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

L. VANESSA SMITH\(^y\) AND DEMOSTHENES N. TAMBAKIS\(^z\)

December 2003

ABSTRACT

Have the emerging market financial crises of recent years had a significant 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 significant 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. fiscal position may be affecting the persistence properties of default-free bond liquidity premia.

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

JEL classification codes: C22; F30; G10

\(^y\)Corresponding author: Cambridge Endowment for Research in Finance (CERF), Judge Institute of Management, Trumpington Street, Cambridge CB2 1AG, UK. E-mail: v.smith@cerf.cam.ac.uk; Tel: +44 (0)1223-760 578; fax: +44(0)1223-339701

\(^z\)Pembroke College, Cambridge CB2 1RF, UK and CERF. We would like to thank Victoria Ilyassova for excellent data assistance. The usual disclaimer applies.
1. INTRODUCTION

Use of the word 'contagion' to describe the international transmission of financial crises has become fraught with controversy, to the extent that some recent authors have seen it 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 financial 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 effects of fundamentals. This definition is closely related to 'true' contagion, as defined in Kaminsky and Reinhart (2000), arising after controlling for common shocks and all possible interconnection channels.

Even with agreement on this definition, there are formidable difficulties in reaching the appropriate set of fundamentals to use as control variables, suggesting that such models may not be effectively 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 specification of the fundamental relationships, while Pesaran and Pick (2003) suggest controlling for fundamental-based market interdependencies using trade flow 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 insufficiently general,\(^1\)

\(^1\)For example, Pericoli and Sbracia (2001) provide an overview of the literature containing different classifications of contagion.
while Pesaran and Pick find that interdependencies are not sufficiently identified in the latent factor framework. However, both methodologies have difficulty 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 effects, focusing on the timing issue between emerging market financial crises and events driven by U.S. fiscal policy and affecting the U.S. Treasury market. We thus approach the issue of international financial contagion from the standpoint of the center vs. periphery debate of Kaminsky and Reinhart (2002). In particular, we ask whether the East Asian financial 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 Furse and Remolona (2002). It is noteworthy that these were adopted against the background of a sustained upbeat fiscal environment, which led to the Treasury's debt buyback program in March 2000. Subsequently, however, the U.S. fiscal position began reverting in response to the macroeconomic slowdown of 2000-2001 and the impact of September 11, 2001.

These developments, along with financial 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 off-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 differential,
typically between the rst off-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.

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 be related to the emerging market nancial crises and U.S.
iscal 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

²Krishnamurthy (2003) and Longsta® (2003), among others, have documented the signi®cant change 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 different 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)\(^3\) henceforth LKSN\(^\dagger\) for detecting break points in time series persistence to ask whether they were significantly affected 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 coefficient. 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 find the recursive tests to be favourable.

We find that both U.S. Treasury bond on/off-spreads had a significant 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

$$y_{t+1} = d_{t+1} + u_{t+1}; \quad d_t = z_t^\theta$$

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

\[5\]
where \( t = 1; \ldots; T \), \( z_t = [1; t]^0 \) and \( \bar{\gamma} = [-\bar{\gamma}; -1]^0 \). We restrict attention to \( \bar{\gamma} = 0 \), without loss of generality. Lag polynomial \( \hat{A}(L) \) is of known order \( p_i = 1 \), where the roots of \( 1 - \hat{A}(L) = 0 \) lie outside the unit circle. The errors follow a martingale difference sequence and the first \( p_i = 1 \) values of \( y_t \) are assumed to exist.

The standard null hypothesis \( H_{11} \) is that \( y_t \) is \( I(1) \) throughout, or \( \bar{\gamma} = 1 \). The alternative is that \( y_t \) undergoes a change in persistence from \( I(0) \) to \( I(1) \) at observation \( \bar{\gamma}T \) in forward time,

\[
\begin{align*}
\bar{\gamma} &< 1; \quad t \cdot \bar{\gamma}T \\
\bar{\gamma} &> 1; \quad t > \bar{\gamma}T
\end{align*}
\] (2)

or from \( I(1) \) to \( I(0) \), implying the time-reversed series \( y_t = y_{T-t+1}, t = 1; 2; \ldots; T \) changes from \( I(0) \) to \( I(1) \) at observation \( (1 - \bar{\gamma})T \), where the break fraction \( \bar{\gamma} \) 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, \( y_t^d = y_t - \hat{\gamma}_0 \hat{\gamma} \); \( t = 1; 2; \ldots; T \), an ADF regression with no trend is ran on \( \bar{\gamma} y_t^d \) using only the first \( \lfloor \bar{\gamma}T \rfloor \) observations for varying break fraction \( \bar{\gamma} \),

\[
\bar{\gamma} y_t^d = b \bar{\gamma} y_{t-1}^d + \sum_{j=1}^{p_i} \hat{A}_j(\bar{\gamma}) \bar{\gamma} y_{t-j}^d + \bar{\xi}_t; \quad t = 1; 2; \ldots; \lfloor \bar{\gamma}T \rfloor
\] (3)

where \( \lfloor \bar{\gamma} \rfloor \) is the integer part of \( \bar{\gamma}T \) and \( \bar{\gamma} \) belongs to a closed interval \( \alpha \) in \((0; 1)\):
In this setting, weighted-symmetric estimation of \( \frac{1}{2} (\eta) \) proposed originally in Fuller (1976) minimizes

\[
Q(\mu) = \sum_{t=p+1}^{T} \left( \begin{array}{c} 1 \end{array} \right) \left( \begin{array}{c} \eta_T \end{array} \right) \left( \begin{array}{c} 1 \end{array} \right) \left( \begin{array}{c} \hat{A} \end{array} \right) \left( \begin{array}{c} 1 \end{array} \right) \left( \begin{array}{c} y^d_{t+1} \end{array} \right) + \sum_{t=1}^{p} \left( \begin{array}{c} w_t \end{array} \right) \left( \begin{array}{c} 1 \end{array} \right) \left( \begin{array}{c} 1 \end{array} \right) \left( \begin{array}{c} \hat{A} \end{array} \right) \left( \begin{array}{c} 1 \end{array} \right) \left( \begin{array}{c} y^d_{t+1} \end{array} \right) + \sum_{j=1}^{p} \left( \begin{array}{c} \hat{A} \end{array} \right) \left( \begin{array}{c} 1 \end{array} \right) \left( \begin{array}{c} y^d_{t+j+1} \end{array} \right)
\]

over all \( \eta \), and \( \mu = (\frac{1}{2} \hat{A}, \hat{A}_1, \hat{A}_2, \ldots, \hat{A}_{p+1}) \) with \( w_t \) defined as:

\[
w_t = \begin{cases} 0; & t < p + 1 \\ 1; & t \geq p + 1 \end{cases}
\]

The statistic for testing the alternative hypothesis \( H_{01} \) is given by

\[
WS^f \inf (\eta) = \inf_{\eta_2} \eta_2 + WS^f (\eta);
\]

where \( f \) denotes the recursive test in forward time and \( \eta_2 \) is the break fraction minimizing 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 \( \eta^d \).

\[4^\text{Note that if a trend is included in the regression the denominator for } \hat{B}(\hat{\xi}) \text{ becomes } [\hat{\xi}T]; p; 3; \]

\[7\]
\[ c_y = \epsilon_y \epsilon_{j+1} + \sum_{i=1}^{\pi} \epsilon_{j+i} \epsilon_j + \epsilon; \quad t = 1, 2, \ldots, (1 - \xi)T \] (7)

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

\[ WS^r_{\inf}(\xi) = \inf_{\xi_2} WS^r(\xi), \] (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 \( \xi_1 \). The statistic is then the pairwise minimum of \( WS^f_{\inf} \) and \( WS^r_{\inf} \)

\[ \min(WS^f_{\inf}, WS^r_{\inf}); \] (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

5Note 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/o® 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 $y_t = c + a y_{t-1} + \varepsilon_t$, where $\varepsilon_t = h_t^{-1/2} v_t$, $h_t = \hat{h}_0 + \hat{h}_1 \varepsilon_{t-1}^2 + \hat{h}_2 h_{t-1}$.

**TABLE 2 HERE**

The $\hat{h}_1$ coefficient estimate captures short-run variation in volatility. This is estimated at approximately $0.15$ for both series, while $\hat{h}_1 + \hat{h}_2$ (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). Nonstationarity is not rejected for the 5-year on/o® spread, while it is for the 1-year spread. However, when the WS

---

*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 nd that the observed size distortions can be eliminated.\footnote{For a Monte Carlo assessment of these ndings for WS-based tests see Smith and Tambakis (2003).}

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 $\omega T$. 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.\footnote{We let $\omega$ 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.} The results in Table 4, Panels A and B contain both the non-White-corrected, $WS^f_{Inf}$, and White-corrected, $WS^f_{Infw}$, versions. Results for the reverse and \textquotedblleft two-sided\textquotedblright test statistics, respectively $WS^r_{Inf}$ and $\min(WS^f_{Inf}; WS^r_{Inf})$, are also included.

\textbf{TABLE 4 HERE}

For the 1-year Treasury bill on/o spread, the $WS^f_{Inf}$ and $\min(WS^f_{Inf}; WS^r_{Inf})$
tests both reject the unit root null at the 0.05 level. Supporting this outcome, the null is not rejected using $WS_r^{\text{inf}}$. The White-corrected statistics, $WS_f^{\text{inf}}$ and $WS_w^{\text{inf}}$, point in the same direction. For the 5-year spread, both $WS_f^{\text{inf}}$ and $\min(WS_f^{\text{inf}}; WS_r^{\text{inf}})$ reject the null at the 0.05 level. The rejections are less significant in the White-corrected case. The $WS_w^{\text{inf}}$ statistic rejects the null at the 0.10 level, marginally missing the 0.05 critical value, while $\min(WS_w^{\text{inf}}; WS_w^{\text{inf}})$ 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 $WS_f^{\text{inf}}$.

**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/o® 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/o® 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 Longsta® (2003).
Therefore, with regard to financial contagion our findings suggest that there was no significant volatility spillover from the East Asian financial 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 financial market uncertainty especially at the short end of the Treasury yield curve and affected the persistence of the relevant liquidity premium. In that respect, it could also be argued that the change in persistence in both on/off spread time series may currently be sustained by growing financial market uncertainty concerning the future course of the U.S. fiscal position.

4. CONCLUDING REMARKS

This paper studied the likely spillover from the Asian and Russian/LTCM financial 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/off spreads, a significant break from I(0) to I(1) was found in the late 1990s. It was suggested that while financial market uncertainty the current reversal of the U.S. fiscal position following the earlier debt reduction initiative may have affected 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/off yield spreads as a stationary explanatory variable in factor models of credit spreads.
References


Figure 1. U.S. Treasury bond on/off spread levels: 1991-2002.
## Appendix A

### Simulated critical values

<table>
<thead>
<tr>
<th>Panel A</th>
<th>Statistic</th>
<th>$T$</th>
<th>0:01</th>
<th>0:05</th>
<th>0:10</th>
<th>Panel B</th>
<th>Statistic</th>
<th>$T$</th>
<th>0:01</th>
<th>0:05</th>
<th>0:10</th>
</tr>
</thead>
<tbody>
<tr>
<td>WS</td>
<td></td>
<td>100</td>
<td>-3.124</td>
<td>-2.552</td>
<td>-2.235</td>
<td>WS$_w$</td>
<td>100</td>
<td>-2.857</td>
<td>-2.299</td>
<td>-2.007</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>400</td>
<td>-3.080</td>
<td>-2.543</td>
<td>-2.222</td>
<td></td>
<td>400</td>
<td>-2.733</td>
<td>-2.225</td>
<td>-1.949</td>
<td></td>
</tr>
<tr>
<td>min</td>
<td></td>
<td>500</td>
<td>-4.162</td>
<td>-3.586</td>
<td>-3.309</td>
<td>min$_w$</td>
<td>500</td>
<td>-3.770</td>
<td>-3.252</td>
<td>-2.993</td>
<td></td>
</tr>
</tbody>
</table>

Note: Statistics $\min(WS_f, WS_r)$ and $\min(WS_w, WS_r)$ are respectively denoted by $\min$ and $\min_w$. Beyond $T = 500$ critical values for WS-based tests appeared to converge.
### TABLE 1


<table>
<thead>
<tr>
<th>Statistics</th>
<th>1Y</th>
<th>5Y</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean</td>
<td>-1.632</td>
<td>0.684</td>
</tr>
<tr>
<td>Std. Dev.</td>
<td>6.311</td>
<td>4.202</td>
</tr>
<tr>
<td>Max</td>
<td>28.90</td>
<td>12.00</td>
</tr>
<tr>
<td>Min</td>
<td>-26.50</td>
<td>-14.80</td>
</tr>
<tr>
<td>Skewness</td>
<td>1.518</td>
<td>-0.771</td>
</tr>
<tr>
<td>Kurtosis</td>
<td>8.437</td>
<td>4.947</td>
</tr>
<tr>
<td>Jarque-Bera</td>
<td>820.76</td>
<td>152.18</td>
</tr>
</tbody>
</table>

### TABLE 2

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

<table>
<thead>
<tr>
<th>Series</th>
<th>c</th>
<th>a</th>
<th>A_0</th>
<th>A_1</th>
<th>A_2</th>
</tr>
</thead>
<tbody>
<tr>
<td>1Y</td>
<td>0.591</td>
<td>0.825</td>
<td>0.508</td>
<td>0.168</td>
<td>0.770</td>
</tr>
<tr>
<td></td>
<td>(0.103)</td>
<td>(0.023)</td>
<td>(0.147)</td>
<td>(0.028)</td>
<td>(0.039)</td>
</tr>
<tr>
<td>5Y</td>
<td>0.126</td>
<td>0.900</td>
<td>0.086</td>
<td>0.131</td>
<td>0.846</td>
</tr>
<tr>
<td></td>
<td>(0.051)</td>
<td>(0.015)</td>
<td>(0.026)</td>
<td>(0.023)</td>
<td>(0.024)</td>
</tr>
</tbody>
</table>

Note: Standard errors are in parentheses.

### TABLE 3

ADF and WS tests for whole sample

<table>
<thead>
<tr>
<th>Panel A</th>
<th>Series</th>
<th>ADF</th>
<th>WS</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>1Y</td>
<td>-3.009^b</td>
<td>-3.168^a</td>
</tr>
<tr>
<td></td>
<td>5Y</td>
<td>-1.409</td>
<td>-1.711</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Panel B</th>
<th>Series</th>
<th>ADF_w</th>
<th>WS_w</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>1Y</td>
<td>-2.102</td>
<td>-1.680</td>
</tr>
<tr>
<td></td>
<td>5Y</td>
<td>-1.299</td>
<td>-1.209</td>
</tr>
</tbody>
</table>

Note: a; b; c denote 0.01; 0.05 and 0.10 significance levels.
### TABLE 4
Recursive WS tests for a change in persistence

<table>
<thead>
<tr>
<th>Panel A</th>
<th>Series</th>
<th>WS$_f^{inf}$</th>
<th>Break date</th>
<th>WS$_r^{inf}$</th>
<th>Break date</th>
<th>min($;$)</th>
<th>Break date</th>
</tr>
</thead>
<tbody>
<tr>
<td>1Y</td>
<td>-3.602$^a$</td>
<td>03/03/99</td>
<td>-2.808</td>
<td>n/a</td>
<td>-3.602$^a$</td>
<td>03/03/99</td>
<td></td>
</tr>
<tr>
<td>5Y</td>
<td>-3.603$^b$</td>
<td>30/07/97</td>
<td>-2.371</td>
<td>n/a</td>
<td>-3.603$^b$</td>
<td>30/07/97</td>
<td></td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Panel B</th>
<th>Series</th>
<th>WS$_w^{inf}$</th>
<th>Break date</th>
<th>WS$_w^{inf}$</th>
<th>Break date</th>
<th>min$_w$($;$)</th>
<th>Break date</th>
</tr>
</thead>
<tbody>
<tr>
<td>1Y</td>
<td>-3.282$^a$</td>
<td>24/03/99</td>
<td>-1.692</td>
<td>n/a</td>
<td>-3.282$^a$</td>
<td>24/03/99</td>
<td></td>
</tr>
<tr>
<td>5Y</td>
<td>-2.927$^c$</td>
<td>27/05/98</td>
<td>-1.555</td>
<td>n/a</td>
<td>-2.927</td>
<td>n/a</td>
<td></td>
</tr>
</tbody>
</table>

Note: Statistics min(WS$_f^{inf}$;WS$_r^{inf}$) and min(WS$_w^{inf}$;WS$_w^{inf}$) are respectively denoted min and min$_w$. Break dates are reported only when the null is rejected. The significant 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

<table>
<thead>
<tr>
<th>Panel A</th>
<th>Series</th>
<th>WS$_f^{inf}$</th>
<th>Pre-break</th>
<th>Post-break</th>
</tr>
</thead>
<tbody>
<tr>
<td>1Y</td>
<td>-3.602$^a$</td>
<td>-1.847</td>
<td></td>
<td></td>
</tr>
<tr>
<td>5Y</td>
<td>-3.603$^b$</td>
<td>-1.736</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Panel B</th>
<th>Series</th>
<th>WS$_f^{inf}$</th>
<th>Pre-break</th>
<th>Post-break</th>
</tr>
</thead>
<tbody>
<tr>
<td>1Y</td>
<td>-3.282$^a$</td>
<td>-1.651</td>
<td></td>
<td></td>
</tr>
<tr>
<td>5Y</td>
<td>-2.927$^b$</td>
<td>-1.261</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>