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Working Paper

Testing for a Unit Root with Near-Integrated Volatility

Tinbergen Institute Discussion Paper, No. 01-077/4

Provided in Cooperation with:

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Suggested Citation: Boswijk, H. Peter (2001) : Testing for a Unit Root with Near-Integrated Volatility, Tinbergen Institute Discussion Paper, No. 01-077/4, Tinbergen Institute, Amsterdam and Rotterdam

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TI 2001-077/4

Tinbergen Institute Discussion Paper

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Testing for a Unit Root with Near-Integrated Volatility*

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August 2000; revised August 2001

Abstract

This paper considers tests for a unit root when the innovations follow a near-integrated GARCH process. We compare the asymptotic properties of the likelihood ratio statistic with that of the least-squares based Dickey-Fuller statistic. We first use asymptotics where the GARCH variance process is stationary with fixed parameters, and then consider parameter sequences such that the GARCH process converges to a diffusion process. In both cases, we find a substantial asymptotic local power gain of the likelihood ratio test for parameter values that imply heavy tails in the unconditional innovation distribution. An empirical application to the term structure of interest rates in the Netherlands illustrates the proposed procedures.

1 Introduction

A well-known property of financial time series is that their conditional variance displays variation over time, such that persistent periods of high variation are followed by low-volatility periods. This phenomenon, known as *volatility clustering*, is modelled in the econometrics literature either by GARCH (generalized autoregressive-conditional heteroskedasticity) type models (see Bollerslev *et al.*, 1994, for an overview) or by stochastic volatility models, see e.g. Shephard (1996). When applied to daily financial returns data, both classes of models display a high degree of persistence, and hence a low degree of mean-reversion in the volatility process. Such processes are referred to as *near-integrated*, since their characteristic polynomial has a root close to but not necessarily equal to unity. Boswijk (1999) considers asymptotic distribution theory for likelihood based estimators of the volatility parameters in near-integrated exponential GARCH (EGARCH) models and stochastic volatility models.

In the present paper we study the effect of such near-integrated volatility processes on testing for an autoregressive unit root in the level of the process itself (instead of its volatility). This problem is relevant in finance, for example when models for the term structure of interest rates depend on the presence and degree of mean-reversion in the short rate. A typical model for the short rate is the one by

*Helpful comments from and discussions with Peter Reinhard Hansen, André Lucas and Anders Rahbek are gratefully acknowledged.

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Vasicek (1977), which is essentially a first-order autoregression with constant volatility. When applied to daily or weekly interest rates, the hypothesis of a unit root (i.e., no mean-reversion) often cannot be rejected, and a possible explanation of this is that least-squares based tests are not powerful enough to discover the (weak) mean-reversion. Since interest rates clearly do not have a constant volatility, a likelihood-based testing procedure which takes this phenomenon into account might be expected to yield more efficient estimates and hence more powerful tests.

Previous work in this area is by Ling and Li (1997, 1998) and Rahbek (1999), who consider tests for a unit autoregressive root in models with GARCH errors. They find that the maximum likelihood estimator of the mean-reversion parameter has a limiting distribution that is a weighted average of a Dickey-Fuller-type distribution and a normal distribution. They consider GARCH processes with fixed parameters in the stationarity region, whereas in this paper we study the case where the volatility parameters approach the unit root bound. Therefore, we consider parameter sequences such that the autoregressive root in the volatility process approaches unity as the sample size increases. This allows us to use the results of Nelson (1990) on continuous-time diffusion limits of GARCH processes. The present paper is also closely related to Hansen (1992b, 1995), who considers ordinary least-squares, generalized least-squares and adaptive estimation of regressions with non-stationary volatility.

The outline of the remainder of the paper is as follows. In Section 2, we define the model and hypothesis, and the parameter sequences that will be used in the asymptotic analysis. Section 3 analyses the likelihood function, the score and the information, and their asymptotic distribution under the relevant probability measures. We study the asymptotic distributions of the Dickey-Fuller test statistic, based on least-squares estimation, and the likelihood ratio test statistic, both under the null hypothesis and under local alternatives. Section 4 provides numerical evidence on the local power of these tests. In Section 5 we investigate the relevance of these local power results in finite samples. Section 6 contains an empirical illustration concerning the term structure of interest rates in the Netherlands, and Section 7 concludes.

2 The Model

Consider a univariate first-order autoregressive process with GARCH(1,1) innovations:

$$\Delta X_t = \gamma(X_{t-1} - \mu) + \varepsilon_t, \quad t = 1, \dots, n, \quad (1)$$

$$\varepsilon_t = \sigma_t \eta_t, \quad (2)$$

$$\sigma_t^2 = \omega + \alpha \varepsilon_{t-1}^2 + \beta \sigma_{t-1}^2, \quad (3)$$

$$\eta_t \sim \text{i.i.d. } N(0, 1), \quad (4)$$

where $\Delta X_t = (X_t - X_{t-1})$, and where X_0 , ε_0 and σ_0^2 are fixed.

The parameter γ describes the degree of mean-reversion. If $-2 < \gamma < 0$, then X_t reverts back to its mean μ . The null hypothesis that we wish to test is the unit root hypothesis, or equivalently the no-mean-reversion hypothesis

$$\mathcal{H}_0 : \gamma = 0, \quad (5)$$

which is tested against the alternative $\gamma < 0$. The model (1) has a restricted constant term, such that under the null hypothesis the process does not contain a drift. Other specifications of the deterministic component in X_t can be considered, including a restricted linear trend term (to test a random walk with drift against a trend-reverting autoregression), but this is not considered explicitly here. Similarly, the model can be extended to allow for more lags in (1).

The (nonnegative) parameters ω , α and β characterize the dynamics of the volatility process. If $\alpha + \beta < 1$, then the variance reverts back to its mean $\sigma^2 = \omega/(1 - \alpha - \beta)$, and if $\alpha + \beta = 1$ then the variance follows a random walk (with drift if $\omega \neq 0$). The asymptotic distribution of the test statistics considered in the next section will depend on what we assume about the parameter of interest γ , but also on assumptions about the volatility parameters (ω, α, β) . We consider two alternative assumptions (in all cases $\omega > 0, (\alpha, \beta) \geq 0$):

Assumption 1 For all $n \geq 1$, $\gamma_n = \kappa/n$ and $\alpha + \beta < 1$, with $(\kappa, \mu, \omega, \alpha, \beta)$ fixed.

Assumption 2 For all $n \geq 1$, $\gamma_n = \kappa/n$, $\alpha_n + \beta_n = 1 + \lambda/n$, $\omega_n = \varpi/n$ and $\alpha_n = \zeta/\sqrt{2n}$, with $(\kappa, \mu, \varpi, \lambda, \zeta)$ fixed, such that $\varpi > 0$, $\zeta > 0$ and $\lambda < \zeta^2/2$.

Under Assumption 1, the process X_t is near-integrated with stationary volatility. The unit root null hypothesis requires $\kappa = 0$, and values $\kappa \neq 0$ define the local alternatives. Under Assumption 2, the variance process is also near-integrated. One possible motivation for these parameter sequences is that the model (1)–(4) is viewed as a discrete-time approximation, for varying n but over a fixed time interval, of the continuous-time diffusion process defined below in Lemma 2, see Nelson (1990).

We conclude this section with two lemmas that describe the limiting behaviour of X_t under each of the two possible assumptions.

Lemma 1 Under Assumption 1, and as $n \rightarrow \infty$,

$$\left(\frac{1}{\sigma\sqrt{n}} \sum_{t=1}^{\lfloor \cdot \rfloor} \varepsilon_t, \frac{1}{\sigma\sqrt{n}} X_{\lfloor \cdot \rfloor} \right) \xrightarrow{\mathcal{L}} (W(\cdot), U(\cdot)), \quad (6)$$

in $D[0, 1]^2$, where $\sigma^2 = \omega/(1 - \alpha - \beta)$, $W(\cdot)$ is a standard Brownian motion process on $[0, 1]$, and $U(\cdot)$ is an Ornstein-Uhlenbeck process on $[0, 1]$:

$$dU(s) = \kappa U(s) ds + dW(s), \quad U(0) = 0. \quad (7)$$

The proof of this lemma is given in Ling and Li (1998, Theorem 3.3) for $\kappa = 0$, in which case $U(\cdot)$ reduces to $W(\cdot)$. This is extended to the case $\kappa \neq 0$ by writing $X_{\lfloor \cdot \rfloor}$ as a continuous functional of the partial sum of ε_t .

Lemma 2 Under Assumption 2, and as $n \rightarrow \infty$,

$$\left(\frac{1}{\sqrt{n}} \sum_{t=1}^{\lfloor \cdot \rfloor} \eta_t, \frac{1}{\sqrt{2n}} \sum_{t=1}^{\lfloor \cdot \rfloor} (\eta_t^2 - 1), \frac{1}{\sqrt{n}} X_{\lfloor \cdot \rfloor}, \sigma_{\lfloor \cdot \rfloor}^2 \right) \xrightarrow{\mathcal{L}} (W_1(\cdot), W_2(\cdot), Y(\cdot), V(\cdot)), \quad (8)$$

in $D[0, 1]^4$, where $(W_1(\cdot), W_2(\cdot))$ is a standard bivariate Brownian motion process on $[0, 1]$, and $(Y(\cdot), V(\cdot))$ is the solution to the system of stochastic differential equations

$$dY(s) = \kappa Y(s)ds + V(s)^{1/2}dW_1(s), \quad (9)$$

$$dV(s) = [\lambda V(s) + \varpi]ds + \zeta V(s)dW_2(s), \quad (10)$$

with $Y(0) = 0$ and $V(0) = \sigma_0^2$.

The proof of this lemma follows from Nelson (1990, Theorem 2.2 and Section 2.3). The difference again is that Nelson considers the case $\kappa = 0$, but the extension of his proof to the present case is straightforward. If the process $Y(s)$ is discretely sampled at times $s = t/n$, and we define $X_t = \sqrt{n}Y(t/n)$, $t = 0, 1, \dots, n$, then the actual process generating X_t may be approximated by (1)–(4) under Assumption 2; the approximation error will vanish as $n \rightarrow \infty$, see Nelson (1990). An alternative (Euler) approximation would lead to a discrete-time stochastic volatility-type model, but we choose to work with the GARCH model because it has a closed-form expression for the likelihood function, which simplifies the construction of likelihood-based test statistics considered in the next section.

3 Likelihood Analysis

The statistical analysis of model (1)–(4) is given in Ling and Li (1997, 1998) and Rahbek (1999), but will be briefly repeated here.

It will be convenient to introduce the parameter vector $\delta = (\gamma, -\gamma\mu)'$ and $Z_t = (X_{t-1}, 1)'$, such that (1) becomes $\Delta X_t = \delta' Z_t + \varepsilon_t$, and the null hypothesis is $\mathcal{H}_0 : \delta = 0$. The full parameter vector is $\theta = (\delta', \omega, \alpha, \beta)'$, and the log-likelihood function is

$$\ell(\theta) = \sum_{t=1}^n \ell_t(\theta) = \sum_{t=1}^n -\frac{1}{2} \left(\log 2\pi + \log \sigma_t^2(\theta) + \frac{\varepsilon_t^2(\delta)}{\sigma_t^2(\theta)} \right), \quad (11)$$

where $\varepsilon_t(\delta) = \Delta X_t - \delta' Z_t$, and where it should be noted that $\sigma_t^2(\theta)$ depends on the volatility parameters (ω, α, β) , but also, via ε_{t-1}^2 , on the regression parameters δ . The log-likelihood is conditional on σ_0 and ε_0 , which are not observed. In practice, they may be replaced by suitable estimates (we will assume that this has an asymptotically negligible effect).

The unrestricted parameter space for θ is $\Theta = \mathbb{R}^2 \times \mathbb{R}^+ \times \mathbb{R}^+ \times \mathbb{R}^+$, and the restricted parameter space defined by the null hypothesis is $\Theta_0 = (0, 0) \times \mathbb{R}^+ \times \mathbb{R}^+ \times \mathbb{R}^+$. Define $\hat{\theta} = \operatorname{argmax}_{\theta \in \Theta} \ell(\theta)$ and $\tilde{\theta} = \operatorname{argmax}_{\theta \in \Theta_0} \ell(\theta)$, the unrestricted and restricted maximum likelihood estimators, respectively. The likelihood ratio statistic for the null hypothesis is

$$LR = -2 \left(\ell(\tilde{\theta}) - \ell(\hat{\theta}) \right). \quad (12)$$

We will compare the performance of this test with that of Dickey and Fuller's (1981) F -statistic:

$$\Phi_1 = \frac{n-2}{2} \frac{\sum_{t=1}^n \Delta X_t Z_t' \left(\sum_{t=1}^n Z_t Z_t' \right)^{-1} \sum_{t=1}^n Z_t \Delta X_t}{\sum_{t=1}^n (\Delta X_t - \hat{\delta}'_{LS} Z_t)^2}, \quad (13)$$

with $\hat{\delta}_{LS} = (\sum_{t=1}^n Z_t Z_t')^{-1} \sum_{t=1}^n Z_t \Delta X_t$; this is a monotonic transformation of the likelihood ratio statistic for \mathcal{H}_0 under the restriction $\alpha = \beta = 0$ (i.e., homoskedastic innovations).

Define the score vector $S(\theta) = \partial \ell(\theta) / \partial \theta$ and the observed information matrix $J(\theta) = -\partial^2 \ell(\theta) / \partial \theta \partial \theta'$. Conventional Taylor series expansions (corresponding to a quadratic approximation of the log-likelihood function) result in

$$\begin{aligned} LR &= \hat{\theta}' E_1 [E_1' J(\theta_0)^{-1} E_1]^{-1} E_1' \hat{\theta} + o_P(1) \\ &= (n^{-1} \nu' + S(\theta_0)' J(\theta_0)^{-1} E_1) [E_1' J(\theta_0)^{-1} E_1]^{-1} (E_1' J(\theta_0)^{-1} S(\theta_0) + n^{-1} \nu) + o_P(1), \end{aligned} \quad (14)$$

where θ_0 is the true value (which is a sequence under Assumption 1 or 2), $E_1 = [I_2 : 0]'$ is a selection matrix such that $\delta = E_1' \theta$, and ν is the normalized distance between the true and hypothesized value of δ :

$$\nu = n \left(\begin{pmatrix} \kappa/n \\ -\kappa\mu/n \end{pmatrix} - \begin{pmatrix} 0 \\ 0 \end{pmatrix} \right) = \kappa \begin{pmatrix} 1 \\ -\mu \end{pmatrix}. \quad (15)$$

Therefore, we need to find an expression for $S(\cdot)$ and $J(\cdot)$, and evaluate their joint asymptotic behaviour under either Assumption 1 or 2.

Let $\xi = (\omega, [\alpha + \beta], \alpha)'$, the (linearly transformed) GARCH parameters, and $w_t(\theta) = (1, \sigma_{t-1}^2(\theta), \varepsilon_{t-1}^2(\delta) - \sigma_{t-1}^2(\theta))'$. The following results are useful ingredients for the score vector:

$$\frac{\partial \ell_t(\theta)}{\partial \sigma_t^2(\theta)} = \frac{1}{2\sigma_t^2(\theta)} \left(\frac{\varepsilon_t^2(\beta)}{\sigma_t^2(\theta)} - 1 \right) = \frac{1}{2\sigma_t^2(\theta)} (\eta_t^2(\theta) - 1), \quad (16)$$

$$\frac{\partial \sigma_t^2(\theta)}{\partial \delta} = \beta \frac{\partial \sigma_{t-1}^2(\theta)}{\partial \delta} - 2\alpha \varepsilon_{t-1}(\delta) Z_{t-1} = -2\alpha \sum_{i=1}^{t-1} \beta^{i-1} \varepsilon_{t-i}(\delta) Z_{t-i}, \quad (17)$$

$$\frac{\partial \sigma_t^2(\theta)}{\partial \xi} = \beta \frac{\partial \sigma_{t-1}^2(\theta)}{\partial \xi} + w_t(\theta) = \sum_{i=0}^{t-1} \beta^i w_{t-i}(\theta), \quad (18)$$

where $\eta_t(\theta) = \varepsilon_t(\delta) / \sigma_t(\theta)$. Here we use the fact that a fixed start-up value for σ_0^2 implies $\partial \sigma_0^2 / \partial \delta = 0$ and $\partial \sigma_0^2 / \partial \xi = 0$. Thus we find

$$S_\delta(\theta) = \frac{\partial \ell(\theta)}{\partial \delta} = \sum_{t=1}^n \left(Z_t \frac{\varepsilon_t(\delta)}{\sigma_t^2(\theta)} - \frac{\alpha}{\sigma_t^2(\theta)} (\eta_t^2(\theta) - 1) \sum_{i=1}^{t-1} \beta^{i-1} \varepsilon_{t-i}(\delta) Z_{t-i} \right), \quad (19)$$

$$S_\xi(\theta) = \frac{\partial \ell(\theta)}{\partial \xi} = \sum_{t=1}^n \left(\frac{1}{2\sigma_t^2(\theta)} (\eta_t^2(\theta) - 1) \sum_{i=0}^{t-1} \beta^i w_{t-i}(\theta) \right). \quad (20)$$

Expressions for the blocks $J_{\delta\delta}$, $J_{\delta\xi}$ and $J_{\xi\xi}$ of the information matrix can be derived from this. We shall not give explicit expressions here, but only provide their limiting behaviour in the next lemma, see Ling and Li (1998).

Lemma 3 *Under Assumption 1, and as $n \rightarrow \infty$,*

$$\frac{1}{\sqrt{n}} \sum_{t=1}^{\lfloor n \rfloor} \left(\frac{\varepsilon_t}{\sigma_t^2} - \frac{\alpha}{\sigma_t^2} (\eta_t^2 - 1) \sum_{i=1}^{t-1} \beta^{i-1} \varepsilon_{t-i} \right) \xrightarrow{\mathcal{L}} \tau B(\cdot), \quad (21)$$

in $D[0, 1]$, jointly with Lemma 1, where

$$\tau^2 = E \left[\frac{1}{\sigma_t^2} + 2\alpha^2 \sum_{i=1}^{\infty} \beta^{2(i-1)} \frac{\varepsilon_{t-i}^2}{\sigma_t^4} \right], \quad (22)$$

and $(W(\cdot), B(\cdot))$ is a bivariate vector Brownian motion process with $\text{var}(W(1)) = \text{var}(B(1)) = 1$ and $\text{cov}(W(1), B(1)) = \rho = \frac{1}{\sigma\tau}$. Letting $D_{1n} = \text{diag}((\sigma n)^{-1}, n^{-1/2})$ and $F(s) = (U(s), 1)'$,

$$D_{1n}S_{\delta} \xrightarrow{\mathcal{L}} \tau \int_0^1 F(s)dB(s), \quad (23)$$

$$D_{1n}J_{\delta\delta}D_{1n} \xrightarrow{\mathcal{L}} \tau^2 \int_0^1 F(s)F(s)'ds, \quad (24)$$

Furthermore,

$$n^{-1/2}S_{\xi} \xrightarrow{\mathcal{L}} N(0, \Sigma), \quad n^{-1}J_{\xi\xi} \xrightarrow{P} \Sigma, \quad n^{-1/2}D_{1n}J_{\delta\xi} \xrightarrow{P} 0, \quad (25)$$

where Σ is a positive definite matrix.

These results leads to the following theorem, the proof of which is given in the Appendix:

Theorem 1 Under Assumption 1, and as $n \rightarrow \infty$,

$$\begin{aligned} LR &\xrightarrow{\mathcal{L}} \left(\int_0^1 F(s) \left[dB(s) + \frac{\kappa}{\rho} U(s) ds \right] \right)' \left[\int_0^1 F(s)F(s)' ds \right]^{-1} \\ &\quad \times \left(\int_0^1 F(s) \left[dB(s) + \frac{\kappa}{\rho} U(s) ds \right] \right), \end{aligned} \quad (26)$$

$$\begin{aligned} 2\Phi_1 &\xrightarrow{\mathcal{L}} \left(\int_0^1 F(s)[dW(s) + \kappa U(s) ds] \right)' \left[\int_0^1 F(s)F(s)' ds \right]^{-1} \\ &\quad \times \left(\int_0^1 F(s)[dW(s) + \kappa U(s) ds] \right). \end{aligned} \quad (27)$$

The limiting distribution of LR under the null hypothesis ($\kappa = 0$) depends on the nuisance parameter ρ . In practice this nuisance parameter can be estimated consistently by $\hat{\rho} = 1/\sqrt{\hat{\sigma}^2 \hat{\tau}^2}$, where $\hat{\sigma}^2 = \hat{\omega}/(1 - \hat{\alpha} - \hat{\beta})$ and $\hat{\tau}^2$ is the sample analog of (22). Although we have not been able to obtain an explicit formula for ρ in terms of α and β , an approximation yields

$$\rho(\alpha, \beta) \approx \tilde{\rho}(\alpha, \beta) = \sqrt{\frac{(1 - \alpha - \beta)(1 - \beta^2)}{(1 - \alpha - \beta + \alpha^2)(1 - \beta^2 + 2\alpha^2)}}, \quad (28)$$

which is obtained by replacing $\varepsilon_{t-i}^2/\sigma_t^2$ in (22) by 1, and using $E(1/\sigma_t^2) \approx (1 - \alpha - \beta + \alpha^2)/\omega$, which corresponds to the continuous-record stationary distribution of $1/\sigma_t^2$ obtained by Nelson (1990). In order to check the accuracy of this approximation, we estimate the expectation in (22) by the average, over 1000 Monte Carlo replications¹, of the sample mean corresponding to (22) with a sample size of 10,000. This is done for $\alpha + \beta \in \{0.1, 0.2, \dots, 0.9\}$ and $\alpha/(\alpha + \beta) \in \{0.1, 0.2, \dots, 1\}$. It appears

¹All numerical results have been obtained using Ox versions 2.2 and 3.0, see Doornik (2001).

that (28) somewhat underestimates the true correlation; from a log-linear regression of the actual ρ 's on $\tilde{\rho}(\alpha, \beta)$, we obtain the following adjusted approximation:

$$\hat{\rho}(\alpha, \beta) = \tilde{\rho}(\alpha, \beta)^{0.64}, \quad (29)$$

which is quite accurate, with a regression standard error of about 1%.

Next, the estimate of ρ can be used to obtain an asymptotic p -value, either by Monte Carlo simulation or by the Gamma approximation proposed by Boswijk and Doornik (1999). The power function depends, in addition to ρ , only on κ (it is invariant to σ). In the next section, we compare the power functions of the two statistics for two cases.

Consider now the asymptotic behaviour of the score vector and information matrix under Assumption 2:

Lemma 4 *Under Assumption 2, and as $n \rightarrow \infty$,*

$$\frac{1}{\sqrt{n}} \sum_{t=1}^{\lfloor n \rfloor} \left(\frac{\varepsilon_t}{\sigma_t^2} - \frac{\alpha}{\sigma_t^2} (\eta_t^2 - 1) \sum_{i=1}^{t-1} \beta^{i-1} \varepsilon_{t-i} \right) \xrightarrow{\mathcal{L}} \int_0^1 V(u)^{-1/2} dW_1(u), \quad (30)$$

in $D[0, 1]$, jointly with Lemma 2. Letting $D_{2n} = \text{diag}(n^{-1}, n^{-1/2})$ and $G(s) = (Y(s), 1)'$,

$$D_{2n} S_\delta \xrightarrow{\mathcal{L}} \int_0^1 G(s) V(s)^{-1/2} dW_1(s), \quad (31)$$

$$D_{2n} J_{\delta\delta} D_{2n} \xrightarrow{\mathcal{L}} \int_0^1 G(s) G(s)' V(s)^{-1} ds. \quad (32)$$

Furthermore, there exist non-singular norming matrices D_{3n} such that

$$D_{3n} S_\xi = O_P(1), \quad D_{3n} J_{\xi\xi} D_{3n} = O_P(1), \quad D_{2n} J_{\delta\xi} D_{3n} \xrightarrow{P} 0. \quad (33)$$

A proof is given in the Appendix. Note that the limiting Riemann integral in (32) is the quadratic variation of the stochastic integral in (31). The suitably normalized information matrix is block-diagonal in the limit, because the cross-variation between the two parts of the score vector is zero in the limit. These results imply:

Theorem 2 *Under Assumption 2, and as $n \rightarrow \infty$,*

$$LR \xrightarrow{\mathcal{L}} \left(\int_0^1 G(s) V(s)^{-1/2} [dW_1(s) + \kappa V(s)^{-1/2} Y(s) ds] \right)' \left[\int_0^1 G(s) G(s)' V(s)^{-1} ds \right]^{-1} \\ \times \left(\int_0^1 G(s) V(s)^{-1/2} [dW_1(s) + \kappa V(s)^{-1/2} Y(s) ds] \right), \quad (34)$$

$$2\Phi_1 \xrightarrow{\mathcal{L}} \left(\int_0^1 G(s) [V(s)^{1/2} dW_1(s) + \kappa Y(s) ds] \right)' \left[\int_0^1 G(s) G(s)' ds \int_0^1 V(s) ds \right]^{-1} \\ \times \left(\int_0^1 G(s) [V(s)^{1/2} dW_1(s) + \kappa Y(s) ds] \right). \quad (35)$$

The theorem is proved in the Appendix. The results are closely related to those obtained by Hansen (1992b, 1995), who considers ordinary least-squares, generalized least-squares and adaptive estimation of regressions with non-stationary volatility. Note that the likelihood ratio statistic is asymptotically equivalent to a Wald statistic based on weighted least-squares with known $\{\sigma_t^2\}$. Hansen shows that when the process generating the non-stationary volatility is unknown, it may be estimated non-parametrically, without loss of efficiency relative to a parametric likelihood analysis.

Both distributions in Theorem 2 depend on nuisance parameters, even under the null hypothesis ($\kappa = 0$). In principle they are affected by all volatility parameters (ϖ, λ, ζ) , although parameter variations that only affect the scale of $V(s)$ will leave the distributions in (34) and (35) unaffected. From Nelson (1990), it appears that the function λ/ζ^2 is most relevant, since it determines the stationary distribution of the volatility process. Unfortunately these parameters are not consistently estimable.

A possible solution to this nuisance parameter problem is to use the *conditional* asymptotic null distribution of the two test statistics, given the volatility process $V(\cdot)$. Although this process depends on the parameters (ϖ, λ, ζ) , these are variation independent of the parameter of interest κ , so that conditioning on V does not entail a loss of information on κ . In other words, V is S-ancillary for κ , see Barndorff-Nielsen (1978). Clearly the asymptotic distributions in (34) and (35) for fixed V will depend on the realization of V , and hence cannot be tabulated. However, given the independence between W_1 and V the conditional distribution is quite easy to simulate. In practice this will involve replacing the volatility process by its estimate $\hat{V}_n(s) = \hat{\sigma}_{[sn]}^2$, with $\{\hat{\sigma}_t^2\}$ the filtered estimate of $\{\sigma_t^2\}$ based on the maximum likelihood estimates of the GARCH parameters. The results of Nelson and Foster (1994) suggest that $\hat{V}_n(\cdot)$ converges in probability to $V(\cdot)$ in $D[0, 1]$, which in turn would imply that the estimated conditional distribution given $\{\hat{\sigma}_t^2\}_{t=1}^n$ converges to the true conditional distribution given $\{V(s), s \in [0, 1]\}$.

Before we proceed, it is of interest to discuss the difference of the two types of asymptotic approximation in Theorems 1 and 2 as $\alpha + \beta$ approaches 1, so that the unconditional variance σ^2 diverges. For fixed parameter values (Assumption 1) the approximation (28) suggests that $\rho \downarrow 0$ as $\alpha + \beta \uparrow 1$, which is confirmed by the fact that $\rho = 1/(\sigma\tau)$, and σ diverges whereas τ remains finite. This implies, first, that the limiting distribution of the *LR* statistic will approach the $\chi^2(2)$ distribution under the null hypothesis, because B and F become independent. Secondly, it shows that the local power will increase, and in fact approach 1 for all κ , because the non-centrality parameter in (26) is essentially κ/ρ . This suggests that in such cases the likelihood ratio test has infinite power superiority over the least-squares based test. Note, however, that Theorem 1 is only valid under Assumption 1, which involves the condition $\alpha + \beta < 1$; the quality of the asymptotic approximation might deteriorate as $\alpha + \beta \uparrow 1$. More importantly, if $\alpha \downarrow 0$ at the same time as $\alpha + \beta \uparrow 1$, then the above arguments are no longer valid, since $\lim_{\alpha \downarrow 0, \alpha + \beta \uparrow 1} \rho(\alpha, \beta)$ does not exist. This implies that for parameter values with $\alpha + \beta$ close to 1 and α close to 0, which are typically encountered with daily financial returns, this asymptotic approximation will not be reliable, and we should turn to the continuous-record asymptotic approximation implied by Assumption 2 instead.

Under Assumption 2, then, it is allowed that $\alpha + \beta = 1$ and hence $\lambda = 0$; no discontinuity in

the limit theory is to be expected around $\lambda = 0$, as long as $\lambda < \zeta^2/2$, which is the condition for strict stationarity of the limiting diffusion process $V(s)$. The main difference between the cases $\lambda < 0$ and $0 \leq \lambda < \zeta^2/2$ is that in the former case the disturbances ε_t have finite variance, whereas in the latter case the unconditional variance is infinite, since the limiting distribution of ε_t is Student's $t(2 - 4\lambda/\zeta^2)$, see Nelson (1990). For $\lambda = 0$ the limiting distribution of LR will not be $\chi^2(2)$ under the null hypothesis, since $V(s)^{-1/2}Y(s) = V(s)^{-1/2} \int_0^s V(u)^{1/2}dW_1(u)$ and $W_1(s)$ are not independent for $\lambda = 0$. As ζ increases however, the variation in $V(s)$ increases, and one might expect that the behaviour of $V(s)^{-1/2}Y(s)$ will be dominated by $V(s)$, such that it becomes independent of $W_1(s)$. From the expressions in (34) and (35) it is not clear that the relative power advantage of LR will increase with ζ ; this will be investigated in the next section.

In summary, the results in this section indicate we may expect a power gain of the likelihood ratio test over the Dickey-Fuller test when a large value of $\alpha + \beta$ (implying persistent volatility) is combined with a large value of α (implying a large short-run variation in the volatility). In the next section we investigate whether these predictions are reflected in the asymptotic local power behaviour of the tests, and in Section 5 we turn to the finite sample behaviour of the procedures.

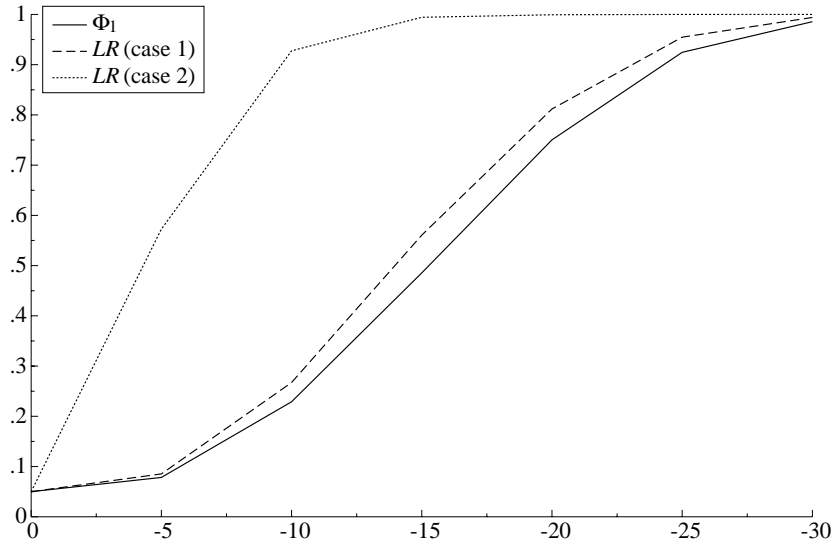
4 Local Power

In this section we provide some numerical evidence on the local power of the two alternative test statistics. First, we consider the case of stationary volatility (Assumption 1). We consider two sets of GARCH parameters:

1. $\alpha = 0.05$, $\beta = 0.9$ and $\sigma^2 = \omega/(1 - \alpha - \beta) = 1$, which implies $\rho = 0.967$ (the value of $\rho = 1/(\sigma\tau)$ is obtained by Monte Carlo simulation, as described in Section 3). These parameter values correspond to a relatively smooth GARCH process with strong persistence, as typically found in empirical data sets of daily returns. The high value of the correlation coefficient suggests that the power difference between the LR and Φ_1 test will be relatively small in this case.
2. $\alpha = 0.35$, $\beta = 0.6$ and $\sigma^2 = 1$, which implies $\rho = 0.570$. Again this leads to a rather slowly mean-reverting GARCH process, but now the higher value of α leads to more short-run variation in the volatility. The low value of ρ leads us to expect more power gains for the LR test in this case.

Figure 1 displays the local power function of the Φ_1 test, which is the same for both parameter combinations, and that of the LR statistic for each data-generating process. Note that the local power of Φ_1 is the same as the local power of LR when $\alpha = \beta = 0$ and hence $\rho = 1$, i.e., when there are no GARCH effects. All results are obtained by Monte Carlo simulation, using a discretization (1000 equidistant points) of the processes and integrals, and with 10,000 replications for the power calculations, and 100,000 replications for the critical values. The 5% critical values are 9.109 for $2\Phi_1$, and 9.028 (case 1) and 7.528 (case 2) for LR .

Figure 1: Local power of Φ_1 and LR with stationary volatility (size = 5%).



As expected, the power gain of the LR test relative to the least-squares-based Φ_1 is very small when $(\alpha, \beta) = (0.05, 0.9)$. This suggests that for such GARCH processes, one might as well use the conventional test. For the second parameter combination, however, the power gain is much larger. Therefore, these results confirm the prediction in the previous section that only when the volatility process has itself a high volatility (corresponding to a high value of α), the likelihood ratio test yields a substantial power gain over the least-squares based Dickey-Fuller test.

Next, we consider the local power function when the volatility process is near-integrated. In this case we consider four parameter configurations:

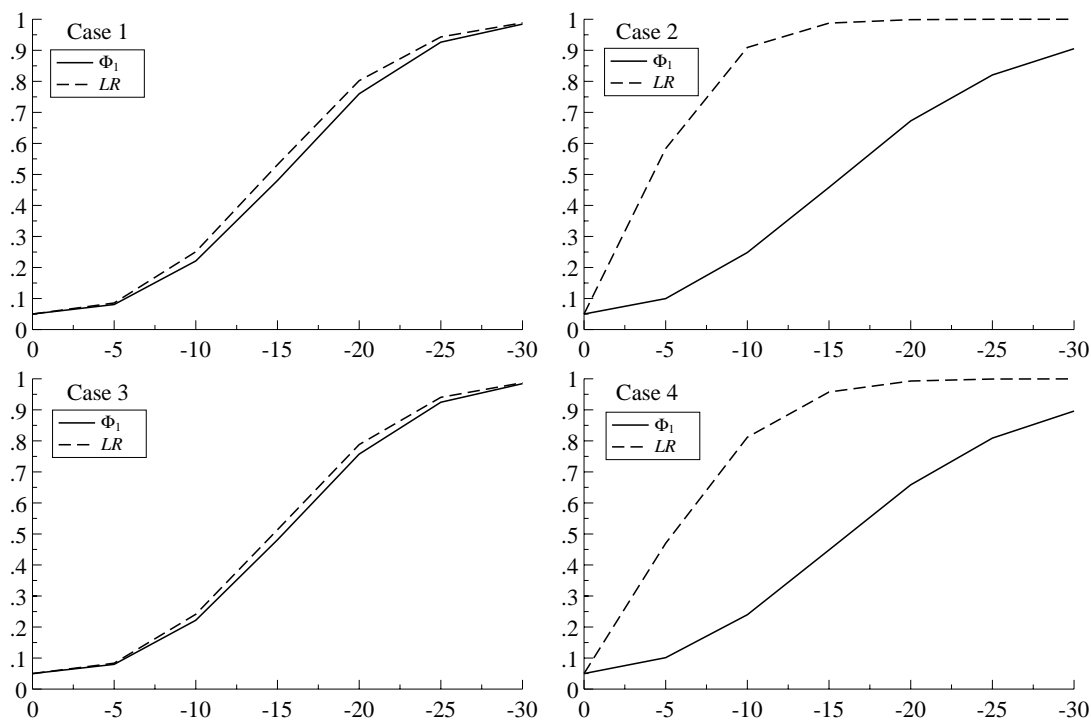
1. $\lambda = -100$, $\zeta = \sqrt{10}$. This corresponds to the first case $((\alpha_n, \beta_n) = (0.05, 0.9))$ considered above, with $n = 2000$.
2. $\lambda = -100$, $\zeta = 7\sqrt{10}$. For $n = 2000$, this corresponds to the second case above $((\alpha_n, \beta_n) = (0.35, 0.6))$, which leads to more variation in the volatility process.
3. $\lambda = -40$, $\zeta = 2$. This is a process with less mean-reversion in the volatility than case 1, but with the same value of $-\lambda/\zeta^2 = 10$; it corresponds to $(\alpha_n, \beta_n) = (0.05, 0.9)$ for $n = 800$. Therefore, we expect roughly the same results as in case 1.
4. $\lambda = -40$, $\zeta = 14$. This is comparable to case 2 (same $-\lambda/\zeta^2 = 10/49 \approx 0.2$), but with less mean-reversion in the volatility, and corresponds to $(\alpha_n, \beta_n) = (0.35, 0.6)$ for $n = 800$.

In all cases we set $\varpi = -\lambda$, such that $V(s)$ reverts to 1, but the results are invariant to ϖ , as long as the starting value is chosen appropriately. We use a fixed start-up value $V(0) = 1$, which corresponds to the expectation of the stationary distribution of $V(s)$; alternatively, one could draw $V(0)^{-1}$ from its stationary $\Gamma(1 - 2\lambda/\zeta^2, 2\varpi/\zeta^2)$, distribution, see Nelson (1990). An important difference between

cases 1 and 3 on the one hand, and cases 2 and 4 on the other hand, is the existence of higher moments of ε_t . Nelson's (1990) limiting $t(2 - 4\lambda/\zeta^2)$ distribution for the disturbances implies that ε_t has no finite integer moments beyond the variance in cases 2 and 4, whereas it has much higher moments (up to $2 - 4\lambda/\zeta^2 = 42$) in cases 1 and 3.

For each test, we perform a conditional and an unconditional version. The unconditional version involves Monte Carlo simulation of the 5% critical value for the given parameter combination (based on 100,000 replications), and defining the local power as the rejection frequency (based on 10,000 replications) at this critical value. In practice this is infeasible, since the volatility parameters are not known and not consistently estimable, but obtaining the critical values is of interest to investigate how sensitive they are to parameter variations. The 5% critical values for LR in the four cases are 9.095, 7.642, 9.139 and 8.072, respectively, whereas for $2\Phi_1$ they are 9.168, 10.179, 9.189 and 10.628, respectively. This confirms that the critical values are very similar for cases 1 and 3, and also for cases 2 and 4. Furthermore, it demonstrates that whereas the critical values of LR decrease with $-\zeta^2/\lambda$, the critical values of the Dickey-Fuller test *increase* with this quantity, so that the Φ_1 test will have the highest size distortions (if conventional critical values are used) in situations where the volatility-of-volatility parameter ζ is relatively high. In the conditional version of the test, we simulate the p -value (based on 1000 replications), for each of the 10,000 realizations of the test statistic, conditional on the actual volatility process for that realization, and reject when this p -value is less than 0.05. In practice the local power of the conditional and unconditional versions of the test turns out to be almost identical, so we only report the conditional versions in Figure 2.

Figure 2: Local power of Φ_1 and LR with near-integrated volatility (size = 5%).



The results clearly show the expected power gain of the LR test for cases 2 and 4, whereas the two tests are almost equivalent in cases 1 and 3. Furthermore, the power functions in cases 1 and 3 are very close to the corresponding case 1 in Figure 1 (using fixed-parameter asymptotics), and similarly the behaviour in cases 2 and 4 resembles the corresponding case 2 in Figure 1. The effect of λ is much weaker, although we see that the power is slightly lower for the small λ cases. Because the critical values are also fairly close to the corresponding fixed-parameter cases, the two types of asymptotics are largely in agreement. As one might expect, this agreement would break down for parameter values such that $\alpha + \beta = 1$, and hence $\lambda = 0$. The fixed-parameter asymptotic analysis, although not strictly applicable anymore, would suggest that $\rho = 0$, which would imply χ^2 critical values and an infinite power gain of the LR test. However, additional simulations indicate that the behaviour of the tests under $\alpha_n + \beta_n = 1$ and $\alpha_n = \zeta/\sqrt{2n}$ depends very much on ζ , comparable to cases 1–4 in Figure 2; only when $\zeta \rightarrow \infty$ the null distribution of LR approaches the $\chi^2(2)$ distribution, and the local power becomes 1 for all κ .

5 A Monte Carlo Experiment

In this section we consider the finite-sample behaviour of the tests in a small-scale Monte Carlo experiment. We consider $n \in \{250, 2000\}$, which would correspond to approximately 1 and 8 years of daily financial data. Here $n = 250$ may be considered a small sample for GARCH estimation; usually a number of years of daily data are considered. Next, we continue to consider the two near-integrated cases $(\alpha, \beta) = \{(0.05, 0.9), (0.35, 0.6)\}$; note that these are chosen the same for both sample sizes, so that one might expect a relatively better approximation by the stationary (fixed-parameter) asymptotic distributions for larger sample sizes. Furthermore we consider $\gamma_n = \kappa/n$ with $\kappa \in \{0, -5, -20\}$, to study both the size and power properties of the tests.

For the LR^2 and Φ_1 tests we compute two types of p -values, the first based on fixed-parameter asymptotics, and the second based on near-integrated asymptotics, conditional on the estimated $\{\hat{\sigma}_t^2\}$. For the fixed-parameter asymptotic p -values, we use the Gamma approximation of Boswijk and Doornik (1999), in combination with $\hat{\rho}(\alpha, \beta)$ given in (28)–(29), where α and β are replaced by their unrestricted ML estimates. Finally, we also consider QLR , the quasi-likelihood ratio test based on the assumption that $\varepsilon_t \sim \text{i.i.d. } t(\nu)$. This test is included to see whether the same power gain can be obtained by a test that correctly specifies the marginal distribution of the disturbances, although it misspecifies the volatility dynamics. From, e.g., Lucas (1997), it follows that when ε_t is indeed i.i.d. $t(\nu)$, the asymptotic distribution of QLR is the same as that of LR in (26), but with ρ the correlation between ε_t and $-\partial \log f_\nu(\varepsilon_t)/\partial \varepsilon_t$, where f_ν denotes the $t(\nu)$ density. It can be shown that for $\nu > 2$, $\rho^2 = (\nu^2 + \nu - 6)/(\nu^2 + \nu)$. Replacing ν by the unrestricted ML estimate, this yields again a p -value for this test using the Gamma approximation.

The results, based on 2000 replications, are given in Table 1, and give rise to the following conclusions. First, we see that the fixed-parameter asymptotic p -values give a better approximation for the small α generating process, but a worse approximation for the $\alpha = 0.35$ case. In the latter case,

²In the likelihood function, the volatility process is initialized using $\hat{\sigma}_0^2 = \hat{\varepsilon}_0^2 = n^{-1} \sum_{t=1}^n \Delta X_t^2$.

QLR has a rather large size distortion for $n = 250$, but this seems to vanish as n increases. In fact, for $n = 2000$ there hardly seem to be any substantial size distortions left, with the exception of the Φ_1 test using ordinary critical values when α is large. The power behaviour of the LR and Φ_1 tests is as predicted by the asymptotic analysis. For small α , the power of the tests is virtually identical, so that there is not much gain in using the LR test. When α is large on the other hand, the power gain is quite clear, especially for the larger sample size. Finally, we note that for this case the power curve of the Student t -based QLR test lies between that of Φ_1 and LR : although there is a clear gain in fitting the marginal tail behaviour of the disturbances by a Student's t instead of a normal distribution, we see that the misspecification of the conditional variance gives this test a power disadvantage relative to the GARCH-based LR test.

Table 1: Rejection frequencies of LR , Φ_1 and QLR

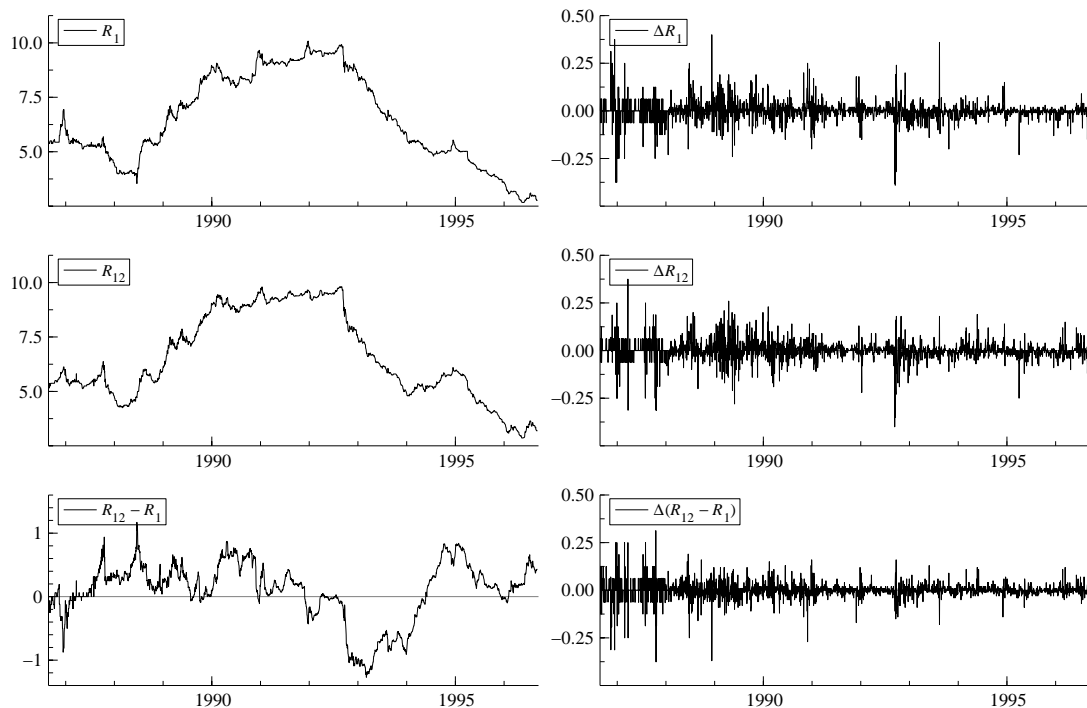
		$n = 250$			$n = 2000$		
		$\kappa = 0$	$\kappa = -5$	$\kappa = -20$	$\kappa = 0$	$\kappa = -5$	$\kappa = -20$
$\alpha = 0.05,$ $\beta = 0.9$	LR , fixed	0.059	0.086	0.784	0.055	0.082	0.794
	LR , near-int.	0.067	0.094	0.809	0.061	0.091	0.813
	Φ_1 , fixed	0.059	0.086	0.775	0.054	0.078	0.752
	Φ_1 , near-int.	0.067	0.095	0.798	0.058	0.085	0.770
	QLR	0.062	0.085	0.760	0.054	0.079	0.755
$\alpha = 0.35,$ $\beta = 0.6$	LR , fixed	0.074	0.248	0.823	0.049	0.448	0.989
	LR , near-int.	0.063	0.249	0.926	0.040	0.417	0.990
	Φ_1 , fixed	0.099	0.150	0.755	0.073	0.127	0.739
	Φ_1 , near-int.	0.070	0.120	0.689	0.052	0.114	0.705
	QLR	0.115	0.190	0.754	0.055	0.234	0.913

6 An Empirical Application

In this section we apply the GARCH-unit root test to the term structure of interest rates in the Netherlands. In particular, we analyse daily observations on the 1-month and 12-month Amsterdam InterBank Offered Rate (AIBOR) over the 10-year period from September 4, 1986 through September 4, 1996, obtained from Datastream (note that the AIBOR has ceased to exist as a separate interest rate since the start of the Euro and the associated EURIBOR). Accounting for missing observations, this gives a total of 2610 observations. The two separate interest rates R_1 and R_{12} , the yield spread ($R_{12} - R_1$), and the first difference of these series are depicted in Figure 3.

We observe that the two interest rates display very slow (if any) mean reversion. The yield spread seems to display stronger mean reversion, which would imply cointegration between the two rates if they are individually integrated of order 1. See Campbell *et al.* (1997, Chapter 10) for a discussion of the connection between term structure models and cointegration of interest rates.

Figure 3: Daily 1-month and 12-month AIBOR and their difference, level and first difference, September 4, 1986 through September 4, 1996.



The graphs of the first differences of the three series clearly display volatility clustering, suggesting that the unit root test analysed in this paper might be useful. Another property that is evident, in particular in the first few years, is that interest rate changes are essentially discrete (they are often equal to zero), so that GARCH models with Gaussian innovations will be necessarily misspecified in this respect, and the corresponding LR test is based on a quasi-likelihood.

Exactly the same data-set was analysed in a bivariate cointegration model by Boswijk and Lucas (2001). They found that allowing for non-normality of the innovations in the likelihood function leads to more evidence of cointegration or even stationarity than by assuming the innovations to be independent Gaussian (and hence using the Johansen (1991) likelihood ratio test). We now consider whether a similar result is obtained by allowing explicitly for the volatility clustering in the data.

Because of the short-run dynamics in the mean displayed by the data, we base the tests on a sixth order autoregressive specification for the mean, rewritten as

$$\Delta X_t = \delta' Z_t + \sum_{i=1}^5 \phi_i \Delta X_{t-i} + \varepsilon_t, \quad (36)$$

where $Z_t = (X_{t-1}, 1)'$, augmented with a GARCH(1,1) specification for the variance. In all cases five lagged differences is sufficient (and in two cases it is necessary) to obtain standardized residuals free of autocorrelation. Although the asymptotic properties have only been derived in this paper for the first-order autoregressive model, we expect the same asymptotic distribution theory to apply to tests for $\delta = 0$ in (36). In all cases the squared standardized residuals display no significant autocorrelation,

indicating that the GARCH(1,1) specification is sufficient to capture the volatility clustering. The test results are given in Table 2.³

Table 2: Test results for interest rate data

	Φ_1	$p_1(\Phi_1)$	$p_2(\Phi_1)$	LR	$p_1(LR)$	$p_2(LR)$	$\hat{\alpha}$	$\hat{\beta}$	$\hat{\rho}$
R_1	0.286	0.980	0.994	8.691	0.057	0.061	0.090	0.867	0.927
R_{12}	0.228	0.988	0.999	5.405	0.123	0.191	0.170	0.823	0.566
$R_{12} - R_1$	2.905	0.213	0.432	14.657	0.001	0.004	0.204	0.796	0.000

Note: The test statistics, $\hat{\alpha}$, $\hat{\beta}$ and $\hat{\rho}$ are defined in the text; $p_1(\cdot)$ denotes the p -value based on fixed-parameter asymptotics, using the Gamma approximation of Boswijk and Doornik (1999), together with (28)–(29); $p_2(\cdot)$ is the simulated p -value based on near-integrated asymptotics, using the estimated volatility process and 10, 000 replications.

Consider first the results for the 1-month AIBOR R_1 . The (augmented) Dickey-Fuller test has a very large p -value, whether one uses the conventional null distribution (applicable for fixed α and $\alpha + \beta < 1$) or the asymptotic null distribution implied by Theorem 2 (using near-integrated asymptotics). On the other hand, the likelihood ratio test based on the GARCH(1,1) likelihood function has a much larger test statistic, with p -values close to 5%. Note that this large difference is somewhat unexpected, since the estimated GARCH parameters are such that the corresponding correlation parameter ρ is very close to one. The small p -value of LR may however not be taken as evidence of mean-reversion, because the unrestricted maximum likelihood estimate of the mean-reversion parameter γ is *positive*, indicating mean-aversion, so that a one-sided test of $\gamma = 0$ against $\gamma < 0$ would not lead to rejection of the null hypothesis, regardless of the significance level used.

A similar conclusion applies to the 12-month AIBOR R_{12} . Again the LR statistic is substantially larger than the Dickey-Fuller statistic, and its p -value is correspondingly smaller, but again this is not evidence of mean reversion because the estimate of γ is positive. Note that the estimates of the GARCH parameters are such that $\hat{\rho} = 0.566$ which would suggest a large power potential for the LR test based on the evidence in the previous sections. Note also that the two p -values of LR differ quite a bit; given the fairly large value of $\hat{\alpha}$, the simulation evidence in the previous section would lead us to trust the p -value from near-integrated asymptotics ($p_2 = 0.191$) better than the fixed-parameter asymptotic p -value $p_1 = 0.123$.

For the yield spread series $R_{12} - R_1$, a different conclusion emerges. The Dickey-Fuller test statistic has a conventional asymptotic p -value larger than 20%, which would lead us not to reject a unit root at conventional significance levels. In fact, because the estimated GARCH model has $\hat{\alpha} + \hat{\beta} = 1$,⁴ we may expect the conventional p -value to be invalid; the p -value of Φ_1 based on Theorem 2 is about 43%, indicating even less evidence of mean-reversion. In contrast, the GARCH-LR test has a p -value

³The empirical results in this section have been obtained using PcGive version 10.0, in particular the Garch package; see Doornik and Hendry (2001).

⁴The results in the table are based on maximization of the likelihood under the inequality constraint $\alpha + \beta \leq 1$; unconstrained maximization leads to very similar results, except that $\hat{\alpha} + \hat{\beta}$ is slightly larger than one, so that we cannot use (28)–(29) anymore.

of $p_2 = 0.004$, indicating quite clear rejection of the null hypothesis (the value of $p_1 = 0.001$ is based on the $\chi^2(2)$ null distribution corresponding to $\rho = 0$, which is a limiting case outside the parameter space where Theorem 1 is applicable). In this case the estimated mean reversion parameter equals -0.0041 (with a standard error of 0.0011), which is very small in absolute value but has the right sign; the estimated mean of the yield spread is $\hat{\mu} = 0.317$ (with a standard error of 0.143).

In summary, we find for these series that the conventional Dickey-Fuller test would lead one to conclude that the two interest rates are integrated of order 1 and not cointegrated (or at least not with cointegrating vector $(1, -1)'$), whereas a GARCH-based likelihood ratio test clearly indicates stationarity of the yield spread and hence cointegration of the two interest rates. Given the economic plausibility of the latter result, this suggests that the GARCH-based test is able to exploit the information on mean-reversion in the data more efficiently.

7 Discussion

In this paper we have investigated likelihood ratio testing for a unit root when the innovations follow a near-integrated GARCH process. We have analysed the asymptotic null distribution and local power function of the likelihood ratio test and the least-squares based Dickey-Fuller test, both under fixed GARCH parameters and under near-integrated sequences. It has been found that the two types of asymptotics are largely in agreement, as long as the sum of the GARCH parameters is less than one. A considerable power gain potential for the LR test has been found to occur with GARCH processes with a large short-run variation in the volatility, corresponding to a heavy-tailed marginal distribution of the innovations. These asymptotic results have been shown to be reflected in the finite sample behaviour of the tests. Furthermore, an empirical application has demonstrated that the possible power gain can lead one to find more evidence in favour of stationarity by explicitly modelling the volatility clustering.

The main limitation of the present analysis is that it applies exclusively to a GARCH(1,1) specification of the volatility. The sensitivity of the inferences to volatility misspecification remains to be investigated. Nelson and Foster's (1994) result that the GARCH-filtered estimator of the volatility process is consistent, as the sampling interval goes to zero, for a large class of volatility processes, suggests that the GARCH-based unit root test may be reasonably robust against deviations of the GARCH(1,1) specification, at least for high-frequency data. On the other hand, if an alternative volatility specification (such as EGARCH or stochastic volatility) is preferred, then one could analyse the corresponding likelihood ratio test. Although the relevant asymptotic theory is not covered by the present paper, the result of Theorem 2 and its proof suggest that very similar results can be obtained for alternative specifications, as long as they have a continuous-time diffusion limit. Special care will have to be taken with the conditional version of the test (conditioning on the estimated volatility process) when the two Brownian motions driving the process are correlated, which occurs for example in EGARCH models with asymmetric (leverage) effects.

The analysis of this paper can be extended in various other directions. One example would be to allow the standardized innovations to follow another distribution than the standard Gaussian, to allow for

the excess kurtosis often found in empirical applications. Another obvious extension is to cointegration analysis in multivariate models. Here the main problem may be to find a flexible but parsimonious multivariate volatility specification; one possible approach is analysed by Li *et al.* (1998). We intend to pursue these extensions in future research.

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Appendix

Proof of Theorem 1. Consider first the limiting distribution of LR . Let $D_{1n}^* = \text{diag}(D_{1n}, n^{-1/2}I_3)$, so that $E_1' D_{1n}^* = D_{1n} E_1'$. Therefore, (14) implies, with $S = S(\theta_0)$ and $J = J(\theta_0)$,

$$LR = [n^{-1}\nu' D_{1n}^{-1} + S' D_{1n}^* (D_{1n}^* J D_{1n}^*)^{-1} E_1] [E_1' (D_{1n}^* J D_{1n}^*)^{-1} E_1]^{-1} \times [E_1' (D_{1n}^* J D_{1n}^*)^{-1} D_{1n}^* S + D_{1n}^{-1} n^{-1} \nu] + o_P(1). \quad (\text{A.1})$$

Lemma 3 yields

$$D_{1n}^* S \xrightarrow{\mathcal{L}} \begin{pmatrix} \tau \int_0^1 F(s) dB(s) \\ N(0, \Sigma) \end{pmatrix}, \quad D_{1n}^* J D_{1n}^* \xrightarrow{\mathcal{L}} \begin{bmatrix} \tau^2 \int_0^1 F(s) F(s)' ds & 0 \\ 0 & \Sigma \end{bmatrix}, \quad (\text{A.2})$$

and clearly $D_{1n}^{-1} n^{-1} \nu \rightarrow (\sigma\kappa, 0)'$. Combining these results gives

$$LR \xrightarrow{\mathcal{L}} \begin{aligned} & \left[(\sigma\kappa, 0) + \tau \int_0^1 dB(s) F(s)' \left(\tau^2 \int_0^1 F(s) F(s)' ds \right)^{-1} \right] \left[\tau^2 \int_0^1 F(s) F(s)' ds \right] \\ & \times \left[\left(\tau^2 \int_0^1 F(s) F(s)' ds \right)^{-1} \tau \int_0^1 F(s) dB(s) + (\sigma\kappa, 0)' \right] \\ & = \left[(\sigma\tau\kappa, 0) \int_0^1 F(s) F(s)' ds + \int_0^1 dB(s) F(s)' \right] \left[\int_0^1 F(s) F(s)' ds \right]^{-1} \\ & \times \left[\int_0^1 F(s) F(s)' ds (\sigma\tau\kappa, 0)' + \int_0^1 F(s) dB(s) \right], \end{aligned} \quad (\text{A.3})$$

and using $\sigma\tau = 1/\rho$, this yields (26). For Φ_1 , we use Lemma 1 together with the continuous mapping theorem to yield

$$D_{1n} \sum_{t=1}^n Z_t Z_t' D_{1n} \xrightarrow{\mathcal{L}} \int_0^1 F(s) F(s)' ds, \quad (\text{A.4})$$

and

$$\begin{aligned} D_{1n} \sum_{t=1}^n Z_t \Delta X_t &= D_{1n} \sum_{t=1}^n Z_t Z_t' D_{1n} D_{1n}^{-1} \delta + D_{1n} \sum_{t=1}^n Z_t \varepsilon_t \\ &\xrightarrow{\mathcal{L}} \int_0^1 F(s) F(s)' ds (\sigma\kappa, 0)' + \int_0^1 F(s) dW(s) \sigma \\ &= \sigma \int_0^1 F(s) [dW(s) + \kappa U(s)] ds. \end{aligned} \quad (\text{A.5})$$

Furthermore,

$$\begin{aligned} \hat{\sigma}_{LS}^2 &= \frac{1}{n-2} \sum_{t=1}^n (\Delta X_t - \hat{\delta}'_{LS} Z_t)^2 \\ &= \frac{1}{n-2} \left(\sum_{t=1}^n \varepsilon_t^2 - \sum_{t=1}^n \varepsilon_t Z_t' D_{1n} \left(D_{1n} \sum_{t=1}^n Z_t Z_t' D_{1n} \right)^{-1} D_{1n} \sum_{t=1}^n Z_t \varepsilon_t \right) \\ &= \frac{1}{n-2} \sum_{t=1}^n \varepsilon_t^2 + o_p(1), \end{aligned} \quad (\text{A.6})$$

which converges in probability to σ^2 . Collecting the results yields (27). \square

Proof of Lemma 4. Write the first term of (30) as

$$\frac{1}{\sqrt{n}} \sum_{t=1}^{\lfloor sn \rfloor} \frac{\eta_t}{\sigma_t} = \int_0^s \sigma_{\lfloor un \rfloor}^{-1} dW_{1n}(u), \quad (\text{A.7})$$

where $W_{1n}(s) = n^{-1/2} \sum_{t=1}^{\lfloor sn \rfloor} \eta_t$. From Lemma 2, $(W_{1n}(\cdot), \sigma_{\lfloor \cdot n \rfloor}^2)$ converges weakly to $(W_1(\cdot), V(\cdot))$. Nelson (1990, Theorem 2.3) shows that $V(\cdot)$ is stationary if $\lambda < \zeta^2/2$ and $\varpi > 0$, and that under those conditions $\sigma_{\lfloor \cdot n \rfloor}^{-2}$ converges weakly to $V(\cdot)^{-2}$ (and hence $\sigma_{\lfloor \cdot n \rfloor}^{-1} \xrightarrow{\mathcal{L}} V(\cdot)^{-1/2}$). Since $\{\eta_t\}$ are i.i.d. $N(0, 1)$, the conditions of Hansen (1992a) apply, and $\int_0^s \sigma_{\lfloor un \rfloor}^{-1} dW_{1n}(u) \xrightarrow{\mathcal{L}} \int_0^s V(u)^{-1/2} dW_1(u)$.

Write the remainder of (30) as $n^{1/2} \sum_{i=1}^{\lfloor sn \rfloor} v_t / \sigma_t$, where v_t is a martingale difference sequence with variance $2\alpha_n^2 \sum_{i=1}^{\infty} \beta_n^{2(i-1)} E(\varepsilon_{t-i}^2 / \sigma_t^2)$. Using $\varepsilon_{t-i}^2 / \sigma_t^2 = \eta_{t-i}^2 (\sigma_{t-i}^2 / \sigma_t^2)$, and substitution of $2\alpha_n^2 = \zeta^2/n$ and $\beta_n^2 = (1 - \zeta/\sqrt{2n} + \lambda/n)^2 = 1 - 2\zeta/\sqrt{2n} + o(n^{-1/2})$, it follows that the variance of v_t is $O(n^{-1/2})$, so that $n^{-1/2} \sum_{i=1}^{\lfloor \cdot n \rfloor} v_t / \sigma_t \xrightarrow{P} 0$. This proves (30).

The results (31) and (32) follow from (30), together with the result that $(n^{1/2} D_{2n} Z_{\lfloor \cdot n \rfloor}, \sigma_{\lfloor \cdot n \rfloor}^{-1}) \xrightarrow{\mathcal{L}} (G(\cdot), V(\cdot)^{-1/2})$, and the fact that $(\eta_t + v_t)$ has bounded variance, so that again the conditions of Hansen (1992a) for weak convergence to a stochastic integral apply.

For the results on the score and information for ξ , let $e_t = (\eta_t^2 - 1)/\sqrt{2}$, $W_{2n}(s) = n^{-1/2} \sum_{t=1}^{\lfloor sn \rfloor} e_t$ and define $F_{1n}(s) = (1 - \beta_n) \sum_{i=0}^{\lfloor sn \rfloor - 1} \beta_n^i = 1 - \beta_n^{\lfloor sn \rfloor} = 1 - \exp(\lfloor sn \rfloor \log(1 - \zeta/\sqrt{2n} + \lambda/n)) \rightarrow 1$. Then the first component S_ω of S_ξ , properly normalized, satisfies

$$\sqrt{\frac{2}{n}} (1 - \beta_n) S_\omega = \int_0^1 \sigma_{\lfloor sn \rfloor}^{-2} F_{1n}(s) dW_{2n}(s) \xrightarrow{\mathcal{L}} \int_0^1 V(s)^{-1} dW_2(s). \quad (\text{A.8})$$

For the second component $S_{\alpha+\beta}$, we use $F_{2n}(s) = (1 - \beta_n) \sum_{i=0}^{\lfloor sn \rfloor - 1} \beta_n^i \sigma_{\lfloor sn \rfloor - i - 1}^2 / \sigma_{\lfloor sn \rfloor}^2$, which converges weakly to 1, so that

$$\sqrt{\frac{2}{n}}(1 - \beta_n)S_{\alpha+\beta} = \int_0^1 F_{2n}(s)dW_{2n}(s) \xrightarrow{\mathcal{L}} \int_0^1 dW_2(s) = W_2(1). \quad (\text{A.9})$$

Note that $\sqrt{2/n}(1 - \beta_n) = n^{-1}\zeta + o(n^{-1})$, so that $n^{-1}\zeta$ (or n^{-1}) can also be used as a normalization in (A.8) and (A.9). The third part S_α of S_ω satisfies

$$\sqrt{\frac{1 - \beta_n^2}{n}}S_\alpha = \frac{1}{\sqrt{n}} \sum_{t=1}^n \left(\sum_{i=0}^{t-1} \sqrt{1 - \beta_n^2} \beta_n^i \frac{\sigma_{t-i-1}^2}{\sigma_t^2} e_{t-i-1} \right) e_t = \frac{1}{\sqrt{n}} \sum_{t=1}^n u_t e_t, \quad (\text{A.10})$$

Now u_t is a stationary process with $E(u_t^2) = q$, such that $\sqrt{(1 - \beta_n^2)/n}S_\alpha \xrightarrow{\mathcal{L}} N(0, q)$. As $1 - \beta_n^2 = 2\zeta/\sqrt{2n} + o(n^{-1/2})$, the normalization in (A.10) is equivalent to $2^{1/4}\zeta^{1/2}n^{-3/4}$. In summary, letting

$$D_{3n} = \begin{bmatrix} \frac{\zeta}{n} I_2 & 0 \\ 0 & \sqrt{\frac{1 - \beta_n^2}{n}} \end{bmatrix}, \quad (\text{A.11})$$

we have $D_{3n}S_\xi = O_p(1)$. By similar methods, it can be shown that $D_{3n}J_\xi D_{3n}$ converges, and that $D_{2n}J_{\delta\xi} D_{3n}$ converges; the latter converges to zero due to the two parts of the score vector being uncorrelated because $E[\eta_t(\eta_t^2 - 1)] = 0$. \square

Proof of Theorem 2. The result (34) follows from Lemma 4 and (14). Previous derivations show that

$$\begin{aligned} D_{2n} \sum_{t=1}^n Z_t \Delta X_t &= D_{2n} \sum_{t=1}^n Z_t \sigma_t \eta_t + D_{2n} \sum_{t=1}^n Z_t Z_t' \nu / n \\ &\xrightarrow{\mathcal{L}} \int_0^1 G(s) V(s)^{1/2} dW_1(s) + \kappa \int_0^1 G(s) Y(s) ds, \end{aligned} \quad (\text{A.12})$$

and similarly

$$D_n \sum_{t=1}^n Z_t Z_t' D_n \xrightarrow{\mathcal{L}} \int_0^1 G(s) G(s)' ds. \quad (\text{A.13})$$

Finally,

$$\begin{aligned} \frac{1}{n} \sum_{t=1}^n (\Delta X_t - \hat{\delta}'_{LS} Z_t)^2 &= \frac{1}{n} \sum_{t=1}^n \varepsilon_t^2 + o_P(1) \\ &= \frac{1}{n} \sum_{t=1}^n \sigma_t^2 + \frac{1}{n} \sum_{t=1}^n \sigma_t^2 (\eta_t^2 - 1) + o_P(1) \\ &\xrightarrow{\mathcal{L}} \int_0^1 V(s) ds. \end{aligned} \quad (\text{A.14})$$

This proves (35). \square