}} & \underline{\underline{0}} & 1 & 0 & -1 \end{bmatrix}$$

In this representation the first row of β' captures the money demand equation, the second and fourth rows measure the Fisher relations using the two distinct measures of inflation in our system, and the third row captures the term structure spread linking the Federal funds rate and the long-term government rate.

Several alternative normalizations of β yield long-run relations that embody our system of cointegrating relations. For example, an observationally equivalent representation of β contains only a single Fisher relation and a separate relation linking

[4] M1as measured here is augmented by estimates of sweeps into Money Market Deposit Accounts starting in 1995.

[5] The assumption that inflation is an I(1) process is equivalent to assuming that the log of the price level is an I(2) process.

the two inflation measures. Define $R_1 = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 1 & 0 & -1 \end{bmatrix}$. Then $R_1\beta'$ is the matrix of cointegrating vectors reflecting a stationary spread between the inflation rates and a single Fisher equation between the deflator inflation rate and the funds rate. Alternatively, define $R_2 = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & -1 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}$. $R_2\beta'$ is the matrix of cointegrating vectors reflecting a stationary spread between the inflation rates and a single Fisher equation between the CPI inflation rate and the funds rate. Similarly interest elasticity and Fisher relations could be expressed in terms of the long-term rather than the short-term rate. Let $R_3 = \begin{bmatrix} 1 & 0 & b_1 & 0 \\ 0 & 1 & -1 & 1 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & -1 & 1 \end{bmatrix}$. Then $R_3\beta'$ is the matrix of cointegrating vectors reflecting the two Fisher equations and the money demand vector in terms of the long-term interest rate. However, the class of steady-state relations is sufficient to satisfy conditions for identification of cointegration spaces discussed in Hoffman and Rasche (1996)

In our system of r=4 cointegrating relations, r-1 (or three) restrictions satisfy the conditions for identification for each of the cointegrating relations. These restrictions appear with double underscore ($\underline{\underline{0}}$) in the representation for β' depicted above. The "ones" down the main diagonal of the expression represent normalizations, but values in all remaining entries may in principal be tested in empirical analysis.

The Case for Cointegration

One approach to estimating the parameters of the Vector Error Correction model that we have specified is to apply an appropriate estimation technique and to test for the number of cointegration vectors present. In the present situation an alternative approach is available since three of the cointegrating vectors have "known" coefficients, i.e. the presumed stationary weighted sums are prespecified by economic theories. We can take advantage of this information and apply tests that have more power to discriminate between stationarity and nonstationarity than a general testing procedure.

The first step in our approach is to verify that the two real interest rates and the term structure spread are indeed stationary as suggested by economic theory. To do so we construct standard unit root tests on these variables. The results of a Dickey-Fuller regression on the difference between Infdef and the Federal funds rate are shown in Table 1. There is no evidence here to reject nonstationarity of the short-term real interest rate, since the "t-ratio" on $CIV2_{-1}$ is only -2.65. However this is not surprising, since a

Table 1
The Fisher Relation using Infdef
(Dickey-Fuller Regression)
Dependent Variable is $\Delta CIV2_t$; Sample 1956:2 1996:1

	coefficient	standard error	*t*-statistic
C	-0.3720	0.1727	-2.15
$CIV2_{-1}$	-0.1477	0.0557	-2.65
$\Delta CIV2_{-1}$	-0.2821	0.0880	-3.20
$\Delta CIV2_{-2}$	-0.1777	0.0880	-2.02
$\Delta CIV2_{-3}$	-0.1344	0.0853	-1.57
$\Delta CIV2_{-4}$	0.0162	0.0796	0.20
Adjusted R-squared 0.16			
S.E. of regression 1.72			

number of studies have found evidence that the real interest rate shifted sometime after the beginning of the New Operating Procedures (e.g. Huizinga and Mishkin (1986)). Therefore we augment the standard Dickey-Fuller regression with a dummy variable D79, whose value is 0.0 through 79:3 and 1.0 thereafter. The results of estimating this augmented Dickey-Fuller regression are shown in Table 2. Now the results appear consistent with the theoretical presumption that the short-term real interest rate is stationary since the "t-ratio" on $CIV2_{-1}$ is now reduced to -3.75[6] The coefficient on D79 in this regression is negative which is consistent with the conclusion reached by Huizinga and Mishkin that the there was an increase in the real interest rate around the time of the New Operating Procedures. The point estimate of the coefficient on $CIV2_{-1}$ in Table 2, -.29, is substantially different from zero; the size of the "t-ratio" principally reflects the imprecision of the estimates, not a process that is extremely persistent.[7]

Our next test is to examine the stationarity of the term structure spread. Research by Campbell and Shiller (1987,1988) suggests that this spread is stationary, consistent with the implications of a rational expectations hypothesis of the term structure of interest rates. The results shown in Table 3 for CIV3 are consistent with this conclusion, since the "t-ratio" on $CIV3_1$ is -4.00. Hence we have not pursued any further tests on this variable.

Table 2
The Fisher Relation using Infdef
(Dickey-Fuller Regression with Dummy)
Dependent Variable is $\Delta CIV2_t$
1956:2 1996:1

	coefficient	standard error	*t*-statistic
C	-0.3011	0.1718	-1.75
$CIV2_{-1}$	-0.2927	0.0781	-3.75
$\Delta CIV2_{-1}$	-0.1795	0.0950	-1.89
$\Delta CIV2_{-2}$	-0.0991	0.0916	-1.08
$\Delta CIV2_{-3}$	-0.0778	0.0866	-0.90
$\Delta CIV2_{-4}$	0.0520	0.0793	0.66
D79	-0.9240	0.3557	-2.60
Adjusted R-squared 0.19			
S.E. of regression 1.72			

[6] It should be noted that the nonstandard distribution of estimated coefficient of $CIV2_{-1}$ depends upon the form of the deterministic portion of the specified equation, hence the critical values for the "t-ratio" of this coefficient change from the standard Dickey-Fuller tables when the dummy variable is introduced into the estimating equation.

[7] The normalized bias test [T*(ρ-1)] is -46.40 which provides strong evidence against a unit root in the series.

Table 3

Dickey-Fuller Regression
Dependent Variable is $\Delta CIV3_t$
The Term Structure Spread
1956:2 1996:1

	coefficient	standard error	*t*-statistic
C	0.1487	0.0749	1.98
$CIV3_{-1}$	-0.1870	0.0468	-4.00
$\Delta CIV3_{-1}$	0.1991	0.0802	2.48
$\Delta CIV3_{-2}$	0.0255	0.0801	0.31
$\Delta CIV3_{-3}$	0.0627	0.0801	0.78
$\Delta CIV3_{-4}$	0.0927	0.0807	1.15
Adjusted R-squared .08			
S.E. of regression .82			

Third, we have estimated a Dickey-Fuller regression for the difference between the CPI inflation rate (Infcpi) and the Federal funds rate. As in the Fisher relation formed from Infdef, there is no evidence to reject a unit root in this difference from the standard unit root tests. However, the results reported in Table 4 for the specification that includes the dummy variable D79 indicate a "t-ratio" for the estimated coefficient on $CIV4_{-1}$ of -3.68 which again is consistent with a stationary short-term real rate. The estimated coefficient on the dummy variable here is also negative, consistent with a possible increase in real rates in the early 1980s.[8]

[8] As noted above, an alternative specification of the structure of our Vector Error Correction model is a single Fisher effect and a stationary spread between GDP deflator inflation and CPI inflation. In this case we would not expect that the difference in the two measures of inflation would shift in the early 1980s. A standard Dickey-Fuller regression for this difference without the dummy variable produced a "t-ratio" for the estimated coefficient of the lagged levels variable of -5.34

Table 4

Dickey-Fuller Regression with Dummy
Dependent Variable is $\Delta CIV4_t$
The Fisher Relation using InfCPI
1956:2 1996:1

	coefficient	standard error	*t*-statistic
C	-0.1518	0.1636	-0.93
$CIV4_{-1}$	-0.2458	0.0667	-3.68
$\Delta CIV4_{-1}$	-0.2057	0.0892	-2.30
$\Delta CIV4_{-2}$	-0.2648	0.0907	-2.92
$\Delta CIV4_{-3}$	0.1092	0.0856	1.27
$\Delta CIV4_{-4}$	0.0351	0.0793	0.40
D79	-0.8186	0.2946	-2.78
Adjusted R-squared 0.26			
S.E. of regression 1.48			

Johansen and Juselius (1992) have designed an extension to the standard Johansen FIML estimator to deal with the case of mixtures of cointegrating vectors with "known" coefficients and vectors with "unknown" coefficients. We have used this approach to construct an estimate of the interest "semielasticity" of M1 velocity, conditional upon the three cointegrating vectors discussed above. We adopt the unitary income elasticity assumption that was shown to be consistent with the data in Hoffman *et. al*. (1995) and in Hoffman and Rasche (1996). At the same time we recognize that there are a number of both theoretically and empirically plausible values for the income elasticity. Cursory investigation on our part reveals that the basic conclusions of this paper are not significantly influenced by this assumption. The estimated interest semielasticity based on data through 96:1 is .085 as shown in Table 5.[9] The estimates are constructed on

[9] Lucas (1995) plots M1 velocity equations against almost a century of U.S. data assuming semielasticities in the range of 5-9 (.05 - .09 in terms of our scaling of interest rate) in his Figure 2.

quarterly data from 56:2 - 96:1, omitting 79:4 - 81:4. The latter period is omitted since the semilog functional form used in this model does not have adequate curvature to

Table 5

Johansen Estimate of Four Constrained
Cointegrating Vectors

Sample Period 56:2 - 96:1 Excluding 79:4 - 81:4,
Including Dummy Variable D82

m1p	infdef	lrate	infcpi	gdp	funds
1.0	0.0	0.0	0.0	-1.0	-.0805 (.0068)
0.0	1.0	0.0	0.0	0.0	-1.0
0.0	0.0	1.0	0.0	0.0	-1.0
0.0	0.0	0.0	1.0	0.0	-1.0

approximate M1 velocity in a range of nominal interest rates above about 10 percent (see Hoffman and Rasche (1996a)).[10] The Vector Error Correction model allows for three lagged changes in each of the six variables, (a 4th order autoregressive specification in the levels of the data) and includes the D82 dummy variable to allow for shifts in the deterministic drift of the nonstationary processes and for shifts in the equilibrium real interest rate after the New Operating Procedures period.[11]

Having constructed the FIML estimates of the four cointegrating vectors with constraints applied to the coefficients of the three "known" vectors, we can apply the tests

[10] A double log specification (log of M1 velocity and log of nominal interest rates) and an estimated interest elasticity of around 0.5 exhibits the necessary curvature to account for M1 velocity over the full range of observed nominal interest rates in the post war period(see Hoffman and Rasche (1996)). Lucas (1994) plots a double log velocity function with an elasticity between 0.3 and 0.7 against almost a century of U.S. data with a high correlation. Unfortunately the log transformation of interest rates does not allow for a linear model incorporating the rest of our cointegrating vectors. The model used here is best viewed as an approximation to the true M1 velocity function that is appropriate for nominal rates less than 10 percent.

[11] Hoffman and Rasche (1996a) show that there is no significant shift in the constant of the velocity cointegrating vector before and after the New Operating Procedures period. The same conclusion applies to the estimates constructed here, though we have not attempted to constrain the intercept of this cointegrating vector to the same value in the two subperiods.

for cointegration rank developed by Horvath and Watson (1995).[12] These results are reported in Table 6. We construct three tests. The first is for cointegration rank four

Table 6

Horvath-Watson Tests for Cointegration
Sample period 56:1 - 96:1 Excluding 79:4 - 81:4

Three “known” Cointegrating Vectors and One ‘unknown” Cointegrating Vector	149.88
One ‘unknown” Cointegrating Vector on Margin of Three “known” Vectors	27.80
Three “known” Cointegrating Vectors on Margin of One “unknown” Vector	138.34

based on three “known” cointegrating vectors and one “unknown” cointegrating vector. The critical values of this test statistic at the 10 percent, 5 percent and 1 percent levels are 69.40. Hence we fail to reject cointegration rank = 4.

The remaining tests reveal the strength of the evidence regarding the various cointegration vectors in the system. The second test, the test for one “unknown” cointegrating vector on the margin of three “known’ cointegrating vectors. This test provides sharp inference about the existence of a money demand vector in a multivariate system characterized by three known vectors. The critical values for this test statistic at the 10 percent, 5 percent and 1 percent levels are 25.57. Finally we test for three “known” cointegrating vectors on the margin of one “unknown” cointegrating vector. This evidence bolsters the case for the three vectors already revealed in the univariate analysis. We regard this as important, since we consistently have found evidence for a velocity vector in lower dimensional systems (Hoffman and Rasche, 1991; Hoffman, Rasche and Tieslau, 1995; Hoffman and Rasche, 1996a). The critical values for this test

[12] The tables of critical values for these test statistics are extended in Hoffman and Zhou (1997).

statistic at the 10 percent, five percent, and one percent levels are 38.75. Hence on both margins we find evidence of additional cointegration up to rank = 4.

We have examined the robustness of the estimated interest rate semielasticity by reestimating the model recursively starting with samples that end in 85:4. In each regression we have added an additional four observations until the longest sample period ends in 95:4. The results of these estimations are reported in Table 7.

Table 7

Recursive Estimates of Semielasticity of Money Demand
All Samples Exclude 79:4 - 81:4 and
Include Dummy Variable D82

Sample Ending in:	Estimated Semielasticity and Standard Error
85:4	.1103 (.0046)
86:4	.1062 (.0049)
87:4	.1051 (.0047)
88:4	.0893 (.0061)
89:4	.0810 (.0070)
90:4	.0791 (.0070)
91:4	.0778 (.0077)
92:4	.0845 (.0072)
93:4	.0843 (.0068)
94:4	.0824 (.0067)
95:4	.0775 (.0070)

The estimates of the interest rate semielasticity range from .0822 to .1126. There is some tendency for the estimated coefficient to drift lower as the sample length is increased through the mid 80s, but since the end of 1989 the estimates are remarkably stable. Even the introduction of a sizable (estimated) amount of Sweeps in 1995 did not materially affect the estimated coefficient.

3. Model Dynamics and Residuals Analysis

While the cointegrating vectors determine the steady-state behavior of the variables in the Vector Error Correction Model, the dynamic responses of each of the variables to the underlying permanent and transitory shocks are completely determined by the sample data without any restriction. Table 8 indicates the estimated coefficients of the VAR portion of the model. Each column of the table indicates the estimated coefficients for the equation with the dependent variable indicated at the top of the column. The row variables in the left hand column indicate the various regressors. Only a few of the estimated coefficients are significantly different from zero as is apparent from the F-statistics indicated in Table 9. These tests are the conventional computations in VAR analyses that test the maintained hypothesis that all of the coefficients in a particular distributed lag are equal to zero. Again the dependent variable is indicated at the top of each column and the variable in the left hand column indicates the distributed lag that is subject to the exclusion test.

First, note that only the estimated distributed lag coefficients on changes in the funds rate are significantly different from zero in the equation for the growth rate of real GDP. Second, other than its own autoregressive structure, the only significant feedback onto the Federal funds rate comes from lagged growth of real GDP. There are no significant distributed lags in the equation for the long-term interest rate, however as will be seen below, this does not imply that the long rate is modeled as a random walk.

Table 8
Estimated VAR Coefficients in VECM Model
Sample Period 56:2 - 96:1 Excluding 79:4 - 81:4
Including Dummy Variable D82

	dependent variable (equation)					
Variable	Δm1p	Δinfdef	Δfunds	Δinfcpi	Δgdp	Δlrate
C	-.0240	-.9541	.3179	-3.183	-.0196	.0739
D82	-.0083	-.8319	.2993	-.5294	-.0136	.2760
$\Delta m1p_{-1}$	.2558	70.36	-.1583	71.90	-.2315	2.755
$\Delta m1p_{-2}$	.0977	-23.62	1.614	-4.809	.2370	4.399
$\Delta m1p_{-3}$	.1579	-12.83	8.242	-43.91	-.0593	4.524
$\Delta infdef_{-1}$	.0004	-.7905	-.0291	.1453	.0007	.0555
$\Delta infdef_{-1}$	.00003	-.7671	-.0146	.3383	.0007	.0831
$\Delta infdef_{-3}$	.00001	-.8074	.0135	.2472	.0002	.1036
$\Delta funds_{-1}$	-.0038	.4902	.1353	.7563	-.0008	-.0193
$\Delta funds_{-2}$	.0002	.4418	-.4165	.8021	-.0039	-.0883
$\Delta funds_{-3}$	-.0001	-.1229	.1766	.2829	-.0010	.1152
$\Delta infcpi_{-1}$	-.0010	.2303	.0821	-.6035	-.0004	-.0053
$\Delta infcpi_{-2}$	-.0017	.4176	.1349	-.6720	-.0010	-.0089
$\Delta infcpi_{-3}$	-.0017	.5198	.1263	-.4271	-.0020	-.0019
Δgdp_{-1}	-.0321	-6.629	22.94	3.399	.1222	9.850
Δgdp_{-2}	.0116	9.057	19.79	-17.50	-.0348	3.369
Δgdp_{-3}	-.1079	8.153	5.776	2.378	-.0914	6.528
$\Delta lrate_{-1}$	-.0045	.1925	.3522	.4685	.0023	.1216
$\Delta lrate_{-2}$	-.0007	-.2288	-.1935	-.2694	.0010	-.0525
$\Delta lrate_{-3}$	-.0005	-.1476	-.0497	-.3516	.0040	-.1846

(continued)

$\bar{R}^2$	.70	.47	.42	.46	.25	.19
s.e.	.0059	1.10	.62	1.19	.0082	.45
D-W	2.04	1.97	1.72	2.08	2.01	2.04

Table 9

F tests for Exclusion of Lagged Variables
(Equations are shown in columns)

	dependent variable (equation)					
	Δm1p	Δinfdef	Δfunds	Δinfcpi	Δgdp	Δlrate
Δm1p	11.27*	5.26*	.54	5.71*	1.36	1.29
Δinfdef	.26	24.93*	.23	2.40	.45	1.58
Δfunds	6.74*	5.63*	6.98*	14.06*	3.58*	2.01
Δinfcpi	2.86*	1.61	1.75	19.94*	1.44	.04
Δgdp	.95	.40	6.94*	.62	.99	2.41
Δlrate	3.02*	.47	2.16	1.48	1.78	1.65

The structure of the lagged change effects in the GDP deflator inflation and the CPI inflation equations are quite similar. In both equations the estimated coefficients on lagged changes in the dependent variable are significant, but the estimated coefficients on lagged changes in the alternative inflation rate are not. Also lagged changes in real balances and the Federal funds rate enter significantly into both inflation rate equations. The distributed lag structures of these equations also differ strikingly from VAR type estimates of "Phillips Curves" (see for example Fuhrer (1995)). Such equations are specified as distributed lags in changes in an inflation rate (either the GDP deflator or the CPI) and lags on the unemployment rate. In both equations estimated here, the estimated coefficients on the lagged changes in real output are not significant. In addition, as will be discussed below, the level of real output does not enter these equations through the error correction terms.

Finally, changes in real balances are significantly affected by lagged changes in real balances, lagged changes in the CPI inflation rate, and lagged changes in both interest rates, but not by lagged changes in real output. The only equation where lagged changes in real output enter significantly is in the Federal funds rate equation.

The matrix of estimated error correction coefficients is shown in Table 10. In this

Table 10

Estimated Matrix of Error Correction (α) Coefficients
Sample Period 56:2 - 96:1 Excluding 79:4 - 81:4
Including Dummy Variable D82
(t-ratios in parentheses)
(Equations are shown in rows)

	error correction terms			
	$mdciv_{-1}$	$defrrciv_{-1}$	$termciv_{-1}$	$cpirrciv_{-1}$
Δm1p	-.0218 (-3.29)	-.0003 (-.46)	.-.0006 (.-1.31)	-.0014 (2.07)
Δinfdef	-.8099 (-.65)	-.5870 (-4.47)	.1753 (1.93)	.4875 (3.68)
Δfunds	.5196 (.75)	.0419 (.57)	-.0708 (-1.39)	.0709 (.95)
Δinfcpi	-3.081 (-2.30)	.4672 (3.29)	.3142 (3.21)	-.6031 (-4.21)
Δgdp	-.0223 (-2.44)	.0005 (.48)	-.0018 (-2.68)	-.0026 (-2.66)
Δlrate	-.0068 (-.02)	.1264 (2.64)	.0658 (1.99)	-.0158 (-.33)

table the six equations in the order of the previous tables are indicated by the rows and the four error correction terms are indicated by the columns. The first column is the real balance vector (mdciv), the second is the deflator real rate vector (defrrciv), the third is the term structure spread vector (termciv), and the fourth is the CPI real rate vector (ciprrciv). The first important feature of these estimates is that none of the coefficients in the funds rate row of the table are significantly different from zero. The implication of these results is that the Federal funds rate is "weakly exogenous" in this equation system. In the estimated structure, changes in the funds rate are an autoregressive process that is influenced positively by past growth in real output. A systematic investigation of the robustness of this result and its implication for policy analysis is discussed in Section 6. In contrast, in the remaining five equations, at least two of the estimated error correction coefficients are significantly different from zero. Changes in real balances respond significantly to both the lagged real balance vector and the CPI real rate vector. Changes in GDP inflation respond significantly to both lagged real rate vectors, though with opposite signs. Changes in the long-term interest rate respond significantly to the lagged Deflator real rate vector and to the interest rate spread vector. Changes in CPI inflation respond significantly to the lagged values of all four vectors, and real GDP growth responds significantly to all lagged vectors except the lagged Deflator real rate vector.

The importance of the error correction terms that appear in Table 10 provides an indication of how much relevant information is contained in the VECM specifications that presume the data adjust to departures from long-run equilibria. The literature of Vector Error Correction models is filled with references to the matrix of error correction coefficients as "speeds of adjustment" with comparisons to single equation "partial adjustment" models. Such analogies are incorrect. The Vector Error correction model is a reduced form structure and as such all the estimated coefficients, including the elements of the error correction matrix in principal are functions of many of the parameters of the underlying economic model. Indeed, Campbell and Shiller (1987), Rasche (1990) and Hoffman and Rasche (1996a) derive Vector Error Correction representations from economic models in which there are no adjustment dynamics. In these cases the elements of the error correction matrix are functions of the slope parameters in the economic model. Valid inferences about the response patterns of the variables in the system can only be drawn by examining the dynamic behavior of the full system of equations, not from coefficients of individual equations. The system dynamic responses to reduced form shocks are represented by impulse response functions (or dynamic multipliers). If sufficient identifying restrictions are available to identify, or overidentify, an economic model, then the dynamic responses to the stochastic elements in the economic model can be extracted from the reduced form impulse response functions.

Analysis of Residuals

Information on the residuals of each of the six equations of the VECM is presented in Figures 1-6. The upper left graph in each picture exhibits both the actual and fitted values of the levels of the respective variables. The lower left graph indicates the regression residuals standardized by the standard error of estimate of the equation. In all of these graphs the vertical line indicates the break in the sample period from 79:4 - 81:4. The points plotted to the left of the vertical line are from the period 56:2 - 79:3; the points to the right of the vertical line are from the period 82:1 - 96:1.

The upper right graph indicates the dispersion of the estimated residuals in comparison to a standard normal distribution. Finally the lower right graph indicates the first 12 autocorrelation coefficients of the estimated residuals. In cases where the autocorrelation coefficients are more than twice their estimated standard error, the graph will show vertical shading from top to bottom over the autocorrelation coefficient. In the absence of such shading, the presumption is that the estimated autocorrelations are not significantly different from zero.

The first thing to note about the residuals is that in all six cases they are relatively homoskedastic when comparing the observations prior to 79:4 with those of the 82:1 - 96:1 period. The variances of the real balance and long-term interest rate residuals may have increased slightly after 1981, but the variances of real GDP and the GDP Deflator inflation rate may have decreased slightly since that time. Little is changed in the magnitude of CPI and funds rate residual variation between the two periods.

Second note that compared with the distribution generated by a standard normal density, each of the equations exhibits some large outliers. We have examined each of the time series of residuals to determine the observations where an error of more than twice the standard error of estimate for the equation is observed. These observations, along with the standardized residual are recorded in Table 11.

There are 151 observations in the sample from which the estimates are constructed. Therefore, about 7 residuals whose absolute value is greater than 2.0 should be expected. The only equation for which the outliers are inconsistent with this standard is the long-term interest rate equation for which there are 9 residuals whose absolute value is greater than 2.0. Two of these are equal to 2.06, so this does not seem to be a major violation of normality. If the residuals were normally distributed there should be about 1-2 residuals whose absolute value is greater than 2.3. In the estimated real balance equation there are five residuals that fall in this group: those for 59:4, 60:4, 66:3, 89:2 and 92:1. The residual for 75:1 (2.10) is coincident with the breakdown in single equation partial adjustment models specified in levels of real M1 (see Hoffman and Rasche (1996a) for a review of these specifications), but the other periods are not particularly noteworthy in U.S. macroeconomic history.

Of the eight estimated residuals from the estimated GDP deflator change equation that are greater than 2.0 in absolute value, only three exceed 2.3. One of these large residuals, that observed in 78:2, occurs in the same period as a large residual is observed in the real GDP growth rate equation. Interestingly, the sign of the residual in both equations in this period is the same, hence this is not a situation where the model is failing to capture the mix of nominal GDP growth between real growth and inflation, but one where the estimation errors for inflation and nominal GDP reinforce each other and imply an even larger error for nominal GDP. In none of the eight observations of outliers in the GDP inflation equation is there a corresponding outlier in the CPI inflation equation. Hence these large errors do not appear to represent a failure of the model to predict inflation in general, but only to predict how the GDP deflator is constructed as a measure of inflation. The construction of the National Income and Product Accounts in the period in which both the GDP deflator inflation equation and the real GDP growth equation both exhibit a large error in the same direction warrants a close examination.

The remaining huge residuals in the real GDP equation are in 58:1 and 60:4 - coincident with the beginning of recessions. In both cases the equation substantially over predicts real GDP. Since this equation, except for the presence of the error correction terms this equation and a weak feedback from changes in the funds rate, is a univariate

AR process, it is not surprising to observe a large residual at a sharp turning point in the level of the series.

Of the seven large outliers in the CPI inflation change equation, two of them are associated with major price changes in the world energy market. The biggest outlier for this series occurs in 86:2 (-3.34), coincident with the collapse of oil prices. A second large outlier (2.10) occurs in 90:3 coincident with the increase in oil prices associated with the outbreak of the Gulf conflict. Remarkably the CPI inflation equation does not exhibit any large outliers during the 74 and 79-80 oil shocks. A third large outlier in the CPI inflation equation occurs in 71:4 (-2.16) at the time of the imposition of the Nixon wage-price freeze.

The outliers for the two interest rate equations are not particularly numerous (7 for the long rate equation excluding the two residuals of 2.06 and 7 for the funds rate equation), but are remarkable for their size. Seven the observed standardized residuals in each of these two equations are larger than 2.2. In the funds rate equation three of the seven outliers are greater than 3 in absolute value; in the long-term rate equation two are greater than 3. These are large departures from normality. None of the outliers in the two interest rate equations are coincident, and most are not particularly close to each other in time. A large negative residual occurs in the funds rate equation in 74:4, immediately after the Franklin National Bank crisis (May - October , 1974). No major outliers are observed in this series since 1985, in particular none is observed at the time of the stock market crash in autumn 1987.

All but one of the outliers in the long-term rate equation occur after 1981. Thus the three large unexplained shocks to the funds rate in the mid 70s appear to be unique to that rate and were not transmitted through the term structure. Throughout the 80s a major outlier is observed in the long-term rate equation roughly once in each year. There is no immediately apparent explanation for the size of the equation errors during this period.

Table 11

Outlier Analysis for VECM Estimated on 56:2 - 96:1 omitting 79:4 - 81:4

A. Real M1 Equation

Dates	Standardized Residual of Real Balance Equation
59:4	-2.82
60:4	-2.51
66:3	-2.98
75:1	-2.10
85:3	2.06
89:2	-2.47
92:1	2.62

B. GDP Deflator Equation

Dates	Standardized Residual of Deflator Equation
57:1	2.14
57:4	-2.46
62:1	2.01
69:1	-2.33
74:3	2.47
76:1	-2.02
78:2	2.00
86:1	-2.21

Table 11 Continued

Outlier Analysis for VECM Estimated on 56:2 - 96:1 omitting 79:4 - 81:4

C. Funds Rate Equation

Dates	Standardized Residual of Funds Rate Equation
69:2	2.28
73:3	3.64
74:4	-2.57
75:1	-3.29
82:1	3.98
84:4	-2.65
85:2	-2.20

D. CPI Inflation Equation

Dates	Standardized Residual of CPI Inflation Equation
58:1	2.60
67:1	-2.76
71:4	-2.16
77:1	2.39
86:2	-3.34
87:1	2.18
90:3	2.10

E. Real Output Equation

Dates	Standardized Residuals of Real GDP Equation
58:1	-2.93
60:4	-3.29
64:1	2.19
71:1	2.15
78:2	3.13

Table 11 Continued

Outlier Analysis for VECM Estimated on 56:2 - 96:1 omitting 79:4 - 81:4

F. 10 Year Government Rate Equation

Dates	Standardized Residual of Long Rate Equation
58:3	2.06
82:4	-4.54
83:3	2.39
84:2	3.50
86:1	-2.36
87.2	2.26
88:2	2.37
92:3	-2.43
94:2	2.06

Figure 1: VAR Equation for Real M1

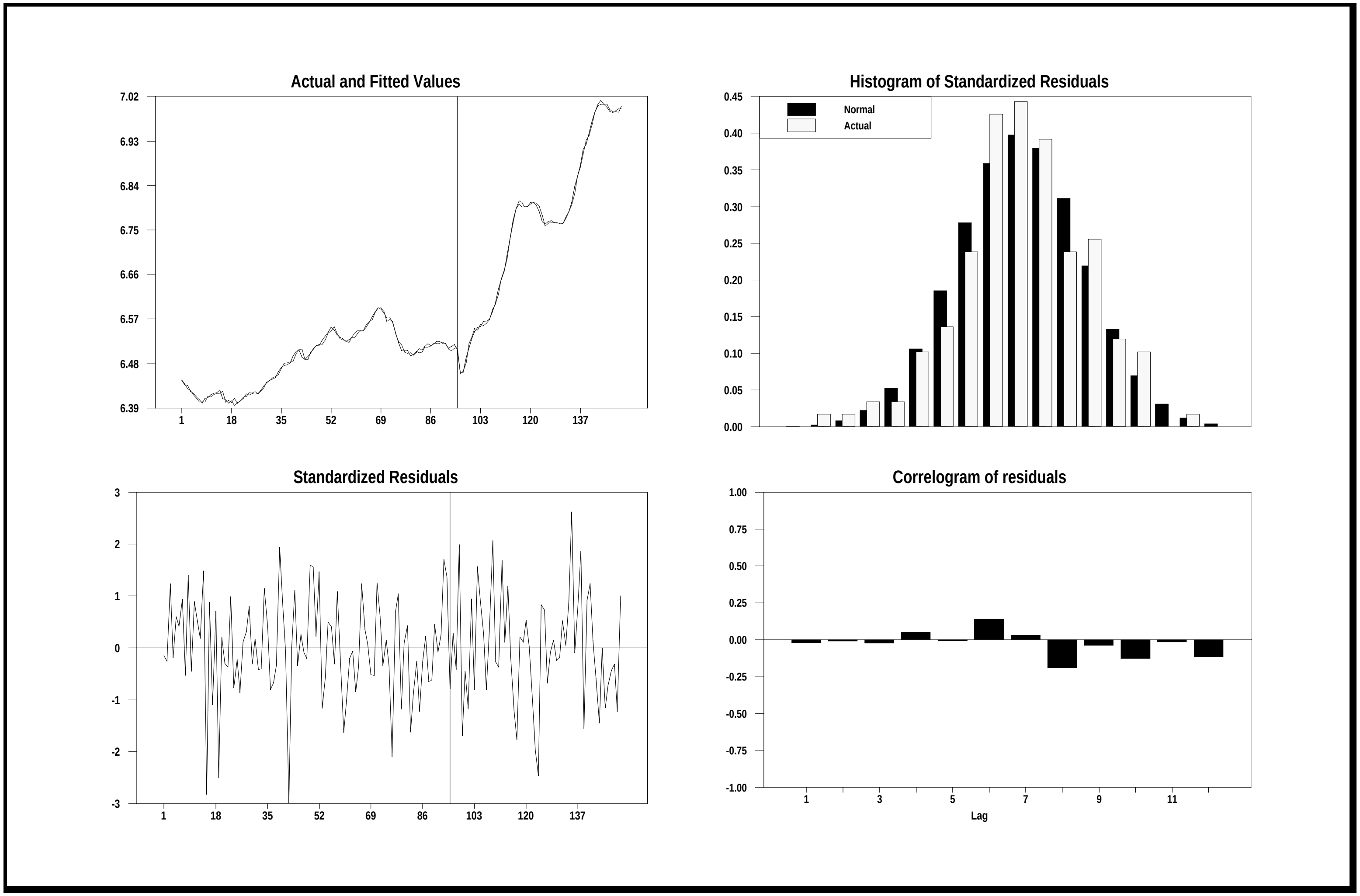


Figure 2: VAR Equation for Deflator Inflation Rate

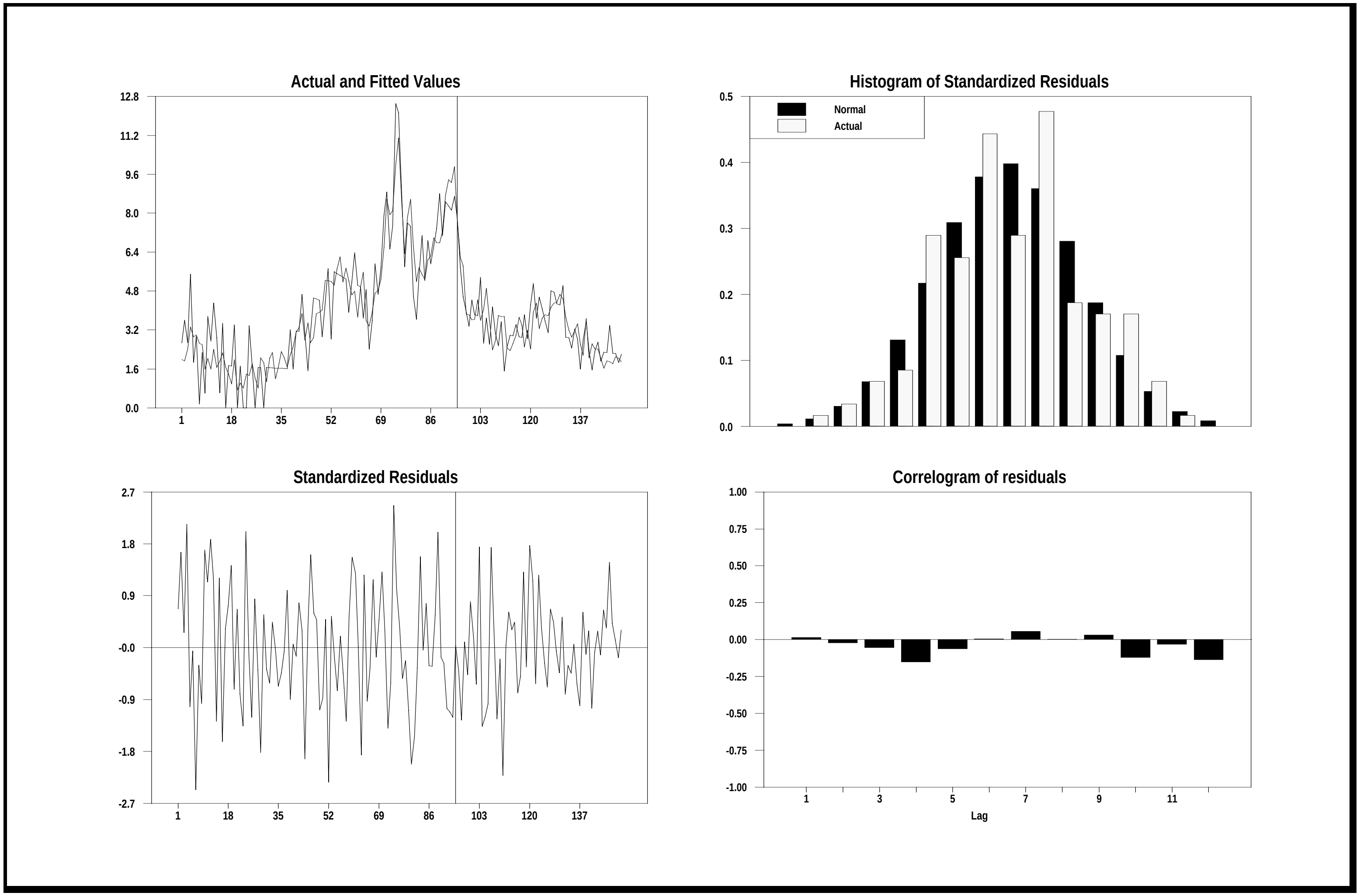


Figure 3: VAR Equation for Funds Rate

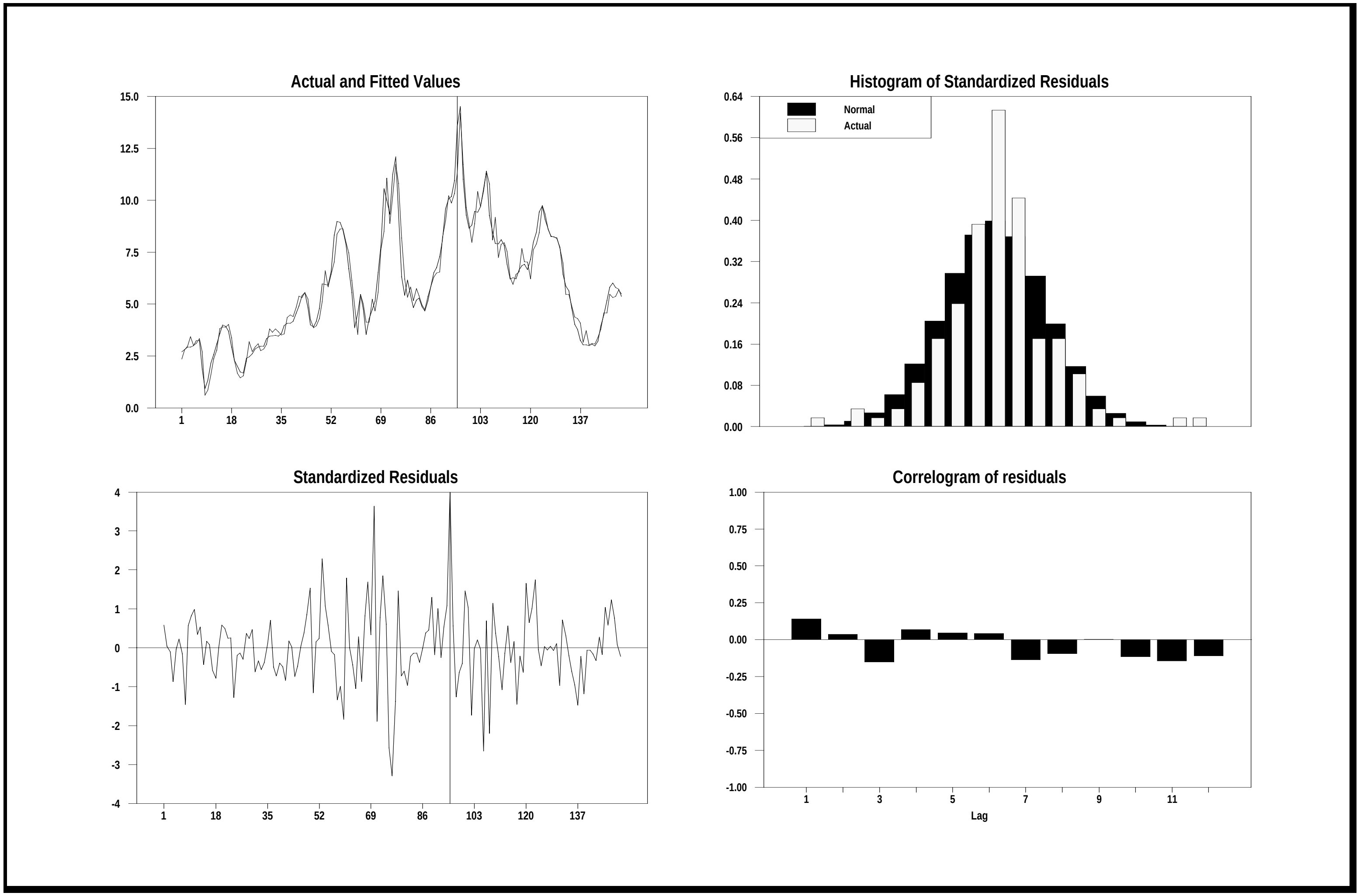


Figure 4: VAR Equation for CPI Inflation Rate

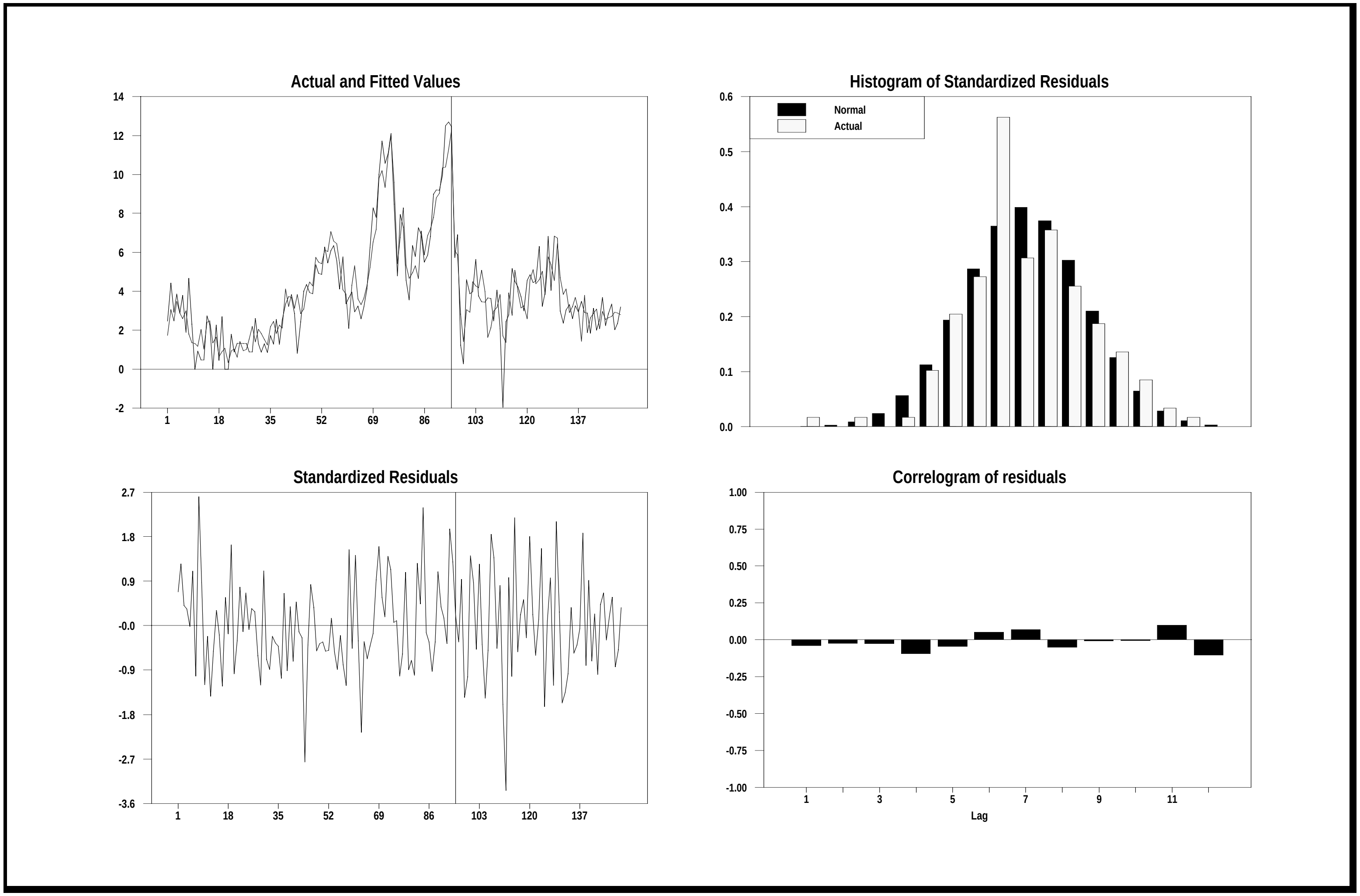


Figure 5: VAR Equation for Real GDP

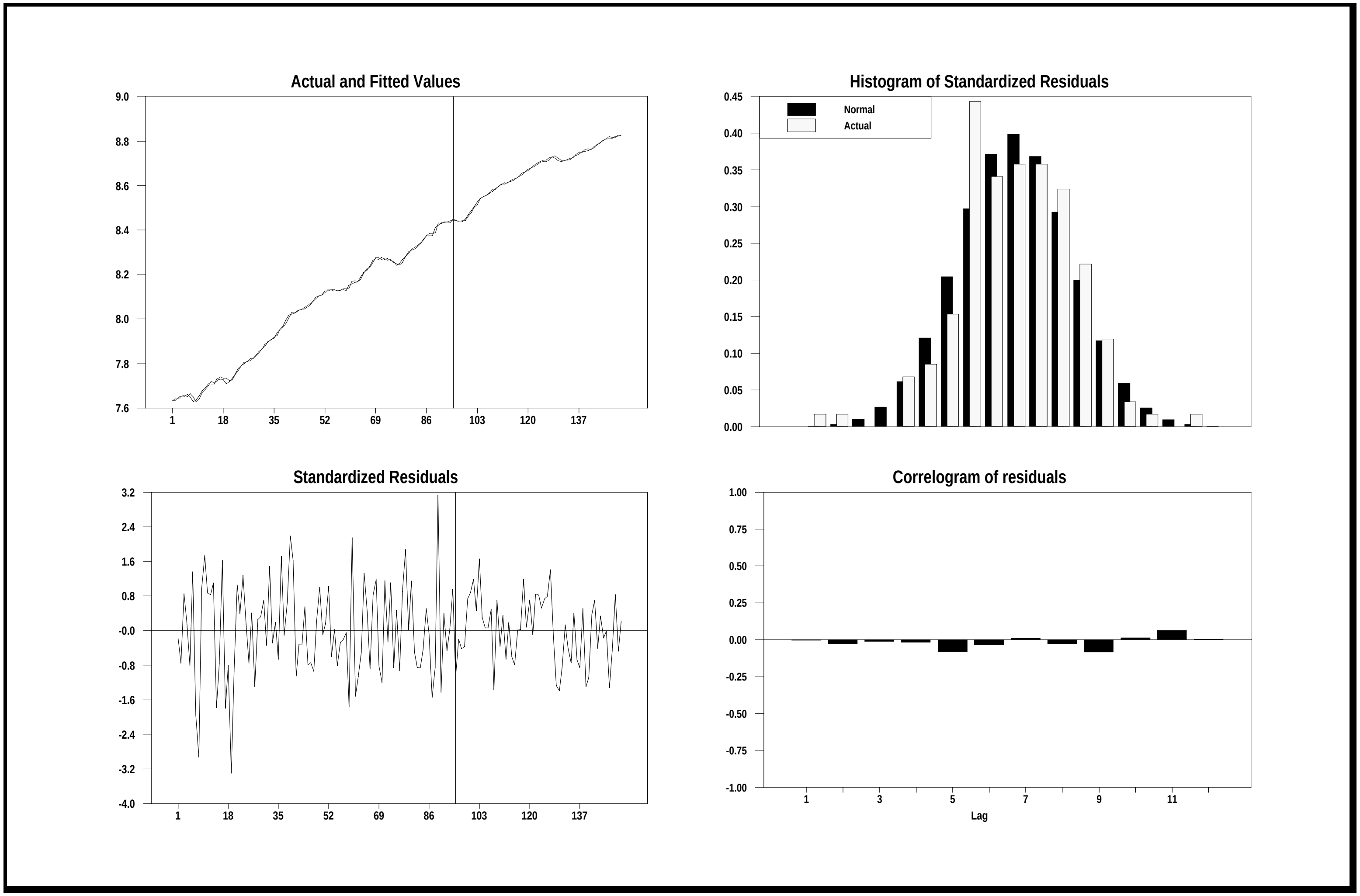
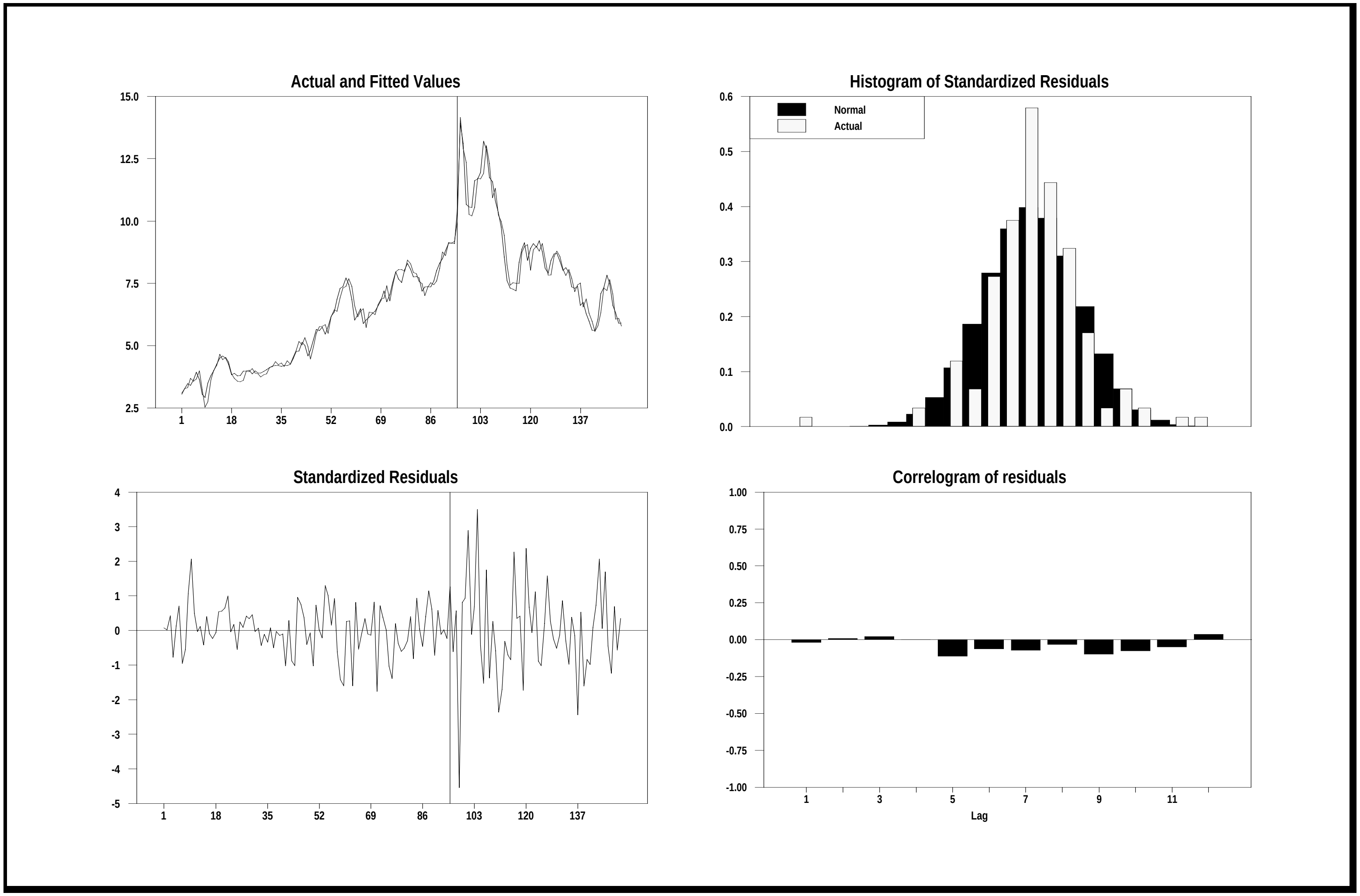


Figure 6: VAR Equation for Long Term Rate



4. Assessing Forecast Performance

Forecasting performance may be gauged in a number of different ways. Papers by Clements and Hendry (1993) and Hoffman and Rasche (1996b) employ measures of system performance, while Clements and Hendry (1993) and Christofferson and Diebold (1996) argue that conventional RMSE criterion may not capture some of the advantages of long-run information into the system. The basic conclusion of this body of literature is that incorporating cointegration may improve forecast performance, but improvement need not show up only at longer horizons as predicted originally by Engle and Yoo (1987). The advantage presumably accrues from the addition of error correction terms in VECM representations. Christofferson and Diebold (1996) contend that conventional RMSE criterion will not capture this forecast advantage at long forecast horizons simply because the importance of the error correction term diminishes with increases in the forecast horizon. For the exercise we have in mind, the relevant issue is forecast performance for a subset of the variables in our system, at various horizons, and the most relevant measure of that performance is the standard mean squared error criterion. We employ RMSFE as a criterion while recognizing that it may not capture all the advantages that the long-term information has to offer.

First, we can measure absolute performance using conventional root mean squared error (RMSE) criterion in an outsample period. The sample period used to provide baseline estimates for this exercise spans 56:2 -87:4. This provides 32 periods for outsample analysis from 88:1 - 95:4. We used the outsample information to measure forecast performance for 1, 2, 4, 8, 12, and 16 period forecast horizons. Over the 32 periods we calculated the 32 one-step ahead forecasts, 31 two-step ahead forecasts, etc. The forecast errors obtained in this exercise are then squared and averaged. The square root of these averages is displayed in Table 12 as the RMSFE's in the outsample period.

Table 12

Root Mean Square Forecast Errors at Various Horizons
(formed over outsample period 88:1 - 95:4)

Horizon	m1p	m1	funds	gdp	lrate
1	.012	.010	1.35	.010	.440
2	.026	.022	2.35	.019	.742
4	.050	.043	2.53	.042	1.31
8	.086	.070	2.01	.081	1.81
12	.107	.078	1.72	.127	1.98
16	.108	.071	2.06	.168	2.17
	Δm1p	Δm1	infdef	infcpi	Δgdp
1	.012	.010	1.22	1.30	.010
2	.016	.014	1.22	1.39	.011
4	.014	.013	0.94	1.45	.015
8	.012	.011	1.30	1.61	.013
12	.013	.013	1.10	1.10	.012
16	.014	.014	0.72	0.81	.010

Note: The entries for m1p, m1, gdp, Δm1p, Δm1, and Δgdp are forecast errors measured in percent. The entries for funds, lrate, infdef, and lifcpi are forecast errors calculated additively from the corresponding rates -- measured as annual rates of return.

We have depicted the RMSFE performance of the model for all of the variables in our system. Moreover, some of the variables in this table are presented in both levels and difference form to allow forecast performance assessment to reveal how well the model predicts both levels and the changes in the series. We also portray the RMSFE for the nominal money stock series that may be imputed from the forecasts of real money balances and the GDP deflator.

The RMSFE's obtained for the levels of both real and nominal money balances follow a similar pattern across forecast horizons. Each begins with about a one percent error at horizon 1 and steadily increases until the root-mean-squared forecast error 16 periods out is about eight percent. Forecast errors for the GDP series follow a similar pattern up to the two year horizon with root-mean-square forecast errors less than seven percent at that point (in line with the RMSFE's for m1 and m1p), but forecast errors then accelerate and reach 13 percent over the 16 period (four year) horizon. The RMSFE's for the two inflation series display a pattern that is quite different from the money and GDP forecasts. RMSFE's for these series tend to be in the vicinity of 1.30 percent (measured additively at annual rates) regardless of the forecast horizon. Indeed the pattern reveals a slight tendency for the RMSFE's to fall as the forecast horizon increases. Federal funds rate forecast errors average just over 100 basis points at short horizons and increase to just over 200 basis points at the two year horizon and remain at about that level for up to four year (16 period) forecasts. The RMSFE's for long-term interest rates display a similar pattern, though short horizon performance is better (only 44 basis points at the one period horizon) and forecast RMSFE's for the long-term interest rates then increase steadily as the horizon lengthens. The RMSFE comparison may also be extended to the growth rates of selected variables in our model. Of course, in the one period horizon, these match the RMSFE performance for the corresponding levels of the variables discussed above. The model produces growth rate forecasts for the money measures that exhibit RMSFE's in the range of 1.2 percent and are quite constant across forecast horizons. Interestingly, though the forecasts for the level of gdp exhibited RMSFE's that increased sharply with increases in the forecasting horizon, the RMSFE's associated with the gdp growth averaged about one percent across all horizons. Of course, while missing by one percent in a one quarter horizon is not particularly satisfying, missing by only one percent at the 4 and 8 period horizons would seem to be a reasonable forecast performance.

Though measuring forecast performance over the eight year out-sample may be revealing, it is useful to analyze performance over two particular periods in the out-sample. We focus, in turn, on the 90-91 recessionary period and the 94-95 expansionary period.

Figure 7 reveals that the six variable VECM model clearly fails to pick up the downturn that occurred in late 1990. By the end of 1990, the forecast errors for GDP growth are in the range of about two percent. This pattern prevails regardless of whether the simulation model is allowed access to actual data through 1987 or all the way through 1989. Interestingly, the forecast errors for growth in 1991 are generally smaller, typically less than one percent. Forecast errors for the Federal funds rate are depicted in Figure 8.

The VECM with information through 1989 produces very accurate forecasts of the funds rate through 1990, but clearly fails to pick up the downturn in rates that occurred in 1991. Comparable forecast errors for the CPI inflation rate appear in Figure 9 for the same period. In 1990 the model performs quite well, producing forecast errors that are less than one percent on average. However, the model clearly over forecasts the slowdown in

Figure 7

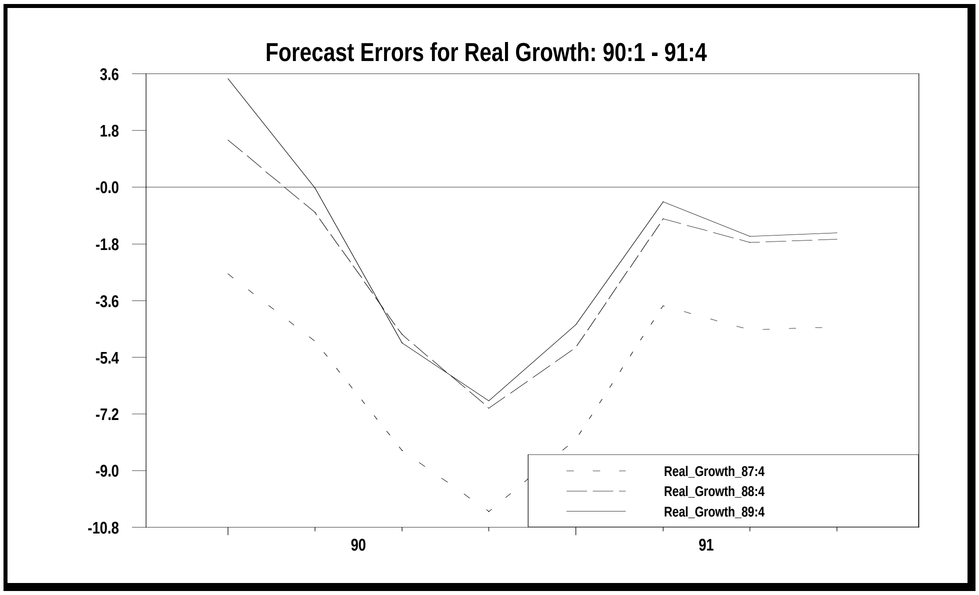


Figure 8

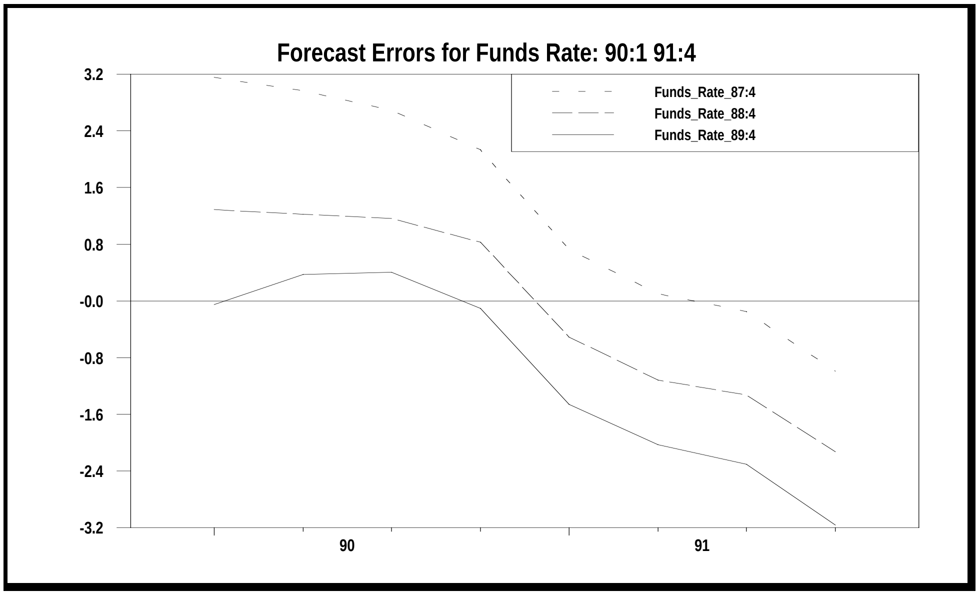


Figure 9

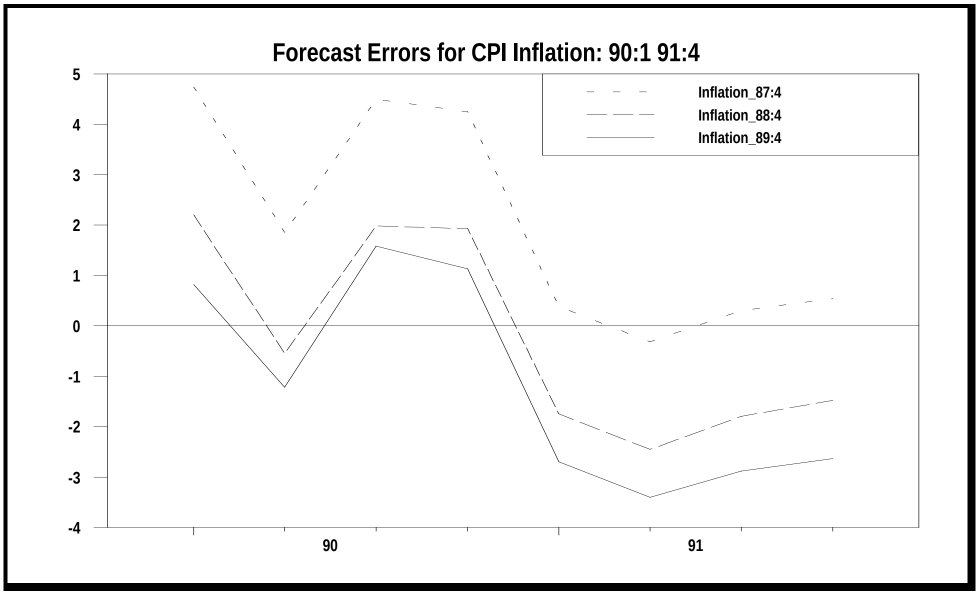
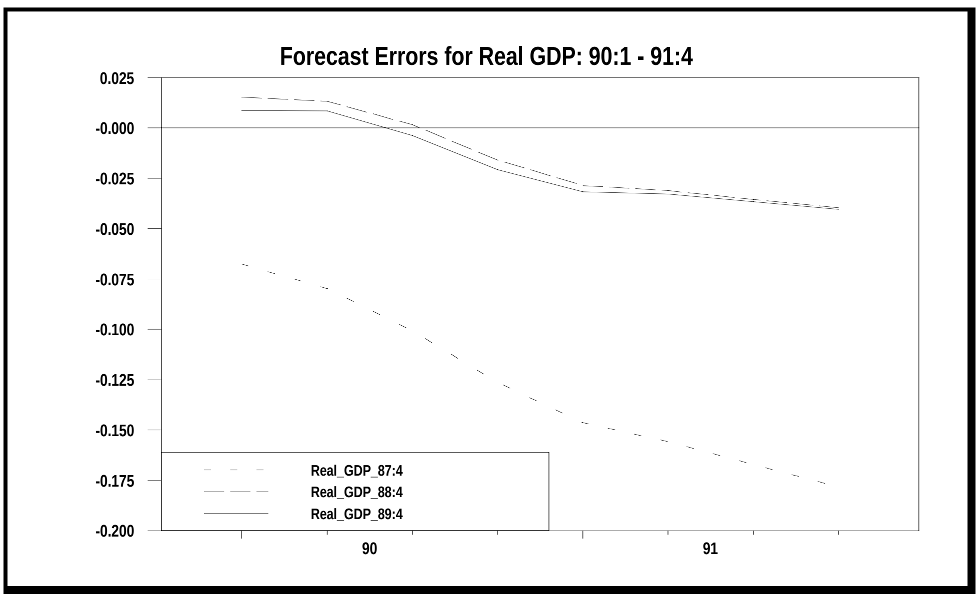


Figure 10



inflation that actually occurred in 1991, resulting in forecast errors of over two percent for forecasts based upon information through 1989. Similar forecast errors for the level of real GDP appear in Figure 10. The Figure reveals that the model's over prediction of GDP growth yields forecast errors for the level of real GDP that are about four percent low by the end of 1991.

The performance of the model over the period 1994 - 1995 appears in Figures 11 for GDP growth. Figure 11 reveals that the model performs quite well in the 94-95 period. Forecast errors real GDP growth average less than 0.5 percent in 1994 and are slightly smaller in 1995. Moreover, the forecast horizon comparisons reveal that essentially the same forecast error pattern is observed regardless of whether the model is allowed to access information through 1993. Figure 12 depicts the forecast errors for the Federal funds rate over the 1994-1995 period. Again, in the first year of the forecast horizon the VECM produces very small Federal funds rate forecast errors and the errors in the second year (1995) average less than two percent. Forecast errors for the CPI inflation rate for the 1994-1995 period appear in Figure 13. Again the model performs quite well. Forecast errors for this variable average about 0.5 percent in 1994 and just under one percent for 1995. Figure 14 depicts the forecasts of the level of real GDP over the 1994-1995 period. The estimates reveal a slight tendency to over predict, though the forecast error for the level of real GDP is less than 1.5 percent by the end of the two year period.

Figure 11

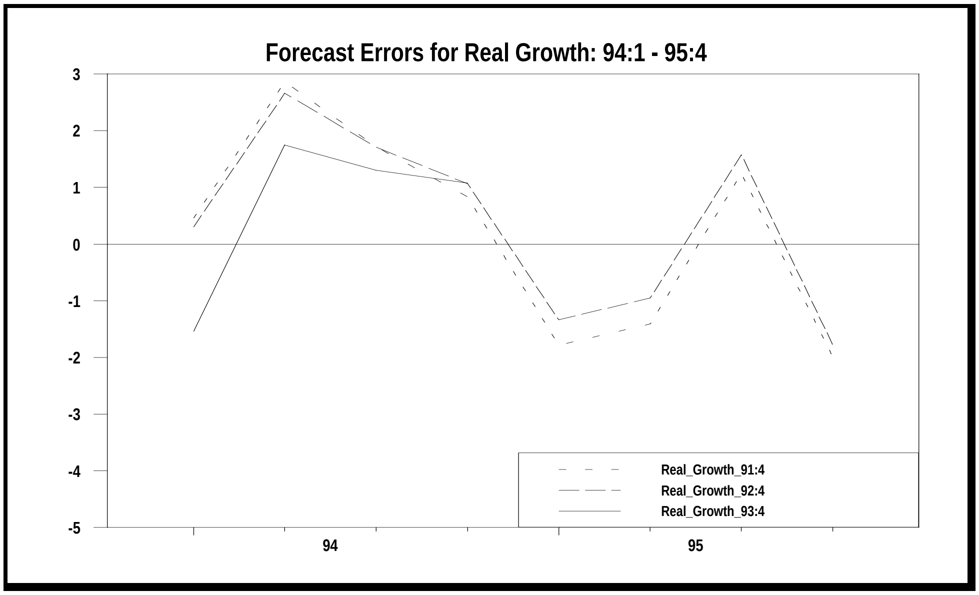


Figure 12

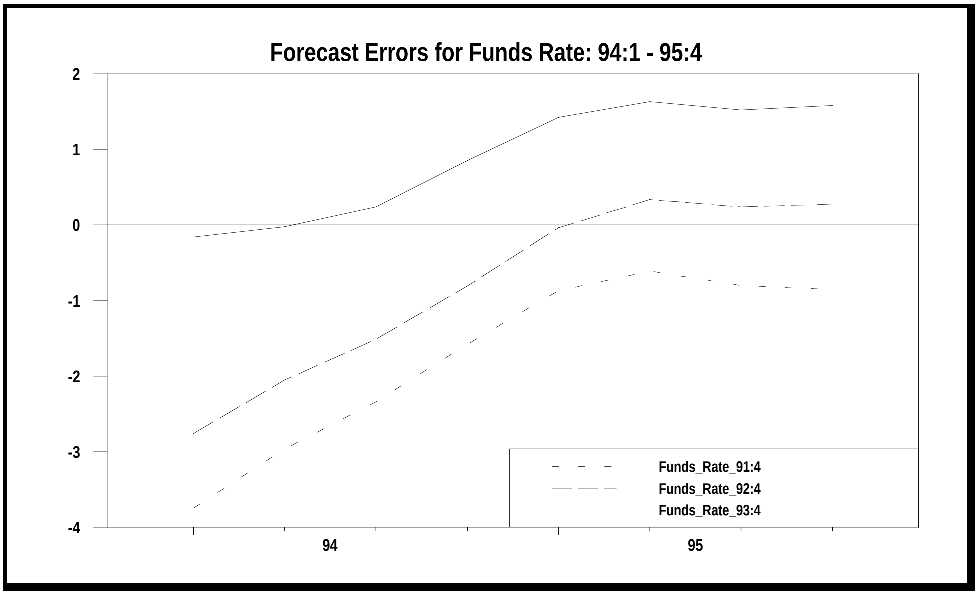


Figure 13

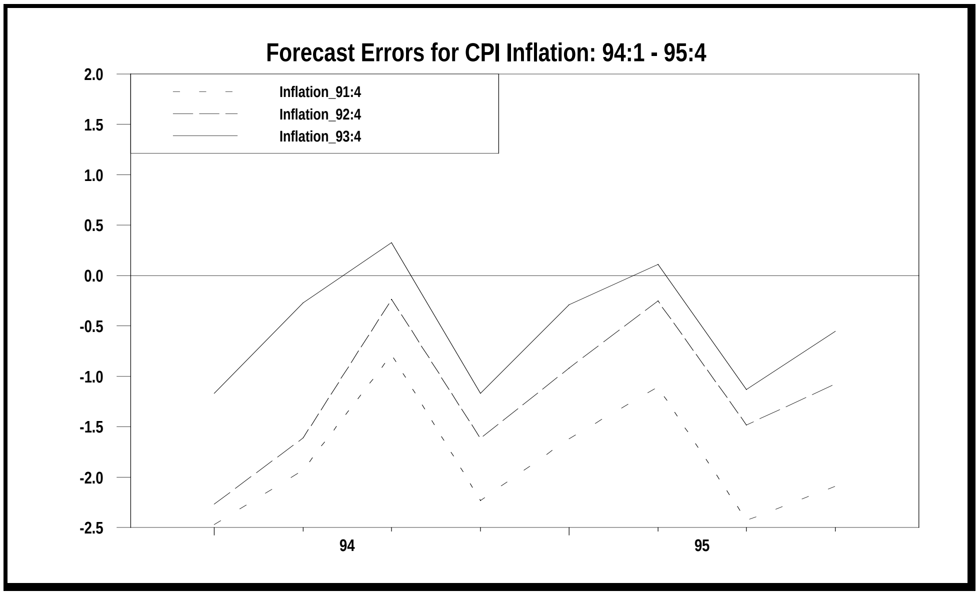
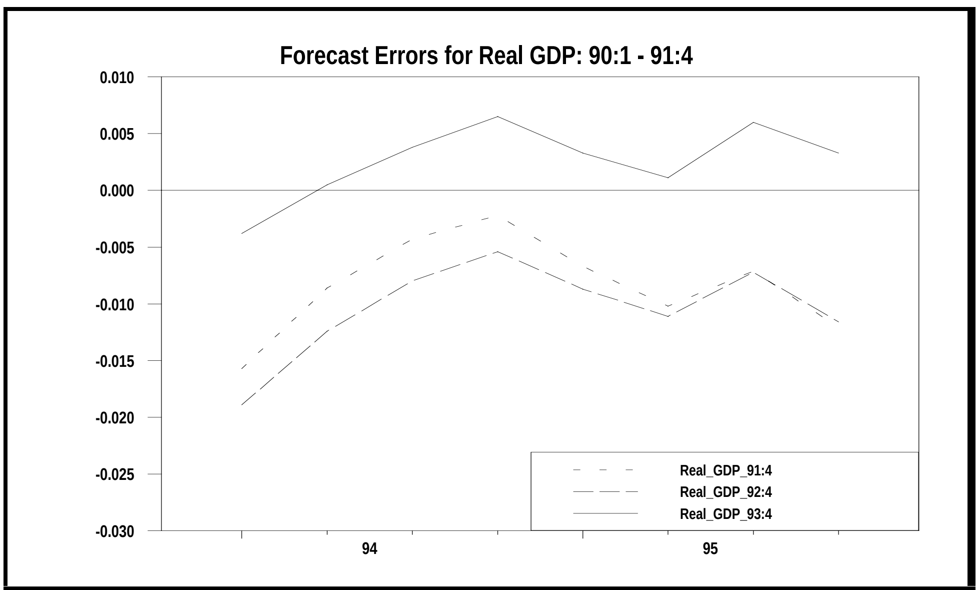


Figure 14



Analysis of forecast performance for GDP growth and CPI inflation in our model reveals that the projections from the VECM fail to capture the slow growth that was observed in the 90/91 period. In contrast, smaller forecast errors were observed for the 1994-1995 expansionary period. Another perspective is obtained by comparing our GDP and inflation projections with those of the FOMC Green Book. Since Green Book forecasts only become public information after a five year lag, we are only able to compare our own performance with Green Book projections for the first of these two periods. For the 1994-1995 period we turn to midpoint of the central tendency range reported by the FOMC in conjunction with the Humphrey-Hawkins hearings. The Green Book forecasts are compiled for a two year forecast horizon and compare quite nicely with the information allowed our VECM from the 1989 forecast. However, the central tendency ranges are based upon February, 1994 for the 1994 forecast horizon and July, 1994 for the 1995 forecast horizon, so the 1995 forecasts allow about six months more information than is available to the VECM in the 1995 period. In each case we have access only to the CPI and real GDP growth forecasts.

Table 13 contains the average annual forecasts for GDP growth and CPI inflation for the years 1990-1991 and 1994-1995 periods. Both Green Book/Humphrey-Hawkins and VECM forecasts are presented along with the actual values observed over the two periods. The reference points for each forecast are 89:4 and 93:4 respectively for the VECM. As indicted in the preceding discussion, the VECM projections overstate both real growth and inflation in the 1990-1991 period. Table 13 reveals that similar forecast errors were realized by the FOMC Green Book projections over the same period. Neither forecast picked up the recessionary period that occurred less than one year after the forecasts were compiled. Also, neither the Green Book nor VECM projections foretold the moderation in inflation that occurred in 1991. Forecast performance for the VECM improved markedly for the 94-95 period. The model predicts moderate growth in 1994 followed by slower growth in 1995. Indeed, the projected growth of real GDP for 1995 of 1.5 percent is remarkably close to the observed 1.52 percent growth rate. The model predicted modest and stable inflation rates that were slightly higher than those observed over the 94-95 period.[13] In contrast, the midpoint of the FOMC central tendency ranges displayed more accuracy with respect to real GDP growth in 1994, but failed to pick up the deterioration in growth that occurred in 1995. Similarly, the midpoint of the central tendency ranges substantially over predicted inflation in both 1994 and 1995.

Table 13

A Comparison of Forecast Performance: Green Book/Humphrey-Hawkins
vs. Forecasts taken from the Six Variable VECM
1990-1991
(annual average growth rates)

Year	Green Book/ Humphrey-Hawkins		VECM		Actual Values	
	Δgdp	infcpi	Δgdp	infcpi	Δgdp	infcpi
1990	1.6	4.4	2.2	5.5	-.22	6.11
1991	2.3	4.5	2.4	5.8	.27	2.92
1994	3.125	4.25	2.5	3.2	3.61	2.61
1995	2.625	4.125	1.5	3.3	1.52	2.63

Our analysis of forecasting performance reveals that a simple VECM characterized by long-run information in the form of Fisher equations, a term structure relationship, and a long-run money demand relation does an adequate job of forecasting over two recent periods, 1990-1991 and 1994-1995. Moreover, when compared with the point forecasts compiled in the Federal Reserve's Green Book or in concert with the Humphrey-Hawkins hearings, the VECM forecast errors were as small as or even smaller. We simply choose the 1990-1991 period and the 1994-1995 period to contrast performance in distinct contractionary and expansionary periods. We leave further investigation of the relative forecast performance of the VECM to future research.

[13] Actual growth for gdp over these periods is measured by multiplying log differences of quarterly data by 400 and then averaging the resulting quarterly growth rates over the indicated periods.

5. The Role of Permanent and Transitory Shocks

The Granger representation theorem demonstrates that the long-run features of a six dimensional system characterized by cointegration rank four may be expressed in terms of two distinct independent stochastic trends and four independent transitory innovations. It is important to understand the role of both types of shock in the behavior of the particular variables.

Isolating the effects of the individual shocks requires differing sets of identifying assumptions, depending on how an extensive analysis is desired. At one level, the assumption that the two different types of shocks, permanent and transitory, are independent is critical to, but not quite sufficient for, the decomposition of each of the time series into a component generated by the permanent shocks and a second component generated by the transitory shocks. A much more extensive set of assumptions is required to define (identify) and isolate the dynamic effects of the individual permanent and transitory shocks. Identification schemes which accomplish this are discussed in the following section.

Two different types of permanent-transitory decompositions have been proposed in the literature. The first, proposed by King, Plosser, Stock and Watson (1991) is a generalization of the Beveridge-Nelson decomposition for a univariate process. King, Plosser, Stock and Watson (KPSW), in addition to assuming that the permanent and transitory shocks are uncorrelated, assume that the permanent shocks are random walks. These two assumptions are sufficient to allow decomposition of each of the time series into a component originating from the permanent shocks, a component originating from the transitory shocks and a constant generated by the initial conditions. The actual values of each of the six time series in the model are plotted along with the estimated permanent component of the series over the period from 82:1 to the end of the sample in Figures 15 - 20. For this decomposition, the initial conditions have been allocated by setting the mean of the transitory component equal to zero.

The trend in the permanent components of GDP and the two inflation rates track the trend in the actual values of these series quite closely over the fourteen year period.

The transitory component of GDP starts substantially negative in the recession year 1982, and remains negative through 1985. It then becomes substantially positive by the late 1980s, before turning negative again during the 90-91 recession. In 1993 and 1994 the permanent component of GDP tracks actual real GDP quite closely, but then declines sharply in early 1995 and remains below the actual real GDP for the remainder of the sample period so that the implied transitory component of real GDP in 1995 is positive.

The permanent component of both measures of inflation not only tracks the trends in these series quite well, but also captures the lower frequency cyclical fluctuations in these series. The only exception to this broad characterization of the data is in 1982-3 and again in 1994-5 for the deflator inflation rate. In the former case, the measured transitory component of deflator inflation is systematically positive; in the latter case it is systematically negative. For both series the measured permanent inflation component captures the upward movement of actual inflation leading up to the Gulf conflict, and the

Figure 15

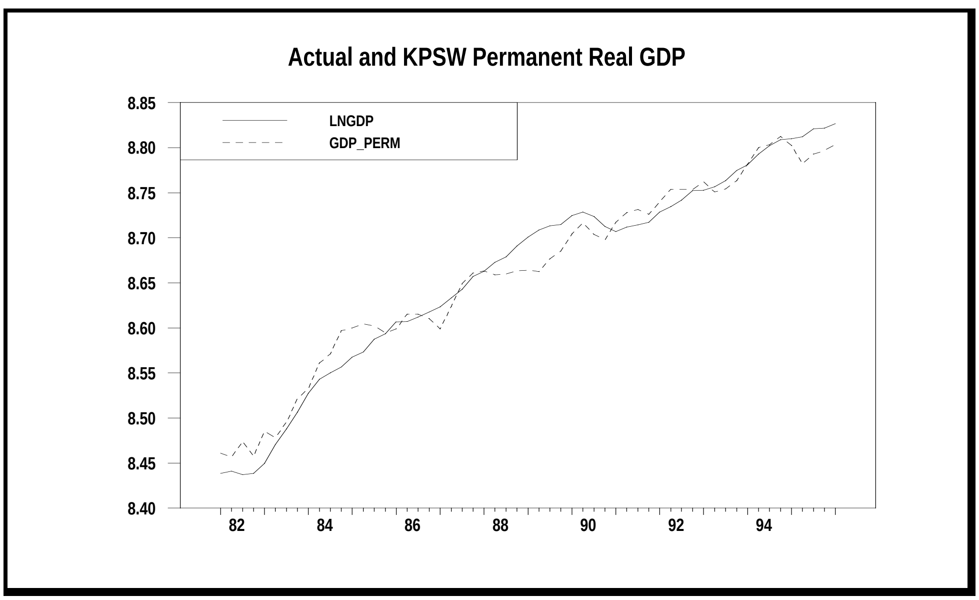


Figure 16

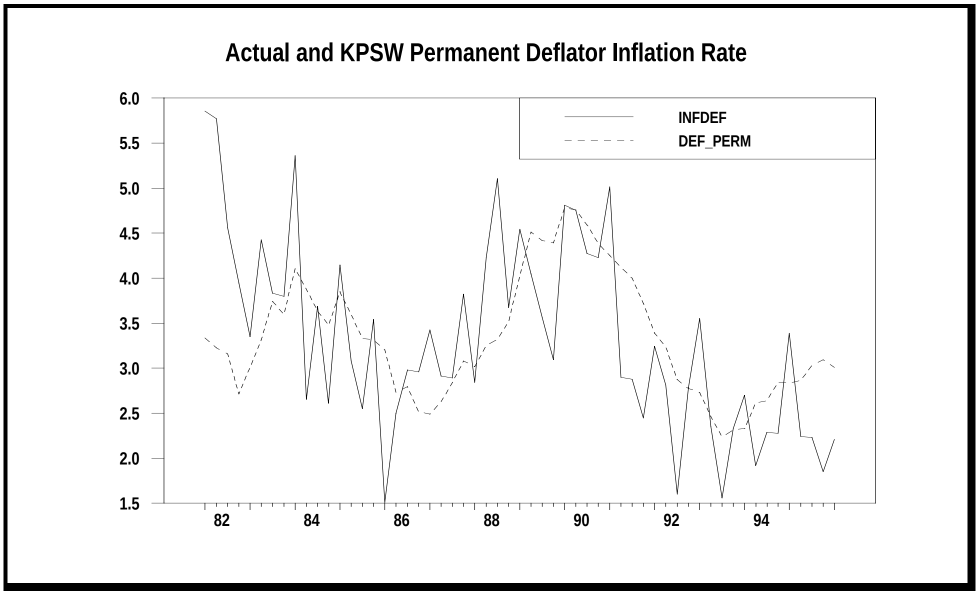


Figure 17

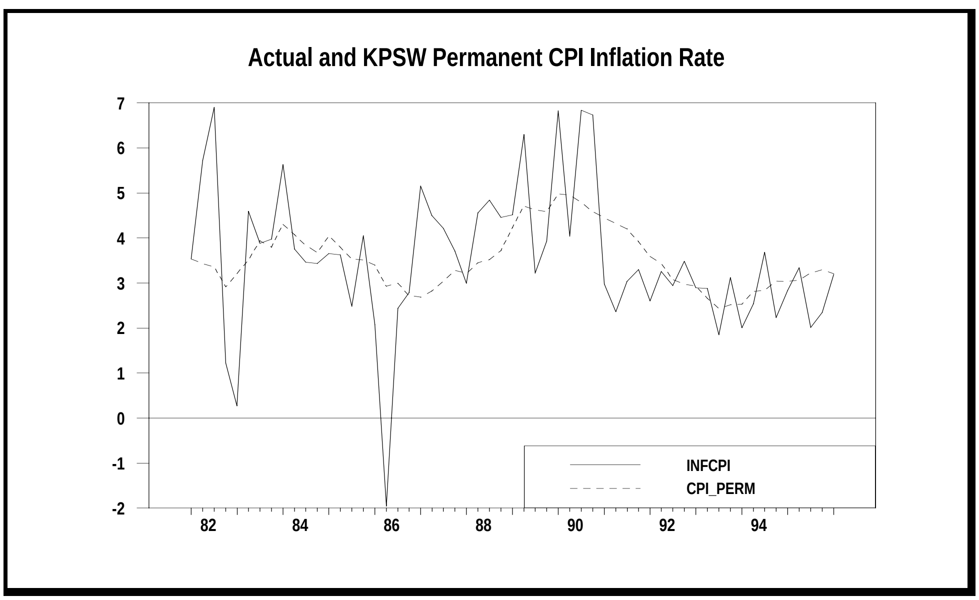


Figure 18

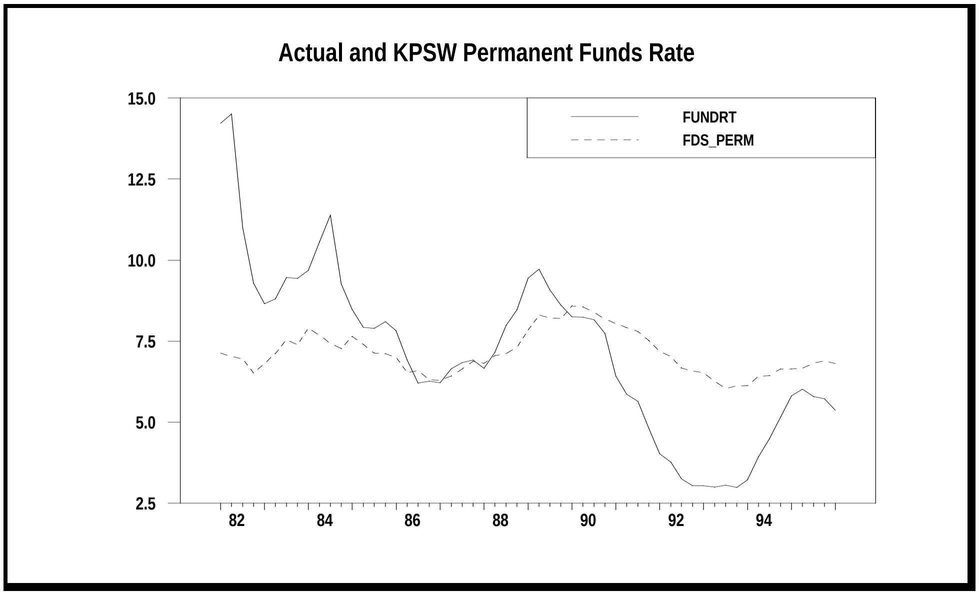


Figure 19

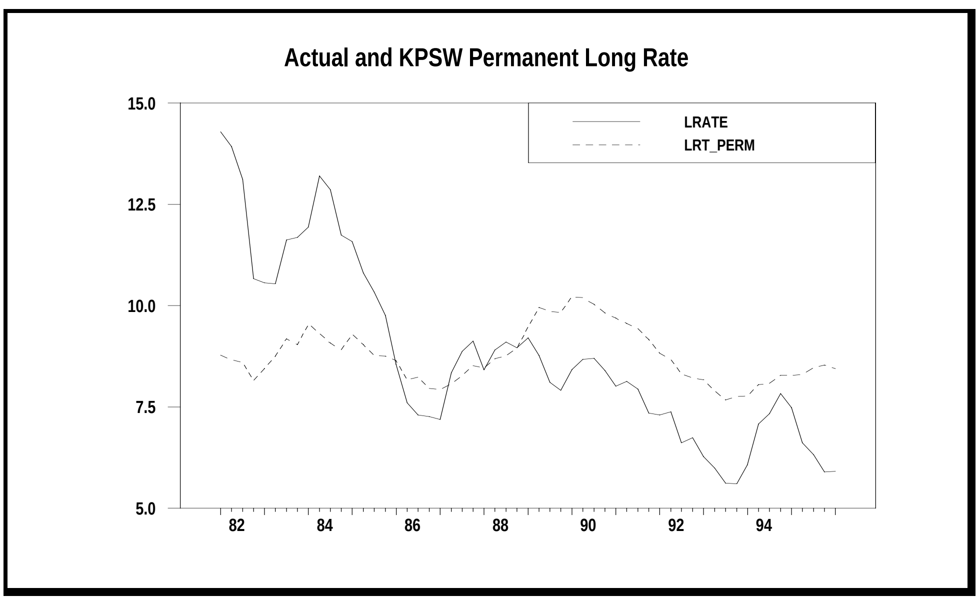
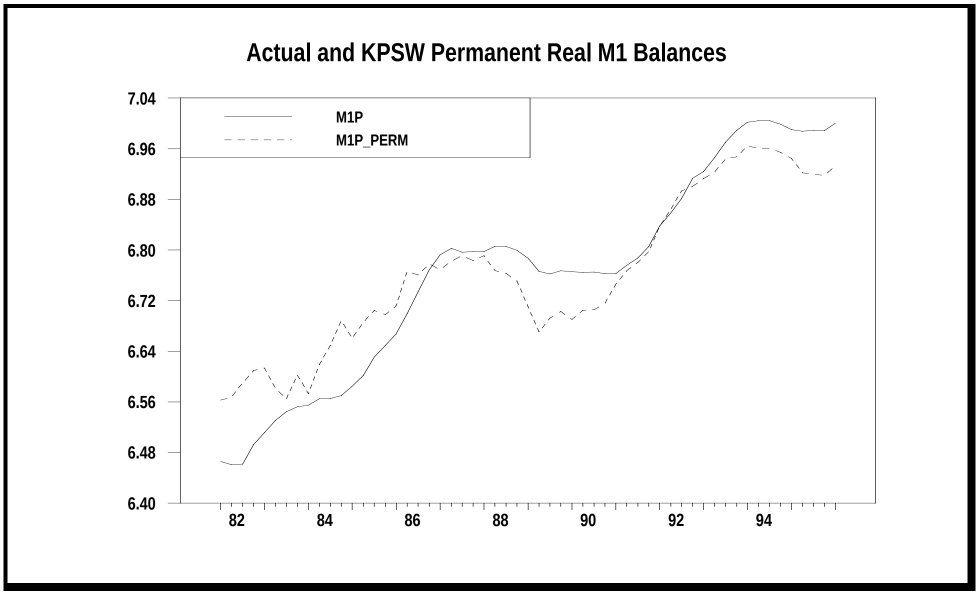


Figure 20



subsequent slowdown in CPI inflation to the neighborhood of three percent. The trends in the permanent components of both measures of inflation must be the same, since the VECM assumes that the cointegrating relationship between the two is stationary around a constant mean over this period. The permanent component of the inflation rate is of most interest to monetary authorities who are concerned about future inflation trends and not quarter-to-quarter transitory fluctuations in the inflation rate.

The measured permanent components of both interest rates reflect the same trends as the two measure of inflation, again because the VECM assumes that the estimated real rate cointegrating vectors are stationary around a constant mean. The trends in the actual interest rate series over this sample period are not fully reflected in the trends in the estimated permanent components of the series. In both cases, the estimated transitory component is positive throughout the 90s. At the end of 1995 the permanent component of the long rate was substantially higher than the actual long rate. In this light, the sharp increase in long-term interest rates in the early months of 1996 is seen as a reversion of such rates to their permanent levels.

The trend in the estimated permanent component of real balances also fails to perfectly replicate the trend in actual real balances from 1982-95. Since the VECM presumes that a linear combination of real balances, real income and nominal interest rates is stationary around a constant mean, the observed behavior of the two real M1 series is just reflecting the patterns observed in the interest rate graphs, given that the trend in the estimated permanent component of real GDP is very close to the trend in actual real GDP.

The KPSW permanent/transitory decomposition is not unique. Park (1990) and Gonzolo and Granger (1991) have proposed alternative decompositions. Both of these approaches identify the permanent/transitory decomposition by the assumption that the permanent components are linear combinations of the observable variables, and that the transitory components are stationary. The two studies differ in the normalizations that they apply to the transitory components. Gonzolo and Granger just measure the transitory component by applying the cointegrating vectors to the current observed values of the data series. Park uses normalized cointegrating vectors to construct what he calls a "common factor" decomposition.[14] These "common factor" permanent components of each series are plotted in Figures 21-26.[15]

The "common factor" of real GDP tracks the trend in actual GDP over the 1982-95 sample period, but has a larger variance than the "permanent component" measured by the King, Plosser, Stock and Watson identifying restrictions, and implies a larger variance for the implied transitory part of real GDP. This higher volatility of both the permanent and transitory components is possible because the "common factor" decomposition does not constrain the two components to be uncorrelated.

The "common factor" of both inflation rates is very similar to the KPSW "permanent components" of these series, with the exception that the "common factor" for the deflator inflation rate is much closer to the actual values of the inflation rate at the end

[14] In estimating the "common factors" we have allocated the constant term in each equation of the VECM so that the mean of the transitory component from 82:1 through the end of the sample period is zero.
[15] The transitory components in the common factor decomposition are computed as $\alpha(\beta'\alpha)^{-1}\beta' X_t$.

Figure 21

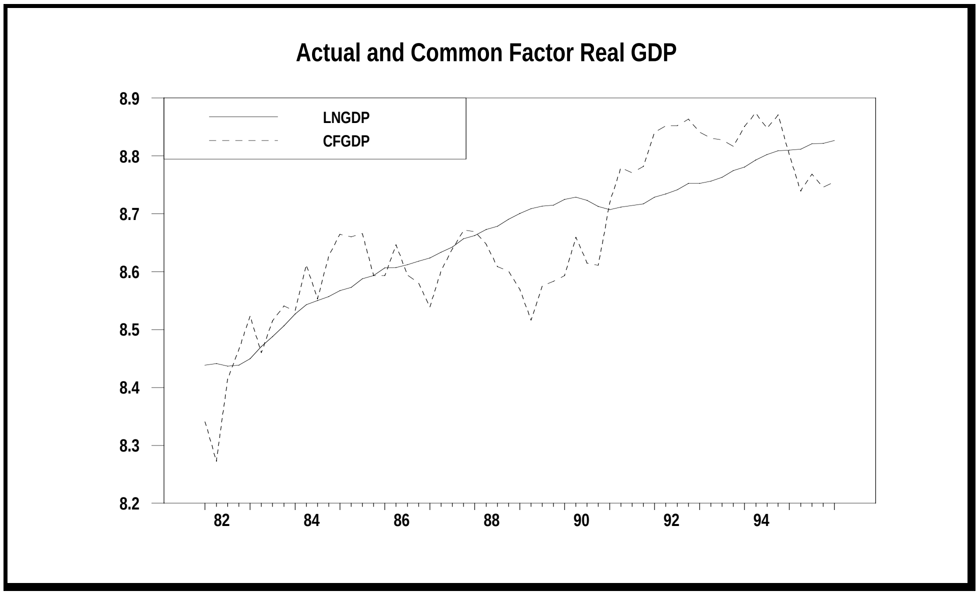


Figure 22

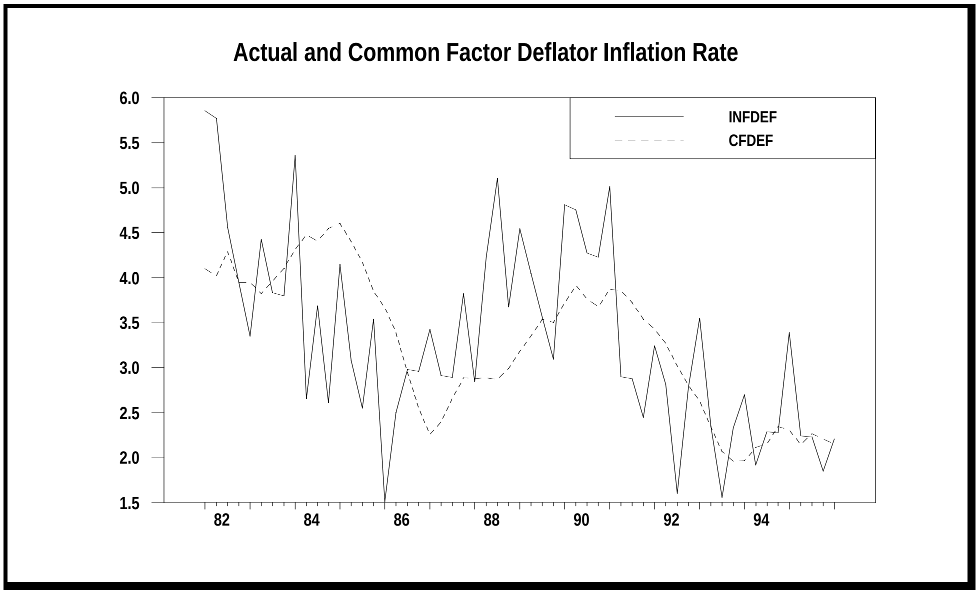


Figure 23

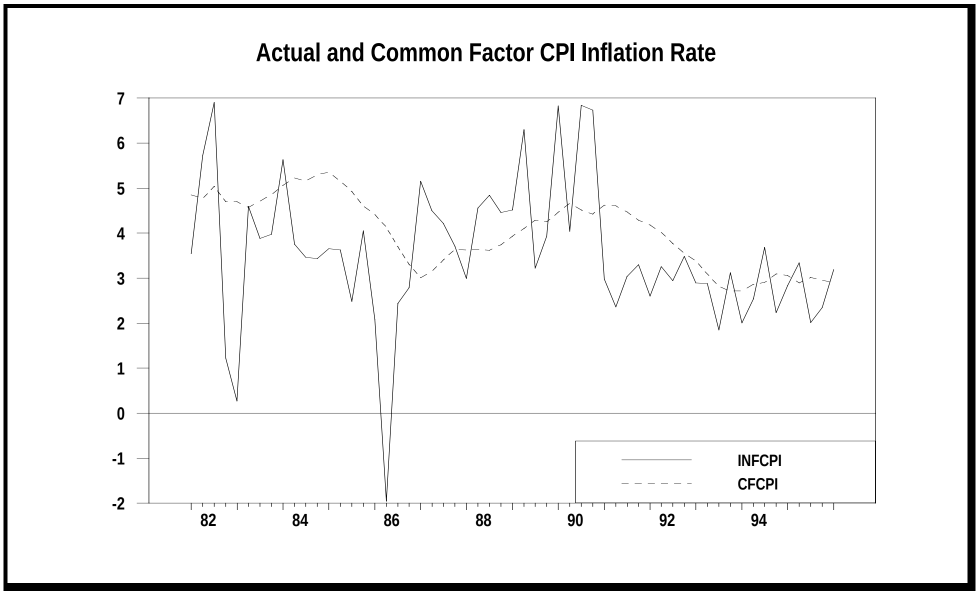


Figure 24

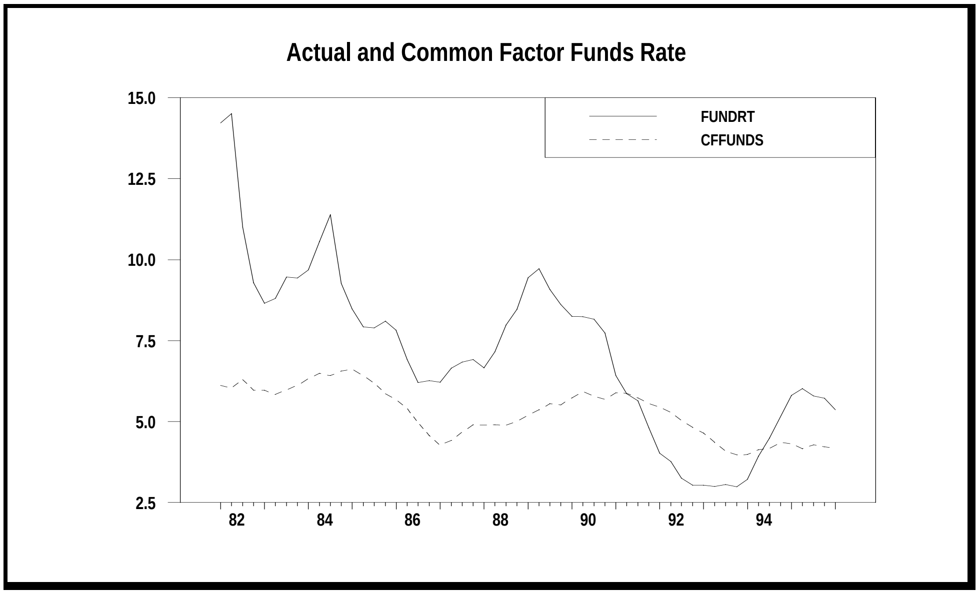


Figure 25

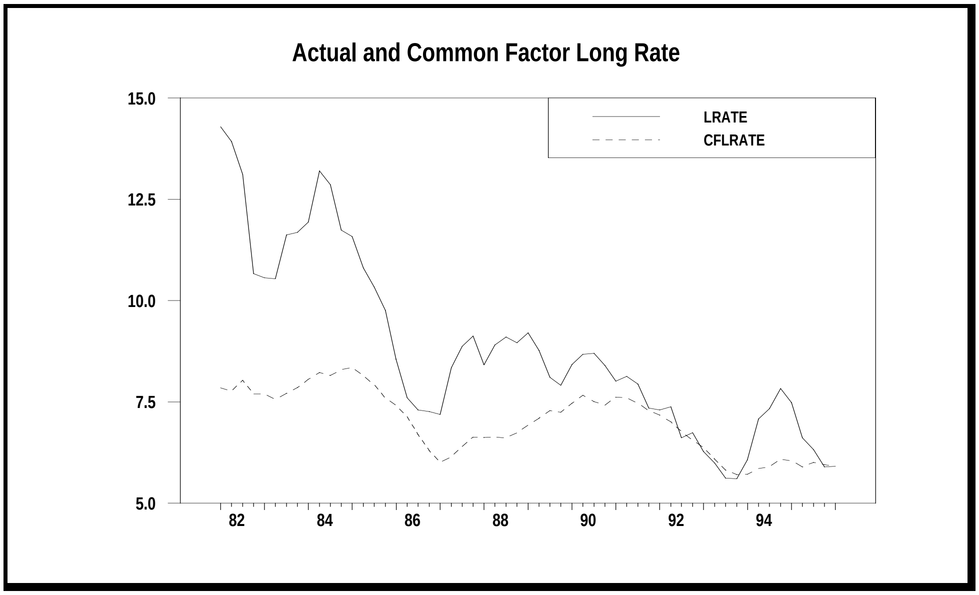
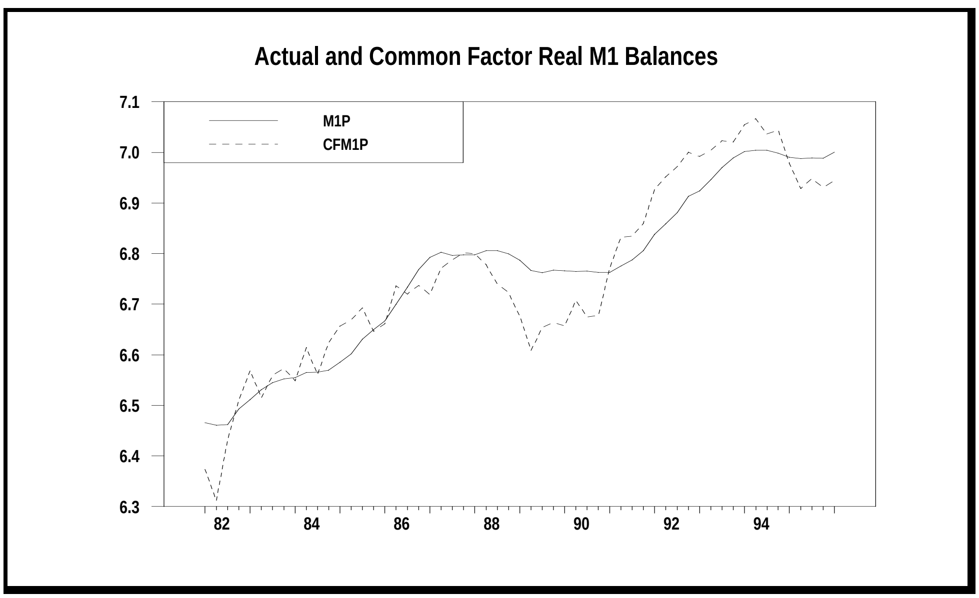


Figure 26



of the sample than is the "permanent component". Given the similarity of the "common factor" and permanent components of the inflation rates, the two measures of the permanent part of the interest rate series are also quite similar. Neither the "common factors" nor the "permanent components" track the time series of actual nominal interest rates with any precision over the last 15 years, and hence the transitory components of the interest rate series show very high serial correlation.

6. Weak Exogeneity

As noted above, the estimated VECM for the sample period ending in 96:1 is suggestive of a weakly exogenous Federal funds rate.[16] We have investigated the accuracy and robustness of this inference using both Wald and likelihood ratio tests. In the former case, we have tested the significance by excluding the error correction terms from the funds rate equation, conditional on the estimated long-run interest rate semielasticity from the unrestricted system. In the latter case, we have reestimated the entire VECM imposing the zero restrictions on the funds rate row of the α (error correction) matrix in addition to the maintained restrictions on the β (cointegrating vectors) matrix.[17] Table 14 contains the estimated χ^2 statistics for various sample periods ending from 85:4 through 96:1 as well as the estimates of the long-run interest semielasticity parameter with and without the weak exogeneity (α) restrictions. Once the sample period is extended through 88:4 (which, coincidentally, includes the first year of the Greenspan era at the Fed), these tests consistently fail to reject weak exogeneity of the funds rate. Imposition of the weak exogeneity restrictions leaves the estimate of the long-run interest semielasticity parameter virtually unchanged.

The statistical implications of weak exogeneity have been discussed extensively (See for example, Engle, Hendry and Richard (1983), or Ericcson and Irons (1995)). The basic idea is that estimation of cointegrating relations may be confined to a subset of the variables that comprises a vector process if some of the elements of the process are

[16] This contrasts with models that we have estimated previously using either the commercial paper rate or the Treasury bill rate where the error correction coefficients in the interest rate equation have always been significantly different from zero. We have not yet determined whether this difference is characteristic of the funds rate or the higher dimension of the present system.

"weakly exogenous". Policy implications may apply if the weakly exogenous variable(s) are policy control variables.

Table 14

Weak Exogeneity Tests for Funds Rate Equation

End of Sample	β: no α restriction	β: Weak Exogeneity	Wald Test $\chi^2_{(4)}$	LR Test $\chi^2_{(4)}$
96:1	.08494	.08461	6.68	6.53
95:4	.08278	.08243	6.65	6.50
94:4	.08754	.08679	6.01	5.88
93:4	.08923	.08828	5.30	5.45
92:4	.08956	.08851	5.57	5.45
91:4	.08217	.08123	8.53	8.24
90:4	.08388	.08330	8.53	8.25
89:4	.08599	.08563	8.22	7.96
88:4	.09370	.09265	8.31	8.00
87:4	.10669	.10354	12.58	11.54
86:4	.10782	.10677	14.09	13.23
85:4	.11262	.11400	16.20	15.06

Our purpose here is to focus on the economic implications of the resulting equation for the funds rate, since the funds rate is the operating target of Federal Reserve Open Market Policy and has been used as the dependent variable in regressions and/or specifications that purport to be descriptive representations of an implicit monetary policy rule (Bernanke and Blinder, 1992; Taylor, 1993).

One possible interpretation of the VECM with a weakly exogenous funds rate is to treat the funds rate as the variable on the left hand side of the monetary policy rule and to assume no feedback from the economy to policy actions within the period of measurement (quarter). The underlying economic model is then of the form of equations

[17] The likelihood ratio tests were constructed using the CATS in RATS estimation routine.

(1) and (3) in Bernanke and Blinder (1992) and the funds rate equation with weak exogeneity is a marginal submodel of the VECM that can be estimated independently of the rest of the reduced form model.[18] The estimated funds rate equation from the VECM is given in Table 15 for the sample period ending in 96:1. Note that none of the estimated

Table 15

Estimated Funds Rate with Weak Exogeneity
56:2 - 91:1

Const	$\Delta m1p_{-1}$	$\Delta m1p_{-2}$	$\Delta m1p_{-3}$	$\Delta infd_{-1}$	$\Delta infd_{-2}$	$\Delta infd_{-3}$	$\Delta fund_{-1}$	$\Delta fund_{-2}$	$\Delta fund_{-3}$
-.3539	-3.130	-3.635	12.871	.0054	.0297	-.0200	.1992	-.4171	.1971
(-3.61)	(-.35)	(-.35)	(1.55)	((.14)	(.70)	(-.61)	(2.45)	(-4.95)	(2.38)
$\Delta infc_{-1}$	$\Delta infc_{-2}$	$\Delta infc_{-3}$	Δy_{-1}	Δy_{-2}	Δy_{-3}	Δrl_{-1}	Δrl_{-2}	Δrl_{-3}	D82
.0483	.0785	.0629	20.40	18.19	12.68	.3958	-.1869	-.0241	-.0997
(1.24)	(1.95)	(1.70)	(3.01)	(2.67)	(1.80)	(2.91)	(-1.35)	(-.17)	(-85)

coefficients on the lagged changes in real balances or the deflator inflation rate are significant. The estimated coefficients on the lagged CPI inflation rate approach significance (at the 5 percent level) and the estimated coefficients on lagged real growth

[18] It should be noted that Bernanke and Blinder estimate their models on monthly data. At that time interval, the assumption that the implicit policy rule does not involve reactions to contemporaneous information on the state of the economy is undoubtedly quite accurate, given the lags in availability of information about the economy. At quarterly frequency such as we are using here and with the incorporation of term structure information into the model, the information lag argument for the identifying restrictions becomes less plausible.

Table 16

Recursive Estimates of Restricted Funds Rate Equation, 1972 - 96

(Omitting 79:4 - 81:4)

	Const	D82	ΔRff_{-1}	ΔRff_{-2}	ΔRff_{-3}	$\Delta infc_{-3}$	$\Delta infc_{-2}$	$\Delta infc_{-3}$	Δlny_{-1}	Δlny_{-2}	Δlny_{-3}	ΔRl_{-1}
96:1	-0.41	-0.00	0.16	-0.41	0.18	0.06	0.10	0.07	24.61	19.72	12.05	0.39
95:4	-0.41	-0.00	0.16	-0.41	0.18	0.06	0.10	0.07	24.59	19.77	11.97	0.39
94:4	-0.42	-0.03	0.13	-0.42	0.16	0.06	0.11	0.08	24.75	20.40	12.71	0.42
93:4	-0.42	-0.07	0.13	-0.43	0.16	0.07	0.11	0.08	24.52	20.07	13.34	0.40
92:4	-0.43	-0.09	0.11	-0.43	0.15	0.07	0.11	0.08	24.90	20.45	13.65	0.42
91:4	-0.43	-0.07	0.10	-0.44	0.15	0.07	0.11	0.09	25.01	20.88	13.40	0.44
90:4	-0.44	-0.08	0.11	-0.44	0.14	0.08	0.12	0.08	24.40	21.77	14.43	0.42
89:4	-0.44	-0.06	0.11	-0.45	0.14	0.08	0.12	0.09	24.08	21.76	14.47	0.43
88:4	-0.44	-0.11	0.09	-0.46	0.13	0.09	0.13	0.08	23.23	22.65	15.38	0.44
87:4	-0.45	-0.19	0.07	-0.46	0.12	0.09	0.13	0.09	23.84	22.14	16.16	0.46
86:4	-0.45	-0.16	0.07	-0.46	0.12	0.08	0.13	0.08	23.90	21.93	15.78	0.50
85:4	-0.45	-0.21	0.07	-0.46	0.11	0.08	0.14	0.08	24.12	21.86	15.48	0.51
84:4	-0.45	-0.20	0.09	-0.52	0.16	0.09	0.14	0.08	24.55	22.06	14.55	0.57
83:4	-0.48	0.10	0.11	-0.48	0.18	0.08	0.12	0.09	25.01	22.06	16.81	0.56
82:4	-0.51	0.37	0.11	-0.50	0.18	0.07	0.19	0.11	25.19	26.16	15.23	0.70
79:3	-0.44		0.27	-0.55	0.15	0.13	0.16	0.11	28.17	14.50	15.77	0.48
78:4	-0.47		0.27	-0.56	0.13	0.13	0.17	0.12	28.56	14.49	17.99	0.47
77:4	-0.48		0.28	-0.58	0.12	0.14	0.17	0.13	32.08	11.70	17.06	0.40
76:4	-0.48		0.27	-0.59	0.12	0.16	0.18	0.13	32.76	11.33	16.61	0.40
75:4	-0.49		0.26	-0.59	0.12	0.16	0.19	0.14	33.99	10.96	16.71	0.38
74:4	-0.45		0.12	-0.47	0.19	0.21	0.23	0.12	36.43	13.03	8.62	0.40
73:4	-0.37		0.11	-0.23	-0.05	0.18	0.19	0.12	33.00	8.94	6.83	0.49
72:4	-0.31		0.31	-0.38	0.05	0.09	0.09	0.09	26.96	-0.31	12.43	0.47

are highly significant.[19] The estimated coefficient on the first lagged change in the long rate is significant as are all the estimated coefficients in the distributed lag on the funds rate. This can be interpreted under the Bernanke-Blinder identifying restrictions of their equations (1) and (3) as a policy rule in which the Fed gradually moves the funds rate target up(down) in response to observed acceleration(deceleration) in inflation, observed real growth above(below) deterministic drift, or observed increases(decreases) in the slope of the yield curve (spread between the long-term rate and the funds rate). Other than the differencing of the variables to account for nonstationarity, this equation is similar in structure to the specification utilized by Bernanke and Blinder in section II of their analysis.

Though not completely in the spirit of VAR analysis, which emphasizes unrestricted lag structures, we have reestimated the funds rate equation omitting the regressors whose estimated coefficients are not significant in the specification in Table 15. Recursive estimates of this equation are shown in Table 16. With the exception of the very shortest samples, which span only 15-16 years, this specification appears to be remarkably robust, particularly in light of the well known instabilities in the estimated monetary policy reaction functions that appear in the literature. It is likely that this reflects the first differences applied to the specification here. Typical monetary policy reaction function or policy rule equations are specified in levels. If the variables in the specification are truly nonstationary, such "spurious regressions" are unlikely to be robust across different samples. Nevertheless, the specification estimated here is not a particularly accurate representation of the data generating process for changes in the funds rate. The adjusted R^2 of the equation in Table 15 is only .41 and the estimated standard error of the residuals is 63 basis points.

[19] Note that the interest rates and inflation rates are expressed in percentage changes at annual rates, while real growth is measured at the first (quarterly) difference of the log of real GDP. Hence to interpret the estimated coefficients in Table 15, those on real growth should be divided by 400 to convert all variables into percentage changes at annual rates.

Consider an economic model of the form of equations (1) and (3) in Bernanke and Blinder (1992), extended to include cointegration and weak exogeneity of the policy variables, P_t.[20]

(3) $$\Delta P_t = D_1 \Delta Y_{t-1} + G \Delta P_{t-1} + v_t$$

(1) $$\Delta Y_t = B_0 \Delta Y_t + C_0 \Delta P_t + B_1 \Delta Y_{t-1} + C_1 \Delta P_{t-1} + F \begin{bmatrix} P_{t-1} \\ Y_{t-1} \end{bmatrix} + u_t$$

For future reference, we can write the reduced form conditional submodel of the Y vector as:

$$\Delta Y_t = (I - B_0)^{-1} C_0 \Delta P_t + (I - B_0)^{-1} B_1 \Delta Y_{t-1} + (I - B_0)^{-1} C_1 \Delta P_{t-1} + (I - B_0)^{-1} F \begin{bmatrix} P_{t-1} \\ Y_{t-1} \end{bmatrix} + (I - B_0)^{-1} u_t$$

Note that under this assumed structure, the parameters of the reduced form conditional submodel are not functions of the parameters of the policy rule equation (3). The reduced form of equations (1) and (3) is:

$$\begin{bmatrix} \Delta P_t \\ \Delta Y_t \end{bmatrix} = \begin{bmatrix} D_1 & G \\ C_0 (I - B_0)^{-1} D_1 + (I - B_0)^{-1} B_1 & C_0 (I - B_0)^{-1} G + (I - B_0)^{-1} C_1 \end{bmatrix} \begin{bmatrix} \Delta P_{t-1} \\ \Delta Y_{t-1} \end{bmatrix} +$$

$$\begin{bmatrix} 0 \\ C_0 (I - B_0)^{-1} F \end{bmatrix} \begin{bmatrix} P_{t-1} \\ Y_{t-1} \end{bmatrix} + \begin{bmatrix} v_t \\ C_0 (I - B_0)^{-1} v_t + u_t \end{bmatrix}$$

Assume that $E[v_t u_t] = 0$ as typical in VAR analysis. Then the covariance matrix of the residuals of the reduced form model is:

$$\Sigma_\varepsilon = \begin{bmatrix} \sigma_1 & \sigma_1 (I - B_0)^{-1'} C_0' \\ C_0 (I - B_0)^{-1} \sigma_1 & C_0 (I - B_0)^{-1} \sigma_1 (I - B_0)^{-1'} C_0' + (I - B_0)^{-1} \sigma_2 (I - B_0)^{-1'} \end{bmatrix}$$

where σ_1 and σ_2 are the covariance matricies of u_t and v_t respectively. Note that $\sigma_{\varepsilon_1 \varepsilon_2} \sigma_{\varepsilon_1}^{-1} = C_0 (I - B_0)^{-1}$. Therefore the covariance restriction on v_t and u_t is sufficient to identify the parameters of the reduced form conditional submodel: $(I-B_0)^{-1}B_1$ and $(I-B_0)^{-1}C_1$, from the reduced form model. Also note that

[20] In this example, as in Bernanke and Blinder, we consider only first order models. The analysis is not affected by this restriction.

$$\sigma_{\varepsilon_2} - \sigma_{\varepsilon_2\varepsilon_1}\sigma_{\varepsilon_1}^{-1}\sigma_{\varepsilon_1\varepsilon_2} = (I - B_0)^{-1}\sigma_2 (I - B_0)^{-1'}$$

which is the covariance matrix of the residuals of the reduced form conditional submodel.

With these parameters identified, it is possible to augment the reduced form conditional submodel with alternative specifications of the policy rule that preserve both the information lag in (3) and the weak exogeneity of the funds rate, and to construct counterfactual policy analyses. Such exercises are, of course, subject to the proviso that the structure of equation (1) is invariant to the alterations in the policy rule.

Additional identifying restrictions can be placed on the matrices $(I-B_0)$ or $(I-B_0)^{-1}$ to extract the individual u_t shocks from the residual matrix of the reduced form conditional submodel, but such information is not required to determine the effects of alternative policy rules given the restrictions already imposed on the system.

7. Innovation Analysis and Impulse Response Functions

The VAR methodology for analyzing the effect of disturbances to the various variables dates from Sims (1980). He states:

> "The 'typical shocks' whose effects we are about to discuss are positive residuals of one standard deviation unit in each equation of the system. The residual in the money equation, for example, is sometimes referred to as the 'money innovation' since it is that component of money which is 'new' in the sense of not being predicted from past values of the variables in the system. The residuals are correlated across equations. In order to be able to see the distinct patterns of movement the system may display it is therefore useful to transform them to orthogonal form. There is no unique best way to do this. What I have done is to triangularize the system with variables ordered as M,Y,U,W,P and PM. Thus the residuals whose values of other variables enter the right-hand-sides of the regressions with a triangular array of coefficients." (p. 21)

The standard approach to construction of the "typical shocks" as defined by Sims was to rearrange the elements of the data vector in the VAR and then decompose the estimated covariance matrix of the VAR residuals, Σ as TDT' using the Cholesky technique , where T is a lower triangular matrix normalized to 1.0 on the principal diagonal and D is a positive definite diagonal matrix.[21] The element t_{ij} of the T matrix (i>j) then measures the impact effect (or impact multiplier) of the j^{th} shock on the i^{th} variable in the VAR.

Frequently, in applications of the Sims approach, it is not acknowledged that Sims' scheme for constructing "typical shocks" is a particular application of a causal chain structure for the identification of economic (or "structural") models first proposed by Wold (1954). In particular, the Sims approach is equivalent to estimating the simultaneous equations model:

$$R(WX_t) = \sum_{i=1}^{k} \Gamma_i (WX_{t-i}) + \varepsilon_t$$

subject to the exactly identifying restrictions that R is lower triangular and the covariance matrix of ε_t is diagonal. W is the permutation matrix that establishes the desired order of the variables in the X_t vector.

It is well known from the literature on simultaneous equation estimation that the system so identified can be estimated by OLS and that this estimator is the FIML estimator of this structure (e.g. Theil (1971), sections 9.6 and 10.7). OLS applied equation-by-equation to the above system will produce an estimate of the R matrix which is the inverse of the T matrix estimated by the Cholesky decomposition approach.[22]

Sims (1980) was extremely critical of the then prevalent approach to identifying large scale econometric models by imposing "incredible" restrictions on the distributed lags of such models, preferring instead to identify models with an approach that links fundamental innovations to distinct variables in a vector process. Of course this mechanism embodies its own "structure" and requires its own identifying restrictions.[23] The identification problem is pervasive and certainly does not vanish when the focus is shifted from structural behavioral relations to structural relations among fundamental innovations. The number of interdependent pieces of information needed to obtain identification is *exactly the same* in either case.

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[21] The process of reordering the elements of the data vector, X_t of the VAR is equivalent to premultiplying the entire VAR by a permutation matrix W, or alternatively premultiplying the covariance matrix of the VAR residuals by W and postmultiplying the covariance matrix by W'.

[22] Altenatively OLS can be applied equation-by-equation to estimate the system $R(WZ_t)= \varepsilon_t$ where Z_t is the vector of estimated VAR residuals.

[23] Interestingly a heated controversy raged in the econometrics literature of the late 1950s and early 1960s over the credibility of the Wold causal chain indentification scheme.

Subsequently, alternative identifying schemes were applied to determine the “typical shocks” of the innovation analyses (Bernanke, 1986; Sims, 1986). These identification schemes retained the assumption of independent “shocks”, but proposed nontriangular decompositions of the VAR residual covariance matrix in order to determine the “typical shocks”. Models so identified have become known as structural VARs. Usually a sufficient number of restrictions were imposed on the impact matrix of the shocks to achieve exact identification. Implicit in each of those sets of identifying restrictions is a corresponding set of exactly identifying restrictions on the contemporaneous interactions in a conventional simultaneous equations model that contains, in each equation, unrestricted distributed lags of order k. Hence, the “structural VAR” approach is equivalent to the estimation of the simultaneous equation model:

$$RX_t = \sum_{i=1}^{k} \Gamma_i X_{t-i} + \varepsilon_t$$

where the Γ_i matrices are unrestricted, the covariance matrix of the ε_t is restricted to a diagonal matrix, and sufficient additional restrictions are imposed on the R matrix to achieve exact identification of the system. The restrictions on R will be more complex that the typical exclusion restrictions discussed in textbook presentations of the simultaneous equation identification problem because each element of the R matrix is a function of all the elements of the matrix of impact coefficients (the inverse of the R matrix).

Recent developments in “structural VAR modeling” have extended the approach in several directions. Gianinni (1992) discusses identification and FIML estimation of three different types of “structural VAR models”: a K model, C model and AB model. The K model is the traditional simultaneous equations approach to identification with diagonality restrictions on the covariance matrix of the “shocks” and linear restrictions on the simultaneous interaction (slope coefficients) of the variables in the VAR (linear restrictions on the R matrix in the above notation). The C model is the Bernanke-Sims approach to identification through diagonality restrictions on the covariance matrix of the “shocks” plus linear restrictions on the matrix of impact effects (multipliers) of the shocks (linear restrictions on R^{-1} in the above notation). The AB model is a hybrid specification that allows combinations of K model and C model type restrictions. All three classes of models can be specified with overidentifying restrictions, and the overidentifying restrictions can be subjected to statistical testing. A RATS procedure, SVAR, is available to perform FIML estimation on such (over)identified models.[24]

The application of innovation analysis and the construction of impulse response functions from Vector Error Correction models is not as advanced as "structural VAR" modeling. In existing applications to these models, the identification is performed within the framework of the Wold causal chain structure that characterized the original Sims work on VARs. Pioneer applications to this type of model are King, Plosser, Stock and Watson (1991) and Warne (1993). The former dealt exclusively with the identification of the dynamic effects of permanent shocks. The latter extended the analysis to consider both permanent and transitory shocks. In these studies, all individual "shocks", whether permanent or transitory, are assumed to be independent. Exact identification is achieved by assuming a block causal ordering in which the permanent shocks are assumed to precede the transitory shocks. Within these two groups, individual shocks are identified with a triangular causal chain structure. The common trends model is overidentified, but KPSW's construction of impulse response functions is derived from a set of exactly identifying restriction, and they do not test nor impose the overidentifying restrictions in their empirical analysis. In principle the innovation analysis of VECMs can be developed under a more general set of identifying or overidentifying restrictions along the lines of the Giannini K, C or AB models, though there does not appear to be such applications in the existing literature. Hoffman and Rasche (1996a) test the overidentifying restrictions among the two common trends in their VECMs involving money demand functions, and where appropriate estimate impulse response functions subject to these overidentifying restrictions. However, these overidentified structures are restrictions imposed within the Wold causal chain framework. The application of more general identifying (or overidentifying) restrictions in a 'structural VECM' framework in principle can lead to the isolation of historical policy rules in the data, and allow for the simulation of models that are augmented by counterfactual policy rules (see Rasche (1995)).

[24] Gali (1992) produces a model with a mixture of identifying restrictions characteristic of a Giannini AB

We have applied the KPSW-Warne-Sims-Wold causal chain identification scheme to derive impulse response functions for the six variable VECM developed above, both with and without the restriction that the funds rate is weakly exogenous. The impulse response functions for a permanent nominal shock, derived from the models estimated through 96:1, are shown in Figures 27-38. In the identification scheme for these models, the permanent nominal shock is ordered first in the Wold causal chain structure. This shock is normalized to have a 1.0 long-run (steady-state) equilibrium impact on both measures of inflation and both nominal interest rates. In principle this shock should have a zero effect on real output, but the sample covariance matrix of the two permanent shocks is not perfectly diagonal as implied by the KPSW common trends hypothesis, so the long-run effect of this shock on real output (~ .011) reflects the insignificant sample covariance between the permanent shocks. We have not imposed the overidentifying restriction that the covariance between the two permanent shocks be zero, even though this restriction is not rejected by the data.[25]

Several things are noteworthy in the graphs of the impulse response functions. First, it is important to remember that the cointegrating vectors impose certain restrictions on the long-run properties of these impulse response functions, but the adjustment patterns and the length of time that it takes for the responses to attain the equilibrium values (constrained or unconstrained) is determined completely by the data through the unrestricted VAR part of the specification. The effect of the permanent nominal shock on real output, velocity, and both inflation rates reaches the steady-state value over very short horizons, eight to 12 quarters. The short-run effect of real output to a positive permanent nominal shock is positive, consistent with macroeconomic models that postulate price or nominal wage "stickiness", but the patterns in Figure 29 and 30 are not consistent with those implied by traditional macroeconomic models of the U.S. economy which generate extremely long periods of adjustment. The response pattern of CPI inflation exhibits a great deal of fluctuation at short horizons, but our experience with computing precision measures for such impulse response functions in previous, smaller

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type model.

[25] Estimation of the impulse response functions under this restriction is a straightforward application of Giannini's SVAR program.

models suggests that these responses are very imprecise. One interpretation consistent with Figure 30 is that the CPI inflation rate effectively jumps to the steady-state increase of 1.0 immediately. Alternatively, Figure 30 could be interpreted as a gradual increase of the CPI inflation rate to the new steady-state rate over a period of 2-3 years. The available data probably are not capable of discriminating between these two alternatives. In any event, any adjustment period of inflation appears to be completed within a three year horizon.

The patterns in Figures 31-32 suggest that the adjustment of nominal interest rates to the steady-state change of 1.00 is accomplished only over a very long horizon when the funds rate is not constrained to be weakly exogenous. With the weak exogeniety constraint imposed, the funds rate approaches its steady-state value in about two years, while the long rate takes about three years. Neither response function exhibits a negative impact or short-run effect with a permanent positive increase in the inflation rate. In contrast to these graphs, the impulse response of the interest rate spread plotted in Figure 33 has completely died out after 10-12 quarters.

Two of the remaining three cointegrating vectors—corresponding to the Fisher relationship (the negative of the real interest rate reponses) and the differential of the two inflation rates (or, equivalently, to two Fisher relationships)—approach zero quite rapidly (Figures 35-36). The real-balance cointegrating vector appears to approach equilibrium

Figure 27

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

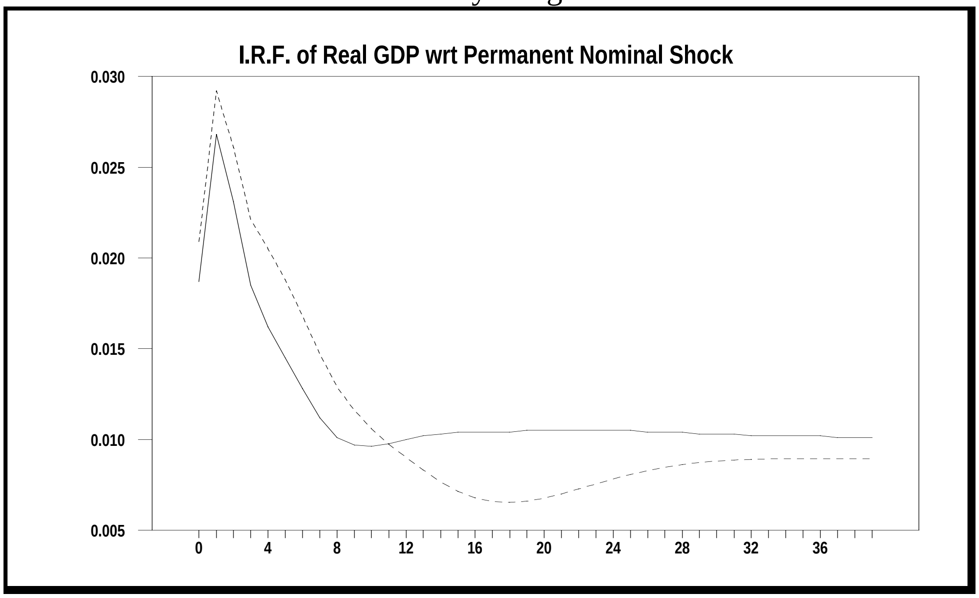


Figure 28

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

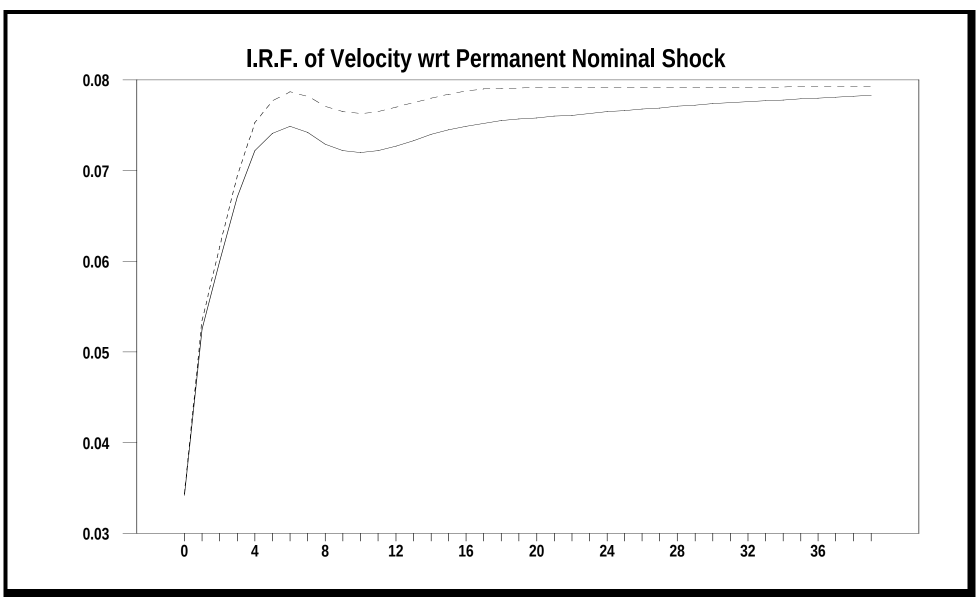


Figure 29

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

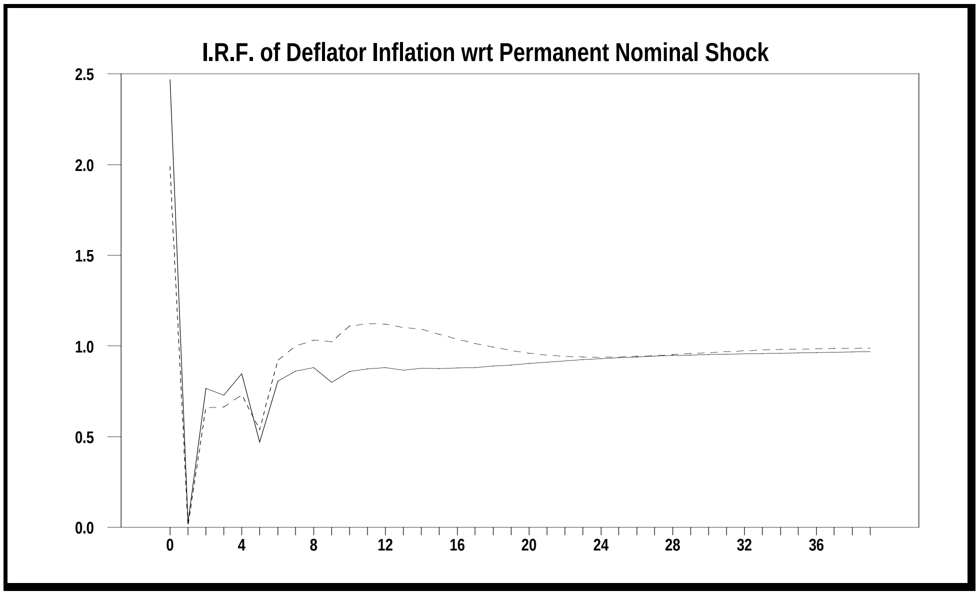


Figure 30

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

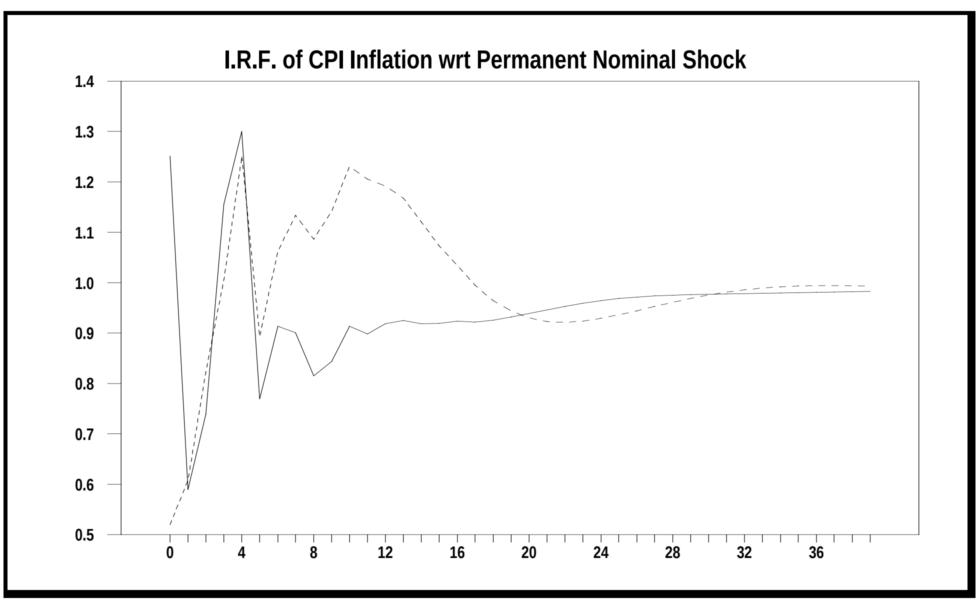


Figure 31

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

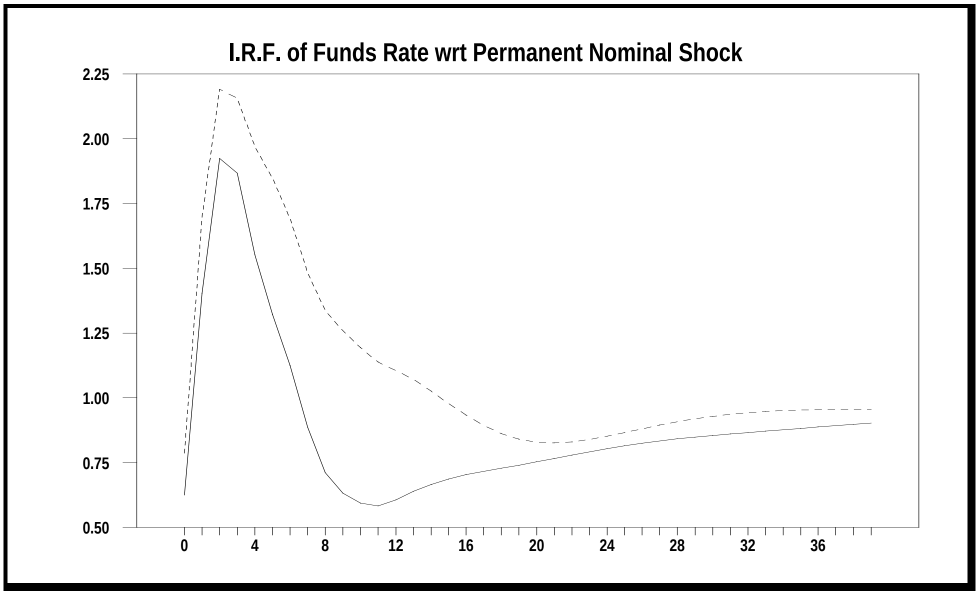


Figure 32

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

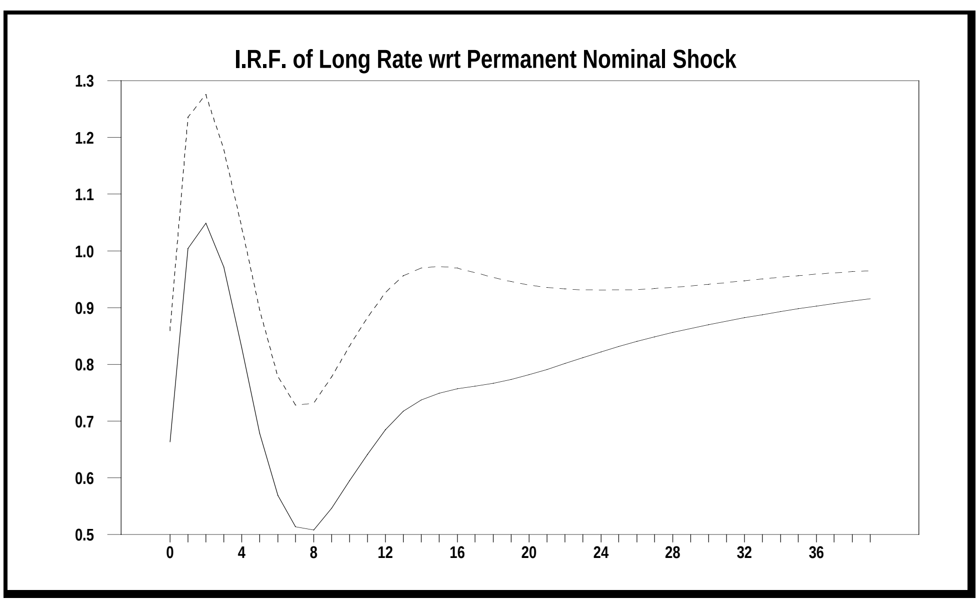


Figure 33

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

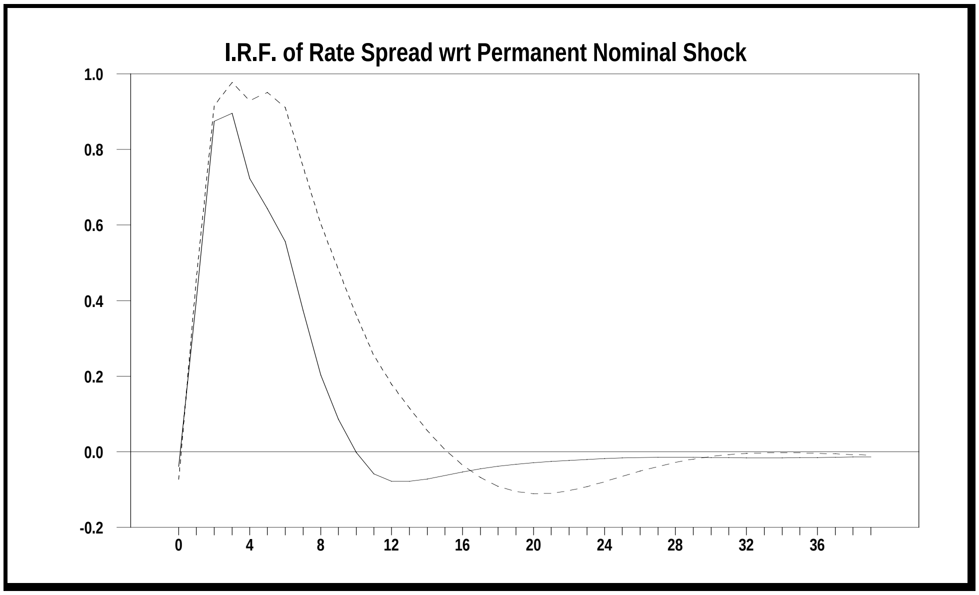


Figure 34

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

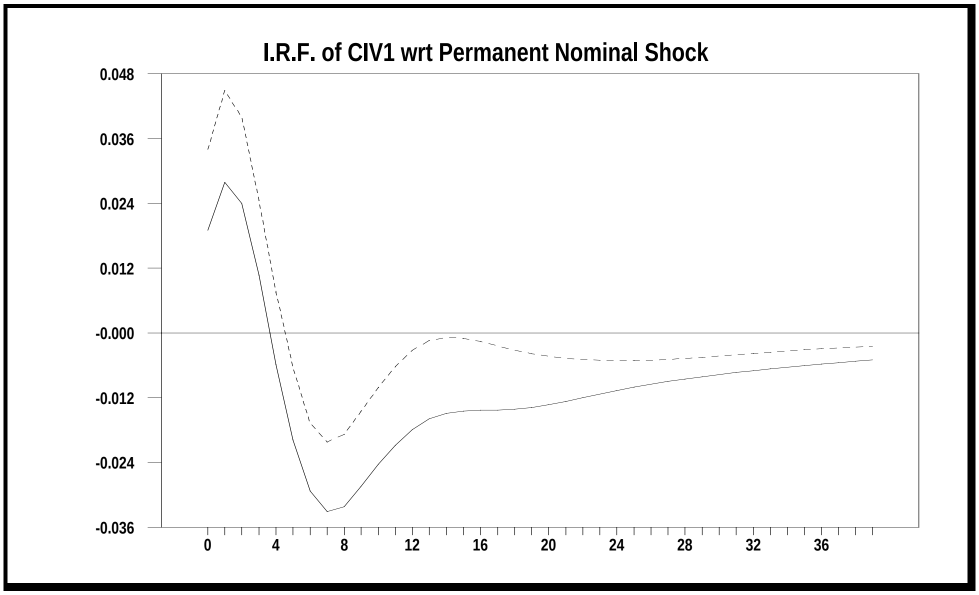


Figure 35

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

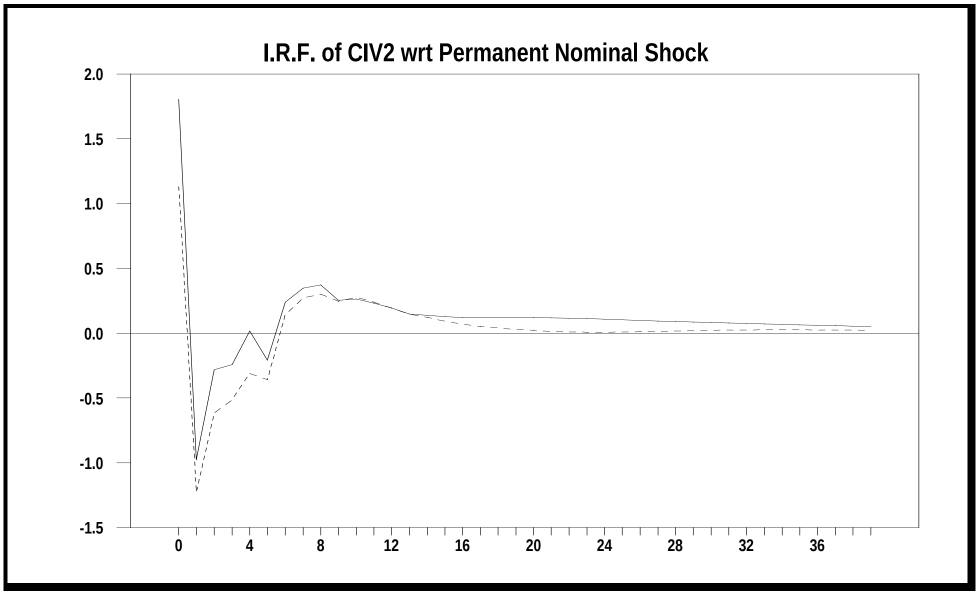


Figure 36

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

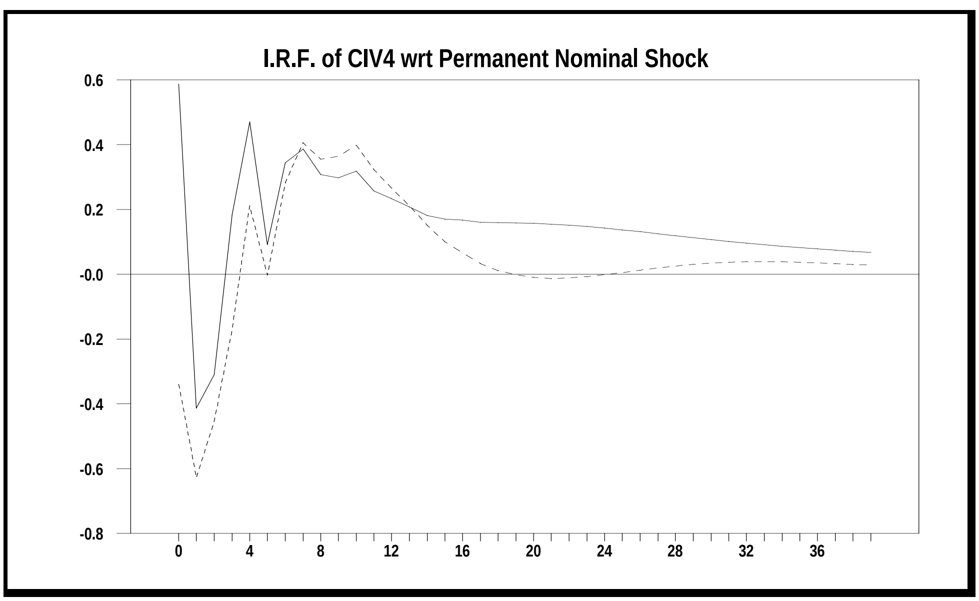


Figure 37

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

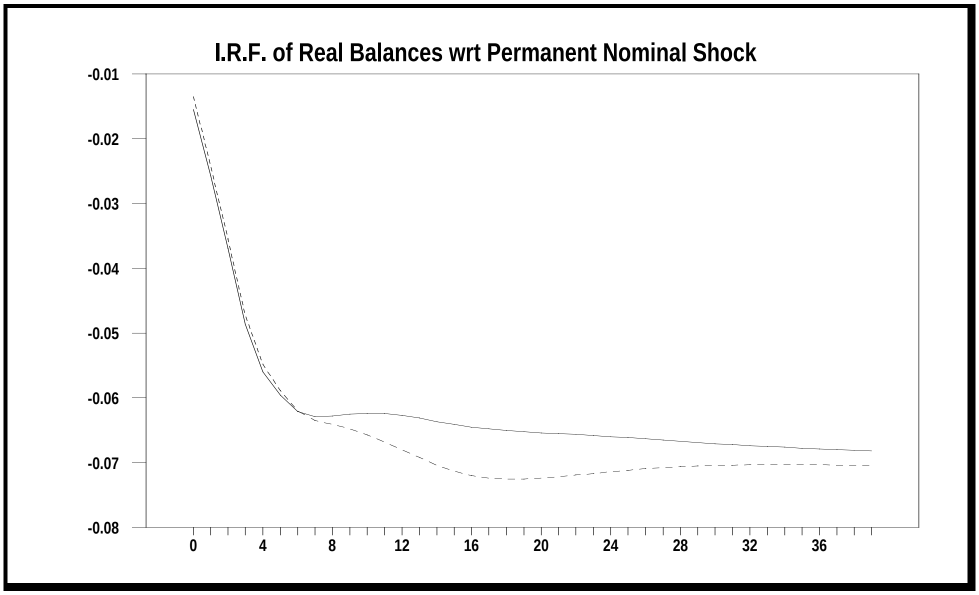
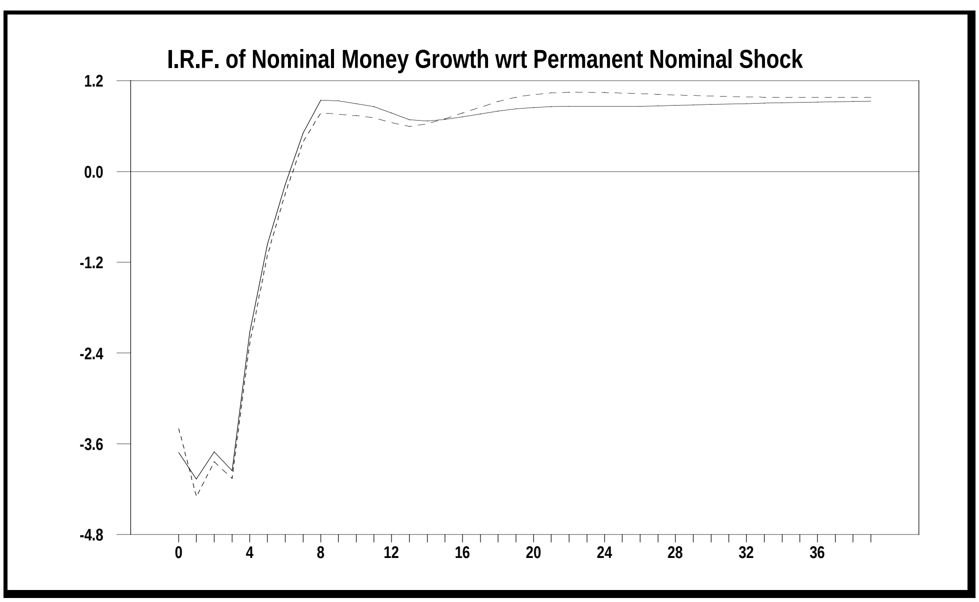


Figure 38

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate



very gradually (Figure 37), but the scale of the graph can be misleading; the deviation from zero is less than one percent after 12 quarters.

The impulse response functions for the real permanent shock are shown in Figures 39-50. The real permanent shock is normalized to have a unitary impact on real output in the steady state. Since this shock is ordered after the nominal permanent shock in the Wold causal chain identification scheme, it cannot have any long run effect on either inflation rate nor on either nominal interest rate. These normalization and identifying restrictions are evident in Figures 39, 41-42, and 50. Since the each of the four cointegrating vectors are stationary, the steady-state effect of any permanent shock on these linear combinations of the variables is zero (Figures 45-48). The impact effect of the permanent real shock on real GDP (Figure 39) is very small < .20). This increases rapidly over a horizon of about two years, overshoots the steady-state value of 1.0 by about 10 percent after about three years, and then settles down very close to the equilibrium effect after about four years. The initial effect of the positive permanent real shock on both inflation rates is substantially negative, consistent with the interpretation of this shock as a technology shock or supply shock. The effects on the inflation rate approach the steady-state values of zero after about three years, but also exhibit a small amount of overshooting of equilibrium for a brief period. The response function for the CPI is relatively smooth, though that of the deflator exhibits some high frequency fluctuations in the initial periods. Our experience with the construction of confidence intervals for such response functions in smaller VECMs with similar cointegrating vectors suggests that such "choppiness" in the impulse response functions is not significant.

The responses of nominal interest rates to this permanent real shock are much more prolonged than either the real output or inflation rate responses (Figures 43-44). It is particularly interesting that the estimated initial effects on short-term and long-term interest rates are in the opposite directions: short-term interest rates move up in response to the positive permanent real shock (which temporarily <u>lowers</u> the observed inflation rate) while long-term interest rates drop in response to this shock in the absence of the weak exogeneity constraint. This result is extremely sensitive, and the initial impulse response function of the long rate is also negative once the constraint is imposed. The initial effect on the funds interest rate is on the order of -40 to -60 basis points. This effect diminishes to zero over a horizon of two years, overshoots the equilibrium by about one-third of the impact effect, and then gradually returns to the steady-state value in the unconstrained specification. Once the weak exogeneity constraint is imposed, the funds rate response function basicly does not exhibit any overshooting.

Figure 39

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

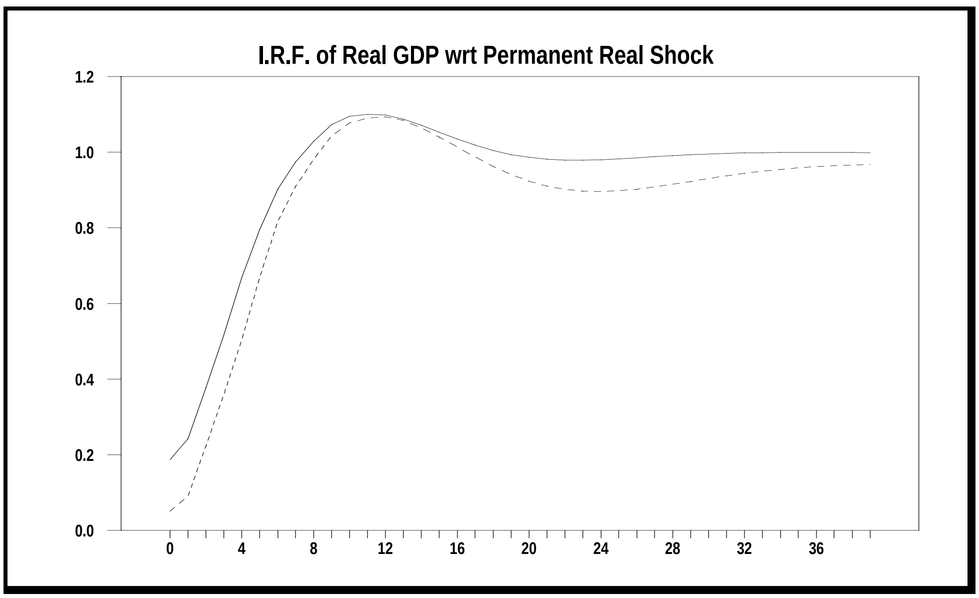
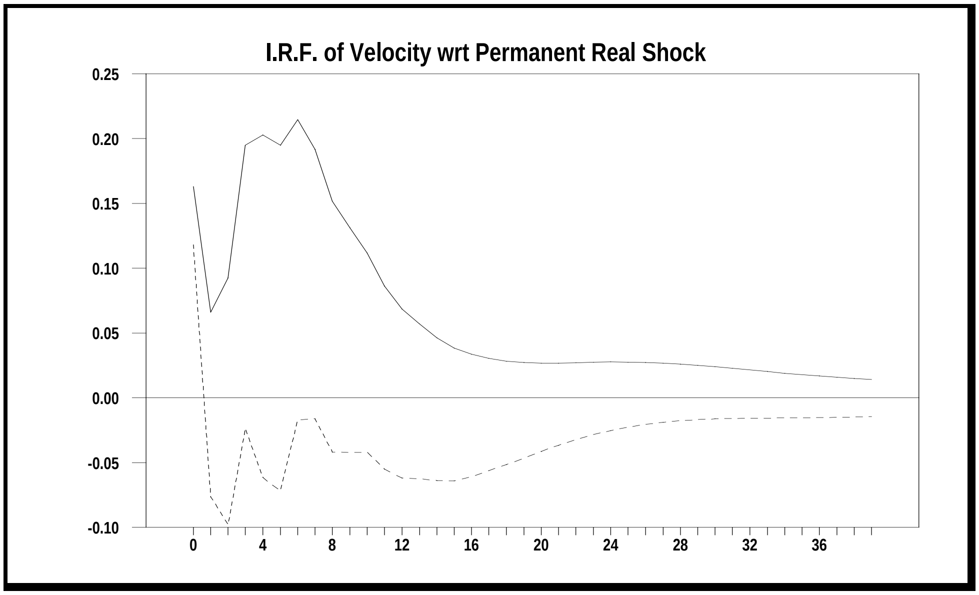


Figure 40

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate



Note: In the above graph, a one percent premanent shock to real GDP is scaled as 1.0 A response of 1.0 represents a unitary elasticity of velocity in response the the permanent real GDP shock.

Figure 41

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

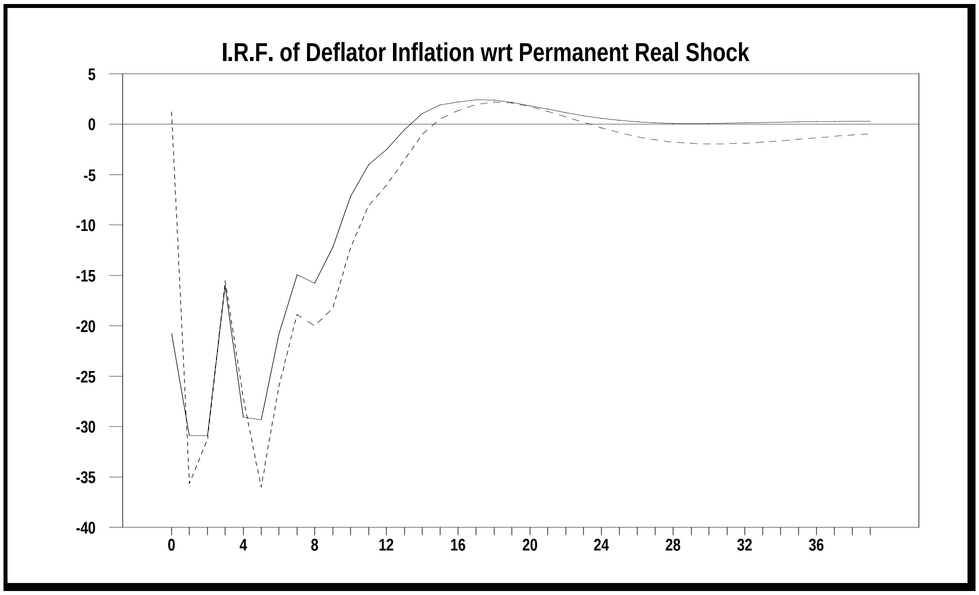
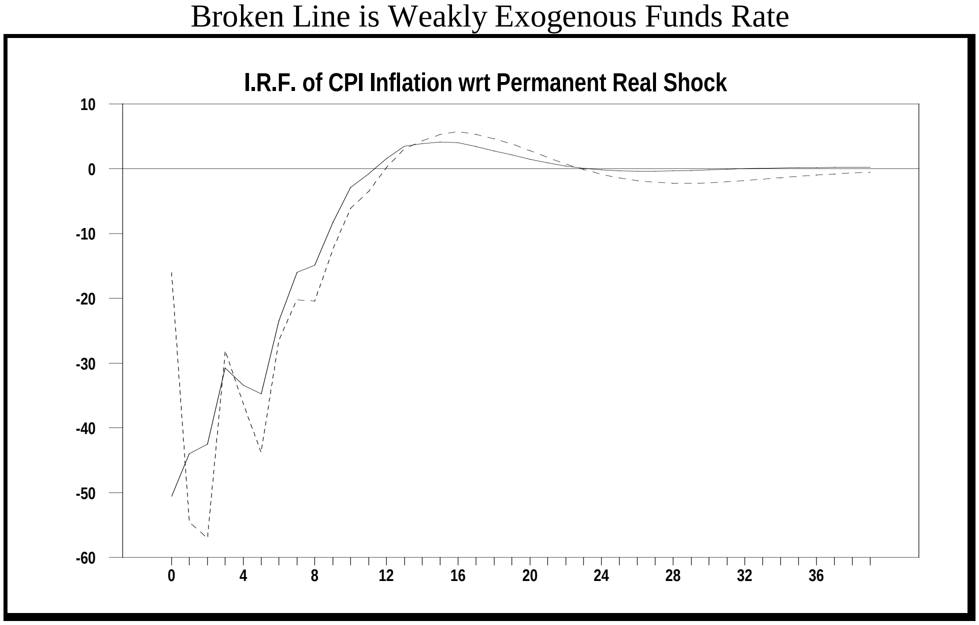


Figure 42

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate



Note: In the above graphs the vertical scale represents the basis point response of the inflation rate to a permanent one percent shock to real GDP.

Figure 43

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

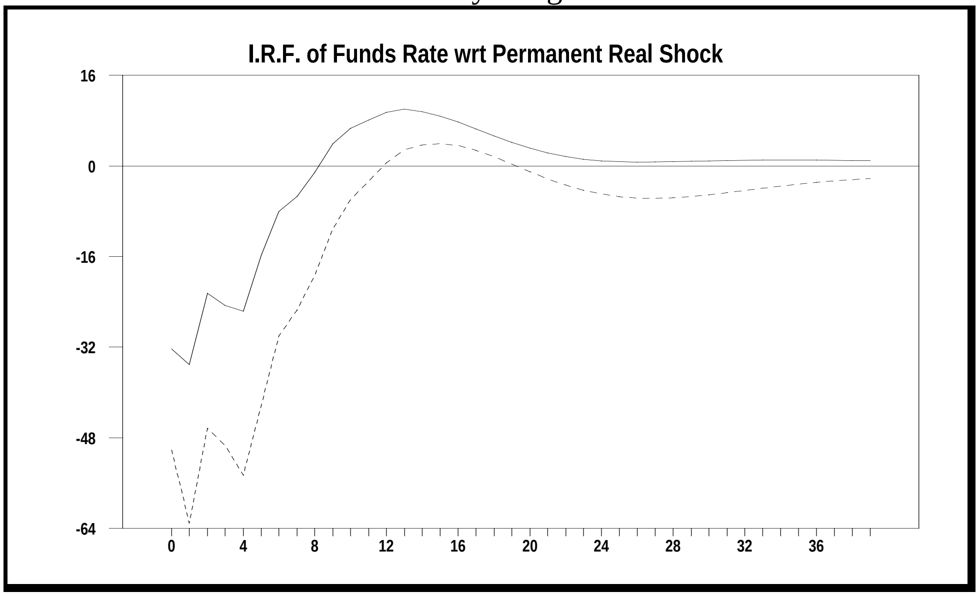
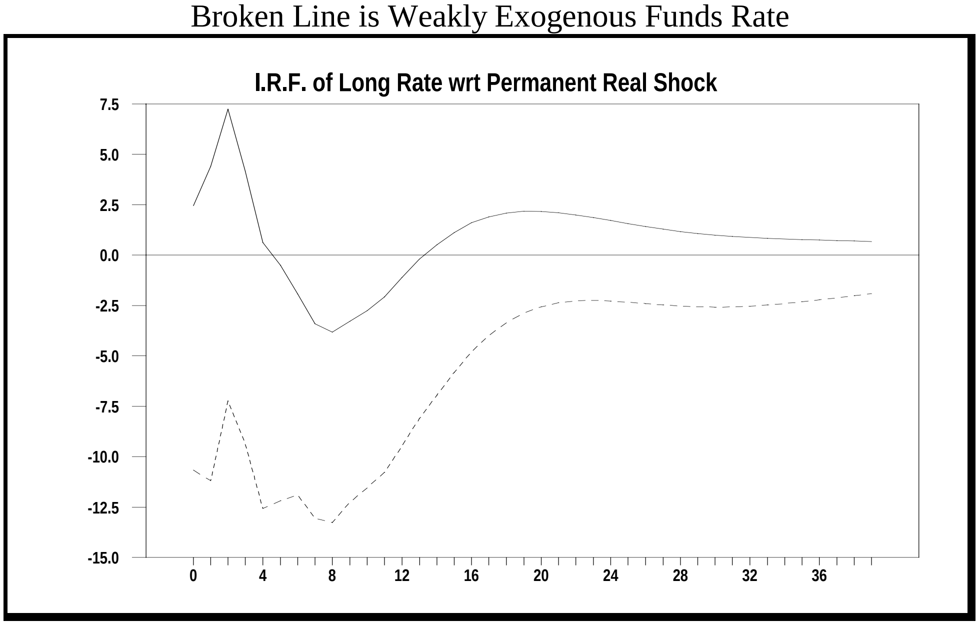


Figure 44

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate



Note: In the above graphs, the vertical scale represents the basis point response of the interest rates to a one percent permanent shock to real GDP.

Figure 45

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

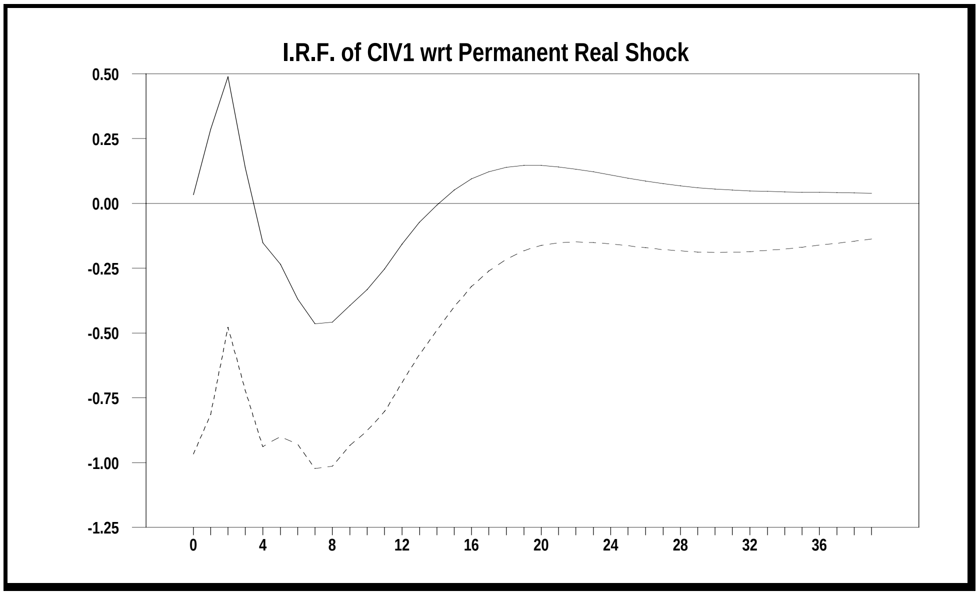
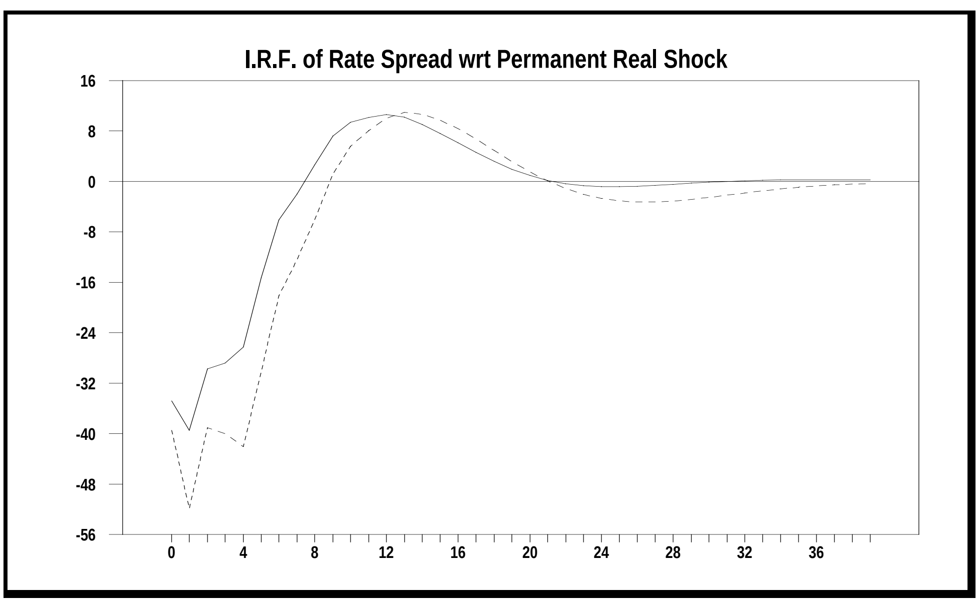


Figure 46

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate



Note: In the above graphs a response of one represents a unitary elasticity of the response of the money demand cointegrating vector to a permanent real GDP shock. The vertical scale on the interest rate graph represents the basis point response to a one percent permanent shock to real GDP.

Figure 47

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

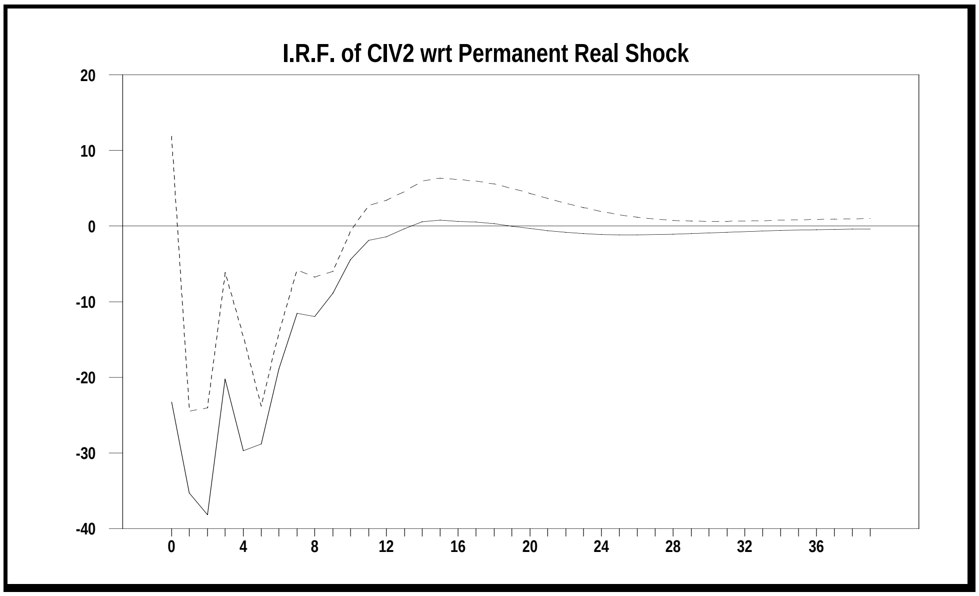
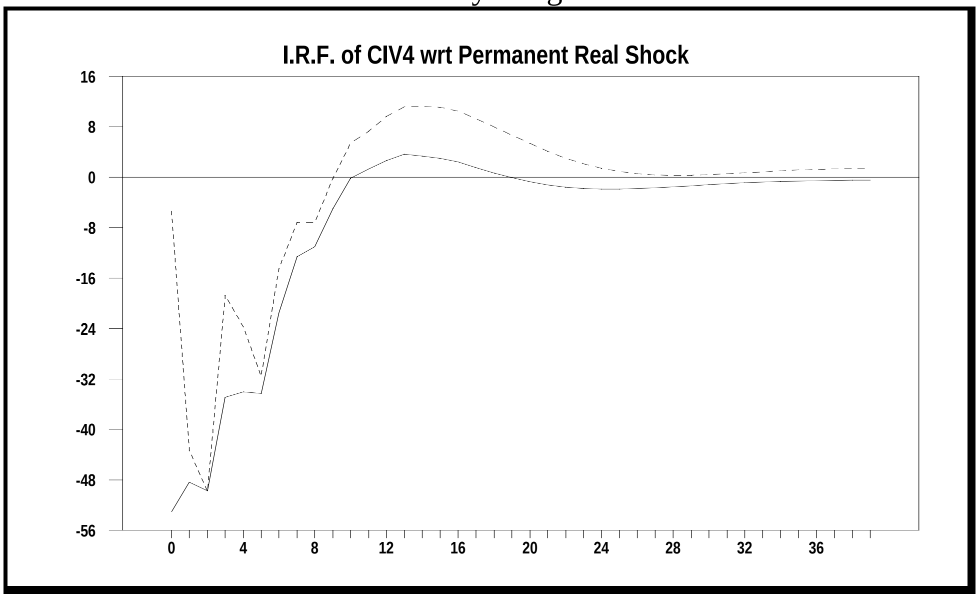


Figure 48

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate



Note: The vertical scale in the above graphs represent the basis point response to a one percent permanent shock to real GDP.

Figure 49

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate

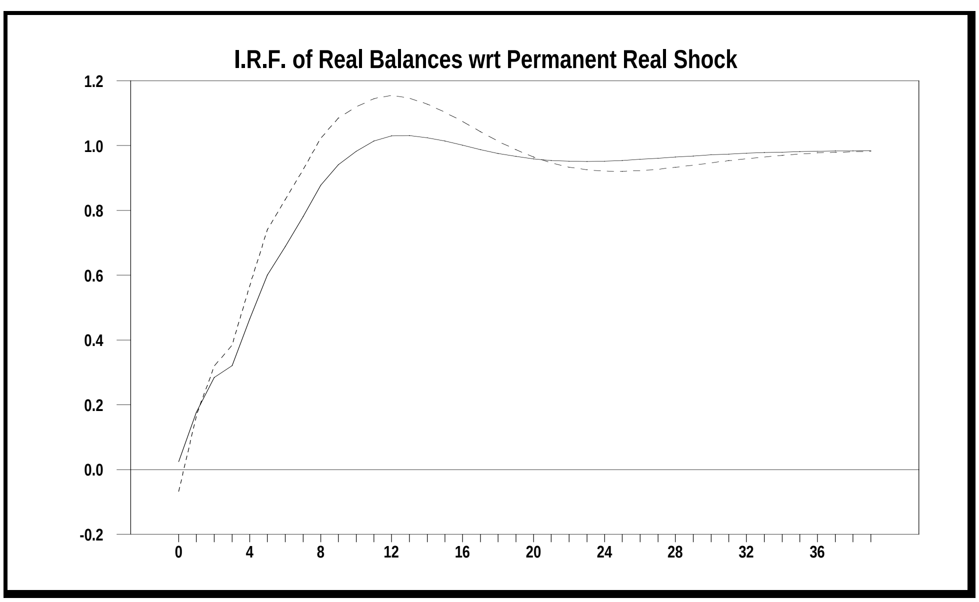
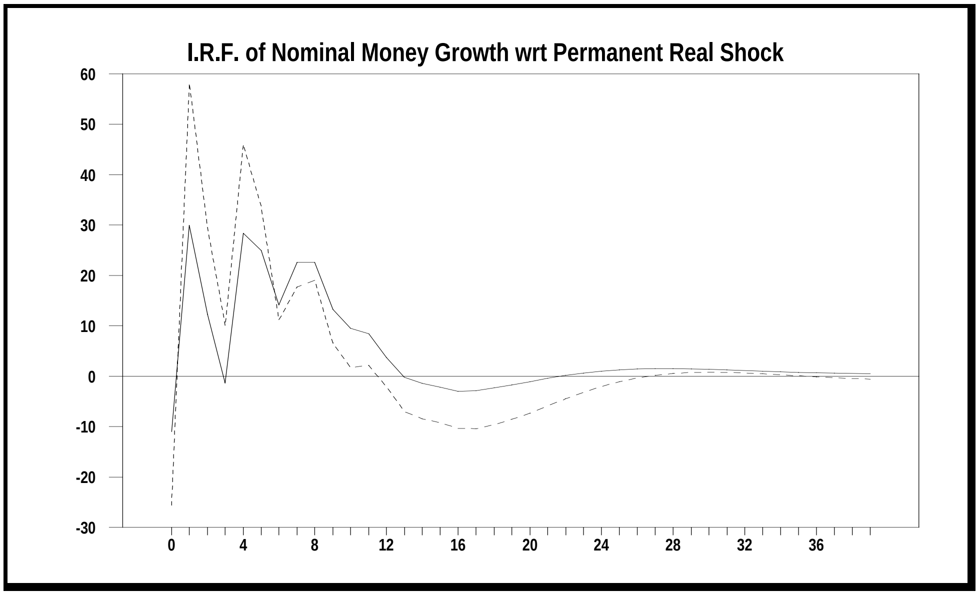


Figure 50

Solid Line without α Constraint;
Broken Line is Weakly Exogenous Funds Rate



Note: In the above graphs a response of one represents a unitary elasticity of the response of the real balances to a permanent real GDP shock. The vertical scale on the money growth graph represents the basis point response to a one percent permanent shock to real GDP.

8. Some Forecasting Results

The following examples illustrate some ex-ante forecasts produced from the VECM model. The first forecast was produced using data, available on August 1, 1996, through the first quarter of 1996. Figure 51 illustrates the actual quarterly data for various series through 96:1 (solid lines) and the forecasts from the model through 97:4. For GDP-deflator and CPI inflation, the historical values are plotted along with the computed and forecast values of their permanent components (the alternating long and short dashed lines). Table 17 provides the numeric forecasts both quarter-by-quarter and on an annual basis. The annual values for the levels of real GDP, nominal M1 (+ Sweeps), and the two interest rate series are four-quarter averages. The annual values for the inflation rates and growth rates are measured on a fourth-quarter to fourth-quarter basis.

Real growth for the second quarter of 1996 is forecast to be very slow, indeed to decline from the rate of 96:1. In retrospect we know that this was a really large forecast error, since the second quarter of 1996 came in exceptionally strong. The forecast is for continued slow real growth (< 2 percent per annum) though the end of 1997. This certainly will be an underestimate of real growth for 96, and appears at this time (January, 1997) to be an underestimate of real growth for 1997 (see discussion of updated forecasts below).

In addition, the forecast did not catch the large increase in long rates and the accompanying increase in the slope of the yield curve that began in the last part of the 96:1 and continued through autumn. Given the size of the standard error of the long-rate equation (45 basis points), errors of this magnitude are not highly unusual, illustrating the difficulty in forecasting nominal interest rates. In comparison, the funds rate is projected to remain essentially constant through 1996 (around 5.10 percent), not far from the actual experience. Thus the long-term rate forecast errors in 1996 are attributable to errors in the implicit short-run term structure relationship. The funds rate is projected to decline to around 4.75 percent by the end of 1997. Taken literally, the model is forecasting a reduction in the fund rate target to around 4.75 percent by the end of 1997. Again, given the standard error of the funds rate equation (62 basis points) such a projection is not inconsistent with no change in the funds rate target over this horizon. M1+Sweeps velocity is forecast to fall by about one percent in 1996, but then to remain constant over 1997.

Finally, the model projects no change in the permanent components of the two inflation rates (at roughly 2.5 percent per annum in terms of the CPI inflation rate) through the end of 1997. Hence the model is not predicting any increase in the long-run rate of inflation over this period.

Updated forecasts incorporating data for 96:2 are plotted in Figure 52. The data used here are those available in December, 1996 and reflect minor revisions from the earliest available estimates for the second quarter of 1996. Forecast values are indicated in Table 18. While real growth for 1996 is revised upward substantially from the forecasts in Table 17, real growth forecasts for 1997 are essentially unchanged (1.46 percent fourth quarter to fourth quarter compared to 1.44 percent previously). The permanent inflation forecast is revised upward by about 20 basis points on each measure. This reflects a revision in the forecast trend in the permanent component of inflation from zero to slightly positive. Real growth for 96:3 is forecast at a strong 3.19 percent which substantially overestimates the current (December, 1996) estimate of actual 96:3 real growth of 2.30 percent. Both long-term and short-term interest rate forecasts are revised upward substantially. the implicit monetary policy reaction function in the reduced form structure implies that the FOMC would have raised its funds rate target to around 5.75 percent for 96:3 and to 6.0 percent in 96:4. In retrospect we know that the funds rate target was maintained at 5.25 percent, though there was considerable public speculation that the FOMC would vote to raise their target in early autumn. Again, given the large standard error of the residuals of the funds rate equation (> 60 basis points) forecast errors of this magnitude will be observed frequently. The 10-year Treasury rate is forecast to reach a peak of around 7.2 percent in 96:4 - 97;1. This clearly overestimates the 1996 experience, though the forecast for 96:3 of 7.01 percent is quite accurate (actual = 6.83 percent), in spite of the large funds rate forecast error. Again there is a forecasting error in the implicit short-run term structure relationship in the same direction as in the previous forecasting experiment.

The third set of forecasts, shown in Figure 53, are based on the information through 96:3 that was available as of January 1, 1997. Here we have extended the forecasting horizon out to the end of 1998 to provide forecasts over the typical horizon of public forecasts that are made towards the end of each calendar year. With the addition of data through 96:3, the forecast of real growth for 1997 is revised upwards substantially and the model now projects fairly strong real growth (2.68 percent) from the fourth quarter of 1996 through the fourth quarter of 1997 (see Table 19). The extension of the forecast into 1998 indicates an even stronger projection of real growth from the fourth quarter of 1997 through the fourth quarter of 1998: 3.04 percent. This upward revision in the forecast of real GDP growth is not accompanied by any indication of a sustained acceleration of inflation. The forecasts of the permanent component of both inflation measures for 96:3 proved to be almost exactly correct. The previous forecasting experiment projected these inflation measures at 2.33 and 2.56 percent per annum for the deflator inflation and CPI inflation respectively. Based on currently available data, the actual values are 2.27 and 2.51 percent. The permanent components of the inflation measures are now forecast at 2.28 and 2.52 percent for fourth quarter of 1996 to fourth quarter of 1997. The corresponding values in Table 18 are 2.52 and 2.75 percent, so there is a small downward revision in the forecast of sustained inflation for 1997. The current projections for 1998 are an additional slight decline in these measures of sustained inflation to 2.15 and 2.39 percent. Hence, the projection of strong real growth over the next two years is <u>not</u> accompanied by any indication of an upward trend in the inflation rate.

The forecasts in Table 19 suggest that the long-term interest rate will remain at roughly the 96:3 level through the first quarter of 1997, then decline slowly to around 6.20 percent by the end of 1998. The Federal funds rate is now projected to remain around 5.25 percent through 96:4 (a forecast which is known at this date to be quite accurate), then to decline by about 25 basis points over each of the three succeeding quarters and then to level off in the 4.50 percent range for the second half of 1997 and all of 1998. This implies that the FOMC, if it follows the implicit monetary policy rule incorporated in the reduced form structure, will cut the funds rate target by 25 basis points in the first three quarters of 1997. Again, recall that the standard error of the residuals of the funds rate equation (> 60 basis points), so the forecast pattern of the funds rate is not inconsistent with an unchanged funds rate target over the first half of 1997.

Figure 51
Based on information through 96:1

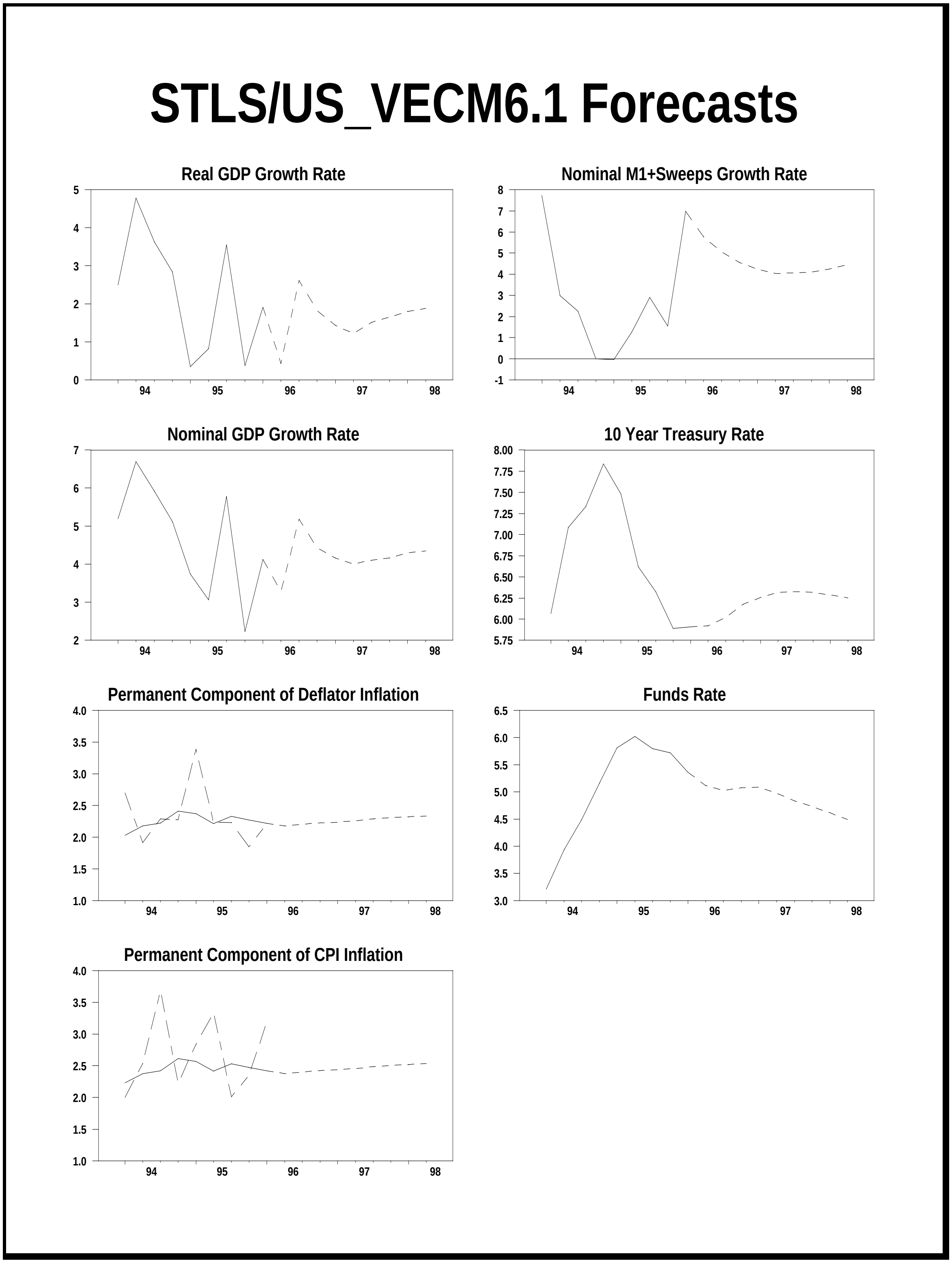


Table 17: Forecasts Based on Information Through 96:1

	Actual 1996:01	1996:02	1996:03	1996:04	1997:01	1997:02	1997:03	1997:04
Real GDP	6813.6	6821.0	6865.7	6897.1	6922.0	6943.2	6969.5	6998.4
Real GDP Growth Rate	1.92	0.43	2.61	1.83	1.44	1.22	1.51	1.65
Nominal GDP Growth Rate	4.13	3.28	5.19	4.43	4.16	4.01	4.11	4.16
Permanent Deflator Inflation	2.22	2.18	2.20	2.23	2.24	2.26	2.29	2.31
Permanent CPI Inflation	2.42	2.38	2.40	2.43	2.44	2.46	2.49	2.51
Nominal M1 + Sweeps	1195.9	1213.3	1228.7	1242.8	1256.0	1268.8	1281.7	1294.9
Growth Rate of M1 + Sweeps	6.98	5.78	5.05	4.56	4.24	4.05	4.05	4.11
10 Year Treasury Rate	5.91	5.92	6.02	6.18	6.26	6.32	6.33	6.32
Federal Funds Rate	5.36	5.11	5.03	5.07	5.09	4.98	4.84	4.73

	Actual 1995	1996	1997
Real GDP	6742.8	6849.3	6958.3
Real GDP Growth Rate	1.28	1.70	1.46
Nominal GDP Growth Rate	3.71	4.26	4.11
Permanent Deflator Inflation	2.30	2.21	2.27
Permanent CPI Inflation	2.50	2.41	2.47
Nominal M1 + Sweeps	1166.6	1220.1	1275.3
Growth Rate of M1 + Sweeps	1.42	5.59	4.11
10 Year Treasury Rate	6.58	6.01	6.30
Federal Funds Rate	5.84	5.14	4.91

Figure 52
Forecasts Based on Information through 96:2

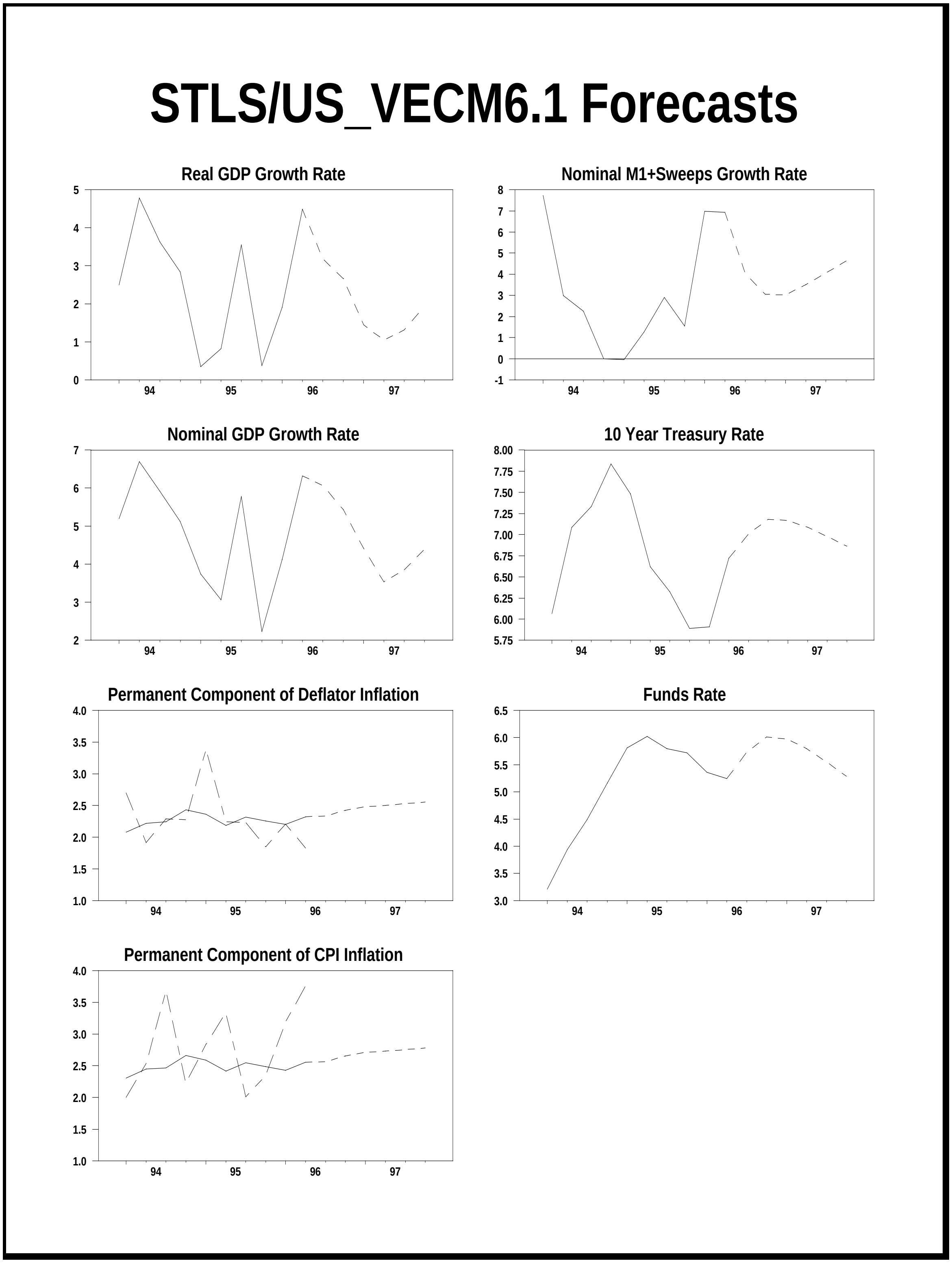


Table 18:Forecast based on information through 96:2

	Actual 1996:02	1996:03	1996:04	1997:01	1997:02	1997:03	1997:04
Real GDP	6890.5	6945.6	6992.1	7017.5	7036.0	7059.2	7093.5
Real GDP Growth Rate	4.49	3.19	2.67	1.45	1.05	1.32	1.94
Nominal GDP Growth Rate	6.32	6.07	5.44	4.42	3.54	3.85	4.40
Permanent Deflator Inflation	2.33	2.33	2.43	2.48	2.50	2.53	2.55
Permanent CPI Inflation	2.56	2.56	2.66	2.71	2.73	2.76	2.78
Nominal M1 + Sweeps	1216.8	1229.1	1238.6	1248.0	1259.0	1271.8	1286.7
Growth Rate of M1 + Sweeps	6.94	4.04	3.05	3.04	3.51	4.06	4.64
10 Year Treasury Rate	6.72	7.01	7.18	7.17	7.09	6.98	6.86
Federal Funds Rate	5.24	5.73	6.01	5.97	5.80	5.55	5.29

	1995	1996	1997
Real GDP	6742.8	6910.5	7051.6
Real GDP Growth Rate	1.28	3.07	1.44
Nominal GDP Growth Rate	3.71	5.49	4.05
Permanent Deflator Inflation	2.28	2.32	2.52
Permanent CPI Inflation	2.51	2.55	2.75
Nominal M1 + Sweeps	1166.6	1220.1	1266.4
Growth Rate of M1 + Sweeps	1.42	5.25	3.81
10 Year Treasury Rate	6.58	6.70	7.02
Federal Funds Rate	5.84	5.59	5.65

Figure 53
Forecasts based on Information through 96:3

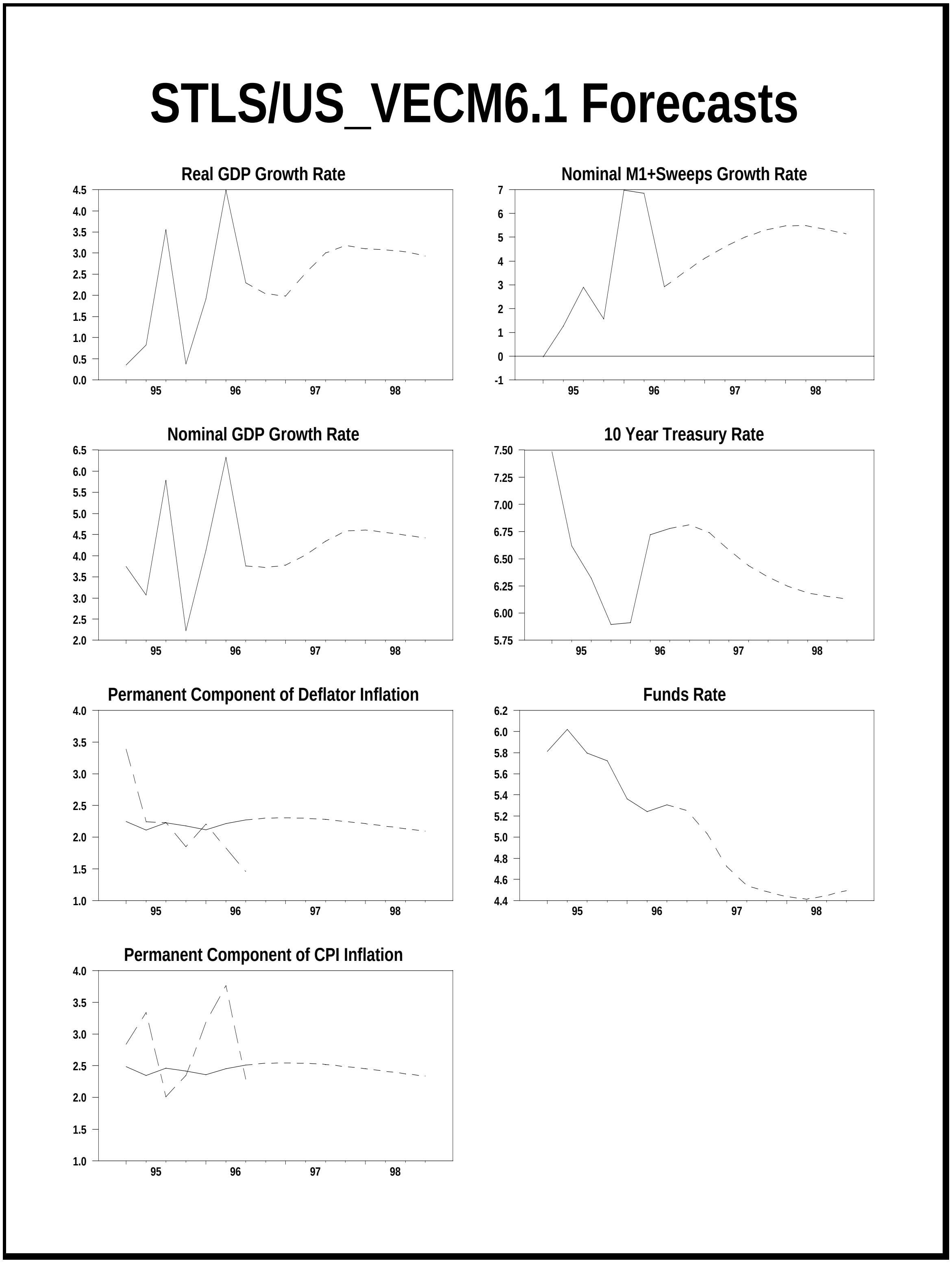


Table 19: Forecasts Based on Information through 96:3

	Actual 1996:03	1996:04	1997:01	1997:02	1997:03	1997:04	1998:01	1998:02	1998:03	1998:04
Real GDP	6930.2	6965.6	7000.3	7044.8	7097.9	7154.6	7210.3	7266.0	7321.3	7375.2
Real GDP Growth Rate	2.30	2.04	1.98	2.53	3.01	3.18	3.10	3.08	3.03	2.93
Nominal GDP Growth Rate	3.76	3.72	3.78	4.02	4.34	4.58	4.61	4.55	4.49	4.43
Permanent Deflator Inflation	2.27	2.30	2.30	2.30	2.28	2.25	2.21	2.17	2.13	2.10
Permanent CPI Inflation	2.51	2.54	2.54	2.54	2.52	2.49	2.45	2.41	2.37	2.34
Nominal M1 + Sweeps	1225.4	1236.3	1249.0	1263.5	1279.5	1296.5	1314.4	1332.6	1350.4	1367.9
Growth Rate of M1 + Sweeps	2.92	3.53	4.11	4.61	5.01	5.30	5.48	5.49	5.33	5.14
10 Year Treasury Rate	6.78	6.81	6.74	6.58	6.44	6.33	6.25	6.19	6.15	6.13
Federal Funds Rate	5.31	5.25	5.03	4.72	4.54	4.49	4.44	4.41	4.45	4.49

	1996	1997	1998
Real GDP	6900.0	7074.4	7293.2
Real GDP Growth Rate	2.69	2.68	3.04
Nominal GDP Growth Rate	4.48	4.18	4.52
Permanent Deflator Inflation	2.23	2.28	2.15
Permanent CPI Inflation	2.47	2.52	2.39
Nominal M1 + Sweeps	1218.5	1272.1	1341.3
Growth Rate of M1 + Sweeps	5.07	4.76	5.36
10 Year Treasury Rate	6.56	6.52	6.18
Federal Funds Rate	5.29	4.70	4.45

9. Possible Extensions

There are a number of directions in which this research effort can be extended. First, current policy discussions focus on real GDP in chained 1992 dollars and the inflation rate in terms of the associated price index for this concept, rather than deflated real GDP and the implicit GDP deflator that are utilized in this VECM. Currently, real GDP in chained 1992 dollars and its corresponding price index are available only from the third quarter of 1959. Some decision on how to measure real GDP prior to the middle of 1959 is necessary if we are to reconstruct the model in terms of the contemporary chained GDP measures. One possibility is to chain, at 59:3, the old fixed-weight real GDP in 1987 dollars (the numbers available prior to the 1996 revision of the National Income and Product Accounts, published through late 1995) to the new real GDP in chained 1992 dollars. This approach preserves the growth rates in the old real GDP measure from the beginning of the sample to 59:3. We have constructed some preliminary estimates of the interest semielasticity of the money demand cointegrating vector utilizing this approach to measure real GDP for the entire sample period. These estimates suggest that the estimate of this coefficient is little changed by the alternative data series. This is not surprising since the chained price index in 59:3 and 95:4 is 23.0 and 105.8 respectively, while the implicit GDP deflator is 22.9 and 105.8 for the same two periods. Changing the price index will leave the trends in real output, real balances, and GDP price index inflation unaffected, and thus will have little impact on the estimated parameters of the cointegrating vectors. We have not undertaken the more extensive analysis to determine if all the properties of the model we have estimated here are invariant to the change in data series.

Some attention should also be given to when to end the break in the sample that begins in 79:4. In this work and our earlier work we have consistently ended the break in the sample with 81:4. This dates back to the study by Rasche (1987) of breaks in the trend of M1 velocity, which always dated a trend break towards the end of 1981, regardless of the degree of time aggregation of the data. For estimation of a "real balance–real income–nominal interest rate" cointegrating vector, this timing of the end of the sample break is not terribly important: if the data from 79:4 through 81:4 are omitted, there is no significant shift in the mean of the cointegrating vector. Once the model is expanded to include real interest rates, as is the case here, it may be more appropriate to also omit the first three quarters of 1982, since the research of Huizinga and Mishkin (1986) dated breaks in real interest rates at both 79:4 and 82:4.

A second extension would be to replace the M1 series with the new St. Louis Adjusted Monetary Base series. The advantage of the latter concept is that it is not as seriously affected by "sweeps" subsequent to 1993.[26] Alternatively, a cointegrating vector in the real Adjusted Monetary Base, real GDP and nominal interest rates could be added to the VECM structure estimated here. Anderson, Hoffman and Rasche (1997) have preliminary estimates indicating the existence of such a cointegrating vector.

A third extension of the VECM model would be to investigate the effect of adding a commodity price variable to the specification. Sims has estimated a number of VARs over the years that utilize such a variable, and he finds that such a variable is useful in clarifying some puzzling dynamic properties of his models. To our knowledge, no one has investigated the role of commodity price variables in vector error correction structures such as that estimated here. One problem that would need to be investigated prior to such an analysis is the correct order of integration of such a commodity price variable.

A fourth extension of the analysis presented here would be to add additional cointegrating vectors involving consumption and investment, such as those specified by King, Plosser, Stock and Watson (1991) and Fama (1992). This would add additional detail to the model forecasts, and would provide a richer menu of transitory shocks. With this broader menu of transitory shocks it may be possible to develop identifying restrictions which better define the short-run structure of the economy.

Within the structure that is estimated here, a number of interesting research initiatives can be pursued. First, a more extensive investigation of the relative forecasting accuracy of the model can be pursued beyond that which we have prepared in section 4 of this study. We have tabulated semiannual forecasts from the FOMC,

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[26] Research presently underway is focused on developing an appropriate RAM adjustment to account for the effective "homebrewed" reduction in reserve requirements that is produced by "sweeping". If this

Administration, CBO, and Blue Chip indicators over the past decade, from Federal Reserve and CBO publications. Semiannual forecasts of the *Wall Street Journal* panel of economists are available. For all of these forecasts, reasonably accurate information on the timing of the date of the forecast can be reconstructed. In addition there is the data base of forecasts maintained by the Federal Reserve Bank of Philadelphia, although ascertaining the exact timing of these forecasts is problematic. Where we can determine the approximate information set available to these forecasters, we can produce VECM forecasts which can be used for comparative forecasting analysis.

Second, we have identified major outliers in each of the VECM residuals in section 3 of this study, but we have not undertaken an extensive historical analysis of these particular dates to see if we can determine any particular relevant omitted variables that may be a source of bias to our estimates. This may be particularly important with respect to the interest rate spread implicit in the long-term rate equation. The extremely large residuals for this equation that occur quite regularly during the 1980s are troublesome and warrant careful analysis.

Third, while the forecasting properties of this simple VECM are comparable to other forecasts or forecasting instruments in the periods that we have studied, the errors are not necessarily small. Recently two analyses have been published which discuss how high frequency information may be incorporated to improve the forecasting properties of quarterly models, at least over short-term forecasting horizons (Miller and Chin, 1996; Ingenito and Trehan, 1996). Reifschneider, Stockton, Wilcox (1996) imply that judgmental forecasters at the Board of Governors consistently improve upon model based forecasts because they can incorporate very high frequency information that is unknown to the forecasting models utilized at the Board of Governors. An interesting and possibly quite productive analysis would examine how high frequency information can be systematically incorporated into VECM forecasting models such as that constructed here and to measure the improvement in forecast error that is derived from alternative approaches to this problem.

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measurement can be accomplished, then the St. Louis Adjusted Monetary Base series should be invariant to the “sweeping” that has occurred in 1994-1996.

Finally, the possibility of utilizing the structure of the Federal funds rate equation that is estimated here to "transplant" alternative monetary policy rules into the vector error correction model for the purpose of constructing conditional forecasting experiments needs careful and extensive investigation.

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