

The Persistence in International Real Interest Rates

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Abstract

In this paper, we investigate the degree of persistence in quarterly postwar tax-adjusted *ex post* real interest rates for 13 industrialized countries using two recently developed econometric procedures. Our results show that international tax-adjusted real interest rates are typically very persistent, with the lower bound of the 95% confidence interval for the sum of the autoregressive coefficients very close to 0.90 for nearly every country. A highly persistent real interest rate has important theoretical implications.

JEL classifications: C22, E43

Key words: Real interest rate; Persistence; Grid bootstrap; Subsampling; Half-life

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Non-technical Summary

The degree of persistence in international real interest rates has potentially important implications for theoretical models in financial economics and macroeconomics. In this paper, we investigate the degree of persistence in quarterly postwar tax-adjusted *ex post* real interest rates for 13 industrialized countries.

Previous research, employing conventional unit root tests, provides limited information on the degree of persistence in real interest rates, as unit root tests concentrate solely on testing the null hypothesis that the sum of the autoregressive (AR) coefficients is unity in an AR representation of a series against the alternative hypothesis that the sum of the AR coefficients is less than unity. In contrast, a confidence interval for the sum of the AR coefficients provides us with a more informative statistical description of a variable's persistence. In this paper, we employ two recently developed econometric procedures, due to Hansen (1999) and Romano and Wolf (2001), in order to estimate 95% confidence intervals for the sum of the AR coefficients in AR representations of international real interest rates. Unlike conventional or bootstrapped confidence intervals, the Hansen (1999) and Romano and Wolf (2001) procedures generate confidence intervals for nearly integrated variables with correct first-order asymptotic coverage for the sum of the AR coefficients. These procedures also have good coverage in finite samples.

Our results indicate a high degree of persistence in quarterly postwar tax-adjusted real interest rates. The lower bound for the 95% confidence interval for the sum of the AR coefficients is often greater than 0.90, while the upper bound is almost always greater than unity, for the countries we consider. Even if one holds to a strong prior belief that real interest rates are stationary, the lower bounds for the 95% confidence intervals for the sum of the AR coefficients indicate that international real interest rates are highly persistent. Our results indicate that a very high degree of persistence in international real interest rates is a "stylized fact" that theoretical models need to address.

1. Introduction

In this paper, we investigate the degree of persistence in international real interest rates using a number of recently developed econometric procedures. The key role played by the real interest rate in prominent theoretical models in financial economics and macroeconomics motivates our interest in the degree of persistence in international real interest rates. These models include the consumption-based asset pricing model of, among others, Lucas (1978); the neoclassical growth model with explicitly optimizing agents, usually associated with Cass (1965) and Koopmans (1965); and various models of the monetary transmission mechanism. The degree of persistence in international real interest rates has potentially important implications for these canonical theoretical models.

Starting with Rose (1988), and in the spirit of Nelson and Plosser (1982), the empirical literature has focused on testing whether international real interest rates contain a unit root. Overall, the evidence is somewhat mixed (Rapach and Weber, 2003). However, conventional unit root tests alone provide limited information on the degree of persistence in real interest rates, as they concentrate solely on testing the null hypothesis that the sum of the autoregressive (AR) coefficients is unity in an AR representation of a series against the alternative hypothesis that the sum of the AR coefficients is less than unity. In contrast, a confidence interval for the sum of the AR coefficients provides us with a more informative statistical description of a variable's persistence. Our primary objective in the present paper is to provide such confidence intervals for international tax-adjusted real interest rates.

There is a well-known difficulty in constructing confidence intervals for the sum of the AR coefficients: conventional asymptotic or bootstrapped confidence intervals are not valid for this key measure of persistence when the data are generated by a nearly integrated process (Basawa et al., 1991). In order to construct confidence intervals for the sum of the AR coefficients that provide correct first-order asymptotic coverage, we employ the recently developed Hansen (1999) grid-bootstrap and Romano

and Wolf (2001) subsampling procedures.¹ Using Monte Carlo simulations, Hansen (1999) and Romano and Wolf (2001) find that their respective procedures for constructing asymptotically valid confidence intervals also provide good coverage in finite samples. In addition to the sum of the AR coefficients, we measure persistence through the half-life, or the number of years required for a shock to a variable to dissipate by one-half, as this is a popular measure of persistence.²

The rest of the paper is organized as follows. We outline the construction of confidence intervals for measures of persistence in Section 2. Section 3 reports measures of the sum of the AR coefficients and half-lives for quarterly postwar tax-adjusted real interest rates in 13 industrialized countries. Section 4 concludes.

2. Econometric Methodology

To fix ideas, consider the following AR(p) process for the variable y_t :

$$y_t = \mu + \alpha_1 y_{t-1} + \alpha_2 y_{t-2} + \dots + \alpha_p y_{t-p} + e_t, \quad (1)$$

for $t = 1, 2, \dots, T$. Andrews and Chen (1994) argue that an informative scalar measure of persistence in the AR process is the sum of the AR coefficients, $\alpha = \sum_{i=1}^p \alpha_i$, as the cumulative impulse response (CIR, the sum of the impulse response function over all time horizons) is related to α via $\text{CIR} = 1/(1-\alpha)$. Andrews and Chen (1994) view α as more informative than the largest root of the AR(p) model, since two AR(p) models with identical largest roots can have very different persistence properties.³

The variable y_t is stationary, or mean-reverting, if $|\alpha| < 1$, while y_t is a unit-root process if $\alpha = 1$. We concentrate on the situation in which $\alpha > 0$, as this is the relevant range for our real interest rate data. We can straightforwardly obtain a point estimate of α by re-arranging equation (1) and using

¹ See Andrews and Chen (1994) for another procedure for generating confidence intervals for the sum of the AR coefficients based on a median-unbiased estimator.

² The half-life is especially popular in the empirical purchasing power parity literature; see, for example, Lothian and Taylor (1996) and Papell (1997).

³ Andrews and Chen (1994) also point out that the spectrum at frequency zero is directly related to α , again making α an informative measure of persistence.

OLS to estimate the familiar augmented Dickey and Fuller (1981, ADF) and Said and Dickey (1984) regression model,

$$y_t = \mu + \alpha y_{t-1} + \sum_{j=1}^k \beta_j \Delta y_{t-j} + e_t, \quad (2)$$

where $\Delta y_t = y_t - y_{t-1}$. The construction of confidence intervals for α is problematic because the asymptotic distribution of the OLS estimator (as well as its rate of convergence) is different in the stationary and unit-root cases. Strictly speaking, if $\alpha < 1$, a confidence interval for α can be constructed through conventional asymptotic methods based on the standard normal distribution. However, the conventional asymptotic procedure fares quite poorly in finite samples, especially when α is near unity. In fact, if we formalize the near-unit-root case in a local-to-unity framework, where $\alpha = 1 + c/T$ with c held constant as $T \rightarrow \infty$, the conventional t -statistic used to construct asymptotic confidence intervals for α has a non-standard distribution. In this case, the conventional confidence interval is not valid asymptotically and will likely perform very poorly in finite samples. We cannot circumvent this difficulty by relying instead on a conventional bootstrap procedure, as the conventional bootstrap will also fail to generate confidence intervals with correct first-order asymptotic coverage (Basawa et al., 1991). The problem is that the asymptotic t -statistic depends on c , and thus α , and is therefore non-pivotal, while the conventional bootstrap procedure implicitly assumes that the t -statistic is pivotal. The confidence interval therefore does not properly control for a Type I error.⁴

Recently, Hansen (1999) and Romano and Wolf (2001) develop procedures for constructing confidence intervals for α with correct first-order asymptotic coverage. The Hansen (1999) grid-bootstrap procedure is an alternative to the conventional percentile- t bootstrap that provides correct first-order asymptotic coverage in a local-to-unity framework. The key difference between the conventional and grid-bootstrap procedures is that the grid-bootstrap computes empirical quantiles for the t -statistics for an entire grid of α values and not just the OLS estimate, $\hat{\alpha}$. More specifically, consider a grid of

⁴ See Section 1 in Hansen (1999) for a more detailed presentation of these issues.

values for α , α_i ($i=1,\dots,B$), covering $\hat{\alpha}$. In order to estimate the data-generating process for each α_i , we estimate equation (2), with α restricted to α_i , using restricted OLS for each α_i . The restricted OLS parameter estimates, together with re-sampled restricted OLS residuals, are used to build up a large number of pseudo-samples (say, 2000) for each α_i . For each of the 2000 pseudo-samples for each α_i , we calculate the t -statistic, $t_i^* = (\hat{\alpha}_i^* - \alpha_i) / s(\hat{\alpha}_i^*)$. We sort the t -statistics, giving us an empirical distribution of t -statistics for each α_i , from which we can calculate the 0.025 and 0.975 quantiles of t -statistics for each α_i . The upper bound for the 95% confidence interval for α is the α_i grid value such that $(\hat{\alpha} - \alpha_i) / s(\hat{\alpha}) = t_{i,0.025}^*$. The lower bound is the α_i grid value such that $(\hat{\alpha} - \alpha_i) / s(\hat{\alpha}) = t_{i,0.975}^*$.⁵ In Monte Carlo simulations, Hansen (1999) finds that the grid-bootstrap procedure generates confidence intervals with good coverage in finite samples.

Romano and Wolf (2001) develop a subsampling procedure for constructing confidence intervals for α that also provides correct first-order asymptotic coverage. This approach re-computes the OLS estimator on smaller blocks, or subsamples, of the observed series. More specifically, we begin with a block of size b and calculate the t -statistic, $\tau_b(\hat{\alpha}_{b,t} - \hat{\alpha}) / \hat{\sigma}_{b,t}$, for each subsample of size b for $t=1,\dots,T-b+1$, where $\hat{\alpha}_{b,t}$ is the OLS estimates of α for the t th block of size b , $\hat{\sigma}_{b,t} = b^{1/2}s(\hat{\alpha}_{b,t})$, and $\tau_b = b^{1/2}$. We generate the empirical approximating distribution for the subsample t -statistics,

$$L_b(x) = \frac{1}{T-b+1} \sum_{t=1}^{T-b+1} 1\{\tau_b(\hat{\alpha}_{b,t} - \hat{\alpha}) / \hat{\sigma}_{b,t} \leq x\}. \quad (3)$$

Let $c_{b,0.025}$ and $c_{b,0.975}$ be the 0.025 and 0.975 quantiles of the subsampling distribution, equation (3). A 95% two-sided equal-tailed confidence interval for α is given by

$$[\hat{\alpha} - (1/\tau_T)s(\hat{\alpha})c_{b,0.975}, \hat{\alpha} - (1/\tau_T)s(\hat{\alpha})c_{b,0.025}], \quad (4)$$

where $\tau_T = T^{1/2}$. Romano and Wolf (2001) also discuss the construction of a two-sided symmetric

⁵ Thus, the grid bootstrap is an application of the no-rejection principle.

subsampling interval. Instead of equation (3), we generate the empirical approximating distribution,

$$L_{b,|\cdot|}(x) = \frac{1}{T-b+1} \sum_{t=1}^{T-b+1} 1\{\tau_b |\hat{\alpha}_{b,t} - \hat{\alpha}| / \hat{\sigma}_{b,t} \leq x\}. \quad (5)$$

Let $c_{b,|\cdot|,0.05}$ be the 0.05 quantile for the empirical distribution, equation (5). A 95% two-sided symmetric confidence interval is given by

$$[\hat{\alpha} - (1/\tau_T)s(\hat{\alpha})c_{b,|\cdot|,0.05}, \hat{\alpha} + (1/\tau_T)s(\hat{\alpha})c_{b,|\cdot|,0.05}]. \quad (6)$$

We follow Algorithm 5.1 (Minimizing Confidence Interval Volatility) in Romano and Wolf (2001, p. 1297) in order to select b . The algorithm proceeds as follows:

1. Compute a subsampling 95% confidence interval for α for each b in $b = b_{small}$ to $b = b_{big}$, yielding the endpoints $I_{b,low}$ and $I_{b,up}$. Set $b_{small} = c_1 T^\eta$ and $b_{big} = c_2 T^\eta$ for $0 < c_1 < c_2$ and $0 < \eta < 1$. Romano and Wolf (2001) recommend $c_1 \in [0.5, 1]$, $c_2 \in [2, 3]$, and $\eta = 0.5$. (We set $c_1 = 1$, $c_2 = 3$, and $\eta = 0.5$.)
2. For each b , compute a volatility index, VI_b , where the volatility index is the standard deviation of the interval endpoints in a neighborhood of b . That is, for a small integer k , let VI_b equal the standard deviation of $\{I_{b-k,low}, \dots, I_{b+k,low}\}$ plus the standard deviation of $\{I_{b-k,up}, \dots, I_{b+k,up}\}$. Romano and Wolf recommend $k = 2$ or $k = 3$. (We set $k = 2$.)

Select the value for b , b^* , with the smallest volatility index and report $[I_{b^*,low}, I_{b^*,up}]$ as the final 95% subsampling confidence interval.

In Monte Carlo simulations, Romano and Wolf (2001) find that their subsampling confidence intervals have good coverage in finite samples, with the two-sided symmetric confidence interval performing somewhat better than the equal-tailed two-sided confidence interval. A potential advantage of the Romano and Wolf (2001) subsampling procedure is that, unlike the Hansen (1999) grid-bootstrap procedure, it does not require the assumption that e_t is independently and identically distributed (*iid*) in

equations (1) or (2), as it is still valid for dependent error processes.⁶

We also consider the half-life as a measure of persistence. The half-life is calculated from the impulse response function and is defined as the number of periods required for a unit shock to dissipate by 0.5: $\sup_{l \in L} \left| \frac{\partial y_{t+l}}{\partial e_t} \right| \geq 0.5$. The half-life is also commonly computed using $\ln(0.5)/\ln(\alpha)$. However, this measure requires the adjustment to be monotonic, which may not be the true for the data. We calculate the half-life through the impulse response function for an $AR(p)$ in levels, as in Cheung and Lai (2000). Interestingly, Inoue and Kilian (2000) show that even in a local-to-unity framework, as long as $p > 1$, conventional bootstrap confidence intervals are valid asymptotically for the individual slope coefficients of the $AR(p)$ model, equation (1), and some functions of these coefficients—including the impulse response function—while they are not valid for the sum of the AR coefficients, α . However, conventional bootstrap confidence intervals for the half-life when $p > 1$ can have very poor coverage in finite samples for persistent series (Gospodinov, 2004). In light of this, we compute percentile grid-bootstrap 95% confidence intervals for the half-life of the impulse response function using the procedure outlined in Gospodinov (2004).

3. Empirical Results

Our quarterly nominal interest rate and inflation rate data are from the June 2000 International Monetary Fund *International Financial Statistics* (IFS) CD-ROM. We measure the nominal interest rate using the long-term government bond yield (IFS series 61..ZF..). One reason for focusing on the long-term government bond yield is that data beginning in the 1960s are available for a much larger number of countries than short-term government bond yield data on the IFS CD-ROM. Another reason for focusing on the long-term bond yield is that it is widely believed that long-term rates are the most relevant for

⁶ The Andrews and Chen (1994) procedure for constructing confidence intervals for α also requires the *iid* assumption for e_t .

saving and investment decisions. We also conducted the analysis (results not reported) for the few countries for which short-term government bond yields are available, and the results are qualitatively the same as those for the long-term government bond yields.⁷ The 13 countries for which we have long-term government bond yield data are Australia, Belgium, Canada, Denmark, France, Ireland, Italy, the Netherlands, New Zealand, Norway, Switzerland, the United Kingdom, and the United States. Our annualized inflation data for the same set of countries are based on consumer price indexes (IFS series 64..ZF..). Our sample begins in 1960:4 and extends to 1998:3. Marginal tax rates on a decadal basis are computed as the sum of the values in columns (4) and (6) of Appendix Table A1 in Paddovano and Galli (2001), and these provide us with marginal tax rates with a degree of temporal variation. The tax-adjusted real interest rate is defined as $y_t = (1 - \tau)i_t - \pi_t$, where τ is the marginal tax rate for the relevant decade, i_t is the nominal interest rate (nominal long-term government bond yield), and π_t is the annualized inflation rate, all at time t .⁸ Existing time-series studies typically do not adjust real interest rates for taxes. Our tax-adjusted real interest rates should provide a more accurate measure of the real interest rate relevant for economic decisions.⁹

Column (3) of Table 1 reports the OLS estimate of α in equation (2) for each country, while column (2) reports the lag length k in equation (2), selected using the modified AIC criterion of Ng and Perron (2001). The OLS point estimates of α are greater than or equal to 0.90 for every country, with the exception of Denmark. Of course, these point estimates are biased downward and are of limited value. In order to provide more informative measures of persistence, we calculate Hansen (1999) grid-bootstrap and Romano and Wolf (2001) subsampling 95% confidence intervals for α for every country. As discussed above, these confidence intervals provide valid asymptotic first-order coverage and appear to

⁷ These results are available on request from the authors.

⁸ We follow the convention in the literature and match up the nominal interest rate for a given quarter with the inflation rate over the following quarter. We are also following much of the extant literature by using the *ex post*, and not the *ex ante*, real interest rate. This allows us to sidestep the thorny issue of specifying explicitly how inflationary expectations are formed, and given that we are interested in measuring persistence—a long-run notion—this should not be crucial. If expectations are rational, actual and expected inflation will only differ by a white-noise error term.

⁹ However, results (not reported) are similar if we use tax-unadjusted real interest rates.

have good coverage in finite samples.

We report the Hansen (1999) grid-bootstrap 95% confidence interval for α in column (4) of Table 1.¹⁰ Observe that the lower bound of the grid-bootstrap confidence interval is greater than or equal to 0.90 for every country, with the exceptions of Denmark, the Netherlands, Switzerland, and the U.K. For Switzerland and the U.K., the lower bound is still quite close to 0.90. The upper bound of the 95% confidence interval is greater than unity for every country, so the data are not inconsistent with a unit root in the tax-adjusted real interest rate for every country. According to the Hansen (1999) grid-bootstrap 95% confidence interval for α , even if one holds to a strong prior belief that real interest rates are stationary (so that $\alpha < 1$), the lower bounds for the 95% confidence intervals still indicate that real interest rates display a high degree of persistence (with the possible exceptions of Denmark and the Netherlands).

Romano and Wolf (2001) equal-tailed and symmetric subsampling 95% confidence intervals for α are found in columns (5) and (6) of Table 1. For the most part, these confidence intervals are similar to the grid-bootstrap confidence intervals. For the equal-tailed intervals reported in column (5), the lower bounds are relatively close to or greater than 0.90, with the exception of Denmark, and the upper bound is greater than unity for every country with the exceptions of the Netherlands and Switzerland (and the upper bound is still very close to unity for these countries). The symmetric intervals, reported in column (6), are similar to the equal-tailed subsampling intervals, although both the upper and lower bounds for the symmetric intervals appear to be somewhat smaller on average than those for the equal-tailed intervals. Nevertheless, with the exception of Denmark, the lower bounds for the symmetric intervals still indicate a high degree of persistence in real interest rates.

An alternative measure of persistence, the half-life based on the impulse response function, is reported in column (7) of Table 1, and the grid-bootstrap 95% confidence interval for the half-life is reported in column (8). As discussed above, we use the Gospodinov (2004) percentile grid-bootstrap

¹⁰ We generated the Hansen (1999) grid-bootstrap confidence intervals using the GAUSS program available from Bruce E. Hansen's home page at <http://www.ssc.wisc.edu/~bhansen/>.

procedure to calculate the 95% confidence interval for the half-life reported in column (8).¹¹ Note that the half-lives are measured in years in Table 1. The confidence intervals are very wide for all of the half-lives, and an infinite upper bound (corresponding to a nonstationary real interest rate) is found for all countries. Again, the data are not inconsistent with a high degree of persistence.

Overall, our results show that international tax-adjusted real interest rates are typically characterized by a high degree of persistence, even when we look at the lower bound of the 95% confidence intervals for α . As an illustration of the theoretical importance of our results, we reconsider the arguments of Rose (1988). Rose (1988) points out that a unit root in the real interest rate poses a serious problem for the canonical consumption-based asset-pricing model. The Euler equation corresponding to the maximization problem for the consumption-based asset-pricing model entails a tight link between the growth rate of consumption and the real interest rate. As Rose (1988) emphasizes, if the real interest rate is nonstationary, while consumption growth is stationary (as it almost surely is), the Euler equation cannot hold over time. Our results generalize Rose's point: it is not just a matter of whether the real interest rate is nonstationary, since even if the real rate is stationary, a high degree of persistence in the real rate implies substantial violations of the Euler condition for considerable periods of time if consumption growth is not as highly persistent. Consider the case of the U.S. Looking back at Table 1, we see that the lower bound for the 95% confidence interval for α is between 0.89 and 0.94 for the U.S. We also calculated a grid-bootstrap 95% confidence interval for α for quarterly postwar U.S. real consumption growth, with consumption defined as the sum of consumer purchases of nondurable goods and services.¹² We obtained a 95% confidence interval for α of [0.10, 0.65]. The upper bound of the 95% confidence interval for α for U.S. consumption growth is well below the lower bound of the 95% confidence interval for α for the U.S. real interest rate. These significant differences in the degree of persistence for the U.S. real interest rate and consumption growth imply sustained violations of the

¹¹ We generated the grid-bootstrap confidence intervals for the half-lives reported in column (8) of Table 1 using the GAUSS program available from Nikolay Gospodinov's home page at <http://alcor.concordia.ca/~gospodin/>.

¹² The data were downloaded from the Federal Reserve Economic Database (FRED), available from the Federal Reserve Bank of St. Louis web page at <http://www.stls.frb.org>.

Euler condition at the center of the consumption-based asset-pricing model.

4. Conclusion

In this paper, we investigate the degree of persistence in international tax-adjusted real interest rates using the recently developed Hansen (1999) grid-bootstrap and Romano and Wolf (2001) subsampling procedures. These procedures generate confidence intervals for the sum of the AR coefficients with correct first-order asymptotic coverage that also have good coverage in finite samples. Our results indicate a high degree of persistence in quarterly postwar tax-adjusted real interest rates. The lower bound for the 95% confidence interval for the sum of the AR coefficients is often greater than 0.90, while the upper bound is almost always greater than unity, for the countries we consider. Grid-bootstrap 95% confidence intervals for the half-lives have lower bounds ranging from approximately two to four years, while the upper bound is infinite for every country. Even if one holds to a strong prior belief that real interest rates are stationary, the lower bounds for the 95% confidence intervals for the sum of the AR coefficients and the half-lives indicate that international real interest rates are still highly persistent. Our results indicate that a very high degree of persistence in international real interest rates is a “stylized fact” that theoretical models need to address.

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Table 1: Point estimates and 95% confidence intervals for measures of real interest rate persistence

(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Country	k_{MAIC}^a	$\hat{\alpha}_{OLS}^b$	Grid bootstrap 95% CI	Subsampling equal-tailed 95% CI	Subsampling symmetric 95% CI	HL_{IRF}^c	Grid bootstrap 95% CI
Australia	4	0.96	[0.93, 1.02]	[0.89, 1.01]	[0.90, 1.02]	4.29	[2.08, ∞]
Belgium	2	0.96	[0.93, 1.02]	[0.91, 1.02]	[0.90, 1.02]	4.38	[2.49, ∞]
Canada	4	0.95	[0.91, 1.03]	[0.92, 1.03]	[0.89, 1.01]	3.18	[1.56, ∞]
Denmark	8	0.86	[0.79, 1.04]	[0.80, 1.07]	[0.75, 0.98]	0.92	[0.88, ∞]
France	5	0.98	[0.96, 1.03]	[0.95, 1.04]	[0.93, 1.03]	8.73	[4.43, ∞]
Ireland	4	0.96	[0.92, 1.04]	[0.93, 1.06]	[0.87, 1.04]	3.11	[1.42, ∞]
Italy	4	0.96	[0.93, 1.02]	[0.92, 1.03]	[0.90, 1.02]	4.15	[1.60, ∞]
Netherlands	4	0.90	[0.84, 1.03]	[0.88, 0.98]	[0.83, 0.98]	0.99	[0.93, ∞]
New Zealand	8	0.97	[0.94, 1.04]	[0.95, 1.05]	[0.91, 1.03]	4.57	[1.57, ∞]
Norway	8	0.96	[0.92, 1.04]	[0.95, 1.03]	[0.89, 1.02]	3.86	[1.28, ∞]
Switzerland	4	0.93	[0.88, 1.02]	[0.90, 0.98]	[0.88, 0.98]	1.52	[1.19, ∞]
United Kingdom	8	0.93	[0.89, 1.03]	[0.88, 1.04]	[0.86, 1.01]	2.36	[1.51, ∞]
United States	7	0.95	[0.92, 1.02]	[0.94, 1.02]	[0.89, 1.01]	3.04	[2.45, ∞]

Notes: ^aLag length k in equation (2) selected using the Ng and Perron (2001) modified AIC.

^bOLS estimate of for the sum of the AR coefficients in equation (2).

^cEstimate of the half-life (measured in years) based on the impulse response function.