Money-Multiplier Sh(S)20.25(oop(6725(o392)3

1 Introduction

An enormous body of literature, dating back to Hume (1752), investigates the e ects that changes in the money supply have on output and prices. Classic studies in the monetarist tradition, including Friedman and Schwartz's (1963) Monetary History of the United States (henceforth, MHUS) and Cagan (1965), go further, by decomposing the money stock into its three "proximate" determinants: the monetary base, the ratio of currency to deposits, and the ratio of reserves to deposits. Their aim was to use this decomposition as part of a "narrative" e ort to pinpoint the fundamental sources of co-movement in money and other key macroeconomic variables – that is, in the language of modern econometrics, to solve the problem of identifying and estimating the e ects of structural disturbances to the economy.

To review the familiar decomposition, let the monetary aggregate be the sum of currency in circulation and deposits . The monetary base (often referred to synonymously as the stock of high-powered money), meanwhile, equals the sum of currency and bank reserves . Now,

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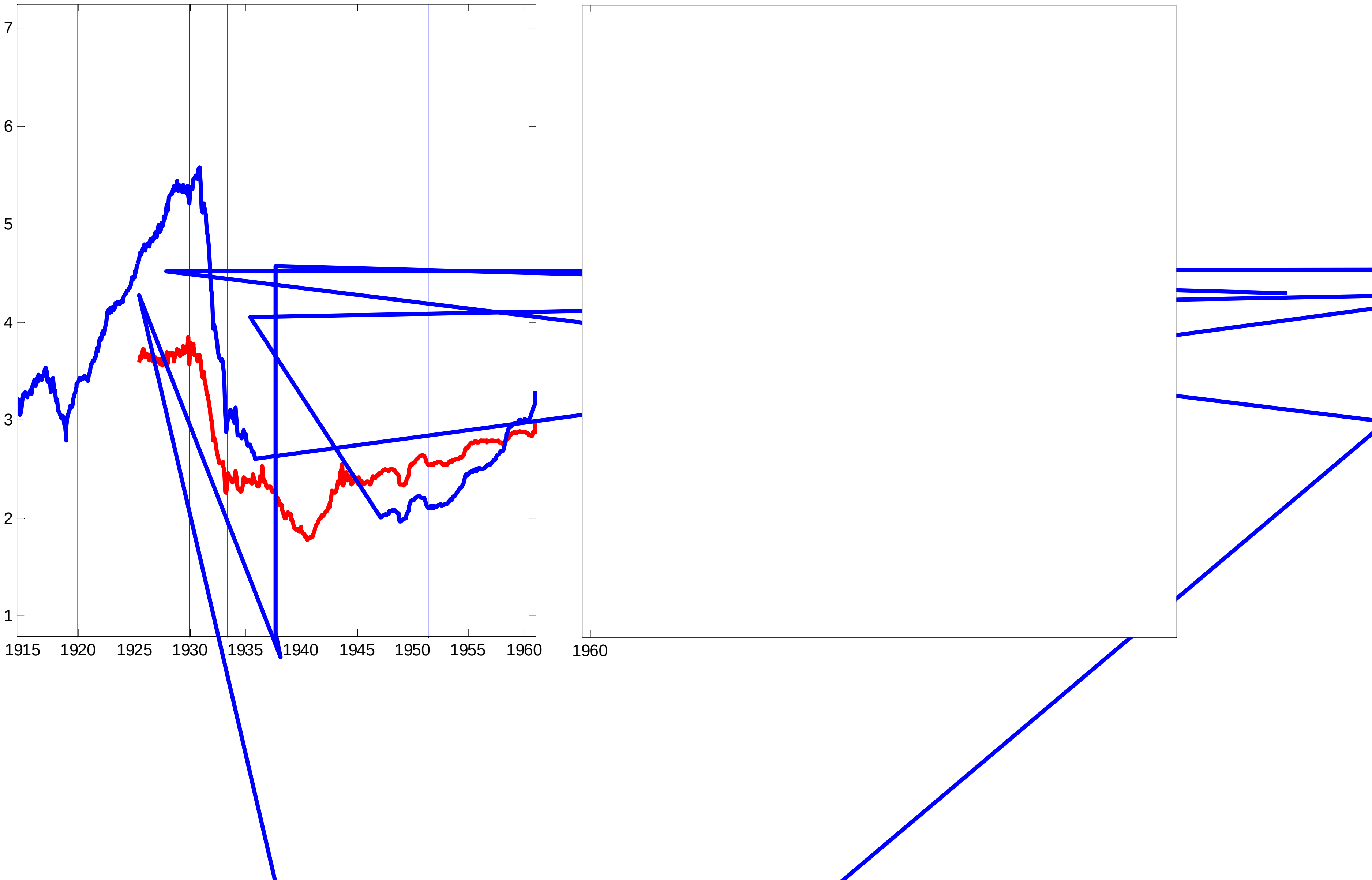
where = is the currency-deposit ratio and = the reserve-deposit ratio and, as indicated by the last equality, the money multiplier depends on both of these ratios.

Friedman and Schwartz (1963) and Cagan (1965) Þnd important roles for the money multiplier and the two ratios on which it depends in their narrative histories. Most famously, Chapter 7 of Friedman and Schwartz's MHUS describes how, beginning in October 1930, the severe contraction characterized initially by a decline in the can be seen in isolation. To cite just one additional example, Chapter 9 of the MHUS describes how the Federal Reserve moved, first in July 1936 and again in January 1937 to raise reserve requirements in several steps by a total of 3 billion dollars, an amount then equal to nearly 25 per cent of the monetary base. Although they were intended by Federal Reserve officials to simply be a "precautionary measure to prevent an uncontrollable expansion of credit in the future" (MHUS, p. 524), Friedman and Schwartz note (p. 527) that their ultimate effect on the money multiplier, working through changes in the reserve ratio , led the money stock to reach "an absolute peak in March" 1937 and to fall "with only minor interruption to the end of the year." Once again, in the MHUS, what was identified as an important, autonomous shift in the money multiplier appeared to be followed by a sharp cyclical contraction.

Despite the prominent role assigned to the money multiplier in these historical studies, and despite the obvious connections between the aims of Friedman and Schwartz (1963) and Cagan's (1965) narrative analyses and the goal of modern econometrics – namely, to learn about the structure of the economy by identifying the exogenous disturbances that drive large cyclical fluctuations in aggregate output and prices – the recent literature features no attempt, to the best of our knowledge, to build on and extend these analyses with the help of more formal, time-series methods.[1] In this paper, we aim to fill this gap in the literature. In particular, we use cointegrated structural VARs, in which fundamental disturbances are identified using long-run restrictions, to re-address the same questions posed by Friedman and Schwartz and Cagan. How important do identified shocks to the two components of the money multiplier – the currency-to-deposit and reserve-to-deposit ratios – appear to be in driving macroeconomic dynamics during the interwar period? Can an analysis based on modern time-series analysis confirm the conclusions of these classic studies? To what extent do movements in the money multiplier continue to be important in explaining movements in aggregate output and prices during the post-World War II era and, in particular, during the period of the Great Inflation of the 1970s which, after the Great Depression and before the financial crisis of 2007-08, represents the most striking period of monetary instability in a long span of United States economic history?

Þnancial system precluded any detailed analysis of the multiplier for the non-M1 components on M2. Here, we can – and do – examine both the M1 and M2-M1 mul-

June 1914-December 1960
(based on Friedman and Schwartz's data)

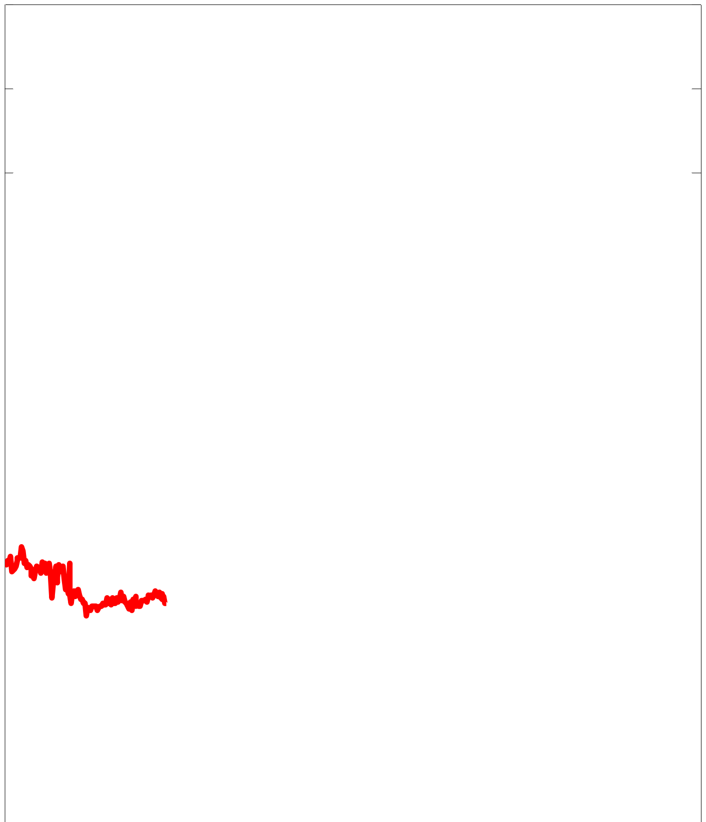


As discussed by Lucas and Nicolini (2015), the rationale for including MMDAs in M1 is that they perform an economic function similar to the more traditional 'checkable deposit' component of the Federal Reserve's o cial M1 series. In fact, Benati, Lucas, Nicolini, and Weber (2017; henceforth, BLNW) show that whereas – in line with, e.g., Friedman and Kuttner (1992) – based on the standard aggregate there is no evidence of a stable long-run demand for M1, evidence of cointegration between velocity and the short rate is very strong based on Lucas and Nicolini's (2015) aggregate.

During either period, the multiplier of M2-M1 had exhibited a strong positive correlation with the short rate. In fact, as we will discuss in Section 4.1, for both periods we detect strong evidence of cointegration between the two series.[2] The most natural explanation for this stylized fact has to do with permanent portfolio shifts out of (mostly) non interest-bearing M1, and into interest-bearing M2-M1, caused by permanent interest rate shocks, whatever their origin (i.e., permanent inßation shocks, or permanent shocks to the real interest rate).

The M1 multiplier, on the other hand, does not exhibit a consistent pattern across sub-periods. During the period January 1919-December 1960 it also exhibits a strong positive correlation with the short rate. It is to be noticed, however, that Þrst, Johansen's tests do not detect cointegration between the two series (see Table 2a); and second – and crucially – the explanation for such a correlation, in terms of direction of causality, is most likely completely di erent from that for the multiplier of M2-M1. In particular, the narrative account of interwar macroeconomic ßuctuations provided by Friedman and Schwartz in chapters 6 to 9 of MHUS the multiplier, to subsequent ßuctuations in the economy, including movements in the short rate. For the multiplier of M2-M1, on the other hand, evidence of cointegration with the short rate suggests that its key driver were the permanent shifts in interest rates caused by shocks to the determinants of the M1 multiplier.

Turning to the period 1959Q1-2008Q3, up until the introduction of MMDAs, in 1982Q4, the M1 multiplier followed a remarkably smooth path, and exhibited, overall, comparatively little variation, ßuctuating between 2.44 and 2.92 (between January 1919 and the attack on Pearl Harbor, on the other hand, it had ßuctuated between 1.78 and 3.85). The natural explanation for such a smooth path up until



bank failures beginning in October 1930, culminating in December of that year with the collapse of the Bank of the United States. The relentless climb in continued through subsequent waves of banking failures in 1931, 1932, and 1933, ending only after Roosevelt's banking holiday of March 1933.

On the other hand, the increase in the reserve-deposit ratio that accompanied, but initially lagged behind, that in went on through June 1940 according to a dramatic series of events outlined in Chapters 7 and 9 of the MHUS. Banks' desire to increase their holdings of reserves before the Bank Holiday of 1933 was the natural response to the series of bank runs and panics that produced the rise in . But this accumulation or reserves continued even after fell back towards more normal levels. As Friedman and Schwartz (p. 348) explain,

Figure 3 United States, June 1914-December 1960: Business-cycle components of log industrial production, the multipliers, and the currency/deposits and reserves/deposits ratios, and average gain and coherence at th

Figure 4 United States, 1959Q1-2016Q4: Business-cycle components of log industrial production, the multipliers, and the currency/deposits and reserves/deposits ratios, and average gain and coherence at the business-cycle frequencies autonomous shifts in and – and therefore in the M1 multiplier – in driving macroeconomic ßuctuations during the interwar period. As we will see in Section 4, our analysis conÞrms indeed Friedman and Schwartz's position. Intuitively, this should lead us to expect to Þnd a strong correlation between real activity and either , , or the M1 multiplier at the business-cycle frequencies during this period. The evidence in Figure 3 conÞrms indeed this conjecture. Both and had exhibited a strong counter-cyclical pattern, whereas the M1 multiplier had been very strongly pro-cyclical. As for the multiplier of M2-M1, the pattern had been strongly pro-cyclical until World War II, and it then turned mainly counter-cyclical after that. By the same token, the bootstrapped distributions of the coherence of the series of interest onto industrial production points towards a sizeable explanatory power of the former for the latter. (It is worth recalling that the coherence, which by construction is bounded between 0 and 1, is nothing but the R-squared in the regression of one variable onto the other at a speciÞc frequency, or within a speciÞc frequency band. By the same token, the gain is the absolute value of the slope coe cient in the same regression.) This is especially clear for and for the M1 multiplier, whereas it is less so for , and especially for the multiplier of M2-M1.

For the period 1959Q1-2008Q3, on the other hand, our evidence in Section 4 suggests that shocks to either or , and therefore to the M1 multiplier, had played a negligible role in driving macroeconomic ßuctuations during those years. The evidence in Figure 4 is, under this respect, mixed. On the one hand, the relationship between the business-cycle components of either of the four series of interest, and the business-cycle component of GDP, is not nearly as strong and clear-cut as for the former period. On thermt-313.e5(r)-251[hae)3.ndst,h-7.6(h)26(w229.4(e24.2(e)-5.3(2-8.7(,)-)-5.2(t)-5.29(e)-2.2(b)-28

Table 1 a United States, January 1919- December 1960: Bootstrapped p-values for Elliot, Rothenberg, and Stock unit root tests [d]

	Lag order:			
	p=3	p=6	p=9	p=12
	In levels, without a time trend			
New York FED discount rate	0.293	0.210	0.298	0.279
High grade bond rate	0.564	0.618	0.457	0.191
BAA rate	0.508	0.506	0.369	0.150
Logarithm of (1 + r)	0.617	0.383	0.083	0.093
Logarithm of (u+ r)	0.803	0.665	0.439	0.460
n	0.617	0.396	0.085	0.084
u	0.789	0.679	0.522	0.468
M1 multiplier	0.799	0.684	0.478	0.475
Multiplier of M2-M1	0.891	0.819	0.691	0.706
	In levels, with a time trend			
Log nominal M0	0.991	0.981	0.942	0.977
Log CPI	0.803	0.798	0.611	0.547
Log industrial production	0.436	0.360	0.106	0.268
Log department store sales	0.977	0.944	0.774	0.727
	In differences, without a time trend			
	p=3	p=6	p=9	p=12
Log CPI	0.000	0.000	0.001	0.000
New York FED discount rate	0.000	0.000	0.000	0.000
High grade bond rate	0.000	0.000	0.000	0.000
BAA rate	0.000	0.000	0.000	0.000
Logarithm of (1 + r)	0.000	0.000	0.003	0.001
Logarithm of (u+ r)	0.000	0.000	0.023	0.022
n	0.000	0.000	0.002	0.001
u	0.000	0.000	0.016	0.018
M1 multiplier	0.000	0.001	0.069	0.082
Multiplier of M2-M1	0.000	0.000	0.017	0.033
Log nominal M0	0.000	0.000	0.012	0.004
Log industrial production	0.000	0.000	0.000	0.000
Log department store sales	0.000	0.000	0.003	0.009

[d] Based on 10,000 bootstrap replications of estimated ARIMA processes.
n= currency/deposits ratio. u= reserve/deposits ratio.

Table 1 b United States, 1959Q1-2008Q3: Bootstrapped p-values for Elliot, Rothenberg, and Stock unit root tests[d] period we end the sample in 2008Q3 as the subsequent explosion in reserves – and therefore in the monetary base – associated with quantitative easing policies, would render any analysis of the money multiplier meaningless.) For all series exhibiting obvious trends the tests are based on models including an intercept and a time trend.[7] These series are the logarithms of nominal M0, the CPI, industrial production, and department store sales for the former period; and of nominal M0, nominal M0 and M1 per capita, and real GDP and consumption per capita for the latter period. For all other series the tests are based on models including an intercept, but no time trend. For log hours worked per capita for the period 1959Q1-2008Q3 – for which visual evidence on the presence or absence of a trend is not clear-cut – we report results from tests based on either model. As for the determinants of the M1 multiplier, we report results both for the levels of   and , and for the logarithms of the numerator and denominator of the multiplier – that is: ln(1+) and ln(+), respectively. The rationale for also reporting results for the two latter variables is that, in Section 5, we will identify permanent shocks to   and   by entering ln(1+) and ln(+) in cointegrated VARs, and then imposing a Cholesky structure on the respective (2×2) block of the long-run impact matrix of the structural shocks. Because of this, we want to be sure that not only   and , but also ln(1+) and ln(+) are I(1).

At the 10 per cent significance level we take as our benchmark throughout the entire paper, the following results emerge from the two tables:

( ) inflation had been I(0) in the former period, whereas it has been I(1) in the latter one.

( ) The monetary base had been I(1) in the former period, whereas it has been trend-stationary in the latter one. (The second result is robust to considering either M0, or M0 per capita.)

(   ) For all other series, the null of a unit root cannot be rejected.[8]

(   ) Finally, for all series, and for either period, tests in differences without a time trend strongly reject the null of a unit root. (This is crucial because a necessary condition for performing Johansen's tests is that the series under investigation do contain a unit root, but that their order of integration is not greater than one.)

Both ( ) and ( ) justify our choice of performing the analysis by sub-sample, rather than for the joint sample 1919-2007 based on annual data.

4.2 Cointegration tests

Tables 2a and 2b report, for either period, results from Johansen's cointegration tests for both the 10-variables systems which will be the focus of our analysis in Section 5,

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[7] The reason for including a time trend is that, as discussed e.g. by Hamilton (1994, pp. 501), the model used for unit root tests should be a meaningful one also under the alternative.

[8] For ln(1+ r) and ln(u+ r) for the period January 1919-December 1960, a unit root is rejected, at the 10 per cent level, based on s = 9 and s = 12 (but not based on the other two lag orders). In either case, we regard the null of a unit root as not having been convincingly rejected, and in what

and several smaller sub-systems. For the period January 1919-December 1960, the 10-variables system features the logarithms of industrial production, department store sales, M0, the CPI, (1+), and (+); the multiplier of M2-M1; and the New York FED discount rate, Moody's BAA corporate bond yield, and the index of yields of high grade public utility bonds. For the period 1959Q1-2008Q3, on the other hand, it features the logarithms of GDP, consumption, and hours per capita; the logarithms of (1+) and (+); and the multiplier of M2-M1, M1 velocity, inßation, the Federal Funds rate, and the 5-year Treasury bill rate.

Following BLNW (2017), we bootstrap the tests[9] via the procedure proposed by Cavaliere et al. (2012; henceforth, CRT). In a nutshell, CRT's procedure is based on the notion of computing critical and p-values by bootstrapping the model which is relevant under the null hypothesis[10] All of the technical details can be found in CRT, which the reader is referred to. We select the VAR lag order as the maximum[11] between the lag orders chosen by the Schwartz and the Hannan-Quinn criteria[12] for the VAR in levels, for a maximum allowed lag order of = 12 for the former period, and = 4 for the latter one.

The following results emerge from the two tables:

( ) In line with BLNW's (2017) results for the U.S. over the entire period since 1915, we detect strong evidence of a long-run demand for M1 for either period. Specifically, for the period 1959Q1-2008Q3 we detect, as BLNW (2017), cointegration between M1 velocity and the short rate. This corresponds to the speciÞcation originally estimated by Selden (1956) and Latané (1960), which is linear in the

	Trace tests of the null of no cointegration against the alternative of h or more cointegrating vectors:			
	h = 1	h = 2	h = 3	h = 4
Baseline 10-variables system	463.621 (0.000)	311.890 (0.000)	213.786 (0.000)	131.868 (0.000
New York FED discount rate and M1 multiplier	13.427 (0.199)			
New York FED discount rate and multiplier of M2-M1	31.223 (8.0e-4)			
New York FED discount rate and high-grade bond rate	19.708 (0.033)			
Log industrial production, log M_1, New York FED discount rate, and log CPI	75.518 (0.000)	33.588 (0.003)		
	Maximum eigenvalue tests of h versus h+1 cointegrating vectors:			
	0 versus 1	1 versus 2	2 versus 3	3 versus 4
Baseline 10-variables system	151.731 (0.000)	98.104 (0.000)	81.918 (0.000)	45.919 (0.216)
New York FED discount rate and M1 multiplier	12.524 (0.121)			
New York FED discoun, New York FED discount rate, and log CPI	41.931 (0.002)	19.365 (0.159)		

[d] Bootstrapped p-values (in parentheses) are based on 10,000 bootstrap replications, based on Cavaliere et al.'s (2012) methodology.

Table 2 b United States, 1959Q1-2008Q3: Results from Johansen's cointegration tests for alternative systems [d]

	Trace tests of the null of no cointegration against the alternative of h or more cointegrating vectors:			
	h = 1	h = 2	h = 3	h = 4
Baseline 10-variables system	360.293 (0.000)	268.117 (0.000)	198.429 (0.000)	147.754 (0.000)
Federal Funds rate and multiplier of M_2-M_1	25.307 (0.010)			
Federal Funds rate and 5-year Treasury bill rate	29.036 (0.001)			
Federal Funds rate and M_1 velocity	21.703 (0.031)			
Federal Funds rate and inflation	16.051 (0.117)			
Logarithms of (1+k), (r+k), and M_1 per capita	30.152 (0.097)			
Log real GDP per capita and log real consumption per capita	22.480 (0.009)			
	Maximum eigenvalue tests of h versus h+1 cointegrating vectors:			
	0 versus 1	1 versus 2	2 versus 3	3 versus 4
Baseline 10-variables system	92.176 (0.003)	69.687 (0.067)	50.675 (0.399)	–
Federal Funds rate and multiplier of M_2-M_1	23.086 (0.008)			
Federal Funds rate and 5-year Treasury bill rate	26.232 (0.001)			
Federal Funds rate and M_1 velocity	18.156 (0.032)			
Federal Funds rate and inflation	9.665 (0.325)			
Logarithms of (1+k), (r+k), and M_1 per capita	23.185 (0.047)			
Log real GDP per capita and log real consumption per capita	21.937 (0.005)			

[d] Bootstrapped p-values (in parentheses) are based on 10,000 bootstrap replications, based on Cavaliere et al.'s (2012) methodology.

The maximum eigenvalue test of one versus two cointegration vectors, on the other hand, does not reject the null, leading us to conclude that, in line with what we would expect ex ante based on economic theory, the system features one, and only one cointegration relationship, i.e., the long-run demand for M1.

( ) Again, as we would expect based on theory, in either period short- and long-term nominal rates are cointegrated.

( ) The same holds, in the latter period, for real GDP and consumption per capita.

( ) Interestingly, for both periods we detect very strong evidence of cointegration between the multiplier of M2-M1 and the short-term rate. This provides statistical support to the visual impression from Figure 1 of a very strong relationship between the two series in either period. As previously discussed, the natural explanation for this pattern has to do with the permanent portfolio shifts out of (mostly) non interest-bearing M1, and into interest-bearing M2-M1, caused by permanent shocks to nominal interest rates, whatever their origin.

( ) In the period 1959Q1-2008Q3 inßation and the short rate have not been cointegrated. In line with King, Plosser, Stock, and Watson (1991), this suggests that, beyond permanent inßation shocks, the unit root component of nominal interest rates has also been driven by permanent shocks to the real interest rate. In what follows we do not report results for this shock because it explains uniformly minor fractions of forecast error variance for all series[13]

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cointegration vectors is motivated by the previously discussed failure of the maximum eigenvalue test to reject the null hypothesis of two versus three cointegration vectors. Considering four, on the other hand, is motivated by ( ) the results from the trace that, in fact, $^{n}_{w}$and $^{u}_{w}$

5.1.2 Characterizing the extent of uncertainty around the estimated objects

We characterize uncertainty around all of the estimated objects of interest – impulse-response functions (IRFs), fractions of forecast error variance (FEVs), and counterfactual paths obtained by killing o the shocks – by bootstrapping the estimated reduced-form cointegrated VAR as in CRT (2012), and imposing upon the bootstrapped data the same identifying restrictions we impose upon the actual data.

We now turn to discussing the evidence.

5.1.3 Evidence

Figures 5 and 6 show the IRFs to the structural shocks, and the fractions of FEV of individual series explained by either shock, together with the 16th, 84th, 5th, and 95th percentiles of the respective bootstrapped distributions, whereas Figures 7 and 8 show results from several counterfactual simulations in which we kill o either individual shocks, or, jointly, $^{n}_{w}$ $^{u}_{w}$ and $^{Po}_{w}$.

IRFs and fractions of FEV Permanent shocks to industrial production have a statistically insignificant impact on all other series at all horizons, with the single exception of department store sales, which can be thought of a proxy for consumption (in the same way as industrial production is a crude proxy for GDP). As for M0, although the point estimate of the long-run impact is positive – as we should expect – it is borderline insignificant at the one-standard deviation confidence level. It is to be noticed, however, that $^{L\ S}_{w}$ explains a small fraction of the FEV of M0 (probably partly reflecting the imperfect approximation it provides to GDP) so that its impact on the base is necessarily imprecisely estimated.

$^{n}_{w}$ and $^{u}_{w}$ explain dominant fractions of the FEV of ln(1+) and ln(+), respectively, thus providing reassurance that the two shocks have been precisely identified. On the other hand, they play a comparatively small role for other series, with the exception of department store sales and the CPI, for which $^{u}_{w}$ plays a significant role, especially at horizons up to 5-6 years ahead ($^{n}_{w}$ on the other hand, plays a negligible role for either series at all horizons); the monetary base, for which both $^{u}_{w}$ and to a lesser extent $^{n}_{w}$ play sizeable roles, especially at long horizons; and the multiplier of M2-M1, which is largely driven by $^{u}_{w}$ at all horizons. Finally, di erent from sales, neither shock is estimated to have played a sizeable role foeAm5(a)257.7Tj/Fei85(a)25324.1(r)-210012.

Figure 5 United States, January 1919-December 1960: Impulse-response functions to the structural shocks, with 16-84 and 5-95 bootstrapped confidence bands

Figure 6 United States, January 1919-December 1960: Fractions of forecast error variance explained by either of the structural shocks, with 16-84 and 5-95 bootstrapped confidence bands in the multiplier of M2-M1, whereas its impact on interest rates is uniformly statistically insigniÞ

lower than the actual ones (although the di erence is never signiÞcant at the ten per cent level).

$^{P_0}_{w}$ played an important role for the base itself – which, absent these shocks, would have been uniformly lower in the last part of the sample – and for nominal interest rates. In particular, the discount rate would have been uniformly higher, by about one percentage point, during most of the 1920s, and it would have been quite signiÞcantly higher, by about 1-2 percentage points, during the entire period between the Wall Street crash and the early 1950s. Evidence for both the BAA rate, and especially the high grade bond rate is even starker, with the counterfactual paths for both rates being uniformly higher than the actual, historical paths during most of the sample period, and most of the time by sizable amounts. This is especially the case for the second half of the 1920s, and for the period between Roosevelt's inauguration and the early 1950s.

These statistical Þndings are fully consistent with the narrative told in the MHUS. Friedman and Schwartz's (1963, p. 332) observe, for example, that changes in high-powered money alone would have produced a steady rise rate, and the high-grade bond rate, all of which would have been, between the early

Figure 9 United States, 1959Q1-2008Q3: Impulse-response functions to the structural shocks, with 16- 84 and 5-95 bootstrapped confidence bands

Figure 10 United States, 1959Q1-2008Q3: Fractions of forecast error variance explained by either of the structural shocks, with 16-84 and 5-95 bootstrapped confidence bands

5.2.2 Evidence

Figures 9 and 10 show the IRFs to the Þve identiÞed structural shocks, and the fractions of FEV of individual series explained by either shock, whereas Figures 11 to 13 show results from several counterfactual simulations in which we kill o either individual shocks, or, jointly, $^{n}_{w}$ and $^{u}_{w}$.

IRFs and fractions of FEV Permanent shocks to hours explain the bulk of hours' ßuctuations, especially at long horizons, but they are largely irrelevant for almost all other series. The notable exception are GDP and consumption, for which $^{k}_{w}$ explain about one-fourth of the FEV at all horizons. As expected, a permanent shock to hours leads to permanent increases in both GDP and consumption (although for this latter variable the long-run impact is borderline insigniÞcant).

Shocks to the reserves/deposits ratio are uniformly irrelevant for all series – including ln(+) – at all horizons, with the fractions of explained FEV being consistently negligible. In the light of this, the fact that the response of hours to a positive innovation to $^{u}_{w}$ is estimated to be, at short-horizons, positive and statistically signiÞcant should be put into perspective: Since $^{u}_{w}$ explains essentially nothing of the variance of hours at any horizon, correctly capturing the impact in small samples is obviously di cult, and this result should therefore be quite heavily discounted.

Shocks to the currency/deposits ratio, on the other hand, played a non-negligible role not only for ln(1+) and ln(+), but also for inßation and especially M1 velocity, explaining, at the 10-year horizon, about half of the FEV of either series. The response of GDP to $^{n}_{w}$

and the 5-year rate, whereas the response of hours, consumption, and GDP is positive, and statistically signiÞcant at short horizons, and insigniÞcant in the long-run (for GDP, the long-run impact is borderline insigniÞcant).

Finally, the residual permanent inßation shock explains about one-Þfth of the FEV of inßation and the Federal Funds rate, and about half of the FEV of the 5-year rate, at all horizons, whereas it plays a negligible role for all other series. The response to $_{w}$ is, as expected, positive and permanent for the Federal Funds rate and the 5-year rate; it is borderline insigniÞcant for M1 velocity; it is negative and statistically sigiÞcant at the short horizons for the multiplier of M2-M1; and it is positive and statistically signiÞcant at the short horizons for hours, GDP, and consumption.

Counterfactual simulations The counterfactual simulations reported in Figure 11 point towards a uniformly negligible role played by either $^{l}_{w}$ $^{u}_{w}$ or $^{k}_{w}$ for all series other than hours. Consistent with the evidence reported in Figure 10, on the other hand, killing o $^{n}_{w}$ produces counterfactual paths which di er from the actual ones by non-negligible amounts for inßation, and especially M1 velocity. For all other series, on the other hand, the di erence is negligible.

Figure 12 reports the conterfactual paths obtained by jointly killin1.2(p)76i2.06y

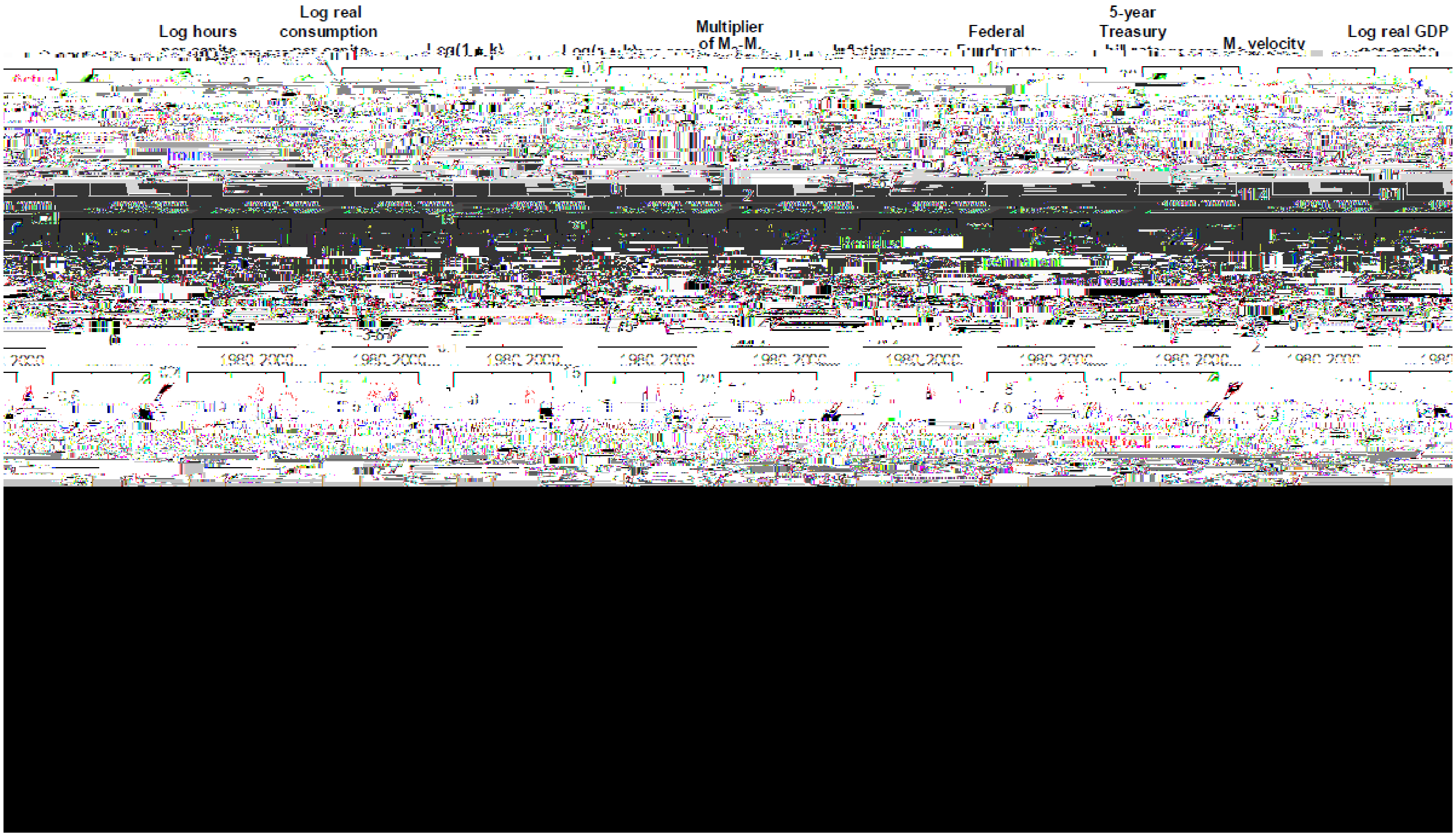

Figure 11 United States, 1959Q1-2008Q3: Counterfactual simulations killing off individual structural shocks, with 16-84 bootstrapped confidence bands

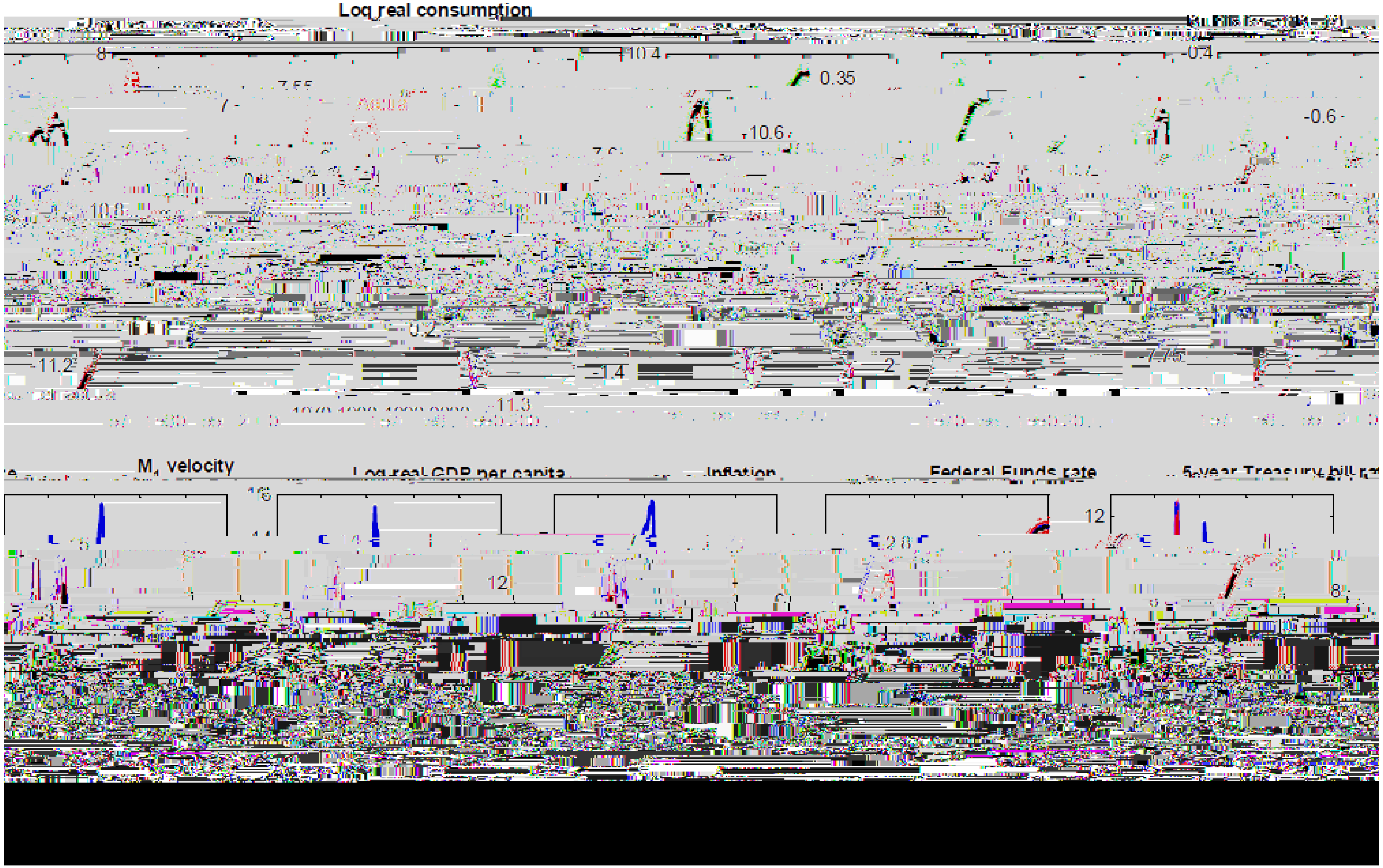

Figure 13 United States, 1959Q1-2008Q3: Counterfactual simulations jointly killing off the permanent shocks to the multiplier of M2-M1, with 16-84 and 5-95 bootstrapped confidence bands

M2-M1 around the time the Great Inßation, we still detect a non-negligible role for a non-monetary permanent inßation shock, which has the natural interpretation of a disturbance originating from the de-anchoring of inßation expectations following the

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A The Data

A.1 Monthly series for the period January 1919-December 1960

Seasonally adjusted series for currency held by the public, demand deposits, bank reserves, and M2 are from Tables A.1 and A.2 of Friedman and Schwartz (1963). We compute high-powered (i.e., base) money as the sum of currency held by the public and bank reserves. A seasonally adjusted series for the industrial production index is from the Board of Governors of the Federal Reserve System. A seasonally adjusted series for the CPI has been constructed by linking the seasonally adjusted CPI series for all urban consumers, all items (acronym is CPIAUCSL) from the U.S. Department of Labor: Bureau of Labor Statistics, which is available since January 1947, to the CPI all items series (NBER series 04128 from NBER Historical database), which is, originally, seasonally unadjusted, and we seasonally adjusted via ARIMA X-12. A seasonally unadjusted series for the discount rate of the Federal Reserve Bank of New York is from the NBER Historical database (acronym is M13009USM156NNBR). The seasonally unadjusted series for Moody's seasoned Baa corporate bond yield is Moody's. A seasonally unadjusted series for the index of yields of high grade public utility bonds for United States is from the NBER Historical database (acronym is M13025USM156NNBR). A seasonally unadjusted series for department store sales is from the NBER Historical database (acronym is M06F2BUSM350NNBR), and it has been seasonally adjusted via ARIMA X-12.

A.2 Quarterly series for the period 1959Q1-2008Q3

A monthly seasonally adjusted M2 series is from the St. Louis FED's website (acronym is M2SL). Monthly seasonally unadjusted series for the Federal Funds rate and the 5-year Treasury bill rate are from the St. Louis FED's website (acronyms are FEDFUNDS and GS5). A monthly seasonally unadjusted series for the St. Louis Source Base (SBASENS) is from the St. Louis Fed's website. The series has been seasonally adjusted via ARIMA X-12 as implemented in EViews. A monthly seasongregate has been kindly provided by Juan-Pablo Nicolini. SpeciÞcally, the series is equal to M1SL from the St. Louis FED's website (converted to the quarterly frequency by taking averages within the quarter) until 1981Q4, and it is equal to M1SL plus MMDAs for the period 1982Q1-2012Q4. As discussed by LucasJr. and Nicolini (2015), the rationale for including MMDAs (which were introduced in 1982) into M1 is that, although they have traditionally been classiÞed as part of the M2-M1 component, in fact, the economic function they perform is very similar to that performed

Appendix

Figure A.1 United States, 1959Q1-2008Q3: Impulse-response functions to the structural shocks, with 16-84 and 5-95 bootstrapped confidence bands (based on the model with 2 cointegration vectors)

Figure A.2 United States, 1959Q1-2008Q3: Fractions of forecast error variance explained by either of the structural shocks, with 16-84 and 5-95 boot

Figure A.3 United States, 1959Q1-2008Q3: Counterfactual simulations killing off the shocks to the M_1 multiplier, with 16-84 bootstrapped confidence bands (based on the model with 2 cointegration vectors)

Figure A.4 United States, 1959Q1-2008Q3: Counterfactual simulations jointly killing off the shocks to the M_1

Figure A.5 United States, 1959Q1-2008Q3: Counterfactual simulations jointly killing off the permanent shocks to the multiplier of M2-M1, with 16-84 and 5-95 bootstrapped confidence bands (based on the model with 2 cointegration vectors)

Figure A.6 United States, 1959Q1-2008Q3: Impulse-response functions to the structural shocks, with 16-84 and 5-95 bootstrapped confidence bands (based on the model with 4.0008 M00000 n-1.145

Figure A.7 United States, 1959Q1-2008Q3: Fractions of forecast error variance explained by either of the structural shocks, with 16-84 and 5-95 bootstrapped confidence bands (based on the model with 4 cointegration vectors)

Figure A.8 United States, 1959Q1-2008Q3: Counterfactual simulations killing off the shocks to the M_1 multiplier, with 16-84 bootstrapped confidence bands (based on the model with 4 cointegration vectors)

Figure A.9 United States, 1959Q1-2008Q3: Counterfactual simulations jointly killing off the shocks to the M_1 multiplier, with 16-84 and 5-95 bootstrapped confidence bands (based on the model with 4 cointegration vectors)

Figure A.10 United States, 1959Q1-2008Q3: Counterfactual simulations jointly killing off the permanent shocks to the multiplier of M2-M1, with 16-84 and 5-95 bootstrapped confidence bands (based on the model with 4 cointegration vectors)