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Population size and the size of government

Abstract

We examine the effect of population size on government size for a panel of 130 countries for the period between 1970 and 2014. We show that previous analyses of the nexus between population size and government size are incorrectly specified and fail to consider the influence of cross-sectional dependence, non-stationarity and cointegration. Using a panel time-series approach that adequately accounts for these issues, we find that population size has a positive long-run effect on government size. This finding suggests that effects of population size that increase government size (primarily due to the costs of heterogeneity, congestion, crime and conflict) dominate effects that reduce government size (primarily due to scale economies).

JEL-Codes: H110, H500.

Keywords: government size, population size, non-stationary, cross-sectional dependence, panel cointegration.

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1. Introduction

Government size (i.e., government spending as a share of GDP) is an important macroeconomic variable. For instance, by increasing the burden of taxation and crowding out private economic activity and investment, government size has been found to negatively affect factor accumulation and productivity (e.g., Dar and AmirKhalkhali, 2002). Consequently, empirical studies on the determinants of economic growth often report a negative correlation between economic activity and government size (for a review, see Bergh and Henrekson, 2011).[1]

Given its potentially substantial economic importance, a considerable theoretical and empirical literature has sought to identify the determinants of government size, investigating, inter alia, the role of a country's level of economic development, its degree of ethnic fragmentation and the nature of its political institutions (for a brief review, see Shelton, 2007: 2234-2240). Among potential determinants of government size, *population size* has gained considerable prominence in the literature.

Indeed, several advantages may allow more populous countries to afford smaller governments. First, more populous countries can capitalize on *scale economies* associated with the provision of public goods (Alesina, 2003). Fixed costs of public goods and increasing returns to scale may make it possible for more populous countries to allocate fewer resources to public spending (in relation to total GDP). For instance, Andrews and Boyne (2009) show that administrative costs are lower in larger local governments for a sample of English communities, a finding consistent with the notion of economies of scale. Second, more populous countries are less likely to be threatened by *foreign aggression*, given that their sheer size discourages war (Alesina, 2003). This in turn allows larger countries to spend comparatively less on defense and security, again negatively affecting government size. Third, more populous countries benefit from comparatively *larger domestic markets*, creating fewer incentives to engage in international trade and competition. Thus, more populous countries are less exposed to the volatility and external risk associated with openness (Alesina and Wacziarg, 1998). By contrast, more open (i.e., smaller) economies face more risk;

[1] For instance, Barro (1990) and Glomm and Ravikumar (1997) discuss the role of government size in economic growth from a theoretical perspective.

they may consequently use government spending to mitigate associated risks, thus increasing the size of government (Rodrik, 1998).

Empirical studies find mixed evidence on the population-government size nexus. In their seminal analysis, Alesina and Wacziarg (1998) find that population size is indeed negatively associated with government size. A similar result is obtained by Benarroch and Pandey (2008). Shelton (2007) also finds that government spending tends to decrease with population size.

However, other empirical studies prove less conclusive and fail to show that more populous countries have smaller governments. Rodrik (1998) reports no statistically significant association between population size and government size. Similarly, Jetter and Parmeter (2015) find that the effect of population size on government size is dependent on empirical choices (e.g., considering the use of specific datasets). Finally, Ram (2009) finds that while population size is negatively related to government size in a pooled OLS setting (thus mimicking the approach of Alesina and Wacziarg, 1998), the relationship between both variables actually becomes positive in a fixed-effects setting.

Indeed, theory suggest that population size may not only have effects that reduce government size. Rather, certain factors may make more populous countries more likely to expand the size of their government. First, the benefits of size (primarily, scale economies) may decrease when public goods provided by government spending are subject to *congestion* (e.g., Oakland, 1972). For instance, congestion is expected to incur administrative costs when it leads to the rationing of public goods (Oakland, 1972). Consequently, the costs of managing congestion may offset or even outweigh the advantages of size due to scale economies. Second, Alesina (2003) argues that more populous countries face higher *costs of heterogeneity of preferences*. For instance, more populous countries exhibit more interest groups and political parties, reflecting the country's (comparatively) high level of heterogeneity (e.g., Murrell, 1984). In turn, a larger number of interest groups and political parties is expected to increase government size, e.g., as (diverse) interest groups and parties will have to accommodate many pet projects to form a winning coalition (e.g., Mueller and Murrell, 1986; Mukherjee, 2003). Third, population size may contribute to costly *social deviance*. For instance, more populous communities tend to experience disproportionately more crime due to reduced social control and solidarity (Chamlin and Cochran, 2004). Reduced social control due to increasing population size may also contribute to other forms of deviance (e.g., corruption), which are expected to require an expansion of the government (more police, establishment of anti-corruption agencies etc.) as a countermeasure, leading to a positive association between population and government size. Finally, population size is a strong positive predictor of *domestic conflict* such as civil war (for a review, see Blattman and Miguel, 2010) and terrorism (for a review, see Krieger and Meierrieks 2011). For instance, increases in population size may result in more conflict by exacerbating resource scarcity, distributional conflicts or environmental degradation (e.g., Blattman and Miguel, 2010; Brückner, 2010). In turn, increased risk of violent conflict can be expected to increase government size, as grievances may have to be met with higher public spending on social policies (education, health, social security etc.) (e.g., Taydas and Peksen, 2012) or with more public spending on security, the police and the military to suppress conflict.

Our discussion of the existing literature on the population-government size relationship can be summarized as follows. First, the theoretical effect of larger population size on government size is a priori unclear: the beneficial effects predicted to reduce government size (scale economies, reduced exposure to international aggression and markets) must be weighed against effects that may stimulate government size (costs due to congestion, heterogeneity, crime, corruption and domestic conflict). Second, the empirical evidence reflects this theoretical ambiguity, with some studies reporting a negative population-government size relationship (e.g., Alesina and Wacziarg, 1998) and others reporting positive or non-significant associations (e.g., Ram, 2009; Jetter and Parmeter, 2015).

Our paper adds to the diverse empirical evidence on the population-government size nexus in two fundamental ways. First, we uncover crucial methodological shortcomings associated with "traditional" approaches to the population-government size nexus. Specifically, traditional approaches rely on pooled OLS and fixed-effects models, leading to issues of *cross-sectional dependence*, *non-stationarity* and (panel) *cointegration*. Corresponding misspecifications result in invalid inferences about the population-government size relationship. Second, we address these methodological shortcomings by employing a novel empirical *panel time-series approach* (the common correlated effects mean-group error-correction model) that accommodates cross-sectional dependence, non-stationarity and (panel) cointegration. To preview our main finding, the estimates from this approach indicate that larger population size is *positively* related to government size, suggesting that the costs of size (due to congestion, crime, conflict etc.) dominate its potential benefits (e.g., from scale economies).

The remainder of this paper is organized as follows. In Section 2 we introduce the data and test the variables measuring population and government size for cross-sectional dependence, non-stationarity and (panel) cointegration, showing that all of these issues matter. Section 3 investigates the effect of population on government size using the "traditional" pooled OLS and fixed-effects approaches. Sources of misspecification when employing this approach are identified and discussed. Section 4 introduces the common correlated effects mean-group error-correction model. We show how this model eliminates various sources of misspecification. Using this model, we provide novel insights into the government size-population size nexus. We also investigate the issues of reverse causation and non-linearity. Section 5 concludes.

2. Data

For the following empirical analyses, we use balanced panel data for 130 countries for the 1970-2014 period. The summary statistics are reported in Table 1. A country list is provided in the appendix.

—Table 1 here—

2.1 Measuring Population and Government Size

Our choice of variables measuring population and government size reflects earlier empirical studies on the nexus between population and government size (e.g., Alesina and Wacziarg, 1998; Ram, 2009; Jetter and Parmeter, 2015). First, *government size* is measured as the share of government consumption at current purchasing power parities. Second, *population size* is measured by a country's population size in millions of inhabitants. Both data series are drawn from the *Penn World Table* (version 9.0) (Feenstra et al., 2015).

2.2 Cross-Sectional Dependence, Panel Unit Roots and Panel Cointegration

As emphasized in the introduction, we suspect that both data series are affected by cross-sectional dependence and non-stationarity, with the latter raising the possible issue of (panel) cointegration.

As we shall discuss below in more detail, disregarding these issues may contribute to misspecifications and incorrect inferences regarding the population-government size nexus.

Cross-Sectional Dependence. Cross-sectional dependence refers to the interdependency of variables of interest between countries, where this interdependency may be due to, e.g., common shocks (e.g., economic booms or recessions) or spillover effects (Sarafidis and Wansbeek, 2012). For our case, population size may be correlated across countries due to, e.g., common exposure to economic and ecological shocks (e.g., natural disasters). Similarly, government size may exhibit cross-sectional dependence due to, e.g., tax competition or regional arms races. If not accounted for, cross-sectional dependence in the panel data may lead to correlation in the residuals, consequently affecting estimation efficiency and the validity of inference (Sarafidis and Wansbeek, 2012).

Non-Stationarity. Variables that trend over time are often found to be non-stationary (i.e., containing a unit root). For our data, it is plausible that both population size and government size are non-stationary. For instance, global population size has obviously exhibited a long-run positive trend over the last several decades (the so-called "population explosion"). If a regression model includes two (or more) non-stationary variables, this may give rise to the spurious regression problem, as shown in a pioneering study by Granger and Newbold (1974). This proves problematic because significance tests on the regression coefficients from spurious regressions are invalid (Granger and Newbold, 1974; Kao, 1999). That is, when regression models include non-stationary variables, it is possible that significance tests indicate a "significant" relationship between variables when in fact none exists. Importantly, the problem of spurious regression also matters in the panel setting (e.g., Kao, 1999).

Panel Cointegration. When two variables are non-stationary and integrated of the same order, they may be cointegrated (Engle and Granger, 1987). Cointegration refers to the existence of a stationary linear combination of two non-stationary variables. Disregarding (panel) cointegration is expected to result in misspecification, leading to incorrect inferences (e.g., Granger, 1986; Engle and Yoo, 1987; MacDonald and Kearney, 1987). Accounting for cointegration allows for inferences about the long-run relationship between non-stationary variables, while also considering any short-run dynamics (Engle and Granger, 1987). For our case, it seems plausible that population and government size are cointegrated, sharing a stable long-run (cointegrating) relationship – which may either be positive or negative – along the theoretical lines discussed above, while short-run deviations (e.g., due to excessive public spending during recessions) from the long-run equilibrium may still exist.

Tests for Cross-Sectional Dependence, Panel Unit Roots and Panel Cointegration. To examine whether our variables of interest are indeed subject to cross-sectional dependence and non-stationarity, we run a series of statistical tests.

First, we test for cross-sectional dependence by employing Pesaran's (2004) CD-test, which tests the null hypothesis of cross-sectional independence against the alternative of cross-sectional dependence. Importantly, the CD-test is robust to non-stationarity (Pesaran, 2004), which may also matter to the variables we examine. Second, to investigate the data series' stationarity properties, we employ two different panel unit root tests, the Im-Pesaran-Shin test (IPS test) (Im et al., 2003) and the CADF test developed by Pesaran (2007). For both tests, the null hypothesis is that the investigated series contain unit roots (i.e., are non-stationary) versus the alternative that (a fraction of) the series are stationary. Importantly, both tests account for cross-sectional dependence.[2]

As shown in Table 2, both data series are indeed affected by cross-sectional dependence (Panel A), meaning that observations for government and population size are not independent across countries. As argued above, such interdependencies may be explained by exposure to, e.g., common shocks or spillover effects. For example, economic crises that transcend national boundaries, international economic integration and politico-economic cooperation, competition or hostilities between nation-states may play a role in this context.

The results of the panel unit root tests (Panel B) indicate that both data series are non-stationary in levels but stationary after first-differences are taken. These findings prove highly intuitive. First, the global population doubled between 1970 and 2014, from 3,682 to 7,349 million. This development may be due to medical advances, advances in hygiene and other socio-economic

[2] The IPS test does so by demeaning the data. For the CADF test, cross-sectional averages of lagged levels and first-differences of the investigated series are added to the standard augmented Dickey-Fuller regressions that are used to investigate non-stationarity.

factors that have allowed many developing countries to enter a stage of demographic transition with (relatively) low death but high birth rates. Second, trends towards larger governments are widely discussed in the literature, e.g., by Peltzman (1980), Holcombe (2005) and Durevall and Henrekson (2011). For instance, ratchet effects (where government size grows during times of crises but does not revert back to pre-crisis levels once the crisis is over) may explain a positive trend in government size (Holcombe, 2005).

—Table 2 here—

Given that both series are found to be non-stationary and integrated of the same order, the series may also be cointegrated, sharing a long-run equilibrium relationship. To assess whether this is the case, we employ the test for panel cointegration developed by Westerlund (2005). Here, we test the null hypothesis that the investigated series are cointegrated against the alternative that they are not. When employing the test, we subtract the cross-sectional averages from the series to mitigate the influence of cross-sectional dependence, which is warranted given the results reported above. As shown in Table 2 (Panel C), the different variants of the panel cointegration test unanimously suggest that population and government size are indeed cointegrated.[3]

3. Pooled OLS and Fixed-Effects Regressions

3.1 Empirical Approach

Having introduced and pre-tested the data, we begin our empirical analysis of the population-government size nexus by running a series of regressions using pooled OLS and fixed-effects approaches, following previous empirical efforts that have studied the effect of population on government size in such frameworks (e.g., Alesina and Wacziarg, 1998; Ram, 2009; Jetter and Parmeter, 2015). As in these studies, we consider a series of empirical specifications of the following form:

$$\mathrm{GOV}_{it} = \alpha_0 + \alpha_1 \mathrm{POP}_{it} + \alpha_2 X' + \theta_i + \varphi_t + \varepsilon_{it} \qquad (1)$$

Here, we relate population size (*POP*) to government size (*GOV*) for country *i* at year *t*. Both data series are measured as discussed above and log-transformed to remain comparable to previous

[3] Employing alternative panel cointegration tests by Pedroni (1999, 2001) and Westerlund (2007) yields the same finding (results available upon request).

empirical studies, while also being less affected by outliers.[4] Equation (1) also includes an intercept (α_0) and an idiosyncratic error term (ε). Furthermore, we include country fixed-effects (θ) when employing the fixed-effects estimator to account for (time-invariant) unobserved heterogeneity. Finally, a simple way to account for one potential source of cross-sectional dependence, common shocks, is to amend an empirical model by a set of year dummies (φ), as we do for some variants of (1). However, such an approach may not be sufficient to entirely expunge the cross-sectional dependence.

With respect to equation (1), inferences about the population-government size nexus are only valid when cross-sectional dependence and non-stationarity (and thus panel cointegration) are not influential. However, the pre-tests reported in Table 2 suggest that these assumptions may not be justified. Consequently, if cross-sectional dependence and non-stationarity are indeed influential in (1) but not accounted for, they will be "captured" in the regression residuals (i.e., the ε_t series). Consequently, below we subject the regression residuals to a number of diagnostic tests to examine whether misspecification issues are indeed present. In the presence of misspecification issues, the results from (1) will be misleading and potentially lead to incorrect inferences about the population-government size nexus.

3.2 Empirical Results

Our regression results are reported in Table 3. Employing the usual baseline specification of Alesina and Wacziarg (1998), Ram (2009) and Jetter and Parmeter (2015), we find that population size exerts a negative and statistically significant effect on government size. As in Ram (2009) and Jetter and Parmeter (2015), the estimated effects are much larger in the fixed-effects setting. These results are consistent with the arguments put forth by Alesina and Wacziarg (1998) and Alesina (2003) regarding the benefits of population size in reducing size of government, e.g., in the form of scale economies and reduced relative exposure to international markets.

—Table 3 here—

[4] Also, first-differences of log-transformed data series approximate their growth rates, facilitating the interpretation of results when first-differences are taken.

However, the diagnostics concerning cross-sectional independence and stationary residuals reported in Table 3 are clearly worrisome. First, tests of the regression residuals for unit root presence strongly indicate that the residuals are non-stationary.[5] As discussed above, non-stationary residuals may imply a spurious regression (e.g., Kao, 1999). They also suggest that a cointegrating relationship between population and government size ought to be modelled. Second, the majority of CD-test results indicate that the residuals are affected by cross-sectional dependence.[6] As discussed above, this may affect the validity of inference (Sarafidis and Wansbeek, 2012). In sum, the diagnostics reported in Table 3 indicate that the empirical results from a "traditional" approach to the population-government size nexus shown in Table 3 are likely misleading.

Table 3 also reports some "naïve" ways to remedy the misspecification issues. First, taking first-differences of both variables is expected to produce stationary variables and thus stationary residuals. Second, employing standard errors developed by Driscoll and Kraay (1998) ought to aid statistical inference, as these standard errors are not only robust to heteroskedasticity and autocorrelation, but also to general forms of cross-sectional dependence. Employing these remedies in models (5) and (6) of Table 3, we find that population size no longer exerts a statistically significant effect on government size, suggesting that the relationship between the two variables may indeed be spurious. However, the models in first-differences – though free of non-stationary residuals – discard valuable information about the long-run (cointegrating) relationship between population and government size. As argued above, incorrectly disregarding (panel) cointegration may be another source of misspecification and may therefore also lead to incorrect inferences (e.g., Granger, 1986, Engle and Yoo, 1987; MacDonald and Kearney, 1987).

4. Panel Time-Series Approach

[5] We only report the CADF-test results but the IPS-tests yield the same conclusion (results available upon request).

[6] The inclusion of year dummies can ameliorate the issue of cross-sectional dependence in the pooled OLS setting. However, their inclusion is not sufficient to account for cross-sectional dependence when fixed-effects models – which are preferred as they better reflect the panel structure of the data – are run.

4.1. Empirical Approach

Given the misspecification issues that plague the "traditional" pooled OLS and fixed-effects regression frameworks, in this section we employ a modelling approach that is able to account for cross-sectional dependence, while producing stationary residuals and incorporating a long-run (cointegrating) relationship between population and government size. In detail, we use the panel time-series approach of Pesaran (2006) and Chudik and Pesaran (2015), the (dynamic) *common correlated effects (mean-group) error-correction model*.[7] Below, we introduce this model in several steps, showing how these steps relate to misspecification issues that plague the "traditional" pooled OLS and fixed-effects regression frameworks.

As a first step, we account for non-stationarity and cointegration by considering the following *error-correction model* (*ECM*):

$$\Delta\mathrm{GOV}_{it} = \alpha_{0i} + \rho\left(\mathrm{GOV}_{i,t-1} - \beta POP_{i,t-1}\right) + \gamma^{p}\Delta POP_{it} + \gamma^{g}\Delta GOV_{i,t-1} + \epsilon_{it} \quad (2)$$

Here, government size and population size are first-differenced (indicated by the first-difference operator Δ) to achieve stationarity. Besides the intercept (α_0) and well-behaved error term (ε_{it}), equation (2) also includes the error-correction term $\rho(GOV_{i,t-1}-\beta POP_{i,t-1})$ which corresponds to the stationary linear combination of the levels of government and population size, allowing us to examine the long-run relationship between these variables (Engle and Granger, 1987).

We can reparametrize equation (2) to:

$$\Delta\mathrm{GOV}_{it} = \pi_{0i} + \pi^{EC} GOV_{i,t-1} + \pi^{P}\, POP_{i,t-1} + \pi^{p}\, \Delta POP_{it} + \pi^{g}\Delta GOV_{i,t-1} + \epsilon_{it} \quad (3)$$

Here, if the regression coefficient π^{EC} is statistically significant and lies between [0; -1] (implying dynamic stability), a long-run (cointegrating) equilibrium exists, where the exact value of π^{EC} indicates the speed of adjustment to it. π^{P} indicates the long-run effect of population size (in levels) on government size. An alternative way to measure this long-run effect is to recover β_i from equation (2) by $\beta_i = -\pi^{P}/\pi^{EC}$. Finally, π^{p} and π^{g} allow us to directly gauge the short-run effects of lags of the first-differences of population and government size on present values of government size (in first-differences).

[7] A highly instructive introduction to and application of this empirical method is provided by Eberhardt and Presbitero (2015).

As a final step, we add the cross-sectional averages of all variables in the model. Thus, we arrive at:

$$
\begin{aligned}
\Delta\mathrm{GOV}_{it} = \pi_{0i} &+ \pi_i^{EC} GOV_{i,t-1} + \pi_i^{P} POP_{i,t-1} + \pi_i^{p} \Delta POP_{it} + \pi_i^{g} \Delta GOV_{i,t-1} + \epsilon_{it} \\
&+ \pi_{1i}^{CA}\overline{\Delta\mathrm{GOV}_t} + \pi_{2i}^{CA}\overline{\mathrm{GOV}_{t-1}} + \pi_{3i}^{CA}\overline{\Delta\mathrm{POP}_t} + \pi_{4i}^{CA}\overline{POP_{t-1}} \\
&+ \sum_{l=2}^{p} \pi_{5il}^{CA}\overline{\Delta\mathrm{GOV}_{t-l}} + \sum_{l=1}^{p} \pi_{6il}^{CA}\overline{\Delta\mathrm{POP}_{t-l}} \qquad (4)
\end{aligned}
$$

Regarding (4), a number of remarks are necessary:

(i) Combining the first and second lines of equation (4) gives Pesaran's (2006) common correlated effects estimator. The terms in the second line are the cross-sectional averages. As argued by Pesaran (2006), the inclusion of these averages can accommodate cross-sectional dependence. That is, their inclusion provides consistent estimates of the parameters in the first line of equation (4) that are robust to cross-sectional dependence, i.e., unobserved common factors (due to spillover effects, global politico-economic shocks etc.) (Pesaran, 2006).[8]

(ii) Estimation equation (4) includes one lag of the dependent variable; below, we shall also add further lags of the dependent variable (as well as of the explanatory variable) to the model. This dynamic specification is expected to affect the consistency of the common correlated effects mean-group estimates due to endogeneity (Chudik and Pesaran, 2015). Chudik and Pesaran (2015) argue that by adding further lags of the cross-sectional averages, the common correlated effects mean-group estimators perform well again, even when allowing for weakly exogenous regressors in a dynamic setting. These additional lags of the cross-sectional averages are indicated by the third line of equation (4).

(iii) As in the fixed-effects model, we control for unobserved heterogeneity through a country-varying intercept. However, heterogeneity is not necessarily only time-invariant and independent of the explanatory variables (which would be accounted for by an intercept that varies by country). For instance, it is plausible that systematic and time-varying differences exist between countries

[8] The parameter estimates associated with the cross-sectional averages have no meaningful interpretation on their own; thus, we do not report them in our regression tables.

in preferences over welfare spending and redistribution (both of which are expected to increase government size) (e.g., Corneo and Grüner, 2002). Such differences could result in heterogeneous responses in government size with respect to changes in population size. Indeed, Pesaran and Smith (1995) show that the incorrect assumption of parameter homogeneity produces inconsistent and potentially misleading estimates of the regression coefficients. Consequently, to account for more complex forms of heterogeneity, we apply the mean-group approach of Pesaran and Smith (1995).[9] That is, we allow all parameters to vary by country; in contrast, they were set equal across countries in equations (2) and (3). To arrive at the mean-group estimates, we first estimate a series of country-specific regressions and then average the estimated coefficients across countries. The associated standard errors are derived non-parametrically following Pesaran and Smith (1995).

(iv) Baltagi et al. (2000) argue that the bias due to the incorrect assumption of parameter homogeneity needs to be weighed against the efficiency gains from pooling. They argue that allowing for parameter heterogeneity through a mean-group approach – even if warranted on theoretical grounds – may produce inferior results compared to a pooled approach. Therefore, we also estimate equation (4) in a pooled variant (with the short- and long-run coefficients being constrained to be equal across all countries) described in Pesaran (2006), with cross-sectional dependence still being controlled for by the inclusion of cross-sectional averages. To decide whether a heterogeneous or pooled variant is to be preferred, we follow Baltagi et al. (2000) and calculate the *root mean square errors* (*RMSE*) associated with each variant, consequently choosing the variant that minimizes the RMSE.

4.2 Empirical Results

The (dynamic) common correlated effects estimation results are reported in Table 4. Given that a mean-group (heterogeneous) modelling approach yields a smaller RMSE compared to a pooled (homogeneous) approach, we follow Baltagi et al. (2000) and prefer the mean-group over a pooled approach. Consequently, we will only report and discuss the mean-group findings.[10]

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[9] Without the inclusion of cross-sectional averages, the model represented in equation (4) is equivalent to the mean-group model of Pesaran and Smith (1995).

[10] In Table 4 we only report (for the sake of brevity) one pooled-CCE regression result which we compare with an otherwise identically specified MG-CCE result, where the latter yields a smaller

We are most interested in the long-run effect of population size on government size. This effect is calculated (and reported) in two ways. First, we report the long-run coefficient associated with the lag of the level of population size, which corresponds to the average of coefficient π_i^P from equation (4). Second, we report the average long-run coefficient of population size, which is equal to $\beta_i = -\pi_i^P / \pi_i^{EC}$ (using the average coefficients) from equation (2); for this estimate, the standard errors and associated *t*-statistics are calculated using the Delta method.

Independent of the specification of the short-run dynamics, these long-run estimates strongly indicate that population size exerts (on average) a *long-run positive effect* on government size. This finding suggests that population factors that are costly and thus increase government size (e.g., congestion, conflict risk and costs of heterogeneity and social deviance) are more important than factors that negatively affect government size (scale economies, military deterrence and trade effects). Notably, this finding stands in stark contrast to our earlier findings from the "traditional" pooled OLS and fixed-effects approaches, where we found that population size decreases government size. Consequently, our findings also contrast with earlier empirical contributions on the government size-population size nexus. For instance, our findings are not in line with Alesina and Wacziarg (1998), who argue that scale economies lead to a negative association between population and government size.

—Table 4 here—

Contrary to the "traditional" estimates reported in Table 3, the results reported in Table 4 are not affected by misspecification, suggesting that the latter are more trustworthy than the former. First, we are never able to reject the CD-test null hypothesis of cross-sectional independence. That is, by introducing (lags of) cross-sectional averages we are able to account for cross-sectional dependence, as argued by Pesaran (2006) and Chudik and Pesaran (2015). Second, the regression residuals are always found to be stationary. In addition, the long-run estimates are dynamically

RMSE. However, we also compare all other (dynamic) MG-CCE models reported in Table 4 with their pooled counterparts. The calculated RMSE always suggest that a heterogeneous modelling approach is preferred over the homogeneous (pooled) approach (results available upon request).

stable and, given their negative sign and statistical significance, highly indicative of a cointegration relationship.

Considering the short-run effects, lags of (first-differenced) government size predict its present values. By contrast, there are no significant short-run effects of lags of population growth (i.e., first-differenced population size) on the growth of government. As similarly argued by Eberhardt and Presbitero (2015), the lack of significance in short-run effects does not necessarily imply that population growth does not affect the growth of government; rather, the short-run relationship appears to be highly heterogeneous, with dynamics on average cancelling each other out. By contrast, this heterogeneity does not appear to be very influential in the long run.

As a robustness check, we amend equation (4) with additional controls for per capita income and the age dependency ratio (the ratio of those not in the labor force, i.e., children and the elderly, to those in the labor force, i.e., individuals aged between 15 and 65). Data on per capita income is from the Penn World Tables, while the age dependency ratio data is drawn from the *World Development Indicators* (World Bank, 2017). Both variables are frequently named as determinants of government size (e.g., Shelton, 2007), e.g., with per capita income potentially driving government size via Wagner's law and unfavorable demographic conditions (i.e., a large dependency ratio) leading to larger government size due to increased public spending on education, health or old age care. As shown in the appendix (Supplementary Table 1), adding these variables to the model does not change our main finding of a positive (cointegrating) relationship between government and population size. This speaks to, inter alia, Lütkepohl (2007: 322) who argues that a cointegration relationship ought to be robust to model extensions. That is, a cointegrating relationship is expected to hold even when additional variables are added to the model. Consequently, a parsimonious model – which in our case only considers population size and government size and their short- and long-run dynamics – will be sufficient, particularly in the context of cointegration analysis (Lütkepohl, 2007).

4.3 Reverse Causation

The cointegration results in Tables 2 (Panel C) and 4 imply that there exists a long-run relationship between population and government size. If two variables are cointegrated, one variable must Granger-cause the other or there must be Granger causality in both directions simultaneously (Engle and Granger, 1987). That is, while the panel cointegration test results show that population and government size are (Granger-causally) linked, they do not indicate the "direction" of Granger causality. So far, we have assumed – following the existing literature – that Granger causality runs from population size (as the independent variable) to government size (as the dependent variable). However, feedback between both variables may also exist. For instance, government size is expected to correlate with increased public spending on health, education and welfare. Such increased public spending may disincentivize "quantity" over "quality" with respect to childbearing, thus reducing population growth at the macro-level. Conversely, increased welfare spending may also attract international migration, consequently fueling population growth. While the nature of the effect of government size on population size is thus a priori unclear, it is nevertheless necessary to test whether feedback exists, as such an effect would question the validity of the empirical findings reported above.

To investigate whether government size also impacts population size, we consider the following specification:

$$\begin{aligned}\Delta \mathrm{POP}_{it} = {} & \pi_{0i} + \pi_i^{EC} POP_{i,t-1} + \pi_i^{P} GOV_{i,t-1} + \pi_i^{p} \Delta GOV_{it} + \pi_i^{g} \Delta POP_{i,t-1} + \epsilon_{it} \\ & + \pi_{1i}^{CA} \overline{\Delta \mathrm{POP}_t} + \pi_{2i}^{CA} \overline{\mathrm{POP}_{t-1}} + \pi_{3i}^{CA} \overline{\Delta \mathrm{GOV}_t} + \pi_{4i}^{CA} \overline{GOV_{t-1}} \\ & + \sum_{l=2}^{p} \pi_{5il}^{CA} \overline{\Delta \mathrm{POP}_{t-l}} + \sum_{l=1}^{p} \pi_{6il}^{CA} \overline{\Delta \mathrm{GOV}_{t-l}} \qquad (5)\end{aligned}$$

Equation (5) corresponds to equation (4), with the dependent and independent variable being inverted.[11] As above, the inclusion of (lagged) cross-section averages accounts for cross-sectional dependence, while first-differencing and the inclusion of an ECM accounts for non-stationarity and cointegration.

We summarize our empirical findings from (5) in Table 5. Here, we only report the long-run estimates for an effect of government size on population size, given that the short-run estimates – as with the other direction of causality reported in Table 4 – tend to be uninformative.[12] As shown in Table 5, regardless of which lag order of the short-run dynamics we choose, there is never a long-run effect running from government size to population size (Panel A). Here, the diagnostics indicate that the underlying models are specified correctly. By contrast, we previously found that

[11] This approach is also proposed in Eberhardt and Presbitero (2015).

[12] The short-run estimates are available upon request.

population size always exerts a positive and statistically significant long-run effect on government size. For comparison, these findings are also presented in a concise fashion in Table 5 (Panel B). In sum, the empirical results of Table 5 therefore indicate that while (i) cointegration between population and government size exists, (ii) Granger causality runs from population size to government size but (iii) not vice versa, so that (iv) the results reported in Table 4 are not affected by feedback and thus remain valid.

—Table 5 here—

4.4 Further Empirical Analyses

While our dynamic mean-group approach allows for a maximum of country-specific heterogeneity, it may nevertheless be fruitful to also consider whether our main result – population size increases government size in the long run – is also relevant to sub-samples of countries that differ with respect to specific characteristics. Below, we differentiate between (i) (relatively) poor and rich economies and (ii) (relatively) small and large countries, where the latter allows for a non-linear effect of population size on government size.

Economic Development. The nexus between population and government size may be different between rich and poor countries. Here, it is a priori unclear whether the effect of population size on government size is more or less pronounced in richer economies. On the one hand, richer countries tend to be less affected by social deviance (e.g., crime) and violent conflict (e.g., civil wars; see Blattman and Miguel, 2010). Thus, richer countries may have to devote fewer resources to anti-crime and anti-conflict measures as their populations grow, so that the effect of population size on government size may become weaker as the level of economic development increases. On the other hand, richer countries tend to be more open to international trade, e.g., as found in Ram (2009); in turn, increased exposure to trade may create demand for higher government spending to insure against the risks of trade (Rodrik, 1998). Finally, Wagner's law postulates that richer countries are generally more prone to government expansion (e.g., Shelton, 2007), as richer countries are expected to fund public goods (e.g., culture) for which scale effects may be less important.

Non-Linear Effects. The influence of population on government size may differ with the total size of the population, suggesting a non-linear effect of the former on the latter. For instance, congestion costs (which are expected to increase with population size and stimulate government growth) may be negligible below a certain population threshold and therefore may only matter for fairly large countries (Alesina, 2003). Similarly, the costs of heterogeneity and increased conflict risk may only become pronounced above a certain population threshold.

Empirical Results. We run a series of common correlated effects mean-group estimations as specified in equation (4) for various sub-samples. To create these sub-samples, we use the interquartile mean of population size and per capita income. Relying on the interquartile mean provides some protection against outliers; at the same time, it allows us to split the sample into two sub-samples of roughly equal size.

Our empirical results are reported in Table 6. First, we find that population size only increases government size in countries with a population of more than 10 million inhabitants. For countries with less than 10 million inhabitants, there is no significant (positive or negative) long-run effect of population size on government size. This result may indicate that the detriments of population size (which consequently stimulate government growth) only materialize above a certain population threshold, so that population and government size are potentially non-linearly related. Second, we find that population size has a positive long-run effect on government size in relatively rich and poor countries. Thus, a country's level of economic development does not seem to play an obvious role in moderating the population-government size nexus.

—Table 6 here—

5. Conclusion

There are conflicting schools of thought regarding the effect of population size on government size. One school argues that more populous countries benefit from scale economies and reduced exposure to the risks of international conflict and trade and can thus afford smaller governments. Another school of thought argues that more populous countries necessitate larger governments to counter congestion, heterogeneity costs and the ill effects of a larger population size on social deviance and domestic conflict.

Given these conflicting lines of argument, we examine the population-government size nexus for a panel of 130 countries for the 1970-2014 period. We find that "traditional" pooled OLS and fixed-effects approaches to this nexus are incorrectly specified, as they fail to properly account for cross-sectional dependence, non-stationarity and cointegration. Consequently, we employ a panel time-series approach that adequately considers these issues. With this novel empirical approach, we find that population size has a positive long-run effect on government size, suggesting that the effects of population size that promote larger governments (more congestion, increased costs of heterogeneity, social deviance and conflict) dominate effects that reduce government size. As an extension to our empirical analysis, we show that this effect tends to be more important to countries with more than 10 million inhabitants, potentially suggesting a non-linear relationship between population and government size.

Populations in many developing countries (especially in Africa, Asia and Latin America) are expected to grow substantially in the coming decades. In light of our findings, as their populations increase, these countries cannot expect to see their government size shrink relative to GDP. Instead, the opposite appears to be true. Given the large empirical literature linking oversized governments to undesirable socio-economic outcomes (reduced economic growth, crowding-out of private investment etc.), policymakers therefore would do well to pay close attention to the role of population size in determining government size, particularly in developing countries and emerging markets.

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Tables

Variable	N*T	Mean	Std. Dev.	Min.	Max.
Government Size	5,850	2.85	0.50	0.51	4.56
Population Size	5,850	1.99	1.89	-2.95	7.22
Δ Government Size	5,720	0.01	0.14	-1.39	1.63
Δ Population Size	5,720	0.02	0.02	-0.20	0.18
Per Capita GDP	5,850	8.62	1.26	4.96	12.41
Age Dependency Ratio	5,850	4.25	0.29	2.79	4.79

Notes: Δ=First-difference operator. All level data in natural logarithms. The variables "Per Capita GDP" and "Age Dependency Ratio" are only used as part of the robustness checks (see Supplementary Table 1).

Table 1: Summary Statistics

Panel A: Test for Cross-Sectional Dependence		
Variable	CD-Test Statistic (p-value)	Absolute Correlation
(ln) Government Size	33.12 (0.00)***	0.41
(ln) Population Size	543.24 (0.00)***	0.95

Notes: Test robust to non-stationarity and parameter heterogeneity. ****p*<0.01 (rejection of H_0 of cross-sectional independence).

Panel B: Panel Unit Root Tests		
Variable	IPS-Statistic	CADF-Statistic
Level Data		
(ln) Government Size	0.39	-1.54
(ln) Population Size	5.67	-1.74
First-Differenced Data		
Δ (ln) Government Size	-55.17***	-2.58***
Δ (ln) Population Size	-5.81***	-2.28***

Notes: Δ=First-difference operator. All panel unit root tests include country-specific constants as deterministic components. IPS test: lag order chosen by Akaike information criterion (AIC) and cross-sectional averages from the series subtracted to account for cross-sectional dependence. CADF test: lag order p=4 chosen according to rule of thumb $p=int(T^{1/3})$. To eliminate the cross-sectional dependence, standard ADF regressions are augmented with the cross-section averages of lagged levels and first-differences of the individual series. ****p*<0.01 (rejection of H_0 of non-stationarity).

Panel C: Panel Cointegration Test	
Test Variant	VR-Statistic
V1	-6.48***
V2	-2.43***
V3	-6.59***
V4	-2.52***

Notes: H_a for V1 and V2: All panels are cointegrated. H_a for V3 and V4: Some panels are cointegrated. V2 and V4 include secular time trend. All test variants include panel means as deterministic components and subtract cross-sectional averages from to account for cross-sectional dependence ****p*<0.01 (rejection of H_0 of no cointegration).

Table 2: Tests for Cross-Sectional Dependence, Panel Unit Roots and Cointegration

	(1)	(2)	(3)	(4)	(5)	(6)
Econometric Method →	POLS	POLS	FE	FE	POLS	FE
ln(Population Size)	-0.058	-0.056	-0.253	-0.404		
	(0.016)***	(0.016)***	(0.077)***	(0.117)***		
Δ ln(Population Size)					-0.267	-0.222
					(0.175)	(0.242)
Year-Fixed Effects	No	Yes	No	Yes	Yes	Yes
Number of Observations	5,850	5,850	5,850	5,850	5,720	5,720
Root MSE	0.488	0.484	0.336	0.329	0.137	0.136
CADF-statistic	-1.43	-1.27	-1.27	-1.25	-2.52	2.63
(*p*-value)	(1.00)	(1.00)	(1.00)	(1.00)	(0.00)***	(0.00)***
CD-statistic	32.16	-0.10	40.43	-2.62	2.01	2.00
(*p*-value)	(0.00)***	(0.92)	(0.00)***	(0.00)***	(0.04)**	(0.04)**

Notes: Dependent variable=ln(Government Size) in models (1) to (4) and Δ ln(Government Size) in models (5) and (6) Constant not reported. POLS=Pooled OLS estimation. FE=Fixed-effects estimation. Δ=First-difference operator. Cluster-robust standard errors in parentheses for models (1) to (4). Driscoll-Kraay standard errors in parentheses for models (5) and (6). $*p<0.1$, $**p<0.5$, $***p<0.01$.

Table 3: Pooled OLS and Fixed-Effects Estimates

	(1)	(2)	(3)	(4)	(5)
Method →	Pooled CCE	MG-CCE	MG-CCE	MG-CCE	MG-CCE
Short-Run Estimates					
Δ ln(Population Size)	1.042	14.198	3.150	17.031	9.321
	(5.671)	(8.190)*	(31.731)	(31.598)	(13.004)
Δ ln(Population Size) $_{t-1}$	-1.379	-10.051	7.205	31.996	-11.917
	(5.671)	(7.074)	(45.968)	(56.296)	(30.734)
Δ ln(Population Size) $_{t-2}$			-13.330	31.523	12.907
			(26.024)	(51.757)	(31.576)
Δ ln(Population Size) $_{t-3}$				-20.339	-8.279
				(22.393)	(13.433)
Δ ln(Government Size) $_{t-1}$	-0.032	0.172	0.207	0.338	0.159
	(0.087)	(0.021)***	(0.029)***	(0.044)***	(0.028)***
Δ ln(Government Size) $_{t-2}$			0.075	0.169	0.076
			(0.023)***	(0.035)***	(0.021)***
Δ ln(Government Size) $_{t-3}$				0.132	0.043
				(0.024)***	(0.015)***
Long-Run Estimates					
ln(Population Size) $_{t-1}$	-0.308	2.127	2.910	3.404	1.414
	(0.147)	(0.765)***	(1.017)***	(1.245)***	(0.718)**
ln(Government Size) $_{t-1}$	-0.311	-0.630	-0.770	-0.979	-0.666
	(0.147)**	(0.032)***	(0.043)***	(0.061)***	(0.042)***
Long-Run Average Coefficient					
ln(Population Size)		3.376	3.778	3.475	2.213
		(1.208)***	(1.337)***	(1.274)***	(1.076)**
Number of Lags of Cross-Sectional Averages	3	3	3	3	3
Number of Observations	5,330	5,330	5,330	5,330	5,330
Root MSE	0.144	0.088	0.082	0.074	0.086
CADF-statistic	-3.06	-3.46	-3.30	-3.33	-3.21
(*p*-value)	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***
CD-statistic	0.39	1.33	0.49	-0.05	1.20
(*p*-value)	(0.69)	(0.19)	(0.63)	(0.96)	(0.23)

Notes: Dependent variable=Δ ln(Government Size). Constant not reported. MG=Mean-group. CCE=Common correlated effects. Model (5) allows for heterogeneous lag order; i.e., for each panel member and each variable, the largest lag in first-differences is dropped from the regressions if it is insignificant (at the 10%-level). Standard errors (constructed following Pesaran and Smith, 1995) in parentheses. *p<0.1, **p<0.5, ***p<0.01.

Table 4: Common Correlated Effects Error-Correction Estimates

Panel A: Government Size ⇏ Population Size (Long-Run Effect)				
Lag Order	LR Estimate GOV $_{t-1}$	GOV LR Average Coefficient	CADF-Statistic (p-value)	CD-Statistic (p-value)
1	0.001 (0.001)	0.012 (0.013)	-4.05 (0.00)***	1.25 (0.21)
2	0.001 (0.002)	0.008 (0.012)	-4.08 (0.00)***	-1.67 (0.11)
3	-0.001 (0.001)	-0.003 (0.025)	-3.51 (0.00)***	1.32 (0.19)
Panel B: Population Size ⇏ Government Size (Long-Run Effect)				
Lag Order	LR Estimate POP $_{t-1}$	POP LR Average Coefficient	CADF-Statistic (p-value)	CD-Statistic (p-value)
1	2.127 (0.765)***	3.376 (1.208)***	-3.46 (0.00)***	1.33 (0.19)
2	2.910 (1.017)***	3.778 (1.337)***	-3.30 (0.00)***	0.49 (0.63)
3	3.404 (1.245)***	3.475 (1.274)***	-3.33 (0.00)***	-0.05 (0.96)

Notes: Lag order=Number of lags of dependent and independent variable in short-run part of respective model. Short-run results not reported. POP=Population size. GOV=Government size. LR=Long-run. Standard errors (constructed following Pesaran and Smith, 1995) in parentheses. *p<0.1, **p<0.5, ***p<0.01.

Table 5: Weak Exogeneity Tests

	(1)	(2)	(3)	(4)
Short-Run Estimates				
Δ ln(Population Size)	17.229	1.938	-40.502	58.449
	(10.715)	(22.924)	(54.674)	(34.177)*
Δ ln(Population Size) $_{t-1}$	-34.617	8.780	62.836	-100.961
	(29.002)	(52.586)	(79.675)	(80.549)
Δ ln(Population Size) $_{t-2}$	41.181	-12.873	-36.459	82.531
	(25.108)	(51.244)	(56.370)	(83.464)
Δ ln(Population Size) $_{t-3}$	-23.085	5.222	-10.515	-32.220
	(18.555)	(19.306)	(21.442)	(37.325)
Δ ln(Government Size) $_{t-1}$	0.167	0.152	0.267	0.369
	(0.039)***	(0.039)***	(0.063)***	(0.061)***
Δ ln(Government Size) $_{t-2}$	0.085	0.067	0.152	0.170
	(0.031)***	(0.027)**	(0.053)***	(0.042)***
Δ ln(Government Size) $_{t-3}$	0.053	0.035	0.114	0.135
	(0.021)**	(0.021)*	(0.038)***	(0.028)***
Long-Run Estimates				
ln(Population Size) $_{t-1}$	2.245	0.658	3.527	2.515
	(0.970)**	(1.047)	(1.957)*	(1.215)**
ln(Government Size) $_{t-1}$	-0.601	-0.726	-0.914	-0.971
	(0.059)***	(0.059)***	(0.080)***	(0.089)***
Long-Run Average Coefficient				
ln(Population Size)	3.735	0.906	3.856	2.590
	(1.559)**	(1.448)	(2.140)*	(1.241)**
Number of Lags of Cross-Sectional Averages	3	3	2	3
Sub-Sample	POP >10 mill.	POP <10 mill.	GDP > 6,000	GDP < 6,000
Number of Countries	62	68	64	66
Number of Observations	2,542	2,788	2,624	2,706
Root MSE	0.085	0.086	0.057	0.092
CADF-statistic	-3.56	-3.47	-3.50	-3.51
(*p*-value)	(0.00)***	(0.00)***	(0.00)***	(0.00)***
CD-statistic	1.47	-0.92	0.95	0.34
(*p*-value)	(0.14)	(0.36)	(0.34)	(0.74)

Notes: Dependent variable=Δ ln(Government Size). POP=Population Size. GDP=GDP per capita. Dynamic MG-CCE estimates reported. Constant not reported. Standard errors (constructed following Pesaran and Smith, 1995) in parentheses. *p<0.1, **p<0.5, ***p<0.01.

Table 6: Sub-Sample Analysis

Appendix A. Additional Estimates

	(1)	(2)	(3)	(4)	(5)	(6)
Short-Run Lag Order →	1	1	1	2	2	2
Long-Run Estimates						
ln(Population Size) $_{t-1}$	2.623	3.541	4.084	3.121	3.458	3.972
	(0.797)***	(1.095)***	(1.141)***	(0.993)***	(1.376)**	(1.362)***
ln(GDP per capita) $_{t-1}$	-0.151		-1.153	-0.181		-0.208
	(0.049)***		(0.052)***	(0.049)***		(0.052)***
ln(Age Dependency Ratio) $_{t-1}$		1.336	1.256		0.955	0.726
		(0.673)**	(0.713)*		(0.700)	(0.740)
ln(Government Size) $_{t-1}$	-0.702	-0.735	-0.801	-0.829	-0.920	-0.962
	(0.032)***	(0.032)***	(0.032)***	(0.042)***	(0.047)***	(0.047)***
Long-Run Average Coefficient						
ln(Population Size)	3.735	4.817	5.099	3.764	3.759	4.129
	(1.133)***	(1.489)***	(1.427)***	(1.208)***	(1.500)**	(1.410)***
Number of Lags of Cross-Sectional Averages	3	3	3	3	3	3
Number of Observations	5,330	5,330	5,330	5,330	5,330	5,330
Root MSE	0.082	0.084	0.078	0.077	0.077	0.072
CADF-statistic	-3.35	-3.63	-3.64	-3.40	-3.64	-3.80
(*p*-value)	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***
CD-statistic	1.47	1.10	1.54	0.85	0.47	1.19
(*p*-value)	(0.14)	(0.27)	(0.13)	(0.40)	(0.64)	(0.23)

Notes: Dependent variable=Δ ln(Government Size). Short-run estimates not reported. Constant not reported. Mean-group. common correlated effects regression results reported. Standard errors (constructed following Pesaran and Smith, 1995) in parentheses. *p<0.1, **p<0.5, ***p<0.01.

Supplementary Table 1: Common Correlated Effects Error-Correction Estimates with Additional Covariates

Appendix B. List of Countries

Albania	Côte d'Ivoire	Jamaica	Republic of Korea
Algeria	Cyprus	Japan	Romania
Angola	D.R. of the Congo	Jordan	Rwanda
Antigua and Barbuda	Denmark	Kenya	Saint Lucia
Argentina	Djibouti	Laos	Sao Tome and Principe
Australia	Dominican Republic	Lebanon	Saudi Arabia
Austria	Ecuador	Lesotho	Senegal
Bahamas	Egypt	Liberia	Seychelles
Bahrain	El Salvador	Madagascar	Sierra Leone
Bangladesh	Equatorial Guinea	Malawi	South Africa
Barbados	Ethiopia	Malaysia	Spain
Belgium	Fiji	Mali	Sri Lanka
Belize	Finland	Mauritania	St. Vincent and the Grenadines
Benin	France	Mauritius	Sudan
Bhutan	Gabon	Mexico	Suriname
Bolivia	Gambia	Mongolia	Swaziland
Botswana	Germany	Morocco	Sweden
Brazil	Ghana	Mozambique	Switzerland
Brunei	Greece	Myanmar	Syria
Bulgaria	Grenada	Nepal	Tanzania
Burkina Faso	Guatemala	Netherlands	Thailand
Burundi	Guinea	New Zealand	Togo
Cabo Verde	Guinea-Bissau	Nicaragua	Trinidad and Tobago
Cambodia	Haiti	Niger	Tunisia
Cameroon	Honduras	Nigeria	Turkey
Canada	Hungary	Norway	Uganda
Central African Republic	Iceland	Oman	United Arab Emirates
Chad	India	Pakistan	United Kingdom
Chile	Indonesia	Paraguay	United States
China	Iran	Peru	Uruguay
Colombia	Iraq	Philippines	Venezuela
Comoros	Ireland	Poland	Viet Nam
Congo	Israel	Portugal	Zambia
Costa Rica	Italy	Qatar	Zimbabwe