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The persistence of shocks in GDP and the estimation of the potential economic costs of climate change

Richard S. J. Tol
Department of Economics, University of Sussex, Brighton, United Kingdom
Institute for Environmental Studies, Vrije Universiteit, Amsterdam, Netherlands;
Department of Spatial Economics, Vrije Universiteit, Amsterdam, Netherlands
Email: R.Tol@sussex.ac.uk

Francisco Estrada
Centro de Ciencias de la Atmósfera, Universidad Nacional Autónoma de México, Ciudad Universitaria, Circuito Exterior, 04510 Mexico, DF, Mexico; Institute for Environmental Studies, Vrije Universiteit, Amsterdam, Netherlands

Carlos Gay-García
Centro de Ciencias de la Atmósfera, Universidad Nacional Autónoma de México, México

Abstract: Integrated assessment models (IAMs) typically ignore the impact of climate change on economic growth, or simply scale down output and hence the entire future growth. In this manner, IAMs typically assume that the shocks caused by climate change impacts dissipate and have no persistence at all, affecting only the period when they occur. Clearly, this could lead to the underestimation of costs of climate change at global, regional, national and local scales. We adopt an empirical approach for analyzing the observed GDP series for different world regions in order to estimate the persistence of shocks on growth. We interpret the direct impact of climate change as such shocks, and use the estimated models to assess the implications for growth. We compare this to the scaling method pioneered by Nordhaus (Nordhaus and Boyer, 2000). A simple version of the widely used PAGE2002 model (Hope, 2006) is applied to conduct a sensitivity analysis varying the degree of a persistence measure in simulated future GDP. It is shown that when a persistence similar to the observed one is chosen, the economic impacts of climate change are considerably larger in comparison to the "zero persistence" implied by the original scaling method. If the persistence of shocks is ignored, as it is currently done by most IAMs, the economic impacts of climate change can be severely underestimated. Results are not sensitive to the selection of the discount rate. Moreover, it is shown that the original scaling method embedded in most IAMs can be interpreted as assuming an autonomous, costless, extremely large and effective (reactive) adaptation capacity.

JEL Classification: Q54

Key Words: Climate change; economic impacts; dynamic model

I. Introduction

Whether output is DS or TS, there is a consensus in that this and other economic variables show a significant degree of persistence and that shocks tend to affect not only current period but for a number of periods in the future. Roots close to unity in the characteristic equation and high persistence are stylized facts of GDP and other economic time series that are supported by an overwhelming amount of empirical evidence as documented in the specialized literature. Nevertheless, the scaling method pioneered by Nordhaus (Nordhaus and Boyer, 2000) and thereafter used in most of the current IAMs implicitly assume that climate change impacts on the economy have no persistence at all; a fact that is clearly at odds with empirical evidence mentioned above. In this paper we present further empirical evidence on the subject and apply the results to the analysis of the potential economic costs of climate change. Tol and Fankhauser (2005) addressed this problem from a theoretical perspective analyzing the dynamic effects of climate change impacts in future welfare by means of economic growth models. They showed that in addition to the direct impacts of climate change, this phenomenon can have important indirect impacts over capital accumulation, the propensity to save and capital-labor ratio due to climate change's potential health effects. Hallegatte (2005, 2007) stresses the importance of considering the climate and economic dynamics (and feedback processes between these two systems) as well as the short-term socioeconomic constraints in determining the long-term costs of climate change. He argues that the impacts associated to these dynamic processes can be larger than those shown in the traditional assessments of the costs of climate change that have been published. The existence of poverty traps has been also pointed out as a potential mechanism that could have persistent effects over economic growth through its impact on demographic and economic dynamics (Tol, 2011; Hallegatte, 2007).

Here we bridge the discussion regarding the importance of growth dynamics on the assessment of climate change impacts with the empirical time-series analysis literature that has analyzed the dynamics of shocks in macroeconomic series, and with the IAM modeling framework. In this paper we propose a simple manner for incorporating persistence into the IAM analysis of the potential economic impacts of climate change. Our results provide considerable empirical evidence regarding the fundamental role of the memory properties of GDP when assessing the potential economic impacts of climate change. In the first part of this paper, the time-series properties of twelve regional GDP time series are investigated by means of a standard unit root test (Dickey and Fuller, 1979; Said and Dickey, 1984) and a unit root test that allows for a one-time structural change in the trend function (Perron 1989; Perron 1997; Kim and Perron, 2009). Results show that for most of these time series the null hypothesis of a unit root cannot be rejected, indicating that the effects of a shock do not fade even when long horizons are considered. Even for those series where evidence for trend stationarity could be found, it is shown that the vast majority of them are characterized by roots in the characteristic equation close to unity and large values of the sum of the autoregressive coefficients, indicating that shocks are highly persistent. These empirical estimates of the persistence are contrasted to that implied by the scaling method, and to the persistence imparted by the economic growth models in DICE/RICE.

In the second part of this paper, we use a simple version of the widely used PAGE2002 model (Hope, 2006) to conduct a sensitivity analysis varying the degree of persistence in simulated future GDP. It is shown that when a persistence similar to the observed one is chosen, the economic impacts of climate change are considerably larger in comparison to the "zero persistence" implied by the original scaling method that is common to most of the available IAMs and that therefore the costs presented in the literature could be seriously underestimated.

II. How persistent are shocks in GDP?

The dynamic properties of macroeconomic time series have been the subject of a long debate (Nelson and Plosser, 1982; Kim and Perron, 2009). The persistence of shocks has been the chief point of discussion. Can innovations change the long-run path of the economy or do shocks dissipate without altering the growth path? How long do the effects of a shock last? Prior to Nelson and Plosser (1982), analysts separated the business cycle – short-term variations in output – from the secular trend, which is determined by real factors such as technological change, capital accumulation and population growth. In contrast, the business cycle is assumed to be transitory and to be basically produced by monetary factors – stationary fluctuations around a secular trend. That is, economic time series are described as trend stationary (TS) processes. Nelson and Plosser (1982) showed that most macroeconomic time series in the USA could be better represented by difference stationary (DS) processes. A DS process has a stochastic long-term movement that is the sum of stationary innovations. Under this view, current shocks have an infinite persistence and therefore a permanent effect on the long-run path of macroeconomic aggregates.

Campbell and Mankiw (1989; 1987a) address this issue by answering the question: after the occurrence of an unexpected 1% change in output today, should we revise our forecast by more than 1% over a long horizon? If fluctuations in output are dominated by temporary deviations from the natural rate of output, i.e. as in a TS process, the occurrence of a shock should not substantially change our forecast over a five or ten years horizon. The authors find that the forecast would indeed be substantially different: shocks to output are largely permanent and that substantial persistence of output shocks is an important and often neglected feature of post-war economic time series.

Perron (1989) presents a new class of unit root tests that allow for the existence of a one-time structural change in the trend function. Applying these tests to the Nelson and Plosser data set, Perron showed that if only two shocks are considered to have had permanent effects in the various macroeconomic variables (i.e. the 1929 Great Crash and the oil price shock of 1973), then most of these time series are better described as TS processes with a structural break. That is, assuming a unit root would have erroneously attributed too much persistence (infinite persistence) to macroeconomic variables as a result of ignoring the presence of a structural change.

For investigating the persistence of shocks in GDP we selected 12 groups of countries (regions) for the post-war period (1950-2008). All data was taken from the Maddison database (http://www.ggdc.net/maddison/). The annual GDP time series considered are: global, Africa, Asia, Eastern Asia, Western Asia, Eastern Europe, Western Europe, the 12 countries with the largest economies in Western Europe, Western Offshoots (Australia, New Zealand, Canada and the USA), Latin America, the 8 Latin American countries with the largest economies and the countries from the former USSR. Figure A1 shows a plot of the natural log of these time series and, as can be seen, all of them show a clear non-stationary behavior with the potential presence of large structural changes in the slope or on the slope and level of the trend function.

The two most common non-stationary processes are trend stationary and difference stationary (see Appendix). The memory of these processes is very different and distinguishing among them is important for estimating how much persistence could be expected from a shock. In order to investigate the time-series properties of GDP (Figure A1), we first applied the standard Augmented Dickey-Fuller[1] (ADF; Dickey and Fuller, 1979; Said and Dickey, 1984) which consist in estimating the regression:

$$(1) \qquad \Delta y_t = \hat{\mu} + \hat{\beta} t + (\hat{\rho} - 1) y_{t-1} + \sum_{k=1}^{K} \hat{\delta}_k \Delta y_{t-k} + \varepsilon_t$$

and testing the null hypothesis of a unit root ($\rho = 1$) against the alternative hypothesis of trend stationary process ($\rho < 1$). The coefficient ρ equals the sum of the autoregressive coefficients (SAR), one of the most commonly used measures of persistence in macroeconomic time series (Oka and Perron, 2011). Table A1, the null of a unit root cannot be rejected for any of these series at the 10% significance level. The ρ estimates range from 0.97 to 0.87 with a mean value of 0.94.
We selected the number of lagged differences *K* by means of the Schwarz Information Criteria (BIC). Africa is the only series for which the AIC leads to a different conclusion: the rejection of the null at the 5% significance level. Even in this case, the SAR is quite large (0.831). That is, six years after a shock only about 60% of the effect would have dissipated. The cumulative impulse response (CIR) of a shock of magnitude 1 would be $1/(1-\rho)$=5.92.

The CIR values in Table A1 illustrate the high persistence of shocks in GDP, ranging from 7.63 to 32.26 with a mean value of 22.73, implying that the simple scaling method, which ignores these dynamic effects, could provide a very poor estimation of the potential economic impacts of climate change, or of any other type of shock, for that matter.

The sum of the first order autoregressive coefficients is biased towards unity if there is a shift in the trend function (Perron, 1989). In this case, the unit root null is hardly rejected even if the series is composed of a trend (with a break) and *i.i.d.* disturbances. As can be seen in Figure A1, the presence of structural changes in the slope or in both the slope and the intercept of the series is very likely in many of the GDP series. There have been major global and regional macroeconomic shocks during the second half of the 20th century, that may have cause trend breaks.

It is therefore important to test if the results shown in Table A1 are affected by the presence of structural changes. Most of the available tests for structural changes presuppose that the order of integration of the time series is known, while in fact unit root tests are highly sensitive to the presence of structural changes. We applied the Perron and Yabu (2009) procedure for structural change that is valid whether the time series is I(0) or I(1). We only consider models that allow for a break in the slope and in

[1] Because of the obvious trending behavior of the series plotted in Figure 1 we only consider the ADF specification that includes a constant and a linear trend.

the slope and level of the series, respectively. Table A2 shows that the null of no-break in the trend function can be strongly rejected (at 1% or 5% levels) for all series, with the exception of Eastern Asia for which it can only be rejected at the 10% significance level.

As the standard ADF test could erroneously indicate the presence of a unit root, we use the Kim and Perron (2009) unit root test. Table A3 shows that, once the occurrence of a break in the trend function is allowed, the unit root null can be rejected at 5% levels for Latin America and the eight largest Latin American economies and at 10% levels for Asia and the twelve largest economies in Western Europe. The estimated break dates are similar, reflecting the effects of major shocks: the oil price shock of 1973, the 1980s debt crisis, and the disintegration of the USSR in the 1990s. The αs are considerably lower than those shown in Table A1, now ranging from 0.46 to 0.90 with a mean of 0.70, indicating that once structural breaks are accounted for, the persistence of GDP is considerably lower. To exemplify the difference in the persistence of a shock that could be expected from the estimates in Tables A1 and A3, consider the case of the GDP of the twelve largest economies in Western Europe. Without trend break, a unit shock would produce a CIR of 25.57[2]; with trend break, the accumulated response would be 1.85. Note however that both of these long-run accumulated responses are considerably larger that those that would be produced under the "no memory" assumed by most IAMs. The CIR values range from 1.85 to 9.80, with a mean of 4.43, suggesting that certainly the lack of persistence in the scaling method fails to represent a feature of observed dynamics of GDP shocks that could be quite relevant when assessing the potential costs of climate change.

Figure 2 shows the impulse response functions and the accumulated response of a unit shock in the estimated AR(p) models. Panel a ignores structural breaks, and panel b does not. As evident from these figures, the dynamics of a shock and the time it takes for its effects dissipate are very different: shocks dissipate more rapidly in panel b and the cumulative response is much lower. Panel c shows that these phenomena can be approximated with AR(1) models.

In any case, the "zero persistence" of the scaling method used in IAMs is not supported by the observed data. The next sections of this paper illustrate the relevance of including the memory of GDP series when estimating the potential impacts of climate change in a (dynamic) IAM context.

III. Persistence in Integrated Assessment Models

Above, we discuss persistence in GDP data. Here, we turn to persistence in three integrated assessment models: PAGE, FUND, and DICE. These three models are most commonly used to assess the economic impacts of climate change.

III.1 PAGE

[2] That is assuming that the true value of the first order autoregressive coefficient is 0.965 and not 1, although the unit root test in Table 1 indicates that this estimate is not different form the unity. If the coefficient true value is 1, the long-run response would be equal to infinity.

The PAGE model has zero persistence: A unit shock at time *t* has no impact on the level or growth rate of GDP at time *t+s*.

III.2 FUND

In FUND, a shock never dissipates. In an absolute sense, the gap between the growth path with and without shock grows with the growth rate of the economy. In a relative sense, the economy with shock is always X% smaller than the economy without shock.

III.3. DICE

DICE is more complicated than PAGE and FUND. The damage function of DICE is as follows:

(2) $$\Omega_t = 1/\left[1 + D_t\right]$$

(3) $$D_t = \theta_1 T_t + \theta_2 T_t^2$$

where D_t represents the climate damage as fraction of output, θ_1 and θ_2 are the parameters of the damage function, T is global temperature increase over the 1900 level and Ω_t is the scaling factor for output.

The model dynamics impart a certain level of persistence to shocks in GDP. Output Q_t is determined by a Solow growth model[3] specified with a Cobb-Douglas production function of the form:

(4) $$Q_t = A_t K_t^{\gamma} L_t^{1-\gamma}$$

where A_t is the total factor productivity, L_t is population, γ is the elasticity of output with respect to capital and K_t is the capital stock which is determined by the equation for capital accumulation:

(5) $$K_t = \left(1-\delta\right) K_{t-1} + I_{t-1} = \left(1-\delta\right) K_{t-1} + \sigma Y_{t-1}$$

where δ is the depreciation rate, *I* is investment and σ is the savings' rate.

Population L_t increases less than exponentially with a rate of growth g_t^{pop} that declines geometrically over time, leading to a stable population in the long-run. That is, population is:

(6) $$L_t = L_0 \exp\left(\int_0^t g_t^{pop}\right)$$

(7) $$g_t^{pop} = g_0^{pop} \exp\left(-\kappa_t^{pop}\right)$$

Total factor productivity is modeled in a similar way to population, using an exponential growth function with a geometrically declining growth rate:

(8) $$A_t = A_0 \exp\left(\int_0^t g_t^{A}\right)$$

(9) $$g_t^{A} = g_0^{A} \exp\left(-\kappa_t^{A}\right)$$

[3] There are two versions of DICE, Ramsey and Solow. In the Ramsey model, the impact of climate change barely affects the savings' rate, so that the model, for all practical purposes, is *de facto* Solow.

There are two types of dynamics in a growth model: equilibrium dynamics and disequilibrium dynamics. To start with the latter, the speed of convergence to steady state for the Solow model with a Cobb-Douglas production function is $\beta_t = (1-\gamma)(g_t^{pop} + g_t^A + \delta)$. The convergence coefficient β relates to persistence as $\beta_t = (1-\alpha_t^*)$, where α_t^* measures persistence of shocks (see, for example, Lee et al., 1997). The persistence coefficient α^* depends on the elasticity of output γ, the total factor productivity growth rate g_t^A, the population growth rate g_t^{pop} and the depreciation rate δ of capital. Persistence also depends on the length of the time step that is chosen, since the growth rates tend to be larger as the time step becomes larger. For example, using the annual growth rates (instead of decadal) from the DICE base case scenario, the beta convergence coefficient is $\beta_t = (1-0.3)(0.0157 + 0.0038 + 0.1) = 0.08$, indicating that every year 8% of the shock would dissipate. The persistence coefficient is $\alpha^* = 0.916$, quite similar to those shown in Tables 1 and 3. As the length of the time step increases, persistence decreases. In DICE, the time step is 10 years and the persistence is $\alpha^* = 0.04$. The impacts show no dynamics. As such, the combination of the damage functions in DICE and the specification of the economic growth model impose a zero persistence of climate change impacts *within* the 10-year time step and a (reasonable) almost zero persistence *between* time steps. Note that the damage functions are not calibrated or adjusted in anyhow to approximate the effects of persistence within the 10-yer time steps. As a consequence, results would be different if climate change impacts were estimated yearly, allowing for the persistence effects, and then aggregated in 10 year periods than if the damage function is directly applied using a 10-year step.

The equilibrium dynamics are as follows. Labor and total factor productivity are unaffected by changes in output. Capital accumulation is, however. Investment falls by a factor *Ω*, just like output. The equilibrium relationship between capital and output is *K=σY/δ*. The scaling factor for capital is

$$\Theta_t = \frac{\sigma\Omega_{t-1}Y_{t-1} + (1-\delta)K_{t-1}}{\sigma Y_{t-1} + (1-\delta)K_{t-1}} = \frac{\sigma\Omega_{t-1}Y_{t-1} + (1-\delta)\frac{\sigma}{\delta}Y_{t-1}}{\sigma Y_{t-1} + (1-\delta)\frac{\sigma}{\delta}Y_{t-1}} = \frac{\Omega_{t-1} + \frac{(1-\delta)}{\delta}}{1 + \frac{(1-\delta)}{\delta}} = 1 - \delta(1-\Omega_{t-1})$$

(10)

That is, if the impact of climate change is 1% (10%), then the scaling factor for output is 0.99 (0.90). If the depreciation rate is 10%, then impact on capital in the next period is 0.1% (1%). The scaling factor for capital is thus 0.999 (0.990). The scaling factor for output in the next period is the scaling factor for capital raised to its elasticity: 0.999^{γ} (0.990^{γ}). Persistence is therefore

(11) $$\alpha = \frac{1-\Theta_t^{\gamma}}{1-\Omega_{t-1}} = \frac{1-[1-\delta(1-\Omega_{t-1})]^{\gamma}}{1-\Omega_{t-1}} \approx \delta\gamma$$

Therefore, whether we consider the equilibrium or the disequilibrium dynamics of DICE, its persistence is low, much lower than the empirical evidence in Section III.

The differences in persistence, or in the speed of convergence, can also be interpreted as follows. On the one hand, the lack of persistence (or a 100% speed of convergence) in DICE and PAGE2002 can be interpreted as assuming that human and natural systems have an autonomous, costless, extremely large and effective reactive adaptation capacity and resilience[4]. The economy, the society in general and nature are assumed to have the capacity to overcome the effects of a shock in a single period of time without affecting their output/state in any future periods, no matter how large the impact may be and without having to invest anything on it. In contrast, α=1 (or conversely, β=0) would describe a system with a very limited resilience, being never capable of fully recover from a shock. Values of α close or equal to unity, as in FUND, are broadly in agreement with the observed evidence shown above and in the large body of empirical evidence reported in the econometrics literature.

IV. Sensitivity analysis of the estimates of the economic impacts of climate change to different assumptions regarding the persistence of shocks to GDP.

In this section, the sensitivity of the estimates of the costs of climate change to different assumptions regarding the memory of GDP is investigated. We use a simple impact functions, following PAGE2002, for a hypothetical region with an annual GDP growth of 2.5%, and we represent the persistence of shocks in GDP by means of a simple one-period memory equation with persistence *α*. The impact function is as follows:

$$(12) \qquad I_t = \vartheta_1 \left(\frac{\Delta T_t}{2.5} \right)^{\vartheta_2} + \alpha I_{t-1}$$

where *I* represents the economic impacts at time *t*, ΔT is the increment in temperature with respect to its preindustrial value and ϑ_1 and ϑ_2 are parameters. The parameters are represented by triangular probability distributions parameterized for the European Union as shown in Hope (2006), and the increase in temperature at the end of the century is represented by an uniform distribution covering a range from 2.5ºC to 4ºC. Temperature is assumed to increase linearly. All estimates presented were produced by means of simulation experiments of 5,000 realizations and the time-step was chosen to be one year.

Figure 4 shows the mean (dotted) and 5th and 95th percentiles (dashed) for different values of the parameter *α*. The red lines show the estimates produced without considering the persistence of shocks (*α*=0), while the blue ones present those assuming positive values of α. As can be seen, for very small values of α (for example 0.1) the differences in the losses in GDP may be negligible but for those that are closer to what the observed memory of GDP is (say, from 0.5 to 0.9), the differences become very large. For example, the mean economic impact estimated when α=0.5 is close to the 95th percentile for α=0, and for values of 0.8 and 0.9 this estimate is equal and even larger that the 95th percentile. The 90% confidence interval also grows with α.

Table 4 shows the net present value of the impacts of climate change over 100 years, using a 5% discount rate. The net present value is expressed as a percentage of GDP in

[4] Resilience here is understood as a system's capacity and speed to recover after the occurrence of a perturbation (see, for example Adger, 2000).

year 1. If α=0, the present value of the total impacts would range from 5.7% to 25.2% of initial GDP, with a mean value of 14.4%. If *α*=0.85, the mean value increases to 26% of GDP, with a range of 11% to 46% of GDP.

V. Estimates of the economic impacts of climate change under the A2 scenario and for different values of α

In this section we present estimates of the economic impacts of climate change for the world[5] and for 7 world regions using the impact functions of the PAGE2002 model modified to incorporate the persistence of shocks assuming different values of *α*. The regions used are similar to those in the PAGE2002 model; see Table A5.

The increase in global temperature at the end of the century is represented by an uniform distribution covering a range from 2ºC to 5.4ºC, which is the likely range of increase under the A2 emissions scenario (WGI-IPCC, 2007). The regional weights for scaling the impact functions that were used are those from the PAGE2002 model (see Table 5 in Hope, 2006). Regional estimates of temperature increase where produced using the scaling factors obtained from the emulation of the UKMOHADCM3 General Circulation Model of MAGICC/ScenGen[6].

Table 6 presents the net present values for varying values of α. Economic growth is as in the A2 scenario (Grübler et al., 2007; Nakicenovic et al., 2000). World GDP grows by almost 2% per year. The discount rate is 4%. Results are highly sensitive to the memory parameter α. The rightmost two columns show the estimates of the economic impacts of climate change for the world and for the seven world regions above, choosing values of the α parameter for each region that are similar to those in Table 3. Impacts are severely underestimated for *α*=0.

VI. Conclusions.

A stylized fact of macroeconomic and financial time series is the large memory that they posses and the persistence of shocks in these variables. This has been particularly studied at length for GDP series as shown by the economics and econometrics literature. Nevertheless, as shown in this paper, this is a very relevant characteristic that has been ignored or seriously underestimated in most of the IAMs that are available. From the theoretical perspective, this problem was previously addressed by Fankhauser and Tol (2005) using economic growth models. In the present paper, empirical evidence is presented regarding the importance of persistence when estimating of the costs of climate change, and new estimates of the costs of climate change based on a modified version of PAGE2002 are presented.

In this paper, this problem is investigated first by means of a standard unit root test and two unit root tests that allow for a one-time break in the trend function, in order to provide a measure of the observed memory in twelve regional GDP time series and of

[5] Note that world estimates here are simply the sum of the results for the seven regions in Table A5.
[6] http://www.cgd.ucar.edu/cas/wigley/magicc/; the regional scaling factors are: 1.56 for Europe; 1.39 for Latin America; 1,49 for North America/OECD; 1.30 for Africa; 2.04 for North Asia; 1.33 for South Asia and; 1.45 for China.

the persistence of shocks. It is shown that even when allowing for a one structural change in the trend function, the sum of the first order autoregressive coefficients is quite large, leading to the non-rejection of the null hypothesis of a unit root in most cases. These results provide evidence of the importance of the memory of GDP series and of the large persistence of shocks, that is also supported by a vast econometric literature on the subject. As stated by Campbell and Mankiw (1987b) "much disagreement remains over exactly how persistent are shocks to output. Nonetheless, among investigators using post-war quarterly data, there is almost unanimity that there is a substantial permanent effect". By analyzing the DICE/RICE models, we show that even IAMs that treat economic growth endogenously impose a null persistence to climate change impacts. The "zero persistence" implied by IAMs is an extreme assumption regarding the memory of GDP time series that is not supported by the observed data. As is argued in this paper, this zero persistence assumption implies an autonomous, extremely large and effective, costless adaptation capacity to no matter how large impacts that climate change may produce in both human and natural systems. Both of these assumptions are of course not realistic and are in stark contrast with what has been observed.

In the remaining sections of this paper, the impact functions and parameterizations of the PAGE2002 were used to produce simulations experiments that allow to explore the importance of the persistence of climate change shocks in future GDP series. It is shown that the estimates of the costs of climate change are very sensitive to the inclusion of a "memory parameter" and that when this parameter is chosen to be similar to the observed sum of the autoregressive coefficients, estimates of the present value of the costs of climate change are about from 80% to 90% larger than when the persistence or shocks is ignored, as occurs when using the original scaling method. These results are not sensitive to the discount rate that is used.

Estimates of the economic impacts of climate change under the A2 emissions scenario are presented for various values of the memory parameter and for a set of regionally differentiated values based on the sum of the autoregressive coefficients estimated from the regional observed GDP series. Results are dramatically different with respect to the estimates that are produced ignoring the persistence of shocks. Including the persistence of shocks not only leads to larger and presumably more realistic estimates of the cost/benefits of climate change but also points out that regional differences may be larger than previously estimated. This paper presents an easy to implement approach for incorporating the memory of GDP to IAM in order to estimate the potential economic costs of climate change.

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Table 4. Present value of the total climate change economic impacts over a century as a percentage of the GDP in year one. Numbers in parentheses represent the increase (%) in comparison to the estimates produced assuming α=0.

α	5th percentile	Mean	95th percentile
0	5.68%	14.35%	25.21%
0.1	6.40% (12.68%)	16.03% (11.71%)	27.75% (10.08%)
0.5	8.33% (46.65%)	21.09% (46.97%)	37.08% (47.08%)
0.8	10.40% (83.105)	25.39% (76.93%)	44.76% (77.55%)
0.9	11.49% (102.29%)	26.63% (85.57%)	46.56% (84.69%)

Table A5. Concordance of regions.

PAGE2002	Regions in this study
Europe	Europe
Eastern Europe and Former Soviet Union	North Asia
USA	North America/OECD
China	China
India	South Asia
Africa	Africa
Latin America	Latin America
Other OECD	North America/OECD

Table 6. Estimates of global and regional climate change impacts under the A2 emissions scenario and for different values of α.

Region	α=0	α=0.1	α=0.5	α=0.8	α=0.9	α	NPV
Europe	0.11 (0.36, 0.79)	0.12 (0.40, 0.87)	0.53 (0.15, 1.13)	0.64 (0.19, 1.39)	0.67 (0.20, 1.47)	0.5	0.53 (0.15, 1.17)
Latin America	1.65 (0.44, 3.84)	1.80 (0.47, 4.27)	2.43 (0.62, 5.51)	2.92 (0.77, 6.73)	3.06 (0.82, 7.12)	0.65	2.66 (0.69, 6.23)
South Asia	0.70 (2.56, 5.83)	2.82 (0.76, 6.72)	3.73 (0.96, 8.41)	4.48 (1.20, 10.14)	4.73 (1.29, 10.79)	0.8	4.46 (1.15, 10.23)
North Asia	-0.69 (-2.18, 0.04)	-0.76 (-2.41, 0.05)	-1.00 (-3.02, 0.07)	-1.20 (-3.87, 0.08)	-1.28 (-4.04, 0.08)	0.8	-1.22 (-3.89, 0.09)
North America	0.10 (0.02, 0.25)	0.11 (0.02, 0.28)	0.14 (0.02, 0.34)	0.17 (0.03, 0.43)	0.18 (0.03, 0.45)	0.6	0.15 (0.03, 0.38)
Africa	2.02 (0.54, 4.58)	2.23 (0.58, 5.10)	2.97 (0.75, 6.73)	3.56 (0.93, 7.99)	3.74 (1.01, 8.49)	0.9	3.72 (0.98, 8.56)
China	0.94 (0.17, 2.39)	2.12 (0.56, 4.83)	2.45 (0.65, 5.44)	2.74 (0.72, 6.10)	2.81 (0.77, 6.45)	0.8	2.70 (0.73, 6.12)
World	0.52 (0.15, 1.14)	0.62 (0.17, 1.39)	0.80 (0.22, 1.74)	0.96 (0.27, 2.09)	1.01 (0.29, 2.23)	--	0.90 (0.27, 2.00)

Note: Figures denote the number of times of the GDP in year 2000 that the present value of the economic impacts of climate change over this century amounts to. Figures in parentheses give the 95% confidence intervals.

Figure 2. Impulse response (left) and accumulated impulse response (right) functions of AR(p) models based on Table 1 (panel a) and on Table 3 (panel b), while in panel c estimates are based on AR(1) models with the α parameter values shown in Table 3.

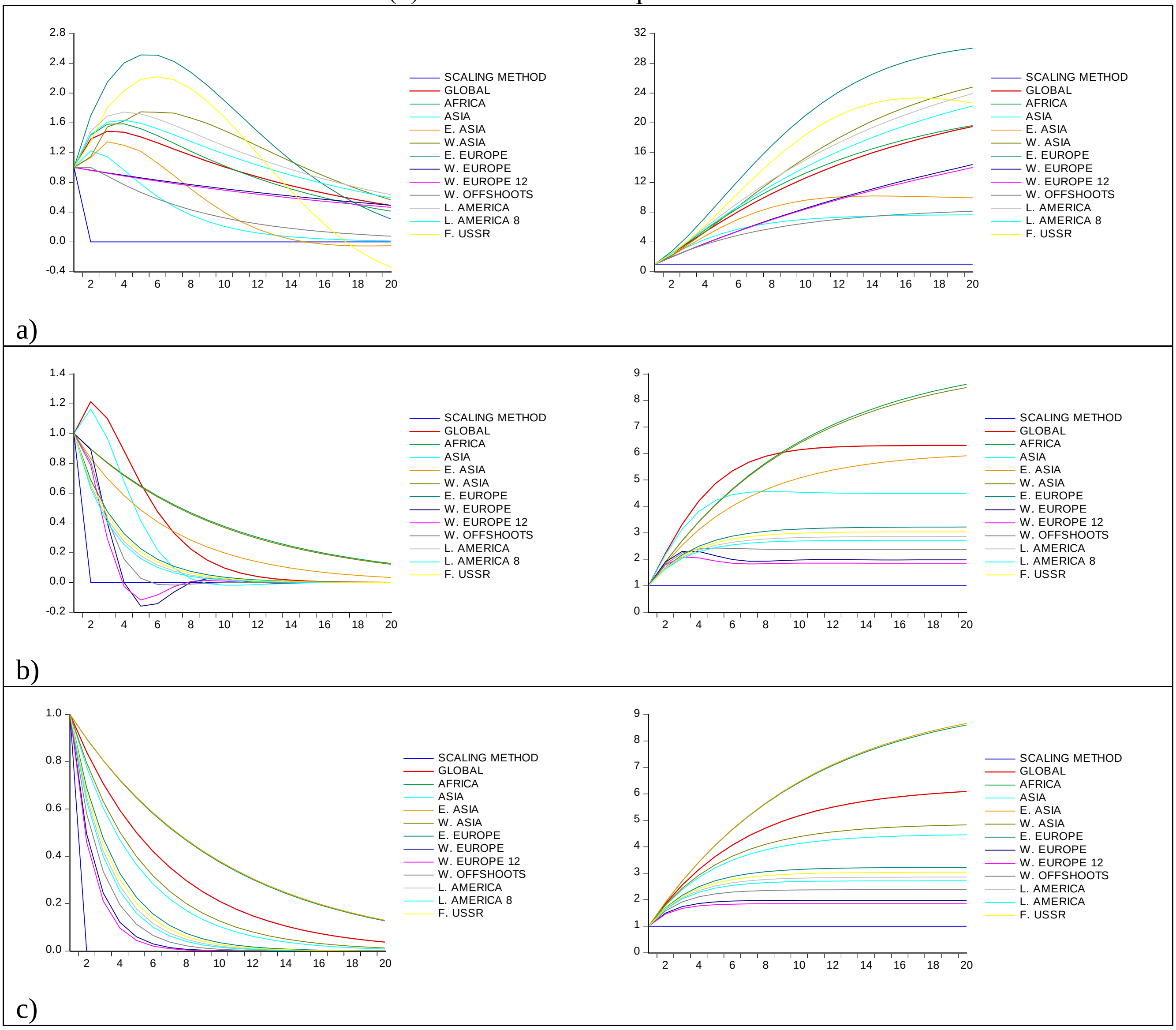
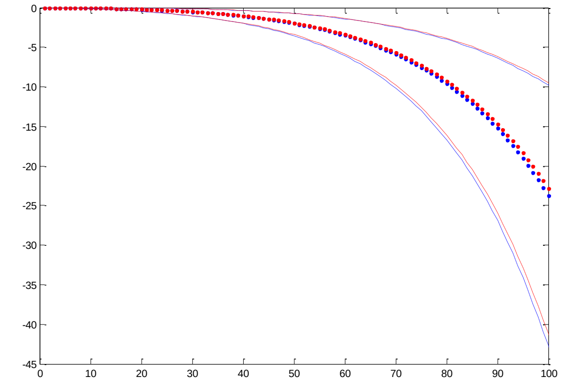
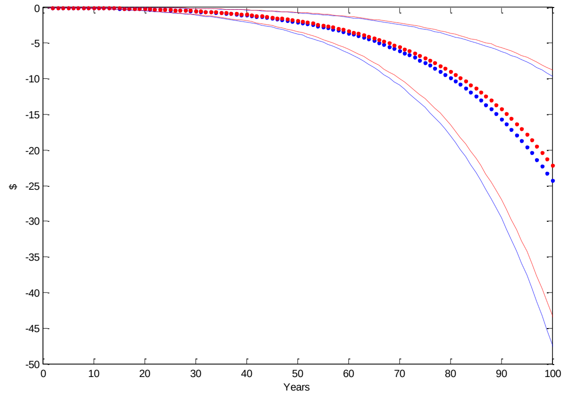


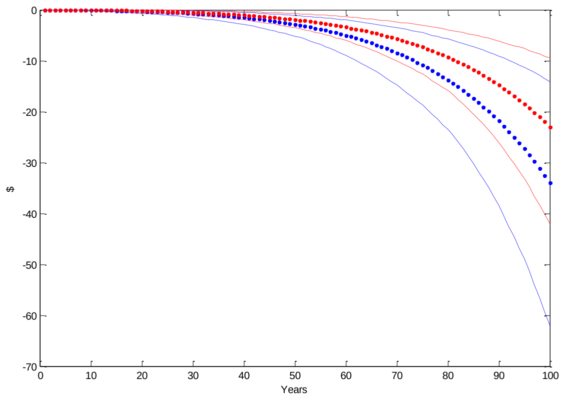
Figure 4. Total economic impacts for different values of parameter α. Dotted lines represent the mean, while the dashed ones the 5th and 95th percentiles. Red lines were produced using α=0, while blue lines use positive values of α.



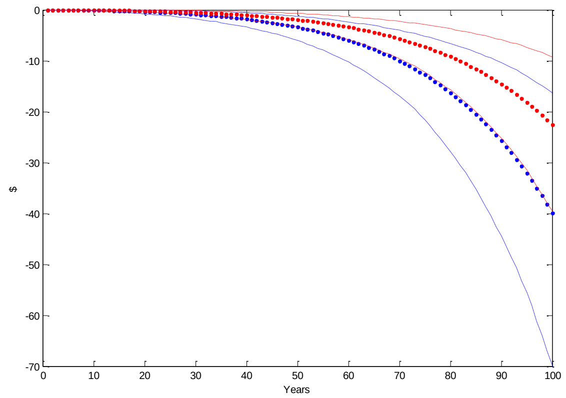
α=0.04 (DICE)



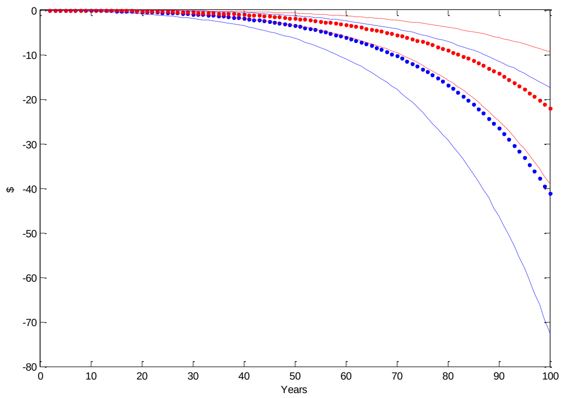
α=0.1



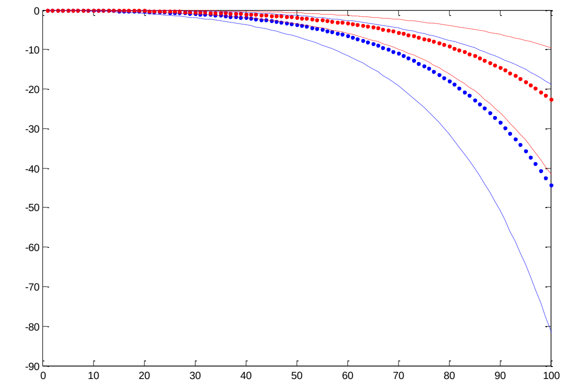
α=0.5



α=0.8



α=0.9



α=1 (FUND)

Appendix Methods and additional results

Methods

Difference stationarity versus trend stationarity

A series that is stationary in levels is said to be integrated of order zero or I(0). An I(1) has to be differenced once to achieve stationarity, while an I(2) series has to be differenced twice. A first order autoregressive process (AR(1)) with a coefficient in the autoregressive term equal to unity and a drift is an example of an I(1) and can be expressed as follows:

$$y_t = \beta + y_{t-1} + e_t$$

$$\Delta y_t = \beta + e_t$$

The solution of this difference equation is

$$y_t = y_0 + \beta t + \sum_{i=0}^{t-1} e_{t-i}$$

where y_0 is the initial condition, βt is a deterministic trend and $\sum_{i=0}^{t-1} e_{t-i} = v_t$ has a stochastic trend produced by the sum of the stationary error term (Maddala and Kim, 1998). Both the mean and variance of this process are time dependent $E(y_t) = \beta t$, $Var(y_t) = E(v_t^2) = t\sigma_e^2$, respectively.

On the other hand, a trend stationary process consists of a deterministic component plus a stochastic stationary process which can range from a simple white noise to a variety of different types of autoregressive and moving average structures such as AR, MA, ARMA. A simple example of this class of process is an AR(1) equation of the form:

$$x_t = \alpha + \beta t + \phi x_{t-1} + e_t$$

where $|\phi| < 1$ is a constant, $e_t \sim i.i.d(0, \sigma^2)$ is white noise, βt is a deterministic time trend and α is the intercept. The deterministic component of this process dominates its long run behavior: variations are transitory and do not change the long run path of the series (Enders, 2009). These processes are mean reverting around a trend function of the form $E(x_t) = \alpha + \beta t$.

Stationarity and structural breaks

Perron (1989; 1997) proposes three different models under the null hypothesis of a unit root:

- Model (A): a "crash" model that allows for an exogenous change in the level of the series;
- Model (B): a "changing growth" model that permits an exogenous change in the rate of growth;

- Model (C): a model that allows for both of these changes to occur at the same time.

The null hypotheses are parameterized as follows:

Model (A) $y_t = \mu + dD(TB)_t + y_{t-1} + e_t$,

Model (B) $y_t = \mu_1 + y_{t-1} + (\mu_2 - \mu_1)DU_t + e_t$,

Model (C) $y_t = \mu_1 + y_{t-1} + dD(TB)_t + (\mu_2 - \mu_1)DU_t + e_t$,

where $D(TB)_t = 1$ if $t = T_b + 1$, 0 otherwise; $DU_t = 1$ if $t > T_b$, 0 otherwise; and $A(L)e_t = B(L)v_t$, $v_t \sim i.i.d.\ (0, \sigma^2)$, with A(L) and B(L) being the *p*th and *q*th order polynomials in the lag operator.

The models under the alternative hypothesis of trend stationarity are:

Model (A) $y_t = \mu_1 + \beta t + (\mu_2 - \mu_1)DU_t + e_t$,

Model (B), $y_t = \mu + \beta_1 t + (\beta_2 - \beta_1)DT_t^* + e_t$,

Model (C) $y_t = \mu_1 +_t \beta_1 t + (\mu_2 - \mu_1)DU_t + (\beta_2 - \beta_1)DT + e_t$,

where $DT_t^* = t - T_b$, $DT_t^* = t$ if $t > T_b$, 0 otherwise. T_b refers to the time of the break[7].

The methodologies of the Perron (1997) and Kim and Perron (2009) are briefly described below. The Perron (1997) tests is a generalization of the Perron (1989) test where the break date is treated as unknown. Here we consider only models B and C and apply the additive outlier approach in which the changes in the trend parameters are assumed to occur rapidly. The break dates are selected by minimizing or maximizing different statistics obtained from the regressions that are specified for each model (B and C) as follows:

$$y_t = z_{t,1}\phi_1 + z(T_b)'_{t,2}\phi_2 + \tilde{y}_t^i \qquad (1)$$

where $z_{t,1} = (1, t)'$, $\phi_1 = (\mu, \beta)'$,

$$z(T_b)_{t,2} = \begin{cases} DT_t^* & (for \quad Model \quad B) \\ (DU_t, DT_t^*) & (for \quad Model \quad C) \end{cases},$$

[7] Note that models A, B and C of the Perron (1989; 1997) and Kim and Perron (2009) correspond to models I, II and III of the Perron and Yabu (2009) test.

$$\phi_2 = \begin{cases} \gamma & (for \quad Model \quad B) \\ (\theta,\gamma)' & (for \quad Model \quad C) \end{cases}$$

and $\{\tilde{y}_t^i\}$, i= B, C are the residuals of the corresponding model. Then for the Perron (1997) test, the following ADF regression is estimated on the residuals $\tilde{y}_t^i$:

$$\tilde{y}_t^i = \alpha\tilde{y}_{t-1}^i + \sum_{i=1}^{k} c_i \Delta\tilde{y}_{t-i}^i + e_t \qquad (2)$$

One disadvantage of the Perron (1997) test is that, contrary to the original Perron (1989) test, the occurrence of a structural change is only allowed under the alternative hypothesis. As such, the rejection of the null of a unit root can occur because of the presence of a large structural change in the trend function even if the noise component is I(1).

To solve this problem, Kim and Perron (2009) proposed a unit root testing procedure that pretests the existence of the break using the Perron and Yabu (2009) procedure, and allows for the occurrence of a structural change both under the null and the alternative hypotheses, as was proposed in the original test of Perron (1989). When the pretest rejects the null of no structural change, the Kim and Perron (2009) test has the same limit distribution as if the break date was known and it has been shown to have greater power, maintain the correct size and offer an improvement over other commonly used methods in small samples.

The testing procedure of the Kim and Perron (2009) unit root test under the additive outlier approach consists in the following steps:

1) Use the estimate of T_b obtained by minimizing the sum of the squared residuals of regression (1) and construct a window around it defined by a lower bound T_l and an upper bound T_h. A window of 6 observations was chosen. Note that the results are not sensitive to the window size (see Kim and Perron, 2009);

2) Create a new data set $\{y^n\}$ by removing the data from $T_l + 1$ to T_h, and shifting down the data after the window by $S(T) = y_{T_h} - y_{T_l}$:

$$y^n = \begin{cases} y_t & if \quad t \leq T_l \\ y_{t+t_h-t_l} - S(T) & if \quad t > T_l \end{cases}$$

3) Perform the unit root test of the corresponding model using the break date T_l and compute the unit root t-test statistic, denoted by $t_\alpha\left(\hat{\lambda}_{tr}^{AO}\right)$, from the following regression:

$$\tilde{y}_t^n = \tilde{\alpha}\tilde{y}_{t-1}^n + \sum_{i=1}^{k} \tilde{c}_i \, \Delta\tilde{y}_{t-i}^n + \tilde{e}_t \qquad (3)$$

where $\hat{\lambda}_{tr} = T_l / T_r$, $T_r = T - (T_h - T_l)$ and $\tilde{y}_t^n$ is the detrended value of y^n.

The appropriate critical values for this test are those in Perron (1989) and in Table 1 of Perron and Vogelsang (1993).

Additional results
Table A1, A2, A3
Figure A1

Table A1. Results of standard ADF tests applied to the natural logs of annual GDP data and CIR values.

Region	K	$\hat{\mu}$	$\hat{\beta}$	$\hat{\rho}$	$\tau(\hat{\rho})$	CIR
Global	1	0.68 [1.79]	0.0014 [1.54]	0.958	-1.71	23.81
Africa	1	0.54 [1.59]	0.0014 [1.49]	0.958	-1.53	23.81
Asia	1	1.84 [2.43]	0.0069 [2.37]	0.869	-2.38	7.63
Eastern Asia	2	1.08 [1.94]	0.0038 [1.93]	0.900	-1.90	10.00
Western Asia	2	0.46 [1.67]	0.0013 [1.16]	0.964	-1.54	27.78
Eastern Europe	1	0.43 [2.06]	0.0007 [1.59]	0.966	-2.01	29.41
Western Europe	0	0.51 [2.20]	0.0003 [0.66]	0.968	-1.97	31.25
Western Europe (12)	0	0.55 [2.36]	0.0004 [0.73]	0.965	-2.14	28.57
Western Offshoots	1	1.63 [1.65]	0.0035 [1.40]	0.889	-1.61	9.01
Latin America	1	0.44 [1.54]	0.0009 [1.09]	0.969	-1.44	32.26
Latin America (8)	1	0.46 [1.54]	0.0097 [1.06]	0.967	-1.42	30.30
Former USSR	2	0.74 [2.60]	0.0008 [1.73]	0.947	-2.57	18.87

Note: *, ** indicate statistical significance at 10% and 5% levels, respectively. The lag length K was chosen using the Schwarz Information Criterion (BIC).

Table A2. Test for structural changes in the trend function

Series	Model	Exp-Wald statistic value
Global	II	5.55[a]
Africa	III	4.52[b]
Asia	II	3.37[a]
Eastern Asia	III	3.00[c]
Western Asia	II	14.42[a]
Eastern Europe	III	17.58[a]
Western Europe	II	4.63[a]
Western Europe (12)	II	3.36[a]
Western Offshoots	II	2.50[b]
Latin America	III	6.53[a]
Latin America (8)	III	5.74[a]
Former USSR	III	19.15[a]

Model II allows a one-time change in the slope of the trend function; Model III permits both a one-time change in the slope and in the intercept of the trend function. A 5% trimming was used for this test. a, b, c denote statistical significance at the 1%, 5% and 10% respectively.

Table A3. Tests for a unit root with a one-time break in the trend function and CIR values.

	T_b	k	$\hat{\mu}$	θ	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\alpha}$	$t_\alpha\left(\hat{\lambda}_{tr}^{AO}\right)$	CIR
Global	1973	1	15.49 **(1952.83)**	--	0.0469 **(97.33)**	-0.0148 **(-21.28)**	0.841	-2.015	6.29
Africa	1983	0	12.20 **(950.79)**	-0.1343 **(-6.83)**	0.0430 **(62.31)**	-0.0118 **(-9.83)**	0.897	-2.947	9.71
Asia	1974	1	13.81 **(1277.37)**	--	0.0568 **(89.32)**	-0.0064 **(-6.78)**	0.777	-3.900[D]	4.48
Eastern Asia	1994	0	10.56 **(621.19)**	-0.2176 **(-6.55)**	0.0388 **(57.03)**	0.0200 **(5.73)**	0.898	-2.415	9.80
Western Asia	1976	0	11.52 **(577.22)**	--	0.0714 **(64.81)**	-0.0407 **(-23.15)**	0.795	-2.258	4.88
Eastern Europe	1990	0	12.23 **(456.34)**	-0.5550 **(-11.76)**	0.0380 **(32.13)**	-0.0061 **(-1.59)**	0.690	-2.527	3.23
Western Europe	1973	1	14.16 **(3032.16)**	--	0.0463 **(162.96)**	-0.0245 **(-59.75)**	0.495	-3.280	1.98
Western Europe (12)	1972	1	14.08 **(2818.99)**	--	0.0458 **(145.46)**	-0.0251 **(-56.77)**	0.460	-3.833[D]	1.85
Western Offshoots	1972	1	14.31 **(1571.10)**	--	0.0386 **(67.27)**	-0.0085 **(-10.56)**	0.580	-3.083	2.38
Latin America	1982	0	12.91 **(1500.47)**	-0.1426 **(-10.99)**	0.0521 **(109.25)**	-0.0245 **(-31.42)**	0.650	-4.240[B]	2.86
Latin America (8)	1982	0	12.76 **(1430.70)**	-0.1414 **(-10.52)**	0.0529 **(107.04)**	-0.0252 **(-31.21)**	0.632	-4.189[B]	2.72
Former USSR	1992	0	13.29 **(362.09)**	-0.8129 **(-11.98)**	0.0349 **(22.68)**	-0.0020 **(-0.32)**	0.671	-2.271	3.04

The regression model for the unit root tests is defined in equations (1) and (3) of the Methods section. The symbols used above have meanings as follows: T_b is the estimated time of the break; k is the number of lagged differences added to correct for serial autocorrelation; $\hat{\mu}$, $\hat{\beta}$, $\hat{\gamma}$ are the regression coefficients of the trend function and the corresponding t-statistic values are shown in parentheses. Bold numbers denote statistical significance at 5% levels. $\hat{\alpha}$ is the sum of the first order autoregressive coefficients and $t_\alpha\left(\hat{\lambda}_{tr}^{AO}\right)$ is the unit root test statistic corresponding to the Kim and Perron (2009) test. A, B, C, D denotes statistical significance at the 1%, 2.5%, 5% and 10% respectively (Perron and Vogelsang (1993) critical values Table 1). Lag length was selected using the Schwarz Information Criteria (BIC).

Figure A1. Natural logs of annual GDP data for the selected regions