Globalisation, Monetary Policy and the Yield Curve

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Abstract

This paper examines how globalisation has affected the transmission between changes in the policy rate and the term structure of interest rates. We first estimate a cointegrated VAR model using US daily data. We find evidence of a structural break coinciding with the beginning of the "Conundrum" period and relate this to increased foreign official holdings of US treasuries. We then estimate the model using monthly data and find evidence for non-stationary term premia. When we remove the global factors from the yields, we find that the idiosyncratic components of the yields cointegrate. We conclude that the main impact of globalisation is on the term premium.

JEL Classification: E52, C32, C38, F41

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1 Introduction

Between June 2004 and June 2005, the ten-year forward rate implied by US treasury securities fell by 150 basis points, despite an almost 200 basis points increase in the Federal Funds Rate. This apparent disconnect between the behaviour of short- and long-term yields puzzled even the chairman of the Federal Reserve Board, who labelled it a "conundrum".[1] Although acknowledging the global nature of the narrowing of yield and risk spreads, Greenspan was sceptical that the conundrum was due to capital flows associated with greater financial globalisation. He suggested that it was unlikely that increased financial integration was responsible as the disconnect was a recent phenomenon, whereas financial flows had been increasing for some time. Since then, there has been considerable evidence supporting the hypothesis that the conundrum was related to an acceleration in foreign official demand for US treasuries (see Warnock and Warnock (2009), Craine and Martin (2009)). While this episode starkly highlighted the potential effects of financial globalisation on the effectiveness of monetary policy, it proved to be only temporary. However, it is likely that the more permanent effects of increased financial integration on the effectiveness of monetary policy are subtle and gradual and therefore difficult to detect. This paper aims to measure these effects.

There are two main channels through which globalisation can influence the effectiveness of monetary policy (Kamin (2010)**?**). First, as economies become more integrated, the influence of external shocks on domestic macroeconomic and financial variables may increase. Therefore, the set of shocks to which central banks must respond increases and the path of the economy becomes more uncertain. Second, globalisation can alter the traditional mechanisms through which monetary policy affects the economy. For example, the effect of policy changes on the exchange rate may become relatively more important.

The central focus of this paper is the impact of globalisation on the interest rate channel of the monetary transmission mechanism. As financial markets become more integrated, long-term interest rates may increasingly be determined by external factors and less by the stance of domestic monetary policy. Indeed, the effects of financial globalisation, such as those linked

[1] Greenspan, Alan, "Federal Reserve Board's semi-annual Monetary Policy Report to the Congress", before the Committee on Banking, Housing, and Urban Affairs, U.S. Senate, February 16, 2005

to the "conundrum" episode, create additonal difficulties for central banks. When the behaviour of bond yields becomes less correlated with domestic macroeconomic conditions, the appropriate monetary response is more uncertain.

For example, if the decline in yields reflects lower inflation expectations on behalf of global investors, then the appropriate policy response may be to reduce policy rates to prevent *ex ante* real interest rates from increasing. However, if lower long-term bond yields are excessively stimulatory, the central bank may instead need to increase policy rates (Kamin (2010). In the limit, the ineffectiveness of an individual country's monetary policy may be complete and Rogoff (2006) **?**argues that, in this case, countries should coordinate their policies and act collectively. However, Woodford (2007)**?** shows that theoretically this may not be the case. He finds in the context of a two-country new-Keynesian model with complete financial market integration and a given path for real activity in both countries, that domestic inflation will always depend on current and expected future domestic monetary policy. Furthermore, even when currencies are perfect substitutes, in which case it is "global" rather than domestic liquidity that matters for aggregate demand, Woodford (2007) shows that a central bank can still control domestic inflation by varying the interest rate on base money.

The interest rate channel of the traditional monetary transmission mechanism suggests that when central banks change the short-term rate (the monetary instrument), they affect long-term rates (the intermediate target) via the expectations hypothesis. The change in long-term rates then influences the interest-rate sensitive components of aggregate demand, and thus inflation (the final target). This mechansim can be tested empirically in the form of restrictions on a cointegrated VAR. For example, a target variable is "controllable" by the Central Bank if it can be made stationary around a particular target value after adjusting the instrument (Johansen and Juselius (2001)). In econometric terms, if the instrument and target are non-stationary variables and they cointegrate, then the target will adjust to any changes in the instrument in the long-run. Although, the transmission of changes in the policy rate to the final target may "exhibit long and variable lags" and is therefore difficult to test empirically, the effect of policy changes on the yield curve should be immediate and thus more amenable to empirical testing.[2]

[2]Identifying the effect of changes in the policy rate on the final target variables is also

Therefore, the main contribution of this paper is an econometric analysis of how globalisation has affected this relation between policy rates and the yield curve. As the Federal Reserve intervenes daily to equate the Federal Funds Rate with the target set by the FOMC, we first test whether cointegrating relations exist between these variables and short term yields, which have traditionally been the focus of empirical testing of the expectations hypothesis. We then consider the spreads between short-, and medium- and long-term yields, and find that they are stationary once a break is included in the cointegrating relations. As this break coincides with the beginning of the conundrum period, we investigate whether long-term yields and foreign official holdings of US treasuries are cointegrated and find evidence of a long-term relation.

We re-estimate the model over a longer sample period using monthly data and test whether the spreads between short and long-term yields are stationary. The expectations hypothesis is strongly rejected and we find evidence for non-stationary term premia. After decomposing each monthly yield in to factors that are common across countries and components that are unique to the US, we find that the US-specific components of the yields now cointegrate. We conclude that monetary policy only has short-term effects on the short-end of the yield curve and that the Federal Funds Rate is mainly adjusting to shocks to longer-term rates. The main effect of globalisation, aside from the effect of foreign official holdings during the conundrum period, is found to be on the term premia associated with the monthly yields.

The paper is structured as follows: section two discusses the existing empirical literature in this area, section three outlines our theoretical framework, section four presents our econometric model and results for daily and monthly yields, section five presents the results from the factor model, and section six concludes.

2 Literature

One of the earliest applications of cointegration analysis was to test the expectations hypothesis for the short-end of the yield curve, where the coin-

complicated by the frequency of observation of the relevant variables. For example, the Federal Reserve intervenes on a daily basis so that the Federal Funds Rate coincides with the Target set by the FOMC. However, inflation is only measured on a monthly basis and the output gap on a quarterly basis.

tegrating relations were the stationary spreads between pairs of yields (see Engle and Granger (1987)**?**, Campbell and Shiller (1987)**?**). Shea (1992) **?**extended this methodology to include yields on US treasuries with maturities as long as twenty five years. He found that the expectations hypothesis appears to hold for short- and medium-term bonds but not for those at the long end of the yield curve. The latter appeared to be influenced by additional common trends, which he suggested were related to liquidity premia. Johansen and Juselius (2001) **?**implicitly test the expectations hypothesis for short-term yields using a cointegrated VAR and find evidence that the spread between yields is stationary.

Johansen and Juselius (2001) also integrate the expectations hypothesis into the monetary transmission mechanism and specify how a monetary control rule can be tested as a restriction on a VAR process. While their primary focus is on whether the Federal Reserve can "control" its ultimate target, the inflation rate, they also examine whether the monetary instrument (the federal funds rate target) cointegrates with an intermediate target (the three- and six month bond yields). They argue that since these yields are influenced by a stochastic trend in shocks to the yields themselves, rather than shocks to the federal funds target, the Federal Reserve cannot "control" its intermediate target. However, they do find that the federal funds rate target has significant short-run effects on the yields. We adopt the Johansen and Juselius (2001) integrated framework for the relation between the monetary instrument and the term-structure, but we focus more on describing the short- and long-term dynamics of the yield curve, rather than on the "controllability" of short-term interest rates with which they are concerned.

More recently, Giese (2008) **?** examines the relation between the short-term and long-term yields on US treasuries using a cointegrated VAR and finds that, although spreads between yields are nonstationary, *differences* between spreads are stationary. In her framework, this implied two stochastic trends relating the level and the slope of the yield curve.[3] Parts of our approach are similar to Giese (2008) in that we use a cointegrated VAR to capture the term structure of interest rates. However, our focus is not primarily on modelling the time-series properties of the level, slope and curvature of the yield curve, but instead on the relation between changes in short-rates,

[3]Litterman and Scheinkman (1991) **?** found that the term structure of interest rates can be described by three factors, interpretated as the "level", "slope" and "curvature" of the yield curve.

particularly the policy rate, and long-term rates. In addition, Giese (2008) does not consider the international dimension of yields, which is of particular interest to us.

Some studies have investigated whether the monetary transmission mechanism has changed in general, while others have specifically focused on the contribution of globalisation. The latter has typically involved examining whether the impact of shocks to the monetary instrument on macroeconomic variables has weakened as the influence of global factors on these variables has increased. Using intra-daily data for the US, Faust et al (2007)**?** find that the response to surprise policy announcements by the FOMC across the term structure did not change over the period 1987 to 2002. Boivin and Giannoni (2008) **?** estimate a Factor Augmented VAR (FAVAR) using US data for the period 1984 to 2005 and find that international factors have become more important for domestic macroeconomic variables especially the long-term interest rate, but that the transmission of monetary policy has not changed.[4]

Boivin et al (2010) **?** again use a FAVAR to determine whether the US monetary transmission mechanism changed between the pre-1979 and post-1984 periods. They find that inflation, particularly expected inflation, and real output responded less to monetary policy shocks in the latter period and they suggest that this may be due to more reactive central banks, as measured by the coefficients in a policy reaction function. Thus, one way in which long-term interest rates may become less responsive to changes in short-term rates, is via the expectations channel. With a constant real interest rate, long-term interest rates will be relatively less responsive if inflation expectations are well anchored. However, it may be globalisation that is driving this decline in expected inflation (see Ciccarelli and Mojon (2010).

A recent set of contributions to the literature has investigated whether the yield curve has global factors. However, these studies do not relate these factors to the monetary transmission mechanism. Diebold, Li and Yue (2008) **?**estimate global factors for four major countries based on monthly data for the period 1985-2002 and find a dominant global level factor (reflecting global inflation) and an important slope factor (reflecting the global business cycle). They find that the importance of global factors increases with the maturity

[4]A FAVAR uses a large dataset of macroeconomic indicators related to real activity, prices, interest rates , stock prices and money and credit aggregates, to extract common factors across countries.

of bonds but, of all the countries in the sample, these factors are least important for the US. Kaminska, Meldrum and Smith (2011) **?** use a 3 country affine term structure model to decompose forward rates into expectations of the short-rate, term-premia and a convexity effect. They find common level and slope factors for the US, UK and euro area over the period 1992-2008 and relate these factors to global inflation and global economic activity, respectively. Local factors are also needed to explain the behaviour of yields, with monetary policy being the most important local factor. While the focus of Kaminska et al (2011) is on modelling the commonalities across yields, we concentrate instead on the country-specific relation between yields of different maturities, once these commonalites have been removed. Therefore, although we also extract common factors, we instead concentrate on the idiosyncratic component.

Several authors have focused on a recent particular episode, the "conundrum", in which globalisation may have affected US long-term interest rates, but not specifically on how it has affected the relation between short- and long-term rates. Warnock and Warnock (2009) **?**use monthly Treasury International Capital System (TIC) data to examine the effect of foreign official purchases of US treasury and agency securites on short- and long-term Treasuty yields. They find that these holdings had no effect on short-term yields but reduced the ten-year yields by up to 90 basis points in 2005. Craine and Martin (2009) **?**use weekly data on foreign official holdings from the Federal Reserve Boards H4.1 release to examine whether this, or other factors such as an increase in the supply of these bonds, and surprise macroeconomic and monetary announcements, caused the fall in long-term yields. They find that the ten-year forward rate was at least 50 basis points lower in 2005 due to increased foreign holdings of US treasuries.

However, others have offered a different explanation for the behaviour of lower long-term interest rates during the conundrum period. Kim and Wright (2006)**?** estimate a three factor affine model of the term structure of US yields and find that the term-premium fell significantly during the period 2004-2005. Similarly, Backus and Wright (2007)**?** argue that the conundrum reflected a fall in the term-premium, although they do not explain which factors may be driving this. They suggest, however, that it could not primarily be due to increased official demand for treasuries, as similar declines in forward rates were observed in other countries. Gürkaynak and Wright (2010) **?**estimate a term structure model with latent factors between 1971 and 2009 and find that term premia were lowest in the period 2004 to 2005. Finally, Smith and Taylor (2009)**?** suggest that the conundrum was due to the perception among market participants that the coefficient on inflation in the Federal Reserve's policy rule had fallen and therefore, that future short-term interest rates would be lower.

Bernanke, Reinhart and Sack (2004) **?**point to another episode when external factors significantly influenced long-term US interest rates. In a term-structure model with inflation and the output gap as factors, they find that treasury yields declined significantly during periods of substantial intervention by the Bank of Japan in the foreign exchange market between 2000 and 2004. The savings glut hypothesis is a closely related explanation as to how globalisation may have affected long-term yields. For example, Byrne et al (2010) **?**show that the correlation between long-term interest rates across countries is actually greater than the correlation between short- and long-term interest rates within countries and that this correlation has been increasing over time. Using data for eight countries from 1988 to 2006, they find that the first principal component of long term interest rates cointegrates with a measure of global foreign exchange reserves, which they suggest highlights the importance of the "global savings glut" in explaining lower long-term interest rates.

The conundrum episode is important in our model of daily yields, as it constitutes a structural break in the relation between short- and long-term rates. However, in testing for a cointegrating relation between long-term rates and foreign official demand for US treasuries, we also contribute to the debate on its origins. We now briefly outline the theoretical framework underpinning the yield curve and how it relates to our econometric model.

3 The Expectations Hypothesis and Monetary Policy

The "strong" form of the expectations hypothesis asserts that the term premium is zero and therefore that yields simply reflect expected short-term rates in the future. Given that the yield curve usually has a positive slope, this would imply ever-increasing future short-term rates. An empirically more reasonable assumption is to allow for a term premium but to assume that it is constant. This is the "weak" form of the EH and implies that any *changes* in yields must be related to *changes* in future expected short-term rates and the expected path of monetary policy.

Let $P_t(n)$ be the price at time t of a zero-coupon bond with n-years to maturity. The annual yield on this bond between t and $t+n$ can be written as:

$$y_t(n) = \frac{1}{n}\ln(P_t(n))$$

Yields on a n-period zero-coupon bond can also be written as the average of forward rates over the duration of the bond:

$$y_t(n) = \frac{1}{n}\sum_{i=1}^{n} f_t^i$$

The strong form of the EH assumes that investors are risk neutral and thus the risk premium that they require to hold long maturity bonds is zero. In this case, the i-period forward rate at time t is the expected short-term interest rate r at time $t+i-1$:

$$f_t^i = E_t r_{t+i-1}$$

The weak form of the EH suggests that investors demand a risk or term premium as compensation for holding long-term bonds.[5] Forward rates therefore include a term premium, b, and this premium may be maturity specific but is assumed to be constant:

$$f_t^i = E_t r_{t+i-1} + b_t^i$$

Therefore, according to the weak form of the EH, yields on zero-coupon bonds can be expressed as the average of expected future short-term rates plus a term premium:

$$y_t(n) = \frac{1}{n}\sum_{i=1}^{n}(E_t r_{t+i-1} + b_t^i)$$

An even weaker form of the EH, which we test in this paper, allows the term premium to be time-varying but assumes that it is stationary. Empirical

[5]For example, this risk may take the form of higher than expected inflation or uncertainty about the path of future policy rates, and thus the resale value of the bond if it is sold prior to maturity.

testing of the EH using Vector Error Correction (VECM) or Cointegrated VAR (CVAR) models focuses on examining whether the spreads between short-term rates and longer-term yields are stationary.

The spread between the yield on an n-period bond and the risk-free rate is given by the weighted average of future changes in the short-rate:

$$
\begin{aligned}
y_t(n) - y_t(1) &= [\frac{1}{n}\sum_{i=1}^{n}(E_t r_{t+i-1})] - r_t + \sum_{i=1}^{n} b_t^i \\
&= (\frac{1}{n} - 1)r_t + \frac{1}{n}\sum_{i=1}^{n=1} E_t r_{t+i} + \frac{1}{n}\sum_{i=1}^{n-1} b_t^i \\
&= \frac{1}{n}\sum_{i=1}^{n-1}(n-i)E_t(\Delta r_{t+i}) + \frac{1}{n}\sum_{i=1}^{n-1} b_t^i
\end{aligned}
$$

As bond yields are generally integrated of order one, then their first-differences, Δr_{t+i}, will be stationary. If the spread $y_t(n) - y_t(1)$ is non-stationary, then the term premium must be non-stationary. However, if pairs of spreads are stationary then the weighted differences between each yield's term premium must also be stationary.[6] For any constant weight c:

$$
\begin{aligned}
[y_t(n) - y_t(m)] - c[y_t(m) - y_t(1)] = \frac{1}{n}\sum_{i=1}^{n-1}(n-i)E_t(\Delta r_{t+i}) - \frac{1+c}{m}\sum_{i=1}^{m-1}(m-i) \\
E_t(\Delta r_{t+i}) + \frac{1}{n}\sum_{i=1}^{n-1} b_t^i - \frac{1+c}{m}\sum_{i=1}^{m-1} b_t^i
\end{aligned}
$$

The weighted difference between spreads expresses the curvature of the yield curve when the yields are not equally spaced along the yield curve, as shown by Giese (2008). This has important implications for empirical testing of the EH. Unless the econometric model allows for cointegrating term premia, it may reject any long-term relation between short-. medium- and longer-term yields.

The expectations hypothesis therefore, provides the theoretical framework for investigating the transmission of changes in the monetary policy

[6]Intuitively, if the term premia coinegrate, this suggests that investors' *relative* preferences over bonds with different maturities are constant.

instrument, for example the Federal Funds Rate or its target, to the rest of the term-structure. This constitutes the first part of the monetary transmission mechanism, where the central bank seeks to "control" its intermediate target through changes in its instrument.

4 A Cointegrated VAR model

The focus of this paper is on how globalisation has affected the relation between short- and long-term rates. Cointegration analysis is particularly useful in the latter context as it allows us to determine whether there is a long-term relation between the policy rate and the term-structure of interest rates. In our daily model we focus on the period around the conundrum episode and examine whether we can detect a change in the nature of this relation and, if so, if it can be related to foreign official demand for US treasuries. We then examine this relation over the period between 1990 and 2007 using monthly data to allow for feedback from the yields to the policy rate. We leave it to the following section to investigate fully how globalisation has affected the monthly yield curve.

4.1 A Daily Model

The Federal Reserve intervenes daily in order to match the prevailing rate in the market for federal funds with the target for the federal funds rate set by the FOMC. Therefore, we first examine whether the federal funds rate closely follows its target and whether the target strongly influences other short-term rates. The presence of a cointegrating relation between the policy and short rates implies that the EH holds for at least the short-end of the yield curve. We then extend the analysis to include medium- and longer-term yields and test whether the spreads between these yields and those at the short-end of the yield curve are stationary, as the EH would imply. I

Daily data for the federal funds rate and its target are from the FRED database of the Federal Reserve Bank of St. Louis, while the yields for three- and six-month, and one-, five- and ten-year zero-coupon US treasuries are taken from the updated dataset of Gurkaynak, Sack and Wright (2007). The daily models are estimated for the period 1 April 2003 to 31 March 2007, the sample being truncated at this latter date due to the significant increase in the volatility of financial markets at the onset of the financial crisis.[7]This provides a sample of 1044 observations. Figure 2 plots these variables over our sample period. From the top panel in Figure 2, it is clear that, not only has the federal funds rate (FFR) tracked its target very closely, but that the three- and six-month yields have also reacted strongly to changes in the policy rate. The lower panel in Figure 2 shows that co-movement with the federal funds rate target is decreasing with the maturity of the bonds, with the five- and ten-year yields appearing invariant to changes in shorter-term rates.

4.1.1 Short-term Yields

To examine the relation between the FFR and target, and the short-end of the yield curve, we estimate a cointegrated VAR with the following form:

$$\Delta x_t = \alpha(\beta', \beta_0)\begin{pmatrix} x_{t-1} \\ 1 \end{pmatrix} + \underset{i=1}{\overset{k-1}{\Gamma_i}} \Delta x_{t-i} + A_0 z_t + \phi D_t + \varepsilon_t \tag{1}$$

where $\varepsilon_t \tilde{\ } iidN(0,\Omega)$, $x_t' = [ffr, G3m, G6m]$ is a vector of endogenous variables, $z_t = [Trg_t]$ is the Federal Funds Rate Target and, as it is set periodically by the FOMC, it is treated as exogenous. The constant, β_0, is restricted to the cointegrating relations. D_t is a vector of dummy variables that control for particularly large shocks affecting the interest rates during the sample period.

Likelihood ratio tests suggested that two lags were sufficient to capture the persistence of short-run effects. The results of the misspecification tests on the VAR(2) model are presented in Table 1. The multivariate test of no first-order serial correlation is not rejected at the five percent level, while the null hypothesis of no ARCH effects is only rejected for the Federal Funds Rate. Normality (Jarque-Bera test) is rejected for the three endogenous variables, although this is due to excess kurtosis rather than skewness (not reported). As VAR parameter estimates have been shown to be robust to ARCH effects and kurtosis, these rejections should not significantly affect our analysis (see Gonzalo and Ng (2001) **?** and Juselius (2006)**?** for a discussion). We also test whether any of any of the endogenous variables may actually be weakly exogenous and this is clearly rejected.[8]

[7]We found a significant increase in ARCH effects when later data were included.

[8]A weakly exogenous variable is affected only by its own shocks in the long run and not by those to other variables. This imlies a zero-row in the matrix of adjustment coefficients.

To determine the rank (r) of the $\Pi = \alpha\beta$ matrix, or the number of cointegrating relations, we consider both the Johansen Trace test and the characteristic roots of the process. Table 1 shows that the trace test suggests $r = 2$ and this choice of rank leaves 0.877 as the largest root in the model. As this is consistent with a stationary process that exhibits slow mean-reversion, it appears to be a reasonable choice of rank. Our finding of two cointegrating relations and one stochastic trend among the three endogenous variables is supportive of the Expectations Hypothesis, which implies that the spreads between non-stationary yields (here, the three- and six-month yields) and the yield with the shortest maturity (here, the FFR) should be stationary.

Table 2 presents the overidentified cointegrating relations and adjustment coefficients, as well as the short-run parameters from the Γ_1 matrix. This restricted model was accepted with a $\chi^2(2)$ statistic of 0.4 and a p-value of 0.72.[9] Further restrictions of exact and non-exact spreads between the FFR and the Target in the first cointegrating relation, $\widehat{\beta}_1$, were rejected. The first cointegrating relation is between the FFR, the Target and the three-month yield, and the FFR is strongly adjusting to this relation. This relation is evident in the top panel of Figure 1 where all three rates display significant long-run (and short-run) comovement. It suggests that the FOMC reacts to short-term rates by targeting the FFR. Although the three-month and six-month yields are not adjusting to this relation, the latter is only marginally insignificant. The second relation, is the exact spread between the three- and six-month yields and thus provides evidence for the EH for the short-end of the yield curve. Both of these yields are strongly adjusting to this relation, implying that any deviation from the spread is quickly corrected by the market. A recursive test of the constancy of the log-likelihood and the cointegrating relations did not reveal any parameter instability over the sample period.[10]

Table 2 also presents information about the short-run effects of the en-

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[9] The restrictions on the CVAR can be formulated by specifying the number of free parameters, s_i in each β vector: $\beta = (H_1\varphi_1, ..., H_r\varphi_r)$ where β is $(p_1 \times r)$, p_1 is the dimension of the CVAR, φ_i are $(s_i \times 1)$ coefficient matrices and H_i are $(p_1 \times s_i)$ design matrices which are used to determine the free parameters in each cointegrating vector. These restrictions are then tested using a likelihood-ratio test (see Juselius (2006)).

[10] The recursive test of the log likelihood estimates the model over a baseline period at the beginning of the sample and then recursively tests whether subsequent observations follow the same model. Under the null hypothesis of constant parameters, the 5% confidence value of the test is 1.36. (see Juselius (2006), chapter 9).

dogenous and exogenous variables. The most important finding given the focus of this paper is that the contemporaneous and lagged changes in the Target have significant effects on nearly all the endogenous variables, as would be expected from changes in monetary policy. Although, changes in the Target by the FOMC have clear short run effects on short-term yields, Johansen and Juselius (2001) argue that for these yields to serve as intermediate targets in the monetary transmission they must also be affected by shocks to the Target in the long run. To see which shocks are driving these yields in the long-term, we can look at the C-matrix shown in Table 3 which is derived from the Moving-Average (MA) representation of the VAR:

$$x_t = C \sum_{i=1}^{t} \varepsilon_i + C^*(L)\varepsilon_t + \widetilde{X}_0 \tag{2}$$

where $C = \widetilde{\beta}_\perp \alpha'_\perp$, $\widetilde{\beta}_\perp = \beta_\perp(\alpha'_\perp \Gamma \beta_\perp)^{-1}$. C^* is a stationary process and .$\widetilde{X}_0$ are the initial conditions. The model contains one stochastic trend, which as $\alpha_\perp$ parameters show, mainly comprises shocks to the three- and six-month yields.The loading parameters show that the FFR is not stongly affected by the the cumulated shocks to the yields. As the yields are not affected by the shocks to the FFR in the long-run, this would imply that these yields cannot serve as intermediate targets according to the methodology of Johansen and Juselius (2001). We have shown that changes in the monetary instrument (the Target) have strong short-run effects on both the FFR and the yields but that, in the long-run, these yields are affected by shocks not related to the conduct of monetary policy.

4.1.2 The Daily Term Structure

The previous section showed that the EH holds in general for short-term yields and therefore that there was a close relation between changes in the policy rates and short-term yields during the conundrum period. As we are interested in examining the transmission of monetary shock across the yield curve, we now extend our analysis to include medium- and long-term yields. In order to limit the dimension of the CVAR and having shown that the homogenous spread between the three- and six-month yields is stationary, we exclude the latter from this model. We re-estimate the model in [1] with a vector of endogenous variables that now includes the FFR, the three-month yield, and the one-, five- and ten-year zero-coupon yields: x_t'=[ffr,$G3m$,$G1y$,$G5y$,$G10y$]. Likelihood tests for the appropriate lag-length in this model suggest that three lags are now needed to account for short-run persistence after including the longer-term yields. Table 4 presents the misspecification tests for this model. Multivariate normality is again rejected, with the univariate tests suggesting that only the ten-year yield is normally distributed. We expect that our parameter estimates are robust to this rejection, as it is again due to excess kurtosis rather than skewness. As in the previous analysis, the estimates should also be robust to the moderate ARCH effects in the FFR. We also find that none of the variables can be considered weakly exogenous.

Table 4 also shows the results of the trace test and the characteristic roots of the process. The trace test suggests a rank of $r = 4$, which implies one unit root process. This leaves a large root of 0.931 in the model. As this may be consistent with a slow mean-reverting process or with a unit root process that cannot be detected due to the low power of the trace test, we also look at the unrestricted adjustment coefficients presented in Table 5. The fourth column of this table shows that one-, five- and ten-year yields are strongly adjusting to a fourth cointegrating relation. Therefore, this evidence suggests $r = 4$.

The over-identified cointegrating relations, adjustment coefficients and short-run parameters are shown in Table 6. This model was accepted with a $\chi^2(9)$ statistic of 15.132 and a p-value of 0.1. Recusive tests on the constancy of the likelihood suggested that the parameters in the model are not constant. Figure 2 illustrates that a break seems to have occurred in the cointegration relations in the middle of 2004, when the dashed R1(t) line representing the parameters of the cointegrating relations exceed the five percent confidence value 1.36. This suggests that a level shift occurred in the relation between yields around this time.

As no such instability was evident in the model with short-term yields, this instability must be related to longer-term yields. This suggests that the model is capturing the Greenspan "conundrum" when the FFR and short-term yields became disconnected from long-term yields. Furthermore, the timing of the break coincides with a period when foreign central banks and finance ministries were rapidly increasing their holdings of mainly long-term US government securites, as documented in Craine and Martin (2009) and Warnock and Warnock (2007). This demand increased the price of long-term US treasuries and resulted in a fall in long-term yields, even though short-term rates were rising. We return to the possible connection between long-term rates and official holdings below. In the current framework, we control for this change in the behaviour of long- relative to short-rates by including a structural break in the model in mid-July 2004.[11]

The first cointegrating relation is the non-exact spread between the three-month yield and the ten-year yield. The break dummy is significant, again indicating that there was a change in the relation between short- and long-rate during this period and that a break or level shift is needed to make this relation stationary. The FFR, and the five- and ten year rates are all significantly adjusting to this spread. The second cointegrating relation is the same relation between the Target, the FFR and the three-month yield as in the model with short-term yields and, as before, only the FFR is adjusting.

The final two relations are the exact spreads between the three-month yield and one- and five year yields, respectively. It is interesting that the break dummy is significant in both relations and it implies that the yields on treasuries with maturites as short as one year also became disconnected from those at the short-end of the yield curve. Adjustment to the one-, five and ten-year spreads is almost identical. The main difference is that the three-month rate appears to only adjust to the one-year spread.

Table 6 also presents the short-run effects of changes in each of the exogenous and endogenous variables. It is clear that changes in the Target by the FOMC have no contemporaneous or lagged effects on treasuries with maturities of one year or more over our sample period. Further, the short-run behaviour of these yields is mainly driven by lagged own-changes and the lagged changes of the other yields, although lagged changes in the three-month yield do have a significant effect on current changes in the one-year yield.

To determine what drives these yields in the long-term we can look at the C-matrix presented in Table 7.With five endogenous variables and $r = 4$, there is one stochastic trend and this mainly comprises shocks to the five- and ten-year yields. The loading coefficients, $\widetilde{\beta}_{\perp}$, show that the FFR does not load strongly on this trend, and therefore we can conclude that it is mainly affected by changes in the target, as shown in Table 6. The other variables load on the trend according to their coefficients from the cointegrating relations. As before, shocks to the FFR have no long-run effects on the yields.

[11] The model is not sensitive to our choice for the specific date of the break. The results are similar when a break is placed at any date between June and August 2004. We chose the middle of July as it is the mid-point of this range and is close to date that the Federal Reserve began its tightening cycle.

4.1.3 Foreign Official Treasury Demand and the Conundrum

We now return to the possible connection between the break in the transmission mechanism and official purchases of long-term treasuries. The cointegrated VAR is an ideal framework to investigate whether a long-term relation exists between long-term yields and official holdings of long-term treasuries. The H4.1 release from the Federal Reserve provides the weekly total of offical foreign holdings of US treasuries at the Federal Reserve Bank of New York and these official holdings, scaled by total US government debt, are plotted in Figure 3.[12] It is clear that official demand for US treasuries began to increase rapidly towards the end of 2003 and beginning of 2004. We therefore test whether a cointegrating relation exists between the five- and ten-year yield and the holding of these bonds by foreign central banks and finance ministries.[13]

Table 8 presents the results of the model with these three endogenous variables, estimated using weekly data covering the period 5 January 2000 to 28 March 2007. The trace test and the unrestricted adjustment coefficients suggested a rank $r = 2$ and thus two cointegrating relations. The over-identified CVAR(2) was accepted with a $\chi^2(2)$ statistic of 4.29 and p-value of 0.117. The first relation is the exact spread between the five- and ten-year yields and is stationary around a constant. Foreign official holdings of US treasuries tend to increase when the yields deviate from this relation. The second relation is between the ten-year rate and official holdings with both of the yields adjusting to this relation. Therefore, both the ten-year rate and foreign official demand for US Treasuries move together in the long-run. The short-run parameters indicate that the lagged change in the five year yields influences the change in the ten-year yield, while the lagged change in the latter has a significant effect on foreign official holdings. The long-run impact

[12] http://www.federalreserve.gov/releases/h41/

[13] The H4.1 release provides the most timely (weekly) data on foreign official holdings of US securites but includes only those holdings held at the Federal Reserve Bank of New York. A more complete estimate of foreign official holdings at the monthly level is available from Bertruat and Tyron (2007)**?** who combine annual TIC survey data with monthly transactions data. However, using monthly data would significantly reduce the number of observations over the sample period and thus, the power of the trace test. We do identify similar cointegrating relations to those in the weekly model when we use monthly data, but given the lower power of the trace test, this model is unlikely to be well-specified. In any case, the correlation between between the H4.1 series and that estimated by Bertraut and Tyron (2007) is 0.85 over our sample period.

matrix indicates that only shocks to foreign official demand have permanent effects on the yields.

Our analysis of the daily data has provided some evidence that yields on treasuries of different maturities are cointegrated, as the expectations hypothesis would predict. We also find a break in the transmission of changes in short-rates to longer-term rates. We show that this can be explained by the close relation between long-term rates and foreign official holdings of US securities. This approach has treated changes in policy rates as exogenous, which is a valid assumption with daily data. However, to allow feedback from longer rates to policy changes, we now look at the transmission between short- and long-term yields using monthly data. In the following section, we then examine whether the nature of the transmission mechanism changes, once the global factors that influence monthly yields have been removed.

4.2 A Monthly model

In the daily model, changes in the policy rate were treated as exogenous and therefore no feedback was permitted from long-term rates. However, at the monthly level Central Banks may react to changes in economic conditions which are reflected in these rates. Further, while the FFR and Target may deviate on a daily basis, they are indistinguishable at monthly frequencies. We therefore re-examine the relation between short- and long-term yields using monthly zero-coupon yield data from Wright (2011) and we use the FFR to reflect changes in the policy rate. The variables are the monthly counterparts of the yields used in the daily model. The sample covers the period from January 1990 to March 2007, which provides 207 observations. Figure 4 plots the FFR, the three- and six-month zero-coupon yields, and the one-, five- and ten-year yields on US treasuries over this period. It is clear that there is strong comovement across yields of all maturities apart from a short-period between 2004 and 2005.

4.2.1 A CVAR Model of the Monthly Term Structure

The estimating equation for the monthly model is similar to (1) except that the exogenous Federal Funds Target, z_t, is now excluded. The vector of endogenous variables, x_t=[ffr_t,$G3m_t$,$G6m_t$,$G1y_t$,$G5y$,$G10y$], contains the six variables plotted in Figure 4. Likelihood ratio tests found that two lags were necessary to capture the short-run persistence. Table 9 presents the results of the misspecification tests and tests for the weak exogeneity of the variables. ARCH effects are not significant in the monthly model, while normality is only rejected for the FFR and short-term yields. As before, this is due to excess kurtosis rather than skewness. First- and second-order autocorrelation is also rejected. We also find that the ten-year yield can be considered to weakly exogenous, while the one-year rate is marginally so. A test of joint exogeneity of both the one- and ten-year yields was strongly rejected and so only the latter is treated as weakly-exogenous.

Table 9 also reports the results of the trace test and the characteristic roots for the rank suggested by that test. The trace test suggests four cointegrating relations which implies two unit root processes. This choice of rank leaves 0.877 as the largest stationary root in the model. For confirmation that this choice is reasonable, we can also look at the unrestricted adjustment coefficients presented in Table 10. It is clear that there is significant adjustment to four cointegrating vectors and thus $r = 4$ appears to be a suitable description of the number of stationary relations in the model.

The four over-identified cointegrating relations, adjustment coefficients and short-run parameters are shown in Table 11. This model was accepted with a $\chi^2(10)$ statistic of 13.0 and a p-value of 0.223. The first relation is the non-exact spread between the FFR and the six-month yield, which is stationary around the constant. Only the FFR adjusts to this relation. The second relation is the spread between the FFR and the three-month yield. It is interesting that the latter is the adjusting variable, given that the former adjusts in the first relation. The third cointegrating relation represents the stationary curvature of the yield-curve at medium- to long-maturites and the three-month through five-year yields are all significantly adjusting to this relation.[14]The final cointegrating relation is the weighted spread between short-and long-term yields.[15] Adjustment to the third and fourth relation is very similar, although the one-year yield is marginally insignificant in the latter.

Table 11 also presents estimates of the short-run parameters in the model. The FFR is mainly influenced by lagged changes in long-term yields, the three-month yield and its own lag in the short-run. The three-month yield is mainly affected by the lagged change in the FFR and one-year yields,

[14]The third cointegration relation has the form: $\beta_3' x_t = 0.57(G5y_t - G1y_t) - 0.43(G10y_t - G5y_t)$.

[15]The fourth cointegrating relation can be expressed as: $\beta_4' x_t = 0.78(G5y_t - ffr) - 0.22(G10y_t - G5y_t) - 0.07$

while short-run persistence in long-term yields is mainly a results of lagged changes in these yields. Interestingly, given the focus of this paper, the FFR has no short-run effect on any yield in our dataset, apart from the yield on the three-month bond.

Tests for the constancy of the log-likelihood did not reveal any parameter instability. This is surprising given our finding of a structural break in the daily model in 2004. However, the conundrum period was a temporary episode and is likely to be less significant at the monthly level in the context of a sample that begins in 1990. We therefore assume that the parameters in the model have remained stable over the sample period.

4.2.2 Common Trends and the Structural MA Model

We now turn to the stochastic trends in the model. Table 12 shows that the first common trend comprises shocks to the weakly-exogenous ten-year rate, while the second common trend is mainly driven by shocks to the six-month and one-year yields and, to a lesser extent, the five-year yield. The loadings to the first common trend are in a narrow range between 0.69 and 1.27 and suggest that each yield is affected similarly by shocks to the ten-year yield. Therefore, the first common trend is likely to be a level factor. The loadings to the second common trend are similar up to one year and then decline. This may indicate a slope factor with positive shocks to the six-month and one and five-year yields increasing the slope and resulting in a large negative effect on short-yields (see Giese (2008)).

As the residuals are correlated in the MA representation, we now attempt to separate the residuals, ε_t, into permanent and transitory shocks, u_t. We therefore form the $\widetilde{C} = CB^{-1}$ matrix from a rotation of the C-matrix where only the $p - r = 2$ linear combinations of the structural shocks have permanent effects and r shocks have only transitory effects. This gives the "structural MA" representation:

$$x_t = CB^{-1} \sum_{i=1}^{t} u_i + C^*(L)B^{-1}u_t + \widetilde{X}_0 \tag{3}$$

where $\varepsilon_t = B^{-1}u_t$ and $\widetilde{X}_0$ are intial conditions. The B matrix shows which linear combinations of the shocks generate the transitory and permanent effects. As there are two permanent shocks in the model, it is necessary to impose a zero-restriction on one of the shocks in order to identify the $\widetilde{C}$ matrix. As the ten-year rate was found to be weakly exogenous this suggests that cumulated shocks to the ten-year rate are not correlated with the shocks to the other yields. Therefore, we place the zero-restriction on the ten-year yield in the second permanent shock.

The (normalised) structural impact matrix is presented in Table 13. The first $r = 4$ columns indicate the long run effects of transitory shocks, while the last $p - r = 2$ columns indicate the long-run effects of the permanent shocks. The coefficients indicate how each variable loads on the permanent shock. The loadings on the first permanent shock are in a relatively narrow range between 0.617 and 1 and therefore this shock may represent a level factor. The loadings on the second permanent shock are similar up to the one-year yield but are much smaller on the five-year yield. This suggests that the second permanent shock represents a slope factor.

Table 13 also presents the rotation matrix, B, which indicates the extent to which the permanent and transitory shocks are influenced by the shocks to each variables. As we are primarily interested in the permanent shocks, we focus on the last two rows of the B-matrix which represent the first and second permanent shocks, respectively. The first permanent shock mainly comprises shocks to medium- and long-term yields and that the second permanent shock is mainly determined by shocks to the spread between the six-month and one-year yields. Therefore, although each permanent shock does not permit a perfectly obvious interpretation, it appears reasonable to conclude that the two permanent shocks and common trends represent shocks to the level and slope of the yield curve.

5 Globalisation and the Monthly Term Structure

We now attempt to determine how globalisation has affected the monthly term-structure, and specifically, the transmission of changes in monetary policy across the term-structure. In the daily model, we found a break in the relation between short and long rates that strongly coincided with the sharp increase in the purchase of long-term treasuries by foreign central banks. However, no such break was evident when monthly data was used. In any case, the impact of globalisation on the transmission between short- and long yields may be gradual and therefore a clear break in the transmission may be difficult to detect. The aim of this section is to separate the domestic factors from the global factors that influence each yield and analyse whether the Expectations Hypothesis holds when only domestic factors are considered. Further, if globalisation dampens the effect of changes in monetary policy on yields of every maturity, we should observe a much stronger influence of changes in monetary policy on yields across the term structure, once global factors have been excluded.

5.1 Global Factors and Idiosyncratic Components

We extract the global and US-specific factors from each yield using a factor model of the form:

$$X_t = \Lambda F_t + e_t \tag{4}$$

where X_t is a $T \times N$ matrix of yields for N countries, F_t is a $r \times T$ vector of factors common to all countries and Λ is a $N \times r$ matrix of factor loadings with $r \leq N$. e_t is a $T \times N$ vector of idiosyncratic components, which can exhibit weak cross-sectional dependence and weak serial correlation. We estimate F_t by conducting principal component analysis on X_t and obtain the idiosyncratic component from a regression of the yields on the estimated factors.

As the yields are non-stationary and it is not known whether this non-stationarity has a pervasive (common) or an idiosyncratic source, the factors will not be consistently estimated by principal components analysis in a panel with a cross-sectional dimension of the size in our sample. Similarly, the dimension is not sufficiently large to use the Bai and Ng (2004)**?** common-idiosyncratic (I-C) methodology to obtain the number of factors as this requires $N \geq 40$ for consistent estimation. Our approach is to adopt the Bai and Ng (2004) method of conducting principal components analysis on the first differences of the yields and then re-cumulating the estimates to form the factors. However, we ascertain the appropriate number of factors from the proportion of variance explained rather than from the information criteria that Bai and Ng (2004) propose.

Zero-coupon yields on government bonds for eight countries with the same maturities as in the previous section are taken from the dataset of Wright (2011) **?** and the sample period is chosen to coincide with that of the previous section.[16]The yields are plotted in Figures 5a through 5e. It is clear that there has been a downward trend in long-term yields across all countries since the beginning of the 1990s, whereas short-term yields only exhibited a clear downward trend in the first half of that decade.

We find that three factors are sufficient to account for the majority of variation in each yield as additional factors have eigenvalues less than one. Table 14 reports the loadings of each US yield on each of the common factors affecting that yield, as well as the proportion of each yield's variance that is unique to the US. The first factor has the largest effect on yields and increases with maturity. The second factor mainly affects medium-term maturities, while the third factor primarily influences the three-month, one- and ten-year yields. The proportion of each yield's variance that is unique to the US declines with maturity, indicating that globalisation has a larger effect on long-term relative to short-term yields.

We also examine whether the source of the non-stationarity of yields is due to factors, idiosyncratic components, or both. The order of integration of the idiosyncratic component of each yield has important implications for our analysis. As we are primarily interested in whether there is a stronger relation between short- and long-term yields when global influences on yields are removed, the finding of stationary idiosyncratic components would indicate that there is no long-run relation between the US-specific component of yields on government bonds of different maturities. This would imply that the long-run relation between yields is being driven by factors common to other countries.

Therefore, we conduct Augmented Dickey-Fuller unit root tests on the factors and idiosyncratic-components to determine whether the nonstationarity is pervasive (common to all countries) or idiosyncratic (unique to the US).[17] The results of these tests are reported in Table 15 and show that the source of nonstationarity is both pervasive and idiosyncratic. The null hypothesis of a unit root on the levels of the factors and idiosyncratic component is only rejected for the second factor of the six-month, and one- and five-year yields. Table 15 also reports the results of unit root tests on the first differences and they are found to be stationary.

[16]The sample includes the following countries: Australia, Canada, Germany, Japan, New Zealand, Switzerland, United Kingdom and United States.

[17]We also conducted Phillips-Perron unit roots tests and obtained similar results to the ADF tests.

5.2 A CVAR Model of Idiosyncratic Components

Our finding of nonstationary idiosyncratic components allows us to continue with our analysis and examine the long-run relation between the US-specific components of zero-coupon treasury yields. Therefore, we re-estimate (1) with the estimated idiosyncratic components as the endogenous variables in the model. We also include the Federal Funds Rate in order to measure changes in monetary policy, as in the monthly model in the previous section. The model is estimated over the period February 1990 to March 2007, providing 206 observations. Likelihood ratio tests indicated that two lags were again sufficient to capture short-run persistence at monthly frequencies. Table 16 presents the results of the misspecification tests on this model. The model appears to be well specified apart from the rejection of normality for the FFR and the three-month yield. This is again due to excess kurtosis rather than skewness and therefore should not affect the estimated parameters. Table 16 shows that none of the variables can be considered weakly exogenous, in contrast to the monthly model of non-factored yields in the previous section, which found that the ten-year yield was exogenous. This indicates that the exogeneity is due to global factors.

Table 16 also reports the results of the trace test and the characteristic roots in the model. The former suggests that there are at least two cointegration relations. However, the unrestricted coefficients reported in Table 17 show that there is strong adjustment to five cointegrating relations. A rank of five, leaves 0.931 as the largest stationary root in the model, which may correspond to a slow mean reverting process. On balance, and particularly given the large t-statistics on the adjustment coefficients to the five relations, we choose $r = 5$.

The over-identified cointegrating relations, adjustment coefficients and short-run parameters are presented in Table 18. The model is accepted with a $\chi^2(8)$ statistic of 12.8 and a $p-$value of 0.117. The five cointegrating vectors are the five exact spreads relative to the FFR. This suggests that we observe greater evidence for the expectations hypothesis once we focus on the idiosyncratic components of the yields. Each variable is adjusting, again suggesting that the FFR is not a weakly exogenous variable, but instead reacts to changes in the yields. Most of the short-term persistence in the model comes from the FFR and three-month yield at the short- to medium-term maturities and the ten-year yield at medium- to longer-term maturites. Finally, Table 19 indicates that the stochastic trend in the idiosyncratic components comprises shocks to the spreads between the yields. Shocks to the FFR do not have permanent effects, consistent with our earlier findings.

Comparing the model with non-factored yields in the previous section with the model with idiosyncratic components, we can see the effects of finanical globalisation. In the former, there is no clear evidence for the expectations hypothesis, except perhaps at the short-end of the yield curve. In addition, the stochastic trends are given by shocks to the level and slope of the yield curve. When we focus on the US-specific components of yields, we find much stronger evidence for the expectations hypothesis while the stochastic trend is mainly given by shocks to the spreads between the idiosyncratic components. The short-run effects of changes in the FFR are similar in both models, indicating that the short-term effects of monetary policy are not influenced by global factors.

6 Conclusion

We examine the relation between policy rates and the term-structure of interest rates and analyse how globalisation has affected this relation. Using a cointegrated VAR with daily data on US yields, we find that the spreads between short and long-term yields are stationary but that the parameters of the cointegrating relations changed during the conundrum period. We find that changes in the Federal Funds Rate Target have significant short-run effects on the short-end of the yield curve but that, in the long-run, yields are mainly influenced by shocks to long-term yields. We then consider whether the change in the behaviour of long-term yields during the conundrum period may be related to foreign official demand for US Treasuries. We find clear evidence of a long-run relation between long-term yields and foreign offical Treasury holdings, which is consistent with the findings of Warnock and Warnock (2009) and Craine and Martin (2009).

To account for the possibly more gradual effects of financial globalisation, we re-estimate the cointegrated VAR model using monthly yield data since 1990. We first show that, although the spreads between short-term yields are stationary, the spreads between short- and long-term yields are non-stationary. This is consistent with the findings of Shea (1992) and Giese (2008), who show that the yield curve is characterised by more than one stochastic trend at the long-end of the term structure and suggest that the failure of the the Expectations Hupothesis is due to non-stationary term premia. We show that the weighted-spreads between yields, which describe the curvature of the yield curve, are stationary and therefore, that term premia at different maturities cointegrate. The monthly yields in our model are influenced by two stochastic trends which have the approximate interpretation of level and slope factors.

We then attempt to remove the effects of globalisation by obtaining each yield's idiosyncratic component from a regression of the yields on three factors common to a sample of eight countries. When we re-estimate the cointegrated VAR with these idiosyncratic components we find that the Federal Funds Rate and each of the yields are cointegrated. The stochastic trend driving the yields now comprises shocks to the spreads between yields. As there is no clearly exogenous idiosyncratic component, the economic interpretation of this stochastic trend is unclear. However, as the Expectations Hypothesis holds for all idiosyncratic components in this model, the term premium associated with each yield must now be stationary. This suggests that the stochastic trend related to term premia is influenced by global factors. Indeed, Wright (2011) shows that the downward trend in long-term yields across countries is the result of a decline in term premia due to lower uncertainty about inflation.

Our results are also consistent with the findings of Diebold et al (2008) who find a global level factor relating to "global inflation" and a global slope factor relating to a "global business cycle". However, Diebold et al (2008) model these factors as stationary but persistent processes. Therefore, an interesting extension of our analysis would be to examine how our estimated common factors, which are mainly non-stationary, load on measures of global inflation, for example as constructed by Ciccarelli and Mojon (2010)**?**, and global real economic activity.

How do our findings relate to the globalisation and effectiveness of monetary policy debate? Woodford (2007) shows that a central bank will always be able to control domestic inflation regardless of the level of finanical integration because it can influence the opportunity cost of holding money. Our results show that changes in the policy rate have significant short-run effects on the short-end of the yield curve and therefore that, in the context of Woodford's model, monetary policy would still be effective in controlling inflation. However, Kamin (2010) suggests globalisation may make the conduct of monetary policy more complex because it increases the range of shocks to which a central bank must respond. We find that, in the long run, yields are mainly affected by shocks to the level and slope of the yield curve. In terms of the yield curve, the main impact of globalisation appears to be on the term premium, possibly related to uncertainty about global inflation or real output. It is in this sense that global factors have important implications for the conduct of monetary policy.

References

Backus, David K., and Jonathan H. Wright. 2007. "Cracking the Conundrum." *Brookings Papers on Economic Activity*, 38(1): 293–329.

Bai, Jushan, and Serena Ng. 2004. "A PANIC Attack on Unit Roots and Cointegration." *Econometrica*, 72(4): 1127–1177.

Bernanke, Ben S., Vincent R. Reinhart, and Brian P. Sack. 2004. "Monetary Policy Alternatives at the Zero Bound: An Empirical Assessment." *Brookings Papers on Economic Activity*, 35(2): 1–100.

Bertaut, Carol C., and Ralph Tryon. 2007. "Monthly Estimates of U.S. Cross-Border Securities Positions." *FRB International Finance Discussion Paper*, , (910).

Boivin, Jean, and Marc P Giannoni. 2008. "Global Forces and Monetary Policy Effectiveness." National Bureau of Economic Research, Inc NBER Working Papers 13736.

Boivin, Jean, Michael T. Kiley, and Frederic S. Mishkin. 2010. "How Has the Monetary Transmission Mechanism Evolved Over Time?" In *Handbook of Monetary Economics*. Vol. 3 of *Handbook of Monetary Economics*, , ed. Benjamin M. Friedman and Michael Woodford, Chapter 8, 369–422. Elsevier.

Byrne, Joseph P., Giorgio Fazio, and Norbert Fiess. 2010. "Interest Rate Co-movements, Global Factors and the Long End of the Term Spread." Business School, Economics, University of Glasgow Working Papers 2010-10.

Campbell, John Y, and Robert J Shiller. 1987. "Cointegration and Tests of Present Value Models." *Journal of Political Economy*, 95(5): 1062–88.

Ciccarelli, Matteo, and Benoit Mojon. 2010. "Global Inflation." *The Review of Economics and Statistics*, 92(3): 524–535.

Craine, Roger, and Vance Lindsay Martin. 2009. "Interest Rate Conundrum." *The B.E. Journal of Macroeconomics*, 9(1): 8.

Diebold, Francis X., Canlin Li, and Vivian Z. Yue. 2008. "Global yield curve dynamics and interactions: A dynamic Nelson-Siegel approach." *Journal of Econometrics*, 146(2): 351–363.

Engle, Robert F, and Clive W J Granger. 1987. "Co-integration and Error Correction: Representation, Estimation, and Testing." *Econometrica*, 55(2): 251–76.

Faust, Jon, John H. Rogers, Shing-Yi B. Wang, and Jonathan H. Wright. 2007. "The high-frequency response of exchange rates and interest rates to macroeconomic announcements." *Journal of Monetary Economics*, 54(4): 1051–1068.

Giese, Julia V. 2008. "Level, Slope, Curvature: Characterising the Yield Curve in a Cointegrated VAR Model." *Economics - The Open-Access, Open-Assessment E-Journal*, 2(28): 1–20.

Gonzalo, Jesus, and Serena Ng. 2001. "A systematic framework for analyzing the dynamic effects of permanent and transitory shocks." *Journal of Economic Dynamics and Control*, 25(10): 1527–1546.

Gurkaynak, Refet S., Brian Sack, and Jonathan H. Wright. 2007. "The U.S. Treasury yield curve: 1961 to the present." *Journal of Monetary Economics*, 54(8): 2291–2304.

Gürkaynak, Refet S., and Jonathan Wright. 2010. "Macroeconomics and the Term Structure." C.E.P.R. Discussion Papers CEPR Discussion Papers 8018.

Johansen, Soren, and Katarina Juselius. 2001. "Controlling Inflation in a Cointergrated Vector Autoregressive Model with an Application to US Data." European University Institute Economics Working Papers ECO2001/02.

Juselius, Katarina. 2006. *The Cointegrated VAR Model: Methodology and Applications.* Oxford University Press.

Kaminska, Iryna, Andrew Meldrum, and James Smith. 2011. "A global model of international yield curves: no-arbitrage term structure approach." Bank of England Bank of England working papers 419.

Kamin, Steven B. 2010. "Financial globalization and monetary policy." Board of Governors of the Federal Reserve System (U.S.) International Finance Discussion Papers 1002.

Kim, Don H., and Jonathan H. Wright. 2005. "An arbitrage-free three-factor term structure model and the recent behavior of long-term yields and distant-horizon forward rates." Board of Governors of the Federal Reserve System (U.S.) Finance and Economics Discussion Series 2005-33.

Litterman, R., and J. Scheinkman. 1991. "Common factors affecting bond returns." *Journal of Fixed Income*, 1: 49–53.

Rogoff, Kenneth S. 2006. "Impact of globalization on monetary policy." *Proceedings, Federal Reserve Bank of Kansas City*, 265–305.

Shea, Gary S. 1992. "Benchmarking the Expectations Hypothesis of the Interest-Rate Term Structure: An Analysis of Cointegration Vectors." *Journal of Business & Economic Statistics*, 10(3): 347–66.

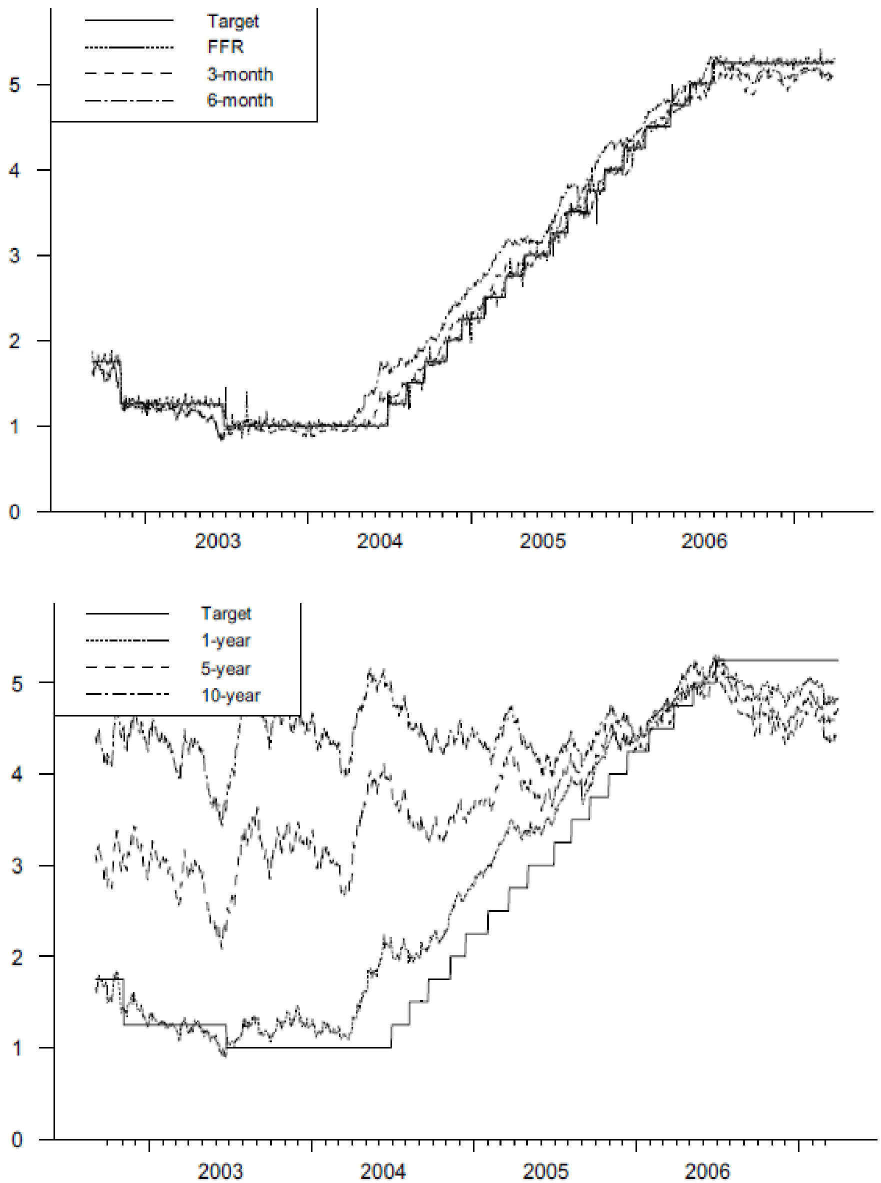
Smith, Josephine M., and John B. Taylor. 2009. "The term structure of policy rules." *Journal of Monetary Economics*, 56(7): 907–917.

Warnock, Francis E., and Veronica Warnock. 2009. "International Capital Flows and U.S. Interest Rates." *Journal of International Money and Finance*, 28: 903–919.

Woodford, Michael. 2007. "Globalization and Monetary Control." In *International Dimensions of Monetary Policy. NBER Chapters*, 13–77. National Bureau of Economic Research, Inc.

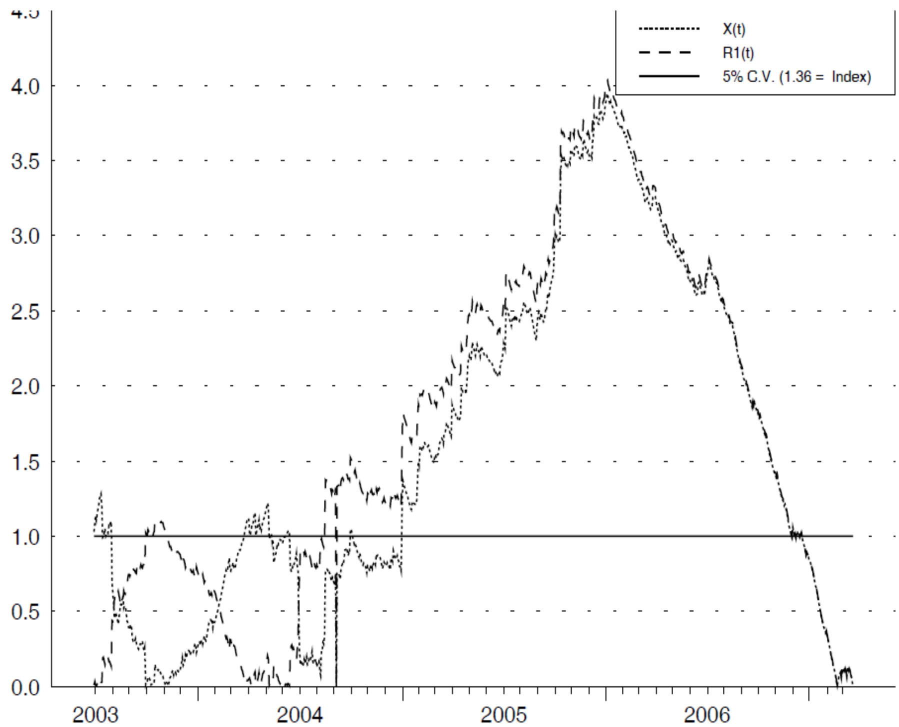
Wright, Jonathan H. 2011. "Term Premia and Inflation Uncertainty: Empirical Evidence from an International Panel Dataset." *American Economic Review*, 101(4): 1514–34.

Figure 1: Daily Federal Funds Rate, Target and Zero-coupon US Yields



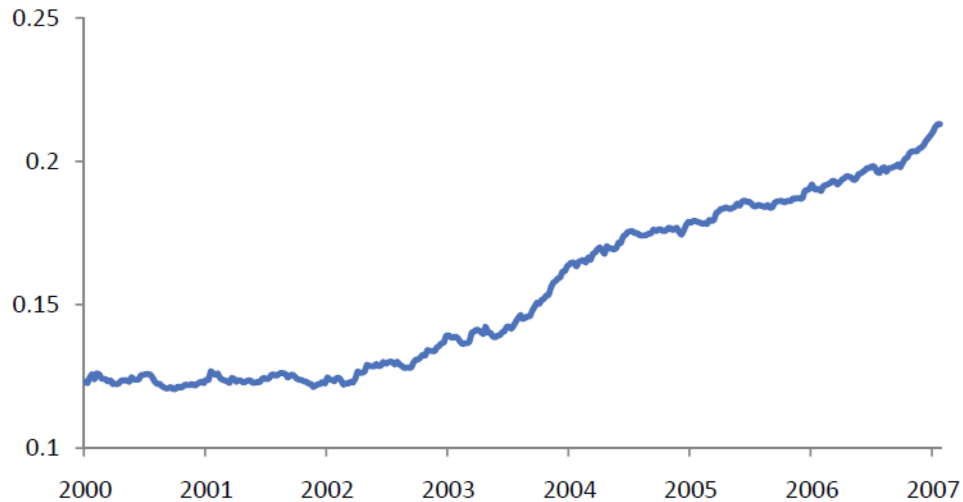
Source: Daily Federal Funds Rate and Target for the period 1 April 2003 to 31 March 2007 are taken from the FRED database. The daily 3-month, 6-month, 1-year, 5-year and 10-year zero-coupon yields are taken from the dataset of Gurkaynak, Sack and Wright (2007).

Figure 2: Test for Constancy of the Log-Likelihood for the daily model with long-term yields



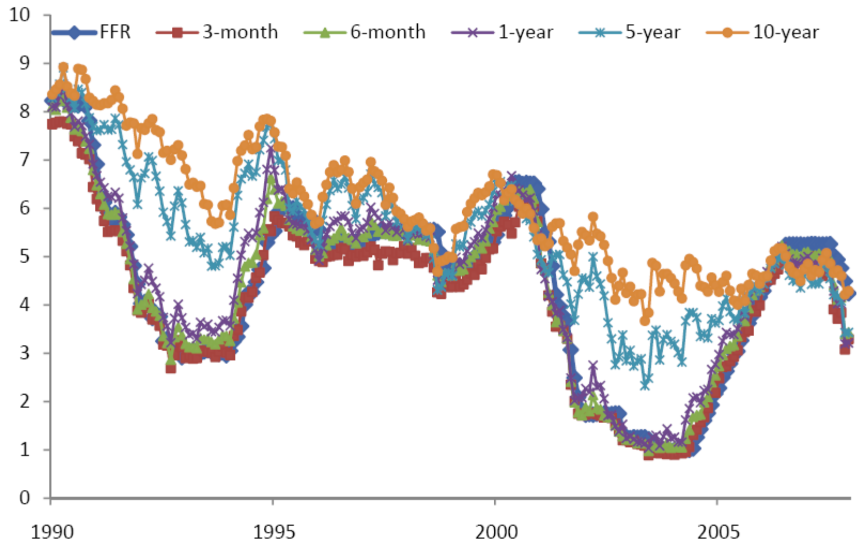
Note: X(1) measures the constancy of the log-likelihood for the full model while R(1) measures the constancy of the "concentrated" model where the short-run effects are concentrated out of the log-likelihood.

Figure 3: Foreign Official Holdings of US treasures as a Share of Total US Government Debt



Note: Foreign official holdings of US treasury securities from the Federal Reserve H4.1 release. Total US government debt is taken from FRED database.

Figure 4: Monthly US Zero-coupon Yields



Note: zero-coupon yields on US treasuries of selected maturities from Wright (2011)

Figure 5a: 3-month zero-coupon Yields

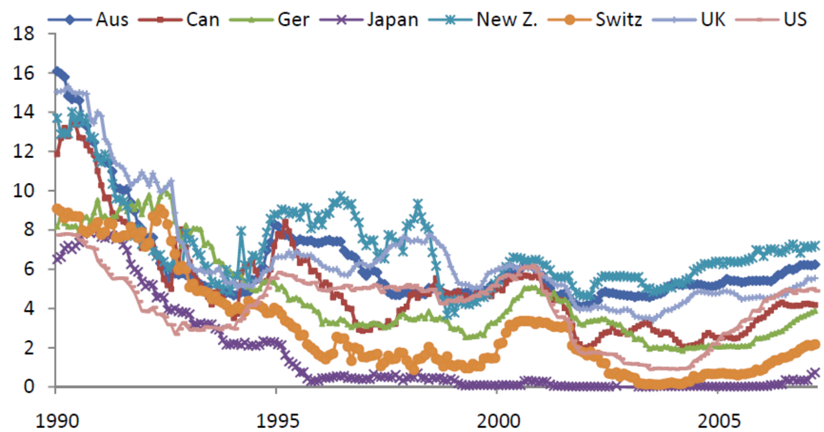


Figure 5b: 6-month zero-coupon Yields

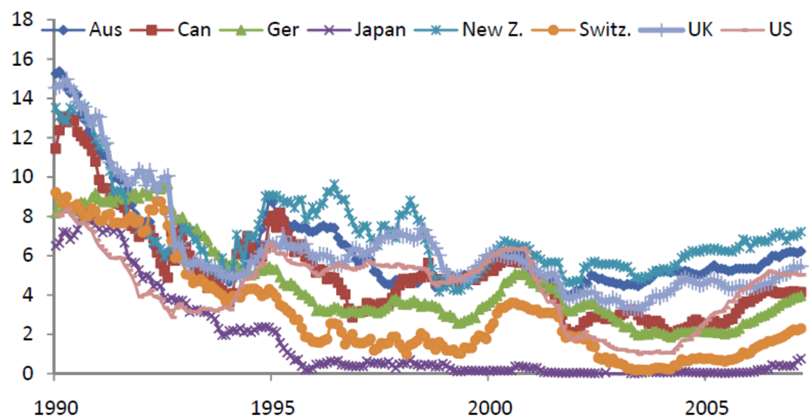


Figure 5c: 1-year zero-coupon Yields

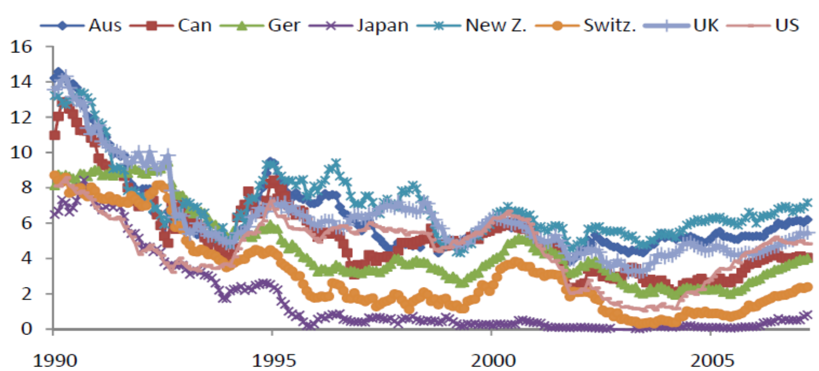


Figure 5d: 5-year zero-coupon Yields

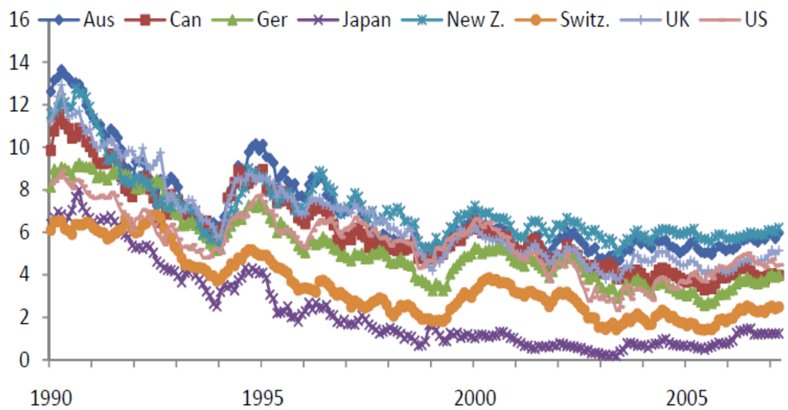
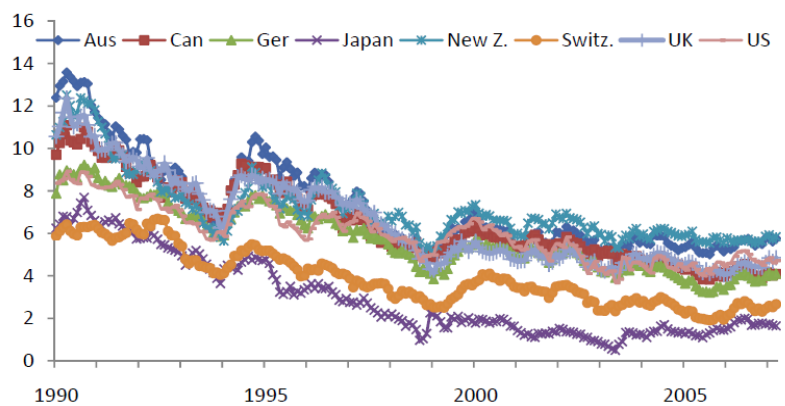


Figure 5e: 10-year zero-coupon Yields



Source: zero-coupon yield data for Figures 3a-3e are from Wright (2011).

Table 1: Misspecifcation tests, characteristic roots and weak exogeneity for daily model with short-term yields

Univariate tests	Δffr_t	$\Delta G3m_t$	$\Delta G6m_t$
ARCH (2)	35.4 [0.000]	5.5 [0.071]	1.1 [0.569]
J.-B.(2)	78.7 [0.000]	58.3 [0.000]	38.3 [0.002]
Trace test	295 (42)	66 (25)	**10** (12)
$Ch.Roots_{(r=2)}$	1.0	0.877	0.381
$W.Exogeneity_{\chi^2(r=2)}$	210.5 [0.000]	44.5 [0.000]	20.0 [0.000]
Multivariate tests			
Autocorrelation	$LM(1)$: $\chi^2(9) = 17.8$ [0.081]		
Normality	$\chi^2(6) = 93.1$ [0.001]		

Note: Tests on the daily model including the federal funds rate from the *FRED* database,and the three-and six-month zero-coupon yields from the dataset of Gurkaynak, Sack and Wright (2007). p-values are in brackets.

Table 2: Cointegrating relations, adjustment coefficients and short-term parameters for daily model with short-term yields

Cointegrating Vectors, β			Adjustment coefficients, α		
Var.	$\widehat{\beta}_1$	$\widehat{\beta}_2$	Eq.	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$
ffr_t	**1.0**	0.0	Δffr_t	$\mathbf{-0.51}$ (−17.8)	0.00 (0.1)
$G3m_t$	$\mathbf{-0.15}$ (−7.6)	**1.0**	$\Delta G3m_t$	0.02 (1.2)	$\mathbf{-0.05}$ (−8.5)
$G6m_t$	0.0	$\mathbf{-1.0}$	$\Delta G6m_t$	0.03 (1.7)	$\mathbf{-0.04}$ (−6.1)
Trg_t	$\mathbf{-0.86}$ (−43.8)	0.0			
Constant	−0.01 (−1.4)	**0.01** (4.7)			

Short-Run Parameters

	$\mathbf{\Gamma}_1$				$\mathbf{A}_0$
	Δffr_{t-1}	$\Delta G3m_{t-1}$	$\Delta G6m_{t-1}$	ΔTrg_{t-1}	ΔTrg_t
Δffr	**0.06** (2.4)	−0.01 (−0.0)	$\mathbf{-0.26}$ (−3.6)	0.07 (1.7)	0.**335** (8.0)
$\Delta G3m$	0.02 (1.2)	0.04 (1.0)	−0.04 (−1.0)	$\mathbf{-0.08}$ (−3.5)	$\mathbf{-0.06}$ (−2.8)
$\Delta G6m$	0.01 (0.4)	**0.2** (4.8)	$\mathbf{-0.18}$ (−4.3)	$\mathbf{-0.05}$ (−2.3)	$\mathbf{-0.05}$ (−2.1)

Note: Overidentified cointegrating relations, adjustment coefficients and short-run parameters. t-statistics are in parentheses and significant coefficients are in bold-face. The data are daily zero-coupon yields for the period 1 April 2003 to 31 March 2007 from the dataset of Gurkaynak, Sack and Wright (2007).

Table 3: The long-run impact matrix, common trends ($\alpha_\perp$) and loadings ($\tilde{\beta}_\perp$) for daily model of short-term yields

Long-Run Impact matrix, C					
	$\widehat{\varepsilon}_{ffr}$	$\widehat{\varepsilon}_{g3m}$	$\widehat{\varepsilon}_{g6m}$	$\tilde{\beta}_\perp$	$\alpha_\perp$
ffr	0.02 (1.0)	$-$**0.51** (-1.9)	**0.67** (2.4)	0.67 (2.4)	0.03 (1.1)
$G3m$	0.13 (1.0)	$-$**3.44** (-1.9)	**4.47** (2.4)	4.47 (2.4)	-0.77 (-8.0)
$G6m$	0.13 (1.0)	$-$**3.44** (-1.9)	**4.47** (2.4)	4.47 (2.4)	1.0

Note: Long run impact of shocks to the federal fund rate, three- and six-month treasury zero-coupon yields, along with the common trends and corresponding loadings. t-statistics are in parentheses and significant coefficients are in bold-face.

Table 4: Misspecifcation tests, characteristic roots and weak exogeneity for daily model with long-term yields

Univariate tests	Δffr_t	$\Delta G3m_t$	$\Delta G1y_t$	$\Delta G5y_t$	$\Delta G10y_t$
ARCH (3)	8.3 [0.049]	7.1 [0.061]	5.9 [0.113]	3.2 [0.362]	4.0 [0.259]
J.-B.(3)	134.2 [0.000]	65.4 [0.000]	56.1 [0.000]	6.6 [0.037]	5.2 [0.076]
Trace test	361 (79)	112 (54)	43 (35)	**19** (20)	7 (9)
$\underset{(r=4)}{Ch.Roots}$	1.0	0.931	0.901	0.844	0.442
$\underset{\chi^2(r=4)}{W.Exogeneity,}$	252.7 [0.000]	56.8 [0.000]	19.5 [0.001]	11.4 [0.023]	11.4 [0.023]
<u>Multivariate tests</u>					
Autocorrelation	$LM(1)$: $\chi^2(25) = \underset{[0.071]}{35.2}$		$LM(2)$: $\underset{[0.129]}{33.1}$		
Normality	$\chi^2(10) = \underset{[0.000]}{150.5}$				

Note: Daily model includes the federal funds rate from the *FRED* database, and the three-month, one-, five- and ten-year zero-coupon yields from the dataset of Gurkaynak, Sack and Wright (2007). p-values are in brackets.

Table 5: Unrestricted adjustment coeficients for daily model with long-term yields

Unrestricted Adjustment coefficients, $\widehat{\alpha}^u$					
	$\widehat{\alpha}_1^u$	$\widehat{\alpha}_2^u$	$\widehat{\alpha}_3^u$	$\widehat{\alpha}_4^u$	$\widehat{\alpha}_5^u$
Δffr	**0.02** (15.6)	**0.003** (2.2)	0.001 (0.3)	-0.002 (-0.3)	-0.002 (-0.3)
$\Delta G3m$	**0.003** (4.1)	$\mathbf{-0.01}$ (-6.7)	**0.002** (2.8)	-0.001 (-0.4)	-0.002 (-0.5)
$\Delta G1y$	0.001 (0.7)	-0.002 (-1.3)	**0.003** (2.2)	$\mathbf{-0.004}$ (-3.1)	0.002 (1.4)
$\Delta G5y$	-0.002 (-1.0)	0.002 (1.2)	**0.01** (3.2)	$\mathbf{-0.01}$ (-3.0)	0.001 (0.5)
$\Delta G10y$	-0.002 (-1.5)	0.002 (1.4)	**0.004** (2.7)	$\mathbf{-0.01}$ (-3.4)	-0.001 (-0.1)

Note: Daily model includes the federal funds rate and target taken from the *FRED*, database and the three-month, one-, five- and ten-year zero-coupon yields from the dataset of Gurkaynak, Sack and Wright (2007). t-statistics are in parentheses and significant coefficients are in bold-face.

Table 6: Cointegrating relations, adjustment coefficients and short-run parameters for daily model with long-term yields

Cointegration Vectors, β				
Var.	$\widehat{\beta}_1$	$\widehat{\beta}_2$	$\widehat{\beta}_3$	$\widehat{\beta}_4$
ffr_t	0.0	**1.0**	0.0	0.0
$G3m_t$	**1.0**	**−0.21** (−10.3)	**1.0**	**1.0**
$G1y_t$	0.0	0.0	**−1.0**	0.0
$G5y_t$	0.0	0.0	0.0	**−1.0**
$G10y_t$	**−0.84** (−13.2)	0.0	0.0	0.0
Trg_t	0.0	**−0.8** (−40.4)	0.0	0.0
break	**−3.91** (−3.2)	0.0	**−0.52** (3.8)	**2.77** (3.6)
Constant	**−0.75** (−2.6)	0.0	0.0	0.0

Adjustment Coefficients, α				
Eq.	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$	$\widehat{\alpha}_3$	$\widehat{\alpha}_4$
Δffr_t	**−0.05** (−3.9)	**−0.6** (−15.2)	**−0.05** (−2.8)	**−0.08** (−3.8)
$\Delta G3m_t$	0.01 (−0.6)	−0.03 (−1.3)	**−0.02** (2.6)	−0.01 (−0.9)
$\Delta G1y_t$	−0.01 (−1.4)	−0.01 (−0.1)	−0.01 (−0.8)	−0.02 (−1.3)
$\Delta G5y_t$	**−0.05** (−3.2)	0.02 (0.5)	**−0.04** (−2.0)	**−0.08** (−3.0)
$\Delta G10y_t$	**−0.04** (−3.0)	0.04 (1.0)	**−0.04** (−1.9)	**−0.07** (−2.8)

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Table 6 (continued)

Short-run adjustment coefficients, Γ_1							A_0
Δ	ffr_{t-1}	$G3m_{t-1}$	$G1y_{t-1}$	$G5y_{t-1}$	$G10y_{t-1}$	Trg_{t-1}	Trg_t
ffr_t	**0.07** (1.9)	**−0.3** (−4.5)	0.2 (1.9)	−0.06 (−0.3)	−0.06 (−0.5)	−0.01 (−0.3)	**0.40** (8.4)
$G3m_t$	0.02 (1.1)	0.03 (1.0)	0.06 (1.2)	−0.003 (−0.1)	−0.06 (−0.9)	**−0.13** (−5.7)	**−0.05** (−2.3)
$G1y_t$	0.03 (1.0)	**0.11** (2.1)	**0.16** (2.0)	**−0.33** (−2.7)	**0.2** (2.0)	−0.03 (−0.8)	0.01 (0.2)
$G5y_t$	−0.01 (−0.1)	0.12 (1.6)	**0.22** (1.9)	**−0.39** (−2.1)	**0.3** (2.0)	−0.04 (−0.7)	−0.01 (−0.1)
$G10y_t$	−0.01 (−0.1)	0.07 (1.0)	0.1 (1.0)	**−0.35** (−2.1)	**0.34** (2.4)	−0.02 (−0.4)	−0.05 (−0.8)

Note: Over-identified relations for the daily model. t-statistics are in parentheses and significant coefficients are in bold-face.

Table 7: The long-run impact matrix, common trends ($\alpha_\perp$) and loadings ($\tilde{\beta}_\perp$) for daily model with long-term yields

Long-Run Impact Matrix, C							
	$\widehat{\varepsilon}_{ffr}$	$\widehat{\varepsilon}_{g3m}$	$\widehat{\varepsilon}_{g1y}$	$\widehat{\varepsilon}_{g5y}$	$\widehat{\varepsilon}_{g10y}$	$\tilde{\beta}_\perp$	$\alpha'_\perp$
ffr_t	0.01 (1.4)	-0.01 (-0.2)	0.11 (1.1)	$\mathbf{-0.44}$ (-4.4)	$\mathbf{-0.46}$ (5.1)	0.5	0.03
$G3m_t$	0.06 (1.4)	-0.07 (-0.2)	0.51 (1.1)	$\mathbf{-2.1}$ (-4.4)	$\mathbf{-2.19}$ (5.1)	2.2	-0.03
$G1y_t$	0.06 (1.4)	-0.07 (-0.2)	0.51 (1.1)	$\mathbf{-2.1}$ (-4.4)	$\mathbf{-2.19}$ (5.1)	2.2	0.23
$G5y_t$	0.06 (1.4)	-0.07 (-0.2)	0.51 (1.1)	$\mathbf{-2.1}$ (-4.4)	$\mathbf{-2.19}$ (5.1)	2.2	-0.96
$G10y_t$	0.08 (1.4)	-0.08 (-0.2)	0.61 (1.1)	$-2.\mathbf{5}$ (-4.4)	$\mathbf{-2.61}$ (5.1)	2.6	1.0

Note: Long run impact of shocks to the federal fund rate, three-month, and one-, five and ten-year zero-coupon treasury yields, together with the common trends and corresponding loadings . The model is estimated for the period 1 April 2003 to 31 March 2007 using the dataset of Gurkaynak, Sack and Wright (2007). The federal funds rate is taken from the *FRED* database. t-statistics are in parentheses and significant coefficients are in bold-face.

Table 8: Equilibrium Correction and long-run impacts from model with Foreign Official Holdings of US treasuries

Cointegrating relations, β			Adjustment coefficients, α		
Var.	$\widehat{\beta}_1$	$\widehat{\beta}_2$	Eqn.	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$
$G5y$	$\mathbf{-1.0}$	0.0	$\Delta G5y$	-0.01 (-1.6)	**0.03** (3.9)
$G10y$	**1.0**	$\mathbf{-2.93}$ (-41.5)	$\Delta G10y$	-0.01 (-1.7)	**0.03** (3.5)
official	0.0	**1.0**	Δ*official*	**0.001** (8.6)	-0.001 (-1.1)
constant	**2.39** (6.8)	0.0			

Short-Run Parameters, Γ_1			
	$\Delta G5y$	$\Delta G10y$	Δ*official*
$\Delta G5y$	0.32 (1.6)	-0.25 (-1.2)	-0.93 (-0.7)
$\Delta G10y$	**0.36** (1.9)	-0.29 (-1.4)	-0.7 (-0.6)
Δ*official*	0.01 (1.2)	$\mathbf{-0.02}$ (-2.1)	-0.06 (-0.9)

Long-Run Impact Matrix, C				loadings	CT_1
	$\widehat{\varepsilon}_{G5y}$	$\widehat{\varepsilon}_{G10y}$	$\widehat{\varepsilon}_{official}$	$\tilde{\beta}_{\perp}$	$\alpha'_{\perp}$
$G5y$	-0.35 (-1.2)	0.42 (1.3)	**0.34** (2.1)	0.42	-0.85
$G10y$	-0.35 (-1.2)	0.42 (1.3)	**0.34** (2.1)	0.42	1.0
official	-1.03 (-1.2)	1.23 (1.3)	**0.99** (2.1)	1.23	0.82

Note: Cointegration relations, adjustment coefficients, short-run parameters and long-run impact matrix from model with the five- and ten-year yields and foreign official holdings of US securities. The model is estimated using weekly data from 5 January 2000 to 28 March 2007. Official holdings data is taken from Federal Reserve's H4.1 release. Yield data is from Gurkaynak, Sack and Wright (2007)**?**. t-statistics are in parentheses and significant coefficients are in bold.

Table 9: Misspecifcation tests, characteristic roots and weak exogeneity for the monthly model

Univariate tests	Δffr_t	$\Delta G3m_t$	$\Delta G6m_t$	$\Delta G1y_t$	$\Delta G5y_t$	$\Delta G10y_t$
ARCH (2)	5.2 [0.073]	0.2 [0.891]	2.9 [0.236]	4.1 [0.132]	0.2 [0.913]	0.6 [0.725]
J.-B.(2)	35.9 [0.000]	19.3 [0.000]	21.3 [0.002]	6.1 [0.055]	1.6 [0.449]	1.9 [0.389]
Trace test	272 (104)	156 (77)	92 (54)	48 (35)	**17** (20)	3 (9)
Ch.Roots $(r=4)$	1.0	1.0	0.877	0.811	0.531	0.487
W.Exogeneity $\chi^2(r=4)$	120.5 [0.000]	52.7 [0.000]	23.9 [0.000]	10.8 [0.056]	11.1 [0.030]	7.0 [0.225]
Multivariate tests						
Autocorrelation	LM(1): $\chi^2(36) =$ 48.9 [0.081]			LM(2): 45.2 [0.141]		
Normality	$\chi^2(12) =$ 93.1 [0.001]					

Note: Tests on the monthly model including the federal funds rate from the *FRED* database, and the three-and six-month, one-, five- and ten-year zero-coupon yields from Wright (2011). p-values are in brackets.

Table 10: Unrestricted adjustment coeficients for the monthly model

Unrestricted Adjustment coefficients, $\widehat{\alpha}^u$						
Eq.	$\widehat{\alpha}_1^u$	$\widehat{\alpha}_2^u$	$\widehat{\alpha}_3^u$	$\widehat{\alpha}_4^u$	$\widehat{\alpha}_5^u$	$\widehat{\alpha}_6^u$
Δffr_t	**0.08** (12.8)	0.01 (0.7)	**0.01** (2.2)	0.01 (1.0)	-0.01 (-0.1)	0.01 (0.1)
$\Delta G3m_t$	**0.05** (5.1)	$-$**0.06** (-5.4)	-0.01 (-0.4)	**0.03** (2.8)	-0.01 (-0.3)	0.01 (0.8)
$\Delta G6m_t$	**0.03** (2.8)	$-$**0.03** (-2.9)	**0.02** (2.1)	**0.03** (2.2)	-0.01 (-1.2)	0.013 (1.2)
$\Delta G1y_t$	0.02 (1.7)	$-$**0.031** (-2.2)	0.02 (1.6)	0.01 (0.6)	-0.02 (-1.4)	0.02 (1.4)
$\Delta G5y_t$	-0.01 (-0.6)	$-$**0.05** (-2.4)	**0.04** (2.1)	-0.03 (-1.4)	-0.01 (-0.3)	0.03 (1.5)
$\Delta G10y_t$	-0.02 (-1.2)	-0.02 (-0.8)	**0.04** (2.0)	-0.02 (-1.2)	0.008 (0.5)	0.03 (1.6)

Note: Monthly model includes the federal funds rate taken from the *FRED*, database and the three- and six-month, one-, five- and ten-year zero-coupon yields from Wright (2011). The model is estimated for the period January 1990 to March 2007. t-statistics are in parentheses and significant coefficients are in bold-face.

Table 11: Cointegrating relations, adjustment coefficients and short-run parameters for the monthly model

Cointegration vectors, β				
Var.	$\widehat{\beta}_1$	$\widehat{\beta}_2$	$\widehat{\beta}_3$	$\widehat{\beta}_4$
ffr_t	$\mathbf{-1.01}$ (-91.5)	$\mathbf{-0.93}$ (97.9)	0.0	$\mathbf{-0.78}$ (-47.0)
$G3m_t$	0.0	**1.0**	0.0	0.0
$G6m_t$	**1.0**	0.0	0.0	0.0
$G1y_t$	0.0	0.0	$\mathbf{-0.57}$ (-22.3)	0.0
$G5y_t$	0.0	0.0	**1.0**	**1.0**
$G10y_t$	0.0	0.0	$\mathbf{-0.43}$ (-16.9)	$\mathbf{-0.22}$ (-13.1)
Constant	**0.13** (4.3)	0.0	0.0	$\mathbf{-0.07}$ (-3.7)

Adjustment Coefficients, α				
Eq.	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$	$\widehat{\alpha}_3$	$\widehat{\alpha}_4$
Δffr_t	**0.7** (4.5)	-0.11 (-1.4)	-0.05 (-0.5)	0.14 (1.4)
$\Delta G3m_t$	0.22 (0.8)	$\mathbf{-0.55}$ (-3.9)	$\mathbf{-0.48}$ (-2.8)	$\mathbf{-0.38}$ (-2.1)
$\Delta G6m_t$	-0.28 (-0.9)	-0.07 (-0.4)	$\mathbf{-0.46}$ (-2.4)	$\mathbf{-0.49}$ (-2.4)
$\Delta G1y_t$	-0.18 (-0.5)	-0.04 (-0.2)	$\mathbf{-0.32}$ (-1.9)	-0.36 (-1.7)
$\Delta G5y_t$	-0.56 (-1.0)	0.04 (0.1)	$\mathbf{-0.5}$ (-1.9)	$\mathbf{-0.53}$ (-1.9)
$\Delta G10y_t$	0.00 (0.0)	0.00 (0.0)	0.00 (0.0)	0.00 (0.0)

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Table 11 (continued)

Short Run Adjustment Coefficients, Γ_1						
Δ	ffr_{t-1}	$G3m_{t-1}$	$G6m_{t-1}$	$G1y_{t-1}$	$G5y_{t-1}$	$G10y_{t-1}$
ffr_t	$-0.$**1** (−1.9)	**0.17** (2.2)	−0.33 (−1.8)	0.26 (1.6)	**−0.28** (−2.5)	**0.18** (2.0)
$G3m_t$	**−0.17** (−1.9)	−0.04 (−0.3)	0.39 (1.2)	**−0.48** (−1.9)	0.26 (1.3)	−0.09 (−0.5)
$G6m_t$	−0.04 (−0.4)	0.12 (0.8)	−0.01 (−0.1)	−0.09 (−0.3)	0.01 (0.1)	0.1 (0.5)
$G1y_t$	−0.12 (−0.8)	0.15 (0.7)	0.07 (0.2)	−0.06 (−0.1)	−0.22 (−0.7)	0.32 (1.3)
$G5y_t$	−0.3 (−1.6)	0.14 (0.5)	0.11 (0.2)	0.24 (0.4)	**−0.7** (−2.1)	**0.67** (2.1)
$G10y_t$	−0.15 (−0.9)	0.16 (0.7)	−0.24 (−0.4)	0.54 (1.1)	**−0.73** (−2.1)	**0.61** (2.1)

Note: over-identified relations for the monthly model including the federal funds rate taken from the *FRED*, database and the three- and six-month, one-, five- and ten-year zero-coupon yields from Wright (2011). The model is estimated for the period January 1990 to March 2007. t-statistics are in parentheses and significant coefficients are in bold-face

Table 12: Common trends and loadings in the monthly model

Common Trends, $\alpha_\perp$			Loadings, $\tilde{\beta}_\perp$:	
	CT_1	CT_2	CT_1	CT_2
ffr	0.0	−0.05	0.79	−1.89
$G3m$	0.0	−0.08	0.74	−1.75
$G6m$	0.0	1.0	0.80	−1.91
$G1y$	0.0	−0.79	0.69	−2.35
$G5y$	0.0	−0.33	0.93	−1.26
$G10y$	1.0	0.0	1.27	0.26

Note: two common trends comprising shocks to the federal fund rate, three-month, and one-, five and ten-year zero-coupon treasury yields, together with the loadings to both common trends .t-statistics are in parentheses and significant coefficients are in bold-face.

Table 13: The structural long-run impact matrix and rotation matrix

Structural Long-Run Impact Matrix, $\tilde{C} = CB^{-1}$						
	T_1	T_2	T_3	T_4	P_1	P_2
ffr	0.0	0.0	0.0	0.0	0.935	0.830
$G3m$	0.0	0.0	0.0	0.0	0.869	0.772
$G6m$	0.0	0.0	0.0	0.0	0.945	0.839
$G1y$	0.0	0.0	0.0	0.0	1.000	1.000
$G5y$	0.0	0.0	0.0	0.0	0.837	0.575
$G10y$	0.0	0.0	0.0	0.0	0.617	0.000
Rotation Matrix, B						
	$\widehat{\varepsilon}_{ffr}$	$\widehat{\varepsilon}_{g3m}$	$\widehat{\varepsilon}_{g6m}$	$\widehat{\varepsilon}_{g1y}$	$\widehat{\varepsilon}_{g5y}$	$\widehat{\varepsilon}_{g10y}$
T_1	1.000	0.006	−0.180	0.110	0.114	−0.148
T_2	−0.443	1.000	0.012	−0.449	0.711	−0.745
T_3	−0.110	−0.446	1.000	−0.571	0.391	−0.360
T_4	−0.125	0.101	−0.494	1.000	−0.962	0.434
P_1	−0.162	0.192	−0.240	0.846	−0.987	1.000
P_2	−0.038	−0.088	1.000	−0.818	−0.298	0.470

Note: Long run impact of structural shocks to the federal fund rate, three-month, and one-, five and ten-year zero-coupon treasury yields, together with rotation matrix of orthogonal transitory and permanet shocks. t-statistics are in parentheses and significant coefficients are in bold-face.

Table 14: Unique US variance and loadings of US yields to common factors

Yield	factor 1	factor 2	factor 3	Unique var.
$G3m$	0.665	0.066	-0.256	0.49
$G6m$	0.734	-0.092	-0.086	0.45
$G1y$	0.752	-0.209	-0.144	0.37
$G5y$	0.809	-0.226	-0.075	0.29
$G10y$	0.834	0.016	-0.178	0.27

Note: estimates of loadings of each US yield on the three common factors and the proportion of variance of each yield explained by the idiosyncratic component using data from Wright (2011).

Table 15: Tests of stationarity on common factors and US idiosyncratic component

Yield		factor 1	factor 2	factor 3	Idiosyn.
$G3m$	level	-0.51	-1.38	-1.18	-0.52
	diff.	-6.14***	-7.97***	-11.78***	-6.43***
$G6m$	level	-0.57	-1.7*	-1.16	-0.59
	diff.	-6.74***	-8.34***	-9.34***	-6.68***
$G1y$	level	-0.65	-1.8*	-1.19	-0.73
	diff.	-7.29***	-9.12***	-10.3***	-7.38***
$G5y$	level	-1.23	-2.07**	-1.26	-1.36
	diff.	-8.64***	-9.48***	-9.86***	-9.19***
$G10y$	level	-1.61	-1.58	-2.35**	-1.39
	diff.	-9.32***	-10.83***	-10.56***	-9.47***

Note: Augmented Dickey-Fuller unit root tests on the levels and first-differences of each of the common factors and idiosyncratic components. Lag length for the tests are given by 4(min(N,T)/100)^1/4 as in Bai and Ng (2004). *,**,*** indicate rejection of a unit-root at the ten, five and one percent levels respectively.

Table 16: Misspecifcation tests, characteristic roots and weak exogeneity for the monthly model of idiosyncratic components of US yields

Univariate tests	Δffr_t	$\Delta G3m_t$	$\Delta G6m_t$	$\Delta G1y_t$	$\Delta G5y_t$	$\Delta G10y_t$
ARCH (2)	5.1 [0.083]	5.42 [0.066]	3.88 [0.143]	2.49 [0.288]	1.46 [0.482]	0.83 [0.662]
J.-B.(2)	11.83 [0.003]	9.4 [0.039]	5.18 [0.079]	1.33 [0.515]	1.71 [0.425]	2.55 [0.286]
Trace test	199 (118)	109 (89)	**57** (63)	35 (43)	22 (26)	9 (12)
$Ch.Roots_{(r=5)}$	1.0	0.934	0.901	0.901	0.834	0.781
$W.Exogeneity_{\chi^2(r=5)}$	19.7 [0.001]	22.39 [0.000]	12.5 [0.021]	12.43 [0.023]	14.2 [0.014]	13.8 [0.017]
<u>Multivariate tests</u>						
Autocorrelation	LM(1): $\chi^2(36) =$ 48.6 [0.078]			LM(2): 43.3 [0.187]		
Normality	$\chi^2(12) =$ 86.1 [0.003]					

Note: Tests on model idiosyncratic components of the three-and six-month, one-, five- and ten-year zero-coupon yields estimated using Wright (2011), along with the federal funds rate from the *FRED* database. p-values are in brackets.

Table 17: Unrestricted adjustment coefficients from monthly model of idiosyncratic components

Unrestricted Adjustment coefficients, $\widehat{\alpha}^u$					
Eq.	$\widehat{\alpha}^u_1$	$\widehat{\alpha}^u_2$	$\widehat{\alpha}^u_3$	$\widehat{\alpha}^u_4$	$\widehat{\alpha}^u_5$
Δffr_t	**0.28** (3.9)	-0.08 (-1.0)	$-\mathbf{0.23}$ (-1.9)	**0.11** (2.6)	-0.05 (-1.7)
$\Delta G3m_t$	**0.27** (4.5)	-1.0 (-1.5)	0.09 (0.8)	-0.03 (-0.79)	$-\mathbf{0.07}$ (-2.8)
$\Delta G6m_t$	0.02 (0.25)	$-\mathbf{0.26}$ (-3.4)	-0.08 (-0.7)	-0.03 (-0.9)	$-\mathbf{0.08}$ (-2.7)
$\Delta G1y_t$	$-\mathbf{0.21}$ (-2.3)	$-\mathbf{0.23}$ (2.2)	-0.12 (0.8)	0.05 (1.0)	$-\mathbf{0.11}$ (-3.0)
$\Delta G5y_t$	$-\mathbf{0.35}$ (-2.8)	-0.13 (-0.1)	$-\mathbf{0.57}$ (-2.6)	-0.12 (-1.6)	$-\mathbf{0.12}$ (-2.3)
$\Delta G10y_t$	$-\mathbf{0.28}$ (-2.4)	-0.06 (-0.5)	$-\mathbf{0.62}$ (-3.1)	$-\mathbf{0.12}$ (1.9)	-0.08 (-1.6)

Note: Monthly model of idiosyncratic components of three- and six-month and one-, five- and ten-year US zero-coupon treasury yields. The model is estimated for the period January 1990 to March 2007. t-statistics are in parentheses and significant coefficients are in bold-face.

Table 18: Cointegrating relations, adjustment coefficients and short-run parameters for model with US idiosyncratic components

Cointegration vectors, β					
Var.	$\widehat{\beta}_1$	$\widehat{\beta}_2$	$\widehat{\beta}_3$	$\widehat{\beta}_4$	$\widehat{\beta}_5$
ffr_t	**1.0**	**1.0**	**1.0**	**1.0**	**1.0**
$G3m_t$	**−1.0**				
$G6m_t$		**−1.0**			
$G1y_t$			**−1.0**		
$G5y_t$				**−1.0**	
$G10y_t$					**−1.0**
Constant	0.0	0.0	**−0.66** (−18.3)	**−0.45** (−15.2)	0.0

Adjustment coefficients, α					
Eq.	$\widehat{\alpha}_1$	$\widehat{\alpha}_2$	$\widehat{\alpha}_3$	$\widehat{\alpha}_4$	$\widehat{\alpha}_5$
Δffr_t	0.1 (1.6)	**−0.17** (−2.1)	**−0.41** (−3.8)	**0.15** (1.9)	−0.05 (−1.0)
$\Delta G3m_t$	−0.003 (−0.1)	**0.24** (3.6)	**−0.34** (−3.9)	0.07 (1.0)	−0.01 (−0.3)
$\Delta G6m_t$	**−0.18** (−2.9)	**−0.25** (−3.3)	−0.07 (−0.7)	−0.003 (0.03)	−0.01 (−0.2)
$\Delta G1y_t$	**−0.27** (−3.5)	**0.23** (2.3)	0.09 (0.7)	0.05 (−0.4)	−0.09 (−1.4)
$\Delta G5y_t$	**−0.32** (−3.0)	**0.25** (1.9)	0.02 (0.1)	**0.33** (2.3)	**−0.26** (−3.1)
$\Delta G10y_t$	**−0.28** (−2.8)	0.21 (1.7)	0.03 (0.15)	**0.27** (2.0)	**−0.2** (−2.5)

(continued on next page)

Table 18: (continued)

Short Run Adjustment Coefficients, Γ_1

Δ	ffr_{t-1}	$G3m_{t-1}$	$G6m_{t-1}$	$G1y_{t-1}$	$G5y_{t-1}$	$G10y_{t-1}$
ffr_t	**0.3** (4.3)	0.22 (1.4)	−0.27 (−0.8)	0.24 (0.8)	−0.04 (−0.2)	0.04 (0.2)
$G3m_t$	0.07 (1.2)	−0.1 (−0.7)	0.2 (0.6)	−0.09 (−0.4)	0.02 (0.1)	0.03 (0.2)
$G6m_t$	**0.11** (1.9)	**0.25** (1.9)	0.03 (−0.1)	−0.18 (−0.7)	−0.16 (−0.7)	0.2 (1.1)
$G1y_t$	0.08 (0.9)	**0.41** (2.1)	0.09 (0.2)	0.14 (0.4)	−0.34 (−1.6)	**0.41** (1.9)
$G5y_t$	−0.03 (−0.3)	0.35 (1.3)	0.43 (0.8)	0.04 (0.1)	−0.63 (−1.6)	**0.76** (2.4)
$G10y_t$	0.02 (0.2)	0.15 (0.6)	0.48 (0.9)	−0.03 (−0.1)	−0.49 (−1.3)	**0.58** (2.0)

Note: over-identified relations for the model including the federal funds rate and idiosyncratic components of the three-month, one-, five- and ten-year zero-coupon US treasury yields estimated from Wright (2011). The model is estimated for the period January 1990 to March 2007. t-statistics are in parentheses and significant coefficients are in bold-face

Table 19: The long-run impact matrix, common trends and loadings for the model with US idiosyncratic components

Long-Run Impact matrix, C								
	$\widehat{\varepsilon}_{ffr}$	$\widehat{\varepsilon}_{g3m}$	$\widehat{\varepsilon}_{g6m}$	$\widehat{\varepsilon}_{g1y}$	$\widehat{\varepsilon}_{g5y}$	$\widehat{\varepsilon}_{g10y}$	$\widetilde{B}_{\perp}$	$\alpha_{\perp}$
ffr	-0.52 (-0.3)	3.21 (1.8)	$\mathbf{-6.1}$ (-2.0)	**4.76** (2.1)	$\mathbf{-3.47}$ (-2.0)	**2.95** (1.9)	-6.1	0.1
$G3m$	-0.33 (-0.6)	3.21 (1.8)	$\mathbf{-6.1}$ (-2.0)	**4.76** (2.1)	$\mathbf{-3.47}$ (-2.0)	**2.95** (1.9)	-6.1	-0.52
$G6m$	-0.33 (-0.6)	3.21 (1.8)	$\mathbf{-6.1}$ (-2.0)	**4.76** (2.1)	$\mathbf{-3.47}$ (-2.0)	**2.95** (1.9)	-6.1	1.0
$G1y$	-0.33 (-0.6)	3.21 (1.8)	$\mathbf{-6.1}$ (-2.0)	**4.76** (2.1)	$\mathbf{-3.47}$ (-2.0)	**2.95** (1.9)	-6.1	-0.78
$G5y$	-0.33 (-0.6)	3.21 (1.8)	$\mathbf{-6.1}$ (-2.0)	**4.76** (2.1)	$\mathbf{-3.47}$ (-2.0)	**2.95** (1.9)	-6.1	0.56
$G10y$	-0.33 (-0.6)	3.21 (1.8)	$\mathbf{-6.1}$ (-2.0)	**4.76** (2.1)	$\mathbf{-3.47}$ (-2.0)	**2.95** (1.9)	-6.1	-0.48

Note: Long run impact of shocks to the federal fund rate and the idiosyncratic components of the three- and six-month, and one-, five and ten-year zero-coupon US treasury yields, together with the common trends and their corresponding loadings. t-statistics are in parentheses and significant coefficients are in bold-face.