BANCA D'ITALIA
EUROSISTEMA

Temi di discussione

del Servizio Studi

Testing for trend

by Fabio Busetti and Andrew Harvey

Number 614 - February 2007

TESTING FOR TREND

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Abstract

The paper examines various tests for assessing whether a time series model requires a slope component. We first consider the simple t-test on the mean of first differences and show that it achieves high power against the alternative hypothesis of a stochastic nonstationary slope as well as against a purely deterministic slope. The test may be modified, parametrically or nonparametrically to deal with serial correlation. Using both local limiting power arguments and finite sample Monte Carlo results, we compare the t-test with the nonparametric tests of Vogelsang (1998) and with a modified stationarity test. Overall the t-test seems a good choice, particularly if it is implemented by fitting a parametric model to the data. When standardized by the square root of the sample size, the simple t-statistic, with no correction for serial correlation, has a limiting distribution if the slope is stochastic. We investigate whether it is a viable test for the null hypothesis of a stochastic slope and conclude that its value may be limited by an inability to reject a small deterministic slope. Empirical illustrations are provided using series of relative prices in the euro-area and data on global temperature.

JEL Classification: C22, C52.
Keywords: Cramér-von Mises distribution, stationarity test, stochastic trend, unit root, unobserved component.

Contents

[*] Bank of Italy, Economic Research Department

[+] Cambridge University, Faculty of Economics

1 Introduction[1]

The question of whether a time series exhibits a trend is an important one. In other words, does the series show a persistent upward or downward movement over time that can be extrapolated into the future? If no trend is present, first differences of the series have zero mean and a time series model will have an eventual forecast function for the level that is constant.

The main issues are best understood by considering a model that specifically contains a stochastic trend, namely

$$y_t = \mu_t + \upsilon_t, \quad t = 1, ..., T \tag{1}$$

$$\mu_t = \mu_{t-1} + \beta_{t-1} + \eta_t, \qquad \eta_t \sim NID(0, \sigma_\eta^2), \tag{2}$$

$$\beta_t = \beta_{t-1} + \zeta_t, \qquad \zeta_t \sim NID(0, \sigma_\zeta^2), \tag{3}$$

where μ_t is the level of the trend, β_t is the slope, υ_t is a zero mean Gaussian stationary component and the notation $NID\,(0, \sigma^2)$ denotes normally and independently distributed with mean zero and variance σ^2. If both variances σ_η^2 and σ_ζ^2 are zero, the trend is deterministic. When only σ_ζ^2 is zero, the slope is fixed and the trend reduces to a random walk with drift. Allowing σ_ζ^2 to be positive, but setting σ_η^2 to zero gives an *integrated random walk* (IRW) trend, which when estimated tends to be relatively smooth. For all variants of the model, the eventual forecast function is a linear trend with slope b_T, where b_T is the estimator of β_T.

Most of the literature has focussed on models with a deterministic slope, that is $\sigma_\zeta^2 = 0$, and has investigated tests of the null hypothesis that this slope, β, is zero. If the notion of a stochastic slope is entertained, the question is whether or not to include this component, β_t, in the model. More formally the joint hypothesis to be tested is $H_0 : \beta_0 = 0$ and $\sigma_\zeta^2 = 0$.

Testing may be carried out nonparametrically or by fitting a model. When an unobserved components, or structural time series model (STM), is used,

[1]Earlier version of this paper were presented at the Frontiers in Time Series Analysis meeting in Olbia, the NSF/NBER Time Series conference in Heidelberg and the conference on Unit root and Cointegration Testing in Faro; we are grateful to a number of participants for helpful comments. We would like to thank in particular Peter Phillips, Robert Taylor, Jesus Gonzalo, Paulo Rodrigues and two anonymous referees. The views expressed here are those of the authors, not the Bank of Italy. Andrew Harvey gratefully acknowledges hospitality and financial support from the Bank of Italy.

v_t and the level and slope disturbances, η_t and ζ_t, respectively, are usually assumed to be mutually independent. If the model is estimated with a fixed slope, the null hypothesis that the slope is zero may be tested using its '$t-$statistic'. We will call this test $t_\beta(STM)$. If we allow for the possibility of the slope being stochastic, a t-test may be carried out by estimating a model and then setting σ_ζ^2 to zero while fixing the other parameters at their estimated values. (Since one of the variance parameters may be concentrated out of the likelihood function and re-estimated, the other variances are fixed relative to it). We will call this test $t_\beta^*(STM)$. Either way the alternative hypothesis is the general one of *permanent* slope, that is a nonstationary component that is deterministic in a limiting case. In neither case is it necessary to assume that σ_η^2 is positive.

The aforementioned t-tests may also be carried out within an ARIMA framework. The tests would be essentially as in Kim *et al.* (2003), except that they do not consider the possibility that it might be used against a stochastic trend.

Testing for trend can also be approached nonparametrically. There are a number of ways in which this may be done, but the most straightforward possibility is to base a test on the average change - growth rate if in logarithms - in the series, that is

$$\widehat{\beta} = T^{-1} \sum_{t=1}^{T} \Delta y_t = (y_T - y_0)/T, \tag{4}$$

where - to simplify notation - it is assumed that y_0 is also observed. The statistic can be regarded as a *contrast* between the first and last observations. The (asymptotic) distribution of $\widehat{\beta}$ depends on whether the trend is deterministic, integrated of order one, denoted $I(1)$, or $I(2)$. If the trend is $I(1)$, $T^{1/2}\widehat{\beta}$ converges to a limiting normal distribution with variance σ_L^2, where σ_L^2 is the long-run variance of Δy_t. This suggests a nonparametric test $t-$statistic in which $T^{1/2}\widehat{\beta}$ is divided by a consistent estimator of σ_L. Unfortunately, when the trend is deterministic, there is a problem because $\sigma_L^2 = 0$; the nonparametric tests proposed by Vogelsang (1998) do not suffer from this problem, nor does the local-to-unit root approach of Canjels and Watson(1997) to constructing confidence interval for β. On the other hand, if the trend is $I(2)$, $T^{-1/2}\widehat{\beta}$ has a limiting normal distribution. As well as providing part of the proof that tests such as the one based on $T^{1/2}\widehat{\beta}$ are consistent against the alternative of a stochastic slope, this result also suggests the possibility of a test in which the slope is stochastic under the null.

When the trend is deterministic, $Var(\widehat{\beta})$ is of $O(1/T^2)$. However, least squares regression on a time trend yields an estimator of β which converges at a faster rate: it needs to be multiplied by $T^{3/2}$ to yield a limiting distribution. Hence a t-test on the slope from this regression - which may be carried out parametrically or nonparametrically depending on whether a model for υ_t is specified - will be more powerful than a test based on $\widehat{\beta}$.

Although the nonparametric $t-$test based on $T^{1/2}\widehat{\beta}$ has power against a changing slope, the IRW trend model suggests an alternative approach based on signal extraction whereby the average change in the series is measured by subtracting the estimated level of the trend at the beginning from the level at the end. We can then test the significance of the change by reference to the root mean square error (RMSE) of the difference; see Visser and Molenaar (1995) and Harvey (2001). When σ^2_ζ is zero, the deterministic trend model is obtained and the test is the parametric t-test referred to at the end of the previous paragraph. As well as providing a test as to whether a model should include a slope, contrasting the beginning and end of a particular series is also important for measuring how much it has actually gone up or down over the period in question.

The IRW model offers a different perspective on testing for trend by assuming that an extracted trend should be slowly changing and smooth. While this view of the world contrasts with the fixed slope I(1) paradigm, it should be noted that a simple IRW model provides the statistical rationale for the Hodrick-Prescott filter in macroeconomics and is closely related to the cubic spline in statistics. The contrast between different ways of modeling trend can be illustrated by an example concerned with global warming. Figure 1 shows the global annual surface land and marine air anomalies with respect to the 1950-79 average; see Parker *et al* (1995). A deterministic trend[2] with an AR(1) plus white noise stationary part, has a slope of 0.0054, while a simple random walk plus drift with an additive irregular component gives an estimated slope of 0.0065. An IRW trend fits almost as well as a random walk plus drift trend and does even better if the irregular is replaced by an AR(1). The slope at the end of the series is 0.0215 and this contrasts sharply with the estimates for the fixed slopes in the random walk plus drift and the deterministic trend models. The implications for forecasting are clear from plots of the deterministic and smooth trends. Diagnostic tests and good-

[2] All the models were fitted using the STAMP package of Koopman et al (2000).

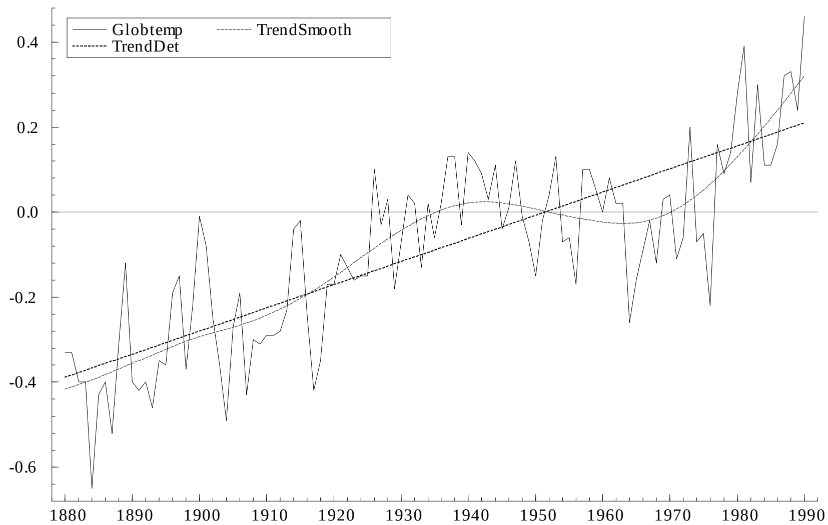

Figure 1: Global temperature anomaly series with smooth and deterministic trends plotted against year.

ness of fit statistics offer no clear guidance as to the choice of trend: the deterministic trend model gives the best fit, but only just.

The plan of the paper is as follows. Section 2 considers the null hypothesis of no slope against the alternative of a permanent slope. We note that this hypothesis may be tested by a variant of the stationarity test where it is assumed that there is no slope under the null hypothesis. The test may be carried out parametrically using the residuals from a fitted model or nonparametrically using the approach of Kwiatkowski et al (1992). As such it offers an alternative to the t-test on the slope. Section 3 provides the limiting representation of these tests and shows their consistency. Their performance is compared with a number of other tests, including the t-PS^1 test of Vogelsang (1998). Section 4 investigates how the $t-$test fares when the trend is deterministic. Section 5 picks up on the result that when the trend is $I(2)$, $T^{-1/2}\widehat{\beta}$ has a limiting distribution. It is shown that the standardized slope, $\widehat{\beta}/\widehat{\sigma}$, where $\widehat{\sigma}^2$ is the sample variance of the first differences, has a known (nonstandard) limiting distribution and so can be used to construct a nonparametric test of the null hypothesis of a stochastic slope. Section 6 provides empirical illustrations with the series of global temperature and data of relative prices in the euro area. Section 7 concludes.

2 Testing against stochastic and deterministic slope

Taking first differences in (1), and assuming for simplicity of notation that y_0 is also observed, gives

$$\Delta y_t = \beta_{t-1} + \omega_t, \qquad t = 1, 2, ..., T \tag{5}$$

where $\omega_t = \eta_t + \Delta \upsilon_t$. More generally, assuming that ω_t is a zero mean weakly dependent process with long run variance $\sigma_L^2 > 0$ allows us to obtain asymptotic representations for nonparametrically modified tests of the null hypothesis that there is no slope, that is $\beta_t = 0$ for all t.

Consider the model given by (5) and (3). If the slope is assumed to be deterministic, that is $\sigma_\zeta^2 = 0$, a nonparametric test of the hypothesis that it is zero can be set up in first differences using the test statistic

$$t_\beta(m) = T^{\frac{1}{2}} \widehat{\beta} / \widehat{\sigma}_L(m) \tag{6}$$

where the denominator is the square root of an estimator of the long-run variance based on m lags. Various options for the kernel and guidelines for choosing m may be found in Andrews (1991). Here we use a Bartlett window.

The test can be carried out by fitting a parametric model and testing the significance of the estimate of the fixed slope β. Either an ARMA model can be fitted to first differences or a structural time series model may be estimated. The latter may have some attraction when the series is such that it is natural to include components like cycles and seasonals in a model for the levels. As noted in the introduction there is a variant in which the model is fitted with allowance made for a stochastic slope. These parametric tests are still valid if $\sigma_\eta^2 = 0$ in (3) and they have the same asymptotic representation as $t_\beta(m)$, given in proposition 1 below.

If $\omega_t \sim NID(0, \sigma_\omega^2)$ and $\zeta_t \sim NID(0, \sigma_\zeta^2)$, the locally best invariant (LBI) test of the null hypothesis of a deterministic slope against the alternative of a nonstationary stochastic slope, that is $H_0 : \sigma_\zeta^2 = 0$ against $H_1 : \sigma_\zeta^2 > 0$ is to reject for large values of

$$\zeta^1 = T^{-2}\widehat{\sigma}^{-2}\sum_{t=1}^{T}\left(\sum_{i=1}^{t}(\Delta y_i - \widehat{\beta})\right)^2 = T^{-2}\widehat{\sigma}^{-2}\sum_{t=1}^{T}\left(y_t - y_0 - t\widehat{\beta}\right)^2, \qquad (7)$$

where $\widehat{\sigma}^2$ is the sample variance of the first differences. This is the test of Nyblom and Mäkeläinen (1983) applied to first differences; Bailey and Taylor (2002) show that it is still optimal when ω_t and ζ_t are correlated. In deriving the test from the LBI principle, one initially obtains the summations running in reverse, that is from $t = T$ to i, but, as a consequence of fitting the slope it can be shown that the two statistics are identical. Asymptotically, ζ^1 has the Cramér-von Mises distribution, denoted CvM_1, under the null hypothesis.

If the above test statistic is formed (with reverse partial sums) without subtracting $\widehat{\beta}$ it will be LBI against $H_1 : \sigma_\zeta^2 > 0$ for zero initial conditions, that is $\beta_0 = 0$. However, it is also a consistent test against a deterministic slope. Its asymptotic distribution under the null is a different member of the Cramér-von Mises family, denoted CvM_0; see Nyblom (1989). The statistics constructed with forward and reverse partial sums, ζ_F^0 and ζ_R^0 respectively, are no longer identical, but have the same asymptotic distribution under the null hypothesis. Although ζ_R^0 is the LBI test, assuming that β_0 is zero is somewhat arbitrary, which is why ζ_F^0 is considered as well.

To be strictly LBI, the statistics ζ_F^0 and ζ_R^0 should have the sample mean square rather than the sample variance in the denominator. However, the limiting distribution under the null hypothesis is the same. The local asymptotic distribution is also the same, but because dividing by the sample variance makes the test statistic bigger, the power is higher in small samples and there appears to be no adverse effect on size.

Parametric and nonparametric forms of the above LBI tests may be constructed. In the nonparametric case, $\widehat{\sigma}_L^2(m)$ replaces $\widehat{\sigma}^2$ just as in the $t-$test. A parametric statistic can be constructed by fitting an ARIMA or structural time series model.

3 Limiting representations and local asymptotic power

For a model given by (5) and (3) in which ω_t is a weakly dependent process with positive long run variance, σ_L^2, and ζ_t need not be Gaussian, we consider the limiting behavior of the non-parametrically modified statistics $t_\beta(m)$, $\zeta_F^0(m)$, $\zeta_R^0(m)$ and $\zeta^1(m)$, where $\zeta^1(m)$ is as in (7) with $\widehat{\sigma}$ replaced by $\widehat{\sigma}_L(m)$, and similarly for $\zeta_F^0(m)$ and $\zeta_R^0(m)$, and $m \to \infty$ with $m^2/T \to 0$. The following proposition provides the limiting distribution of the statistics under the local alternative hypothesis

$$H_{1,T} : \beta_0 = c_d \sigma_L / \sqrt{T},\ \sigma_\zeta^2 = c_s^2 \sigma_L^2 / T^2, \tag{8}$$

where c_d, c_s are fixed constants; the proof is contained in the appendix.

Proposition 1 *Consider the model (5)-(3) with ω_t being a weakly dependent process as in Stock (1994, p.2745), with $\sigma_L^2 > 0$. Let $W_0(r)$ and $W_1(r)$ be independent standard Wiener processes for $r \in [0,1]$. Then under $H_{1,T}$,*

$$t_\beta(m) \xrightarrow{d} V(1; c_d, c_s), \tag{9}$$

$$\zeta_F^0(m) \xrightarrow{d} \int_0^1 V(r; c_d, c_s)^2 dr, \tag{10}$$

$$\zeta_R^0(m) \xrightarrow{d} \int_0^1 \left(V(1; c_d, c_s) - V(r; c_d, c_s)\right)^2 dr, \tag{11}$$

$$\zeta^1(m) \xrightarrow{d} \int_0^1 V^*(r; c_s)^2 dr, \tag{12}$$

where

$$V(r; c_d, c_s) = W_0(r) + c_d r + c_s \int_0^r W_1(s) ds,$$

$$V^*(r; c_s) = W_0(r) - rW_0(1) + c_s \int_0^r \left(W_1(s) - \int_0^1 W_1(u) du \right) ds.$$

Remark 1 *For $c_s = 0$, the limiting distribution of $t_\beta^2(m)$ is a noncentral chi-square with one degree of freedom and noncentrality parameter equal to c_d^2, while that of $\zeta^1(m)$ is a standard Cramér-von Mises, CvM_1, distribution. For $c_s = c_d = 0$, $t_\beta(m)$ is asymptotically standard normal, while the limiting distributions of $\zeta_F^0(m)$ and $\zeta_R^0(m)$ are CvM_0.*

Consistency of the tests follows after showing that the statistics diverge under the fixed alternative hypotheses of deterministic and stochastic slope. This is done in the following proposition; the proof is in the appendix. Note that the test based on $\zeta^1(m)$ is not consistent against a purely deterministic slope.

Proposition 2 *Under the fixed alternative hypothesis of deterministic slope, $\beta_0 \neq 0$ and $\sigma_\zeta^2 = 0$, $t_\beta(m)$ is of $O_p\left(T^{\frac{1}{2}}\right)$, $\zeta_F^0(m)$ and $\zeta_R^0(m)$ are of $O_p(T)$, while $\zeta^1(m) \xrightarrow{d} \int_0^1 V^*(r;0)^2 dr$ as in proposition 1. Under the fixed alternative hypothesis of stochastic slope, $\beta_0 = 0$ and $\sigma_\zeta^2 > 0$, $t_\beta(m)$ is of $O_p\left((T/m)^{\frac{1}{2}}\right)$ while $\zeta_F^0(m)$, $\zeta_R^0(m)$, $\zeta^1(m)$ are of $O_p(T/m)$.*

The asymptotic representations given in proposition 1 can be used to compare the power of the tests against local deviations from the null hypothesis in the direction of deterministic and/or stochastic slope. Some results are reported in table 1 in terms of the percentage of rejections. Specifically, we have generated 50,000 replications of the limiting random variables defined in (9) to (12) by replacing the continuous time Wiener processes W_0 and W_1 by their discrete counterparts (dividing the unit interval into 1000 parts) and computing the rejection probabilities for tests run at the 5% level of significance.

In addition we report the rejection frequencies for the tests $t\text{-}PS^1$ and $T^{-1}W$ of Vogelsang (1998), and for the $t\text{-}PS$ test of Zambrano and Vogelsang (2000), here denoted as $t\text{-}PS^1_{FD}$, which is similar to $t\text{-}PS^1$ but computed on the constant term of first differences. The power of these tests is computed by direct simulation of the data generating process (5) to (3) with $\omega_t \sim NID(0,1)$, for $T = 1000$ under the (local) alternative hypothesis $\beta_0 = c_d/\sqrt{1000}$, $\sigma_\eta = c_s/1000$; given that T is large and the alternative hypothesis is scaled by the sample size these rejection frequencies can be seen as approximation of the local limiting power. Recall that the $t\text{-}PS^1$ test is asymptotically valid whether or not y_t contains a unit root, while $T^{-1}W$ requires a unit root in y_t. Note that while the non-parametric versions of our tests $t_\beta(m), \zeta_F^0(m), \zeta_R^0(m), \zeta^1(m)$ require a unit root in y_t, this is not the case for the parametric tests $t_\beta(STM)$ and $t_\beta^*(STM)$ based on estimating an unobserved component model, or their analogues based on ARIMA models. The limiting power of $t_\beta(STM)$, $t_\beta^*(STM)$, and their ARIMA analogues, is the same as that of $t_\beta(m)$.

As expected the $t_\beta(m)$ test is most powerful against a deterministic slope. For example for $c_d = 2$ (and $c_s = 0$), its local asymptotic power is 0.518, as opposed to 0.441 and 0.443 for the tests based on $\zeta_F^0(m)$ and $\zeta_R^0(m)$ respectively. Note that the asymptotic power of the $\zeta^1(m)$ test against a deterministic slope is always equal to its size. The $\zeta_R^0(m)$ test achieves the highest power against a stochastic slope starting at zero, that is $\beta_0 = c_d = 0$;

indeed it corresponds to the LBI test for this case. Thus with $c_s = 5$, the power of the $\zeta_R^0(m)$ test is 0.569 while that of $t_\beta(m)$ is only 0.524. However, the $t_\beta(m)$ test dominates both the $\zeta_F^0(m)$ and $\zeta^1(m)$ tests, for which the powers are 0.436 and 0.310 respectively. On the other hand, the power of $\zeta_F^0(m)$ is slightly greater than that of $\zeta_R^0(m)$ when c_d is high and c_s is not too large. Of course, if β_{T+1} rather than β_0 had been assumed to be zero, the powers of $\zeta_F^0(m)$ and $\zeta_R^0(m)$ would have been interchanged. The $\zeta^1(m)$ test is invariant to c_d and it is dominated by all the other tests except when $c_s = 50$. However, this may be useful insofar as a non-rejection by $\zeta^1(m)$ and rejection by the other tests is an indication of deterministic slope. Overall, it seems that the $t_\beta(m)$ test is the best compromise. Even when $c_d = 0$ and $c_s = 50$ its power is only a little below those of the $\zeta_F^0(m)$ and $\zeta_R^0(m)$ tests. Furthermore there may be occasions when a one-sided alternative is plausible, in which case the $t_\beta(m)$ would become even more powerful.

The t-PS^1 test of Vogelsang (1998) has low power against the alternative of deterministic slope and has virtually no power against a stochastic slope. The t-PS^1_{FD} test is more powerful than t-PS^1 but less powerful than $T^{-1}W$. However, $T^{-1}W$ is clearly dominated by the $t_\beta(m)$ test.

4 Finite sample behaviour for a deterministic trend with highly persistent disturbances

In this section we compare the $t_\beta(m)$ test with the t-PS^1, $T^{-1}W$ tests of Vogelsang (1998) and with the t-PS^1_{FD} of Zambrano and Vogelsang (2000) when the data generating process consists of a trend with a deterministic slope plus a highly persistent $AR(1)$ process, that is

$$y_t = \alpha + \beta_T t + v_t, \qquad t = 1, 2, ..., T \tag{13}$$

$$v_t = \rho\varepsilon_{t-1} + \eta_t, \quad \eta_t \sim NID(0,1), \tag{14}$$

$$\beta_T = c_d/\sqrt{T}, \tag{15}$$

where $c_d = 0, 0.5, 1, 1, 1.5, 2, 3$, $\rho = 1, .975, .95, .90$, and $T = 100$. When $\rho = 1$ the series Δy_t is $I(0)$, while it is overdifferenced when $\rho < 1$. Overdifferencing implies that the long run variance of Δy_t is equal to zero but $t_\beta(m)$ will still use the estimator $\hat{\sigma}_L^2(m)$: the question is whether the test might continue to work reasonably well in finite samples when ρ is close to one.

Table 1: Simulated local asymptotic power of the tests (x100)

	c_d	0	0.5	1.0	2.0	3.0	4.0
	$t_\beta(m)$	4.9	7.7	16.7	51.8	85.2	97.9
$c_s = 0$	$\zeta_F^0(m)$	5.0	7.3	14.5	44.1	78.0	95.4
	$\zeta_R^0(m)$	4.9	7.1	14.6	44.3	77.8	95.3
	$\zeta^1(m)$	4.9	4.9	4.9	4.9	4.9	4.9
	$t\text{-}PS^1$	4.8	5.5	6.9	12.2	18.2	22.8
	$T^{-1}W$	5.3	7.3	13.7	39.2	68.9	89.4
	$t\text{-}PS^1_{FD}$	5.7	8.0	14.8	40.5	69.5	88.4
	$t_\beta(m)$	26.4	28.5	34.1	52.4	72.6	87.7
$c_s = 2.50$	$\zeta_F^0(m)$	19.2	21.3	27.4	47.5	70.2	87.2
	$\zeta_R^0(m)$	30.1	31.5	35.5	49.9	66.7	81.8
	$\zeta^1(m)$	12.7	12.7	12.7	12.7	12.7	12.7
	$t\text{-}PS^1$	6.5	6.7	7.4	10.0	13.3	17.0
	$T^{-1}W$	18.9	20.1	24.4	39.5	58.6	75.9
	$t\text{-}PS^1_{FD}$	14.9	16.7	21.3	37.4	57.8	76.5
	$t_\beta(m)$	52.4	53.1	54.8	60.5	68.6	77.3
$c_s = 5$	$\zeta_F^0(m)$	43.6	44.7	46.9	55.5	66.5	77.9
	$\zeta_R^0(m)$	56.9	57.4	58.4	62.4	68.0	74.8
	$\zeta^1(m)$	31.0	31.0	31.0	31.0	31.0	31.0
	$t\text{-}PS^1$	5.6	5.8	6.1	6.9	8.6	10.4
	$T^{-1}W$	35.0	35.5	37.2	43.6	52.7	62.9
	$t\text{-}PS^1_{FD}$	24.8	25.6	27.9	36.7	48.5	61.6
	$t_\beta(m)$	74.3	74.2	74.5	75.6	76.8	78.9
$c_s = 10$	$\zeta_F^0(m)$	70.5	70.7	71.1	72.9	75.6	79.3
	$\zeta_R^0(m)$	78.7	78.9	79.0	79.6	80.6	82.0
	$\zeta^1(m)$	61.3	61.3	61.3	61.3	61.3	61.3
	$t\text{-}PS^1$	2.7	2.6	2.8	3.0	3.5	3.9
	$T^{-1}W$	48.0	48.3	48.6	50.4	52.8	56.4
	$t\text{-}PS^1_{FD}$	33.2	33.3	34.2	37.3	42.2	48.0
	$t_\beta(m)$	94.8	94.8	94.9	94.8	94.8	94.8
$c_s = 50$	$\zeta_F^0(m)$	98.8	98.8	98.7	98.8	98.8	98.8
	$\zeta_R^0(m)$	99.1	99.1	99.1	99.2	99.1	99.1
	$\zeta^1(m)$	99.1	99.1	99.1	99.1	99.1	99.1
	$t\text{-}PS^1$	0.1	0.1	0.1	0.1	0.1	0.1
	$T^{-1}W$	56.1	56.2	56.1	56.1	56.1	56.1
	$t\text{-}PS^1_{FD}$	36.7	36.7	36.7	36.8	36.9	37.2

The above model is as in Table IV of Vogelsang (1998, p.141), except that - to save space- we report results only for a pure AR(1) process instead of a more general ARMA(1,1). It is also the data generating process for the simulation results of Kim *et al.* (2003, Table II, p.543). Here we essentially replicate the experiments of Vogelsang (1998) and Kim *et al.* (2003), adding the rejection frequencies for the $t_\beta(m)$ and the $t\text{-}PS^1_{FD}$ tests to the results already reported in those papers. Additionally, as a benchmark, we report the rejection frequencies of the unfeasible GLS t-test (denoted t^*_{GLS}), obtained by computing the t-statistic on the trend coefficient after applying the Cochrane-Orcutt transformation to equation (13) taking ρ as known. Table 2 contains the simulated percentage rejections over 10,000 Monte Carlo replications. The random number generator of the matrix programming language Ox 2.20 was used.

For the case of $\rho = 1$ we obtain rejection frequencies similar to the case of $c_s = 0$ in table 2, confirming that the local limiting power analysis of the previous section provides a good approximation in finite samples. Notice that in this case $m = 0$ would be the correct choice for the long run variance estimator (and that $t_\beta(0)$ coincides with t^*_{GLS}); $m = 4, 8$ renders the t-test somewhat oversized.

When $\rho < 1$ all tests but the unfeasible GLS are conservative, in the sense that the actual size is well below the nominal 5%. As c_d increases the power becomes non-negligible. It is interesting to notice that the rejection frequencies of $t_\beta(m)$ increase with m, since adding extra lags tends to reduce the estimate of the long run variance, a reflection of the overdifferenced nature of Δy_t. Despite being asymptotically valid, the $t\text{-}PS^1$ test still appears less preferable than $t_\beta(m)$ when there is high persistence in the data ($\rho = 0.975$ and 0.95); in these cases its power is very low for $c_d \geq 2$ (the $T^{-1}W$ and the $t\text{-}PS^1_{FD}$ tests display more power but not as high as $t_\beta(m)$)[3]. The $t\text{-}PS^1$ test is more attractive when $\rho = 0.90$, since it has non-negligible power for very small values of c_d, although it also displays a big probability of not rejecting the null hypothesis when $c_d > 2$. Figure 2 provides the graphs for the empirical power functions of the tests $t_\beta(4)$, $T^{-1}W$, $t\text{-}PS^1$ and $t\text{-}PS^1_{FD}$ for $\rho = 1, .975, .95, .9$.

The figures for $T^{-1}W$, $t\text{-}PS^1$reported in table 3 are also very similar to those of Table IV of Vogelsang (1998). Vogelsang additionally reports the

[3] For $\rho = 0.975$ and $c_d \geq 2$ the $t_\beta(m)$ test rejects more frequently than the infeasible GLS test, but this is a finite sample effect.

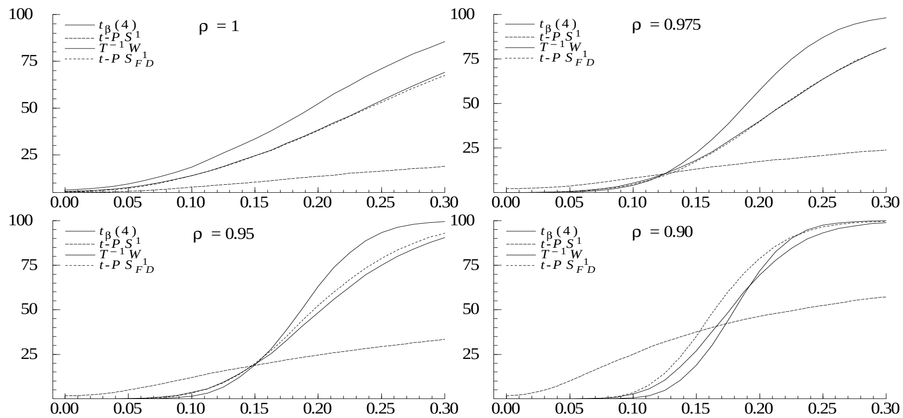

Figure 2: Empirical power functions of tests plotted against β_T. The data generating process is a deterministic linear trend model with $AR(1)$ errors (from $NID(0, 1)$ innovations). The sample size is $T = 100$; β_T is the magnitude of the deterministic slope.

rejection frequencies for the feasible GLS (Bonferroni) test of Canjels and Watson (1997) based on a local-to-unit root approximation. His results show that in nearly all cases the Bonferroni GLS test is dominated by either $T^{-1}W$ or t-PS^1. Vogelsang then graphs (cf. figures 2-5, p.138-139) the asymptotic power functions of t-PS^1 and $T^{-1}W$ based on the sequence of local alternatives (15) for a local-to-unit root data generating mechanism. The power displayed in those graphs appears much greater than its finite sample approximation reported in his Table IV and in our table 2 and our figure 2. We believe that his graphs refer to one-sided tests, as opposed to two-sided tests, and this would explain the apparent contradiction.[4]

Table II of Kim et al. (2003) contains useful additional results for the data generating process (13)-(15). In particular the rejection frequencies from fitting an ARIMA(1,1,1) model, which in our notation corresponds to the parametric $t_\beta(ARIMA)$ test, are reported, showing that the parametric test clearly dominates. It is never oversized for $\rho < 1$ and it generally displays

[4] Bunzel and Vogelsang (2005) recently proposed a modification of the tests of Vogelsang (1998) based on the fixed-bandwidth asymptotic framework. The power improvement of the new test appears to be of the order of $0.05 - 0.1$ for $\rho = 0.90$ (cf. their figure 7, p.390); we thank a referee for pointing out this reference.

greater power. For example, for $\rho = 0.90$ and $c_d = 1$, the power of the parametric test equals 31% against 25.2% of t-PS^1. However further from the alternative, e.g. for $c_d = 2$, the power properties of $t_\beta(ARIMA)$ are largely comparable with those of $t_\beta(m)$ reported in our table 2. The $t_\beta(ARIMA)$ test is oversized for $\rho = 1$, the empirical size being 11% for a sample of 100 observation, due to the fact that the ARIMA(1,1,1) structure is now overparametrized. Kim et al. (2003) also confirm previous results of Sun and Pantula (1999) on the degree of oversizing of the test obtained from estimating the correctly specified linear trend plus AR(1) model in finite samples: the empirical size of a nominal 5% test would be equal to 27% and 19% for ρ equal to 0.95 and 0.90 respectively.

In summary, the $t_\beta(m)$ test is conservative when the trend is deterministic and this has implications for its power. Nevertheless it still has good power properties when the data are highly persistent and, in such circumstances, it may be preferable to tests specifically designed to be robust to whether the series are I(0) or I(1).

5 Permanent slope as the null

When the slope is stochastic, that is σ_ζ^2 in (3) is positive, the standardized slope, $\widehat{\beta}^* = \widehat{\beta}/\widehat{\sigma}_\omega$, where $\widehat{\sigma}_\omega^2 = T^{-1}\sum_{t=1}^T(\Delta y_t - \widehat{\beta})^2$, has an asymptotic distribution given by

$$\widehat{\beta}^* = \widehat{\beta}/\widehat{\sigma}_\omega = t_\beta(0)/\sqrt{T} \xrightarrow{d} \left[\int_0^1 \underline{W}(r)^2 dr\right]^{-1/2} \int_0^1 W(r)dr\,, \qquad (16)$$

where $\underline{W}(r) \equiv W(r) - \int_0^1 W(r)dr$. No correction for serial correlation is needed. Note that the limiting distribution of $T^{-1/2}\widehat{\beta}$ is $N(0, 1/3)$.

The above result suggests the setting up of a test of the null hypothesis of stochastic slope, that is $\sigma_\zeta^2 > 0$, by rejecting for small values of $\left|\widehat{\beta}^*\right|$. The 1%, 5%, 10% critical values, based on the asymptotic distribution, are 0.024, 0.118 and 0.239 respectively. The proof of (16) follows as a special case of theorem 2 of Vogelsang (1998), though the motivation for his test, and the way it is used, is completely different. Our test derives from a proposal made by Bierens (2001) in the context of testing nonstationary cycles. Here the test can be regarded as a test (in differences) at frequency

Table 2: Simulated finite sample power of tests against β_T (x100)

ρ	c_d	GLS^*	$t_\beta(0)$	$t_\beta(4)$	$t_\beta(8)$	$t\text{-}PS^1$	$T^{-1}W$	$t\text{-}PS^1_{FD}$
1	0.0	5.3	5.3	6.4	8.0	5.3	5.6	5.4
	0.5	8.3	8.3	9.5	11.1	5.8	7.7	7.4
	1.0	16.8	16.8	18.6	21.0	8.0	13.9	14.0
	1.5	31.6	31.6	33.6	35.4	10.6	24.6	24.6
	2.0	50.5	50.5	52.3	54.5	13.7	38.3	38.0
	3.0	84.9	84.9	85.4	86.2	19.0	69.1	67.5
0.975	0.0	5.0	0.0	0.0	0.0	2.2	0.1	0.0
	0.5	6.5	0.1	0.3	0.7	3.8	0.7	0.6
	1.0	11.7	2.1	4.0	6.5	8.3	5.3	4.6
	1.5	20.0	16.1	22.1	27.3	13.2	18.1	17.5
	2.0	30.7	51.7	57.6	62.7	17.5	40.2	40.0
	3.0	57.5	98.1	98.1	98.0	24.0	81.1	81.1
0.95	0.0	5.0	0.0	0.0	0.0	1.7	0.0	0.0
	0.5	11.7	0.0	0.0	0.1	4.9	0.1	0.1
	1.0	30.7	0.3	1.4	3.8	11.9	3.5	3.2
	1.5	57.5	9.1	19.1	27.9	18.9	19.6	19.9
	2.0	81.4	51.1	63.1	70.9	24.6	48.2	52.4
	3.0	99.1	99.7	99.5	99.5	33.5	90.5	93.1
0.90	0.0	5.0	0.0	0.0	0.0	1.7	0.0	0.0
	0.5	30.7	0.0	0.0	0.0	10.1	0.0	0.0
	1.0	81.4	0.0	0.5	3.1	25.2	2.7	3.4
	1.5	99.1	3.7	18.8	36.3	37.6	27.0	35.0
	2.0	100.0	48.4	71.9	84.1	46.4	69.7	79.0
	3.0	100.0	100.0	100.0	99.9	57.3	99.0	99.7

zero. However, we have made a slight modification in that Bierens constructs the denominator without subtracting the mean. In the present context this leads to the statistic $\widehat{\beta}^{\dagger} = \widehat{\beta}/(\Sigma(\Delta y_t)^2/T)^{1/2}$. The asymptotic distribution then has $W(r)$ in the denominator as well as in the numerator. This makes virtually no difference to the 5% critical value which is the same as before to three decimal places. Since the statistic is smaller than the one based on $\widehat{\beta}^{*}$ it might be thought more likely to reject. However, as will be seen shortly, this appears to make virtually no difference in practice.

What if the slope is purely deterministic? Then

$$p\lim\widehat{\beta}^{*} = \beta/\sigma_{\omega}, \qquad \text{and} \qquad p\lim\widehat{\beta}^{\dagger} = \beta/\sqrt{\sigma_{\omega}^{2} + \beta^{2}}. \tag{17}$$

In both cases the null is unlikely to be rejected unless the size of the deterministic slope is small relative to σ_{ω}^{2}. Specifically, at the 5% level of significance, the null is rejected by $\widehat{\beta}^{*}$ with probability one as $T \to \infty$ only if $|\beta| < 0.118\sigma_{\omega}$.

To evaluate the properties of the $\widehat{\beta}^{*}$ and $\widehat{\beta}^{\dagger}$ tests, a series of Monte Carlo experiments were carried out for the model (5)-(3) with $\omega_t \sim NID(0, \sigma_{\omega}^{2})$ and $\beta_0 = 0$. Table 3 shows probabilities of rejection at the 5% level of significance, estimated with 10,000 replications, over different values of $q^{1/2} = \sigma_{\zeta}/\sigma_{\omega}$ for $T = 100$. Results for the augmented Dickey-Fuller (ADF) test with m lags, denoted $ADF_0(m)$, are given so as to provide a benchmark; $ADF_1(m)$ indicates the inclusion of a constant. In practice, small values of q are most likely to arise, so the case of $q = 0.01$ ($q^{1/2} = 0.1$) is of particular importance. When q is small, the size of the ADF test is well above the nominal 5%. The reason for this is well-known - the reduced form of second differences contains a moving-average root close to the unit circle and hence the autoregressive approximation is poor. On the other hand the $\widehat{\beta}^{*}$ and $\widehat{\beta}^{\dagger}$ tests do rather well in that the rejection probability is 0.17 for $q = 0.01$ while when $q = 0$ it shoots up to 0.76.

We also carried out simulations for $\beta_0 = 0.1, 0.2$.and 0.5. For non-zero q the rejection probabilities of the standardized slope tests and $ADF_0(m)$ are changed very little; $ADF_1(m)$ is unaffected anyway. (Indeed it can be shown that the local asymptotic distributions -as in Phillips and Perron(1988) -are independent of β_0). When $q = 0$, there is a sharp change from $\beta_0 = 0.1$, where the probability of rejection is quite high, to $\beta_0 = 0.2$, where it is low; see Busetti and Harvey (2002) for details. This is exactly what one would expect given the probability limit in (17).

Table 3: Percentage rejections for a random walk with stochastic drift data generating process, T=100

$q^{\frac{1}{2}}$	0	0.1	0.25	0.5	1
$\widehat{\beta}^{\dagger}$	76.1	16.8	8.0	5.8	5.0
$\widehat{\beta}^{*}$	75.9	16.7	8.0	5.8	4.9
$ADF_0\,(5)$	100.0	71.6	24.7	8.8	5.3
$ADF_1\,(5)$	99.7	77.8	26.7	9.4	6.1

6 Empirical illustrations

6.1 Global temperature

The nonparametric test statistics for the null of no slope in the temperature data of figure 1 are as follows: $t_{\beta}(10) = 1.36$, $\zeta_F^0(10) = .308$ and $\zeta_R^0(10) = 1.055$. None of the tests rejects. Vogelsang's (nonparametric) t-PS^1 test, on the other hand, rejects at the 5% level of significance - but not at the 1% level; see also Fomby and Vogelsang (2002). Reversing the null hypothesis as in section 5 tells a consistent story in that $\widehat{\beta}^{*} = .052$ and $\widehat{\beta}^{\dagger} = 0.046$, so the null of a permanent slope is rejected. As regards parametric tests, fitting a random walk plus drift with an additive irregular component to the levels of the observations gives a $t-$statistic of 1.870. This is close to rejection at the 5% level of significance. A one-sided test would reject and this might be reasonable as it corresponds to a hypothesis of an upward trend in temperature. An IRW trend fits almost as well as a random walk plus drift. The $t_{\beta}^{*}(STM)$ test is then based simply on fitting a deterministic trend and $t-$statistic of 13.69 indicating a massive rejection of the null of no slope.

6.2 Inflation differentials between Italy and the euro area

One of the Maastricht requirements for joining the European Monetary Union was that the country's inflation differentials with respect to the three best performers had to be less than 1.5 percentage points in the average of 1997. This ensured convergence in the rates of inflation for the economies that joined the monetary union. However, with the start of the EMU inflation differentials have somehow begun to widen again; cf.Busetti *et al* .(2006) for detailed evidence. Here as an example we consider the dynamics of Italian inflation vis-a-vis the other EMU countries. Denote by $y_{t,i}$ the log-price differential

$$y_{t,i} = \log P_t - \log P^*_{t,i}$$

where P_t is the Italian consumer price index (CPI) and $P^*_{t,i}$ is the CPI in country $i =$ Germany, France, Spain, The Netherlands, Belgium, Austria, Greece, Finland, Portugal, Ireland, Luxemburg. Figure 3 shows the dynamics of the relative prices $\exp(y_{t,i}) = P_t/P^*_{t,i}$ over the period 1998M1-2003M6, using seasonally adjusted data rebased to one in 1998M1. The graph shows that over this period inflation in Italy has been cumulatively higher than in France and Germany by around 5 percentage points; at the same time it has been substantially lower than in Ireland, Portugal and Greece. The presence of a drift in $y_{t,i}$ would indicate some kind of divergence in inflation rates following the inception of the EMU. Table 4 reports the results of the tests. A star indicates (at least) 10% rejection in the two-sided test of the null hypothesis of no drift for $t_\beta(4)$, $\zeta^0_R(4)$, $\zeta^0_F(4)$, $t\text{-}PS^1$ and a 10% rejection of the null of permanent drift for $\widehat{\beta}^*$. For the non-parametric long run variance estimator we have used a bandwidth parameter $m = 4$, the nearest integer in the formula $4\,(T/100)^{0.25}$; similar results however have been obtained for other values of m between 2 and 6 but to save space are not reported.

The t-test $t_\beta(4)$ and the stationarity test $\zeta^0_R(4)$ are those displaying more frequent rejections of the null hypothesis of zero inflation differentials. All tests agree on the stability of inflation differentials between Italy and Luxemburg and they provide strong evidence against convergence with Germany, France on one side and Greece, Portugal on the other. Contrary to the Vogelsang's $t\text{-}PS$ test, the non-parametric t-test would also imply non convergence with Belgium, Austria, Finland and Ireland. Actually, given the graph of the data it seems surprising that the $t\text{-}PS^1$ statistic fails to recognize a trend in the Italy-Ireland and Italy-Austria relative prices. However this could perhaps be explained in terms of the results of section 2.2, namely that the $t\text{-}PS^1$ test has no power against the presence of a stochastic slope. Finally, the evidence from the standardized drift test based on $\widehat{\beta}^*$, where rejection is interpreted as stability of inflation differentials, is somewhat inconclusive.

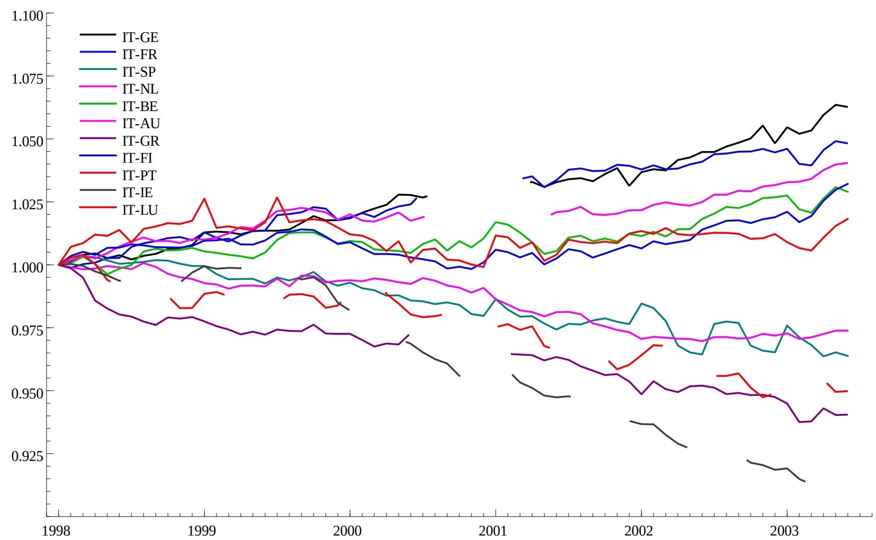

Figure 3: Relative prices between Italy and other countries in the euro area (1998=1)

Note that a possible objection to applying the nonparametric t_β and ζ^0 tests to the rate of inflation is that they are invalid if the price levels have converged. This is because the long-run variance is zero when there is overdifferencing. However, if the price levels contrasts are stationary but persistent, the evidence in sub-section 2.3 suggests that the $t-$test is conservative and so will tend not to reject the null hypothesis of stability. The same can be expected for the ζ^0 tests.

7 Conclusion

The t_β test is designed to test against a deterministic slope, but it is also consistent against the alternative hypothesis of a non-stationary stochastic slope. Overall, it seems to be the best option for testing the null hypothesis of no slope against the alternative of permanent slope. The parametric $t-$tests, which can be carried out in an unobserved components or ARIMA framework,

Table 4: Tests of trend in (log) relative prices

	$t_\beta(4)$	$\zeta_R^0(4)$	$\zeta_F^0(4)$	t-PS (10%)	$\widehat{\beta}^*$
IT-GE	4.974*	8.517*	7.345*	10.091*	0.361
IT-FR	3.495*	3.201*	5.171*	7.493*	0.317
IT-SP	-1.785*	1.781*	0.822	-18.284*	-0.145*
IT-NL	-1.573	1.375*	1.114	-0.022	-0.194*
IT-BE	1.801*	1.192	0.868	1.642	0.187*
IT-AU	3.188*	3.232*	2.925*	0.162	0.401
IT-GR	-2.998*	2.133*	3.647*	-4.498*	-0.336
IT-FI	1.771*	1.706*	0.426	0.000	0.236*
IT-PT	-2.234*	2.079*	1.555*	-2.939*	-0.224*
IT-IE	-3.293*	5.143*	3.830*	-0.003	-0.472
IT-LU	0.608	0.163	0.145	-0.160	0.051*

do not need the series be integrated of order one under the null hypothesis. The nonparametric test t-PS^1 proposed by Vogelsang (1998) also gets around this problem, but a price is paid in terms of power. Furthermore we find that when there is a highly persistent stationary component attached to a deterministic trend, the nonparametric $t_\beta-$test suffers little size distortion and in general is still more powerful than t-PS^1.

The standardized slope provides a simple nonparametric test of the null hypothesis of permanent slope. Unfortunately, a question mark hangs over the test because of its inability to reject a small deterministic slope.

We note that most of the literature on testing for trend in econometrics has focussed on deterministic trends and series that are stationary in first differences. Yet the assumption that underlying slopes, or growth rates, are constant is a strong one and is often implausible. This point is brought home by the global warming example. Using more recent data in Jones et al (1999), we find that an integrated random walk trend, with an AR(1) disturbance, provides excellent unconditional forecasts over a post-sample period from 1991 to 2004. On the other hand a deterministic trend model with AR(1) and white noise stationary components grossly underpredicts. A model in which the AR(1) is replaced by a random walk also underpredicts, but not by quite so much[5].

[5]The sum of absolute forecast errors from unconditional predictions for the smooth trend model is 7.11, for the deterministic trend it is 17.84 and for the random walk plus

Finally, the application of the tests to the series of the logarithm of relative prices provides evidence against the hypothesis of zero inflation differential between Italy and the majority of the countries in the euro area.

drift it is 12.99.

APPENDIX

Proof of Proposition 1

Define the independent partial sum processes $S_{0,[Tr]} = T^{-\frac{1}{2}} \sum_{j=1}^{[Tr]} \omega_j$, $r \in [0,1]$, where $[Tr]$ denotes the integer part of Tr, and $S_{1,[Tr]} = c_s^{-1} T^{\frac{1}{2}} \sum_{j=1}^{[Tr]} \zeta_j$. Under the local alternative hypothesis $H_{1,T}$ of (8), $S_{0,[Tr]}$ and $S_{1,[Tr]}$ satisfy a multivariate invariance principle such that

$$T^{-1/2} \left(S_{0,[Tr]},\, S_{1,[Tr]} \right) \Rightarrow \sigma (W_0(r), W_1(r)),\ r \in [0,1], \tag{18}$$

jointly, where $W_0(r)$ and $W_1(r)$ are independent standard Wiener processes. Then

$$T^{-\frac{1}{2}} \sum_{t=1}^{[Tr]} \Delta y_t = T^{-\frac{1}{2}} \sum_{t=1}^{[Tr]} \omega_t + c_d \sigma_L [Tr]/T + T^{-\frac{1}{2}} \sum_{t=1}^{[Tr]} \sum_{j=1}^{t-1} \zeta_j \Rightarrow \sigma_L V(r; c_d, c_s)$$

where, for $r \in [0,1]$, $V(r; c_d, c_s) = W_0(r) + c_d r + c_s \int_0^r W_1(s) ds$, and $T^{\frac{1}{2}} \widehat{\beta} \Rightarrow \sigma_L V(1; c_d, c_s)$. As $\widehat{\sigma}_L^2(m) \xrightarrow{p} \sigma_L^2$, see Stock (1994, page 2799), an application of the continuous mapping theorem (CMT) gives (9)-(10)-(11).

As $\Delta y_t - \widehat{\beta} = \left(\omega_t + \beta_{t-1} \right) - T^{-1} \sum_{j=1}^{T} \left(\omega_j - \beta_{j-1} \right)$, by (18) and the CMT we obtain that, under $H_{1,T}$, $T^{-\frac{1}{2}} \sum_{t=1}^{[Tr]} \left(\Delta y_t - \widehat{\beta} \right) \Rightarrow \sigma_L V^*(r; c_s)$, where $V^*(r; c_s) = W_0(r) - r W_0(1) + c_s \int_0^r \left(W_1(s) - \int_0^1 W_1(u) du \right) ds$. Thus an application of the CMT delivers (12). Note that, since the statistic $\zeta^1(m)$ is constructed with demeaned first differenced, its limiting distribution is not influenced by the presence of a (local or fixed) slope.

Proof of Proposition 2

Under $\beta_0 \neq 0$ and $\sigma_\zeta^2 = 0$, it is easy to see that $\widehat{\beta}$ is square-root consistent and that the partial sum process $\sum_{t=1}^{[Tr]} \Delta y_t$ is of $O_p(T)$. Since it still holds that $\widehat{\sigma}_L^2(m) \xrightarrow{p} \sigma_L^2$, we immediately obtain that $t_\beta(m)$ is of $O_p\left(T^{\frac{1}{2}}\right)$, while $\zeta_F^0(m)$ and $\zeta_R^0(m)$ are of $O_p(T)$. Then, since the limiting distribution of $\zeta^1(m)$ is unaffected by the presence of a slope, it readily follows that $\zeta^1(m) \xrightarrow{d} \int_0^1 V^*(r; 0)^2 dr$.

Under $\beta_0 = 0$ and $\sigma_\zeta^2 > 0$, the partial sum processes $\sum_{t=1}^{[Tr]} \Delta y_t$ and $\sum_{t=1}^{[Tr]} \left(\Delta y_t - \widehat{\beta}\right)$ are of $O_p\left(T^{\frac{3}{2}}\right)$, while $\widehat{\beta} = O_p\left(T^{\frac{1}{2}}\right)$. Then, since $\widehat{\sigma}_L^2(m) = O_p\left(Tm\right)$, compare Kwiatowski et al. (1992, p. 168), it follows that $t_\beta(m)$ is of $O_p\left((T/m)^{\frac{1}{2}}\right)$ while $\zeta_F^0(m)$, $\zeta_F^0(m)$, $\zeta^1(m)$ are of $O_p\left(T/m\right)$.

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