Corporate Bond Yields in the Transmission Mechanism of Monetary Policy

Hans–Martin Krolzig & Isaac Sserwanja*
School of Economics, Keynes College, University of Kent, Canterbury

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Abstract

Financial markets are a major channel through which monetary policy influences the economy. The policy transmission mechanism in the United States is analysed within a cointegrated vector error-correction model framework, with emphasis on the role played by the corporate bond market. The corporate bond yield is found to significantly react to monetary policy shocks, both directly and indirectly. It is also a significant transmitter of these shocks to the general economy via its effect on inflation and output. Impulse response analysis shows corporate bonds reacting faster and deeper to a contractionary monetary shock than Treasuries, making the corporate-Treasury bond yield spread a potential target in implementing monetary policy.

Keywords: Bond yields; Monetary Policy; Cointegration; Error-correction.
JEL classification: E4; E52; G1.

1 Introduction

The transmission mechanism of monetary policy has been extensively investigated in the literature. Nevertheless, most studies accord little or no consideration to the interaction between monetary policy and corporate bonds. One possible explanation is that this interaction can largely be inferred from looking at government bond yields. However, for a deeper understanding of this interaction, one has to look beyond the intermediated link by Treasuries. Not least because corporate bond yields capture credit market conditions as faced by non-government market participants and so provide a more direct pulse of the economy. Government bond yields on the other hand do not represent (direct) borrowing costs in the business world but the cost of borrowing by the government. This investigation therefore looks at the monetary policy transmission mechanism and the corporate bond market in more detail beyond the association with Treasuries. The emphasis is on long-term relationships. The two are brought together under the question: what role does the latter play in the former? In answering this question, we test the hypothesis that corporate bond market interaction with monetary policy does not always mirror that of government bonds. The literature is rich in the approaches employed in discussing monetary policy

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and bond markets. The VAR methodology is a popular framework for type of analysis, with the federal funds rate as the monetary policy proxy for US focussed analyses. This study retains this this tradition.

Does monetary policy affect corporate bonds? Are these effects stronger, do they last longer than on government bonds? To what extent do these effects differ from what is observed from government bond yields? What do corporate bonds tell us about the economy that government bonds do not? In a way, this paper attempts to re-focus the bond yield - monetary policy interaction question away from government bonds and towards corporate bonds which are more sensitive to the cost of borrowing by businesses. These questions are explored by analysing the US corporate and government bond markets and how they respond to US monetary policy. Long-maturity interest rates are linked to the federal funds rate via the expectations hypothesis of the term structure of interest rates and to similar-maturity Treasuries via the risk premium. According to the primary interest rate channel[1] of the transmission mechanism of monetary policy, if one assumes price stickiness, then contractionary policy in the form of an increase in the federal funds rate results in higher real interest rates. The higher user cost of capital leaves firms facing higher bond yields, leading to a fall in investment and ultimately lower aggregate demand and output.

An early attempt in Cook and Hahn (1989) found evidence of strong movements in US Treasury yields in response to changes in the federal funds rate target in the 1970s. However, the response falls with increasing maturity. Kuttner (2001), Gurkaynak et al. (2005), and Beechey (2007) among others, have also linked shocks to the federal funds rate to movements in US market interest rates, in some cases including maturities as long as 30 years. Using structural VAR-based analysis, Ludvigson et al. (2002) find the federal funds rate effect on asset values significant but transitory. On the other hand, Roley and Sellon (1995) found that for the 1987-1995 period, the long-term government bond rate response to an increase in the federal funds rate target was statistically insignificant. This suggests that long rates response to policy varies over the business cycle and depends on market perceptions of the persistence of policy. Further VAR analysis in Edelberg and Marshall (1996) and Evans and Marshall (1998), respectively based on 1947-1995 and 1964-1995 datasets, also found no significant response to policy from long-term rates of maturities above 10 years, with Evans and Marshall (1998) asserting that long-term rates are unaffected by monetary policy actions.

More recently, Wright (2011) also employs a structural VAR to investigate the effect of non conventional monetary policy shocks - quantitative easing, for example - on long-term interest rates when the federal funds rate is near the zero lower bound. Policy shocks are found to have a significant effect on 10-year Treasuries, Aaa-rated, and Baa-rated corporate bond yields. However, the effect on corporate bond yields is a little more than half that on Treasuries, with all effects dissipating within a few months. That Wright (2011) finds the effect on corporate bonds less than on Treasuries contradicts our results. However, this could be explained by two factors. Firstly, our study uses the federal funds rate as the policy instrument while Wright (2011) has quantitative easing (QE) which involved buying government bonds, and Federal Reserve Board announcements on the future path of policy - keeping the federal funds rate near the zero lower bound for the foreseeable future - as the main policy instruments. Secondly, Wright (2011) focusses on a two-year period beginning in December 2008, a time when credit markets were tightening and not functioning normally. This study is based on over forty years of data.

Gilchrist et al. (2009) find that for the 1990-2008 period, credit market shocks were a significant con-

[1] See Bernanke and Blinder (1992), Taylor (1995), and Kuttner and Mosser (2002).

tributor to economic fluctuations. They estimate that real interest rate shocks in US corporate bond markets accounted for 30% of the forecast error variance decomposition in US economic activity. Spreads on longer maturity bonds from firms with higher credit ratings were found to be rich in information relevant for improving forecasting. The forecast usefulness of corporate bond yields is also pointed out in Gilchrist and Zakrająsek (2012), in which a credit spread with significant predictive power for future economic activity is constructed out of individual corporate bond prices. Meeks (2010) looks at evidence from corporate spreads and defaults to study the link between credit market shocks and output fluctuations. Among other findings, credit shocks were found significant in accounting for increases in bond spreads both before and during the 2007-09 financial crisis. Increases in bond spreads were however mainly due to higher default risk, perhaps as a result of the crisis. However, very little of the variation in output appears to be linked to variation in bond spreads. Cenesizoglu and Essid (2012) look at the monetary policy effects on credit spreads over the business cycle, finding that lower-rated corporate bond yields widen relative to high rating yields following a contractionary policy shock. This finding is in line with theories on imperfect capital markets. From this non-exhaustive review of the literature, it is clear that corporate bonds are an important asset class. They paint a picture of prevailing conditions in credit markets, and are information-rich assets, at least for the purposes of forecasting economic activity.

The rest of the paper is organised as follows. §2 discusses some of the economic theory underpinning the monetary policy-bond market relationship, as well as introducing the data. Next, §3 presents the econometric methodology, implements the empirical analysis, and presents the results. This section presents theory details on the VAR model and relevant variants that are employed in formulating a congruent model to answer the title question. Implications of the final model are discussed in §4, with respect to responses to a monetary policy impulse. §5 concludes.

2 Data and the theoretical framework

An empirical analysis of the role of corporate bond in the transmission mechanism of monetary policy needs the correct variables and a theoretical framework to analyse the data. At least two of the variables are naturally the corporate bond yield and the effective federal funds rate, the monetary policy instrument. The effective federal funds rate closely tracks the federal funds rate target, the policy rate set at meetings of the Federal Open Market Committee. Inflation is included in the dataset due to its importance to investors as well as being a target of US monetary policy. The forth variable is inflation expectations. Including both inflation and expectations of inflation is a notable departure from the literature justified by the need to mitigate the price-puzzle problem where inflation rises in response to a contractionary policy shock. Manufacturing capacity utilisation proxies the business cycle. It would have been preferable to use industrial capacity utilisation but the available series is too short. An equal-maturity government bond is included to compare corporate and government bond response to monetary policy shocks, bringing the total of variables to six.

2.1 Data properties

Quarterly time series data are collected on the variables in table 1 for the period 1960Q1-2012Q4. The government bond yield is the 30-year Treasury constant maturity rate. Where 30-year Treasury data are missing, splicing is used in conjunction with the 20-year Treasury series to approximate the missing data points, see appendix for details on the splicing procedure. The corporate bond is Moody's Baa-rated seasoned corporate bond yield with a maturity as close to 30 years as possible.

Table 1 *Time series and model variables*

Time series	Description	Code: Source	Variable
π_t^e	Expected change in prices during the next 12 months	Table 19: Survey of Consumers, Reuters/University of Michigan	π_t^e
P_t	GDP implicit price deflator, Index 2005=100, SA	GDPDEF:FRED database	π_t
Y_t	Capacity utilisation (%), Manufacturing (SIC) SA	G17/B00004: Federal Reserve System	y_t
I_t	Effective federal funds rate, NSA	FEDFUNDS: FRED database	i_t
$GS30_t$	Government bond yield, NSA	GS30: FRED database	r_t
Baa_t	Corporate bond yield, NSA	Baa: FRED database	b_t

$\pi_t = \Delta log(P_t)400$. All series are transformed based on equation $x = log(1+X/100)100$ for variable X, with the exception of inflation and capacity utilsation.

A visual inspection of the time series in figure 1 which is juxtaposed over National Bureau of Economic Research (NBER) recession dates reveals some properties about the data. As expected, output - represented by capacity utilisation scaled by 100 - is falling during all eight recessions covered by the data. The end of the recession is marked by a pick up in output. Capacity utilisation appears trend-stationary, with a downward trend indicating falling US manufacturing capacity use over the last five decades. One possible explanation for the trend is the net growth in US imports over the sample period, resulting in falling market share for US manufacturers.

The middle panel of figure 1 shows all three interest rates are trending upwards for the first twenty years, then followed by a downward trend in the early 1980s. This change in interest rate trajectory is explained by a re-alignment of US monetary policy by the Federal Reserve Board (FRB) under chairman Paul Volcker to combat high inflation in the 1970s. We also see that the federal funds rate is always lower at the end a recession than at its beginning.[2] In other words, monetary policy is accommodative during recessions. However, there are also cases when monetary policy has reacted in a way not consistent with the conventional view of supporting the economy in a recession. For example, the rise in the federal funds rate during the 1974Q1-1975Q1 and 1981Q4-1982Q4 recessions could arguably have prolonged these downturns, or even created the second one. It is to the credit of the Federal Reserve Board that such counter-intuitive movements in the federal funds rate have not been seen since the inflation-bursting late 1970s and early 1980s. Both bond yields track each other closely, suggesting that where monetary policy targets the government yield curve, this effect is transmitted to corporate bonds as well. In addition, both yields move in the same direction for most of the time. The exceptions are periods when investor flight-to-safety tends to drive yields on Treasuries relatively lower. Such periods are synonymous with recessions. A notable divergence is observed during the 2008Q1-2009Q2 recession with the Treasury yield falling while the corporate bond yield rises due to the massive flight to safety resulting from the 2007-09 financial crisis. It is as if the credit crunch crisis incapacitated the ability

[2]Even more striking in the data is that almost every recession is preceded by a negative term spread, $r - i < 0$, during the last five quarters or so. Only the 1960Q3-1961Q1 recession is the exception. Is the negative liquidity premium forecasting recessions?

of monetary policy to link Treasuries and corporate bonds. The last recession notwithstanding, these two bond yields track each other closely over the sample and are obvious candidates for cointegration. The behaviour of expected inflation and actual inflation appear theory-consistent as one would expect. Inflation expectations lead actual inflation but the two variables also move clearly together, and so make the second cointegration candidate. Prices grow at a lower rate at the end than at the beginning of recessions. There is, however, price stickiness evidenced in continued price growth well after recessions have started. The results of Federal Reserve Board efforts to reduce inflation are also evident in lower and less volatile inflation post early 1980s. Broadly speaking, recessions coincide with falling inflation except the stagflation of 1974Q1-1975Q1, which might have its roots in the 1973 oil crisis. Recessions also coincide with falling output, and a falling federal funds rate, generally speaking. While one would expect coincidence between falling inflation and rising bond yields during recessions, this is not observed in the data. Bond yields rise or fall depending on the recession. It is further observed that variables rise gradually but fall more sharply around recessions.

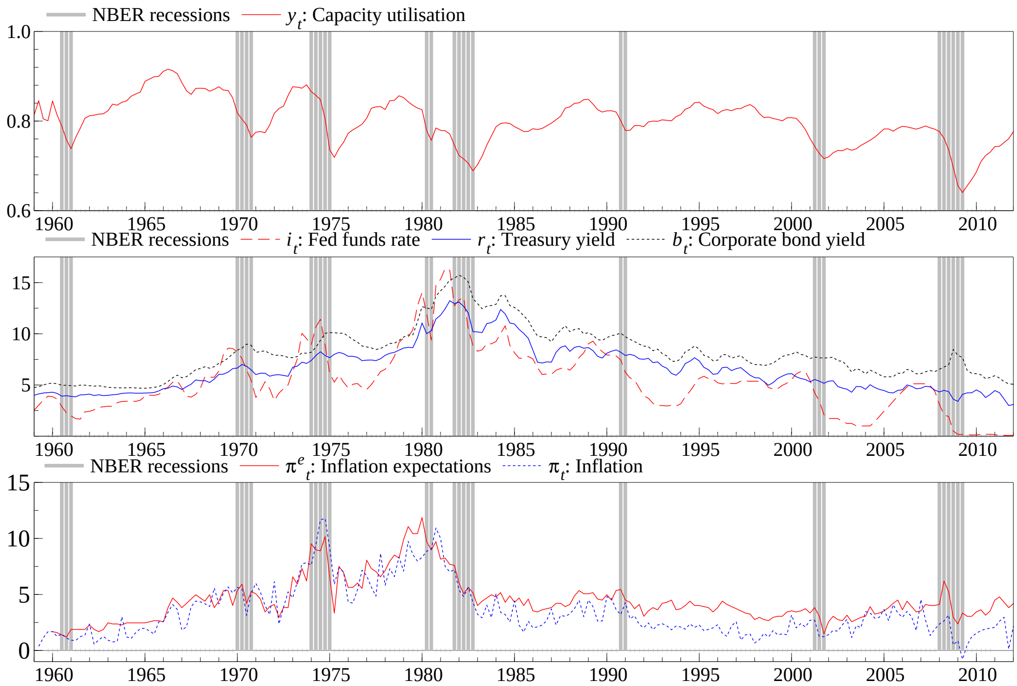

Figure 1 *Variable time series and NBER recessions*

Time series unit root profiles

Every series is tested for the presence of a unit root using the Augmented Dickey-Fuller (ADF) test. This avoids estimating spurious regressions given that part of the estimation is implemented using the ordinary least squares (OLS) method. Test results in table 2 confirm that capacity utilisation is indeed trend-stationary while all other variables have unit roots. The five variables with unit roots become stationary after first-lag differencing. This means our 6-dimensional system is integrated of order one, $I(1)$.

Table 2 *Augmented Dickey-Fuller test for a unit root*

	$\Delta\pi_t^e$	$\Delta\pi_t$	Δy_t	Δi_t	Δr_t	Δb_t
μ	0.37*	-0.53*	8.37**	0.29*	0.21*	0.12
	(0.013)	(0.043)	(0.000)	(0.012)	(0.042)	(0.166)
γ	-0.001	0.001	-0.005**	-0.002	-0.001*	-0.0007
	(0.194)	(0.354)	(0.001)	(0.053)	(0.043)	(0.097)
β	0.92	0.93	0.90	0.93	0.98	0.999
Δx_{t-j} lags	4	3	3	5	5	1
	(0.041)	(0.009)	(0.021)	(0.007)	(0.048)	(0.000)
t-adf (β)	-2.65	-1.94	-5.14**	-3.43	-1.40	-0.50
5 % CV	-2.88	-2.88	-3.43	-3.43	-3.43	-1.94
Unit Root?	Yes	Yes	No	Yes	Yes	Yes

Sample 1961Q2-2012Q4. Variable x_t is tested for the presence of a unit root based on equation $\Delta\, x_t = \mu + \gamma\, t + \beta\, x_{t-1} + \sum_{j=1}^{p} \phi_j\, \Delta\, x_{t-j} + \varepsilon_t$. $\gamma = 0$ is imposed if insignificant. The null hypothesis, $H_0 : \beta = 0$ which implies presence of a unit root is rejected when t-adf $<$ CV.

2.2 Some economic theories on variable relationships

The theoretical framework is motivated by the choice of variables as well as the need to understand the long-term behaviour of the system. Relationships from economic theory that are expected to hold in the data include the term structure of interest rates defining the federal funds rate and the Treasury yield, the risk premium defined by the yield difference between corporate and government bond yields, the spread between actual and expected inflation and finally the Fisher hypothesis defining the real (short-term) interest rate as the nominal federal funds rate less expected inflation. These relationships (partly) motivate the restrictions imposed to identify the long-run structure of the system through cointegration analysis.

The expectations hypothesis of the term structure of interest rates (EHT)

The 30-year maturity of the bonds in this study is dictated by the availability of data on corporate bond yields. These data start in 1960 and covering several business cycles, which is in line with the perspective of cointegration analysis focus on the long-run. In addition, investment involves a time horizon therefore economic activity (long-term investors, in particular) is sensitive to long-term interest rates such as those on corporate bond yields, Treasuries and mortgages. US monetary policy sets the level of the federal funds rate, which is linked to longer maturity rates via the yield curve. The EHT links short-term rates and long-term rates. The EHT posits that the long-term interest rate is essentially the average of expected short-term interest rates over the duration of the long-term bond, plus a risk premium. This definition is re-stated in compact form in (1).

$$r_t = \frac{1}{T}\sum_{j=0}^{T-1} \mathsf{E}_t i_{t+j} + \psi_t. \tag{1}$$

r_t is the yield on a long-term bond, i_t is the time t expectation of the short-term interest rate at time $t+j$, and ψ_t the risk premium. How policy affects long-term interest rates largely depends on how that policy affects market expectations of future short-term rates and the liquidity premia. Empirical evidence as to the validity of the EHT is mixed but Kuttner (2001) finds interest rate response to the target rate consistent with EHT theory. Where the EHT holds, it is expected that there exists a statistically testable long-run relationship between short and long-run interest rates. This hypothesis is tested and confirmed below.

The risk premium

Exposure to risk matters to investors and this is especially true when the investment choice involves financial assets of differing risk profiles. The Treasury yield less corporate bond yield spread defines a risk premium. This definition is narrow but suffices for our use of the spread. Elton et al. (2001) and Houweling et al. (2005) argue that other factors in addition to risk explain this spread. These include, but are not limited to, the tax structure and a liquidity premium. we specify the risk premium in (2), where b_t is the corporate bond yield, r_t the government bond yield and ϕ_t the risk premium.

$$b_t = r_t + \phi_t. \tag{2}$$

The term structure of inflation

Another candidate for a cointegrating relation is the inflation - inflation expectations differential. If economic agents are rational, then they should not consistently err by over-estimating or under-estimating inflation. An implication of this is that the actual inflation - inflation expectations differential is stationary, a hypothesis confirmed by the ADF test in table 3. Because the two measures are not based on coincident periods, we do not think of $(\pi - \pi^e)$ as representing rational expectations. Rather, it can be a random walk with future inflation forecast by present inflation.

The Fisher hypothesis (FH)

The Fisher hypothesis defining the real (short-term) interest rate is also tested for stationarity. It decomposes the nominal federal funds rate into a real part and the expected rate of inflation. Given by (3), i_t is the nominal federal funds rate, ρ_t the real short term rate, and π_t^e expected inflation. Unit root tests in table 3 confirm all four relations as stationary and therefore, strong contenders for long-run restrictions in the analysis that follows.

$$i_t = \rho_t + \pi_t^e. \tag{3}$$

Testing the stationarity of variable spreads has the advantage of potentially aiding system identification, especially the long-run relations. Spreads such as the risk premium and the liquidity premium, often define stable long-run relationships. Unit root tests for a selection of spreads are summarised in table 3.

3 Econometric modelling

The analysis in this is focussed on the title question of characterising what corporate bonds contribute to the transmission of monetary policy. We begin with a general model specification from which a parsimonious structural model is retrieved. The implementation follows the data-driven approach in Krolzig (2003) which may be classified under 'general-to-specific' modelling, see Krolzig (2001). The

Table 3 *Unit root tests on spreads*

Spread	$(r-i)_t$	$(b-r)_t$	$(\pi-\pi^e)_t$	$(i-\pi^e)_t$
μ	0.15^*	0.16^{**}	-0.05	0.17
	(0.04)	(0.00)	(0.73)	(0.09)
γ	0.002	0.0005	-0.003	-0.0007
	(0.19)	(0.19)	(0.07)	(0.63)
β	0.81	0.91	-0.36	0.93
Δx_{t-j} lags	1^*	1^*	3^{**}	3^*
	(0.00)	(0.01)	(0.00)	(0.00)
t-adf(β)	-3.96^{**}	-3.18^*	-2.04^*	-2.01^*
5%*CV*	-2.88	-2.88	-1.94	-1.94
Unit Root?	No	No	No	No

unrestricted VAR model in (4) is specified for our six variables of interest: inflation expectations (π_t^e), inflation (π_t), capacity utilisation (y_t), the federal funds rate (i_t), the Treasury bond yield (r_t), and the corporate bond yield (b_t). The model then undergoes statistical testing to systematically decide on the lag-length, deterministic terms such as a trend and/or a constant, and other misspecification tests.

The Johansen (1991) trace test for cointegration is run to determine the number of cointegrating relations in the system. These stable, linear relationships characterise the long-run structure of the system. By estimating a cointegrated VAR model with economically meaningful restrictions on the relevant parameters, the cointegration relations are identified. We see, for example, that the corporate bond yield - Treasury yield risk premium is accepted as one of five stable relations. This cointegration analysis is in the spirit of Juselius (2006). The next stage involves identification of the short-term structure of the model and this involves determining the instantaneous causal structure among variables. This is important for retrieving the structural VECM (SVECM) from the estimated reduced-form model. Following Demiralp and Hoover (2003), the PC algorithm - a causal search algorithm - is run to determine the contemporaneous causal order among variables. These contemporaneous relationships between variables based on this ordering are added to a re-specified system and a structural model is estimated. The SVECM model undergoes both system and single-equation reduction until a congruent parsimonious SVECM (PSVECM) is obtained as the final representation of the data generating process. The single-equation reduction process involves dropping statistically *in*significant parameters from the SVECM. It is the resulting PSVECM that is used in impulse response analysis to inform on how the system responds to shocks.

3.1 The VAR model

A general specification of a K-dimensional reduced-form unrestricted VAR model of order p is given by (4) where $\boldsymbol{x}_t = (\pi_t^e,\, \pi_t,\, y_t,\, i_t,\, r_t,\, b_t)'$, $\boldsymbol{\mu}$ is a $K \times 1$ vector of intercepts, γ a $K \times 1$ vector of trend-coefficients, $\boldsymbol{A}_j$ a $K \times K$ matrix of coefficients. ε_t is a Gaussian vector of white noise processes each with zero mean and $\boldsymbol{\Sigma}$ their variance-covariance matrix. This VAR specification includes p lags.

$$\boldsymbol{x}_t = \boldsymbol{\mu} + \boldsymbol{\gamma} t + \sum_{j=1}^{p} \boldsymbol{A}_j \boldsymbol{x}_{t-j} + \boldsymbol{\varepsilon}_t, \quad \boldsymbol{\varepsilon}_t \sim \mathsf{NID}(\boldsymbol{0}, \boldsymbol{\Sigma}). \tag{4}$$

We need the correct lag length p to estimate (4) and of course, the estimated model should have autocorrelation-free residuals. (Enders (2010), p.397) suggests an initial lag length of $T^{1/3}$ with T being the number of observations. With just over 200 observations, we therefore estimate (4) with six lags for a start. The lag length can be selected based on an information criterion of the resulting model, for example the Akaike (AIC), Schwarz (SC) or Hannan-Quinn (HQ) information criterion. From table 4, 6 lags are chosen by the AIC, 2 lags by the HQ, and 1 lag by the SC. The VAR(1) residuals are severely autocorrelated. Also, estimating a VAR(1) would imply a VECM(0) with no short-run dynamics terms. A VAR(2) is not any better. It detects only four cointegration relations in the data yet five are expected, the four relations in section 2.2 and a trend-stationary capacity utilisation. The VAR(2) also has severely serially correlated residuals. This leaves a VAR(6), based on which the Johansen trace test (details shortly) correctly identifies the expected five cointegrating relations. Furthermore, the LM test ($\chi^2(36)$) null that the VAR(6) residuals are not serially correlated is not rejected with 0.416 probability. The six lag choice is also supported by the likelihood ratio (LR) test on lag selection in table 4.

Table 4 *Lag-length determination*

Lag	LR	AIC	SC	HQ
0	NA	21.07644	21.27029	21.15484
1	2120.546	10.71612	11.49155**	11.02973
2	139.0225	10.34156	11.69856	10.89**
3	86.11828	10.22808	12.16665	11.0121
4	60.68511	10.24045	12.76059	11.25968
5	62.00342	10.23363	13.33533	11.48806
6	65.76133**	10.192**	13.87498	11.68135

Sample: 1961Q2-2012Q4. **chosen lag length at the 5% significance level.

In addition to vector autocorrelation, the VAR(6) residuals from all six constituent equations are also individually tested for residual autocorrelation of up to order six. Results in table 5 (p-values are in brackets) are clear that autocorrelation is not a problem except for inflation residuals. The autocorrelation problem could be remedied by increasing the lag length, introducing new variables or both steps. However, either remedy would substantially increase the number of model parameters to be estimated and thus reduce the degrees of freedom. It is therefore decided to proceed with the current VAR(6) specification, with the intention of ensuring that the final parsimonious model has autocorrelation-free residuals.

Table 5 *Individual equation autocorrelation tests*

Test		π_t^e	π_t	y_t	i_t	r_t	b_t
AR 1-6:	$F_{(6,162)}$	2.25* (0.0440)	4.08** (0.0008)	0.93 (0.4782)	0.87 (0.5189)	0.64 (0.6977)	0.98 (0.4438)

**(*) reject the null of non-autocorrelated residuals at the 5%(1%) significance level.

3.2 Cointegration

3.2.1 Testing for cointegration

We know from table 2 that vector series $\boldsymbol{x}_t = (\pi_t^e, \pi_t, y_t, i_t, r_t, b_t)'$ is $I(1)$, with capacity utilisation being trend-stationary and the other level variables having unit roots. These facts, together with the evident co-movement among variables in figure 1, are strongly suggestive and supportive of the cointegration analysis that follows. Here, cointegration between variables means the existence of linear, stable, long-run relationships between (some of) those variables. The number of these relations in our 6-dimensional system is determined using the Johansen (1991) trace test based on equation (5).

$$\Delta \boldsymbol{x}_t = \boldsymbol{\alpha}(\boldsymbol{\mu} + \gamma t + \boldsymbol{\beta}' \boldsymbol{x}_{t-1}) + SR + \varepsilon_t, \quad \varepsilon_t \sim \mathsf{NID}(\mathbf{0}, \boldsymbol{\Sigma}). \tag{5}$$

In 5, $\boldsymbol{\mu}$ is a vector of intercepts and γ a vector of trend-coefficients both of which are restricted to the cointegration space. $\boldsymbol{\alpha\beta}' = \boldsymbol{\Pi}$ is the matrix characterising the long-run structure of the system with $\boldsymbol{\beta}'\boldsymbol{x}_t$ as the cointegrating relations. SR are short-run dynamics. The trace test works by determining the rank, r, of matrix $\boldsymbol{\Pi}$. There are three possibilities. (i): $r = 0$ corresponding to no cointegration relations in the system. With $r = 0$ and $\boldsymbol{x}_t \sim I(1)$, a VAR in first-lag differences is sufficient to model the data. (ii): $0 < r \leq K-1$ is a reduced-rank with r long-run relations. This second result suggests that a VECM specification with cointegration relations given by $\boldsymbol{\beta}'\boldsymbol{x}_t$ (details below) is a valid representation of the data generating process. Finally, (iii): $r = K$ which is a full rank, implies that all vectors in $\boldsymbol{\Pi}$ are linearly independent and therefore no cointegration. For $0 \leq r \leq K-1$, the null hypothesis $H_0 : rank(\boldsymbol{\Pi}) \leq$ r is tested against the alternative hypothesis of $H_1 : rank(\boldsymbol{\Pi}) >$ r. The sequential testing stops when H_0 cannot be rejected, with r designated the number of cointegrating relations. The trace in table 6 finds five cointegration relations at 5% significance level, thereby confirming our earlier hypothesis. We will identify all five relations in the following section but first, we will carry out a robustness check on the long-run structure by checking the stability of the five relations.

Table 6 *Johansen trace test for cointegration*

$rank(\Pi) \leq$ r	eigenvalue	trace test	p-value
0	0.199	146.18	(0.000) **
1	0.140	100.58	(0.005) **
2	0.112	69.552	(0.014) *
3	0.085	45.185	(0.027) *
4	0.073	26.971	(0.034) *
5	0.053	11.323	(0.078)

Stability of cointegration relations

According to Hansen and Johansen (1999), testing for the rank constancy is equivalent to testing for cointegration stability. To this effect, the system is estimated recursively and the five eigenvalue pathes analysed. Where these pathes are deemed stable over the sample, the rank, and therefore the cointegration relations are judged stable. Figure 2 plots the eigenvalue pathes for the cointegrating relations produced under recursive estimation of the system. These are estimated under the R-form whereby only long-run parameters are re-estimated with short-run parameters fixed to their full-sample values. The first 90 observations are used for initialisation. All five eigenvalues are fairly stable and consistently above zero, suggesting a stable rank and hence stable cointegration relations. The trace test was also applied to smaller samples from 1961Q3-2006Q3 upwards and still, five cointegration relations were found. It seems that even the 2007-09 financial crisis and the related Great Recession did not significantly alter the long-run structure. With these robustness results, we are doubly confident that there are five stable, long-run linear relationships in the data which we now proceed to identify.

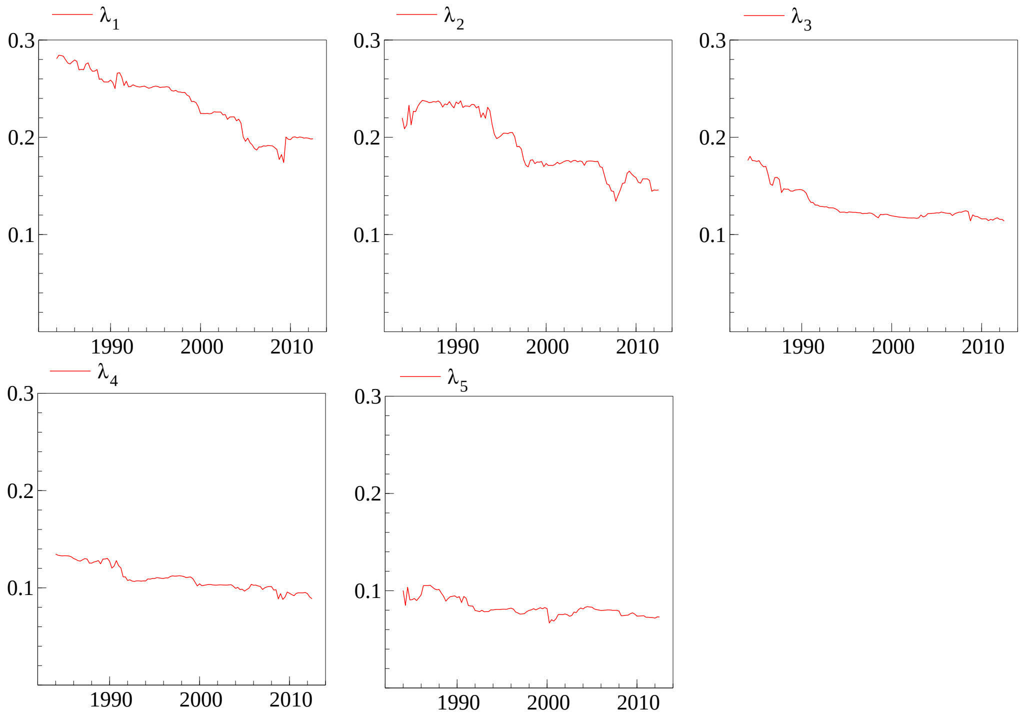

Figure 2 *Eigenvalue pathes for cointegration relations*

3.2.2 Identification of cointegration relations

Augmenting (5) with short-run dynamics gives the cointegrated VAR model in (6).

$$\Delta \boldsymbol{x}_t = \boldsymbol{\alpha}(\boldsymbol{\mu} + \boldsymbol{\gamma} t + \boldsymbol{\beta}' \boldsymbol{x}_{t-1}) + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \boldsymbol{x}_{t-j} + \varepsilon_t, \quad \varepsilon_t \sim \mathsf{NID}(\mathbf{0}, \boldsymbol{\Sigma}). \tag{6}$$

Cointegration relations are identified by imposing (over)identifying restrictions on the long-run matrix $\boldsymbol{\beta}$ to obtain $\boldsymbol{\beta}^{\boldsymbol{r}}$. The restrictions imposed here are the economic relationships motivated by the discussion in section 2.2 above. The restricted system generates a new value for the log-likelihood function which can be compared to the one generated under no restrictions using standard asymptotic critical values to determine the statistical acceptability of restrictions.

The five cointegration relations, which together identify the long-run structure of the system, could be identified as the liquidity premium, the risk premium, capacity utilisation, the term structure of inflation, and the Fisher hypothesis relation.[3]

1. *Liquidity premium:* $(r-i)_t^* \sim I(0)$

[3] (*) represents a constant and/or trend, when significant in a relation.

Relation one given by the first row of β^r defines the spread between the 30-year maturity government bond yield and the federal funds rate. These are interest rates faced by the US government so should practically incorporate no risk of default. We thus call this spread the liquidity premium to capture the difference between the two positions on the US government's yield curve. This term spread was found to be $I(0)$ and is individually accepted as a long-run relation with 0.95 probability.

2. *Risk premium:* $(b-r)_t^* \sim I(0)$

   The long-run relation in row two is of particular interest because it involves the corporate bond yield, a variable whose role in the transmission of monetary policy we are investigating. We call relation two the risk premium. It is extra compensation for investing in a Baa-rated corporate bond as opposed to a government bond of equal maturity. We use Baa rather than Aaa-rated corporate bonds because Aaa rated bonds tend to behave like government bonds for prolonged periods, although this behaviour decouples during periods of economic contraction such as recessions. Baa is probably more representative of the corporate bond market as whole. As stated above, this definition of the risk premium is quite narrow since it ignores other factors such as taxation that partly explain this spread. For our purposes however, the risk premium classification suffices and this relation is on its own accepted with a 0.18 p-value.

3. *Capacity utilisation:* $y_t^* \sim I(0)$

   The third relation is capacity utilisation which we earlier found to be stationary around a trend. Deviations from trend could be construed as deviations from trend-output and therefore a proxy for the output gap. Restrictions specifying capacity utilisation as one of the five stationary relations are imposed and comfortably accepted with 0.20 probability.

4. *Actual inflation and expected inflation spread:* $(\pi-\pi^e)_t^* \sim I(0)$

   Relation four captures the dynamics of the interaction between actual and inflation expectations. We assume economic agents to be rational, that is, they do not persistently make systematic mistakes. This implies a stable relation between the two inflation measures. With a 0.26 p-value, the restriction for the long-run stability of this relation is also not rejected.

5. *Fisher hypothesis:* $(i-\pi^e)_t^* \sim I(0)$

   The final cointegrating relation is the Fisher hypothesis defining the real (short-term) interest rate. The Fisher hypothesis is easily accepted with 0.89 probability. $(i)-(v)$ imply neutrality of monetary policy in the long-run. Nominal variables are $I(1)$. Real variables, including the real interest rate, and the term structures of interest rates and inflation, are $I(0)$. The five cointegration vectors are plotted in figure 3. This supports the specified structure as a valid representation of the long-run of the DGP.

The β^r matrix whose rows represent the five cointegration relations is shown in Table 7.The overidentifying restrictions defining the five relationships are jointly accepted with 0.5375 probability. Also, each vector passes the LR test individually.

Table 7 *Detailed β^r matrix*

	π_t^e	π_t	y_t	i_t	r_t	b_t	Constant	trend	LR test	p-value
H_1	0	0	0	-1	1	0	0.38 (0.37)	-0.01 (0.003)	0.004	0.9506
H_2	0	0	0	0	-1	1	-1.10 (0.21)	-0.005 (0.002)	1.8211	0.1772
H_3	0	0	1	0	0	0	-87.85 (1.71)	0.07 (0.01)	1.62	0.2034
H_4	-1	1	0	0	0	0	1.03 (0.23)	0	2.66	0.2643
H_5	-1	0	0	1	0	0	-6.12 (1.12)	0.04 (0.01)	0.021	0.8896
$\mathsf{H}_1,\ldots,\mathsf{H}_5$									5.049	0.5375

Standard errors for deterministic term are in brackets.

3.2.3 Equilibrium-correction and weak exogeneity

The specified long-run relations will inevitably be shocked into disequilibrium during the course of time. A natural question to ask therefore is, what happens for such a relation to get back to its equilibrium path? Which variables adjust to restore cointegration? Five cointegration relations in a 6-dimensional system mean that at most one variable is expected to be weakly exogenous for the long-run. This is the variable which does not adjust to restore equilibrium. It is determined by sequentially testing whether coefficients in a particular row of the loading matrix, α, can be restricted to zero. Test results in the final column of table 8 suggest inflation or more strongly, inflation expectations. We expect one weakly exogenous variable. The 58.7% p-value for the joint test strongly suggests inflation expectations are weakly exogenous. However, this is contradicted by the individual test results (row 2 of table 8) which have inflation expectations reacting to capacity utilisation. At this stage, weak exogeneity test results are preliminary until the final model is specified. Indeed, inflation expectations will be shown to adjust to the inflation - expected inflation relationship in the final PSVECM, and therefore not weakly exogenous.

The long-run matrix, $\mathbf{\Pi}$

Matrix $\mathbf{\Pi}$ in table 9 gives a preliminary view of how the system responds to movements in variables between periods. Only inflation, the federal funds rate and Treasury yield equations have significant self-stabilising feedback effects. Inflation expectations only rise in response to an increase in output. Inflation, on the other hand, is sensitive to all but a shock to the federal funds rate. That inflation does not react to the federal funds rate is exactly what is expected, after all, monetary policy is forward-looking whereas this inflation measure is backward-looking. Output adjusts downwards after an increase in the federal funds rate in the previous quarter. While the federal funds rate is not reacting to (past) inflation or inflation expectations, it strongly reacts to both long-term bond yields, yields which are a good indicator of the future path of inflation. Monetary policy makers are thus forward-looking in reacting to future rather than past prices.

The federal funds rate also responds to movement in capacity utilisation. The change in the government bond yield between time t and $t-1$ is sensitive to the level of the corporate bond yield at time $t-1$. In contrast, the corporate bond yield is not reacting to the Treasury yield but to the federal funds rate only. The federal funds rate rises to cool an over-heating economy. As inflation falls, the yield on

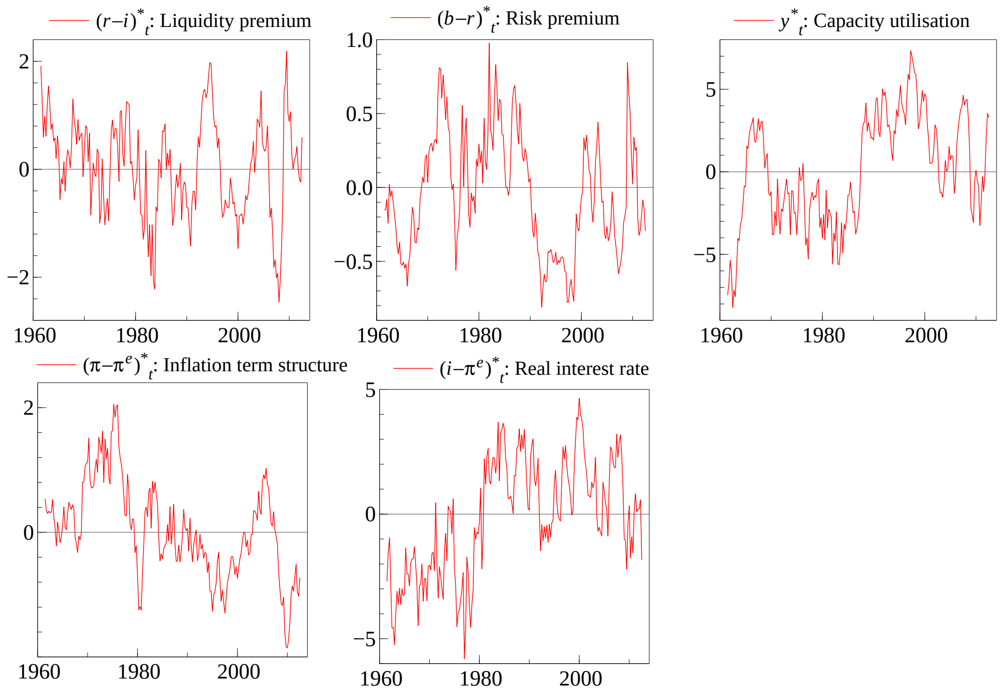

Figure 3 *Cointegration relations*

corporate bonds rises followed by an increase in Treasury yields. It is certainly intriguing that the risk premium affects the government bond but not the corporate bond yield. One would expect the Treasury yield to affect the corporate bond yield, and not the other way around. A possible explanation is that corporate bonds react faster than Treasuries to changes in credit market conditions. After all, the corporate bond market is populated by profit-maximising and comparatively shorter-term minded investors. On the other hand, the government bond market has private investors, the Federal Reserve Board and foreign central banks as major participants. The US government bond market can indeed be assumed to have more longer-term minded investors. The following analysis of contemporaneous effects indeed shows corporate bond yields influencing the Treasury bond yield, further supporting the result in table 9.

3.3 Contemporaneous relationships between variables

The long-run structure of the system is now well-specified so we proceed to determine the contemporaneous interactions between variables by analysing the correlation structure among estimated VECM(5) residuals. We estimate the SVECM once we determine these instantaneous effects. Such a structural model should improve innovation accounting results such as the impulse response analysis exercise to be implemented shortly. The β is super-consistent so we identify the contemporaneous structure conditional on the reduced-form VECM just estimated above. Put differently, the parameter set in matrix β^r are held fixed.

The causal pattern is determined with the PC algorithm, implemented in the computer package Tetrad IV. Developed in Pearl (2000) and Spirtes et al. (2001), this graph theoretic algorithm searches for the causal structure based on conditional independence patterns between variables. It is essentially a

Table 8 *Loading matrix, α, and weak exogeneity tests*

	$(r-i)^*_{t-1}$	$(b-r)^*_{t-1}$	y^*_{t-1}	$(\pi-\pi^e)^*_{t-1}$	$(i-\pi^e)^*_{t-1}$	Test for weak exogeneity ^{w}
$\Delta\pi^e_t$	-0.013 (0.083)	0.147 (0.187)	0.060** (0.025)	0.128 (0.084)	-0.051 (0.036)	3.74 (0.5873)
$\Delta\pi_t$	-0.184 (0.108)	0.527** (0.244)	0.061** (0.033)	-0.316** (0.110)	-0.161** (0.046)	9.67 (0.085)
Δy_t	0.145 (0.100)	-0.009 (0.226)	-0.011 (0.030)	-0.070 (0.102)	-0.080** (0.043)	15.43* (0.009)
Δi_t	0.097 (0.075)	0.455** (0.170)	0.067** (0.023)	-0.068 (0.077)	-0.076** (0.032)	15.51* (0.008)
Δr_t	-0.092** (0.040)	0.202** (0.090)	0.014 (0.012)	-0.028 (0.040)	-0.061** (0.017)	10.81 (0.055)
Δb_t	-0.158** (0.035)	-0.063 (0.078)	-0.001 (0.010)	-0.030 (0.035)	-0.048** (0.015)	20.90** (0.001)

Standard errors are in brackets. ^{w} $\chi^2(5)$ test statistics and p-values for weak exogeneity.

data-based technique for determining the order of causality. The algorithm takes as input argument the VECM(5) residual variance-covariance matrix from equation (6) and calculates the residual correlation structure. It is divided into two steps, the elimination and orientation stages. It begins with a graph in which every variable is linked to every other variable by an undirected edge. This graph represents $\boldsymbol{B}$, the matrix of contemporaneous effects among variables with parameters $b_{ii} = 1$ and b_{ij} the contemporaneous effect of variable j on i. At the elimination stage, non-contemporaneously related variables are de-linked. An edge is deleted if (i) the unconditional correlation between those two variables is insignificant, or (ii) the correlation between the two variables, conditional on every other variable, is insignificant.

The algorithm is implemented in the computer package Tetrad IV which uses Fisher's *z-statistic* to test for correlation significance and therefore, a contemporaneous relationship between variables.[4] Dependent variables remain linked by an edge. At the orientation stage, the direction of each remaining edge is determined and this represents causality between the two variables. The ordering represented by the final directed graph is then mapped into the instantaneous effects matrix, $\boldsymbol{B^r}$, which is used to specify the structural model (SVECM) in 7. Zero entries in $\boldsymbol{B^r}$ can be interpreted as members of the set of restrictions necessary for system identification while non-zero entries are contemporaneous effects to be estimated together with other short-run parameters. The algorithm identifies six contemporaneous relations at the 5% significance level. They include: $(\pi^e \to \pi)_t$, $(r \to \pi^e)_t$, $(r \to i)_t$, $(b \to r)_t$, $(y \to r)_t$, and $(y \to i)_t$. All six are represented in table 10 and are mapped into contemporaneous effects matrix $\boldsymbol{B^r}$ below.

These contemporaneous relationships are fairly robust. On changing the significance level to 10% for whether a pair of variables are correlated, the structure remains unchanged except for an undirected edge between inflation and inflation expectations. Some relationships are exactly what one would expect. For example, the causal direction from a real variable like capacity utilisation to the federal funds rate and the Treasury yield is theoretically consistent. The causal effect from long-term bond yields to inflation

[4]The algorithm is specified to avoid cycles in the causal structure and the significance of a contemporaneous relationship is at the 5% level.

Table 9 *The long-run matrix,* $\mathbf{\Pi}$

	π^e_{t-1}	π_{t-1}	y_{t-1}	i_{t-1}	r_{t-1}	b_{t-1}
$\Delta\pi^e_t$	-0.077	0.128	0.060^{**}	-0.038	-0.160	0.147
	(0.107)	(0.084)	(0.025)	(0.067)	(0.200)	(0.187)
$\Delta\pi_t$	0.477^{**}	-0.316^{**}	0.061^{**}	0.023	-0.711^{**}	0.527^{**}
	(0.139)	(0.110)	(0.033)	(0.087)	(0.261)	(0.244)
Δy_t	0.150	-0.070	-0.011	-0.225^{**}	0.155	-0.009
	(0.129)	(0.102)	(0.030)	(0.080)	(0.241)	(0.226)
Δi_t	0.144	-0.068	0.067^{**}	-0.174^{**}	-0.358^{**}	0.455^{**}
	(0.097)	(0.077)	(0.023)	(0.061)	(0.182)	(0.170)
Δr_t	0.089	-0.028	0.014	0.031	-0.293^{**}	0.202^{**}
	(0.051)	(0.040)	(0.012)	(0.032)	(0.096)	(0.090)
Δb_t	0.077	-0.030	-0.001	0.110^{**}	-0.094	-0.063
	(0.044)	(0.035)	(0.010)	(0.029)	(0.083)	(0.078)

** are significant effects. Standard errors are in brackets.

Table 10 *System contemporaneous causal structure as found by Tetrad*

$$
\begin{tikzcd}
\boldsymbol{i}_t & \boldsymbol{y}_t \arrow[l] \arrow[d] & & \\
\boldsymbol{b}_t \arrow[r] & \boldsymbol{r}_t \arrow[ul] \arrow[r] & \boldsymbol{\pi}^e_t \arrow[r] & \boldsymbol{\pi}_t
\end{tikzcd}
$$

expectations also makes sense given that long-maturity bonds proxy future inflation quite well. However, the causal direction from corporate to Treasury bond yields is surprising. The same direction of causal effect between government and corporate bond yields was observed for the equilibrium adjustment in table 9. There, as here, corporate bond yields appear to affect government bond yields. Perhaps corporate bonds are more sensitive to economic shocks and therefore adjust faster and transmit these shocks to the Treasuries. We now use this instantaneous structure to estimate a structural model.

Table 11 *Contemporaneous feedback matrix*

$$
\boldsymbol{B}^r \boldsymbol{x}_t = \begin{pmatrix} 1 & . & . & . & . & . \\ 0 & 1 & . & . & . & . \\ b_{ry} & b_{rb} & 1 & . & . & . \\ b_{iy} & 0 & b_{ir} & 1 & . & . \\ 0 & 0 & b_{\pi^e r} & 0 & 1 & . \\ 0 & 0 & 0 & 0 & b_{\pi\pi^e} & 1 \end{pmatrix} \begin{pmatrix} y_t \\ b_t \\ r_t \\ i_t \\ \pi^e_t \\ \pi_t \end{pmatrix}
$$

Note: . indicates Choleski-type upper-triangular zero constraints;
0 indicates over-identifying zero constraints.

3.4 The structural VECM

Pre-multiplying (6) by contemporaneous causal matrix $\boldsymbol{B}^r$ gives structural system (7). Zero entries in $\boldsymbol{B}^r$ represent insignificant instantaneous effects while non-zero b_{ij}s are parameters to be estimated.

$$\boldsymbol{B}^r\Delta\boldsymbol{x}_t = \tilde{\boldsymbol{\alpha}}(\boldsymbol{\mu}+\boldsymbol{\gamma}t+\boldsymbol{\beta}^{r\prime}\boldsymbol{x}_{t-1}) + \sum_{j=1}^{p-1}\tilde{\boldsymbol{\Gamma}}_j\Delta\boldsymbol{x}_{t-j}+\tilde{\boldsymbol{\varepsilon}}_t, \quad \tilde{\boldsymbol{\varepsilon}}_t \sim \mathsf{NID}(\boldsymbol{0},\boldsymbol{B}^r\boldsymbol{\Sigma}\boldsymbol{B}^{r\prime}), \tag{7}$$

with $\boldsymbol{\Omega} = \boldsymbol{B}^r\boldsymbol{\Sigma}\boldsymbol{B}^{r\prime}$ being a $K\times K$ diagonal matrix. This 6-dimensional SVECM is still over-parameterised with several insignificant regressors that eat away at the degrees of freedom without adding much to the model. We thus fine-tune (7) further by systematically dropping insignificant terms until we are left with a parsimonious model. Model reduction of the SVECM is carried out using the *Gets* procedure, a General-to-Specific model reduction procedure implemented here in the *PcGets* software package. Each equation is reduced by eliminating insignificant terms while ensuring that what remains is a congruent representation of the data generating process. Where necessary, impulse dummies are introduced to control for outlier effects. See Krolzig (2003) for further details on the model reduction process.

We now discuss the final parsimonious structural vector equilibrium-correction model or PSVECM for short. Figure 4 below and model diagnostics in Table 12 indicate that the estimated system fits the data reasonably well and is sufficiently specified for this type of model to proceed with further analysis.

Table 12 *PSVECM misspecification tests*

Test		$\Delta\pi_t^e$	$\Delta\pi_t$	Δy_t	Δi_t	Δr_t	Δb_t
AR 1-5:	$F_{(5,j)}$	2.16 (0.0712)	1.01 (0.4117)	0.46 (0.8041)	2.20 (0.0559)	2.07 (0.0712)	0.81 (0.5431)
Normality:	$\chi^2(2)$	20.07** (0.0000)	0.97 (0.6161)	0.39 (0.8233)	12.06** (0.0024)	10.81** (0.0045)	10.37** (0.0056)
ARCH 1-4:	$F_{(4,n)}$	2.88** (0.0029)	0.73 (0.5734)	1.90 (0.1129)	2.76* (0.0290)	0.49 (0.7402)	1.15 (0.3321)
Hetero:	$F_{(k,m)}$	20.07** (0.0000)	1.44 (0.1041)	2.17** (0.0018)	3.16** (0.0000)	1.66* (0.0282)	1.40 (0.0892)

**reject H_0 at the 5% level. p-values are in () brackets.

The equations of the PSVECM are discussed below:

Inflation expectations

$$\begin{aligned}
\widehat{\Delta\pi_t^e} = {} & \underset{(0.052)}{0.161}\ (\pi-\pi^e)^*_{t-1} - \underset{(0.066)}{0.165}\ \Delta\pi^e_{t-1} - \underset{(0.058)}{0.112}\ \Delta\pi_{t-1} - \underset{(0.052)}{0.164}\ \Delta\pi_{t-2} - \underset{(0.050)}{0.163}\ \Delta(\pi-\pi^e)_{t-3} \\
& - \underset{(0.050)}{0.257}\ \Delta(\pi-\pi^e)_{t-4} - \underset{(0.045)}{0.117}\ \Delta\pi_{t-5} + \underset{(0.038)}{0.136}\ \Delta y_{t-1} + \underset{(0.034)}{0.121}\ \Delta y_{t-5} + \underset{(0.057)}{0.185}\ \Delta i_{t-2} \\
& + \underset{(0.117)}{0.559}\ \Delta r_t - \underset{(0.728)}{3.488}\ \mathrm{I1975Q1}_t + \underset{(0.506)}{1.727}\ \Delta^4\mathrm{I1975Q3}_t + \underset{(0.607)}{1.915}\ \mathrm{I2008Q2}_t \\
& \hat{\sigma} = 0.645,\ R^2 = 0.523
\end{aligned} \tag{8}$$

Long-term bond yields proxy future inflation so the fact that a shock to the government bond is transmitted to inflation expectations within the same quarter is not surprising. This effect is both statistically

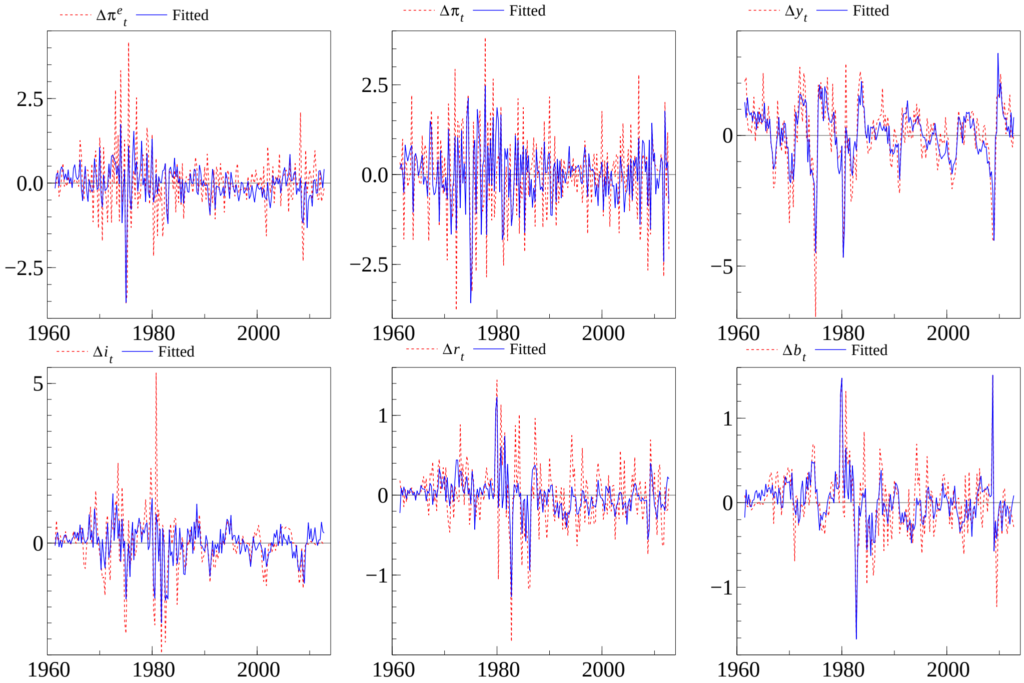

Figure 4 *PSVECM actual and fitted data plots*

and economically significant at about 60 basis points for a 100 basis point increase in the Treasury yield. The change in inflation expectations is also reacting to its own past as well as to past changes in inflation. Inflation expectations are in fact error-correcting through this persistence stretching back for five lags. That present inflation expectations are influenced by past expectations, and to the extent that expected inflation feeds into actual inflation, shows that it is beneficial for the Federal Reserve Board to operate in a world of anchored inflation expectations. Non-inflation-related short-run effects are dominated by capacity utilisation and monetary policy effects. A direct corporate bond yield effect is absent except via its near one-for-one contemporaneous effect on the government bond. We discuss this effect shortly. The long-run effect is also characterised by error-correction via the inflation - expected inflation relation, the sole long-run relation affecting expected inflation. Adjustment towards equilibrium is at 16% per quarter. The dummies seem to capture effects of the 1973-75 recession and oil price shocks in the 1970s and 2008. Average deviation between estimated and actual inflation expectations stands at 0.65 percentage points, with 52% of the variation in expected inflation explained.

Inflation

$$
\begin{aligned}
\widehat{\Delta\pi_t} \quad = \quad & -\underset{(0.042)}{0.102}\ (r-i)^*_{t-1} - \underset{(0.064)}{0.336}\ (\pi-\pi^e)^*_{t-1} - \underset{(0.024)}{0.071}\ (i-\pi^e)^*_{t-1} + \underset{(0.069)}{0.527}\ \Delta\pi^e_t + \underset{(0.070)}{0.107}\ \Delta\pi^e_{t-5} \\
& -\underset{(0.061)}{0.367}\ \Delta(\pi-\pi^e)_{t-1} - \underset{(0.060)}{0.313}\ \Delta\pi_{t-2} - \underset{(0.050)}{0.253}\ \Delta(\pi-\pi^e)_{t-3} - \underset{(0.202)}{0.678}\ \Delta(b-r)_{t-2} \\
& +\underset{(0.833)}{3.392}\ \mathrm{I1974Q3}_t + \underset{(0.836)}{2.702}\ \mathrm{I1980Q4}_t - \underset{(0.811)}{2.371}\ \mathrm{I2011Q4}_t \\
& \hat{\sigma} = 0.799,\ R^2 = 0.585
\end{aligned} \tag{9}
$$

Inflation also adjusts to maintain equilibrium in the inflation - expected inflation long-run relation but at twice the speed (33.6% quarterly) as inflation expectations. That inflation adjusts towards $(\pi - \pi^e)$ more than twice as fast as inflation expectations and the fact that $(\pi - \pi^e)$ is the only cointegration relation affecting inflation expectations according to (8) could explain the weak exogeneity result of inflation expectations in table 8.[5] In addition, inflation reacts to disequilibrium in the liquidity premium and the Fisher relations although considerably slowly at quarterly speeds of 10% and 7%, respectively. Furthermore, signs for the direction of these effects are also theory-consistent. A higher real interest rate depresses investment, with the lower output and demand that follow subsequently leading to lower inflation. When $\pi > \pi^e$ ($\pi < \pi^e$), inflation adjusts downwards (upwards) as expected. The liquidity premium and inflation are moving in opposite directions such that an increase in the former (relative fall in the federal funds rate, for example) shows the FOMC stimulating economic activity at the time when inflation is facing downward pressure. If it is r rising relative to the federal funds rate then future inflation is expected to be higher.

Both short-run and long-run effects are dominated by the two inflation measures and the two bond yields. The 0.527 contemporaneous effect of inflation expectations on inflation is very significant and is further support for the hypothesis that inflation expectations lead inflation. A rise in the bond yield-defined risk premium often signals a slowdown and has a negative effect on prices. Effect signs and dates on the dummies point to oil price shocks in 1973-74, 1979-80 triggered by the Iranian revolution, and the aftermath of the 2008Q1-2009Q2 Great Recession as likely explanations. Compared to inflation expectations, the equation for inflation has a better goodness of fit (59%) but higher deviation (0.799 percentage points) between actual and estimated values.

Capacity utilisation

$$
\begin{aligned}
\widehat{\Delta y_t} = {} & \underset{(0.049)}{0.175}\ (r-i)^*_{t-1} - \underset{(0.026)}{0.060}\ (i-\pi^e)^*_{t-1} - \underset{(0.089)}{0.432}\ \Delta\pi^e_{t-2} - \underset{(0.093)}{0.368}\ \Delta\pi^e_{t-3} - \underset{(0.082)}{0.351}\ \Delta\pi^e_{t-4} \\
& + \underset{(0.065)}{0.341}\ \Delta y_{t-1} - \underset{(0.067)}{0.126}\ \Delta y_{t-2} + \underset{(0.058)}{0.148}\ \Delta y_{t-3} + \underset{(0.065)}{0.238}\ \Delta(i-\pi^e)_{t-1} + \underset{(0.246)}{0.861}\ \Delta r_{t-1} \\
& - \underset{(0.257)}{1.384}\ \Delta b_{t-1} - \underset{(0.244)}{0.739}\ \Delta(b-r)_{t-3} - \underset{(0.607)}{2.289}\ \Delta\mathrm{I1980Q2}_t \\
& \hat{\sigma} = 0.840,\ R^2 = 0.631
\end{aligned}
\tag{10}
$$

Panel II of figure **??** in the appendix shows the liquidity premium leading capacity utilisation rather convincingly at least until the early 1990s. The dynamics of (10) are such that the term structure effect on capacity utilisation at 0.175% is about three times that of the real interest rate at 0.06%. Output is therefore affected more by the $(r-i)$ term spread than by the real (short-term) interest rate. From an investment perspective, it is certainly more sensible for a firm to decide the level of capacity to use based on expected economic conditions according to the term structure than on the present real rate. Rising long interest rates relative to short-term rates may signal higher future inflation, prompting higher capacity use in anticipation of higher demand. On the other hand, higher credit costs from rising real rates depress investment, reducing output with the lower demand leading to falling capacity use. These two relation effects show that monetary policy influences output. Short-term lags are dominated by inflation expectations, the Treasury yield and corporate bond yield effects, further highlighting the forward-looking nature of capacity utilisation dynamics. The dummy is probably due to the recession

[5]The loadings matrix in table 8 also showed inflation (unlike inflation expectations) reacting significantly and strongly to the $(\pi - \pi^e)$ relation.

1980Q2-1980Q3. Equation (10) fits the data reasonably well for this type of model with 63% of the variation in capacity utilisation explained.

Monetary policy (the FFR)

$$\begin{aligned}
\widehat{\Delta i_t} \quad = \quad & \underset{(0.033)}{0.202}\ (r-i)^*_{t-1} + \underset{(0.095)}{0.305}\ (b-r)^*_{t-1} + \underset{(0.013)}{0.062}\ \mathrm{y}^*_{t-1} + \underset{(0.047)}{0.112}\ \Delta\pi^e_{t-2} + \underset{(0.044)}{0.116}\ \Delta\pi^e_{t-3} \\
& + \underset{(0.031)}{0.089}\ \Delta\pi_{t-2} + \underset{(0.037)}{0.216}\ \Delta y_t - \underset{(0.029)}{0.087}\ \Delta y_{t-4} + \underset{(0.057)}{0.331}\ \Delta i_{t-1} + \underset{(0.052)}{0.234}\ \Delta i_{t-3} \\
& + \underset{(0.052)}{0.253}\ \Delta i_{t-5} + \underset{(0.101)}{0.549}\ \Delta r_t + \underset{(0.108)}{0.239}\ \Delta r_{t-1} - \underset{(0.098)}{0.368}\ \Delta r_{t-2} + \underset{(0.103)}{0.167}\ \Delta r_{t-5} \qquad (11) \\
& + \underset{(0.356)}{1.695}\ \Delta\mathrm{I1973Q3}_t + \underset{(0.553)}{4.193}\ \mathrm{I1980Q4}_t \\
& \hat{\sigma} = 0.473,\ R^2 = 0.723
\end{aligned}$$

Monetary policy interacts with the corporate bond yield via the risk premium long-run relation according to (11). That the risk premium has a positive effect on the federal funds rate is initially surprising. After all, the risk premium often increases around economic slowdowns against which the FOMC responds with a rate cut. To explain the +0.305 coefficient, imagine that an increase in expected future inflation triggers a widening of the risk premium as investors demand higher yields on corporate bonds.[6] This expectation of higher inflation in turn elicits an increase in the federal funds rate from the FOMC to control inflation. Two more cointegration relations have significant effects on US monetary policy: the liquidity premium, via which the federal funds rate corrects disequilibrium at 20.2% per quarter, and capacity utilisation. Thinking of relations one and two in terms of what they imply for inflation, then their effect size relative to that of capacity utilisation is illustrative of a Taylor rule effect where monetary policy responds stronger to inflation than to output. The effect of capacity utilisation on the federal funds rate is small compared to the first two but its presence together with a contemporaneous effect is evidence of the importance of output in US monetary policy implementation.

Short-term dynamics are dominated by own-lag effects with considerable persistence in which near-past changes in the federal funds rate have a positive effect on the present change. Inflation expectations either directly or as proxied by the Treasury also contribute significantly to short-term effects thereby highlighting a major concern of monetary policy. Past inflation barely has any effect. The effect of the Treasury bond is much larger than that of inflation expectations. Perhaps the Treasury bond is signalling market sentiment on future inflation better than the 12-month-ahead survey-based estimate. The apparent absence of a direct corporate bond effect among short-run terms is noted but explained by the very substantial contemporaneous effects from corporate bonds to Treasuries to the federal funds rate. That is, a shock to the corporate bond is transported to the federal funds rate within a quarter. With a 72% goodness of fit, our equation for monetary policy explains a lot of the variation in the monetary policy instrument.

[6] Impulse response analysis in section 4 will show the two bonds responding asymmetrically to a monetary policy shock with the corporate bond adjusting faster and further than the government bond.

Government bond yield

$$
\begin{aligned}
\widehat{\Delta r_t} = {} & \underset{(0.012)}{0.030}\ (r-i)^*_{t-1} + \underset{(0.035)}{0.205}\ (b-r)^*_{t-1} + \underset{(0.005)}{0.010}\ y^*_t - \underset{(0.016)}{0.025}\ \Delta\pi^e_{t-1} + \underset{(0.014)}{0.097}\ \Delta y_t - \underset{(0.014)}{0.023}\ \Delta y_{t-1} \\
& + \underset{(0.019)}{0.048}\ \Delta i_{t-2} + \underset{(0.017)}{0.054}\ \Delta i_{t-3} + \underset{(0.042)}{0.958}\ \Delta b_t - \underset{(0.040)}{0.099}\ \Delta b_{t-1} + \underset{(0.051)}{0.239}\ \Delta(b-r)_{t-3} \\
& - \underset{(0.037)}{0.072}\ \Delta b_{t-4} - \underset{(0.188)}{0.683}\ \text{I1986Q2}_t - \underset{(0.207)}{1.979}\ \text{I2008Q4}_t \\
& \hat{\sigma} = 0.181,\ R^2 = 0.806
\end{aligned} \tag{12}
$$

We again see the ubiquitous significance of the term structure of interest rates relation among the long-run forces driving the system, albeit with a considerably smaller effect. In fact, all variables except inflation expectations are reacting to the term structure relation. When the federal funds rate rises to control inflation, for example, the yield on the 30-year maturity government bond rises to maintain compensation in the form of a term premium. Of course, there are exceptions[7] to this dynamic - usually before a recession - when the term premium actually turns negative. The positive parameter, +0.03, implies the Treasury yield does not error-correct via the term structure. However, this does not contradict the error-correction in the Treasury yield implied by long-run matrix $\mathbf{\Pi}$ in table 9. One only has to look at the much larger error-correction via the risk premium relation. At 20.5% per quarter, the Treasury yield adjusts to restore cointegration with the corporate bond much faster than towards the federal funds rate. In addition to the much smaller effect of 10 basis points from capacity utilisation, we observe the same three long-run forces driving the government bond as do the federal funds rate: the term premium, the risk premium, and output.

Short-run effects are populated by the same variables as the long-run, with the two inflation measures barely having an impact. The corporate bond is very influential, including a near one-for-one contemporaneous effect. The contemporaneous effect from the corporate bond to the government bond is unexpected but this is what the data is telling us. The expectation that the government bond influences the corporate bond is seen only with a lag. The first dummy is not easily explainable but the 2008Q4 dummy must surely be due to the 2007-09 global financial crisis when extensive market uncertainty and investor flight to safe investments drove yields on Treasuries lower. The estimated equation fits the data quite well with 81% of the variation in the Treasury explained.

Corporate bond yield

$$
\begin{aligned}
\widehat{\Delta b_t} = {} & - \underset{(0.018)}{0.120}\ (r-i)^*_{t-1} - \underset{(0.008)}{0.037}\ (i-\pi^e)^*_{t-1} - \underset{(0.024)}{0.044}\ \Delta\pi^e_{t-1} + \underset{(0.016)}{0.021}\ \Delta\pi_{t-2} + \underset{(0.019)}{0.070}\ \Delta y_{t-2} \\
& - \underset{(0.029)}{0.062}\ \Delta i_{t-1} - \underset{(0.030)}{0.079}\ \Delta i_{t-2} + \underset{(0.026)}{0.115}\ \Delta i_{t-5} + \underset{(0.076)}{0.460}\ \Delta r_{t-1} + \underset{(0.052)}{0.127}\ \Delta r_{t-3} \\
& - \underset{(0.054)}{0.227}\ \Delta r_{t-5} - \underset{(0.078)}{0.155}\ \Delta b_{t-1} + \underset{(0.077)}{0.104}\ \Delta(b-r)_{t-2} + \underset{(0.189)}{0.782}\ \Delta\text{I1980Q1}_t \\
& - \underset{(0.276)}{1.130}\ \text{I1982Q4}_t + \underset{(0.261)}{1.588}\ \text{I2008Q4}_t \\
& \hat{\sigma} = 0.257,\ R^2 = 0.584
\end{aligned} \tag{13}
$$

[7] See panel I of figure **??** in the appendix.

Equation (13) provides further evidence of interaction between corporate bonds and monetary policy, which acts through the term structure and the real short-term interest rate relations. Loosening monetary policy in the form of lowering the federal funds rate triggers a downward correction in the government bond yield. Improving credit market conditions eventually reach the corporate bond market, signalling impending expansion in economic activity and thus causing the yield to fall by 12 basis points after the initial 1% point hike in the term spread. The real interest rate provides a second but weaker long-run route through which monetary policy affects corporate bonds. There are several significant short-run effects dominated by the government bond. We saw in equation (12) that the corporate bond exerts significant influence on the government bond, including contemporaneously. Now we see short-term dynamics of the Treasury bond affecting corporate bonds. There is, one could say, a strong dynamic interaction between the two bond yields. Interestingly, the sign on the 2008Q4 dummy is opposite that on the same dummy in the Treasury equation thus underscoring the divergence of these bond yields during the 2008Q1-2009Q2 recession.[8] This particular dummy is probably capturing turmoil in financial markets such as Lehman Brothers Bank filing for bankruptcy on 15 September 2008.

Overall, the presence of the term structure of interest rates relation in all but equation (8) points to its potency in influencing economic conditions. As part of the risk premium, the corporate bond yield affects the federal funds rate, and therefore monetary policy. It also reacts to monetary policy via the term structure and real interest rates relations, as well as through short-run adjustments in the federal funds rate. These direct and indirect effects show clear interaction between monetary policy and corporate bonds. Furthermore, corporate bond yields transmit monetary policy shocks to inflation and output among other variables. This is true because of their sensitivity to the policy instrument as well as a close link to government bonds, making corporate bonds a viable conduit of policy.

The focus now switches to the impact of monetary policy shocks on other variables and the role that the corporate bond plays in transmitting these shocks to inflation and output.

4 Effects of a monetary policy shock

Impulse responses are estimated for a 100 basis points unanticipated increase to the federal funds rate. The impulse responses derived from the PSVECM are plotted in figures 5 and 6. The former shows cointegration relation responses while figure 6 shows the responses of the variables. Also included in each plot is the 95% confidence interval constructed via bootstrapping with 2000 repeated random sampling with replacement of residuals from equations (8)-(13), following Hall (1992). When computing these impulse responses, cointegrating relations are assumed to be super-consistent and are therefore held constant. We further include a set of impulse responses in figure 7 based on the SVECM specification to demonstrate the benefits of the *Gets* model reduction process. This results in a parsimonious but still congruent model when insignificant terms are dropped. The difference between the non-*Gets*-reduced SVCEM and the *Gets*-reduced PSVECM is obvious from the significantly wider SVECM confidence intervals compared to those from the PSVECM which are more concise.

[8] This is actually quite pronounced in the middle graph of figure 1. Higher risk in the corporate market, coupled with investor 'flight-to-safety' into US Treasuries should drive the spread wider. The asymmetric implication of the effects of the 2008Q4 dummy (-1.979 and 1.588) for the two bond markets is also telling: investors accepted a lower Treasury yield and demanded a higher corporate bond yield.

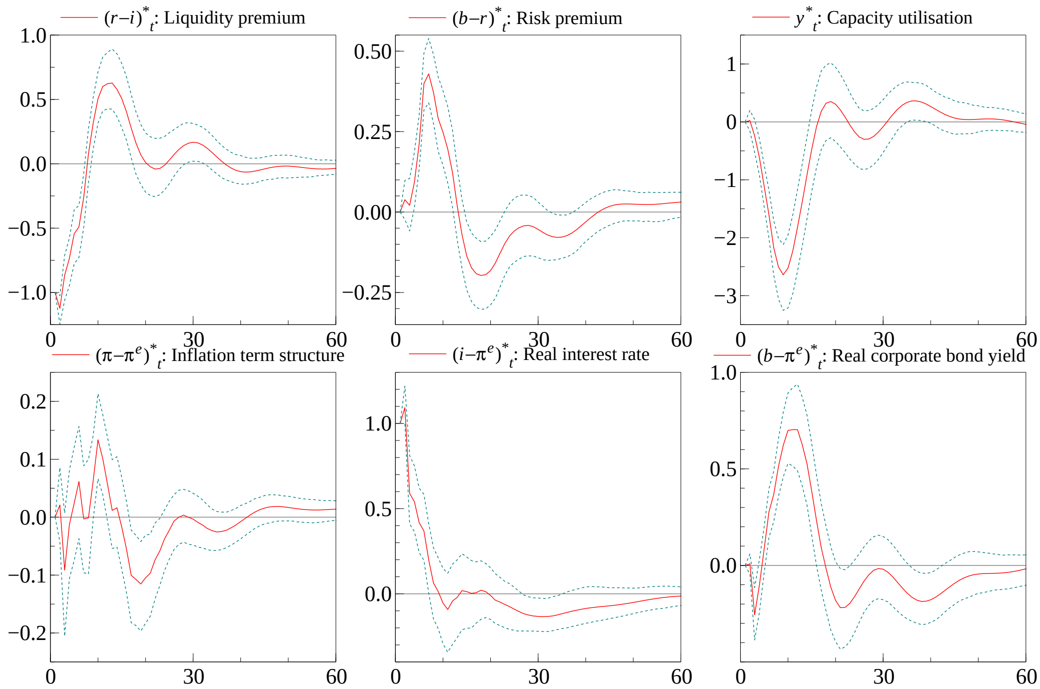

Figure 5 *PSVECM impulse response functions for long-run relations*

1. *Cointegration relations response:* $(r-i)^*_t$, $(b-r)^*_t$, y^*_t, $(\pi-\pi^e)^*_t$, $(i-\pi^e)^*_t$

   As expected, a shock to monetary policy only has a temporary effect on cointegration relations. According to figure 5, the liquidity premium falls by 102 basis points instantaneously due to delayed reaction of the Treasury yield to the shock. Ensuing quarters see adjustment in both the federal funds rate and the Treasury yield as the $(r-i)^*_t$ relation returns to equilibrium within 16 quarters. The rise in the risk premium peaks at 42 basis points with a return to its long-run path after 42 quarters. The $(b-r)^*_t$ reaction is because the corporate bond yield reacts faster, and by more than the Treasury bond, thereby enforcing tighter monetary conditions in the economy. In fact, according to figure 7 the initial reaction of the Treasury bond yield is downwards.

   According to equation (10), the effect of a federal funds rate shock on capacity utilisation is both direct - via the liquidity premium, short-term real interest rate, and other short-term dynamics - and indirect at various lags. As a result of the contractionary policy shock, demand falls so firms reduce capacity under utilisation with a recovery within 16 quarters from the moment of the shock. The increase in $(\pi-\pi^e)^*_t$ can be explained in the context of the interaction between inflation and inflation expectations. Actual prices are assumed sticky, so contractionary policy lowers inflation expectations - which react faster in the short-term - thereby increasing $(\pi-\pi^e)_t$. The $(\pi-\pi^e)_t$ differential is restored to its long-run path in 24 quarters. The real federal funds rate, like the liquidity premium, rises 100 basis points instantaneously by construction with the federal funds rate adjustment to return to equilibrium within 9 quarters. Opposite responses by inflation expectations and the corporate bond means that the real corporate bond yield shoots up

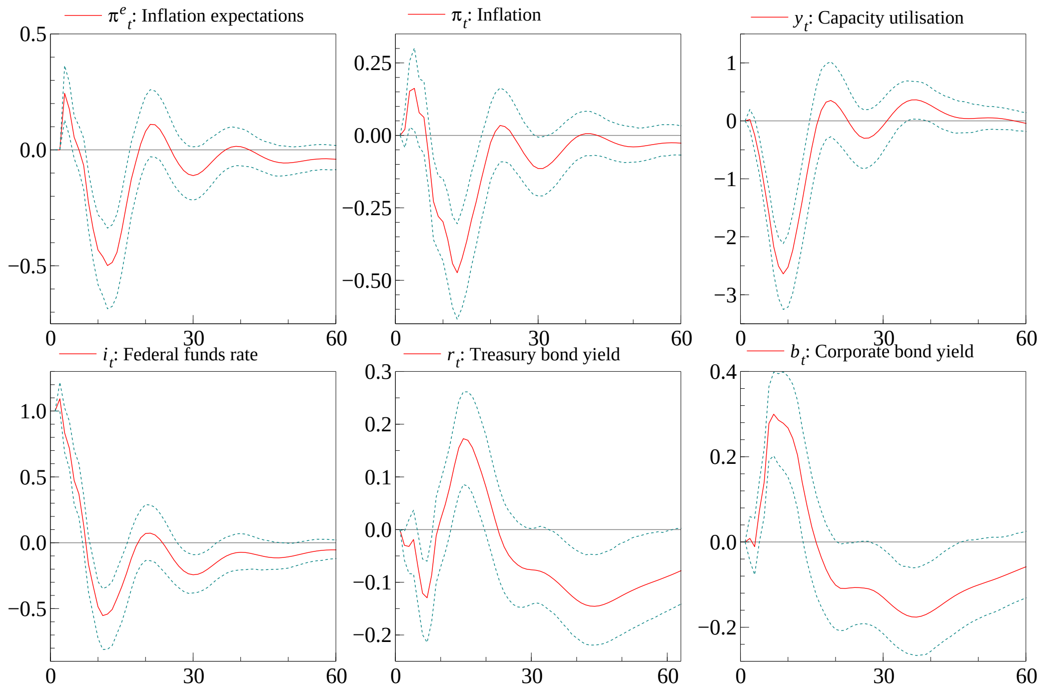

Figure 6 *PSVECM impulse response functions for levels*

by 70 basis points within 10 quarters, before the shock effect dissipates after another 5 quarters.

2. *Level variable responses.*

   According to figure 6, level variable responses to a policy shock are consistent with theory if one ignores the mild price puzzle in the responses of inflation measures.[9] The shock triggers a 50 basis point fall in inflation expectations within 12 quarters. Inflation falls by 47 basis points within 13 quarters. These effects on both inflation measures disappear after 18.5 and 20.5 quarters, respectively. Clearly, inflation expectations respond faster and deeper to the shock than does inflation. The Federal Reserve Board can thus lean against inflation through tighter policy. The Treasury bond yield reacts by first falling before peaking at 17 basis points above its initial position within 15 quarters. The corporate bond on the other hand rises more and faster by 30 basis points.

That the federal funds rate shock elicits these reactions is evidence that monetary policy can and does influence economic conditions. The reaction of corporate bond yield is stronger and faster than that from the government bond yield, suggesting corporate bonds as a viable vehicle through which monetary policy can influence the economy. Indeed, the Federal Reserve Board did just that when it intervened in the corporate bond market as part of the QE program to support credit markets due to the effects of the last financial crisis and Great Recession.

[9]For both inflation equations, (8) and (9), we see that the federal funds rate has a (net) positive effect on inflation and its expectations. This is the source of the price puzzle.

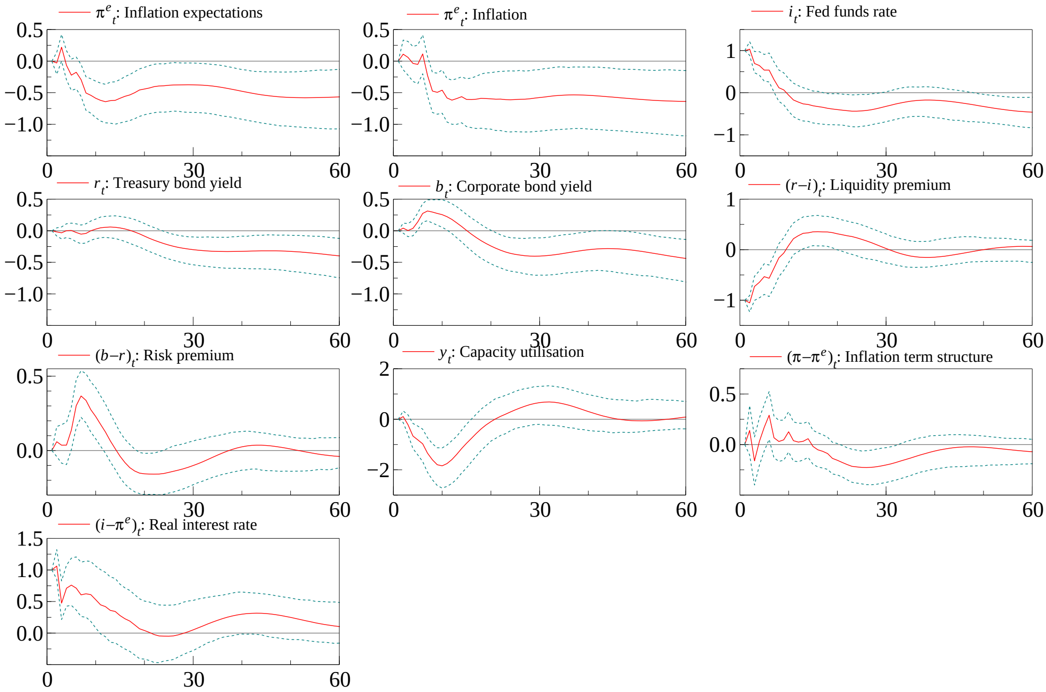

Figure 7 *SVECM Impulse response functions*

5 Conclusion

The contribution that corporate bond yields make to the transmission mechanism of monetary policy is investigated within a 6-dimensional cointegrated VAR framework. Inflation expectations lead inflation and react faster to shocks possibly due to price stickiness. Capacity utilisation shows economically significant sensitivity to the term structure of interest rates, evidence of the effect of the path of inflation on output. The federal funds rate is also significantly responsive to the term structure as well as to the risk premium. The corporate bond yield interacts with monetary policy both in the long-run and the short-run at several lags. The 30-year corporate bond adjusts towards the term structure of interest rates relationship. The faster and higher rise of the corporate bond yield relative to the Treasury in response to a rise in the federal funds rate ensures the efficacy of monetary policy in affecting the risk premium in the economy. In the opposite direction, policy is also reaction to movements in corporate bond yields via the corporate-Treasury bond yield risk premium cointegraton relation. Through significant effects on other macroeconomic variables, such as inflation expectations, inflation, and output, corporate bonds also transmit monetary policy to the rest of the US economy.

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Appendix

Measuring the 30-year bond yield series

Data is available for all series from $1960M1$ except for the government bond yield (GS30) series which starts in $1977M2$. GS30 data is also missing between 2002M3 and 2006M1, during which no new 30-year maturity Treasuries were issued. We exploit the co-movement in GS30_t and GS20_t to reconstruct a government bond series for the full sample periods by splicing the missing GS30 observations. Monthly data is used for splicing so as to capture at as close a frequency as possible data details at the start and end dates for which data is missing than would otherwise be possible were quarterly data to be used. Changes in the two yields should be close at lower frequencies.

The long-term behaviour of the $GS20_t$ and $GS30_t$ series lends credence to the assumption that these two series are linked. Their graphs below confirm this close relationship. Looking at the ranges $1977M2 - 1986M12$, and $1993M10 - 2002M2$ in Figure 8 for which both $GS20_t$ and $GS30_t$ data are available, it is evident that these two series closely track each other. The cross plot of their first differences in the upper right panel of Figure 8 is further support for this argument. The spliced $GS30_t$ series is then transformed to quarterly frequency by taking simple averages.

Based on these properties of the GS20 and GS30 time series, we propose the following target function for splicing the missing values from (i) $t = 1977M1$ to $1959M1$ and (ii) $t = 2002M2$ to $2002M1$:

$$\begin{aligned} \min_{\{GS30_t^*\}} \quad & \frac{1}{2}\sum_{t=\tau_0}^{\tau_1} \left(\Delta GS30_t^* - \Delta GS20_t\right)^2 \\ \text{s.t.} \quad (i) \quad & GS30_{\tau_1}^* = GS30_{\tau_1}, \\ (ii) \quad & GS30_{\tau_0-1}^* = GS30_{\tau_0-1} \text{ and } GS30_{\tau_1}^* = GS30_{\tau_1}, \end{aligned} \tag{14}$$

where τ_0 denotes the month with the first missing value and τ_1 the month with the first observation of $GS30$ after the gap.

Due to the location of the missing data, the splicing takes two forms.

(i) Backward iteration for $t = \tau_0 = 1977M1$ to $\tau_1 - 1 = 1959M1$:

$$\begin{aligned} GS30_t^* &= GS30_{t+1}^* - \Delta GS20_{t+1} \\ & \quad \text{with } GS30_{1977M2}^* = GS30_{1977M2}. \end{aligned} \tag{15}$$

For the $1959M1 - 1977M2$ period, the backward iteration (15) emerges, which approximates the month-on-month change in the 30-year Treasury yield by the month-on-month change in the 20-year Treasury yield.

(ii) Forward iteration for $t = \tau_0 = 2002M2$ to $\tau_1 - 1 = 2002M1$:

$$\begin{aligned} GS30_t^* &= GS30_{t-1}^* + \Delta GS20_t + \delta, \\ & \quad \text{with } \delta = \frac{1}{\tau_1 - \tau_0 + 2}\left(\Delta_{\tau_1-\tau_0+1} GS30_{\tau_1} - \Delta_{\tau_1-\tau_0+1} GS20_{\tau_1}\right) \\ & \quad \text{and } GS30_{\tau_0-1}^* = GS30_{\tau_0-1}. \end{aligned} \tag{16}$$

For the $2002M3 - 2006M1$ period, (16) equates the growth rate in the 30-year Treasury yield to

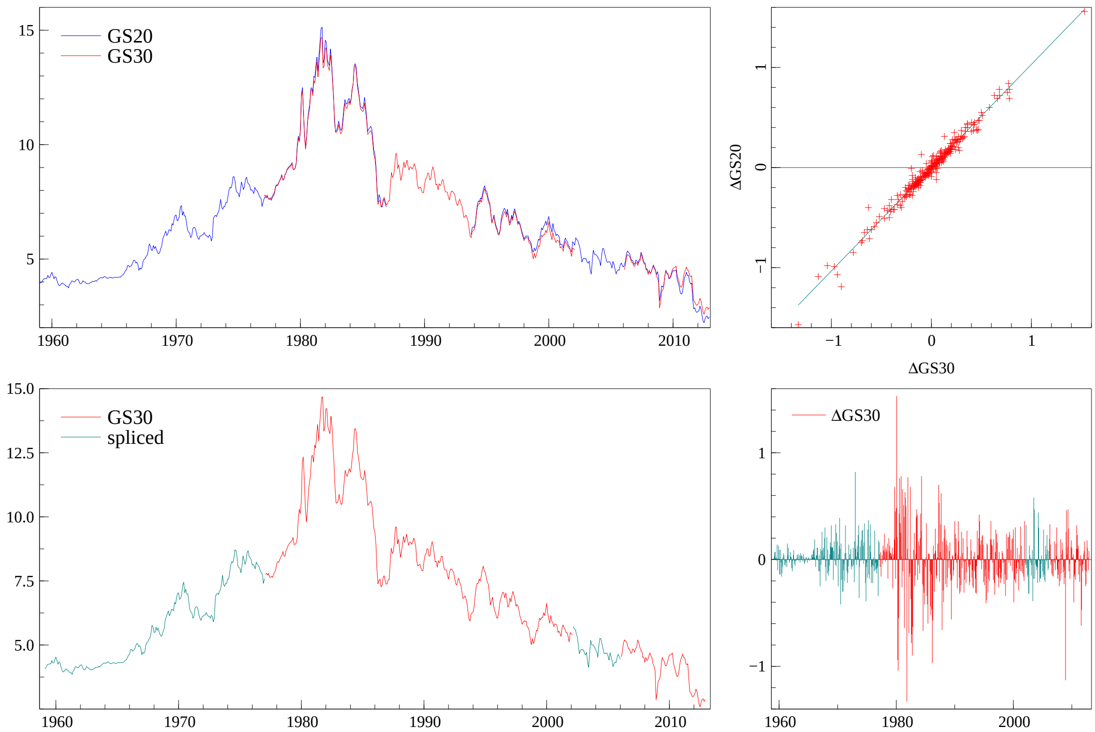

Figure 8 *Splicing of the GS30 series*

the sum of the growth rate of the 20-year Treasury yield and the difference in the two growth rates over the period data is missing, weighted by the number of periods for which data is missing.

Fiscal Policy, Interest Rates, and Output: Equilibrium-Correction Dynamics in the US Economy

Hans–Martin Krolzig & Isaac Sserwanja*

School of Economics, Keynes College, University of Kent, Canterbury

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Abstract

Joint modelling of fiscal and monetary policies should elucidate on their interaction. We construct an eight-dimensional parsimonious structural vector equilibrium correction model (PSVECM) of the US macro economy over the last five decades. The fiscal deficit is found to be one of the five cointegration vectors, constraining fiscal policy in the long-run. In contrast, the share of the government sector is found not to be mean reverting. To overcome the common problem of ad-hoc assumptions regarding the direction of instantaneous causality, we use graph-theoretical methods to identify the causal structure of the model from the data. Model reduction procedures allow us to control for the curse-of-dimensionality inhibiting such high-dimensional vector autoregressive systems. Impulse-response analysis of the parsimonious system facilitated the precise measurement of the dynamic Keynesian fiscal multiplier, where we distinguish between deficit-spending and balanced-budget government spending shocks (as in the so-called Haavelmo, 1945, theorem). Our estimates of the long-run multiplier are 1.62 in case of bond-financed spending and 1.77 for a tax-financed spending shock, with both being significant greater than one at a 95% confidence level. Monetary policy is neutral in the long-run except for the level of output where its effect is permanent. Increasing the federal funds rate by a percentage point is followed by falling tax revenues while government spending is largely unchanged, thus inflating the fiscal deficit in the short-run and medium-run.

Keywords: Cointegrated VAR; Gets modelling; graph theory; deficit; balanced budget multiplier; spending multiplier; fiscal policy; monetary policy.

JEL classification: C32; E63; E52; G12; H62.

1 Introduction

One outcome of the 2007-2009 global financial crisis and the the subsequent Great Recession is the re-emergence of fiscal policy as a powerful tool in the management of economic policy in major economies around the world. Falling output was followed by falling tax revenues, which when coupled with bailing out the financial system explains part of the worsening budget deficits in several countries. Given

*Corresponding author email: hm.krolzig@gmail.com. We are grateful to Alan Carruth and the participants at the MMF Annual Conference, Cardiff; the VfS Annual Conference, Münster; and the NBER-NSF Time Series Workshop, Vienna, for useful suggestions.

a global economy that is yet to fully recover, and is characterised by high debt levels, deficits, and unconventional monetary policies including near-zero interest rates, a natural question ought to be asked: how does the budget deficit interact with interest rates? This empirical study answers this question and more through the analysis of US data. Our contribution to the literature is twofold. First, we analyse the interaction between fiscal policy and monetary policy and how they have affected long-maturity bond yields over the last fifty years. Second, we measure fiscal multipliers over the same period, thereby characterising the extent to which fiscal policy affects output. A characteristic of our approach is the emphasis of theoretically-consistent long-run relationships which supports the structural interpretation we attach to our results. For example, we explicitly include the deficit as a cointegration relation to allow us to study the dynamic feedback between this important budget constraint and the rest of the economy. Our results have a Keynesian flavour. From a Keynesian perspective, increasing government spending - spending, henceforth - or cutting taxes can be used to stimulate aggregate demand. This may be especially necessary during recessions. According to Ilzetzki et al. (2013), fiscal expansion, if well-targeted to optimise the fiscal multiplier, can stimulate aggregate demand and subsequently take an economy out of a downturn faster than if no intervention were undertaken. In fact, empirical evidence in Feldstein (1982) and Eisner (1992), among others, links budget deficits and output. This is in opposition to the Ricardian equivalence hypothesis in Barro (1989) that budget deficits do not affect output. The deficit-output debate is one with real and potentially long-lasting economic consequences. Alesina and Ardagna (2010) look at fiscal stimuli and adjustments in Organisation of Economic Co-operation and Development (OECD) countries between 1970 and 2007. They find: (i) tax-cut-based fiscal stimuli more likely to increase growth compared to increasing spending, and (ii) fiscal adjustment in the form of a cut in spending and no tax increase is more likely to lower the deficit and the debt-to-GDP ratio than one based on a tax increase. Counter arguments posit that fiscal expansion - often entailing an increase in government debt levels - may be necessary to stop a recession from getting worse. It is important to study fiscal and monetary policies simultaneously not least because governments (to an extent) react to changes in their borrowing costs, as demonstrated in de Groot et al. (2014). That there is a fiscal reaction to an increase in the interest rate at which government borrows can improve budget discipline, de Groot et al. (2014) argues.[1]

1.1 Fiscal policy and output

One branch of the literature on fiscal policy focusses on its relationship with output. Studies on this often include, among other aspects, estimations of spending or tax multipliers. Blanchard and Perotti (2002) combine a structural VAR modelling and event-study approach to investigate the response of output and its components to tax and spending shocks in the postwar US. Large fiscal events such as the 1975 tax cut are used in the event-study section. Using institutional information and the elasticity of fiscal variables to economic activity to achieve identification, their findings are largely in line with standard Keynesian theory predictions. Positive spending shocks increase private consumption and output. Positive tax shocks have a negative effect on output. Increasing spending or taxes reduces investment spending by the private sector. Afonso and Souza (2012) look at 1970Q3-2007Q4 US data

[1]Markets do not however impose 'budgetary discipline', at least in a timely manner, as was evident when for several years the rates at which some peripheral Eurozone countries borrowed were markedly lower given their budget positions. de Groot et al. (2014) argue for fiscal rules to complement markets as an enforcer of fiscal sustainability.

and include the government inter-temporal budget constraint in modelling the macroeconomic effects of fiscal policy. They include the response of fiscal variables to government debt dynamics. The observed response is also Keynesian in nature. A positive spending shock has a positive but small effect on gross domestic product dynamics, private consumption, private investment, and asset markets. Stock prices react quickly to fiscal policy shocks but house prices react more persistently. When the shock is to revenue, a crowding-out effect is detected as a positive response in private investment is quickly eroded. It is claimed that taking debt dynamics into account makes GDP and long-term interest rates more responsive, in addition to more persistent fiscal policy effects. Rossi and Zubairy (2011) find fiscal shocks are more relevant in explaining output fluctuations in the medium term while monetary shocks are associated more with business cycle fluctuations. The approach in Rossi and Zubairy (2011) is interesting in that it simultaneously studies fiscal and monetary shocks on the macroeconomy. We too take this approach to jointly model fiscal and monetary policy, justified by the idea that the two policies can (indirectly) interact via their effects on common variables. This is certainly the case for the US where monetary policy has low unemployment as one of its goals while at the same time government-run fiscal policies do affect the level of unemployment.

Favero (2002) jointly models monetary and fiscal policy in Germany, France, Italy, and Spain, observing that not jointly modelling the two policies causes some estimated parameters to be insignificant and ignores important interactions between the two. He finds that (i) 1970s monetary policy by non-German authorities was not able to stabilise inflation. Inflation stabilisation was achieved in the 1980s and 1990s when domestic monetary policy was tied to German monetary policy, (ii) stable inflation has been achieved despite a lack of fiscal discipline, and (iii) the interaction between fiscal and monetary authorities depends on the response of spending and taxes to interest payments on public debt. Afonso and Souza (2011) focus on the link between fiscal policy and asset markets for the US, UK, Germany, and Italy. Fiscal shocks are identified following Blanchard and Perotti (2002) above. Like Favero and Givazzi (2008), the specification in Afonso and Souza (2011) includes a government debt variable to capture effect of debt dynamics feedback on fiscal policy. For the US, spending shocks have a positive and persistent effect on house prices but a negative effect on stock prices. Shocks to government revenues have a negative impact on house prices, and a small but positive impact on stock prices. In fact, fiscal policy shocks only have a minor effect on the behaviour of stock markets and house prices. Perotti (2004) studies fiscal policy effects in the US, West Germany, the UK, Canada, and Australia, finding spending and tax cut effects on GDP and its components weakened after 1980. While relatively large positive effects are observed for private consumption, there is no response from private investment. Fatas and Mihov (2001) find that consumption and employment increase in response to increases in spending.

One way to analyse fiscal policy effects is by measuring the Keynesian fiscal multiplier. The spending multiplier is the increase in output due to an increase in spending. Consensus among economists on the size of fiscal multipliers is elusive. Part of the disagreement can be attributed to how fiscal shocks are identified and whether the researcher takes a Keynesian or classical view. Ricco (2014) argues that ignoring policy anticipation effects and assuming perfect information results in wrong estimations of fiscal multipliers. Estimates of the spending multiplier lie between 0.5 and 2.5, according to a survey by Chinn (2013). On whether fiscal multipliers are dependent on prevailing economic conditions, Ramey and Zubairy (2014) do not find evidence that spending multipliers are different based on slack in the economic system. In fact, they argue that "spending multipliers were not necessarily higher than average during the Great Recession" when interest rates were near zero. Some recent studies on fiscal multipliers during recessions include but are not limited to Auerbach and Gorodnichenko (2013), Auerbach and Gorodnichenko (2012), Bachmann and Sims (2012). Fatas and Mihov (2001) find a spending multiplier of more than one, driven by an increase in private consumption. The response of investment to spending is insignificant. Auerbach and Gorodnichenko (2012) analyse fiscal multiplier variability over the business cycle. Using regime-switching models, fiscal multipliers are found to be larger during recessions than in expansions. Their analysis of disaggregated spending components finds military spending to have the largest multiplier. They also find that taking into account the predictability of fiscal policy increases multipliers in recessions. Perotti (2004) finds a US spending multiplier of more than one based on pre-1980 data.

1.2 Fiscal policy and interest rates

There are mixed results from studies on the interaction between fiscal policy and interest rates. In addition, the methodologies that have been employed to analyse this interaction have also varied. Event studies, SVARs, and combinations of the two are quite popular. Owing to the approach in this paper, the VAR literature commands the most attention here. This should allow for a clearer comparison of results. Often, bonds are investigated in the paradigm of monetary policy. However, unprecedented government intervention in the economy in response to the Great Recession has made fiscal policy a much more active component in the economic policymaking toolkit. This necessitates analysis of the interaction between long-maturity bond yields and fiscal policy. We include the federal funds rate to represent monetary policy. In a way, we are heeding the caution by Rossi and Zubairy (2011) that "failing to recognize that both monetary and fiscal policy simultaneously affect macroeconomic variables might incorrectly attribute the fluctuations to the wrong source." Favero and Givazzi (2008) analyse the impact of fiscal shocks while allowing the level of public debt to affect the debt service cost, tax, and spending. They explain the absence of fiscal shock effects on long-term interest rates as due to misspecified VARs that exclude a debt feedback, in addition to not endogenising debt dynamics. The level of debt-to-GDP ratio, they argue, impacts long-term interest rate.

Marattin et al. (2011) uses cointegration to study the impact of fiscal shocks on government debt and long-term interest rates in the US, Germany, and Italy for the 1983-2009 period. Using the common trends methodology to distinguish transitory shocks (financial and inflationary) from permanent shocks, (fiscal and monetary) they find that (i) debt accumulation is followed by higher long-term interest rates for Germany and Italy but not for the US, whereby the reaction of the long-term interest rate in the US can be explained by liquidity effects, (ii) fiscal shocks determine both permanent and cyclical components of the long-term interest rate. Debt shocks have asymmetric effects depending on debt-to-GDP ratios for Italy, and Germany but not the US. Laubach (2011) studies fiscal variables and interest rates under two scenarios: under heightened levels of sovereign default risk, and when the risk that the government will default on its debt obligations is of no concern to creditors. Again, under normal default risk conditions, results are Keynesian. Contractionary fiscal policy results in lower interest rates, lower real economic activity, and lower growth in prices as the Federal Reserve Board responds with a cut in (short-term) rates. However, a weak economic environment leads to a decline in the surplus-to-GDP ratio, and a subsequent increase in the debt-to-GDP ratio. Laubach (2009) looks at links between projected deficits and public debt on the one hand, and long-horizon forward rates on the other. Deficit and debt effects on interest rates are significant, both statistically and economically, with parameter estimates that are consistent with the neo-classical growth model. It is claimed that the effects of deficit and (public) debt projections by the Congressional Budgetary Office (CBO) in the US are evident at the longer end of the yield curve. A percentage point increase in the projected deficit-to-GDP ratio results in interest rates rising by 25 basis points. A percentage point increase in the debt-to-GDP ratio leads to interest rates rising by 3 to 4 basis points. Gale and Orszag (2003) give a broad review of the literature on fiscal policy and interest rates. They suggest that studies that incorporate expected deficits tend to find a link between deficits and interest rates. They argue that deficits may increase nominal interest rates because (i) they reduce aggregate savings if foreign capital inflows and domestic private savings rises are insufficient to compensate for the fall in public saving, and (ii) they lead to an increase in the stock of public debt. A higher outstanding amount of government bonds commands a higher interest rate as investors demand higher a premium to hold additional government bonds as opposed to other financial assets. Also, expected future deficits have significant effects on long-term bond yields. They also emphasise the negative long-term effect of deficits on the economy especially via reduced national saving rates that result in lower national income. In analysis of financial variables behaviour around periods of major changes in the fiscal stance in OECD countries for the period 1960-2002, Ardagna (2009) finds that long-term government bond yields fall in periods of budget consolidation but rise during fiscal deterioration. Where fiscal policy is associated with a permanent reduction in public debt, there is a strong effect on long-term government bond yields. Stock markets surge as a result of fiscal adjustments involving expenditure reductions but fall during fiscal expansions. Importantly, these results depend on prevailing fiscal conditions and the type of fiscal policy: a spending cut that permanently reduces government debt during high-deficit periods is associated with bigger falls in interest rates.

Engen and Hubbard (2004) analyse the effect of federal government debt and interest rates. Owing to the dependence of results on model specification, they use various specifications on the same data set. They conclude that an increase in debt of 1% of GDP, ceteris paribus, increases the real long-term interest rate by about 0.03%. In a discussion of a trend in the long-term interest rate and its drivers, Brook (2003) identifies cyclical and portfolio factors as the main drivers of the 2000-2003 fall in real long-term interest rates. For selected European countries, falling inflation and exchange rates could also have lowered the equilibrium real interest rate. Attention is also drawn to US evidence pointing to a causal link from fiscal variables to real long-term interest rate. This link is characterised by both the actual and projected fiscal fiscal positions, underlying the importance of 'expectations' in this relationship. Further, an international dimension is also explored, with US originated shocks having a more noteworthy effect on European and Japanese bonds than vice versa. The aim of this paper is to estimate the size of the fiscal multiplier and to characterise the interaction between fiscal policy and monetary policy in the post-war US economy using a large parsimonious structural vector equilibrium correction model. Our methodological approach of using a VAR model variant is in line with most of the literature and should therefore aid comparing of results.

The paper will proceed as follows: in §2 we will introduce the data set and discuss the creation of the fiscal variables. §3 discusses the econometric methodology. This is then applied to data in §4, where we discuss particular methodological issues and present the empirical findings. The model is presented in §5. It is used in §6 for the impulse response analysis of three policy experiments: (i) a balanced-budget government spending shock, (ii) a deficit-spending shock, and (iii) a federal funds rate shock.

The results will allow us to estimate the dynamic fiscal multiplier.

2 The data

The aim of this paper is to estimate the size of the fiscal multiplier for the post-war US economy using a large parsimonious structural vector equilibrium correction model (PSVECM). We collect quarterly data on the US economy for the 1960Q1-2013Q4 period, with the sample size dictated by data availability. The data include GDP, government spending, net taxes, inflation expectations, the GDP deflator, the personal consumption expenditures (PCE) price index,[2] the federal funds rate, the 30-year maturity government bond yield, and the 30-year maturity Moody's Baa-rated corporate bond yield. Table 1 gives details and transformations of the time series characterising our 8-dimensional system.

Table 1 *Variable details*

Time series Description	Source: Label	Model variables
GDP, SA, AR	BEA: GDP	$y_t = 100\ log(\text{GDP}/\text{GDPDEF})$
Net taxes, SA, AR	BEA: T	$\tau_t = 100\ log(\text{T}/\text{GDPDEF})$
Govt spending, SA, AR	BEA: G	$g_t = 100\ log(\text{G}/\text{GDPDEF})$
GDP deflator, SA	BEA: GDPDEF	
Inflation expectations	University of Michigan*	$\pi_t^e = 100\ log(1+\Pi^e/100)$
PCE deflator, SA	BEA : PCEPI	$\pi_t = 400\ \Delta log(\text{PCEPI})$
Federal funds rate, NSA	FRED: FEDFUNDS	$i_t = 100\ log(1\text{+FEDFUNDS}/100)$
Government bond yield, NSA	FRED: GS30†	$r_t = 100\ log(1\text{+GS30}/100)$
Corporate bond yield, NSA	FRED: Baa	$b_t = 100\ log(1\text{+Baa}/100)$

* The mean of the expected change in prices during the next year from Table 32 in the Survey of Consumers.
† Own calculation of missing data based on GS20, see Krolzig and Sserwanja (2014), for details.

We define net tax and spending in the spirit of Blanchard and Perotti (2002). This should allow for comparing of results with this widely cited study on US fiscal policy. Net taxes are defined as the sum of tax and non-tax receipts, contributions for government social insurance less net transfer payments and net interest payments. We use net tax as opposed to gross tax because we are interested in the government budget constraint. Taking gross tax would ignore the fact that part of the current receipts have to be spent on prior commitments such as interest payments and social security transfers. Government spending is defined as consumption expenditures and gross government investment. We discuss this matter in more detail in the following section.

2.1 Measuring government spending and net taxation

The literature on US fiscal policy from the last decade refers extensively to the Blanchard and Perotti (2002) study, making it a good benchmark to compare our results. They too work with total government data, defining net taxes as "the sum of Personal Tax and Non-tax Receipts, Corporate Profits Tax Receipts, Indirect Business Tax and Nontax Accruals, and Contributions for Social Insurance, less Net

[2] 'Inflation expectations' is a survey-based measure that moves more closely with consumption-based price indices than with the GDP deflator. To exploit this potential long-run relation, we use the PCE index to calculate inflation while the GDP deflator is retained to deflate GDP, net taxes, and spending.

Table 2 *Definition of fiscal variables*

	consumption expenditures	(16)
+	gross government investment	(34)
=	government spending	

	personal current taxes	FG(3)
+	taxes on production and imports	FG(4)
+	taxes on corporate income	FG(7)
+	taxes from the rest of the world	FG(10)
+	contributions for government social insurance	FG(11)
+	income receipts on assets	FG(12)
+	current transfer receipts	FG(16)
−	current transfer payments to persons	FG(24)
−	interest payments	FG(29)
+	personal current taxes	SLG(3)
+	taxes on production and imports	SLG(6)
+	taxes on corporate income	SLG(10)
+	contributions for government social insurance	SLG(11)
+	income receipts on assets	SLG(12)
+	current transfer receipts	SLG(16)
−	federal grants-in-aid	SLG(17)
−	government social benefit payments to persons	SLG(23)
−	interest payments	SLG(24)
=	net taxation	

Note: The numbers in parentheses correspond to tax and spending component lines in Tables 3.1, 3.2 and 3.3 (BEA, June 2014). FG: federal government, SLG: state and local government.

Transfer Payments to Persons and Net Interest Paid by Government. Government spending is Purchases of Goods and Services, both current and capital" (Blanchard and Perotti (2002), p.1336). Our definition of the fiscal variables (see Table 2) follows Blanchard and Perotti (2002), but the two sets of series differ slightly. This difference is inevitable, given the numerous data revisions, and component re-classifications that the National Income and Product Accounts (NIPA) tables have undergone over the last twelve years.[3] It is worth noting that the data cover total government, i.e., federal, state and local governments.

Table 3 *Some statistics: our fiscal series versus Blanchard and Perotti (2002)*

	Study	τ	g	$\tau - g$
mean	B&P (2002)	5.945	6.138	-0.193
	1960Q1-1997Q4	5.816	6.195	-0.379
	1960Q1-2013Q4	6.269	6.673	-0.404
median	B&P (2002)	6.070	6.175	-0.179
	1960Q1-1997Q4	5.953	6.215	-0.360
	1960Q1-2013Q4	6.459	6.909	-0.365
std. dev.	B&P (2002)	0.786	0.833	0.087
	1960Q1-1997Q4	0.786	0.826	0.097
	1960Q1-2013Q4	0.996	1.021	0.186

[3] Blanchard and Perotti (2002) takes 'federal corporate profits tax' data from the Quarterly Treasury Bulletin. Our fiscal data are solely from the NIPA.

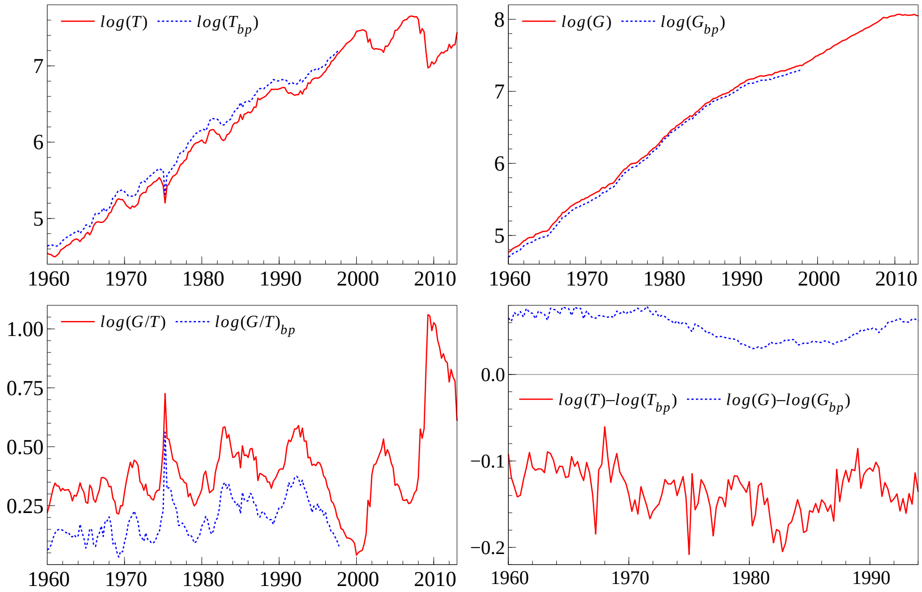

Figure 1 *Our fiscal series versus Blanchard and Perotti (2002)*

The fiscal series in this study are compared to those of Blanchard and Perotti (2002) in figure 1. For the period over which the two samples overlap, we can see from the top right graph that spending in both studies is almost identical. This view is further supported by their difference plot in the bottom graph. It appears that our spending variable is just a level shift up. Differences in tax are more pronounced, with a level shift downwards in our tax measure obvious in the top left and bottom right graphs. This level shift between the two data sets is also clear in the deficit plots given by the lower left graph. These differences notwithstanding, we believe that our tax and spending definitions approximate well those in Blanchard and Perotti (2002), as is evident in Figure 1 and Table 3.

2.2 Inflation

Inflation is represented by the forward-looking inflation expectations measure from the Survey of Consumers run by the University of Michigan/Reuters and an actual inflation measure from the PCE price index. Inflation expectations are included to capture agent expectations of the path of prices in the economy and to reduce the 'price puzzle' problem common in VAR analysis. This puzzle involves a price rise in response to tightening monetary policy, such as an increase in the federal funds rate. According to Figure 2, the relationship between these two measures appears stable over time, making them viable candidates for a cointegration relation. This is supported by the result in Table 6 where $\pi - \pi^e$ is $I(0)$.

2.3 Interest rates

The three interest rates represent the financial sector with the federal funds rate as the monetary policy proxy. The Treasury bond yield is the cost of borrowing for thirty years by the government. The corporate bond yield represents the average interest rate faced by firms when they issue debt maturing in thirty years time. Treasury bond yield data is only available from February 1977 and is also missing

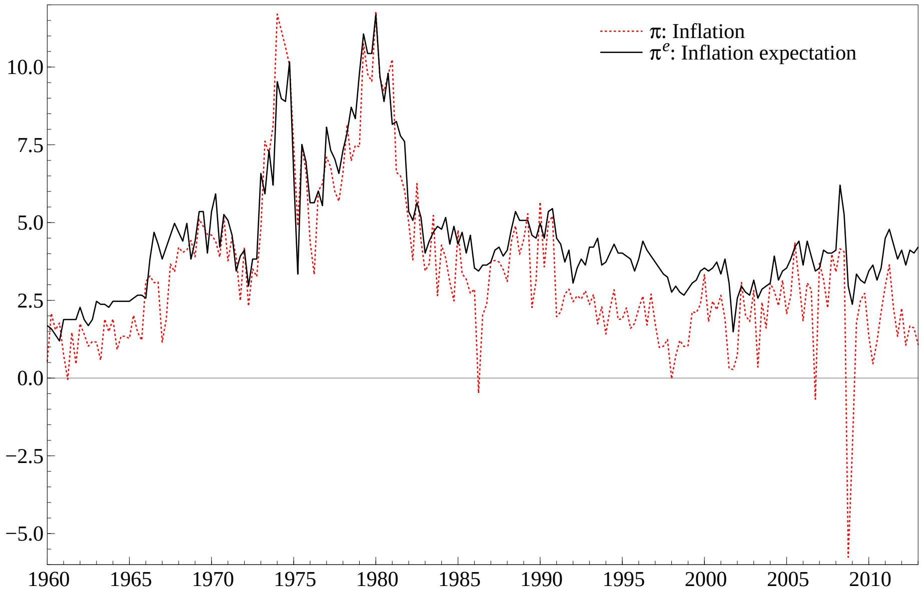

Figure 2 *Inflation and expected inflation*

between 2002M3 and 2006M1, a period during which no 30-year maturity Treasuries were issued by the government. All interest rate series are plotted in Figure 3.

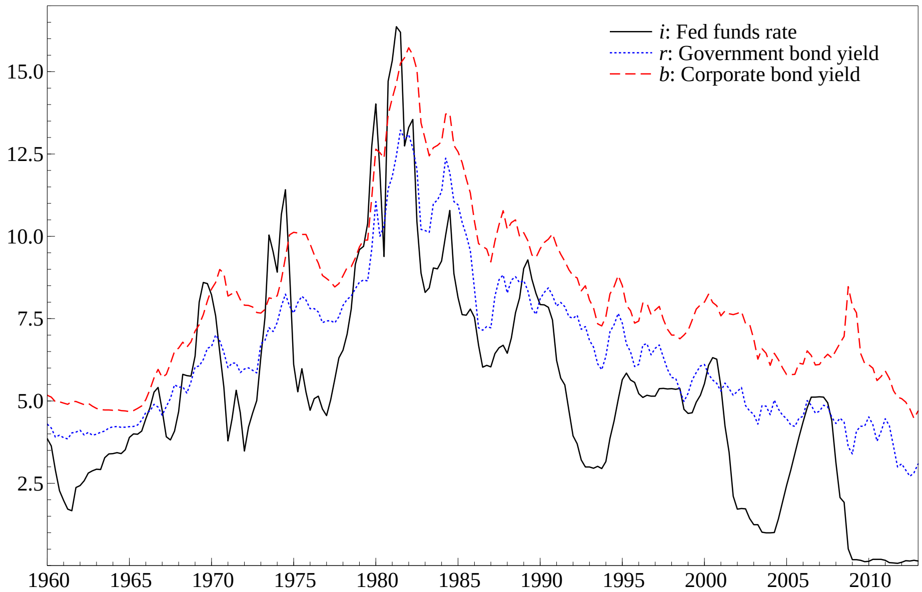

Figure 3 *Interest rates: the federal funds rate, the government bond yield, and the corporate bond yield*

2.4 Unit roots

The unit-root test profiles of variables are in tables 4, 5, and 6. Only the tax share of GDP appears to be $I(0)$. All the other variables are $I(1)$ and so have to be differenced once to make them stationary. We also test for unit roots on a selection of economic theory-inspired relations. This serves as a preliminary check on the (expected) stationarity of prospective cointegration relations. They include: (i) the deficit $(\tau - g)_t$, (ii) the term structure of interest rates $(r - i)_t$, (iii) the risk premium $(b - r)_t$, (iv) the actual-expected inflation difference $(\pi^e - \pi)_t$, and (v) the real short-term interest rate $(i - \pi)_t$. They are all $I(0)$ as expected. That the deficit and $(\tau - y)_t$ are $I(0)$ while $(g - y)_t$ is $I(1)$ is a contradiction because it implies that an $I(0)$ variable cointegrates with an $I(1)$ variable. However, at this stage it is helpful to keep in mind that unit root tests are only suggestive and not definitive given their low power. Furthermore, $(\tau - y)_t$ actually has a unit root (p-value 0.277) if the test is specified with 0 lags as selected by the Schwarz information criterion.[4] Also, the ADF finds the deficit to be $I(0)$ regardless of whether the test lags are chosen based on the Akaike, Schwarz or Hannan-Quinn criteria. Impulse response analysis will confirm that after a shock, the deficit is only ever temporarily away from equilibrium.

Table 4 *Unit root tests on variables in levels*

	μ	γ	π	p	t-adf(π)	5% CV
y_t	8.930 (0.029)*	0.017 (0.065)	-0.024	2 (0.014)*	-2.043 (0.574)	-3.401
$(\tau - y)_t$	-19.047 (0.000)**	-0.024 (0.003)**	-0.109	5 (0.146)	-4.162 (0.006)**	-3.431
τ_t	14.060 (0.000)**	0.044 (0.001)**	-0.081	4 (0.005)**	-3.747 (0.021)**	-3.431
$(g - y)_t$	-6.464 (0.003)**	-0.005 (0.002)**	-0.044	3 (0.001)**	-3.006 (0.133)	-3.431
π^e_t	0.325 (0.023)*	-0.001 (0.230)	-0.069	2 (0.002)**	-2.369 (0.152)	-2.875
π_t	0.334 (0.039)*	-0.002 (0.189)	-0.101	2 (0.001)**	-2.598 (0.095)	-2.875
i_t	0.171 (0.121)	-0.002 (0.053)	-0.009	7 (0.005)**	-1.030 (0.272)	-1.942
r_t	0.087 (0.265)	-0.001 (0.080)	-0.002	5 (0.037)*	-0.484 (0.505)	-2.875
b_t	0.119 (0.138)	-0.001 (0.185)	-0.001	1 (0.000)**	-0.466 (0.512)	-2.942

Sample: 1960Q1-2013Q4. Tests based on $\Delta x_t = \mu + \gamma t + \pi x_{t-1} + \sum_{j=1}^{p} \phi_j \Delta x_{t-j} + \varepsilon_t$ with the lag length selected by the Akaike Information Criteria (AIC). The null hypothesis, $H_0 : \pi = 0$, implies presence of a unit root and is rejected when t-adf $<$ CV.

[4] However, the Hannan-Quinn criterion (HQ) selects 4 lags and $(\tau - y)_t$ is $I(0)$ according to a 0.013 p-value, in line with the Akaike information criteria (AIC) benchmark test specification. $(g - y)_t$ is $I(0)$, regardless of the information criteria used in the test specification.

Table 5 *Unit root tests on variables in differences*

	μ	γ	π	p	t-adf(π)	5%*CV*
Δy_t	0.428 (0.000)**	-0.002 (0.062)	-0.558	1 (0.014)*	-7.168 (0.000)**	-2.875
$\Delta(g-y)_t$	0.005 (0.972)	-0.001 (0.651)	-0.543	2 (0.007)**	-5.913 (0.000)**	-3.431
$\Delta\pi^e_t$	0.098 (0.389)	-0.001 (0.409)	-1.519	1 (0.00)**	-14.491 (0.000)**	-3.431
$\Delta\pi_t$	0.149 (0.452)	-0.001 (0.392)	-2.234	8 (0.006)**	-6.631 (0.000)**	-3.431
Δi_t	0.105 (0.342)	-0.001 (0.242)	-0.914	6 (0.003)**	-6.608 (0.000)**	-3.341
Δr_t	0.002 (0.926)	-0.001 (0.121)	-0.89	4 (0.036)*	-7.255 (0.000)**	-1.942
Δb_t	0.000 (0.999)	-0.000 (0.209)	-0.654	0	-10.219 (0.000)**	-1.942

Table 6 *Unit root tests on prospective cointegration relations*

	μ	γ	π	p	t-adf(π)	5%*CV*
$(\tau-g)_t$	-2.748 (0.003)**	-0.009 (0.179)	-0.067	6 (0.074)	-3.140 (0.025)*	-2.876
$(r-i)_t$	-0.071 (0.486)	0.003 (0.002)	-0.223	8 (0.010)	-5.126 (0.000)**	-3.431
$(b-r)_t$	0.209 (0.000)**	-0.001 (0.057)	-0.130	1 (0.000)**	-4.363 (0.001)**	-2.875
$(\pi^e-\pi)_t$	0.057 (0.703)	-0.004 (0.008)**	-0.440	5 (0.002)**	-3.852 (0.016)*	-3.431
$(i-\pi^e)_t$	0.032 (0.673)	-0.002 (0.097)	-0.059	7 (0.002)**	-1.982 (0.046)*	-1.942
$(i-\pi)_t$	0.229 (0.082)	-0.002 (0.365)	-0.081	2 (0.000)**	-2.357 (0.018)*	-1.942

3 Econometric methodology

In contrast to the existing literature, we follow a data-driven modelling approach that combines the VAR based cointegration analysis of Johansen (1995) and Juselius (2006) with the graph-theoretic approach of Spirtes et al. (2001) implemented in TETRAD for the search for instantaneous causal relations and the automatic general-to-specific model selection algorithm implemented in *PcGets* of Krolzig and Hendry (2001) for the selection of a congruent parsimonious structural vector equilibrium correction model.

3.1 Methodology

The General-to-specific approach implemented in this paper follows the modelling approach of Krolzig (2003) and consists of the following four stages (see Demiralp et al. (2009), for a related approach):

(i) *Specifying the general unrestricted system.*

We commence from a reduced-form vector autoregressive (VAR) model of order p and dimension K, without any equation-specific restrictions, to capture the characteristics of the data:

$$\boldsymbol{x}_t = \nu + \sum_{j=1}^{p} \boldsymbol{A}_j \boldsymbol{x}_{t-j} + \varepsilon_t, \quad \varepsilon_t \sim \mathsf{NID}(\mathbf{0}, \boldsymbol{\Sigma}), \tag{1}$$

where ε_t is a Gaussian white noise process. This step involves the specification of the deterministic terms, selection of the lag length, p, and misspecification tests to check the validity of the assumptions made.

(ii) *Johansen cointegration tests* and *identification of cointegration vectors.*

The Johansen procedure for determining the cointegration rank, r, is then applied to the VAR in (1) mapped to its vector equilibrium-correction mechanism (VECM) representation:

$$\Delta \boldsymbol{x}_t = \boldsymbol{\nu} + \boldsymbol{\Pi} \boldsymbol{x}_{t-1} + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \boldsymbol{x}_{t-j} + \varepsilon_t. \tag{2}$$

For a cointegrated vector process, the reduced-rank matrix, $\boldsymbol{\Pi}$, can be decomposed into a $K \times r$ dimensional loading matrix, $\boldsymbol{\alpha}$, and cointegration matrix, $\boldsymbol{\beta}$, containing the information of the long-run structure of the model, i.e., $\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$. The Johansen procedure delivers unique estimates of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ as a result of requiring $\boldsymbol{\beta}$ to be orthogonal and normalized. These estimates provide a value for the unrestricted log-likelihood function to be compared to the log-likelihood under economically meaningful overidentifying restrictions, $\boldsymbol{\beta}^r$:

$$\Delta \boldsymbol{x}_t = \boldsymbol{\nu} + \boldsymbol{\alpha}(\gamma t + \boldsymbol{\beta}^{r'} \boldsymbol{x}_{t-1}) + \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \Delta \boldsymbol{x}_{t-j} + \varepsilon_t, \tag{3}$$

with $\boldsymbol{\Sigma} = \mathsf{E}[\varepsilon_t \varepsilon_t']$. The empirical modelling procedure for finding the cointegration relations follows Juselius (2006).

(iii) *Graph-theoretic search for instantaneous causal relations.*

The determination of the contemporaneous relationships between the variables has been advanced by modern graph-theoretic methods of searching for causal structure based on relations of conditional independence developed by computer scientists Pearl (2000) and philosophers Spirtes et al. (2001). Following Demiralp and Hoover (2003), who introduced this approach to econometrics, we use the PC algorithm implemented in TETRAD IV (see Spirtes et al. (2005) for details). The PC algorithm exploits the information embedded in the residual variance-covariance matrix, $\hat{\Sigma}$, of the system in (3). A causal structure is represented by a graph with arrows indicating the direction of causality between variables. To detect the directed acyclic graph, the algorithm starts by assuming that all variables are linked to each other through an undirected link. In the elimination stage, connections are first removed between variables which are unconditionally uncorrelated. Then connections are eliminated for variables which are uncorrelated conditional on other variables. Having identified the skeleton of the graph, the orientation step of the algorithm seeks to orient the undirected edges by logical reasoning. This involves the analysis of indirect connections by taking into account the whole graph, considering every pair of variables, exploiting already directed edges and the acyclicality condition.

Once all edges are oriented, a directed acyclic graph (DAG) results. Based on the identified contemporaneous causal structure of the system, the VECM in (3) can be represented as a *recursive* SVECM. By a suitable ordering of the variables of the system, the DAG can be mapped to a lower-triangular contemporaneous matrix, $\boldsymbol{B}^r$, with units on the diagonal and non-zero lower-off-diagonal elements representing the causal links found by the PC algorithm. In contrast to a traditional orthogonalisation with the help of a Choleski decomposition of $\hat{\Sigma}$, this approach results in an overidentified SVECM in the majority of cases. The zero lower-triangular elements of $\boldsymbol{B}^r$ provide testable overidentifying constraints allowing to verify the validity of the selected contemporaneous structure. Most importantly, as the contemporaneous causal structure captured by $\boldsymbol{B}^r$ is data determined, it avoids the problems associated with the ad-hoc nature of orthogonalised structural VAR models.[5]

(iv) *Single-equation reductions of the recursive SVECM.*

Here we consider *Gets* reductions of the SVECM to reduce the complexity of the model and to mitigate the curse of dimensionality. Starting point is the SVECM with the long-run relations $\boldsymbol{\beta}^r$ determined by stage (ii) and contemporaneous structure $\boldsymbol{B}^r$ given by the corresponding directed acyclic graph:

$$\boldsymbol{B}^r \Delta \boldsymbol{x}_t = \boldsymbol{\delta} + \tilde{\boldsymbol{\alpha}} \left(\gamma t + \boldsymbol{\beta}^{r\prime} \boldsymbol{x}_{t-1} \right) + \sum_{j=1}^{p-1} \boldsymbol{\Upsilon}_j \Delta \boldsymbol{x}_{t-j} + \boldsymbol{\eta}_t, \quad \boldsymbol{\eta}_t \sim \mathsf{NID}(\boldsymbol{0}, \boldsymbol{\Omega}), \tag{4}$$

where $\boldsymbol{B}^r$ is the lower-triangular matrix found by Tetrad and $\boldsymbol{\Omega} = \boldsymbol{B}^r \Sigma \boldsymbol{B}^{r\prime}$ is a diagonal variance-covariance matrix. A single-equation based *Gets* reduction procedure such as *PcGets* can be applied to the equations in (4) straightforwardly and, as shown in Krolzig (2001), without loss in efficiency. The parameters of interest are the coefficients collected in the intercept, $\boldsymbol{\delta}$, the adjustment matrix $\tilde{\boldsymbol{\alpha}}$ and the short-run matrices $\boldsymbol{\Gamma}_j$ in the structural VECM. The result is a parsimonious

[5] If the PC algorithm finds a link but has insufficient information to identify if, say, 'A causes B' or 'B causes A', an undirected edge emerges. In this case, there exists a set of contemporaneous causal structures, $\{\boldsymbol{B}^{(i)}\}$, that are all consistent with the data evidence. An additional modelling stage is then required for the selection of $\boldsymbol{B}^r$ and, thus, the identification of the direction of causality. Having applied the model reduction step in (iv) to each SVECM associated with one of the found contemporaneous causal structures, the dominant stable econometric model is finally selected in (v).

structural vector equilibrium correction model denoted PSVECM, which is nested in (4) and defined by the selected $\boldsymbol{\delta}^*$, $\tilde{\boldsymbol{\alpha}}^*$ and $\boldsymbol{\Upsilon}_j^*$ with $j = 1, \ldots, p-1$.

(v) *Selection of the dominant PSVECM*.

If the graph-theoretic search in (*iii*) produces an acyclic graph with at least one undirected edge, the determination of the direction of instantaneous causal relations has to rely on the information from the PSVECMs resulting from the *Gets* reduction of the SVECMs as defined by the set of contemporaneous causal structures. As the PSVECMs are mutually non-nested and the union is usually unidentified, we propose to select the PSVECM with the greatest penalized likelihood. Thus the dominant design of the contemporaneous effects matrix according to information criteria such as Akaike or Schwarz would be used.

4 Empirical findings

We estimate a SVECM and use it to analyse the dynamics of eight key US economic variables spanning more than fifty years. The modelling procedure is general-to-specific. We start with an unrestricted VAR(p) model and determine the lag length p. We then determine the number of cointegrating relations in the system via the trace test from Johansen (1988) and Johansen (1995). These are linear relations among (some, or all of) the series that are stable in the long-run. Next, we apply the PC algorithm from Pearl (2000) and Spirtes et al. (2001) to the residuals from the identified cointegrated VAR to retrieve the contemporaneous relationships structure among the variables. By mapping this contemporaneous structure into the VECM, we estimate the structural VECM. The SVECM still has several insignificant parameters at this stage. We therefore use the *Gets* reduction procedure in Krolzig (2003) to drop insignificant terms, leaving behind a parsimonious structural model (PSVECM) that is a congruent representation of the DGP. It is with this PSVECM that we implement economic analysis. We also carry out impulse response analysis to understand how shocks are propagated in the system.

4.1 The VAR model

We start by estimating the unrestricted 8-dimensional reduced-form VAR(4) model. $\boldsymbol{x}_t = (y_t,\ \tau_t,\ g_t,\ \pi_t^e,\ \pi_t,\ i_t,\ r_t,\ b_t)'$, $\boldsymbol{\mu}$ is the vector of intercepts and has dimension 8×1, $\boldsymbol{\gamma}$ is the 8×1 vector of trend-coefficients, $\boldsymbol{A}_j$ an 8×8 matrix of coefficients at the j^{th} lag. $\boldsymbol{u}_t$ is a Gaussian vector of white noise processes each with zero mean and $\boldsymbol{\Sigma}$ their variance-covariance matrix.

$$\boldsymbol{x}_t = \boldsymbol{\mu} + \boldsymbol{\gamma} t + \sum_{j=1}^{p} \boldsymbol{A}_j \boldsymbol{x}_{t-j} + \boldsymbol{u}_t, \quad \boldsymbol{u}_t \sim \mathsf{NID}(\mathbf{0}, \boldsymbol{\Sigma}) \tag{5}$$

We need the correct lag length p to estimate (5) and of course, the resulting model should pass the usual diagnostic tests, including having autocorrelation-free residuals. The four starting lags are motivated by a need to be consistent with the vast majority of the literature on fiscal and monetary shocks. This choice is also based on the fact that we are working with quarterly data. To check whether the lag length can be reduced, we use various information criteria in Table 7. Clearly, the criteria do not agree on one, three, or four lags. But the choice of four lags is within the general-to-specific spirit since

Table 7 *VAR lag length determination*

Order	logL	LR	AIC	SC	HQ
0	-3606.282		34.1724	34.4258	34.2748
1	-1881.823	3286.233	18.5773	19.7744**	19.0197**
2	-1795.855	157.338	18.3008	20.5804	19.2220
3	-1722.750	128.277	18.2147**	21.5078	19.5456
4	-1669.105	90.083**	18.3123	22.6188	20.0529

Sample: 1960Q1-2013Q4. ** indicates selected lag order. LR: sequential modified LR test statistic (each test at 5% level). AIC: Akaike information criterion. SC: Schwarz information criterion. HQ: Hannan-Quinn information criterion.

any insignificant terms will be dropped in the final model. We therefore choose four lags as indicated by the LR test.

4.2 Cointegration

Cointegration analysis aims to identify any linear relationships in the data that are stable in the long-run. Looking at the time series plots of our eight variables in figures 2, and 3, there is strong co-movement amongst (some of) the variables. If these co-movements are linear and stable, they could represent cointegration relations. This is supported by results in tables 4 and 6 showing that while all but the $(\tau - y)_t$ variable are $I(1)$ in *levels*, some linear combinations of these variables do produce $I(0)$ *relations*. Potential cointegrating relations include: the deficit, the term structure of interest rates, the risk premium, the expected-actual inflation relation, and the real short-term interest rate. A formal test to determine the number of cointegrating relations is given in Table 8. The Johansen trace test, Johansen (1988), is based on (3), which is just the error-correction form of (5) less $\Delta \boldsymbol{x}_{t-j}$ lags. The number of cointegrating relations is given by the rank of matrix $\mathbf{\Pi}$.

Results in Table 8 show that at the benchmark 5% significance level, there are four cointegrating relations in the data. It is vital at this stage that the correct number of cointegrating relations is determined, given its importance to proceeding analysis. One runs the risk of making wrong inference if more relations are specified than do exist in the data. On the other hand, specifying less relations than exist means excluding important information from the analysis, especially in a long-run dynamics study like ours. To this effect, a careful interpretation of the trace test results is warranted. At the traditional 5% significance level, the trace test in Table 8 rejects $H_0^r : rank(\mathbf{\Pi}) \leq r$, for $r = 0, 1, 2, 3$, but not $r \leq 4$, which is not rejected with 0.073 probability, thereby suggesting four cointegration relations in the data. However, we argue that there are five cointegration relations based on the following reasons. First, results in Table 6 suggest a rank of 5, possibly involving relations: $\{(\tau - g)_t$ or $(\tau - y)_t\}$, $(r - i)_t$, $(b - r)_t$, $(\pi^e - \pi)_t$, and $(i - \pi)_t$. Relations 2-5 were identified in Krolzig and Sserwanja (2014). Second, if relation 1 is the deficit, then it can be justified from economic theory as a binding budget constraint for the government. There is empirical evidence, notably Bohn (1998) and Blanchard and Perotti (2002), among others, that the deficit is indeed stationary, albeit at only the 5% significance level. Finally, while a looser criterion for rejecting $r = 4$, 7.3% is still less than 10%, a threshold that is sometimes used in the literature. Certainly, the 18.4% p-value on which the acceptance of rank of 5 is based is a much stronger criterion. We can therefore be reasonably confident that there are 5 linear and stable long-run relations

Table 8 *Cointegration rank test of Johansen (1988) for CVAR with restricted trend and unrestricted constant*

null	trace statistic	5% CV	p-value
$r = 0$	254.208***	187.470	0.0000
$r \leq 1$	187.395***	150.558	0.0001
$r \leq 2$	136.994***	117.708	0.0017
$r \leq 3$	91.129**	88.803	0.0336
$r \leq 4$	61.848*	63.876	0.0732
$r \leq 5$	36.635	42.915	0.1840
$r \leq 6$	14.379	25.872	0.6262
$r \leq 7$	5.129	12.517	0.5778

*,**,*** indicate significance at 10%, 5% and 1% significance, respectively.

in the data. The five cointegration relations indeed suffice to identify the long-run with a 0.1959 joint p-value of the LR test for the over-identifying constraints.

4.3 Contemporaneous causal relationships between variables

There are various ways to identify monetary and fiscal policy shocks as discussed in Perotti (2007). These are often motivated by what the researcher assumes to be the correct transmission mechanism of the policy in question. Choleski decomposition is a popular identification technique.[6] However, this approach is criticised for assuming the causality direction. It ignores agents who anticipate policy changes and alter their actions accordingly. Misidentifying shocks this way could therefore lead to under-estimation of fiscal multipliers, as pointed out in Ricco (2014). One solution to the 'lack-of-anticipation' critique levelled at traditional identification procedures is the 'narrative' approach in Ramey and Shapiro (1998), Edelberg et al. (1999), Burnside et al. (2004), Romer and Romer (2010), Ramey (2011) and in part, Blanchard and Perotti (2002), among others.[7] This route often involves pinpointing particular times in history when tax or spending policy changed but not in response to the prevailing state of the economy. These could be increased spending to finance a war or an unanticipated increase in taxes. These shocks are therefore deemed exogenous. Imposing sign restrictions on impulse responses as in Mountford and Uhlig (2009) is another route to shock identification.

How we identify shocks in this study is motivated by the need to let the data speak. Furthermore, our data-driven identification produces a SVECM that is a valid reduction of a Choleski-identified system since the imposed restrictions pass the likelihood ratio test. To this effect, we use the *conservative* version of PC (CPC) algorithm implemented in the TETRAD software project at Carnegie Mellon University. Details on how this algorithm is used in economics to determine the contemporaneous causal order in SVAR analysis can be found in Demiralp and Hoover (2003). Spirtes and Glymour (1991), Spirtes et al. (2010), Pearl (2000), and Spirtes et al. (2001) provide further information on this algorithm

[6] This involves constructing a lower-triangular matrix of contemporaneous causal effects. The implication is that a shock to variable one is instantaneously transmitted to all the other variables, a shock to variable two is is transmitted to all other variables but only with a lag to variable one, and so on. A criticism of this approach is it depends on the researcher to determine which shock effects are immediate and which act with a lag. Fatas and Mihov (2001) discuss fiscal shocks identified by Choleski decomposition.

[7] The 'event-study' methodology was developed in Romer and Romer (1989) to analyse monetary policy.

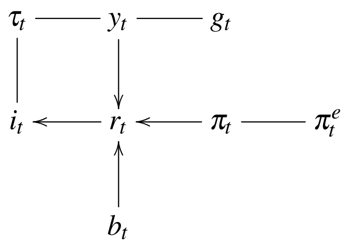

Figure 4 *Instantaneous causality as found by Tetrad*

and its roots in philosophy and computer science. Briefly, it is a graph-theory-based technique that can take as input covariance matrix data and output a directed acyclic graph (DAG) representing the contemporaneous causal pattern in the data generating process. It determines the causal structure among variables by analysing their conditional *in*dependence based on tests of conditional correlation.[8]

We use as input for the algorithm the residuals from the VECM(3) after identifying the long-run structure. The significance level for the dependence test is set to 10%.[9] The graph below gives the causal pattern found by the algorithm. In all, there are 8 significant within-quarter interactions among the variables. However, only half of these are oriented, with the algorithm unable to orient the remaining 4. These have to be oriented via other means.

We orient the remaining four links by analysing the 12 model permutations that result from the non-oriented interactions. The Akaike (AIC), Hannan-Quinn (HQ), and Schwarz (SC) information criteria all choose model 5. However, the long-run (Π) matrix for PSVECM(5) has a positive eigenvalue, suggesting instability. This instability results in explosive impulse responses. On the other hand, Model 6, despite not being preferred by any of the likelihood-based model ranking measures, generates five eigenvalues lying in the $(-1,0)$ interval and 3 zero eigenvalues. This is what is expected according to the five cointegrating relations identified earlier. In fact, model 6 impulse response functions are quite reasonable. Further, Model 6 falls within the 5-8 range of models that are theoretically consistent. We therefore map the causal structure underlying model 6 into matrix $\boldsymbol{B}^r$ and estimate the PSVECM.

The identified causal structure can be mapped into the contemporaneous matrix $\boldsymbol{B}^r$ reported in Table 10. The variables are ordered to ensure that $\boldsymbol{B}^r$ is lower triangular and can be compared to Cholesky decomposition. Matrix $\boldsymbol{B}$ is not unique, but its choice does not affect further analysis. The overidentifying restrictions imposed by the DAG results in a highly parsimonious design.

4.4 Computer-automated Gets single-equation reductions of the SVECM

As high-dimensional VAR models suffer from the curse of dimensionality, a model reduction is required to generate meaningful impulse responses. We are using a general-to-specific model selection algorithm implemented in *PcGets* see Krolzig and Hendry (2001). The starting point is the structural VECM with

[8] If r_{ij} is the unconditional correlation coefficient between variables i and j, then the correlation of i and j conditional on k is $r_{ij|k} = (r_{ij} - r_{ik}r_{jk})/(\sqrt{1-r_{ik}^2}\sqrt{1-r_{jk}^2})$. The significance of the correlation coefficient is determined using Fisher's z-statistic with $H_0 : r_{ij|k} = 0$ and $z = ln((1+r_{ij|k})/(1-r_{ij|k})) \sim \mathsf{N}(0, 1/(T-3))$

[9] The TETRAD 5.1.0-3 manual (p.89) suggests a 5% significance level for dependence for samples of less than 500 observations. With 211 observations, this would suggest the use of $\alpha = 5\%$. However, the overidentifying restrictions imposed by the resulting SVECM are rejected by the likelihood ratio test. Setting $\alpha = 10\%$ gives a SVECM that passes the LR test.

Table 9 *Determining the causal order of undirected links*

PSVECM	Causal Order				AIC	SC	HQ
M_1	y←g,	y←τ,	τ→i,	$\pi^e \leftarrow \pi$	-7.656 (3)	-5.623 (5)	-6.834 (2)
M_2	y←g,	y←τ,	τ→i,	$\pi^e \rightarrow \pi$	-7.561 (9)	-5.607 (7)	-6.771 (7)
M_3	y→g,	y←τ,	τ→i,	$\pi^e \leftarrow \pi$	-7.574 (6)	-5.620 (6)	-6.784 (6)
M_4	y→g,	y←τ,	τ→i,	$\pi^e \rightarrow \pi$	-7.479 (11)	-5.605 (8)	-6.722 (10)
M_5	y←g,	y→τ,	τ←i,	$\pi^e \leftarrow \pi$	-7.684 (1)*	-6.699 (1)*	-6.882 (1)*
M_6	y←g,	y→τ,	τ←i,	$\pi^e \rightarrow \pi$	-7.589 (4)	-5.683 (2)	-6.819 (3)
M_7	y←g,	y→τ,	τ→i,	$\pi^e \leftarrow \pi$	-7.661 (2)	-5.517 (9)	-6.794 (5)
M_8	y←g,	y→τ,	τ→i,	$\pi^e \rightarrow \pi$	-7.566 (7)	-5.501 (10)	-6.732 (9)
M_9	y→g,	y→τ,	τ←i,	$\pi^e \leftarrow \pi$	-7.588 (5)	-5.650 (3)	-6.805 (4)
M_{10}	y→g,	y→τ,	τ←i,	$\pi^e \rightarrow \pi$	-7.493 (10)	-5.634 (4)	-6.742 (8)
M_{11}	y→g,	y→τ,	τ→i,	$\pi^e \leftarrow \pi$	-7.565 (8)	-5.468 (11)	-6.717 (11)
M_{12}	y→g,	y→τ,	τ→i,	$\pi^e \rightarrow \pi$	-7.470 (12)	-5.453 (12)	-6.655 (12)

Figure in parenthesis gives a model's ranking out of 12.

$$
\begin{tikzcd}
\tau_t & y_t \arrow[l] \arrow[d] & g_t \arrow[l] & \\
i_t \arrow[u] & r_t \arrow[l] & \pi_t \arrow[l] & \pi_t^e \arrow[l] \\
 & b_t \arrow[u] & &
\end{tikzcd}
$$

Figure 5 *Instantaneous causality: selected directed acyclic graph*

Table 10 *Instantaneous causality: contemporaneous feedback matrix*

$$
\boldsymbol{B}^r \boldsymbol{x}_t = \begin{pmatrix}
1 & . & . & . & . & . & . & . \\
b_{yg} & 1 & . & . & . & . & . & . \\
0 & 0 & 1 & . & . & . & . & . \\
0 & 0 & b_{\pi\pi^e} & 1 & . & . & . & . \\
0 & 0 & 0 & 0 & 1 & . & . & . \\
0 & b_{ry} & 0 & b_{r\pi} & b_{rb} & 1 & . & . \\
0 & 0 & 0 & 0 & 0 & b_{ir} & 1 & . \\
0 & b_{\tau y} & 0 & 0 & 0 & 0 & b_{\tau i} & 1
\end{pmatrix}
\begin{pmatrix}
g_t \\ y_t \\ \pi_t^e \\ \pi_t \\ b_t \\ r_t \\ i_t \\ \tau_t
\end{pmatrix}
$$

Note: . indicates Choleski-type upper-triangular zero constraints;
0 indicates over-identifying zero constraints.

the identified long-run relations $\boldsymbol{\beta}$ from the Johansen (1995)-Juselius (2006) approach and contemporaneous structure $\boldsymbol{B}^r$ given by the corresponding directed acyclic graph:

$$\boldsymbol{B}^r\Delta\boldsymbol{x}_t = \boldsymbol{\delta} + \boldsymbol{\alpha}\delta t + \boldsymbol{\alpha}\boldsymbol{\beta}'\boldsymbol{x}_{t-1} + \sum_{j=1}^{p-1}\boldsymbol{\Upsilon}_j\Delta\boldsymbol{x}_{t-j} + \boldsymbol{\omega}_t, \quad \boldsymbol{\omega}_t \sim \mathsf{NID}(\mathbf{0},\boldsymbol{\Omega}), \tag{6}$$

where $\boldsymbol{B}^r$ is the lower-triangular matrix found by TETRAD and $\boldsymbol{\Omega}$ is a diagonal variance-covariance matrix. A single-equation based *Gets* reduction procedure such as *PcGets* can be applied to the equations in (6) straightforwardly and, as shown in Krolzig (2001), without a loss in efficiency. The parameters of interest are the coefficients collected in the intercept, $\boldsymbol{\delta}$, the adjustment matrix $\tilde{\boldsymbol{\alpha}}$ and the short-run matrices $\boldsymbol{\Upsilon}_j$ in the structural VECM. The result is a parsimonious structural vector equilibrium correction model denoted PSVECM, which is nested in (6) and defined by the selected $\boldsymbol{\delta}^*$, $\tilde{\boldsymbol{\alpha}}^*$ and $\boldsymbol{\Upsilon}_j^*$ with $j = 1,\ldots,p-1$.

5 The model

We have estimated a parsimonious structural model (PSVECM) to analyse interaction between fiscal policy and interest rates regarding their effect on output. Misspecification tests in Table 11 reveal that some of the model residuals are heteroscedastic. However, our model passes most of the usual diagnostic tests. Critically, all equations pass the autocorrelation test (up to 4^{th} order). Also, none suffers ARCH effects. The RESET test is also passed although the tax equation only passes at the lower 1% threshold. The few violations notwithstanding, our model seems satisfactory for use in analysis.

Table 11 *PSVECM misspecification tests*

Test	Δy_t	$\Delta\tau_t$	Δg_t	$\Delta\pi_t^e$	$\Delta\pi_t$	Δi_t	Δr_t	Δb_t
AR 1-4:	0.99	0.86	1.58	1.09	0.87	1.09	0.68	0.46
	(0.4149)	(0.4904)	(0.1820)	(0.3644)	(0.4829)	(0.3644)	(0.6056)	(0.7666)
Normality	0.084	28.48**	2.79	17.21**	3.51	19.68**	9.87**	8.70*
	(0.9587)	(0.0000)	(0.2480)	(0.0002)	(0.1725)	(0.0001)	(0.0072)	(0.0129)
ARCH 1-4:	1.96	2.41	0.74	2.14	0.71	2.32	0.75	1.85
	(0.1026)	(0.0509)	(0.5628)	(0.0769)	(0.5843)	(0.0590)	(0.5624)	(0.1203)
Hetero:	1.55*	4.84**	0.53	0.92	1.43	3.538**	1.30	1.68*
	(0.0426)	(0.0000)	(0.9487)	(0.5595)	(0.1378)	(0.0000)	(0.1428)	(0.0176)
RESET:	0.002	6.76*	0.57	1.24	0.99	1.10	3.79	0.14
	(0.9650)	(0.0100)	(0.4510)	(0.2664)	(0.3210)	(0.2962)	(0.0530)	(0.7100)

*,** indicates rejection at 5% and 1% significance level, respectively. Numbers in () are p-values.

The model dynamics of output, government spending and net taxes are captured in equations (7)-(9), along with their standard errors. Most coefficients are theoretically-consistent, especially those on long-run relations which are our main interest. The dynamics of real GDP are determined by (7), which shows that the deficit has a negative effect on GDP growth. A 1% increase in G/T is followed by a 0.0079% fall in real GDP in the next quarter, ceteris paribus. While statistically significant, this magnitude is arguably of little economic relevance. On the other hand, increasing the liquidity premium by a percentage point elicits a 0.17% rise in real output. After spending rises by 1%, GDP increases by

0.21% contemporaneously.

$$
\begin{aligned}
\Delta\widehat{y}_t \; = \; & \underset{(0.1)}{0.98} - \underset{(0.003)}{0.0079}\,(g-\tau)_{t-1} + \underset{(0.03)}{0.17}\,(r-i)_{t-1} + \underset{(0.06)}{0.14}\,\Delta y_{t-1} + \underset{(0.008)}{0.038}\,\Delta\tau_{t-2} + \underset{(0.04)}{0.21}\,\Delta g_t \\
& - \underset{(0.04)}{0.1}\,\Delta g_{t-3} - \underset{(0.06)}{0.12}\,\Delta\pi^e{}_{t-2} - \underset{(0.05)}{0.11}\,\Delta\pi^e{}_{t-3} - \underset{(0.03)}{0.076}\,\Delta\pi_{t-1} + \underset{(0.03)}{0.047}\,\Delta\pi_{t-2} + \underset{(0.06)}{0.18}\,\Delta i_{t-1} \\
& - \underset{(0.05)}{0.18}\,\Delta i_{t-2} + \underset{(0.2)}{0.52}\,\Delta r_{t-1} - \underset{(0.2)}{0.96}\,\Delta b_{t-1} + \underset{(0.4)}{1.6}\,\Delta i1971Q1_t + \underset{(0.6)}{2.7}\,\mathrm{i1978Q2}_t \\
& \hat{\sigma} = 0.56,\; R^2 = 0.58
\end{aligned} \tag{7}
$$

Equation (8) suggests taxes are highly endogenous, reacting to four out of five cointegration relations. Taxes are error-correcting via the deficit with a 1% increase in G/T leading to a 0.042% rise in real tax. In other words, an increase in the deficit is followed by a tax rise.

$$
\begin{aligned}
\Delta\widehat{\tau}_t \; = \; & \underset{(0.01)}{0.042}\,(g-\tau)_{t-1} - \underset{(0.17)}{0.45}\,(r-i)_{t-1} - \underset{(0.35)}{2.4}\,(b-r)_{t-1} + \underset{(0.2)}{0.57}\,(\pi-\pi^e)_{t-1} + \underset{(0.3)}{2.2}\,\Delta y_t \\
& + \underset{(0.3)}{1.2}\,\Delta y_{t-1} - \underset{(0.05)}{0.09}\,\Delta\tau_{t-1} - \underset{(0.7)}{1.82}\,\Delta b_{t-1} - \underset{(2.5)}{18.7}\,\Delta i1975Q2_t - \underset{(3.5)}{18.3}\,\mathrm{i2008Q2}_t \\
& + \underset{(3.5)}{15.5}\,\mathrm{i2013Q1}_t + \underset{(2.4)}{13.6}\,\Delta i2013Q2_t \\
& \hat{\sigma} = 3.42,\; R^2 = 0.62
\end{aligned} \tag{8}
$$

The liquidity premium has a negative effect on taxes, which is surprising since a positive liquidity premium indicates expected increase in economic activity, and therefore an increase in tax collections. One explanation of the sign is that inflation from a relatively higher government bond causes the *real* tax fall. A rise in $(b-r)$, which tends to widen during periods of low economic activity, leads to lower taxes. Combining long-run effects from the liquidity and risk premia, we infer that the government bond yield has a positive effect on tax. For the final long-run relation effect, when inflation rises above inflation expectations by a percentage point, (real) taxes rise by 0.57%. Short-run effects are dominated by output and some dummies. Dummies Δi1975Q2 and i2008Q2 capture the 1975 (temporary) tax cut and the 2007-08 tax rebate. i2013Q1 and Δi2013Q2 are likely due to effects from expiration of the Bush tax cuts of 2001 and 2003.[10] Spending is also error-correcting via the deficit. Interestingly, the deficit is the only long-run relation affecting spending. At a speed of 1.6% per quarter, spending adjusts slower than tax (4% adjustment speed per quarter) to maintain the stationary relationship.

$$
\begin{aligned}
\Delta\widehat{g}_t \; = \; & \underset{(0.1)}{1.1} - \underset{(0.003)}{0.016}\,(g-\tau)_{t-1} - \underset{(0.01)}{0.034}\,\Delta\tau_{t-1} + \underset{(0.06)}{0.18}\,\Delta g_{t-1} - \underset{(0.07)}{0.22}\,\Delta\pi^e{}_{t-2} - \underset{(0.08)}{0.26}\,\Delta\pi^e{}_{t-3} \\
& + \underset{(0.05)}{0.11}\,\Delta\pi_{t-3} - \underset{(0.3)}{0.063}\,\Delta r_{t-3} - \underset{(0.3)}{0.41}\,\Delta b_{t-3} + \underset{(0.6)}{2.3}\,\Delta i1963Q3_t + \underset{(0.9)}{2.9}\,\mathrm{i1965Q3}_t + \underset{(0.9)}{3.7}\,\mathrm{i1967Q1}_t \\
& \hat{\sigma} = 0.85,\; R^2 = 0.35
\end{aligned} \tag{9}
$$

Equations (10) and (11) describe inflation dynamics in the system. Both expected inflation and actual inflation error-correct, with the later adjusting about four times faster (63%) than the former

[10]The 'Economic Growth and Tax Relief Reconciliation Act of 2001' and the 'Jobs and Growth Tax Relief Reconciliation Act of 2003' also known as the 'Bush tax cuts' expired. The highest marginal tax rate also increased from 35% to 39.5%.

(15%) per quarter. In Krolzig and Sserwanja (2014) we find a comparable quarterly adjustment speed of 16% for inflation expectations. The Federal Reserve Board's preference and targeting of low, stable inflation since the early 1980s has probably anchored inflation expectations and created a high degree of hysteresis in this variable.

$$
\begin{aligned}
\Delta\widehat{\pi}^e{}_t \quad = \quad & \underset{(0.1)}{0.54} - \underset{(0.07)}{0.36}\ (b-r)_{t-1} + \underset{(0.04)}{0.15}\ (\pi-\pi^e)_{t-1} - \underset{(0.05)}{0.41}\ \Delta\pi^e{}_{t-1} - \underset{(0.05)}{0.38}\ \Delta\pi^e{}_{t-2} + \underset{(0.03)}{0.067}\ \Delta\pi_{t-3} \\
& - \underset{(0.1)}{0.23}\ \Delta r_{t-3} + \underset{(0.6)}{3.3}\ \mathrm{i1973Q1}_t + \underset{(0.4)}{2.7}\ \Delta 4i1974Q1_t - \underset{(0.6)}{4}\ i1975Q2_t + \underset{(0.6)}{2.3}\ \mathrm{i1977Q1}_t \\
& + \underset{(0.6)}{2.2}\ \mathrm{i2008Q2}_t \\
& \hat{\sigma} = 0.56,\ R^2 = 0.58
\end{aligned}
\tag{10}
$$

Actual inflation, on the other hand, is more likely to be driven by economic fundamentals and so will change as they do. The risk premium is another long-run relation affecting both inflation measures, but differently. When inflation expectations are low during recessions, the risk premium rises as investors move from riskier financial assets and into safer government bonds. However, in an expansion, actual inflation rises and as investors move into equities, corporate bond yields rise relative to government bond yields. Dummies i2008Q2 and i2008Q4 are quite telling for they are likely capturing effects of the 2008 oil price shock, the financial crisis and Great Recession that took hold in late 2008. Oil prices rose steeply around the summer of 2008.

$$
\begin{aligned}
\Delta\widehat{\pi}_t \quad = \quad & - \underset{(0.039)}{0.083}\ (r-i)_{t-1} + \underset{(0.05)}{0.19}\ (b-r)_{t-1} - \underset{(0.07)}{0.63}\ (\pi-\pi^e)_{t-1} - \underset{(0.06)}{0.25}\ \Delta y_{t-3} + \underset{(0.08)}{0.96}\ \Delta\pi^e{}_t \\
& + \underset{(0.09)}{0.31}\ \Delta\pi^e{}_{t-1} - \underset{(0.06)}{0.11}\ \Delta\pi_{t-1} - \underset{(0.9)}{7.4}\ \mathrm{i2008Q4}_t \\
& \hat{\sigma} = 0.89,\ R^2 = 0.49
\end{aligned}
\tag{11}
$$

Equations (12), (13), and (14) represent the financial markets. As the monetary policy equation, (14) has a Taylor Rule effect in the interaction between the federal funds rate and the $(\pi-\pi^e)$ relation. Interaction with the liquidity premium captures the expectations hypothesis of the term structure of interest rates. Interestingly, the federal funds rate is sensitive to the deficit. Monetary policy is clearly accommodative by over 30 basis points during recessions which is when the deficit tends to rise. The bond equations are both error-correction via the risk premium, with the Treasury bond doing so at a faster speed. In addition, they are also both sensitive to the real interest rate, especially the corporate bond. Restrictive monetary policy can therefore act through the real short-term interest rate to stabilize long-term interest rates, and therefore inflation.

$$\begin{aligned}
\Delta\widehat{b}_t \;=\; & \underset{(0.07)}{0.23} - \underset{(0.02)}{0.13}\,(r-i)_{t-1} - \underset{(0.04)}{0.11}\,(b-r)_{t-1} - \underset{(0.01)}{0.05}\,(i-\pi)_{t-1} + \underset{(0.03)}{0.065}\,\Delta y_{t-3} + \underset{(0.004)}{0.011}\,\Delta\tau_{t-1} \\
& + \underset{(0.004)}{0.014}\,\Delta\tau_{t-2} - \underset{(0.004)}{0.010}\,\Delta\tau_{t-3} - \underset{(0.018)}{0.030}\,\Delta g_{t-2} - \underset{(0.02)}{0.054}\,\Delta\pi_{t-1} - \underset{(0.01)}{0.044}\,\Delta\pi_{t-2} - \underset{(0.03)}{0.08}\,\Delta i_{t-1} \\
& - \underset{(0.03)}{0.101}\,\Delta i_{t-2} + \underset{(0.08)}{0.515}\,\Delta r_{t-1} - \underset{(0.08)}{0.194}\,\Delta b_{t-1} + \underset{(0.06)}{0.146}\,\Delta b_{t-2} + \underset{(0.3)}{1.1}\,\text{i1979Q4}_t \qquad (12) \\
& + \underset{(0.2)}{0.90}\,\Delta i1980Q1_t + \underset{(0.3)}{1.8}\,\text{i2008Q4}_t \\
& \hat{\sigma} = 0.25,\; R^2 = 0.58
\end{aligned}$$

$$\begin{aligned}
\Delta\widehat{r}_t \;=\; & - \underset{(0.05)}{0.24} + \underset{(0.03)}{0.2}\,(b-r)_{t-1} - \underset{(0.02)}{0.064}\,(\pi-\pi^e)_{t-1} - \underset{(0.006)}{0.022}\,(i-\pi)_{t-1} + \underset{(0.02)}{0.068}\,\Delta y_t + \underset{(0.01)}{0.029}\,\Delta g_{t-1} \\
& - \underset{(0.01)}{0.038}\,\Delta g_{t-2} - \underset{(0.02)}{0.097}\,\Delta\pi^e{}_{t-1} - \underset{(0.02)}{0.043}\,\Delta\pi^e{}_{t-2} + \underset{(0.01)}{0.04}\,\Delta\pi_t + \underset{(0.02)}{0.047}\,\Delta\pi_{t-1} + \underset{(0.01)}{0.037}\,\Delta\pi_{t-2} \\
& + \underset{(0.02)}{0.039}\,\Delta i_{t-2} + \underset{(0.02)}{0.035}\,\Delta i_{t-3} + \underset{(0.06)}{0.24}\,\Delta r_{t-1} + \underset{(0.04)}{0.88}\,\Delta b_t - \underset{(0.06)}{0.25}\,\Delta b_{t-1} - \underset{(0.1)}{0.56}\,\Delta i1980Q2_t \\
& - \underset{(0.2)}{1.5}\,\text{i2008(4)}_t \\
& \hat{\sigma} = 0.18,\; R^2 = 0.81 \qquad (13)
\end{aligned}$$

$$\begin{aligned}
\Delta\widehat{i}_t \;=\; & - \underset{(0.001)}{0.0033}\,(g-\tau)_{t-1} + \underset{(0.02)}{0.086}\,(r-i)_{t-1} + \underset{(0.03)}{0.088}\,(\pi-\pi^e)_{t-1} + \underset{(0.05)}{0.27}\,\Delta y_{t-1} - \underset{(0.04)}{0.074}\,\Delta g_{t-1} \\
& - \underset{(0.03)}{0.069}\,\Delta g_{t-3} + \underset{(0.05)}{0.098}\,\Delta\pi^e{}_{t-1} + \underset{(0.05)}{0.16}\,\Delta\pi^e{}_{t-2} + \underset{(0.04)}{0.13}\,\Delta\pi^e{}_{t-3} + \underset{(0.05)}{0.37}\,\Delta i_{t-1} - \underset{(0.05)}{0.1}\,\Delta i_{t-2} \\
& + \underset{(0.05)}{0.12}\,\Delta i_{t-3} + \underset{(0.09)}{0.62}\,\Delta r_t + \underset{(0.2)}{0.55}\,\Delta r_{t-1} - \underset{(0.1)}{0.51}\,\Delta r_{t-2} - \underset{(0.2)}{0.68}\,\Delta b_{t-1} + \underset{(0.1)}{0.33}\,\Delta b_{t-2} \qquad (14) \\
& + \underset{(0.3)}{1.5}\,\Delta i1973Q3_t + \underset{(0.5)}{4.8}\,\text{i1980Q4}_t - \underset{(0.3)}{2.7}\,\Delta i1981Q4_t + \underset{(0.6)}{3.3}\,\text{i2009(1)}_t \\
& \hat{\sigma} = 0.46,\; R^2 = 0.71
\end{aligned}$$

Overall, fiscal and monetary variables interact both in the short and long-run. This interaction is both direct - the $(r-i)$ effect in the tax equation, for example - as well as indirect through other variables.

6 Three policy experiments and the fiscal multiplier

Before we consider the impulse response analysis of the model, we look into the measurement problems involved with estimating the fiscal multiplier. The focus is then on the system impulse responses for a ten-year period. Confidence intervals are also included, constructed via bootstrap with 5000 replications following Hall (1992). Based on these results, we will finally provide estimates of the fiscal multiplier before concluding by considering a monetary policy shock with focus on its consequences for fiscal policy.

6.1 Measuring the Keynesian fiscal multiplier

The aim of this paper is to estimate the size of the multiplier for the post-war US economy using a large-scale parsimonious structural vector equilibrium correction model. For the measurement of the fiscal multiplier, we will rely on the methods of impulse response analysis. We shall consider the deficit-spending and the balanced-budget multiplier. The spending multiplier gives the change in GDP due to a change in spending. The balanced budget multiplier gives the change in GDP due to a simultaneous change in spending and tax that leaves the deficit unchanged. If both spending and revenue change by the same amount, the Haavelmo theorem posits that national income will also change by the same amount, suggesting a multiplier of one. Matthiessen (1966) questioned the validity of the 1:1 multiplier, while still claiming it to be non-zero. This is in contrast to a deficit-spending multiplier which is greater than one. Let us briefly refer to the Keynesian multiplier. Consider a closed economy with a Keynesian consumption function and autonomous investment spending. The goods market equilibrium is given by

$$Y = \bar{I} + \bar{C} + c(Y - T) + G \tag{15}$$

such that effective demand is

$$Y = \frac{\bar{I} + \bar{C} - cT + G}{1 - c}. \tag{16}$$

In case of deficit-financed government spending shock, we have $dT = 0$, such that the fiscal multiplier results as in Keynes (1936):

$$M = dY/dG = \frac{1}{1-c}. \tag{17}$$

For a tax-financed government spending shock, $dT = dG$, Haavelmo (1945) showed that

$$M = dY/dG = 1. \tag{18}$$

Since our model is formulated in logs with $y = \log Y$, $g = \log G$, $\tau = \log T$, the multiplier will appear as an elasticity. Therefore, it is worth noting that we should expect (i) under deficit spending ($\Delta T = 0$):

$$\mu \equiv \frac{dy}{dg} \equiv \frac{dY/Y}{dG/G} = \frac{1}{1-c}\frac{G}{Y}, \tag{19}$$

and (ii) for the tax-financed spending $dT = dG$ ('Haavelmo Theorem')

$$\mu \equiv \frac{dy}{dg} \equiv \frac{dY/Y}{dG/G} = \frac{G}{Y} \approx 0.2 \tag{20}$$

Allowing for dynamic responses to shocks requires focussing on the dynamic multiplier. For the Robertson lag in the consumption function, $C_t = \bar{C} + c(Y_{t-1} - T_{t-1})$, for example, we would get for deficit spending:

$$M(0) = 1, \quad M(h) \underset{h\to\infty}{\longrightarrow} \frac{1}{1-c}, \tag{21}$$

$$\mu(0) = 1, \quad \mu(h) \underset{h\to\infty}{\longrightarrow} \frac{1}{1-c}\frac{G}{Y}. \tag{22}$$

Alternatively, we could take into account the cumulated effects of the shock when measuring the dynamic multiplier.

$$\begin{aligned} M_t(h) &= \frac{\sum_{i=0}^{h} dY_{t+i}}{\sum_{i=0}^{h} dG_{t+i}} = \frac{\sum_{i=0}^{h} Y_{t+i} dy_{t+i}}{\sum_{i=0}^{h} G_{t+i} dg_{t+i}} \\ &\approx \overline{G/Y}^{-1} \mu_t(h) = \overline{G/Y}^{-1} \frac{\sum_{i=0}^{h} dy_{t+i}}{\sum_{i=0}^{h} dg_{t+i}} \end{aligned} \tag{23}$$

where dy_{t+h} and dg_{t+h} denote the responses to a fiscal shock h quarters ago.

The analysis above could be expanded by considering an IS-LM model as a simple framework for exploring the interaction between budget deficits and real output. The Keynesian model posits that fiscal policy influences economic activity by affecting aggregate demand. If there are under-utilised resources in the economy, for example during a recession, then increasing the budget deficit - tax cuts, increased spending, or a combination of the two - should put upward pressure on aggregate demand. As aggregate demand expands via the fiscal multiplier, output increases. Expansionary fiscal policy shifts the IS curve to the right such that the IS and LM schedules intersect at a new equilibrium with higher output and interest rates. The higher interest rate level offsets some of the multiplier effects by crowding out private investment.

6.2 The deficit-spending shock

We start by investigating the effects of a deficit spending shock represented by $\varepsilon_t^g = 1$ and $\varepsilon_t^\tau = 0$. The results are depicted in figures 6 and 7. After spending goes up by 1%, output rises by 0.21% on impact. The rises in output, tax, and spending are all still significant ten years after the initial shock. The responses of inflation and interest rates are similar to those after a Haavelmo shock, suggesting that the spending component dominates the tax component in the Haavelmo shock. According to Figure 7, deficit spending stimulates the economy to the extent that it reduces the risk premium and the real short-term interest rate, albeit temporarily. The (temporary) increase in the liquidity premium also points to the spending shock having a stimulating effect on economic activity.

6.3 The balanced-budget government spending shock

In addition to a deficit-spending only or tax only shocks, we are also interested in the effect of a *simultaneous* shock to both fiscal variables. As mentioned above, where this is done such that the deficit remains unchanged, we have a balanced budget shock. The Haavelmo theorem theorem predicts a multiplier of 1. Such an exercise contributes to the on-going public discussion on how to stimulate the economy: changing spending, taxes, or a mixture of the two. The balanced-budget shock requires that $T\nabla\tau = G\nabla g$, where for T and G we use the respective sample means $\bar{T}$ and $\bar{G}$. This reduces to $\nabla\tau = 1.63\nabla g$. Figure 8 is for impulse responses to a balanced budget shock. GDP, tax, and spending all rise significantly over the next 10 years. However, there is no change of economic significance in inflation. There is a temporary increase in the Treasury yield relative to the fed funds rate. This steepening in the yield curve is seen reaction of the liquidity premium in Figure 9 which, as expected, has all changes in long-run relations as temporary. The reduction in the deficit is notable for it takes nearly 15 years to return to its long-run path.

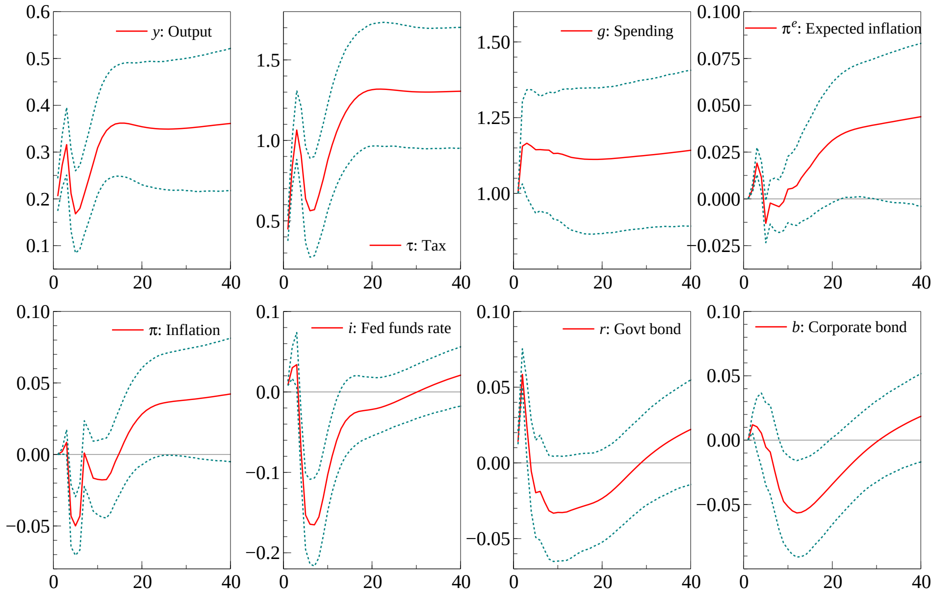

Figure 6 *Level variables response to a* 1% *increase in spending*

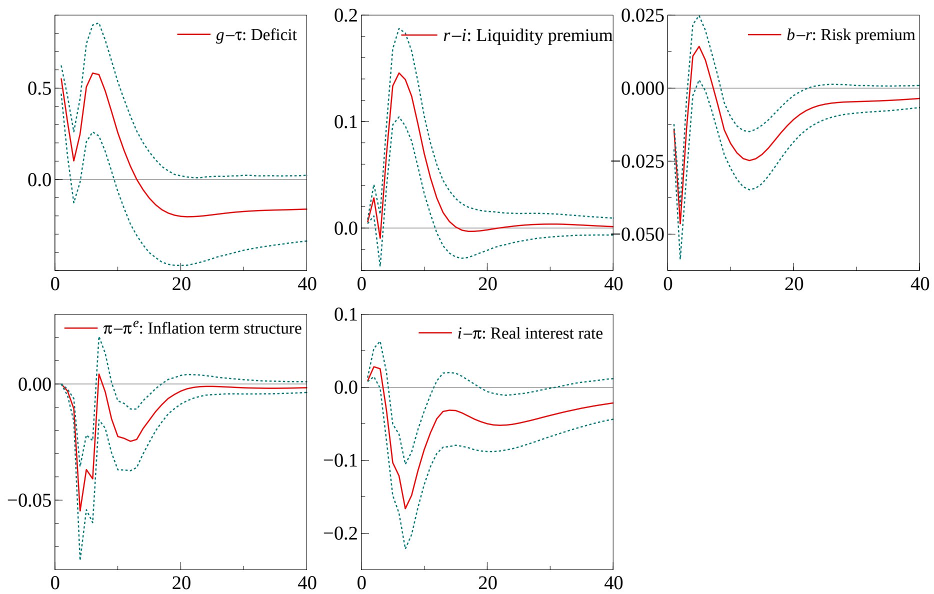

Figure 7 *Long-run relation response to a* 1% *increase in spending*

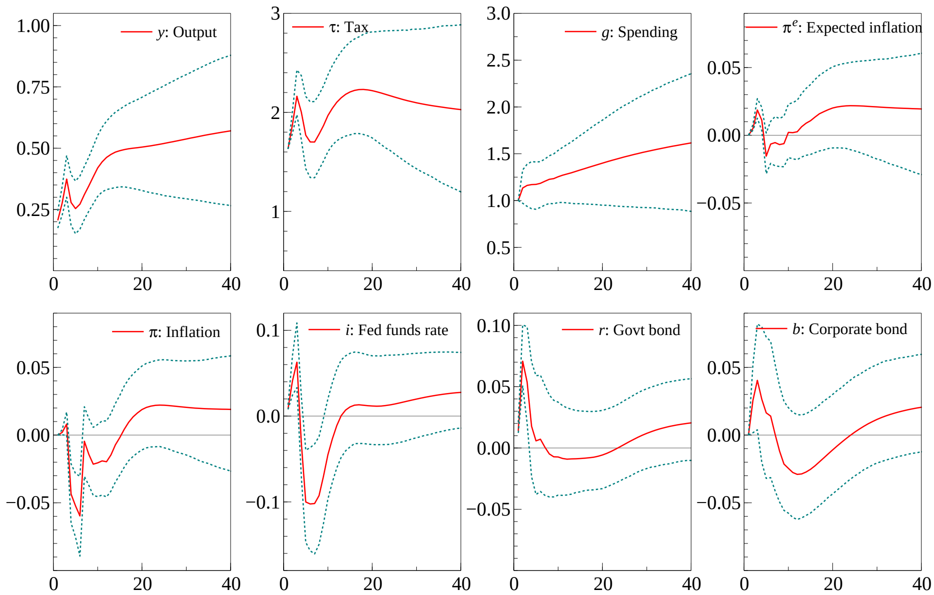

Figure 8 *Level variable response to a tax-financed spending shock*

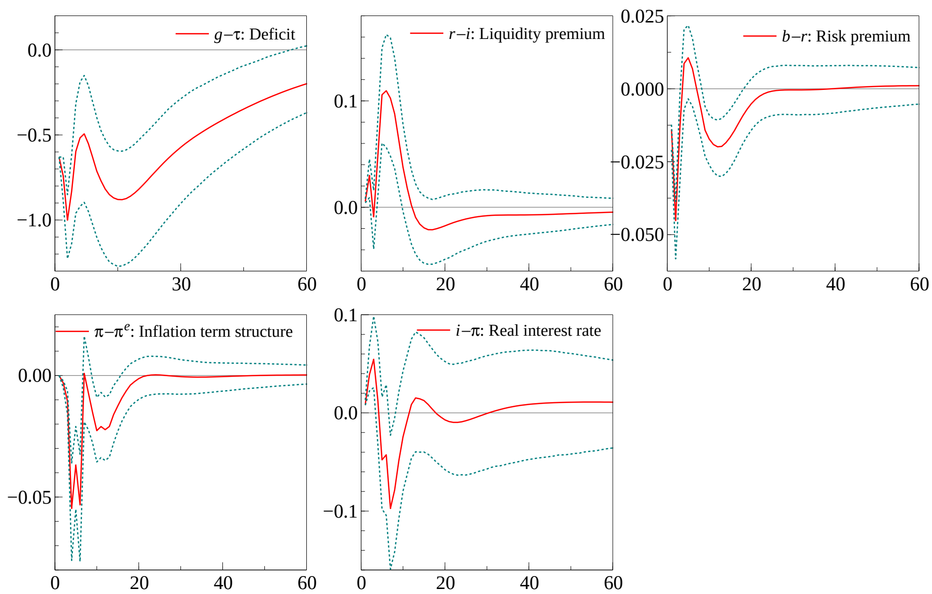

Figure 9 *Equilibrium correction after a tax-financed spending shock*

6.4 The fiscal multiplier

Based on the impulse responses derived in the previous two subsections, we are now in a position to estimate the dynamic multiplier for both the deficit-financed and tax-financed spending shocks. We propose to define the dynamic multiplier, $M_t(h)$, as the ratio of the cumulative change in GDP to the cumulative change in government spending up to h quarters after the shock:

$$M_t^X(h) = \frac{\sum_{i=0}^h \nabla Y_{t+i}}{\sum_{i=0}^h \nabla G_{t+i}} \approx \frac{\sum_{i=0}^h Y_{t+i}\nabla y_{t+i}}{\sum_{i=0}^h G_{t+i}\nabla g_{t+i}} \approx \left(\overline{G/Y}\right)^{-1} \frac{\sum_{i=0}^h \nabla y_{t+i}}{\sum_{i=0}^h \nabla g_{t+i}}, \tag{24}$$

where ∇g_{t+i} and ∇y_{t+i} denote responses of log government spending and log GDP to a fiscal shock i quarters ago. Since $g_t - \tau_t \sim I(0)$, note that $\overline{G/Y}$ is defined here only as the sample mean.

Table 12 *Fiscal multipliers*

	1 qrt	4 qrt	8 qrts	12 qrts	20 qrts	peak (qrts)
Deficit-spending multiplier						
K&S (2014)	1.04**	0.75**	1.24**	1.60**	1.60**	1.64**(14)
B&P (2002)	0.84**	0.45	0.54	1.13**	0.97**	1.29**(15)
Haavelmo multiplier						
K&S (2014)	1.04**	1.09**	1.57**	1.85**	1.81**	1.87**(13)

** significant response at 5% level.

The top panel of Table 12 compares multipliers in this study against those in Blanchard and Perotti (2002) over a five year horizon after a shock to spending. Our multiplier is significant over the 20 quarters, unlike that by Blanchard and Perotti (2002) which fluctuates from being significant in quarter one, insignificant in quarters four to eight, and significant again after quarter twelve. Our multiplier converges to 1.62, so our model estimates that a \$1 increase in government spending results in a \$1.62 increase in GDP. This is in contrast to the 1.29 multiplier found by Blanchard and Perotti (2002). However, the Haavelmo multiplier converges to an even higher level of 1.77. The balanced budget stimulus thus generates a higher increase in output (15 cents more) than a purely spending stimulus. A plot of the two multipliers can be seen in Figure 10.

Both multipliers start at value close to one, meaning that the increase in output is almost entirely due to the increase in government demand. After three to four years the multiplier peaks at values of 1.64 and 1.87, respectively. These are quite close to the long-run multipliers. Interestingly, the Haavelmo multiplier exceeds the deficit-spending multiplier. This is in sharp contrast to the original Keynesian theory, predicting a Haavelmo multiplier of one.

6.5 A monetary policy shock

Joint modelling of fiscal policy and monetary policy captures the effects of the actions of one policy maker on economic variables that are of interest to another policy maker: the effect of fiscal policy on monetary policy, and vice versa. We have indeed seen above that spending and tax shocks affect the federal funds rate in the long-run, according to equation (14). Monetary policy also affects fiscal variables not just in the short-run but also when the liquidity premium relation affects taxes, for example. These effects are augmented by results from impulse response analysis where there are significant, albeit

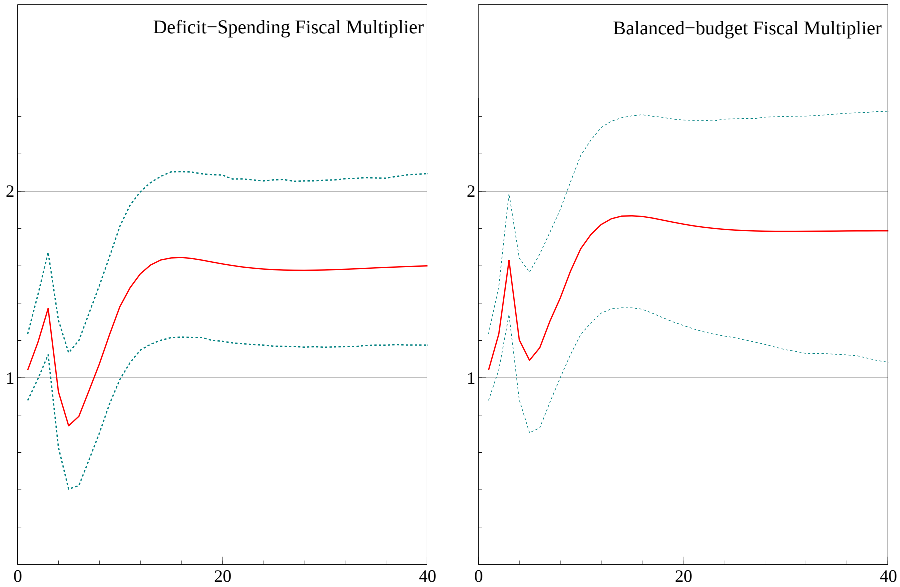

Figure 10 *Haavelmo shock versus deficit-spending shock*

temporary, changes in tax revenues after a monetary policy shock. For other variable responses, inflation falls temporarily and with a lag. The Treasury yield suffers no significant changes but the corporate bond yield rises by 31 bps. A falling liquidity premium and rising real short-term interest rate point to monetary policy's potency in cooling an over-heating economic activity, at least temporarily. Such a contractionary policy is helped by an increase in the risk premium.

The impulse responses in Figures 11 and 12 show a pronounced and long-lived effect of a tightening of monetary policy (in form of a 1% point increase in the federal funds rate) on the fiscal deficit. This suggests that the expansionary stance of the Federal Reserve Board since the global financial crisis has been supportive in controlling the deficit.

7 Conclusion

We have investigated the dynamics of fiscal and monetary policy with respect to their effect on output, inflation, and long-term interest rates. Jointly modelling monetary and fiscal policy allowed us to capture the extent of the interaction between the government and the central bank insofar as one's actions affect the other. This was in addition to the effect of their actions on output, inflation, and long-run interest rates. In particular, we investigate the balanced budget multiplier, the idea that a simultaneous increase in spending and taxes that leaves the deficit unchanged should have a one-to-one effect on output. This is indeed found to be the case based on a balanced budget multiplier of 1.77 that we estimated. We estimated a deficit spending multiplier to be 1.62. We also found fiscal policy directly affects monetary policy via the deficit, with the federal funds rate falling in response to an increase in the deficit. The effects of monetary policy tightening were qualitatively in line with findings in the literature with output and inflation falling, while corporate longer maturity bond yields rise. The temporary increase in the risk premium suggested that the corporate bond yield rises faster and by more than the government bond, evidence that the central bank can alter economic conditions by targeting the risk premium. That

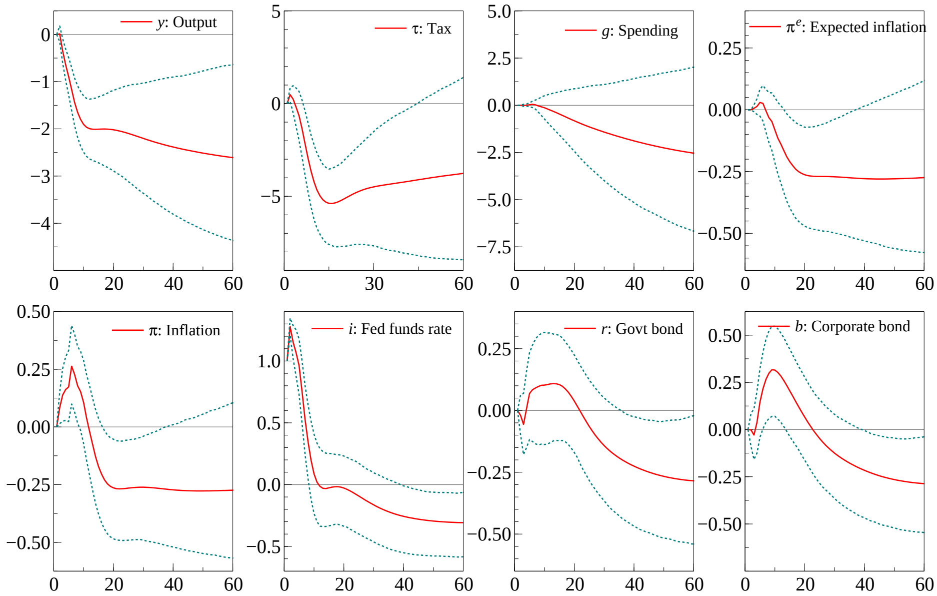

Figure 11 *Level response to a* 1% *point increase in the fed funds rate*

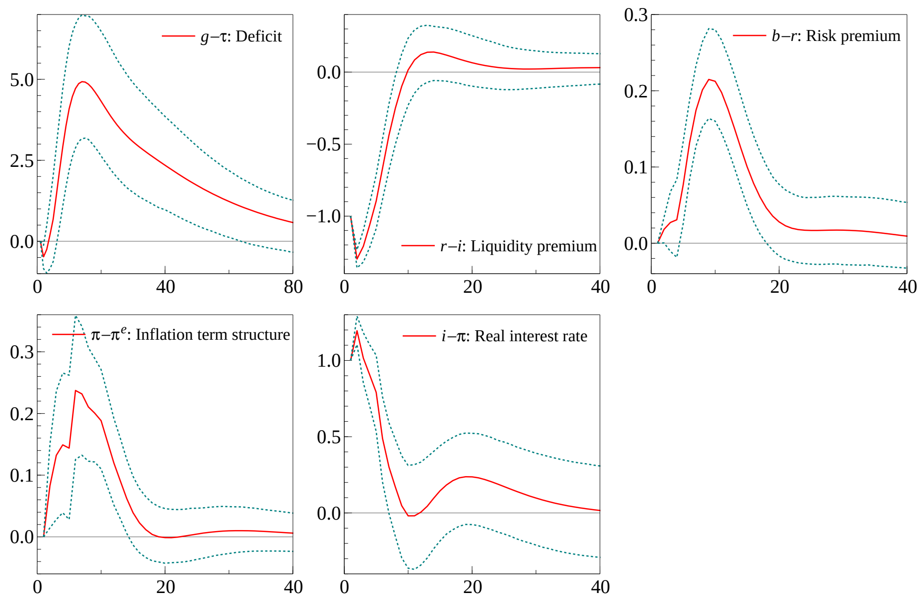

Figure 12 *Long-run relation response to a* 1% *point increase in the fed funds rate*

the Haavelmo shock significantly affects output offers a powerful tool in the use of taxes and spending for economic stabilisation. Such a policy offers a viable compromise between proponents for tax or spending-based stabilisation of the economy.

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