Did the Asian and Russian/LTCM Financial Crises A®ect U.S. Treasury Liquidity Premia?

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ABSTRACT

Have the emerging market ¯nancial crises of recent years had a signi¯cant impact on the dynamics of the U.S. Treasury bond market? This paper applies the recursive break test procedure of Leybourne et al. using weighted-symmetric estimation to detect a single change in persistence in the 1 and 5-year Treasury bonds' liquidity premia. It is found that a signi¯cant change in both series from I(0) to I(1) occurred in the late 1990s. For the longer maturity, this is clearly linked to the timing of the Russian/LTCM crises in 1998, while for the shorter maturity there is no strong relation to the Asian currency crises. Our results also suggest that the Treasury's earlier debt management policy changes and recent uncertainty about the future U.S. ¯scal position may be a®ecting the persistence properties of default-free bond liquidity premia.

Keywords: U.S. Treasury bonds; time series persistence; emerging market crises; ¯nancial contagion

JEL classi¯cation codes: C22; F30; G10

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1. INTRODUCTION

Use of the word 'contagion' to describe the international transmission of ¯nancial crises has become fraught with controversy, to the extent that some recent authors have seen ¯t to avoid using the word entirely; see Favero and Giavazzi (2002) and Rigobon (2003). The term often evokes an emotive response among analysts of international ¯nancial markets, and there is no general agreement over its use.[1] In that respect, Eichengreen, Rose and Wyplosz (1996) proposed that contagion refers to the association of excess returns in one country with excess returns in another country after controlling for the e®ects of fundamentals. This de¯nition is closely related to 'true' contagion, as de¯ned in Kaminsky and Reinhart (2000), arising after controlling for common shocks and all possible interconnection channels.

Even with agreement on this de¯nition, there are formidable di±culties in reaching the appropriate set of fundamentals to use as control variables, suggesting that such models may not be e®ectively operational. Recent empirical research has suggested two alternative means. Dungey et al. (2003) propose the use of latent factor models, which do not require the exact speci¯cation of the fundamental relationships, while Pesaran and Pick (2003) suggest controlling for fundamental-based market interdependencies using trade °ow data and examining contagion as transmissions above that. Each approach contains an implicit criticism of the other. The Dungey et al. framework suggests that the interdependencies captured in the data are insu±ciently general,

[1] For example, Pericoli and Sbracia (2001) provide an overview of the literature containing ¯ve di®erent classi¯cations of contagion.

while Pesaran and Pick ¯nd that interdependencies are not su±ciently identi¯ed in the latent factor framework. However, both methodologies have di±culty in identifying fundamental contagion from other transmissions | a problem highlighted by Dornbusch, Claessens and Park (2000).

In this paper we take a simpler approach to volatility spillover e®ects, focusing on the timing issue between emerging market ¯nancial crises and events driven by U.S. ¯scal policy and a®ecting the U.S. Treasury market. We thus approach the issue of international ¯nancial contagion from the standpoint of the center vs. periphery debate of Kaminsky and Reinhart (2002). In particular, we ask whether the East Asian ¯nancial crises of 1997-98 and Russian/LTCM liquidity crises of 1998 led to changes in the time series properties of U.S. Treasury bond liquidity premia, or whether these were primarily driven by U.S. considerations. In that respect, commentators have noted that U.S. budget surpluses in the late 1990s led to a staged contraction in the supply of Treasury bonds with a series of debt management policy changes since 1998, notably a reduction in their issuance frequencies; see Boni and Leach (2002), Fleming (2000, 2002) and Fur¯ne and Remolona (2002). It is noteworthy that these were adopted against the background of a sustained upbeat ¯scal environment, which led to the Treasury's debt buyback program in March 2000. Subsequently, however, the U.S. ¯scal position began reverting in response to the macroeconomic slowdown of 2000-2001 and the impact of September 11, 2001.

These developments, along with ¯nancial market uncertainty | particularly during Russia's default and LTCM's near-collapse in autumn 1998 | have spawned new research on the dynamics of the Treasury market. In principle, Treasury bonds whose remaining time to maturity and other characteristics are similar should trade at approximately the same price. However, less liquid (older, or o®-the-run) bond yields are often higher than their more liquid (most recent, or on-the-run) counterparts, especially at the longer end of the term structure. Thus, researchers have interpreted this yield di®erential, typically between the ¯rst o®-the-run and the on-the-run issues at each maturity, as a time-varying liquidity premium which is expected to be mean-reverting by market e±ciency.[2]

Our main empirical aim is to examine the issue of possible changes in the persistence of the U.S. Treasury's 1-year bill and 5-year note weekly on/o® spreads. The timing of any such changes can the be related to the emerging market ¯nancial crises and U.S. ¯scal policy-driven events since 1997. We are interested in these particular maturities because both were signi¯cantly a®ected by the debt management policy changes. These included: January 1998, when the 3-year note was discontinued; May 1998 and February 2000, when the 5-year note and 1-year bill's auction frequencies were reduced from monthly to quarterly; and February and October 2001, when the 1-year and 30-year maturities were discontinued. The Treasury also increased issue sizes, leading to more liquidity through lower inventory costs. The issuance frequency reductions, coupled with greater ¯nancial market uncertainty due to the East Asian ¯nancial crises and

[2] Krishnamurthy (2003) and Longsta® (2003), among others, have documented the signi¯cance of the on/o® spread across the term structure. As pointed out by Goldreich at al (2003), if the yield curve is sloping we would expect on and o®-the-run securities to have di®erent yields even in the absence of any liquidity e®ect. Any yield curve e®ects would tend to have a greater impact on shorter than longer maturities. Thus, lowering the issuance frequency of maturities at the shorter end of the term structure would cause greater exposure to interest rate risk, and potentially a®ect the time series properties of their on/o® spreads.

Russian/LTCM liquidity crisis, raises the question of whether there was a change in persistence of Treasury bond liquidity premia from a stationary, I(0), to a nonstationary, I(1), process. Determining the location and direction of such changes is a key issue for policy makers and market forecasters alike; see Kim (2000) and Newbold et al. (2001).

To that end, we apply the recursive econometric methodology of Leybourne et al. (2003b) | henceforth LKSN | for detecting break points in time series persistence to ask whether they were signi¯cantly a®ected by the above events. The null hypothesis is that the data is I(1) throughout, and the alternative is a change from I(0) to I(1) at some point in the series. The LKSN procedure is extended by adopting weighted-symmetric (WS) estimation of the unit root coe±cient. Under stationary alternatives and OLS detrending, this estimation method yields a more powerful unit root t-test than standard Dickey-Fuller and its Generalised Least Squares (GLS)-detrended version proposed by Elliott et al. (1996).[3] LKSN develop GLS-based recursive and sequential unit root tests for detecting a single possible change in persistence under the alternative. The tests allow for an unknown breakpoint and, in their general form, unknown direction of change in persistence. Based on Monte Carlo evidence, they ¯nd the recursive tests to be favourable.

We ¯nd that both U.S. Treasury bond on/o® spreads had a signi¯cant change in the persistence in the late 1990s. For the 5-year note spread, the switch date is ambiguous, occurring in summer 1997 or spring 1998 depending on whether the White-corrected

[3]On related power gains see Leybourne et al. (2003a) and Pantula et al. (1994).

version of the tests is used to allow for the strong GARCH effects in the data. The earlier date could correspond to the onset of the Thai currency crisis and the later one to the U.S. Treasury's debt management policy changes. For the 1-year bill, we detect a significant switch from I(0) to I(1) in the persistence of the corresponding liquidity premium occurring in March 1999. The timing of this switch follows the Russian/LTCM liquidity crises in the third quarter of 1998, and precedes the reduction in the 1-year maturity's issue frequency in February 2000. Moreover, the break date is robust to adjusting (White-correcting) the test statistics. Ongoing debate about the future course of the Federal budget deficit and the consequent financial market uncertainty may be also contributing to the higher persistence of both series. Therefore, the results offer weak support for the presence of contagion from Asia to the Treasury liquidity premium, but there is a significant impact of the Russia/LTCM liquidity-driven events.

The remainder of the paper is organized as follows. Section 2 presents the model; Section 3 applies the recursive test procedure to U.S. Treasury bond on/off spreads from 1991-2002; Section 3 discusses the empirical findings and Section 4 provides some concluding remarks.

2. THE MODEL

Let the true data generating process for time series y_t be

$$y_{t+1} = d_{t+1} + u_{t+1} \quad ; \quad d_t = z_t'\beta \tag{1}$$

$$u_t = \alpha u_{t-1} + \phi(L)\Delta u_{t-1} + \epsilon_t \, ,$$

where $t = 1, \ldots, T$, $z_t = [1, t]'$ and $\beta = [\beta_0, \beta_1]'$. We restrict attention to $\beta_1 = 0$, without loss of generality. Lag polynomial $\phi(L)$ is of known order $p - 1$, where the roots of $1 - \phi(L) = 0$ lie outside the unit circle. The errors follow a martingale difference sequence and the first $p - 1$ values of y_t are assumed to exist.

The standard null hypothesis H^{11} is that y_t is $I(1)$ throughout, or $\alpha = 1$. The alternative is that y_t undergoes a change in persistence from $I(0)$ to $I(1)$ at observation $\tau^* T$ in forward time,

$$\begin{aligned} |\alpha| &< 1, \quad t \le \tau^* T \\ \alpha &= 1, \quad t > \tau^* T \end{aligned} \tag{2}$$

or from $I(1)$ to $I(0)$, implying the time-reversed series $\tilde{y}_t = y_{T-t+1}$, $t = 1, 2, \ldots, T$ changes from $I(0)$ to $I(1)$ at observation $(1 - \tau^*)T$, where the break fraction τ^* is unknown. The respective alternative hypotheses are denoted H^{01} and H^{10}.

Our test statistics are constructed as follows. After detrending the series by OLS, $y_t^d = y_t - \hat{\beta}_0(\tau)$; $t = 1, 2, \ldots, T$, an ADF regression with no trend is ran on Δy_t^d using only the first $[\tau T]$ observations for varying break fraction τ,

$$\Delta y_t^d = \hat{b}(\tau) y_{t-1}^d + \Sigma_{j=1}^{p-1} \hat{\chi}_j(\tau) \Delta y_{t-j}^d + \hat{\varepsilon}_t; \quad t = 1, 2, \ldots, [\tau T] \tag{3}$$

where $[\cdot]$ is the integer part of τT and τ belongs to a closed interval Λ in $(0, 1)$:

In this setting, weighted-symmetric estimation of $\rho(\tau)$ — proposed originally in Fuller (1976) — minimizes

$$Q(\theta) = \sum_{t=p+1}^{[\tau T]} w_t \left(\Delta y_t^d - \rho(\tau) y_{t-1}^d - \sum_{j=1}^{p-1} \phi_j(\tau) \Delta y_{t-j}^d \right)^2 \tag{4}$$

$$+ \sum_{t=1}^{[\tau T]-p} (1 - w_{t+1}) \left(\Delta y_t^d - \rho(\tau) y_{t+1}^d + \sum_{j=1}^{p-1} \phi_j(\tau) \Delta y_{t+j+1}^d \right)^2 \tag{5}$$

over all τ, and $\theta = (\rho, \phi)$, $\phi = \{\phi_1, \phi_2, \ldots, \phi_{p-1}\}$ with w_t defined as

$$w_t = \begin{cases} 0; & 1 \leq t < p+1 \\ (t-p)/([\tau T] - 2p + 2); & p+1 \leq t < [\tau T] - p + 2 \\ 1; & [\tau T] - p + 2 \leq t \leq [\tau T]. \end{cases}$$

The t-statistic associated with $\hat{\rho}(\tau)$ under the null is $WS(\tau) = \frac{\hat{\rho}(\tau)}{\sqrt{\widehat{var}(\hat{\rho}(\tau))}}$, where $\widehat{var}(\hat{\rho}(\tau)) = \hat{\sigma}^2(\tau) h_{PP}$, the estimated error standard deviation is $\hat{\sigma}(\tau) = \frac{Q(\hat{\theta})}{[\tau T] - p - 2}$ and h_{PP} is the $[1;1]$ element of the $\left(\frac{\partial^2 Q(\theta)}{\partial \theta \partial \theta'}\right)^{-1}$ matrix.[4]

The statistic for testing the alternative hypothesis H^{01} is given by

$$WS^{f\ \inf}(\tau) = \inf_{\tau \in *} WS^f(\tau); \tag{6}$$

where f denotes the recursive test in forward time and τ^* is the break fraction minimising equation (6). When the alternative hypothesis is a switch from $I(1)$ to $I(0)$, this test statistic can be applied to the first-difference of time-reversed series $\tilde{y}_t^d$

[4]Note that if a trend is included in the regression the denominator for $\hat{\sigma}(\tau)$ becomes $[\tau T] - p - 3$.

$$\Delta \tilde{y}_t^d = \hat{\alpha}(\tau)\tilde{y}_{t-1}^d + \Sigma_{j=1}^{p-1}\hat{\beta}_j(\tau)\Delta\tilde{y}_{t-j}^d + \hat{\varepsilon}_t\,; \qquad t = 1, 2, ..., (1-\tau)T. \tag{7}$$

Let the t-ratio for $\hat{\alpha}(\tau)$ be $WS^r(\tau)$. The statistic for testing H^{11} against H^{10} then is

$$WS^{r\ inf}(\tau) = \inf_{\tau \in \Lambda} WS^r(\tau)\ , \tag{8}$$

with r denoting the test on the time-reversed series.

If one is a priori uncertain about the direction of change in persistence, a "two-sided" test can be constructed whose null is I(1) throughout against the alternative of a change from I(0) to I(1) or vice versa at break fraction τ^*. The statistic is then the pairwise minimum of $WS^{f\ inf}$ and $WS^{r\ inf}$

$$\min(WS^{f\,inf}; WS^{r\,inf}). \tag{9}$$

Following LKSN and existing asymptotic results in Park and Fuller (1995), the $WS^{f\,inf}$ and $WS^{r\,inf}$ tests will be consistent only under the alternative for which they are designed. Thus, the $\min(WS^{f\,inf}; WS^{r\,inf})$ test will also be consistent under H^{01} or H^{10}. Moreover, all test statistics can be shown to estimate the break fraction consistently against the true alternative. These results also imply that the ADF and non-recursive WS tests are inconsistent under a break in persistence, as the random walk component of the series will dominate these statistics and render them $O_p(1)$. In the remainder of this paper, WS refers to the statistic using the non-recursive weighted-symmetric estimation procedure, and WS-based tests refer to both the recursive and the non-recursive statistics.

3. CHANGING PERSISTENCE IN U.S. TREASURY BOND LIQUIDITY PREMIA

Our sample period extends from 17.6.1991 to 31.12.2002, that is 504 weekly observations on the levels of the 1-year Treasury bill on/o® spread | the yield di®erential between the ¯rst o®-the-run and the on-the-run issues | and 592 for the 5-year Treasury note on/o® spread. Wednesday observations are selected from the daily data to address day-of-the-week e®ects. In°ation-indexed and callable bond issues are excluded, as are holidays and observations more than 30 basis points, re°ecting posting errors. The data source is GovPX, posting price, yield and volume information on 5 out of the 6 U.S. Treasury interdealer brokers.[5]

Figure 1 shows the two series in basis points and Table 1 summarizes their distributional properties.

FIGURE 1 AND TABLE 1 HERE

Both spreads are tightly distributed around their mean until the late 1990s, when they become more volatile, and there is signi¯cant excess kurtosis and GARCH e®ects. From 2000 onwards, the volatility of both on/o® spreads is strikingly greater. This is

[5] Note that the non-reporting broker, Cantor-Fitzgerald, is relatively more important for long maturities.

related as much to the reduction in the maturities' issuance frequency, implying more interest rate risk, as to investors uncertain outlook regarding future U.S. fiscal policy. Also, since early 2001 — after which the 1-year bill was discontinued — the 5-year on/off spread displays a negative trend, likely due to the sharp inversion in the U.S. yield curve. Table 2 reports estimates of the AR-GARCH process $y_t = c + ay_{t-1} + \varepsilon_t$, where $\varepsilon_t = h_t^{1/2} v_t$, $h_t = \phi_0 + \phi_1 \varepsilon^2_{t-1} + \phi_2 h_{t-1}$.

TABLE 2 HERE

The ϕ_1 coefficient estimate captures short-run variation in volatility. This is estimated at approximately 0.15 for both series, while $\phi_1 + \phi_2$ (persistence in volatility) is around 0.95. However, these results should be treated with caution as in the presence of a break the estimate of a could be biased, affecting GARCH parameter estimates. Table 3, Panel A reports the results of ADF and WS tests for the whole sample. Critical values for these and subsequent tests, based on 20,000 replications, are given in Appendix A.

TABLE 3 HERE

The lag order p is selected using the sequential 0.10 level t-tests for the longest lag coefficient's significance. We use the same p for WS-based tests as selected from the standard ADF regressions in equations (6) and (7).[6] Nonstationarity is not rejected for the 5-year on/off spread, while it is for the 1-year spread. However, when the WS

[6] A trend term is not included in the regressions on market efficiency grounds.

statistic is corrected for GARCH in Table 3, Panel B, non-stationarity is not rejected for either series; WS_w denotes the White-corrected statistic.

In general, standard unit root tests are asymptotically valid in the presence of conditional heteroskedasticity. The robustness of unit root limit theory to conditional heterogeneity was noted by, among others, Phillips (1987) and Phillips and Perron (1988). However, simulations reported in Kim and Schmidt (1993) and Seo (1999) indicate that DF statistics tend to overreject the null hypothesis when GARCH errors are persistent, and to decrease towards nominal size at a very slow rate as T increases. The former authors also consider DF t-ratios using White's (1980) heteroskedasticity-consistent covariance matrix estimator, and find that the observed size distortions can be eliminated.[7]

Turning to the recursive tests, the discussion in Section 1 suggested that the alternative hypothesis is a change in persistence from $I(0)$ to $I(1)$ at observation $\tau^* T$. Hence, the $WS^{f\,\inf}$ test in equation (6) is applied to the series, as it is expected to be more powerful than the two-sided $\min(WS^{f\,\inf}; WS^{r\,\inf})$ test.[8] The results in Table 4, Panels A and B contain both the non-White-corrected, $WS^{f\,\inf}$, and White-corrected, $WS_w^{f\,\inf}$, versions. Results for the reverse and "two-sided" test statistics, respectively $WS^{r\,\inf}$ and $\min(WS^{f\,\inf}; WS^{r\,\inf})$, are also included.

TABLE 4 HERE

For the 1-year Treasury bill on/off spread, the $WS^{f\,\inf}$ and $\min(WS^{f\,\inf}; WS^{r\,\inf})$

[7] For a Monte Carlo assessment of these findings for WS-based tests see Smith and Tambakis (2003).

[8] We let τ vary between 0.15 and 0.85 in 0.01 increments. Note that LKSN use GLS-detrending and trim at 0.20 employing the usual ADF statistics.

tests both reject the unit root null at the 0.05 level. Supporting this outcome, the null is not rejected using $WS^{r\,inf}$. The White-corrected statistics, $WS_w^{f\,inf}$ and $WS_w^{r\,inf}$, point in the same direction. For the 5-year spread, both $WS^{f\,inf}$ and $\min(WS^{f\,inf}; WS^{r\,inf})$ reject the null at the 0.05 level. The rejections are less significant in the White-corrected case. The $WS_w^{f\,inf}$ statistic rejects the null at the 0.10 level, marginally missing the 0.05 critical value, while $\min(WS_w^{f\,inf}; WS_w^{r\,inf})$ is narrowly not rejecting at 0.10.

Finally, in Table 5 we report the results of WS tests for the pre- and post-break subsamples. The significant break dates are those given in Table 4, as determined by the forward-based recursive test $WS^{f\,inf}$.

TABLE 5 HERE

The pre-break and post-break subsamples are, respectively, stationary and non-stationary at the 0.01 and 0.05 levels, both with and without the White-correction, providing further support for our postulated alternative hypothesis of a break in the two liquidity premia's persistence from stationary to non-stationary.

3. INTERPRETATION AND DISCUSSION

From Table 4, the switch from $I(0)$ to $I(1)$ according to the non-White-corrected statistics is found in July 1997 for the 5-year on/off spread and March 1999 for the 1-year spread. The corresponding breaks under the White-corrected statistics are in May 1998 and March 1999. Thus, the evidence on the likely impact from world events on the 5-year liquidity premium is sensitive to whether the test statistic employed allows for the presence of GARCH or not. It could be argued that the earlier break date, obtained under non-White correction, coincides with the outbreak of the Asian financial crises, which started with the devaluation of the Thai baht. The later break date under White-correction may reflect the reduction in the frequency of 5-year Treasury bond auctions from monthly to quarterly in May 1998.

In contrast, the break date identified by the recursive unit root tests for the 1-year liquidity premium is unambiguous regardless of the adjustment for GARCH. The switch from I(0) to I(1) in March 1999 occurs in the aftermath of the Russian/LTCM liquidity crises in the third quarter of 1998, long before the reduction in the 1-year maturity's issue frequency in February 2000. The fact that this change was triggered against a background of sustained expectations of future Federal budget surpluses suggests that the impact of the Russia/LTCM crises on investor behavior was very strong.

We tentatively conclude that the East Asian financial crises of 1997-98 — triggered by currency upheaval — did not impact upon the U.S Treasury bonds' on/off spread behavior as much as the subsequent Russian/LTCM crises of 1998. The implication is that the uncertain outcome of these events and likely adverse consequences to the economy have made the Treasury liquidity premium more volatile, particularly at the 5-year maturity. To the extent that the latter events have been widely characterized as liquidity crises affecting the U.S. financial system, a further implication is that the on/off spread time series between adjacent security issues is indeed a good proxy for the time-varying liquidity premium, as recently argued by Krishnamurthy (2003) and Longstaff (2003).

Therefore, with regard to ¯nancial contagion our ¯ndings suggest that there was no signi¯cant volatility spillover from the East Asian ¯nancial crises to the U.S. Treasury market. However, the liquidity-driven events of Russia's default and associated LTCM fallout in 1998 may have contributed to ¯nancial market uncertainty | especially at the short end of the Treasury yield curve | and a®ected the persistence of the relevant liquidity premium. In that respect, it could also be argued that the change in persistence in both on/o® spread time series may currentlly be sustained by growing ¯nancial market uncertainty concerning the future course of the U.S. ¯scal position.

4. CONCLUDING REMARKS

This paper studied the likely spillover from the Asian and Russian/LTCM ¯nancial crises to U.S. Treasury liquidity premia. The recursive unit root tests of Leybourne et al. (2003b) for detecting a single change in time series persistence were applied to the whole sample, and non-recursive weighted-symmetric tests were ran on the pre- and post-break subsamples. Examining the behavior of the 1-year and 5-year U.S. Treasury bonds' on/o® spreads, a signi¯cant break from I(0) to I(1) was found in the late 1990s. It was suggested that while ¯nancial market uncertainty | the current reversal of the U.S. ¯scal position following the earlier debt reduction initiative | may have a®ected the persistence properties at the longer Treasury maturity, the shorter maturity's change in persistence is clearly related to the Russian/LTCM liquidity crises in autumn 1998. The results also serve to caution analysts regressing on/o® yield spreads as a stationary explanatory variable in factor models of credit spreads.

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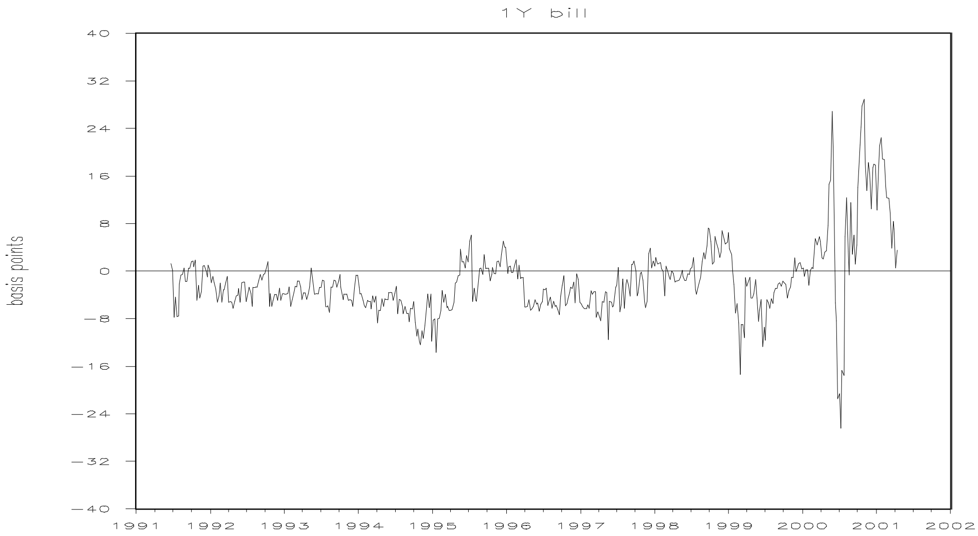

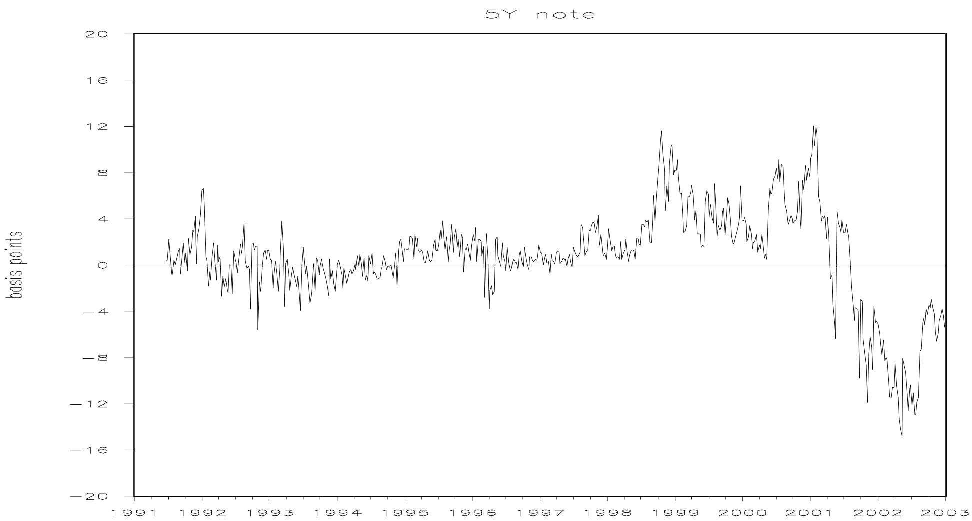

Figure 1. U.S. Treasury bond on/o® spread levels: 1991-2002.

Appendix A

Simulated critical values

Panel A Statistic	T	0:01	0:05	0:10	Panel B Statistic	T	0:01	0:05	0:10
ADF	500	-3.420	-2.875	-2.578	ADF_w	500	-3.453	-2.905	-2.598
WS	100	-3.124	-2.552	-2.235	WS_w	100	-2.857	-2.299	-2.007
	250	-3.160	-2.554	-2.255		250	-2.796	-2.262	-1.982
	350	-3.111	-2.538	-2.255		350	-2.737	-2.232	-1.971
	400	-3.080	-2.543	-2.222		400	-2.733	-2.225	-1.949
	500	-3.109	-2.540	-2.228		500	-2.745	-2.220	-1.942
$WS^{f\,inf}$	500	-3.909	-3.325	-3.030	$WS_w^{f\,inf}$	500	-3.529	-3.004	-2.729
$WS^{r\,inf}$	500	-3.943	-3.323	-3.033	$WS_w^{r\,inf}$	500	-3.578	-3.003	-2.721
min	500	-4.162	-3.586	-3.309	min_w	500	-3.770	-3.252	-2.993

Note: Statistics $\min(WS^{f\,inf}; WS^{r\,inf})$ and $\min(WS_w^{f\,inf}; WS_w^{r\,inf})$ are respectively denoted by min and min_w. Beyond $T = 500$ critical values for WS-based tests appeared to converge.

TABLE 1
On/o® bond spread statistics: 1991-2002

Statistics	1Y	5Y
Mean	-1.632	0.684
Std. Dev.	6.311	4.202
Max	28.90	12.00
Min	-26.50	-14.80
Skewness	1.518	-0.771
Kurtosis	8.437	4.947
Jarque-Bera	820.76	152.18

TABLE 2
AR(1)-GARCH(1,1) maximum likelihood parameter estimates

Series	c	a	Á$_0$	Á$_1$	Á$_2$
1Y	-0.591	0.825	0.508	0.168	0.770
	(0.103)	(0.023)	(0.147)	(0.028)	(0.039)
5Y	0.126	0.900	0.086	0.131	0.846
	(0.051)	(0.015)	(0.026)	(0.023)	(0.024)

Note: Standard errors are in parentheses.

TABLE 3
ADF and WS tests for whole sample

Panel A		
Series	ADF	WS
1Y	-3.009[b]	-3.168[a]
5Y	-1.409	-1.711
Panel B		
Series	ADF$_w$	WS$_w$
1Y	-2.102	-1.680
5Y	-1.299	-1.209

Note: a; b; c denote 0:01, 0:05 and 0:10 signi¯cance levels.

TABLE 4

Recursive WS tests for a change in persistence

Panel A Series	$WS^{f\ inf}$	Break date	$WS^{r\ inf}$	Break date	min(:;:)	Break date
1Y	-3.602[b]	03/03/99	-2.808	n/a	-3.602[b]	03/03/99
5Y	-3.603[b]	30/07/97	-2.371	n/a	-3.603[b]	30/07/97
Panel B Series	$WS_w^{f\ inf}$	Break date	$WS_w^{r\ inf}$	Break date	$\min_w$(:;:)	Break date
1Y	-3.282[b]	24/03/99	-1.692	n/a	-3.282[b]	24/03/99
5Y	-2.927[c]	27/05/98	-1.555	n/a	-2.927	n/a

Note: Statistics $\min(WS^{f\ inf}; WS^{r\ inf})$ and $\min(WS_w^{f\ inf}; WS_w^{r\ inf})$ are respectively denoted min and $\min_w$. Break dates are reported only when the null is rejected. The signi¯cant break points are 395 (03=03=99), 314 (30=07=97), 398 (24=03=99), 357 (27=05=98):

TABLE 5

WS tests for subsamples

Panel A	$WS^{f\ inf}$	
Series	Pre-break	Post-break
1Y	-3.602[a]	-1.847
5Y	-3.603[a]	-1.736
Panel B	$WS_w^{f\ inf}$	
Series	Pre-break	Post-break
1Y	-3.282[a]	-1.651
5Y	-2.927[b]	-1.261