Markov Regime Switching and Unit-Root Tests

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We investigate the power and size performance of unit-root tests when the data undergo Markov regime switching. All tests, including those robust to a single break in trend growth rate, have low power against a process with a Markov-switching trend. Under the null hypothesis, we find that previously documented size distortions in Dickey–Fuller-type tests caused by a single break in trend growth rate or variance do not generalize to most parameterizations of Markov switching in trend or variance. However, Markov switching in variance can lead to overrejection in tests allowing for a single break in the level of trend.

KEY WORDS: Business-cycle asymmetry; Deterministic trend; Heteroscedasticity; Stochastic trend.

For the past 20 years the question of whether various economic time series have a unit root or are (trend) stationary has generated much research. Using standard tests, many researchers are unable to reject the unit-root null hypothesis for macroeconomic and financial time series such as gross domestic product (GDP), interest rates, and exchange rates (Nelson and Plosser 1982). Perron (1989) argued that the evidence in favor of unit roots has been overstated, because standard tests have low power against trend-stationary alternatives with structural breaks in trend level or growth rate. Perron remedied this problem by modifying the augmented Dickey–Fuller test with dummy variables to account for a single structural break. Christiano (1992), Banerjee, Lumsdaine, and Stock (1992), and Zivot and Andrews (1992) extended this methodology to endogenous estimation of the break date, while Lumsdaine and Papell (1997) considered a test robust to two structural breaks. Hereafter we will refer to this class of tests as Perron-type tests. Leybourne, Mills, and Newbold (1998) and Hamori and Tokihisa (1997) demonstrated a converse problem—that standard unit-root tests reject too often when there is a single structural break in trend or variance under the null hypothesis.

Although most of the literature has focused on the effects of a fixed number of structural breaks on unit-root tests, there is a growing consensus that the number of regime changes in economic time series might be better modeled as arising from a probabilistic process. To this end, many authors have successfully used Hamilton's (1989) Markov-switching model to capture regime change in a diverse set of macroeconomic and financial time series. It is thus natural to ask what effects Markov-switching regime change might have on unit-root tests, including the Perron-type tests developed to mitigate the effects of a fixed number of structural breaks.

Examples in which this issue might be relevant are not hard to find. Evans and Wachtel (1993) suggested an I(1) Markov-switching trend model for prices after standard unit-root tests on the price level failed to reject. Garcia and Perron (1996) argued for an I(0) Markov-switching trend and variance model of inflation, and real interest rates based on unit-root tests performed by Perron (1990) suggested that these series were I(0) if one break in the level of trend is allowed. Finally, many studies that employ a Markov-switching variance or trend growth rate simply assume a unit root in the series of interest without any pretesting, most likely because unit-root tests from previous studies suggest that the series are I(1). Examples include Hamilton's (1989) original article for gross national product (GNP), Cecchetti and Mark (1990) for consumption and dividends, and Engel (1994) for the nominal exchange rate.

In this study we investigate the effects of several types of Markov regime switching on unit-root tests, focusing on regime change in trend growth rate and variance, the form of structural change most often considered in the macroeconomics and finance literature. The literature surrounding structural breaks and unit-root tests provides insight into the size and power effects of a fixed number of breaks in trend growth rate on standard unit-root tests. However, it is not clear that these results generalize to the case of endogenous, Markov-switching breaks in trend. Perhaps the closest to addressing this question is the work of Balke and Fomby (1991), who demonstrated that standard unit-root tests continue to have low power when a series has endogenous, probabilistic breaks in trend growth rate. However, the process driving their breaks is an independent Bernoulli process, not a Markov-switching process, and they did not consider the performance of Perron-type tests. With regards to regime change in variance, several authors have considered the effects of the generalized autoregressive conditional heteroscedasticity (GARCH)-type heteroscedasticity on unit-root tests—

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for example, Pantula (1988), Kim and Schmidt (1993), Seo (1999), and Hecq (1995), the latter considering the effects on Perron-type tests. However, the effects of Markov switching in variance have not been considered. The only studies we are aware of investigating the effects of a Markov regime change in a testing framework are those of Evans and Lewis (1993) and Hall, Psaradakis, and Sola (1997), who concluded that Markov switching in trend growth rate or in the cointegrating vector will weaken the evidence in favor of cointegration in a bivariate system.

This article is organized as follows: In Section 1 we evaluate the performance of unit-root tests when the true data-generating process undergoes regime switching in trend growth rate but is otherwise I(0). In line with previous literature, we find that standard unit-root tests do a poor job of distinguishing this model from an integrated process. However, we also find that Perron-type tests have low power in this case. The Markov-switching trend model has often been used to model business-cycle asymmetry. Thus, we also consider alternative Markov-switching models of business-cycle asymmetry, in particular a model by Kim and Nelson (1999) that allows regime switching in the transitory component. Unit-root tests have very good power against this generating process, indicating that the true nature of nonlinearities in the business cycle is very important for what effects these nonlinearities have on unit-root tests. Finally, we briefly consider a model with Markov-switching autoregressive parameters. Such a model, with one regime an I(1) process and the other stationary, has been used by several authors—for example, Ang and Bekaert (1998)—to model interest rates. Standard tests have very low power against this process for empirically plausible parameterizations. In Section 2 we evaluate the performance of unit-root tests when the true data-generating process is I(1) in addition to the Markov switching. The size distortions pointed out in the literature for a single break in trend growth rate or variance do not generalize to most parameterizations of Markov switching. However, similar to the findings of Hecq (1995) for integrated GARCH (IGARCH) errors, Markov switching in variance can cause significant overrejection in Perron-type tests that allow for a single structural break in level. Section 3 concludes.

1. THE POWER OF UNIT–ROOT TESTS AGAINST REGIME–SWITCHING ALTERNATIVES

1.1 Regime Switching in the Trend Component

In this section we investigate the power of unit-root tests, including Perron-type tests, when the true process is I(0) conditional on a Markov-switching trend growth rate. To begin, consider the following data-generating process:

$$
\begin{aligned}
y_t &= \tau_t + c_t \\
\tau_t &= \mu_t + \tau_{t-1} \\
\mu_t &= \mu_1 S_t + \mu_0(1 - S_t) \\
\phi(L)c_t &= \varepsilon_t,\ \varepsilon_t \sim \text{iid}(0, \sigma_\varepsilon^2),
\end{aligned}
\tag{1}
$$

where S_t is a discrete, unobserved state variable that takes on the value 0 or 1, τ_t is a trend component with a switching growth rate, and $\phi(L)$ is a lag polynomial with either all roots outside the unit circle or one root on the unit circle and the rest outside. In this article, we consider the case in which S_t is first-order Markov switching. Here, the value of S_t at time t depends only on its value at time $t-1$, such that $P(S_t = 1 \mid S_{t-1} = 1) = p_{11}$ and $P(S_t = 0 \mid S_{t-1} = 0) = p_{00}$.

The model in (1) is a version of the models given by Hamilton (1989) and Lam (1990). Hamilton (1989) restricted one root of $\phi(L)$ to unity; that is, c_t had a stochastic trend. We will consider Hamilton's version of (1) in Section 2. Lam (1990) generalized Hamilton's model to allow c_t to (possibly) be a stationary autoregressive process. In this section we consider the performance of unit-root tests in this case, where all roots of $\phi(L)$ lie outside the unit circle. Here, innovations do not have permanent effects in the periods between shifts in the growth rate of trend. For some intuition into how unit-root tests will perform at distinguishing this model from the I(1) null, consider the alternative representation of the Markov trend function, τ_t:

$$\tau_t = \tau_0 + \mu_0{}^* t + (\mu_1 - \mu_0)\sum_{j=1}^{t} S_j.$$

Then, setting $\tau_0 = 0$,

$$
\begin{aligned}
y_t &= \mu_0{}^* t + RT_t + c_t \\
RT_t &= RT_{t-1} + (\mu_1 - \mu_0)^* S_t.
\end{aligned}
\tag{2}
$$

Here y_t is written as the sum of a deterministic trend, $\mu_0{}^* t$, a stochastic trend, RT_t, and a stationary component, c_t. The stochastic trend is introduced because the effects of the discrete shocks from the switching trend, $(\mu_1 - \mu_0)^* S_t$, are permanently reflected in the level of RT_t. This stochastic trend is different from an integrated process in the traditional sense that it does not necessarily change each period. It is similar to the integrated case in that first-differencing y_t eliminates the stochastic trend, leaving only a Markov-switching mean.

To assess the power of unit-root tests against the process given in (2), we perform Monte Carlo simulations for both standard and Perron-type unit-root tests. We parameterize the experiments based on the observation that the tests should do a better job of identifying the alternative given by (2) when the proportion of the variance of changes in y_t given by the stochastic trend, RT_t, is smaller rather than larger. The variance of innovations to RT_t is given by $(\mu_1 - \mu_0)^2(p - p^2)$, where $p = E(S_t = 1) = (1 - p_{00})/(2 - p_{00} - p_{11})$. Lam (1990) found that 37% of the variance of growth rates in real U.S. GNP is due to RT_t. We thus chose parameter values that will yield this 37% proportion when $p_{11} = 0.5$ and $p_{00} = 0.95$, the transition probability estimates found by Lam. These parameter values are $\mu_0 = 1$, $\mu_1 = -1.5$, $\sigma_\epsilon^2 \sim \text{N}(0, 0.4)$, and $\phi(L) = 1$. For each unit-root test, 1,000 Monte Carlo simulations were performed with two sample sizes, $T = 200$ and $T = 500$, and the initial values of S_t and y_t set equal to 0. To set p_{11} and p_{00}, we appeal to an existing literature (Hamilton 1989; Lam 1990; Diebold and Rudebusch 1996; Engel 1994), which has found for various monthly and quarterly series that one state is highly persistent, generally having a transition probability above 0.9, while the other is somewhat less persistent, although still usually having a transition probability of

0.5 or greater. We thus consider the following values of p_{00}: 0.9, 0.95, 0.98 and of p_{11}: 0.5, 0.6, 0.7, 0.8, 0.9, 0.95, 0.98.

1.1.1 Augmented Dickey–Fuller Test. We first consider the power of the augmented Dickey-Fuller, hereafter ADF, test (Dickey and Fuller 1979; Said and Dickey 1984) against the alternative hypothesis given in (2). We consider the ADF test based on the t statistic associated with the null hypothesis $\rho = 1$ from the test regression

$$y_t = c + \rho y_{t-1} + \beta t + \sum_{j=1}^{k} \phi_j \Delta y_{t-j} + \eta_t \tag{3}$$

with the lag length, k, chosen by the backward lag-length selection procedure given by Campbell and Perron (1991) with a maximum lag length, $\bar{k}$, set equal to the lower integer bound of $T^{1/3}$ as suggested by Said and Dickey (1984).

As would be expected from the existing literature, the ability of the ADF tests to distinguish the regime-switching trend-stationary alternative given in (2) is quite poor. Table 1 shows the rejection probabilities for the 5% nominal-size ADF test. For the $T = 200$ case, the test never rejects above 35%, only rejects above 20% for 6 of the 21 combinations of the transition probabilities considered, and often rejects in the 5–10% range. The test tends to perform better when one transition probability dominates the other; for example, for the values of the transition probabilities estimated by Lam (1990) for U.S. real GDP, $p_{00} = .95$ and $p_{11} = 0.5$, the test rejects with a 31% frequency. This is because the variance of innovations to the stochastic trend, RT_t, is smaller the larger the difference between the transition probabilities; that is, $(\mu_1 - \mu_0)^2(p - p^2)$ is a decreasing function of $|p_{00} - p_{11}|$. Intuitively, as one state becomes increasingly dominant, the process more closely resembles one with constant trend growth rate. For the larger sample size, the ADF test has even lower power, rejecting at 10% or less frequency in all cases. This is not surprising because the larger sample size gives the ADF test more opportunity to detect the stochastic trend, RT_t.

1.1.2 Perron-Type Tests. Since the influential work of Perron (1989), a large number of unit-root tests that allow for structural breaks in trend growth rate or level under the alternative have been developed. The objective of this research program is to develop tests with higher power against broken-trend-stationary alternatives. These tests are robust to a fixed number of structural breaks, usually one. However, there has been some argument in the literature that when there are multiple structural breaks in trend growth rate it may be sufficient to simply account for the largest of these breaks; see, for example, Garcia and Perron (1996, p. 113). We are thus interested in whether such tests provide increased power against an alternative with a Markov-switching trend growth rate. Here we consider two such tests that assume a single break in the growth rate of the trend function occurring at an unknown date, one given by Perron (1994, 1997), hereafter the Perron test, and the other given by Zivot and Andrews (1992), hereafter the ZA test. The Perron test assumes a single break in trend growth rate under both the null and alternative hypotheses and specifies the break as an additive outlier, meaning that the full effects of the break are immediately reflected. The test is based on the regressions in equations (3a) and (3b) of Perron (1997). The ZA test assumes a single break in trend growth rate under only the alternative hypothesis and specifies the break as an innovational outlier, meaning that the full effects of the change are felt over time. The test is based on the regression in equation 2′ of Zivot and Andrews (1992). For both tests the date of the structural break was estimated as the date that provides the most evidence against the null hypothesis; see Zivot and Andrews (1992) for details.

Tables 2–3 contain the rejection frequencies for 5%-nominal-size Perron and ZA tests. Interestingly, the Perron test performs worse than the ADF test for many of the cases considered. For example, when $T = 200$, the ADF test rejects more frequently for 17 of the 21 combinations of transition probabilities. For the transition probabilities estimated by Lam (1990) for real GDP, $p_{00} = .95$ and $p_{11} = 0.5$, the Perron test rejects 16% of the time versus 31% for the ADF test. The ZA test performs somewhat better, rejecting more frequently than the ADF test for 18 of the 21 combinations of transition probabilities considered when $T = 200$. However, the difference is not decisive: In 10 of these 18 cases, the ZA test is within 15% of the ADF test. In addition, the ZA test only rejects more than 40% of the time on four occasions and for over half the cases rejects at a less than 25% frequency. For the Lam (1990) transition probability estimates for real GDP, the ZA test rejects at a 19% frequency versus 31% for the ADF test. When $T = 500$ the tests have even lower power, usually rejecting at close to their nominal size.

Table 1. Empirical Power of a 5% Augmented Dickey-Fuller Test: True Process Has Markov Switching in Trend Growth Rate

T							
200				500			
	Power				Power		
p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$	p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$
0.5	0.13	0.31	0.33	0.5	0.05	0.06	0.08
0.6	0.24	0.31	0.16	0.6	0.05	0.05	0.09
0.7	0.15	0.08	0.25	0.7	0.08	0.10	0.07
0.8	0.06	0.05	0.12	0.8	0.05	0.06	0.04
0.9	0.06	0.04	0.20	0.9	0.05	0.04	0.06
0.95	0.06	0.05	0.10	0.95	0.05	0.06	0.06
0.98	0.14	0.11	0.08	0.98	0.09	0.09	0.03

Table 2. Empirical Power of a 5% Perron (1994,1997) Test: True Process Has Markov Switching in Trend Growth Rate

	T						
	200			500			
	Power				Power		
p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$	p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$
0.5	0.14	0.16	0.22	0.5	0.05	0.06	0.07
0.6	0.13	0.15	0.24	0.6	0.03	0.05	0.06
0.7	0.10	0.11	0.19	0.7	0.04	0.10	0.08
0.8	0.05	0.05	0.12	0.8	0.03	0.05	0.05
0.9	0.04	0.03	0.09	0.9	0.05	0.06	0.03
0.95	0.05	0.03	0.09	0.95	0.04	0.04	0.05
0.98	0.11	0.09	0.17	0.98	0.06	0.06	0.02

1.2 Regime Switching in the Transitory Component

Models with two-state Markov switching in trend growth rate, such as that discussed in the previous section, have been used extensively to model business-cycle asymmetry. One reason for its popularity is the ability of a regime-switching trend growth rate to capture the empirical observation that recessions are steeper and shorter than expansions. However, one implication of the two-state Markov-switching trend model is that recessions have permanent effects on the level of output; that is, the economy never recovers output lost during a recession. Many authors have provided evidence that this implication is not consistent with the data; instead, following steep, short recessions the economy seems to undergo a high-growth recovery phase to gain back what was lost; see, for example, Friedman (1969, 1993), Wynne and Balke (1992, 1996), and Sichel (1994). In other words, the business cycle is better characterized with three phases rather than two. Recently, Kim and Nelson (1999) used Markov regime switching in the transitory component of real GDP to capture this pattern of business-cycle asymmetry. Here we consider a trend-stationary version of their model:

$$\begin{aligned} y_t &= \tau_t + c_t \\ \tau_t &= \mu + \tau_{t-1} \\ \phi(L)c_t &= \gamma^* S_t + \varepsilon_t,\ \varepsilon_t \sim \text{iid}(0, \sigma_\varepsilon^2), \end{aligned} \tag{4}$$

where $\phi(L)$ has all roots outside the unit circle. Here, unlike the model in (1), the average growth rate of the deterministic trend, μ, is constant. Instead, regime switching occurs in the transitory component, c_t. If $\gamma < 0$, when $S_t = 1$ the level of the series is driven down into a steep recession. However, the recession is not permanent because past shocks from γ disappear through the autoregressive dynamics in the transitory component, causing a high growth recovery phase once S_t returns to 0. In the words of Friedman (1969, 1993), the economy is "plucked" downward during recession, bouncing back to trend following the recession.

The results of Kim and Nelson (1999) suggest that a model specifying recessions as "plucking" episodes provides as good as or better description of U.S. real GDP than a model with regime shifts in the trend component. However, given that the regime switching in (4) works through the transitory component, we would expect unit-root tests to have much better power against this alternative than the model in Section 1.1. To investigate this, we perform a Monte Carlo experiment with the ADF test. We parameterize the simulation based on the percentage of the variance of c_t coming from the "plucks" γ. Kim and Nelson (1999) found this percentage to be approximately 80% for real GDP for estimated transition probabilities of $p_{11} = .95$ and $p_{00} = .70$. When $\phi(L) = 1$, this percentage is given by

$$\frac{\gamma^2(p-p^2)}{\gamma^2(p-p^2)+\sigma_\varepsilon^2}. \tag{5}$$

Table 3. Empirical Power of a 5% Zivot–Andrews (1992) Test: True Process Has Markov Switching in Trend Growth Rate

	T						
	200			500			
	Power				Power		
p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$	p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$
0.5	0.24	0.19	0.73	0.5	0.07	0.10	0.13
0.6	0.14	0.43	0.43	0.6	0.05	0.07	0.10
0.7	0.40	0.25	0.35	0.7	0.01	0.06	0.10
0.8	0.10	0.36	0.35	0.8	0.10	0.10	0.14
0.9	0.16	0.09	0.20	0.9	0.05	0.07	0.10
0.95	0.12	0.09	0.18	0.95	0.04	0.07	0.10
0.98	0.24	0.28	0.33	0.98	0.12	0.14	0.10

To meet the 80% metric when $p_{11} = .95$ and $p_{00} = .70$, we parameterize the simulation with $\gamma = -1.0$ and $\varepsilon_t \sim \mathrm{N}(0, .04)$. We set $\mu = 0.8$, the average growth rate of real GDP over the Kim and Nelson sample. Again, we perform 1,000 Monte Carlo trials for the same range of transition probabilities as in Section 1.1.

Table 4 contains the rejection frequencies for the 5% ADF test. As expected, the ADF test performs very well, rejecting at close to 100% for the most empirically relevant values of the transition probabilities. For example, for the estimated transition probabilities found by Kim and Nelson for real GDP, $p_{11} = .95$ and $p_{00} = .70$, the ADF test rejects at a 99% frequency when $T = 200$ and a 100% frequency when $T = 500$. The power remains above 50% in all but one of the 21 combinations considered for $T = 200$ and in all cases for $T = 500$.

The differing performance of unit-root tests for the model in (1) versus the model in (4) is important in answering the question of whether real GDP has a unit root. If we believe that business-cycle nonlinearities are shifts in trend as shown by Lam (1990), these shifts will have significant deleterious effects on the power of unit-root tests, including Perron-type tests. If, however, these nonlinearities are better characterized as Friedman's "plucks," the power of unit-root tests will be unaffected. Instead, the only remaining sort of structural change relevant to unit-root tests will be long-run breaks, such as the much-discussed productivity slowdown. In this case Perron-type tests will still have an advantage over standard tests such as the ADF test. This points us to the importance of determining the true nature of business-cycle nonlinearities for deciding what classes of unit-root should be used in studies of real GDP.

1.3 Regime-Switching Autoregressive Coefficients

To this point we have investigated Markov switching taking the form of discrete disturbances to the trend or transitory component of a time series. Another popular formulation is Markov switching in the autoregressive parameters of a time series, an example of which is

$$
\begin{aligned}
y_t &= \mu_t + \rho_t y_{t-1} + \varepsilon_t \\
\mu_t &= \mu_1 S_t + \mu_0 (1 - S_t) \\
\rho_t &= \rho_1 S_t + \rho_0 (1 - S_t) \\
\varepsilon_t &\sim \mathrm{iid}(0, \sigma^2_{\varepsilon t}) \\
\sigma^2_{\varepsilon t} &= \sigma^2_{\varepsilon 1} S_t + \sigma^2_{\varepsilon 0} (1 - S_t).
\end{aligned} \tag{6}
$$

In (6), y_t follows a first-order autoregressive [AR(1)] process in which the autoregressive parameter, the constant term, and the variance of the error term all switch between two regimes. A popular version of (6) in the empirical literature specifies y_t to be I(1) in one regime and I(0) in the other; for example, $\rho_0 = 1$ and $|\rho_1| < 1$. Ang and Bekaert (1998) demonstrated that, as long as the I(0) regime has positive probabilities of occurring and persisting, in this case $(1 - p_{00}) \neq 0$ and $p_{11} \neq 0$, y_t is covariance stationary. This occasionally integrated model has been usefully employed to model interest rates. For example, Ang and Bekaert (1998) pointed out that the U.S. Federal Reserve tends to move short-term interest rates in a very persistent fashion during low-inflation periods. However, during high-inflation times, Federal Reserve interest-rate changes become less persistent and have higher variance.

For our purposes, we are interested in the ability of unit-root tests to distinguish the occasionally integrated model from the I(1) null hypothesis. To investigate this issue, we perform Monte Carlo simulations with the ADF test when the generating process is (6). We parameterize the Monte Carlo experiments to mimic the pattern of Federal Reserve interest-rate movements discussed previously. Thus, when $S_t = 0$ (low-inflation times), y_t is a random walk with no drift; that is, $\rho_0 = 1$, $\mu_0 = 0$, and $\varepsilon_t | S_t = 0 \sim \mathrm{N}(0, .25)$. When $S_t = 1$ (high-inflation times), y_t is a stationary AR(1) with positive mean and $\varepsilon_t | S_t = 1 \sim \mathrm{N}(0, 2.0)$. One would expect that unit-root tests would perform worse for more persistent values of the autoregressive parameter when $S_t = 1$. Thus, we consider three pairs of μ_1, ρ_1—(1.0, 0.8); (0.5, 0.9); (0.25, 0.95). In these pairs, μ_1 is altered to maintain a constant mean of 5 for y_t in the stationary state.

Tables 5–7 present the Monte Carlo simulations for the three pairings of μ_1, ρ_1, and the sample sizes $T = 200$ and $T = 500$. As would be expected, the tests perform better as ρ_1 decreases, as p_{11} increases relative to p_{00} [the less time that is spent in the I(1) state], and the larger the sample size

Table 4. Empirical Power of a 5% Augmented Dickey–Fuller Test: True Process Has Markov Switching in the Transitory Component

T = 200				T = 500			
	Power				Power		
p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$	p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$
0.5	1.00	1.00	1.00	0.5	1.00	1.00	1.00
0.6	1.00	1.00	0.99	0.6	1.00	1.00	1.00
0.7	1.00	0.99	0.99	0.7	1.00	1.00	1.00
0.8	0.99	0.98	0.97	0.8	1.00	1.00	1.00
0.9	0.95	0.87	0.79	0.9	1.00	1.00	0.99
0.95	0.88	0.70	0.53	0.95	1.00	0.99	0.94
0.98	0.81	0.52	0.30	0.98	1.00	0.94	0.62

Table 5. Empirical Power of a 5% Augmented Dickey–Fuller Test: True Process Has Markov Switching in the Autoregressive Parameters and $\mu_1 = 1.0$, $\rho_1 = 0.8$

T = 200				T = 500			
	Power				Power		
p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$	p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$
0.5	0.12	0.09	0.08	0.5	0.43	0.14	0.08
0.6	0.17	0.09	0.09	0.6	0.50	0.19	0.09
0.7	0.20	0.12	0.09	0.7	0.63	0.26	0.10
0.8	0.28	0.17	0.12	0.8	0.74	0.35	0.13
0.9	0.48	0.30	0.19	0.9	0.93	0.64	0.23
0.95	0.64	0.47	0.27	0.95	0.99	0.83	0.41
0.98	0.81	0.66	0.47	0.98	1.00	0.96	0.73

[the more data available for the test to detect the I(0) state]. In general, however, the tests perform very poorly for empirically plausible parameterizations. Of the 63 power statistics reported for the $T = 200$ cases, the test has power greater than 50% on only 3 occasions (all for the smallest value of ρ_1) and greater than 20% on only 17 occasions (10 of these for the smallest value of ρ_1). As the sample size increases, the performance of the test is fairly good for the lowest value of ρ_1 considered but is still poor for larger values of ρ_1. For example, Ang and Bekaert (1998) showed that the regime switches in U.S. interest rates roughly correspond to business-cycle frequencies. Depending on the frequency of the data, this corresponds to values of p_{00} between 0.9 and 0.95 and values of p_{11} between 0.5 and 0.9. For $\rho_1 = .9$ and $T = 500$, the ADF test has power greater than 40% over this range of transition probabilities on only one occasion.

2. REGIME–SWITCHING I(1) PROCESSES AND THE SIZE OF UNIT–ROOT TESTS

2.1 Regime Switching in the Trend Component and Variance

In Section 1.1 we were interested in the ability of unit-root tests to distinguish a process that was I(0) with a Markov-switching trend growth rate from an I(1) process. Here we will investigate what deleterious size effects a Markov-switching trend growth rate and variance in an otherwise I(1) process might have on unit-root tests. Consider the following model motivated by Hamilton (1989):

$$\begin{aligned} y_t &= \tau_t + c_t \\ \tau_t &= \mu_t + \tau_{t-1} \\ \mu_t &= \mu_0(1 - S_t) + \mu_1 S_t \\ \phi(L)c_t &= \varepsilon_t,\ \varepsilon_t \sim \text{iid}(0, \sigma^2_{\varepsilon t}) \\ \sigma^2_{\varepsilon t} &= \sigma^2_{\varepsilon 1} S_t + \sigma^2_{\varepsilon 0}(1 - S_t). \end{aligned} \tag{7}$$

Again, S_t is first-order Markov switching and τ_t is a deterministic trend component with a switching growth rate. As Hamilton (1989) did, we specify $\phi(L)$ to have one root on the unit circle and all other roots outside the unit circle so that shocks to y_t in between the Markov-switching trend breaks have permanent effects on the level of the series. We also allow the variance of the error term to undergo regime switching.

To simplify matters, we set $\phi(L) = (1 - L)$. The model in (7) can then be written with a constant growth rate and serially

Table 6. Empirical Power of a 5% Augmented Dickey–Fuller Test: True Process Has Markov Switching in the Autoregressive Parameters and $\mu_1 = 0.5$, $\rho_1 = 0.9$

T = 200				T = 500			
	Power				Power		
p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$	p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$
0.5	0.08	0.07	0.07	0.5	0.15	0.09	0.05
0.6	0.10	0.10	0.07	0.6	0.19	0.09	0.07
0.7	0.11	0.10	0.10	0.7	0.27	0.13	0.09
0.8	0.15	0.12	0.09	0.8	0.39	0.17	0.10
0.9	0.22	0.17	0.13	0.9	0.64	0.35	0.16
0.95	0.32	0.27	0.21	0.95	0.86	0.61	0.32
0.98	0.42	0.37	0.29	0.98	0.95	0.85	0.59

Table 7. Empirical Power of a 5% Augmented Dickey–Fuller Test: True Process Has Markov Switching in the Autoregressive Parameters and $\mu_1 = 0.25$, $\rho_1 = 0.95$

	T						
	200				500		
	Power				Power		
p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$	p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$
0.5	0.07	0.06	0.08	0.5	0.10	0.08	0.06
0.6	0.09	0.08	0.09	0.6	0.09	0.07	0.05
0.7	0.09	0.08	0.09	0.7	0.12	0.08	0.06
0.8	0.09	0.11	0.11	0.8	0.15	0.11	0.09
0.9	0.11	0.13	0.10	0.9	0.29	0.15	0.13
0.95	0.17	0.14	0.15	0.95	0.44	0.27	0.19
0.98	0.17	0.17	0.17	0.98	0.59	0.45	0.34

correlated, conditionally heteroscedastic errors:

$$\Delta y_t = \mu + e_t$$
$$e_t = (\mu_1 - \mu)S_t + (\mu_0 - \mu)(1 - S_t) + \varepsilon_t. \tag{8}$$

To make $E(e_t) = 0$, choose $\mu = (\mu_1 - \mu_0)p + \mu_0$. Substituting in the chosen expression for μ, we arrive at the autocovariance function

$$\begin{aligned}\operatorname{cov}(e_t, e_{t-k}) &= (\mu_1 - \mu_0)^2 E(S_t - p)(S_{t-k} - p)\\ &= (\mu_1 - \mu_0)^2 \operatorname{cov}(S_t, S_{t-k}).\end{aligned} \tag{9}$$

Moreover, conditional on S_t, e_t has a time-varying variance due to the heteroscedasticity of ε_t:

$$\operatorname{var}(e_t \mid S_t = j) = \sigma^2_{\varepsilon j}, \qquad j = 0, 1. \tag{10}$$

A result from the theory of Markov processes tells us that $P(S_t = 1 \mid S_{t-k} = 1)$ and $P(S_t = 0 \mid S_{t-k} = 0)$ converge to the unconditional probabilities p and $(1 - p)$ at a geometric rate. Then, noting that $\operatorname{cov}(S_t, S_{t-k}) = (p^* P(S_t = 1 \mid S_{t-k} = 1) - p^2)$, we have $\operatorname{cov}(e_t, e_{t-k}) = (\mu_1 - \mu_0)^2 (p^* P(S_t = 1 \mid S_{t-k} = 1) - p^2) \to 0$ geometrically. Thus, the model in (7) can be written with constant trend growth rate and errors exhibiting serial correlation that dies off geometrically. It should be noted that this result is entirely due to the modeling of breaks in the trend function as endogenous, probabilistic events. It does not hold true in models assuming a fixed number of structural breaks in trend growth rate such as the cases considered by Perron (1989) and Zivot and Andrews (1992) among others.

Several previous studies (e.g., Schwert 1989) have investigated the properties of unit-root tests under various forms of autoregressive moving average (ARMA) innovations. Therefore, we will find an ARMA process a useful alternative representation of e_t. Consider the following stationary AR(1) representation of S_t, given by Hamilton (1989):

$$S_t = (1 - p_{00}) + \theta S_{t-1} + \omega_t$$
$$\theta = -1 + p_{00} + p_{11}, \tag{11}$$

where, conditional on $S_{t-1} = 1$, $\omega_t = (1 - p_{11})$ with probability p_{11} and $\omega_t = -p_{11}$ with probability $1 - p_{11}$ and conditional on $S_{t-1} = 0$, $\omega_t = -(1 - p_{00})$ with probability p_{00} and $\omega_t = p_{00}$ with probability $1 - p_{00}$. Hamilton (1989) showed that the error term, ω_t, has $E(\omega_t) = 0$, $E(\omega_t^2) = \sigma^2_\omega = p_{11}(1 - p_{11})p + p_{00}(1 - p_{00})(1 - p)$ and is uncorrelated in that $E(\omega_t \mid \omega_{t-j}) = 0$ for all four possible values of ω_{t-j} and $j = 1, 2, \ldots$ Using (11), note that

$$\begin{aligned}e_t - \theta e_{t-1} &= d + b\omega_t + \varepsilon_t - \theta\varepsilon_{t-1};\\ d &= (\mu_1 - \mu)(1 - p_{00}) + (\mu_0 - \mu)(1 - p_{11}),\\ b &= (\mu_1 - \mu_0).\end{aligned} \tag{12}$$

The term on the left side of (12) is an AR(1), while the term on the right side has the autocovariance function of an MA(1) in that it is 0 after the first lag. Thus, e_t follows an ARMA(1, 1) process.

To determine the effects of the regime switching in trend growth rate and variance on unit-root tests, we perform Monte Carlo experiments for the three tests discussed in Section 1—the ADF test, the Perron test, and the ZA test. We consider two cases, one in which there is only regime switching in trend growth rate and one in which there is only switching in variance. To parameterize the trend switching case, we set the parameters to yield a specified amount of serial correlation as measured by the first-order autocorrelation of e_t:

$$\operatorname{corr}(e_t, e_{t-1}) = \frac{(\mu_1 - \mu_0)^2 (p^* p_{11} - p^2)}{(\mu_1 - \mu_0)^2 (p - p^2) + p\sigma^2_{\nu 1} + (1 - p)\sigma^2_{\nu 0}}, \tag{13}$$

where the denominator is the unconditional variance of e_t. We set $\mu_0 = 1$, $\mu_1 = 5$, $\varepsilon_t \sim \mathrm{N}(0, \sigma^2_{\varepsilon t})$, and $\sigma_{\varepsilon 0} = \sigma_{\varepsilon 1} = 1$ to yield a value of (13) equal to .50 for $p_{11} = 0.9$ and $p_{00} = 0.7$, the transition probability estimates for U.S. real GNP found by Hamilton (1989). This level of autocorrelation is similar to that found in the existing literature. For example, the value of (13) for U.S. real GNP reported by Hamilton is 0.38 while Engel's (1994) parameter estimates for the Japanese/French exchange rate suggest a value of (13) equal to 0.50.

For the variance switching case, we set $\varepsilon_t \sim \mathrm{N}(0, \sigma^2_{\varepsilon t})$, $\sigma_{\varepsilon 1}/\sigma_{\varepsilon 0} = 3/1$, and $\mu_0 = \mu_1 = 1.0$. This level of heteroscedasticity is quite reasonable for asset prices; for example, Turner, et al. (1989) reported $\sigma_{\varepsilon 1}/\sigma_{\varepsilon 0} = 2.6$ for stock returns, while Engel (1994) reported much higher ratios for several U.S.

exchange rates. However, this level of heteroscedasticity is overstated for series such as real GDP. Thus, our results for the switching variance case have more relevance for financial time series than for macroeconomic quantities. Again, we consider the same range of transition probabilities and sample sizes as in Section 1. Each Monte Carlo experiment is composed of 1,000 trials with initial values of S_t and y_t set equal to 0.

We begin by considering the effects of the Markov-switching trend growth rate in (7). Because this regime switching simply introduces serial correlation into an otherwise I(1) process, we can appeal to the large literature evaluating the effects of serial correlation on unit-root tests. Schwert (1989) demonstrated that the ADF test performs well in the presence of ARMA errors such as those in (12). However, Leybourne et al. (1998) showed that the ADF test tends to overreject the null hypothesis when there is a single break in trend growth rate that occurs early in the sample. Thus, we expect the ADF test to overreject for parameterizations of (7) that yield few breaks, with one occurring early in the sample. The question of interest is for how broad a range of the Markov-switching parameterizations this result holds. Table 8 presents the rejection frequencies for the 5% ADF test. Note that only for $T = 200$ and $p_{11} = 0.98$ are the size distortions pointed out by Leybourne et al. present. For most parameterizations, the ADF test has size close to its nominal size and in general is slightly oversized. This is likely due to the Campbell–Perron lag-selection procedure, which was documented by Hall (1994), to cause slight overrejection.

Next we consider the Perron and ZA tests that allow for a single break in trend growth rate under the alternative. Table 9 contains the rejection frequencies for the 5% Perron test. The Perron test performs similarly to the ADF test for most parametrizations, which is not surprising given that it captures serial correlation in the same way as the ADF test. Notably, the Perron test performs better than the ADF test when $p_{11} = 0.98$. This is most likely because the Perron test is robust to a single break in trend growth rate under the null hypothesis as well as the alternative, making the Leybourne et al. (1998) critique not as relevant. Table 10 demonstrates that the ZA test can be significantly oversized when there are only a small number of breaks—that is, for large values of p_{00} or p_{11}. This is because the distribution of the ZA test is derived assuming a null with no structural change, meaning the presence of a small number of structural breaks will violate this null hypothesis and lead to overrejections. This issue is not as serious for the large sample size; $T = 500$. Both the Perron and the ZA tests perform similarly to the ADF test for this larger sample size.

We now move to the simulations investigating Markov switching in variance. Many authors have investigated the effects of various forms of heteroscedasticity on unit-root tests, including Pantula (1988), Kim and Schmidt (1993), and Seo (1999). Provided that the heteroscedasticity meets certain conditions, given explicitly by Hamori and Tokihisa (1997), heteroscedasticity does not create size distortions for standard unit-root tests. Piger (2000) showed that Markov-switching heteroscedasticity meets these conditions, suggesting that standard unit-root tests should perform well. However, we are still interested in investigating two scenarios. First, Hamori and Tokihisa (1997) showed that a single break in variance causes Dickey–Fuler-type tests to be oversized. Thus, we might expect that certain parameterizations of Markov switching in variance that yield a small number of breaks will cause size distortions in the ADF test. Table 11 demonstrates that this is not the case. The ADF test is reasonably sized for even large values of p_{00} and p_{11}, suggesting that the result of Hamori and Tokihisa fades quickly when more than one break is allowed.

Second, Hecq (1995) pointed out, for the case of IGARCH errors, that periods of high and low variance in an integrated process can lead to the illusion of breaks in the level of trend. Tests that are robust to a structural break in level under the alternative can spuriously detect such breaks and overreject as a result. We thus might expect versions of the Perron and the ZA tests that allow for a break in the level of trend to be oversized in the presence of Markov-switching heteroscedasticity. To investigate this issue, we consider the performance of the Perron test allowing for a single break in the level of trend under both the null and the alternative, based on equations (14) and (17) of Perron (1994), and the ZA test allowing for a single break in the level of trend under the alternative, given by equation 1′ of Zivot and Andrews (1992). As Tables 12–13 make clear, the size distortions can be significant for certain parameterizations of p_{00} and p_{11}. For example, for $T = 200$ the 5% nominal-size Perron test rejects at a greater than 10% frequency in all but one of the 21 combinations of transition

Table 8. Empirical Size of a 5% Augmented Dickey–Fuller Test: True Process Has Markov Switching in Trend Growth Rate

	T						
200				500			
	Size				Size		
p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$	p_{11}	$p_{00} = 0.9$	$p_{00} = 0.95$	$p_{00} = 0.98$
0.5	0.05	0.07	0.05	0.5	0.07	0.06	0.05
0.6	0.07	0.07	0.07	0.6	0.05	0.05	0.05
0.7	0.05	0.07	0.06	0.7	0.05	0.05	0.05
0.8	0.06	0.06	0.07	0.8	0.04	0.06	0.05
0.9	0.06	0.03	0.06	0.9	0.05	0.04	0.05
0.95	0.06	0.05	0.06	0.95	0.05	0.05	0.04
0.98	0.15	0.10	0.06	0.98	0.07	0.07	0.03

Table 9. Empirical Size of a 5% Perron (1994,1997) Test: True Process Has Markov Switching in Trend Growth Rate

	T						
	200			500			
	Size				Size		
p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$	p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$
0.5	0.07	0.05	0.06	0.5	0.05	0.03	0.08
0.6	0.06	0.04	0.06	0.6	0.04	0.04	0.04
0.7	0.05	0.05	0.06	0.7	0.05	0.06	0.04
0.8	0.05	0.06	0.05	0.8	0.04	0.06	0.03
0.9	0.05	0.04	0.06	0.9	0.05	0.05	0.03
0.95	0.05	0.03	0.05	0.95	0.03	0.05	0.02
0.98	0.09	0.05	0.04	0.98	0.05	0.05	0.02

probabilities considered and greater than 15% for 8 of the 21 combinations. The ZA test rejects at a greater than 10% frequency in all but one case and greater than 15% in more than half of the cases when $T = 200$. Both tests perform somewhat better when $T = 500$ but are still oversized.

2.2 Regime Switching in the Transitory Component

In Section 1.2 we discussed how different Markov-switching models of business-cycle asymmetry can have very different implications for the effects of asymmetry on the power of unit-root tests. Here we examine the difference this modeling choice has for the size of unit-root tests. Consider the following I(1) version of the model presented in Section 1.2:

$$
\begin{aligned}
y_t &= \tau_t + z_t + c_t \\
\tau_t &= \mu + \tau_{t-1} \\
z_t &= z_{t-1} + \nu_t,\ \nu_t \sim \mathrm{iid}(0, \sigma_\nu^2) \\
\phi(L)c_t &= \gamma^* S_t + \varepsilon_t,\ \varepsilon_t \sim \mathrm{iid}(0, \sigma_\varepsilon^2),
\end{aligned}
\tag{14}
$$

where $\phi(L)$ has all roots outside the unit circle. Here y_t is the sum of a deterministic trend with constant drift, a random-walk component, and a stationary autoregressive component that, assuming $\gamma < 0$, is "plucked" downward whenever $S_t = 1$. To see the effects the process in (14) might have on the size of unit-root tests, rewrite (14) in first differences assuming $\phi(L) = 1$:

$$
\begin{aligned}
\Delta y_t &= \mu + e_t^* \\
e_t^* &= \nu_t + \Delta\varepsilon_t + \gamma^* \Delta S_t.
\end{aligned}
\tag{15}
$$

The process can thus be written in first differences with constant drift and an error term that is augmented by a Markov-switching component. The Markov switching introduces additional serial correlation into the process—namely, the first difference of S_t. One interesting note is the similarity of this case to the additive-outlier literature discussed by Franses and Haldrup (1994), among others. The parameter γ would correspond to an additive outlier in the case in which S_t was serially uncorrelated as opposed to being a Markov-switching process. As Maddala and Yin (1997) and Vogelsang (1999) pointed out, the first difference of S_t in (15) would then introduce a first-order moving average [MA(1)] component into the first difference of y_t. For smaller values of γ, the additional serial correlation introduced in both the Markov-switching and additive-outlier cases is captured by tests such as the ADF test and does not cause overrejections. However, as the size of γ increases, the contribution of the transitory component to the variance of Δy_t increases relative to the contribution of the stochastic trend component. This can even-

Table 10. Empirical Size of a 5% Zivot–Andrews (1992) Test: True Process Has Markov Switching in Trend Growth Rate

	T						
	200			500			
	Size				Size		
p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$	p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$
0.5	0.09	0.10	0.09	0.5	0.07	0.04	0.05
0.6	0.08	0.09	0.10	0.6	0.07	0.06	0.07
0.7	0.09	0.10	0.13	0.7	0.06	0.09	0.07
0.8	0.09	0.09	0.13	0.8	0.05	0.08	0.09
0.9	0.08	0.09	0.15	0.9	0.11	0.08	0.09
0.95	0.11	0.10	0.19	0.95	0.05	0.07	0.08
0.98	0.24	0.23	0.29	0.98	0.10	0.10	0.08

Table 11. Empirical Size of a 5% Augmented Dickey–Fuller Test: True Process Has Markov Switching in Variance

	T						
	200			500			
	Size				Size		
p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$	p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$
0.5	0.07	0.07	0.05	0.5	0.06	0.05	0.05
0.6	0.08	0.08	0.06	0.6	0.07	0.06	0.04
0.7	0.06	0.07	0.07	0.7	0.05	0.05	0.06
0.8	0.07	0.07	0.08	0.8	0.06	0.06	0.07
0.9	0.07	0.09	0.08	0.9	0.05	0.06	0.05
0.95	0.06	0.08	0.08	0.95	0.06	0.04	0.07
0.98	0.05	0.06	0.08	0.98	0.05	0.06	0.06

tually lead to spurious rejections from unit-root tests if the variance of the transitory component begins to dominate. The question is whether parameterizations of (14) corresponding to U.S. business cycles generate such spurious rejections.

To investigate this issue, we perform Monte Carlo experiments to investigate the performance of the ADF test when the generating process is (14). We parameterize the simulation using parameter estimates from Kim and Nelson (1999) for U.S. real GDP. That is we set $\mu = .8$, $\nu_t \sim \mathrm{N}(0, 0.4)$, $\gamma = -1.1$, $\varepsilon_t \sim \mathrm{N}(0, 0.04)$, and the lag order of $\phi(L)$ equal to 2 with $\phi_1 = 1.26$ and $\phi_2 = -0.46$ Table 14 demonstrates that this level of "plucking" is indeed large enough to cause spurious rejections in the ADF test. These rejections are fairly severe; the 5% ADF test rejects at a more than 10% frequency for all but one of the combinations of the transition probabilities considered in Table 14. For $T = 200$, the rejections climb above 30% for 9 of the 21 cases, while for $T = 500$, rejections are larger than 30% on 8 occasions. Again, this points out that whether nonlinearities in the U.S. business cycle take the form of shifts in trend or "plucks" in the transitory component can have large implications for the performance of unit-root tests applied to U.S. output series.

3. CONCLUSION

We have investigated the performance of unit-root tests when the true process undergoes various types of Markov-switching regime change. We consider both processes that are I(0) and I(1) in the periods between the regime switching. Our main findings are as follows:

1. In line with previous literature, the ADF test does a poor job of distinguishing an I(0) process with Markov-switching breaks in trend growth rate from an I(1) process. Interestingly, however, tests designed to be robust to a single structural break in trend growth rate under the alternative also have very low power in this case.

2. When the true process is I(1) and undergoes Markov switching in both trend growth rate and variance, ADF tests have approximately the correct size for almost all combinations of transition probabilities. This demonstrates that studies documenting size distortions from a single break in trend growth and variance do not generalize to multiple, probabilistic breaks. However, tests robust to a single break in level overreject the null hypothesis when there is Markov switching in variance.

3. When modeling business-cycle asymmetry, an alternative to Markov switching in trend growth rate, as shown by Lam (1990), is to allow for Markow-switching "plucks" in the transitory component of GDP, as shown by Kim and Nelson (1999). The ADF test has good power when these "plucks" occur under the alternative hypothesis. However, the ADF test can be oversized when the regime switching occurs under the null, mainly because the "plucks" increase the contribution of the transitory component to the series. This demonstrates that the true nature of business-cycle asymmetry has serious implications for the performance of unit-root tests on output series.

Table 12. Empirical Size of a 5% Perron (1994, 1997) Test: True Process Has Markov Switching in Variance

	T						
	200			500			
	Size				Size		
p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$	p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$
0.5	0.12	0.13	0.12	0.5	0.11	0.06	0.11
0.6	0.14	0.13	0.14	0.6	0.10	0.09	0.08
0.7	0.12	0.17	0.13	0.7	0.11	0.15	0.11
0.8	0.13	0.17	0.19	0.8	0.11	0.16	0.14
0.9	0.12	0.17	0.21	0.9	0.15	0.10	0.16
0.95	0.10	0.16	0.20	0.95	0.06	0.14	0.18
0.98	0.09	0.12	0.16	0.98	0.06	0.06	0.17

Table 13. Empirical Size of a 5% Zivot–Andrews (1992) Test: True Process Has Markov Switching in Variance

T							
200				500			
	Size				Size		
p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$	p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$
0.5	0.15	0.14	0.14	0.5	0.07	0.11	0.09
0.6	0.15	0.17	0.15	0.6	0.11	0.10	0.11
0.7	0.16	0.19	0.16	0.7	0.08	0.16	0.14
0.8	0.15	0.22	0.21	0.8	0.10	0.20	0.15
0.9	0.14	0.21	0.24	0.9	0.13	0.16	0.20
0.95	0.10	0.17	0.24	0.95	0.10	0.13	0.16
0.98	0.08	0.12	0.19	0.98	0.04	0.11	0.16

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Table 14. Empirical Size of a 5% Augmented Dickey–Fuller Test: True Process Has Markov Switching in the Transitory Component

T							
200				500			
	Size				Size		
p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$	p_{11}	$p_{00}=0.9$	$p_{00}=0.95$	$p_{00}=0.98$
0.5	0.36	0.28	0.16	0.5	0.22	0.19	0.12
0.6	0.40	0.26	0.19	0.6	0.29	0.24	0.15
0.7	0.42	0.35	0.23	0.7	0.37	0.29	0.19
0.8	0.49	0.34	0.24	0.8	0.47	0.37	0.24
0.9	0.46	0.34	0.24	0.9	0.51	0.46	0.28
0.95	0.38	0.27	0.15	0.95	0.48	0.44	0.28
0.98	0.24	0.18	0.09	0.98	0.34	0.27	0.19

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