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The Contribution of US Bond Demand to the US Bond Yield Conundrum of 2004 to 2007: An Empirical Investigation

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Abstract

Although the federal funds rate started rising from mid-2004 US long term rates continued to fall. A likely contributory factor to this conundrum was the contemporaneous increase in US bond demand. Using ARDL-based models, which accommodate structural breaks, this paper estimates the impact of demand on US bond yields in the conundrum period. This impact is shown to have been everywhere significantly negative. The fact that our model fully explains the bond yield conundrum gives support to the hypothesis that the US CDO market was rapidly expanded before 2007 chiefly to absorb the overspill of global demand for safe assets.

Key Words: ARDL modelling; bond yields; bond yield conundrum; bond demand; subprime crisis; structural breaks

JEL Classification: C22; G01; G12; E43

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From 2002 to mid-2007 when the US subprime crisis broke out US bond yields were at unusually low levels. Before mid-2004 these levels could be explained by the greater stability of ‘fundamentals’ and low short term interest rates (the ‘great moderation’), but the persistence of these low yields after that point in time was puzzling. Financial markets expected long term rates to rise in tandem with the rise in the federal funds rate as was the case in previous periods of monetary tightening. This did not happen. On the contrary, not only did long term rates not rise they actually continued to fall[1] (see Figure 1). As Alan Greenspan, the then Chairman of the Federal Reserve, stated before Congress in June 2005: “Among the biggest surprises of the past year has been the pronounced decline in long-term interest rates on U.S. Treasury securities despite a 2-percentage-point increase in the federal funds rate. This is clearly without recent precedent. … Moreover, even after the recent backup in credit risk spreads, yields for … corporate bonds have declined even more than Treasuries over the same period.” (Greenspan 2005, p.1).

What caused this ‘bond yield conundrum’? Considering that its appearance coincided with a marked upswing in investor demand for US bonds (see Figure 2) it is possible that a considerable part of the downward pressure on US bond yields stemmed from that demand (Bernanke et al. 2011). To verify this possibility, a number of empirical studies have focused specifically on the impact of foreign government demand for US Treasuries on long term Treasury yields. Foreign official investor demand began to increase after February 1994 when China devalued its currency, but the rate of increase in that demand accelerated even more sharply after 2003 as many emerging market economy governments sought to preserve part of their increasing commodity revenues and export surpluses in

[1] In June 2005 the long term rate was 73 basis points lower than it was one year before. In December 2006 the rate was still slightly lower, although the federal fund rate was 425 basis points higher than it was 2 ½ years earlier and expected to stay relatively stable above the 4% level until 2015 (Kozicki and Sellon 2005).

safe stores of value. While some studies found no evidence of a long term demand impact on Treasury yields (e.g. ECB 2006; Rudebusch et al. 2006), the majority of recent studies have found evidence of a negative impact, albeit that the estimated size of the impact varied from study to study (e.g. Idier et al. 2007; Bandholz et al. 2009; Craine and Martin 2009; Warnock and Cacdac Warnock 2009).

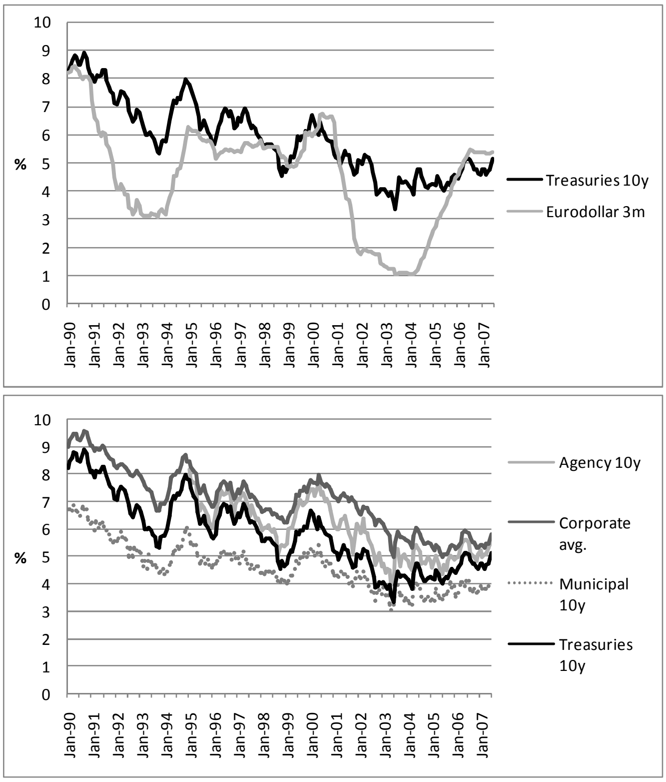

FIGURE 1. LONG AND SHORT TERM INTEREST RATES IN THE US

Notes: The top plot compares the 3-month Eurodollar rate with the 10-year Treasury yield. The bottom plot demonstrates the downward movement of traditional long-term bond yields in the US (Source: Bloomberg 2010, FR Statistical Release H.15 2010).

In this paper we assess the impact of investor demand on long term Treasury yields using an autoregressive distributed lag (ARDL) based econometric model. Since it has become well established that the increase in demand for US bonds stemmed not only from foreign official investors but also from private foreign (mainly European) and domestic investors (Bernanke et al. 2011) we consider the impact of all of these sources of demands on yields. Further, given that the bond yield conundrum applied as much to the other major US bond markets as to the Treasury market, one would have expected an analysis of the impact of demand on long term yields in these other markets. As there has been no such analysis[2], this paper seeks to fill this gap by modeling the impact of investor demand on US agency, and AAA-rated corporate and municipal bond yields using ARDL-based models.

In our view, an important by-product of this econometric assessment of the contribution of demand to the US bond yield conundrum is that it may help to resolve the question as to why the US collateralized debt obligation (CDO) market was allowed to grow in a very short time to a size that was able to trigger widespread financial panic when this market suddenly collapsed in August 2007. The conventional answer to this question is one that places the major burden of responsibility on the US financial system itself. However, there is a minority view that, while the US banks and their associates cannot be absolved from blame in accelerating the rate of production of CDOs right up to mid-2007, the major driving force behind that acceleration was the pressure of demand for US safe assets spilling over from other major US debt security markets (e.g. Caballero and Krishnamurthy 2009; Gros 2009; Lysandrou 2009; Caballero 2010). Clearly, this alternative view, and its ensuing policy implications, would command far more attention were it to be convincingly demonstrated that the rise in foreign and

[2] To our knowledge nearly all existing studies on the conundrum concentrate on the demand from foreign official sources on long term Treasury yields. Exceptions in this regard are: ECB (2006) who test the impact of foreign official purchases on corporate bond yields and agency bond yields (without presenting their models in detail), Xiao and Xiao (2009) who test for the impact of pension funds on the yields of Treasuries and investment grade corporate bonds (without accounting for the demand from foreign sources and other domestic private investors), and Warnock and Cacdac Warnock (2009) who test if the increase in Treasury purchases from foreign sources had a negative effect on the yield of corporate bonds and mortgage rates.

domestic demand for US bonds in the period leading up to the outbreak of the subprime crisis did indeed have a substantial significant negative impact on all highly rated traditional fixed income products.

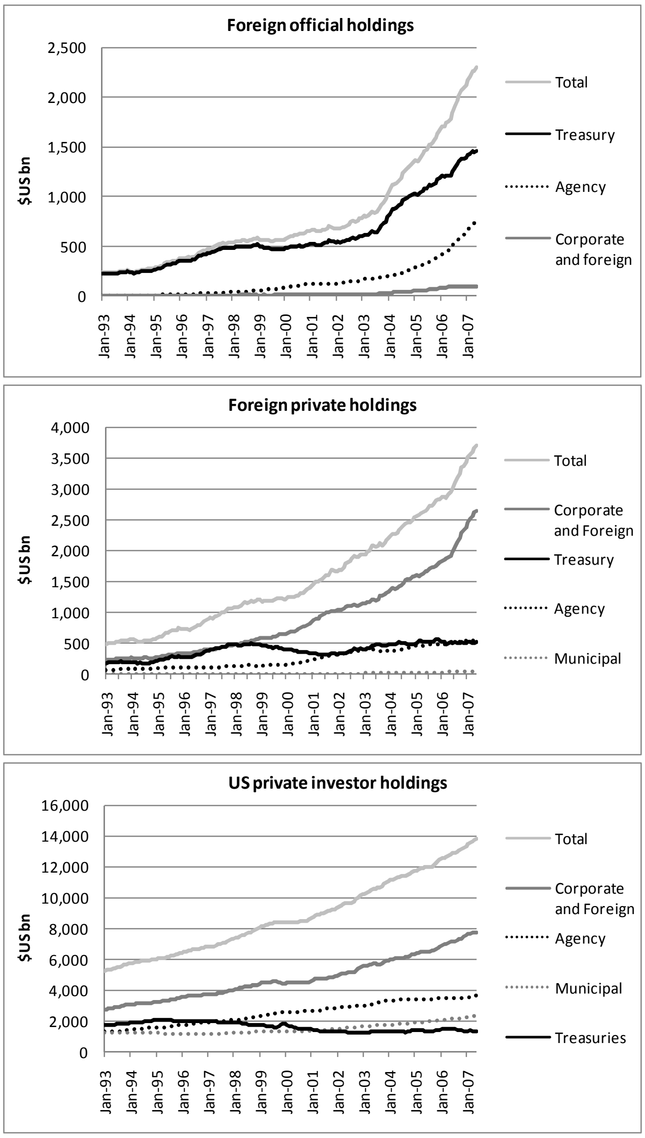

FIGURE 2. US BOND HOLDINGS FROM FOREIGN AND PRIVATE DOMESTIC INVESTORS

Notes: The plots show the US bond holdings of foreign governments (top), foreign private investors (middle) and domestic private investors (bottom), respectively (Source: FR Statistical Release Z.1 2010, Treasury International Capital System 2010).

The layout of this paper is as follows. Section one gives details of model specification, the data used and the chosen sample period. Section two presents and discusses the estimation results while section three briefly comments on their policy implications. Section four concludes.

I. Model specification and model selection

A. Rationale for the models

Any attempt to quantify the impact of demand on bond yields has to begin with a specification of all of the major determinants of yields. According to recent research (see e.g. Rudebusch et al. 2006; Wu 2008) these determinants broadly divide into two groups, those relating to macroeconomic essentials on the one hand and those relating to financial risk on the other. Apart from the short term interest rate, which is usually expected to influence nominal long term yields, inflation and the business cycle are also believed to be important determinants of these yields (see e.g. Bandholz et al. 2009; Warnock and Cacdac Warnock 2009).

Changes in actual inflation can influence expectations about the real value of future coupon payments, the future federal funds rate and long term inflation rates, while changes in long term inflation expectations influence expectations about future short term interest rates and the real par value at maturity. Growth expectations possibly influence long term interest rates because in a boom market participants often expect inflationary pressure and a rise in the federal funds rate to prevent an overheating of the economy and thus request higher yields, and vice versa. As stated, beyond these macroeconomic indicators changes in default risk and volatility can also influence the long term yield (see e.g. Rudebusch et al. 2006). A decrease in the volatility of bond yields, for example, decreases the risk for market participants and thus is expected to lower the yield.

The inclusion of bond demand as a possible determinant of bond yields is not uncontroversial. Investor demand should have no impact on yields in a world where financial markets are frictionless and all assets classes are perfect substitutes (ECB 2006). However, we shall consider the alternative position that financial markets are not frictionless and that bonds have certain distinct properties that enable them to meet investors' needs in ways that other asset classes cannot (for a clear exposition of bond characteristics and their attraction for investors see Krishnamurthy and Vissing-Jorgensen 2007, and Greenwood and Vayanos 2010). In sum, our model can be represented by the following equation:

$$y^l = f(i^s, \pi, \pi^e, y^e, rp, d) \tag{1}$$

where y^l denotes the long term interest rate, i^s the short term interest rate, π current inflation, π^e inflation expectations, y^e growth expectations and rp is a risk premium for the expected default risk and macroeconomic and financial volatility, while *d* denotes investor demand for bonds.

Given that most of these variables are non-stationary according to unreported augmented Dickey-Fuller (ADF) unit root tests[3] stationary vector autoregression (VAR) and vector error correction models (VECM) are preferable to a single equation ordinary least squares (OLS) specification to assess their impact on bond yields. However, existing macro-finance models of the term structure that use VAR and New Keynesian based modeling strategies (see Rudebusch et al. 2006, Eijffinger et al. 2010, Rudebusch 2010) have been criticized because of their no-arbitrage assumption, the difficulty to optimize the likelihood function, the

[3] Even though the ADF test has low power, inspection of data plots and knowledge of the data suggest that most of our variables are intrinsically I(1). Exceptions are the (log of the) ISM-Index, the MOVE Index, and the corporate bond holdings ratio from US individuals, which are stationary according to ADF tests.

overfitting of risk, and the implied homoskedastic yields (Eijffinger et al. 2010)[4]. Moreover, we find it impractical to apply the Johansen test for cointegration due to the large number of variables to be considered in the equilibrium and the large lag lengths required for monthly data (which make the degrees of freedom far too small). Consequently, we use four ARDL models in (unrestricted) error correction mechanism form to test which of the above stated determinants were mainly responsible for the low long term yields of AAA-rated US bonds. The general form of our models is:

$$(2) \qquad \Delta y_t^l = \beta_0 + \sum_{i=0}^{p} \gamma_{1i} \Delta X_{1t-i} + \cdots + \sum_{i=0}^{p} \gamma_{Ki} \Delta X_{Kt-i} + \sum_{i=1}^{p} \alpha_i \Delta y_{t-i}^l + \alpha_0 \Delta y_{t-1}^l + \sum_{k=1}^{K} \beta_k X_{kt-1} + u_t$$

This modeling approach takes into account current and lagged differenced variables to measure short run effects and lagged level variables to account for long run effects, and it allows us to include all of the above stated determinants without losing too many degrees of freedom. Another important advantage of this modeling technique is that, in contrast to VECMs, it produces consistent estimates of the long run coefficients independently of their order of integration (Pesaran and Shin 1995)[5]. This is important in our application given that unit root tests suggest a mixture of I(1) and I(0) variables in our model. We apply Pesaran et al.'s (2001) bounds testing procedure (that corrects for weak endogeneity of regressors) to assess whether the variables in the models cointegrate.

[4] To our knowledge macro-finance models of the term structure of interest rates were not able to solve the bond yield conundrum. For a discussion of macro-finance models of the term structure of interest rates see Diebold et al. (2005) and Rudebusch et al. (2006). See Kim (2007) for a thorough discussion of the shortcomings of this approach.

[5] The ARDL form that we adopt has additional advantages. First, it possesses small sample power dominance in terms of testing cointegration over Engle and Granger type tests and, second, the model corrects for any weak endogeneity of regressors – see, for example, Shin et al. (2011). A further point is that we can simultaneously estimate and test structural breaks in both the long run and short run components of the ARDL model in a simple manner.

B. Data

For each of the above listed determinants a proxy is chosen that is either the same as or similar to that used in previous studies. Considering that most of the relevant data is not available on a daily or weekly basis, monthly data are utilized to ensure sufficient degrees of freedom[6]. As proxies for US long term interest rates of highly rated fixed income securities we take the 10-year Treasury yield (retrieved from the Federal Reserve Statistical Release H.15), the 10-year agency bond yield, and the average yield of Moody's bond index for AAA-rated corporate bonds and for AAA-rated 10-year municipal bonds (all retrieved from Bloomberg).

To account for changes in the US short term interest rate we include the 3-month rate for Eurodollar deposits in London (Federal Reserve Statistical Release H.15). Following Warnock and Cacdac Warnock (2009), we see the Eurodollar rate as a preferable measure for changes in current monetary policy inasmuch as it varies more than the federal funds rate. As a proxy for current and expected inflation we include the trimmed personal consumption expenditure (PCE) deflator, following Bandholz et al. (2009), and the ten year consumer price index (CPI) inflation expectations, as in Warnock and Cacdac Warnock (2009) – data are taken from the Survey of Professional Forecasters and the Philadelphia Fed respectively.

To capture the state of the business cycle, the purchasing manager index from the Manufacturing Survey of the Institute for Supply Management (ISM) is used, as in Bandholz et al. (2009). This is because "[f]inancial market participants have anxiously anticipated the ISM ever since Alan Greenspan once claimed … that he

[6] While monthly series do exist for most of the data some are only available on a quarterly basis and are therefore interpolated to monthly frequency with the "cubic match last" method, which is readily available in EViews. The variables which have been interpolated are: 10-year inflation expectations, domestic bond holdings, and the data on outstanding bonds (with the exception of Treasuries). The available data on the expected deficit-to-GDP ratio are only published twice each year by the CBO and are therefore also interpolated, in line with Warnock and Cacdac Warnock (2009).

placed great emphasis on this report" (Trainer (2006, p. 211)). When the ISM-Index is relatively high (> 50) market participants expect high growth figures and when the ISM-Index is relatively low (< 43) a recession is anticipated. As a proxy for changes in the stock market – which is seen as a good indicator of the business cycle and for shifts in portfolio preferences (Idier et al. 2007) – we employ the Dow Jones Index (retrieved from yahoo finance).

We use the following data to measure changes in default risk perceptions, financial market volatility and macroeconomic uncertainty. Default risk is captured by using data about expected fiscal policy, which is measured by 5-year-ahead deficit-to-GDP expectations as in Laubach (2009)[7] (retrieved from the CBO Budget and Economic Outlooks) and the expected default risk of AAA-rated corporate bonds. The latter is proxied by the Expected Default Frequency (EDF) for AAA-rated corporate bonds (kindly provided by Moody's Analytics UK) as in Krishnamurthy and Vissing-Jorgensen (2007). Analogous to Rudebusch et al. (2006) data from the Merrill Lynch Option Volatility Estimate (MOVE) Index, retrieved from Bloomberg, are used to account for financial market volatility[8]. Furthermore, the 24-month rolling standard deviation of the Eurodollar rate, the Chicago Board Options Exchange Volatility (VIX) Index (retrieved from Bloomberg), and two measures for macroeconomic uncertainty (the 24-month rolling standard deviation of the ISM-Index and of the ten year CPI inflation expectations) are tested for significance, similar to Rudebusch et al. (2006).

To measure the influence of changes in investor demand on bond yields, private and foreign official holdings as a ratio of total outstanding bonds are taken into account, as in Rudebusch et al. (2006) and Bandholz et al. (2009). The

[7] It can be difficult to measure the impact of the actual deficit-to-GDP ratio because automatic stabilizers lead to an increase of deficit levels in recessionary periods, while monetary easing can at the same time be expected to lower the long term yield. Laubach (2009) has therefore proposed using expected deficit-to-GDP ratios as these are more likely to mirror investor's expectations which are important in regard to long term yields.

[8] For non-Treasury bonds the significance of the 24-month rolling standard deviations of changes in the long term yields are tested, similar to Warnock and Cacdac Warnock's (2009) approach, but these proxies are insignificant.

holdings ratio is preferable to mere flow or stock figures because demand pressure can be expected to take place only when investors increase their holdings disproportionally to newly available bonds (i.e. if their holdings ratio increases). The data for changes in the holdings from US banking institutions, US individuals and US institutional investors are retrieved from the Flow of Funds statistics[9]. The data for foreign official and foreign private long term holdings are taken from the Treasury International Capital Reporting System (TIC)[10] because the Flow of Fund statistics do not distinguish between official and private holdings. The amount of total outstanding bonds is retrieved from the Flow of Funds tables and from the Treasury Bulletins (outstanding notes, bonds and TIPS).

C. Sample Period

Most previous studies take the mid-1980s to mid-2000s as their sample period. In contrast, we limit our sample period to that spanning February 1994 to June 2007 (with the exception of the agency bond model where the yield data are only available from 1995 onwards). February 1994 has been chosen as the starting point because the data on foreign official holdings show a structural break at that time – presumably the break can be attributed to the devaluation of the Renminbi from 5.8 ¥/$ to 8.7 ¥/$ between December 1993 and January 1994. Another reason is provided by Thornton (2007) who argues that the relationship between the federal funds rate and long term interest rates changed much earlier than in mid-2000 due to a change of the Federal Open Market Committee's (FOMC) policies in 1988 towards using the federal funds rate as a policy target. Although he identifies 1988 as the break-point, he presents evidence that a structural break might also have occurred in 1994 when the FOMC started to release policy

[9] Federal Reserve Statistical Release: Tables L209, L210, L211 and L212.

[10] Holdings in the TIC data are only reported semi-annually. Therefore, estimations from the Fed about monthly changes in holdings are used. The source for these data is: http://federalreserve.gov/pubs/ifdp/2007/910/ticdata.zip.

statements after its meetings. This change has influenced expectations to a significant degree (Bernanke et al. 2004).

D. Model selection

Treasury yield model.—First, we model the 10-year Treasury yield based on the variables described above. Due to the multitude of potential variables that could be included, a model was constructed which incorporated contemporaneous differenced and level proxies of variables that were significant in the models of previous studies. Thus:

$$\begin{aligned}(3)\quad \Delta y_t^l = {} & \beta_0 + \beta_1(\Delta FO_t) + \beta_2(\Delta FP_t) + \beta_3(\Delta P_t) + \beta_4(\Delta i_t) + \beta_5(\Delta lism_t) + \\ & \beta_6(\Delta \pi_t) + \beta_7(\Delta \pi_t^{10}) + \beta_8(\Delta dow_t) + \beta_9(\Delta move_t) + \beta_{10}(\Delta def_t^5) + \\ & \beta_{11}\left(y_{t-1}^l\right) + \beta_{12}(FO_{t-1}) + \beta_{13}(FP_{t-1}) + \beta_{14}(P_{t-1}) + \beta_{15}(i_{t-1}) + \\ & \beta_{16}(lism_{t-1}) + \beta_{17}(\pi_{t-1}) + \beta_{18}(\pi_{t-1}^{10}) + \beta_{19}(dow_{t-1}) + \beta_{20}(move_{t-1}) + \\ & \beta_{21}(def_{t-1}^5)\end{aligned}$$

where $_t$ indicates the current period, $_{t-1}$ denotes a one month lag, Δ is the difference operator, y^l is the nominal 10-year Treasury yield, *FO* are foreign official holdings as a ratio of total outstanding long term Treasuries, *FP* are foreign private holdings as a ratio of total outstanding long term Treasuries, *P* are US pension funds holdings as a ratio of total outstanding long term Treasuries, *i* is the 3-month Eurodollar rate, *lism* is the log of the ISM-Index, π is the actual PCE inflation rate, π^{10} are 10-year CPI inflation expectations, *dow* is the value of the Dow Jones Index, *move* is the MOVE Index, and def^5 are 5-year deficit-to-GDP expectations.

However, when estimated this model suffered from autocorrelation suggesting that the yield might be influenced by some differenced variables with a time lag.

The monthly frequency of our data suggests consideration of up to twelve lags of each of the variables. However, all twelve differences of all of the variables in (3) could not be included simultaneously. We therefore added the twelve lagged differences of just one variable in (3) and, based on an F-test, excluded the jointly insignificant lags of the differences of this variable. This was repeated in turn for each of the variables in (3), including the dependent variable, until a model that included only significant lags of the differences of all variables was obtained. Finally, all level variables which were not significant at the 5% level were removed from the model.

Variable addition tests were then conducted on the following variables not included in (3): VIX Index and the 24-month rolling standard deviation of the Eurodollar rate, ISM-Index, and ten year CPI inflation – the first lagged levels and twelve lagged differences being considered for each factor. However, all of these variables are jointly insignificant at the 5% level, which is in line with the results of Rudebusch et al. (2006, p. 25) who find that from the volatility variables "[t]he most significant and robust explanatory variable is the implied volatility on longer-term Treasuries." (i.e. the MOVE Index). The resulting model (reported as (i) in Table 2 in the Results section) shows no evident misspecification at the 5%-level in terms of autocorrelation (lags 1...12), non-normally distributed residuals and heteroscedasticity [Arch (lags 1...12) and White tests]. According to Ramsey's Reset test the appropriate functional form is linear and the Wu-Hausman test indicates that all contemporaneous variables are weakly exogenous. Further, the bounds test (with unrestricted intercept) – critical values are taken from Pesaran et al. (2001) – confirms that the level variables are mutually cointegrated irrespective of whether the regressors are I(0) or I(1)[11].

[11] The F-test applied with unrestricted intercept deletes all lagged level terms (but not the intercept) from the model – the number of lagged level terms (excluding y^l_{t-1}) determines the degrees of freedom. For the F-test and t-test the critical

However, unreported CUSUM and CUSUM of Squares Test indicate a structural break. This is in line with the findings of ECB (2006), which reports a structural change in 1999, and Wu (2005) who finds a structural break between 2000 and 2002. Therefore, a Quandt-Andrews breakpoint test (35% trimming) was undertaken. According to this test, the maximum likelihood for a break is in November 1998, although a break is only indicated at the 10% level. However, a Chow breakpoint test finds that a break occurred in November 1998 at the 5% level (Table 1, first column).We note that according to the Chow test no structural break occurred in June 2004, when the conundrum period started. Considering all of these results and those of past studies we believe that it is reasonable to consider the possibility of a break in November 1998.

TABLE 1 — RESULTS BREAKPOINT TESTS

	Treasury model	Corporate model	Agency model	Municipal model
Quandt-Andrews unknown breakpoint test				
Max Likelihood Ratio F-statistic prob.	0.062	0.005	0.000	0.001
Max date	1998:11	1999:02	2001:04	2001:04
Chow breakpoint test				
F-stat. prob. at Quandt-Andrews max date	0.018	0.003	0.000	0.001
F-stat. prob. 2004:06	0.873	0.849	0.186	0.742

Notes: This table shows the results of the Quandt-Andrews unknown breakpoint tests (35% trimmed data, probabilities calculated using Hansen's (1997) method) and the Chow breakpoint tests for all our models. The presented figures are F-statistic probabilities and dates.

To model the structural break, shift variables for all the significant independent variables were created with the value zero before the break and the original value of the variable after the break. All of these shift variables were jointly included in the model. The jointly insignificant variables were subsequently excluded (first the shift variables and then the non-shift variables) to obtain the final parsimonious model. This model (reported as (ii) in Table 2) has a superior fit to the model without a break, no misspecification is evident and its level variables

values corresponding to the I(1) bound are reported in the table because breaching these values confirms cointegration regardless of the variables' order of integration.

are mutually cointegrated[12]. Further, the CUSUM and CUSUM of Squares Test indicate no other structural break after November 1998 (the Quandt-Andrews test cannot be effectively applied in this model because of the shift variables).

Agency, corporate and municipal yield models.—The model selection procedure for the other bond models is essentially the same as that for the Treasury yield model. In addition to the macroeconomic and risk variables that are significant in the Treasury model, the 24-month rolling standard deviation of changes in the long rates for each bond class and the EDF for AAA-rated corporate bonds were tested for significance (again with lags 1...12 for the differenced variables). Furthermore, we also controlled for an increase in foreign and domestic investor demand for each bond class[13]. Having established parsimonious models for agency, corporate and municipal bond yields, breakpoint tests were carried out. In line with the Treasury model, these tests indicated a structural break for each bond class (Table 1). Hence, for each model shift variables were tested for their significance in line with the above described procedure.

The resulting parsimonious models show no evident misspecification, the level variables are mutually cointegrated (reported in Table 3 in the Results section) and, in particular, the CUSUM and CUSUM of Squares Test indicate no further structural breaks. All of our favored models for inference include shift variables

[12] Because the shift variables are related to the non-shift variables the degrees of freedom for the cointegration test are uncertain. One could, for example, either treat the shift and non-shift components of a particular variable as one covariate or two separate variables for calculating degrees of freedom. Following Shin et al. (2011) we consider critical values using degrees of freedom calculated in both of these ways, thereby forming further upper and lower bounds of the test for the already existing upper and lower bounds (related to uncertainty over the variables' orders of integration). If the F-statistic (t-ratio) exceeds (is below) the critical value's bound for I(1) processes treating shift and non-shift components of a variable as one (two) covariate(s) there is unambiguous evidence of cointegration and we use these criteria in our application. We extrapolate some of the critical values reported in Pesaran et al (2001) when the number of variables used to calculate the degrees of freedom exceed 10. We also note that the use of this cointegration test in a model allowing for structural breaks represents one of the novelties of this paper.

[13] Only those investor groups that had significant holdings in June 2007 (i.e. only investor groups with a holdings ratio of above 1%) and that increased their holdings ratio in the respective bond class during the conundrum period were included in each model.

and are discussed in the next section. Due to space limitations only the models that account for the structural break are presented for the agency, corporate and municipal bond yields (all of these models have a superior fit compared to those without a break). The long run solutions for our favored parsimonious dynamic models (with breaks) are reported in Table 4 in the Results section (the Appendix discusses how the equilibrium coefficients and their corresponding standard errors are obtained).

II. Results

A. Treasury yield model

The results of the Treasury model confirm previous findings that an increase in the demand from foreign governments had a negative impact on the long term Treasury yield (Table 2). According to our favored model for inference, model (ii), an increase in foreign government demand had a consistently negative impact on the 10-year US Treasury yield throughout the whole sample period in the short and long run. That is, ceteris paribus, an increase of the foreign official holdings ratio by 1% point had a negative impact on the yield of around 9 basis points (bp) in the long run. This magnitude is similar to the 7 bp impact of total foreign holdings that Bandholz et al. (2009) found in their VECM model. Foreign private investors also had a negative impact in the long run before November 1998 but their impact became insignificant thereafter. The most likely explanation for this change is that although between August 1994 and November 1998 the holdings ratio of foreign private investors increased steadily (by a total of 11% points), after the latter date it began to decline (for example, it declined by 3.5% points in the conundrum period June 2004 to June 2007). Hence, private investors put no further demand pressure on the yield in the post break period.

All the control variables have the expected signs and reasonable magnitudes. The short term interest rate has a positive impact in both the short run and the long run, but after November 1998 this impact becomes much smaller in both cases. This finding supports Greenwald and Stiglitz’s (2003) argument that financial innovation fostered a decoupling of long term interest rates from short term rates. To be specific, we find that, ceteris paribus, before November 1998 a 1% point increase in the short term interest rate leads to a 45 bp increase in the Treasury yield in the long run, with this impact declining to 11 bp after this date. These magnitudes are in line with other studies, e.g. Warnock and Cacdac Warnock (2009) who find that the impact is 37 bp (but who do not consider a possible shift in the relationship between short term and long term interest rates).

Higher growth expectations are also found to lead to an increase in the Treasury yield, but here again the impact becomes smaller after the break: thus, ceteris paribus, in the conundrum period a 1% increase of the ISM Index raised the yield by about 2.5 bp. This result is similar to Bandholz et al. (2009) who report an impact of about 2 bp. In contrast, the long run impact of inflation, stock prices and the volatility of Treasuries on the yield remains unchanged throughout the whole period. Ceteris paribus, a 1% point rise in the PCE deflator increases the yield by 94 bp, a 1000 point increase in the Dow Jones Index raises the yield by 45 bp (in line with Idier et al. 2007) who find that a 1% increase in stock returns has an impact of 42 bp) and an increase of the MOVE Index by 10 points increases the yield by 7 bp in the long term.

TABLE 2 — PARSIMONIOUS MODEL OF THE NOMINAL 10-YEAR TREASURY YIELD

	(i) without break		(ii) with break	
Δ(FOROFFICIAL)	-0.2174***	(-6.44)	-0.2155***	(-6.81)
Δ(FOROFFICIAL(-1))	-0.1273***	(-3.58)	-0.1325***	(-4.11)
Δ(EUR_DOL)	0.3279***	(3.42)	0.7202***	(4.38)
Δ(EURDOL)$^{s11/98}$			-0.5256***	(-2.98)
Δ(EURDOL(-1))	-0.2459***	(-2.62)	-0.1630*	(-1.78)
Δ(LOGISM)	0.9202**	(2.31)	1.0200***	(2.65)
Δ(LOGISM(-1))	1.1464***	(3.17)	1.6376***	(2.86)
Δ(LOGISM(-1))$^{s11/98}$			-1.2539*	(-1.80)
Δ(LOGISM(-4))	1.0097***	(2.83)	0.8844***	(2.66)
Δ(PCE)	0.5404***	(2.72)	0.5403***	(2.90)
Δ(PCE(-9))	-0.6486***	(-2.98)	-0.6432***	(-3.12)
Δ(DOW)	0.0001**	(2.36)	0.0001**	(3.14)
Δ(DOW(-1))	0.0001**	(2.50)		
YIELD(-1)	-0.2835***	(-6.13)	-0.3795***	(-6.63)
FOROFFICIAL(-1)	-0.0224***	(-3.46)	-0.0358***	(-4.67)
FORPRIVATE(-1)	-0.0252**	(-2.07)	-0.0909***	(-6.50)
FORPRIVATE(-1)$^{s11/98}$			0.0924***	(3.62)
EURDOL(-1)	0.0535***	(3.36)	0.1700***	(3.45)
EURDOL(-1)$^{s11/98}$			-0.1277***	(-2.73)
LOGISM(-1)	0.8877***	(2.91)	1.2634***	(4.69)
LOGISM(-1)$^{s11/98}$			0.3038**	(-2.29)
PCE(-1)	0.2431**	(2.28)	0.3578***	(3.28)
CPI10Y(-1)	0.3855**	(2.05)		
DOW(-1)	0.0001***	(3.53)	0.0002***	(5.88)
MOVE(-1)	0.0021**	(2.27)	0.0027***	(2.78)
adj. R-squared	0.58		0.64	
Schwarz criterion	-0.47		-0.54	
Sample: 1994:02 to 2007:06 (161 observations)				

(iii) equilibrium long-run effects of (ii)		
before the break		
FOROFFICIAL	-0.0944***	(-7.14)
FORPRIVATE	-0.2396***	(-5.82
EURDOL	0.4478***	(4.50)
LOGISM	3.3286***	(5.52)
PCE	0.9426***	(3.72)
DOW	0.0005***	(5.56)
MOVE	0.0070***	(2.88)
after the break		
FOROFFICIAL	-0.0944***	(-7.14)
FORPRIVATE	0.0038	(0.07)
EURDOL	0.1113***	(2.85)
LOGISM	2.5283***	(3.61)
PCE	0.9426***	(3.72)
DOW	0.0005***	(5.56)
MOVE	0.0070***	(2.88)

misspecification/cointegration tests	(i)	(ii)
BG(2) prob.	0.16	0.24
BG(12) prob.	0.25	0.36
Jarque-Bera prob.	0.44	0.26
Arch(1) prob.	0.90	0.56
Arch(12) prob.	0.56	0.49
White prob.	0.34	0.61
Ramsey LR prob.	0.87	0.15
Wu-Hausm. prob.	0.85	0.58
Bounds test F-stat.	6.44***	8.20***
Bounds test t-stat.	-6.13***	-6.63***

Notes: This table summarizes the results of our ARDL-model for the nominal 10-year Treasury yield. Where Δ is the difference operator, the number of lags are indicated in parentheses as a suffix to a variable's name, $s^{11/98}$ indicates the shift component of a variable and the date of the structural break (i.e. after November 1998), *YIELD* is the 10-year nominal Treasury yield, *FOROFFICIAL* are foreign official holdings as a ratio of total outstanding long-term Treasuries, *FORPRIVATE* are foreign private holdings as a ratio of total outstanding long-term Treasuries, *EURDOL* is the 3-month Eurodollar rate, *LOGISM* is the log of the ISM-Index, *PCE* is the actual PCE inflation rate, *CPI10Y* are 10-year CPI inflation expectations, *DOW* is the value of the Dow Jones Index, and *MOVE* is the Merrill Lynch Option Volatility Estimate Index. Intercepts are not reported but are included in the models. In each column coefficients and t-statistics (in parenthesis) are reported. Probability values for all misspecification tests are reported in the section headed misspecification/cointegration tests, where BG(x) denotes the probability value of the Breusch-Godfrey test for x order correlation and Arch(x) the probability value of the ARCH heteroskedasticity test with x lags. The 5% critical values for the bounds cointegration test with unrestricted intercept and no trend are (i) F=3.39, t=-4.72, (ii) F=3.50, t=-5.03 [(i) k=8, (ii) k=10 (t), k=7 (F)] – see Pesaran et al. (2001). The significance of a coefficient or test statistic at the 1%, 5% and 10% level of significance is indicated by ***, ** and *, respectively.

In order to make these results more palpable and identify which of the variables included in the Treasury model (ii) were responsible for the 'bond yield conundrum' the marginal cumulative impact (MCI[14]) of each of these variables on the Treasury yield is used. June 2004 to June 2007 is chosen as the reference period for this exercise because it spans the beginning of US monetary tightening and the subsequent debate on the 'bond yield conundrum'. The MCI of each variable depends on the coefficients (including the changes due to the break where applicable) of the differenced and lagged level variables and on the changes in the data of the variable. Thus, the formula for calculating the MCI for each month is:

$$(4) \qquad Impact_{\gamma t} = \beta_{\gamma 1}\Delta\gamma_t + \beta_{\gamma 1}^{shift}\Delta\gamma_t + \cdots + \beta_{\gamma 12}\Delta\gamma_{t-12} + \beta_{\gamma 12}^{shift}\Delta\gamma_{t-12} + \beta_{\gamma 13}\gamma_{t-1} + \beta_{\gamma 13}^{shift}\gamma_{t-1}$$

$$(5) \qquad MCI_{\gamma t} = Impact_{\gamma t} - Impact_{\gamma 2004:05}$$

Figure 3 shows that foreign official demand has the largest negative MCI on the yield in the reference period, which can therefore be seen as mainly responsible for the conundrum, while foreign private demand by contrast had virtually no impact in this period. Our model's finding that the increase in foreign official Treasury holdings depressed the yield by as much as 60 bp during the conundrum period is similar to previous findings: Bandholz et al. (2009) report an impact of 70 bp between 2003 and 2006, Craine and Martin (2009) one of 80 bp between 2004 and 2006, and Warnock and Cacdac Warnock (2009) one of 80 bp between 1984 and May 2005[15]. In addition to foreign official demand, pessimistic

[14] The MCI is the difference in a particular variable's contribution to the yield in any particular period relative to a reference point (in our case May 2004).

[15] These reported impacts are of course influenced by the chosen reference point. If February 1994 is taken as the starting point foreign official demand will be found to have depressed the 10-year Treasury yield by as much as 128 bp in the conundrum period. However, if January 2003 is taken as the starting point the size of the impact is 70 bp, exactly the amount reported in Bandholz et al. (2009).

expectations about the business cycle (ISM Index) and a decrease of the implied yield volatility (MOVE Index) also had a negative impact on the Treasury yield of about 20 bp each and therefore also partly explain the conundrum. Counteracting these factors were the increases in short term interest rates and in core price inflation, both of which had a small positive impact of about 20 bp, and the rise in stock prices, which had a relatively larger positive impact of almost 60 bp.

According to the implied yield of our favored model for inference, which fits the actual Treasury yield remarkably well during the conundrum period, these forces seem to explain the conundrum fully (see Figure 4 for the yield residuals)[16]. Thus, our model improves upon existing Treasury bond models. For example, ECB (2006), Rudebusch et al. (2006), Warnock and Cacdac Warnock (2009), Eijffinger et al. (2010), and Rudebusch (2010) all report that their models overestimate the long term Treasury yield after June 2004, while Bandholz et al.'s (2009) model overvalues the yield throughout the year 2005[17].

[16] The residuals of the yield have been calculated as follows: *actual yield – fitted yield* (where *fitted yield = fitted Δyield + actual yield*$_{t-1}$).

[17] Not all existing studies report their model residuals, see e.g. Idier et al. (2007), and Craine and Martin (2009).

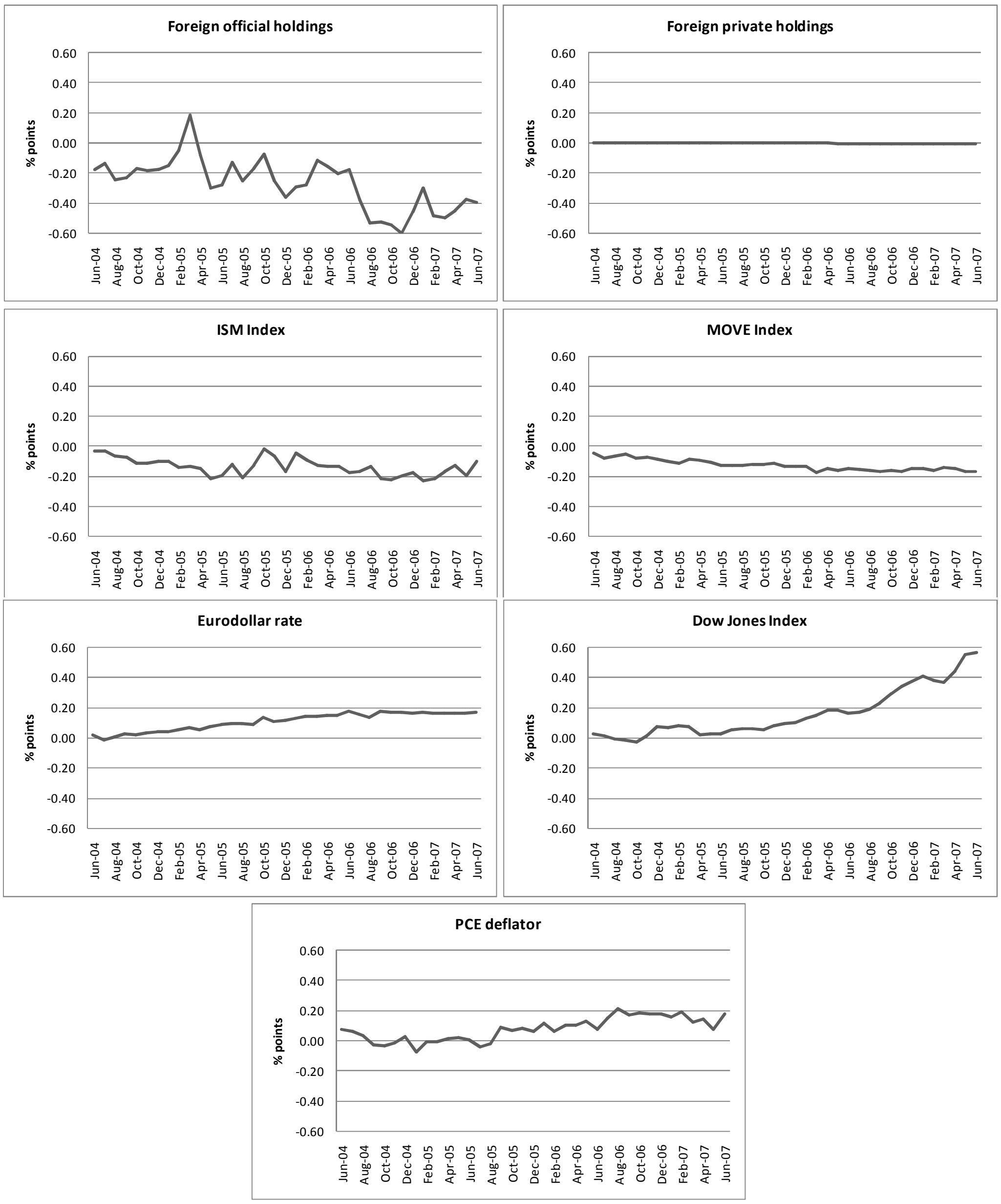

FIGURE 3. VARIABLES' MCIS FOR THE NOMINAL 10-YEAR TREASURY YIELD

Notes: These plots show the marginal cumulative impact of each variable on the nominal 10-year Treasury yield for each month during the conundrum period, according to the results of our Treasury bond model (ii) (see Table 2).

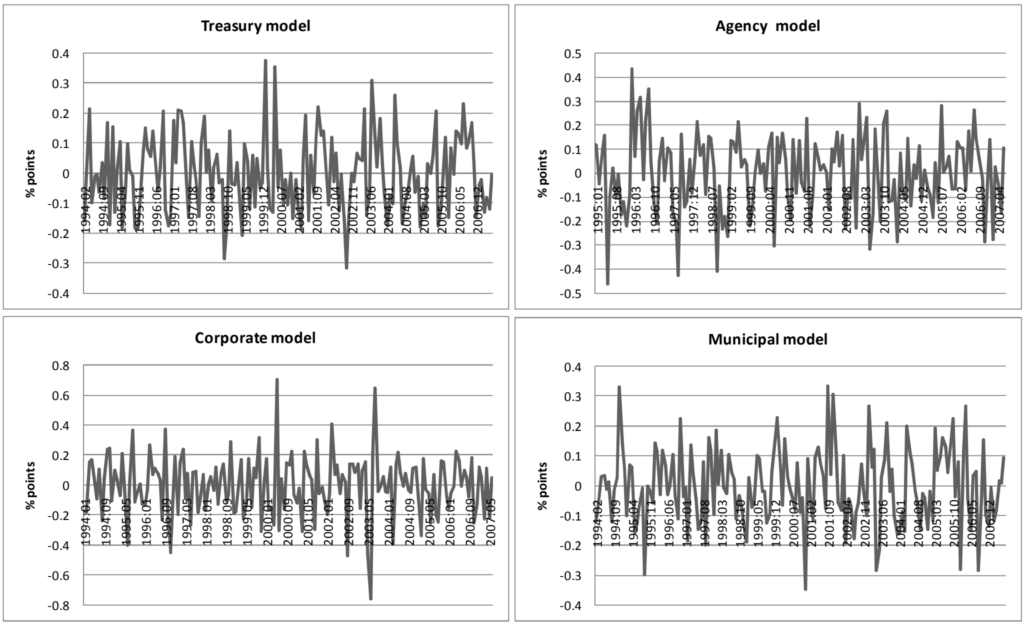

FIGURE 4. YIELD RESIDUALS

Notes: These plots show how well the implied yield values of the respective models fit the respective long-term bond yields.

The reason why our model appears to explain the Treasury yield conundrum better than previous models most likely lies in our different modelling strategy. In contrast to the previous literature, we consider more variables in our model (whilst accounting for non-stationarity) and we model the evident structural break. Indeed, Rudebusch et al. (2006), Rudebusch (2010) and Eijffinger (2010) use a VECM model that does not directly include foreign official demand[18], which our model found to be the most important variable in explaining the conundrum. Furthermore, the above authors do not take into account the

[18] Rudebusch et al. (2006) test if foreign official demand is correlated with the error term of their model, and find no correlation (they use custodial data from the New York Fed (FRBNY) as a proxy for foreign official holdings; this seems not be the best proxy because "… some foreign governments avoid the FRBNY and thus this source is best described as only partial" (Warnock and Cacdac Warnock 2009, p. 905). However, this finding does not imply that the model results would be the same if the variable is fully incorporated in the model.

possibility that the impact of the short term interest rate on the 10-year Treasury yield during the conundrum period was smaller than before November 1998.

The incorporation of this possibility in our model also seems to provide a major explanation of why it fits the yield better than do the models of Bandholz et al. (2009) and Warnock and Cacdac Warnock (2009). These authors' models attribute a higher impact than our model does to the short term interest rate during the conundrum period (with long run coefficients of 0.37 and 0.33, respectively). An additional point is that these authors' studies appear to overestimate the yield either because they do not include a measure for interest rate volatility (Bandholz et al.) or because they use the rolling standard deviation of long yields to proxy the volatility of yields (Warnock and Cacdac Warnock) – in contrast to the MOVE Index, the rolling standard deviation does not indicate a decline in volatility during the conundrum period.

B. Agency, corporate and municipal bond yield models

The results of the agency, corporate and municipal yield models clearly indicate that investor demand also played a major role in explaining the low long term yields of non-Treasury AAA-rated bonds (Table 3 and Table 4). In line with the Treasury yield model, these models fit the data well in the conundrum period (see Figure 4), and all control variables have the expected signs and reasonable magnitudes. However, in some cases the magnitudes differ significantly. Next to noise, the most likely explanation for this observation is that investors do not see these different bond classes as perfect substitutes and therefore ask for different adjustments in prices when conditions are changing. Indeed Previous studies confirm that investors value different bond classes differently even while they may carry the same credit rating (see e.g. Krishnamurthy and Vissing-Jorgenson (2007)).

TABLE 3 — PARSIMONIOUS MODEL OF THE NOMINAL LONG TERM YIELDS OF AAA-RATED NON-TREASURY US SECURITIES

(i) Agency			(ii) Corporate			(iii) Municipal		
Δ(FOROFFICIAL)	-1.7414***	(-5.68)	Δ(YIELD(-1))	0.0956	(1.61)	Δ(YIELD(-1))$^{s04/01}$	0.4644***	(4.57)
Δ(FORPRIVATE)	-0.4600***	(-3.45)	Δ(FORPRIVATE)	-0.3983***	(-9.59)	Δ(YIELD(-1))$^{s04/01}$	0.3334***	(3.59)
Δ(USINDIVIDUALS)	-0.1321***	(-2.98)	Δ(FORPRIVATE(-1))	-0.2464***	(-5.10)	Δ(YIELD(-3))$^{s04/01}$	0.2833***	(3.22)
Δ(EURDOL)	0.4803***	(3.80)	Δ(US INDIVIDUAL(-1))	-0.1792***	(-5.90)	Δ(YIELD(-4))	0.2293***	(3.70)
Δ(LOGISM)	1.4678***	(3.11)	Δ(USBANK(-1))	-0.3478***	(-4.28)	Δ(YIELD(-5))	0.2166***	(3.62)
Δ(PCE)	0.6020**	(2.50)	Δ(EURDOL(-1))	-0.1658***	(-2.59)	Δ(EURDOL)	0.7231***	(8.42)
Δ(PCE(-2))$^{s04/01}$	-0.8739***	(-2.82)	Δ(EURDOL(-8))$^{s02/99}$	-0.3045***	(-4.24)	Δ(EURDOL(-2))$^{s04/01}$	-0.3469**	(-2.34)
Δ(DOW)$^{s04/01}$	0.0003***	(4.95)	Δ(EURDOL(-11))	0.1275**	(2.24)	Δ(EURDOL(-8))$^{s04/01}$	-0.3734***	(-2.87)
Δ(MOVE)$^{s04/01}$	0.0094***	(4.95)	Δ(LOGISM)	0.6734**	(2.39)	Δ(PCE)	0.9009***	(5.03)
YIELD(-1)	-0.4101***	(-7.87)	Δ(LOGISM(-1))	1.5553***	(3.82)	Δ(DOW)s04/01	0.0002***	(3.85)
FOROFFICIAL(-1)	-0.4626***	(-3.27)	Δ(LOGISM(-1))$^{s02/99}$	-1.4720***	(-3.03)	Δ(DOW(-5))	0.0001***	(2.61)
FOROFFICIAL(-1)$^{s04/01}$	0.4027***	(3.32)	Δ(PCE)	0.4381***	(3.29)	Δ(MOVE)	0.0033***	(3.19)
FORPRIVATE(-1)	-0.2168***	(-3.11)	Δ(PCE(-9))	-0.4191***	(-2.91)	Δ(MOVE(-2))	-0.0027***	(-2.96)
USINDIVIDUAL(-1)	-0.0514***	(-3.54)	Δ(DOW)	0.0001***	(3.21)	YIELD10(-1)	-0.5913***	(-8.75)
USPENSION(-1)$^{s04/01}$	-0.1441***	(-3.07)	YIELD(-1)	-0.2273***	(-6.56)	YIELD(-1)$^{s04/01}$	-0.4765***	(-6.14)
EURDOL(-1)	0.2218***	(5.02)	FORPRIVATE(-1)	-0.2113***	(-5.37)	FOREIGN(-1)	-4.9958***	(-6.30)
LOGISM(-1)	1.3153***	(4.34)	FORPRIVATE(-1)$^{s02/99}$	0.1615***	(4.50)	FOREIGN(-1)$^{s04/01}$	4.1318***	(6.05)
PCE(-1)	0.8260***	(6.30)	EURDOL(-1)	0.0664***	(4.83)	USINDIVIDUAL(-1)	-0.1010***	(-4.27)
DOW(-1)	0.0002***	(4.53)	LOGISM(-1)	0.6125***	(3.51)	USINSURANCE(-1)	-0.0747***	(-2.60)
MOVE(-1)	0.0036**	(2.24)	LOGISM(-1)$^{s02/99}$	-0.5380***	(-4.33)	USBANK(-1)	-0.2470**	(-2.07)
MOVE(-1)$^{s04/01}$	0.0061***	(2.80)	PCE(-1)	0.1906***	(2.88)	EURDOL(-1)	0.2054***	(6.09)
adj. R-squared	0.63		CPI10Y(-1)	0.3002**	(2.29)	LOGISM(-1)	0.6388***	(2.73)
Schwarz criterion	-0.09		DOW(-1)	0.0001***	(5.61)	PCE(-1)	0.3050***	(3.66)
Sample: 1995:01 to 2007:06 (150 obs.)			MOVE(-1)	0.0016**	(2.34)	DOW(-1)	0.0001***	(3.71)
			EDFAAA(-1)	3.0898***	(5.60)	MOVE(-1)	0.0055***	(4.82)
			adj. R-squared	0.71		adj. R-squared	0.57	
			Schwarz criterion	-1.19		Schwarz criterion	-0.61	
			Sample: 1994:02 to 2007:06 (161 obs.)			Sample: 1994:02 to 2007:06 (161 obs.)		

Results misspecification/cointegration tests
BG(2) prob.: (i) 0.89, (ii) 0.65, (iii) 0.23 BG(12) prob.: (i) 0.26, (ii) 0.15, (iii) 0.10 Jarque-Bera prob.: (i) 0.44, (ii) 0.99, (iii) 0.54
Arch(1) prob.: (i) 0.61, (ii) 0.86, (iii) 0.41 Arch(12) prob.: (i) 0.56, (ii) 0.15, (iii) 0.87 White prob.: (i) 0.47, (ii) 0.31, (iii) 0.06
Ramsey LR prob.: (i) 0.16, (ii) 0.26, (iii) 0.23 Wu-Hausman Prob.: F-stat. (i) 0.46, (ii) 0.55, (iii) 0.86
Bounds test: F-stat. (i) 8.68***, (ii) 10.41***, (iii) 10.39***; t-stat. (i) -7.87***, (ii) -6.56***, (iii) -9.85***

Notes: This table summarizes the results of our ARDL-models for the nominal 10-year US agency, and AAA-rated corporate and municipal bond yields, respectively. The table notes are the same as in Table 2, with the following exceptions: $s^{x/x}$ indicates the shift component of a variable with the date of the structural break indicated by *x/x* (i.e. after February 1999 and after April 2001), *YIELD* is the 10-year nominal yield of the respective bond class, *FOROFFICIAL* are foreign official holdings as a ratio of total outstanding bonds (i.e. the holdings ratio) of the respective bond class, *FORPRIVATE* is the foreign private holdings ratio of the respective bond class, *FOREIGN* is the foreign holdings ratio of municipal bonds, *USBANK* is the US banking institutions holdings ratio of the respective bond class, *USINDIVIDUAL* is the US individual holdings ratio of the respective bond class, *USINSURANCE* is the US insurance companies holdings ratio of the respective bond class, *USPENSION* is the US pension funds holdings ratio of the respective bond class, and *EDFAAA* is Moody's expected default frequency for AAA-rated corporate bonds. The 5% critical values for a Bounds cointegration test with unrestricted intercept and no trend are (i) F=3.30, t≈-5.20 , (ii) F=3.39, t=-5.03, (iii) F=3.24, t≈-5.20 [(i) k=11 (t), k=9 (F) (ii) k=10 (t), k=8 (F) (iii) k=10 (F), k=11 (t)] – see Pesaran et al. (2001).

TABLE 4 — EQUILIBRIUM LONG RUN IMPACTS ON THE NOMINAL LONG TERM YIELDS OF AAA-RATED NON-TREASURY US SECURITIES

(i) Agency bond yield			(ii) Corporate bond yield			(iii) Municipal bond yield		
before the break			before the break			before the break		
FOROFFICIAL	-1.1282***	(-3.82)	FORPRIVATE	-0.9298***	(-5.39)	FOREIGN	-8.4493***	(-5.64)
FORPRIVATE	-0.5286***	(-3.43)	EURDOL	0.2923***	(5.18)	USINDIVIDUAL	-0.1709***	(-4.33)
USINDIVIDUAL	-0.1253***	(-3.93)	LOGISM	2.6953***	(3.24)	USINSURANCE	-0.1263***	(-2.83)
USPENSION			PCE	0.8387***	(3.15)	USBANK	-0.4178**	(-2.08)
EURDOL	0.5410***	(6.83)	CPI10Y	1.3207**	(2.42)	EURDOL	0.3473***	(8.47)
LOGISM	3.2074***	(4.80)	DOW	0.0005***	(5.56)	LOGISM	1.0805***	(3.17)
PCE	2.0143***	(7.82)	MOVE	0.0071**	(2.51)	PCE	0.5159***	(3.74)
DOW	0.0004***	(4.53)	EDFAAA	13.5957***	(6.33)	DOW	0.0002***	(3.36)
MOVE	0.0088**	(2.20)				MOVE	0.0093***	(4.63)
after the break			after the break			after the break		
FOROFFICIAL	-0.1462*	(-1.73)	FORPRIVATE	-0.2193***	(-5.63)	FOREIGN	-0.8091***	(-7.00)
FORPRIVATE	-0.5286***	(-3.43)	EURDOL	0.2923***	(5.18)	USINDIVIDUAL	-0.0946***	(-5.45)
USINDIVIDUAL	-0.1253***	(-3.93)	LOGISM	0.3282	(0.43)	USINSURANCE	-0.0699***	(-2.92)
USPENSION	-0.3514***	(-3.10)	PCE	0.8387***	(3.15)	USBANK	-0.2314**	(-2.18)
EURDOL	0.5410***	(6.83)	CPI10Y	1.3207**	(2.42)	EURDOL	0.1923***	(7.97)
LOGISM	3.2074***	(4.80)	DOW	0.0005***	(5.56)	LOGISM	0.5983***	(2.88)
PCE	2.0143***	(7.82)	MOVE	0.0071**	(2.51)	PCE	0.2857***	(3.94)
DOW	0.0004***	(4.53)	EDFAAA	13.5957***	(6.33)	DOW	0.0001***	(4.07)
MOVE	0.0237***	(5.91)				MOVE	0.0051***	(4.58)

Notes: This table summarizes the equilibrium results of our ARDL-models for the nominal 10-year US agency, and AAA-rated corporate and municipal bond yields, respectively. The table notes are the same as in Table 2 and Table 3.

Agency bond yield.—Foreign official demand had a negative impact not only on the Treasury yield but also on the agency bond yield both in the short run (Table 3, column 1) and in the long run (Table 4, column 1). However, in contrast to the Treasury yield model, the long run magnitude of the impact declines after the break. This said, it appears that, ceteris paribus, from April 2001 onwards an increase in the holdings ratio by 1% point still reduced the yield by around 15 bp in the long run. The most probable explanation for the shift in the variable's coefficient is the change in the foreign official holdings ratio: while this ratio increased moderately in the pre break period, it increased considerably in the post break period (from 3% in April 2001 to 11% in June 2007). This development in turn helps to explain why market reactions to increases in foreign official holdings in the post break period were comparatively modest relative to the pre-break period given that there was now less scope for price increases (yield decreases) per unit of increase in the foreign official holdings ratio.

By contrast, the impact of foreign private and US individual investors on the long term agency bond yield remained stable throughout the whole sample period. Each 1% point increase in the foreign private holdings ratio led to a decline in the yield of around 53 bp in the long run, while the same increase in the domestic individual holdings ratio lowered the yield by 13 bp. US pension funds only had an impact on the yield after the break: from April 2001 onwards the yield was depressed by 35 bp for each 1% point increase in the pension funds' holdings ratio. The shift in this variable took place because domestic pension funds only increased their holdings ratio significantly in the post break period. Possible explanations as to why the magnitudes of the coefficients of these three investor groups were so different are that they reacted differently to expected changes in the agency yield or that they had different expectations of future yields. US individual investors, for example, might have increased their holdings to a lesser extent than foreign private investors and pension funds when they (rightly) expected the agency yield to decrease and hence put less additional pressure on yields than their counterparts.

The MCI suggests that investor demand was also the main reason for the low long term agency yield during the conundrum period (Figure 5). This is especially true for foreign official investors who, according to our model, depressed the yield by as much as 107 bp. However, it is the case that private foreign and domestic investor demand also helped to reduce the yield, by around 39 bp and 26 bp respectively. This downward pressure on yields, further fuelled by pessimistic expectations about the business cycle and a lowering of the implied yield volatility, was mainly offset by the rise in the short term interest rate and by the increases in stock prices and in core price inflation.

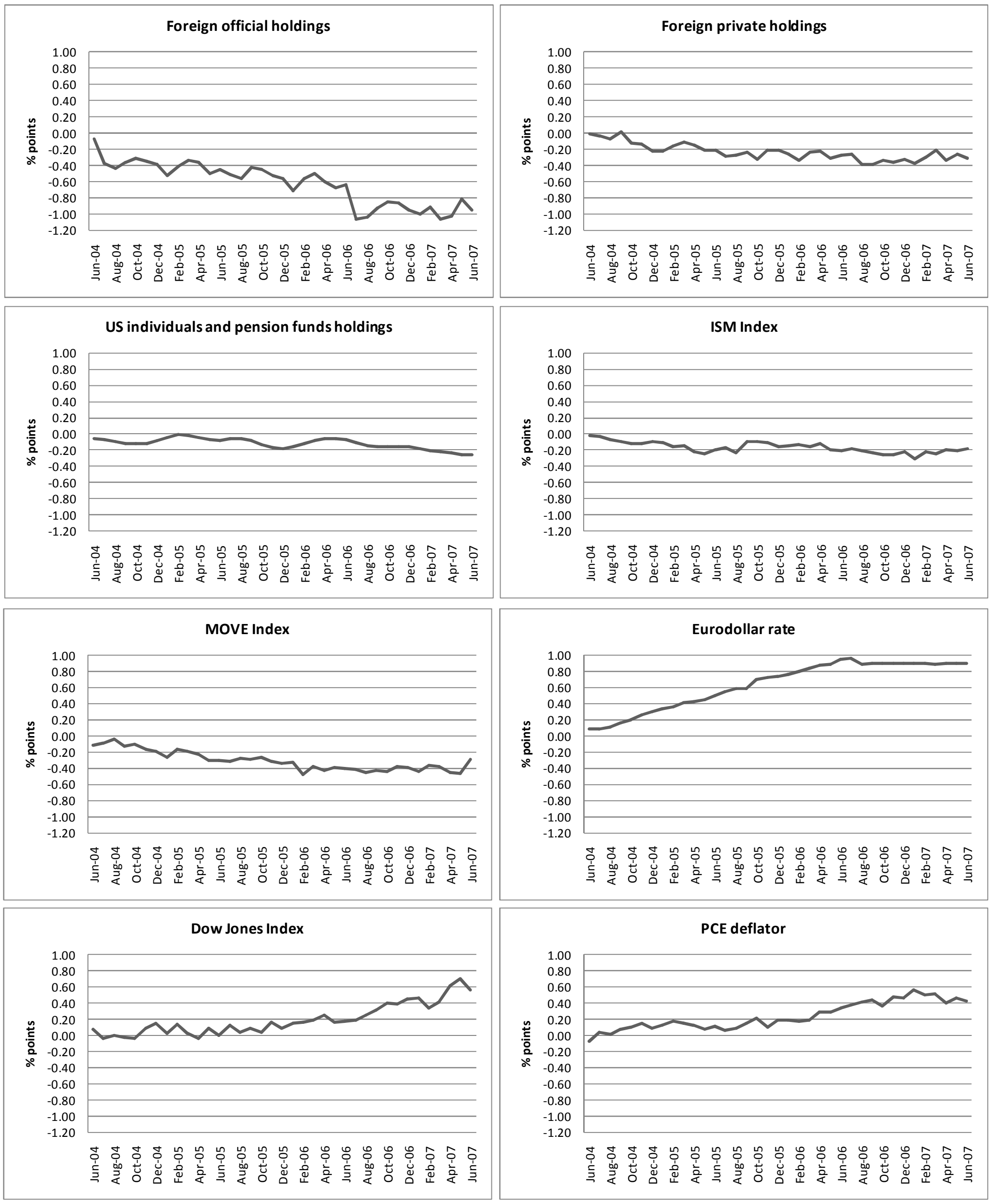

FIGURE 5. VARIABLES' MCIS FOR THE NOMINAL 10-YEAR AGENCY BOND YIELD

Notes: These plots show the marginal cumulative impact of each variable on the nominal 10-year Agency yield for each month during the conundrum period, according to the results of our agency bond model (see Table 3).

Corporate bond yield.—Foreign private investors invested heavily in the corporate bond market between 1994 and mid-2007, their holdings ratio more than doubling (from 11% to 24.5%) during this period, with the result that they put significant downward pressure on AAA-rated corporate bond yields in the short run (Table 3) and in the long run (Table 4). Regarding the long run, in the post break period an increase in the foreign private investors' holdings ratio by 1% point led to a decrease of the yield by about 22 bp (compared to a 93 bp reduction prior to the break). The explanation for the shift in the variable's coefficient is probably the same as that regarding foreign official holdings in the agency bond model inasmuch as the increase in the holdings ratio of foreign private investors mainly took place after the break, in this case after February 1999. Domestic investors also had some negative impact on the yield when they increased their holdings ratio, although only in the short run.

Our proxy for default risk of AAA-rated corporate bonds (EDFAAA)[19] has the expected sign and a reasonable magnitude (as we will see below). A puzzling result is that an increase in growth expectations is not significant in the long run after the break – in the Treasury model the impact of the ISM Index is lower after the break, though it remains highly significant. A possible reason is that in the post break period the increase in investor demand for corporate bonds (which are more attractive in an upswing) and the request for higher yields (due to expected inflationary pressure and an expected rise in the federal funds rate) offset each other when the ISM Index increased and vice versa.

[19] Corporate bonds are the only AAA-rated traditional fixed income asset class which is not directly (Treasuries and municipal bonds) or indirectly (agency bonds) backed by a governmental organization.

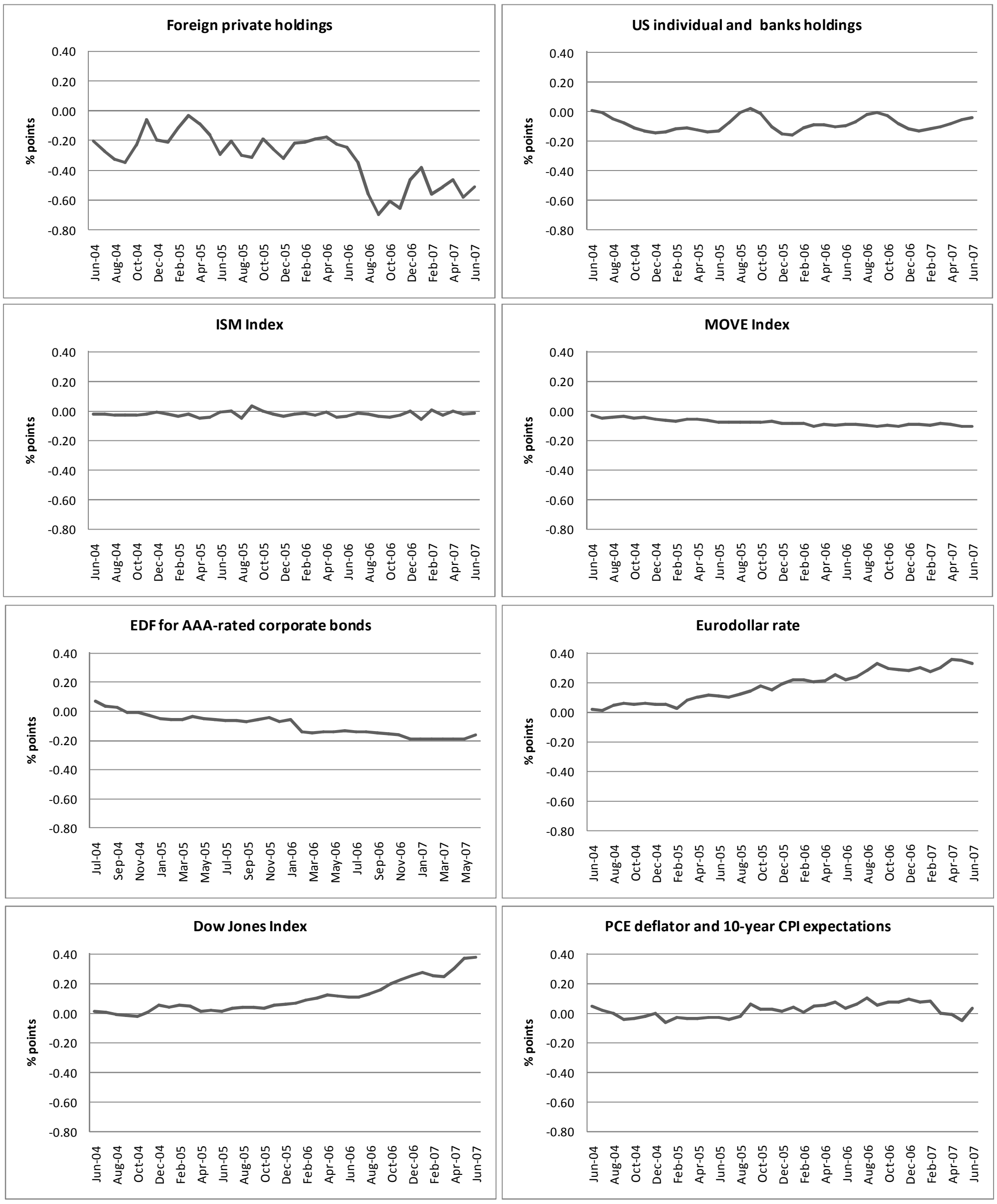

FIGURE 6. VARIABLES' MCIS FOR THE NOMINAL AAA-RATED CORPORATE BOND YIELD

Notes: These plots show the marginal cumulative impact of each variable on the nominal AAA-rated corporate bond yield for each month during the conundrum period, according to the results of our corporate yield model (see Table 3).

Once again, the MCI shows that investor demand was the main suppressing force in the conundrum period (Figure 6). Between June 2004 and June 2007 the yield of AAA-rated corporate bonds was lowered by as much as 69 bp due to demand pressure from foreign private investors, and by as much as 15 bp due to higher demand from domestic investors. Lower yield volatility and a lower default risk for AAA-rated corporate bonds added to this pressure. The main counteracting forces were increases in the Eurodollar rate and increases in stock market prices.

Municipal bond yield.—Finally, an increase in foreign demand for 10-year AAA-rated municipal bonds also had a negative impact on their yield, albeit that it is not clear from the available data whether this demand came mainly from foreign official sources or from foreign private sources. It appears that after the break (April 2001) a 1% point increase in the holdings ratio of foreigners decreased the municipal bond yield by 81 bp in the long run (Table 4). The magnitude of the foreign demand coefficient is smaller after the break, probably for the same reasons that applied to the agency bond market case: market reactions to increases in foreign holdings in the post break period were comparatively more muted given that there was now less scope for price increases (yield decreases) per unit of increase in the foreign holdings ratio. Domestic individual investors, banks and insurance companies also appeared to put downward pressure on the municipal bond yield when they increased their holdings ratios, albeit that the magnitudes of these demand coefficients differ (these differences possibly stemming from differences in expectations or in the reactions to expectations as previously argued).

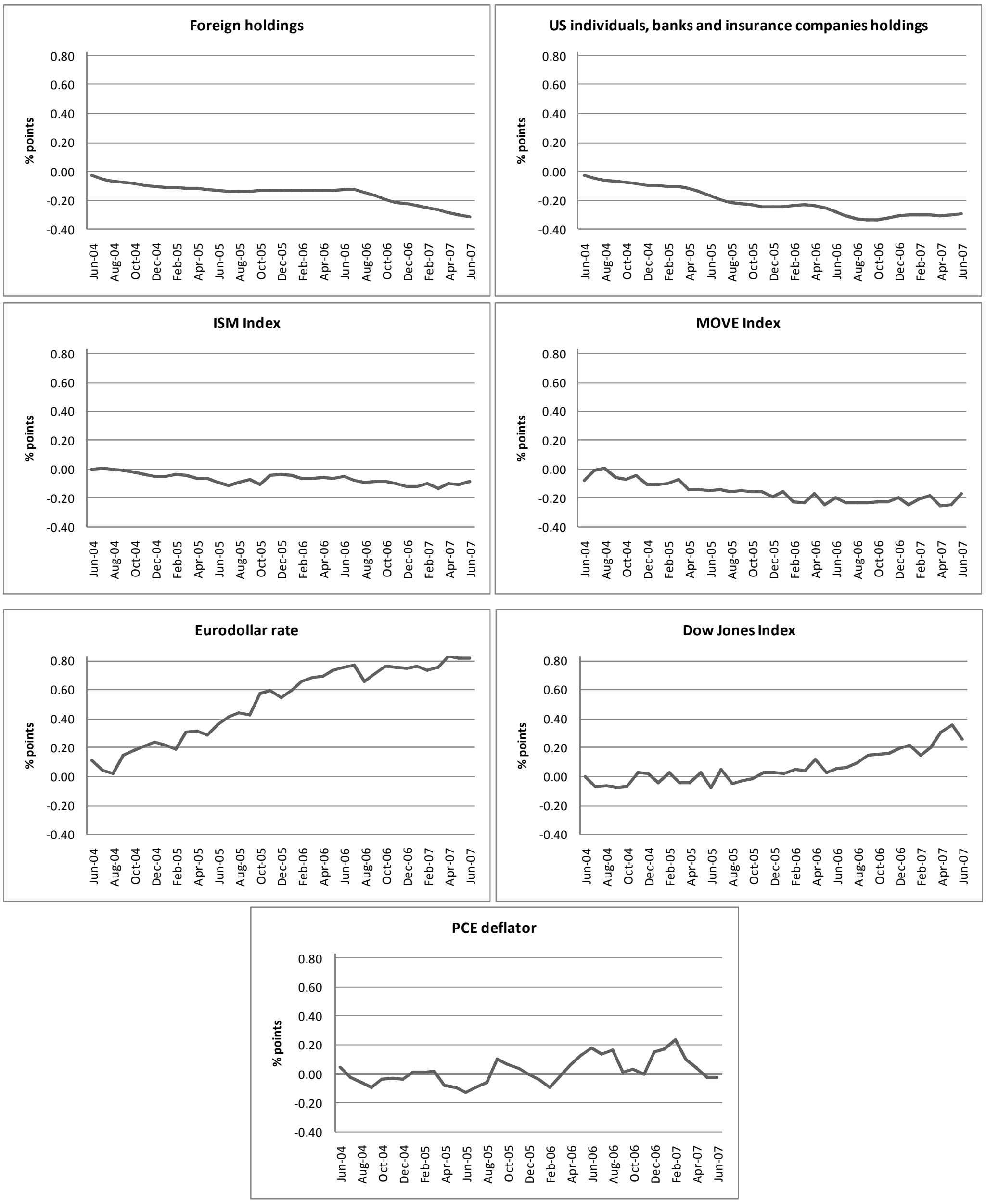

FIGURE 7. VARIABLES' MCIS FOR THE NOMINAL AAA-RATED MUNICIPAL BOND YIELD

Notes: These plots show the marginal cumulative impact of each variable on the nominal 10-year municipal bond yield for each month during the conundrum period, according to the results of our municipal yield model (see Table 3).

The findings for the different variables' MCIs are similar those reported previously (Figure 7), the one main difference being that the low municipal bond yield in the conundrum period seems to be primarily caused by domestic investors who lowered the yield by as much as 34 bp while foreign investors lowered it by no more than 31 bp. This finding is in keeping with the fact that foreign investors do not benefit from the tax advantages of municipal bonds as do domestic investors and are therefore much less active in the municipal bond market. [20] Lower growth expectations and interest rate volatility appear to have added to the downward demand pressure, while increases in the short term interest rate and in stock market prices acted as counter forces.

III. Investor demand and subprime crisis

The US bond yield conundrum has generated much discussion regarding its magnitude and the factors behind it for good reason. As Wu (2008) has argued: "The correct understanding and quantification of the conundrum have direct implications for monetary policy…" (p. 2). While we certainly agree with this argument we also believe that a 'correct understanding and quantification of the conundrum' as manifested in all of the major US bond markets - and not merely in the market for Treasuries – can help to shed more light on the root causes of the recent financial crisis and, in so doing, help guide policy makers in their attempts to prevent a similar crisis on this scale in the future. The logic behind this position is straightforward.

The securities at the epicenter of the financial crisis which broke out in the summer of 2007 were CDOs. The estimated amount of CDOs in 2002 was about $\$^1/_4$ trillion and yet by the time of the crisis that figure had multiplied twelvefold

[20] "Short-maturity municipal yields are equal to the Treasury yield multiplied by one minus the income tax rate, and the ratio between municipal and Treasury yields decreases with maturity." (Ang et al. 2010, p. 566). Hence, the average yield of municipal bonds is normally lower than that of Treasury, corporate and agency bonds with the same maturity.

to about $3 trillion with the bulk of it comprising of triple AAA rated tranches (Blundell-Wignall (2007)). One of the unresolved questions regarding this rapid increase in the CDO market concerns the precise role played by investor demand. Did this demand play a merely passive role? Yields in the other debt securities markets were unusually low in the immediate pre-crisis period and so investors would have been happy to accept the higher yielding CDOs, but was the quest for fees and commissions on the part of the banks and their associates the more important driving force behind the rapid acceleration in CDO production? Or did investor demand play a more active role in the growth of the CDO market? The US financial institutions may have profited handsomely from the creation and distribution of CDOs but were these institutions also under enormous external pressure to do all of this in order to make up for the shortfall in the supply of other US safe assets?

If the answer to the above question is that investor demand did indeed play a secondary role in CDO growth then it is entirely correct for policy makers to concentrate their efforts on rectifying the various institutional and regulatory errors and failures that allowed the US banking system to create the toxic debt securities on so large a scale in such a short time span. However, this policy approach would not on its own prevent future financial crises if it turned out that the demand for extra safe assets was in fact the more important driver behind CDO growth, as recent research from Caballero and Krishnamurthy (2009) suggests. From this alternative, demand-side, perspective on CDO growth “the core policy problem to deal with is how to bridge the safe asset gap without over-exposing the financial sector to systemic risk.” (Caballero, 2010, p. 6). Thus, imposing various new rules and restrictions on the US financial sector’s ability to create debt assets will not only “...not help to deal with the structural problem of excess safe-asset demand.” but will also have the opposite effect of worsening the safe asset gap, the potential “...cost of this policy distortion [being] stronger headwinds for the recovery and the risk that the same pattern of systemically-vulnerable safe-asset creation may migrate to somewhere else in the world that is even less prepared to absorb the systemic risk." (ibid, p. 6-7).

Caballero's take on the major policy lessons of the subprime crisis remains a minority one and a possible reason for this is that to date there has been no comprehensive attempt at econometrically testing the strength of foreign and domestic demand for US safe assets in the pre-crisis era. The crux of the matter is that CDOs are essentially 'second-floor' debt securities, securities backed by securities. Thus, for the demand-pull version of the CDO growth story to be really credible, it has to be convincingly demonstrated that the pressure of aggregate demand for safe stores of value was so great that the combined capacity of all the US 'ground floor' debt securities markets (those for corporate and municipal securities in addition to that for Treasuries) and of the US 'first floor' securities markets (those for agency and other asset backed securities) was simply not large enough to fully accommodate that pressure. We believe that the econometric results generated in this paper amount to such a demonstration insofar as they consistently point to significant and substantial downward demand pressure on all US bond yields in the pre-crisis period.

IV. Conclusion

Our models fully explain the US bond yield conundrum of 2004 to 2007 as found not only in relation to US Treasuries but also in relation to all of the other traditional AAA-rated US debt securities, something that has not been achieved in the previous literature. We attribute this result to the incorporation of a broader set of variables than is usual in our models, this being made possible by the adoption of the ARDL approach, and to the allowance for evident structural change around the time of the millennium (the latter confirming findings of previous authors). It is especially noteworthy that demand variables are found to be the most prominent factor in explaining the unusually low US bond yields during the conundrum period. These findings have substantial policy implications in that they provide strong support for the hitherto underexplored hypothesis that excess safe asset demand on the part of investors rather than excess greed on the part of the banks was the chief force that drove the expansion of the US CDO market well beyond what was prudent.

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Appendix. — approximate equilibrium coefficient standard errors in an (unrestricted) ARDL model with a one-time structural break

Consider the general (unrestricted) ARDL model with error-correction form that allows the coefficients to change at one particular point in time:

(A1) $$\Delta Y_t = \alpha_0 Y_{t-1} + \beta_0 + \beta_1 X_{1t-1} + \cdots + \beta_K X_{Kt-1} + \alpha_0^S (D_t \times Y_{t-1}) + \beta_0^S D_t + \beta_1^S (D_t \times X_{1t-1}) + \cdots + \beta_K^S (D_t \times X_{Kt-1}) + \sum_{i=1}^{p} \alpha_i \Delta Y_{t-i} + \sum_{i=0}^{p} \gamma_{1i} \Delta X_{1t-i} + \cdots + \sum_{i=0}^{p} \gamma_{Ki} \Delta X_{Kt-i} + \sum_{i=1}^{p} \alpha_i^S (D_t \times \Delta Y_{t-i}) + \sum_{i=0}^{p} \gamma_{1i}^S (D_t \times \Delta X_{1t-i}) + \cdots + \sum_{i=0}^{p} \gamma_{Ki}^S (D_t \times \Delta X_{Kt-i}) + u_t$$

where, $D_t = 0$ in the pre-break period and $D_t = 1$ on and after the break point period.

Letting $\boldsymbol{Z}_t' = (Y_t, X_{1t}, \dots, X_{Kt})$, the static equilibrium of (A1) can be derived by applying $\boldsymbol{Z}' = \boldsymbol{Z}_t' = \cdots = \boldsymbol{Z}_{t-p}'$ and recognising that $u_t = 0$ in equilibrium, thus:

(A2) $$Y = -\left(\frac{\beta_0+\beta_0^S D_t}{\alpha_0+\alpha_0^S D_t}\right) - \left(\frac{\beta_1+\beta_1^S D_t}{\alpha_0+\alpha_0^S D_t}\right) X_1 - \cdots - \left(\frac{\beta_K+\beta_K^S D_t}{\alpha_0+\alpha_0^S D_t}\right) X_K$$

The equilibrium in the pre-break and post-break periods are obtained by substituting $D_t = 0$ and $D_t = 1$, respectively, into (A2), thus:

(A3) $Y = \beta_0^* + \beta_1^* X_1 + \cdots + \beta_K^* X_K$, where, $\beta_k^* = -\left(\frac{\beta_k}{\alpha_0}\right), k = 0, 1, \ldots, K$

(A4) $Y = \beta_0^{S*} + \beta_1^{S*} X_1 + \cdots + \beta_K^{S*} X_K$, where, $\beta_k^{S*} = -\left(\frac{\beta_k+\beta_k^S}{\alpha_0+\alpha_0^S}\right), k = 0, 1, \ldots, K$

Since the equilibrium coefficients involve quotients of the coefficients in (A1) we use the formula given in De Boef and Keele (2008) to approximate the standard errors of the equilibrium coefficients, denoted $s_{\beta_k^*}$ and $s_{\beta_k^{S*}}$ for the pre- and post-break periods, respectively. Since the pre-equilibrium coefficients can be obtained from the post-equilibrium coefficients by substituting $\alpha_0^S = \beta_k^S = 0$ into the latter we specify this approximation only for $s_{\beta_k^{S*}}$. That is,

(A5) $$s_{\beta_k^{S*}} = \sqrt{Var(\beta_k^{S*})} = \sqrt{Var\left(\frac{\beta_k+\beta_k^S}{-\alpha_0-\alpha_0^S}\right)}$$

$$= \sqrt{\frac{1}{(-\alpha_0-\alpha_0^S)^2} Var(\beta_k + \beta_k^S) + \frac{(\beta_k+\beta_k^S)^2}{(-\alpha_0-\alpha_0^S)^4} Var(-\alpha_0 - \alpha_0^S) - 2\frac{(\beta_k+\beta_k^S)}{(-\alpha_0-\alpha_0^S)^3} Cov[(\beta_k + \beta_k^S)(-\alpha_0 - \alpha_0^S)]}$$