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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 were incorrectly specified, not accounting for cross-sectional dependence, non-stationarity and cointegration as well as parameter heterogeneity. Using a panel time-series approach that adequately models these issues, we find that population size has a positive long-run effect on government size. This finding suggests that the detrimental effects of population size on government size (primarily due to a greater risk of social conflict) dominate its beneficial ones (primarily due to scale economies). We also show that population size increases government size especially in countries that are vulnerable to social conflict due to ethnic heterogeneity.

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1. Introduction

The size of government (i.e., the share of government spending in relation to a country's 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 growth and government size (for a review, see Bergh and Henrekson, 2011). In addition to depressing economic activity, government size may have further unfavorable effects on the social life. For instance, Bjornskov et al. (2007) show that excessive government consumption is detrimental to life satisfaction.

Given its potentially substantial socio-economic ramifications, a considerable amount of theoretical and empirical contributions has aimed at identifying the determinants of government size (for a brief review, see Shelton, 2007: 2234-2240). Here, *population size* has been named as a potentially important determinant. A priori, however, its effect on government size is unclear.

For one, there are a number of advantages that allow larger countries to potentially afford smaller governments. First, larger countries can capitalize on *scale economies* associated with the provision of public goods (Alesina, 2003). For instance, Andrews and Boyne (2009) show that (per capita) administrative costs are lower in larger local governments for a sample of English communities, indeed implying economies of scale. That is, fixed costs of public goods and increasing returns to scale favor larger countries, making it possible to allocate fewer resources (in relation to total GDP) to public spending. Second, larger 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 less on defense and security, again negatively affecting government size. Third, larger countries benefit from comparatively larger domestic markets, creating fewer incentives to engage in international trade and competition. Thus, larger countries are less exposed to the volatility and external risk openness usually carries (Alesina and Wacziarg, 1998). By contrast, more open (i.e., smaller) economies face more risk; they consequently have to use government spending in the form of social insurance to mitigate associated risks,, increasing the size of government (Rodrik, 1998).

For another, however, population size may also have effects that ultimately lead to an increase in government size. 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 congestion leads to the rationing of public goods (Oakland, 1972). Consequently, the costs of managing congestion may offset or even outweigh the advantages of size in the form of scale economies. Second, and more importantly, larger population size also predicts more *violent social conflict*. For instance, increases in population size are expected to result in more conflict by exacerbating resource scarcity and distributional conflicts as well as environmental degradation (e.g., Blattman and Miguel, 2010; Brückner, 2010). Indeed, larger population size is a strong positive predictor of civil war risk (for a review, see Blattman and Miguel, 2010) and terrorism (for a review, see Krieger and Meierrieks 2011). For instance, the instrumental-variable estimates of Brückner (2010) suggest that increases in population size lead to higher risk of civil conflict for a panel of 37 Sub-Saharan countries over the period 1981-2004. In turn, the increased risk of conflict due to large population size can be expected to increase government size. First, there may be higher public spending (compared to smaller countries) on security and the military to suppress conflict. Second, increased government spending on social policies (education, health, social security etc.) may be used to accommodate grievances that arise due to population pressures so as to counter the risk of rebellion. For instance, Krieger and Meierrieks (2010) and Taydas and Peksen (2012) show that governments can indeed buy internal peace by increasing public spending on welfare policies.

In sum, economic theory is ambiguous about the effect of population size on government size. This ambiguity is also reflected in the empirical evidence. For one, in their seminal analysis Alesina and Wacziarg (1998) find that country size is 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. By contrast, 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 is positive in a fixed-effects setting. 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 methods and datasets). Finally, Rodrik (1998) reports no statistically significant association between population size and government size.

This paper aims at adding to the diverse evidence on the government size-population size nexus. We proceed as follows. In Section 2 we investigate the relationship between government and population size using "traditional" pooled OLS and fixed-effects approaches, thus following the examples of earlier empirical studies discussed above. As a first contribution to the literature, we empirically uncover several methodological shortcomings associated with this "traditional" approach, especially with respect to the roles of cross-sectional dependence, non-stationarity and cointegration as well as slope (parameter) heterogeneity. As a second contribution to the literature, to address these shortcomings we use a novel empirical *panel time-series approach* (the common correlated effects mean-group error-correction model). The estimates from this approach are presented in Section 3 and indicate that larger population size is *positively* related to government size, suggesting that the social conflict channel (where more population size leads to larger governments) dominates other channels (e.g., the scale economies channel) through which more population size would lead to smaller governments. This finding is buttressed by a sub-sample analysis indicating that the positive effect of population size on government size especially matters to ethnically fragmented societies (in which social conflict ought to be more rampant). Section 4 concludes.

2. Pooled OLS and Fixed-Effects Regressions

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 Empirical Model and Data

As in previous empirical efforts such as Alesina and Wacziarg (1998), Ram (2009) and Jetter and Parmeter (2015), we begin our empirical analysis by considering a series of empirical models of the following form:

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

This model includes an intercept (α_0) and an idiosyncratic error term (ε). Most importantly to us, it relates an indicator of government size (*GOV*) to an indicator of population size (*POP*) for country *i* at year *t*. Here, *government size* is measured as the (logged) share of government consumption at current PPPs, while *population size* is a county's (logged) population size in millions; these measures were also used by earlier empirical studies on the nexus between government and population size (e.g., Alesina and Wacziarg, 1998; Ram, 2009; Jetter and Parmeter, 2015). Both data series are drawn from the *Penn World Table* (version 9.0) (Feenstra et al., 2015); they are log-transformed to remain comparable to these previous studies and be less affected by outliers.[1]

While we are primarily interested in the relationship between government and population size, in some specifications we also include a set of year dummies (φ), while country fixed-effects (not included in the pooled regressions) are indicated by θ. Furthermore, some specifications include a vector X with additional controls for per capita income and the age dependency ratio.[2] First, data on (logged) *real per capita income* comes from the *Penn World Table*. Consistent with Wagner's law, we expect richer countries to exhibit larger governments; for instance, richer economies are more diversified and thus require more government activity (e.g., related to the provision of regulation or infrastructure) to function properly (Shelton, 2007). Second, we control for the (logged) *age dependency ratio* (defined as the ratio of dependents, i.e., people younger than 15 or older than 64 to those aged 15-64), using data from the *World Development Indicators* (World Bank, 2016). We expect countries with a larger dependency ratio to have larger governments. For instance, Sanz and Velázquez (2007) show that aging was the main driving force of the growth of government spending in the OECD countries between 1970 and 1998, potentially due to increasing public spending on health, pensions and other forms of social welfare.

With respect to the model represented in (1), we implicitly assume that the estimates are not affected by *cross-sectional dependence,* that the residuals (ε_t) produced by (1) are *stationary* and that population size affects government size uniformly for all countries in the sample (*slope homogeneity*). As we discuss below, these assumptions may not be justified; if violations of these assumptions are not accounted for, the empirical results from model (1) may be misleading.

[1] Also, it is well-known that first-differences of log-transformed data series approximate their growth rates, facilitating the interpretation of results when first-differences are taken.

[2] We also experiment with other controls (e.g., for trade openness or urbanization). Using these controls, however, does not change the main results of our paper.

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., international migration or the diffusion of medical technology, while 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).[3]

Non-Stationarity. Variables that trend over time are often found to be non-stationary (i.e., containing a unit root). For our case, it is plausible that both population size and government size are non-stationary. For instance, global population size is obviously exhibiting a long-run positive trend for the last 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) in the context of time-series data. Here, especially significance tests on the regression coefficients are invalid in the presence of spurious regression (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 it does not actually exist. Importantly, the problem of spurious regression also matters to the panel setting (e.g., Kao, 1999).

Slope Heterogeneity. A final assumption of model (1) that can be challenged is the assumption of slope homogeneity. This assumption would suggest that $\alpha_1=a_i$ for all i in model (1). Indeed, Pesaran and Smith (1995) argue that this assumption is almost always rejected in empirical practice. Pesaran and Smith (1995) show that the incorrect assumption of slope homogeneity produces inconsistent and potentially misleading estimates of the regression coefficients; they instead argue in favor of estimation methods—as used by us in Section 3—that allow for parameter

[3] 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, as we do for some variants of model (1). However, such an approach may not be sufficient to entirely expunge the cross-sectional dependence.

heterogeneity.[4] For our case, assuming parameter homogeneity implies that population size has a uniform effect on government size over all countries considered. Clearly, this is a strong assumption. For instance, in the introduction we argued that an increase in population size could result in more social spending (i.e., larger government size) to accommodate grievances due to increased population pressure. However, there are systematic differences in preferences regarding welfare spending and redistribution (both of which are expected to increase government size) between countries (e.g., Corneo and Grüner, 2002). These differences could result in heterogeneous responses with respect to changes in population size, where countries in which redistribution is favored may expand the government more strongly.

2.2 Tests for Panel Unit Roots and Cross-Sectional Dependence

Before we run model (1), we investigate whether the main data series of interest (government and population size) are affected by non-stationarity and cross-sectional dependence. Given that these issues are not accommodated for in model (1), they are likely to be "captured"—if indeed present—in the regression residuals (i.e., the ε_t series), leading to misspecification issues and potentially incorrect inferences.

We employ two different panel unit tests to assess whether the data series are non-stationary, 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 contains unit roots (i.e., it is non-stationary) versus the alternative that (a fraction of) the series are stationary. Importantly, both tests account for cross-sectional dependence. First, the IPS test does so by demeaning the data. Second, the CADF test amends the standard augmented Dickey-Fuller regressions used to investigate non-stationarity by cross-section averages of lagged levels and first-differences of the investigated series.

As shown in Table 2, the panel unit root tests indicate that both data series are non-stationary in levels but stationary after first-differences have been taken. These findings are highly intuitive.

[4] In a fixed-effects model we allow for unobserved heterogeneity through the intercept. However, this heterogeneity is necessarily time-invariant and independent of the explanatory variables. By contrast, allowing for heterogeneity in the slope parameters also allows us to consider other (more complex) forms of heterogeneity.

First, the global population doubled between 1970 and 2014, from 3,682 to 7,349 million per PENN World Tables data. This development may be due to medical advances, advances in hygiene and other socio-economic factors that have allowed many developing countries to enter a stage of demographic transition with (relatively) low death but high birth rates (for a discussion of the concept of demographic transition, see, e.g., Kirk, 1996). Second, trends towards larger governments have also been discussed in the literature, e.g., by Peltzman (1980) and Holcombe (2005). 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 trending in government size (Holcombe, 2005).

—Table 2 here—

We also test whether the data series are affected by cross-sectional dependence. We employ 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 and parameter heterogeneity (Pesaran, 2004), both of which may also matter to the variables we examine.

As shown in Table 2, both data series are indeed affected by cross-sectional dependence, 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 migration, international economic integration and politico-economic cooperation, competition or hostilities between nation-states may play a role in this context.

2.3 Pooled OLS and Fixed-Effects Regressions Results

The pre-tests reported in Table 2 already suggest that running model (1) without considering non-stationarity and cross-sectional dependence may be inappropriate. For the sake of comparability with previous empirical efforts on the government size-population size nexus (e.g., Alesina and Wacziarg, 1998; Ram, 2009; Jetter and Parmeter, 2015), we nevertheless run the regression model characterized by equation (1) using the pooled OLS and fixed-effects estimator. The results are reported in Table 3. In short, regardless of which method is chosen and which specification is run, we find that country size exerts a negative and statistically significant effect on government size;

the estimated effects are much larger in the fixed-effects setting. These results would be consistent with the arguments of Alesina and Wacziarg (1998) and Alesina (2003) regarding various benefits of country size and their potential (negative) effects on the size of government, e.g., in the form of scale economies or lower risk from an exposure to international markets.

—Table 3 here—

However, in Table 3 we also report diagnostics related to the assumptions we discussed above (cross-sectional independence, stationary residuals and slope homogeneity). First, tests of the regression residuals for unit root presence very strongly indicate that the residuals are non-stationary.[5] As discussed above, non-stationary residuals may imply a spurious regression (e.g., Kao, 1999). Second, the majority of CD-test results indicate that the residuals are affected by cross-sectional dependence.[6] As hinted at above, this may affect the validity of inference (Sarafidis and Wansbeek, 2012). Finally, we also test whether the assumption of slope homogeneity is justified. Baltagi (1981) recommends the Roy-Zellner test over the better known Chow test to examine the null hypothesis of poolability (so that population size has a uniform effect on government size for all countries in the sample) against the alternative that pooling is not appropriate (so that parameter heterogeneity should be allowed for). As reported in Table 3, the null of slope homogeneity of the Roy-Zellner test is always strongly rejected.[7] The incorrect assumption of slope homogeneity may yield inconsistent and potentially misleading estimates of the regression coefficients (Pesaran and Smith, 1995). In sum, the regression diagnostics strongly suggest that the results reported in Table 3 may be affected by various sources of misspecification.

3. 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 are run.

[7] The Chow tests we run for the sake of robustness yield very similar results.

Given that the "traditional" approach to examine the relationship between government and population size in a pooled OLS or fixed-effects setting is likely to suffer from specification issues, producing (potentially) misleading results, in this section we rely on a modelling approach that is able to account for cross-sectional dependence and allow for slope (parameter) heterogeneity as well as to produce stationary residuals while accounting for a long-run (cointegrating) relationship between population and government size. In detail, we use a panel time-series approach of Pesaran (2006) and Chudik and Pesaran (2015), the (dynamic) *common correlated effects mean-group error-correction model*.[8]

3.1. Empirical Model

As a first step, we account for the issue of non-stationarity. As shown in Table 2, population size and government size are integrated of the same order. When this is the case, a cointegration approach is feasible (Engle and Grange, 1987). Cointegration refers to the situation when a stationary linear combination of two non-stationary variables exists; cointegration analysis allows for inferences about the long-run relationship between non-stationary variables (Engle and Grange, 1987). We incorporate the idea of cointegration into our empirical model by considering the following *error-correction model* (*ECM*):

$$\Delta \mathrm{GOV}_{it} = \alpha_0 + \rho\left(\mathrm{GOV}_{i,t-1} - \beta POP_{i,t-1}\right) + \gamma^{p} \Delta POP_{it} + \gamma^{g} \Delta GOV_{i,t-1} + \epsilon_{it} \quad (2)$$

Here, government size and population size are first-differenced (indicated by the first-difference operator Δ) to achieve stationarity; the regression coefficients associated with the first-differences allows us to evaluate the short-run dynamics of the model. 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 and allows us to examine the long-run relationship between these variables.

We can reparametrize equation (2) to:

$$\Delta \mathrm{GOV}_{it} = \pi_0 + \pi^{EC} GOV_{i,t-1} + \pi^{P} POP_{i,t-1} + \pi^{p} \Delta POP_{it} + \pi^{g} \Delta GOV_{i,t-1} + \epsilon_{it} \quad (3)$$

[8] An informative introduction to and application of this empirical method is provided by Eberharrdt and Presbitero (2015).

Here, if the regression coefficient π^{EC} is statistically significant and lies between [0; -1] (implying dynamic stability), a long-run 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} allows 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).

As a final step, we allow for parameter heterogeneity and cross-sectional dependence, arriving 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) Parameter heterogeneity is accounted for in model (4) through the mean-group approach of Pesaran and Smith (1995). That is, we allow the various parameters to be estimated to vary by country i (while they were constrained to be 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).

(ii) Combining the first and second line of equation (4) gives Pesaran's (2006) common correlated effects estimator. The terms in the second line are cross-sectional averages of all variables in the model.[9] As shown by Pesaran (2006), the inclusion of these averages can accommodate cross-sectional dependence, providing consistent estimates of the parameters in the first line of equation

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

(4) that are robust to unobserved common factors (due to spillover effects, global politico-economic shocks etc.).[10]

(ii) Estimation equation (4) includes one lag of the dependent variable; also, further lags of the dependent variable (as well as of the explanatory variable) could be added to the model. This dynamic specification is expected to affect the consistency of the common correlated effects mean-group estimates (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 performs 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).

(iv) Baltagi et al. (2000) argue that the assumption of parameter heterogeneity—even if warranted by poolability tests—may produce inferior results compared to a pooled approach. They suggest that the bias due to the incorrect assumption of parameter homogeneity needs to be weighed against the efficiency gains from pooling. This is why we also estimate equation (4) in a pooled variant described in Pesaran (2006), with cross-sectional dependence still 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.

(v) As stated above, cointegration can be assessed by examining π_i^{EC}; this parameter should be statistically significant and lie between [0; -1]. Furthermore, we report the (unweighted) average *t*-ratios of the error-correction coefficients across countries. We use these *t*-ratios to test for cointegration following Gengenbach et al. (2015); a statistically significant test-statistic would lead us to reject the null hypothesis of no panel cointegration.

(vi) Finally, estimation equation (4) could be amended by additional controls or deterministics in the short- or long-run equation. However, we generally focus on a parsimonious model that only considers population size and government size. This approach is motivated by, e.g., Lütkepohl (2007) who argues in favor of parsimony especially in the context of cointegration analysis. He

[10] 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.

argues that a cointegration relationship is robust to model extensions; that is, a cointegrating relationship between government and population size would—if present—also hold when additional variables are added to the model (Lütkepohl, 2007: 322). As a robustness check, however, we also report one specification that includes our usual controls (per capital income and the age dependency ratio).

3.2 Empirical Results

The (dynamic) common correlated effects estimation results are reported in Table 4. 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 coefficient of 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, *t*-statistics and *p*-values are calculated using the Delta method. The results for both long-run estimates are quite similar and strongly indicate that population size exerts (on average) a *long-run positive effect* on government size. Considering our theoretical considerations, this suggests that population factors that promote government size (congestion and social conflict) are more important than factors that would negatively affect government size (scale economies, deterrence and trade/insurance effects). Notably, our finding is in stark contrast to our earlier findings using 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, they 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. First, we are almost always able to reject the CD-test null hypothesis of cross-sectional dependence. That is, introducing (lags of) cross-sectional averages can capture unobservables and account for cross-sectional dependence, as argued by Pesaran (2006) and Chudik and Pesaran (2015). Second, the regressions residuals are always found to be stationary. In addition to that, the cointegration tests always indicate that a cointegrating relationship exists, with the cointegrating relationship being dynamically stable. Third, a mean-group (heterogeneous) modelling approach yields a smaller RMSE compared to a pooled (homogeneous) approach. Following Baltagi et al. (2000), this suggests that the mean-group approach is preferred over a pooled one.[11]

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.

3.3 Sub-Sample Analysis

In this subsection we want to assess whether the results of Table 4 also hold for specific sub-samples that differ with respect to certain characteristics. For our analysis we consider three such characteristics which we discuss below in more detail:

(i) We investigate the role of *ethnic fractionalization*, using data by Alesina et al. (2003). The fractionalization index of Alesina et al. (2003) reflects the probability that two randomly selected individuals from a population belong to different ethnic groups. Higher levels of ethnic fractionalization may result in more pronounced ethnic cleavages and a greater potential for social conflict, as also argued in, e.g., Alesina et al. (2003) and Esteban and Ray (2008). Thus, it may be more likely that government size increases in response to population pressures to avoid social conflict (e.g., through increased public spending on security or social welfare) when ethnic fractionalization is high. Importantly, higher levels of ethnic fractionalization are not simply a function of country size. For our sample the correlation coefficient between the ethnic

[11] This finding is also in line with the Roy-Zellner test results reported in Table 3. Note that 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 RMSE. However, we also compare all other (dynamic) MG-CCE models reported in Table 4 to their pooled counterparts; the calculated RMSE always suggest that a heterogeneous modelling approach is preferred over the homogeneous (pooled) one (results available upon request).

fractionalization index and population size is r=0.018 (p=0.16). Rose (2006) similarly reports that smaller countries are not necessarily more ethnically homogeneous.[12]

(ii) Second, we consider the *level of population size* itself. Potentially, there is a non-linear effect of population size on government size, so that the former only (positively) affects the latter when a certain threshold of government size has been reached. For instance, congestion costs (which are expected to increase with population size and stimulate government growth) may only matter for fairly large countries but be negligible up to a certain country size (Alesina, 2003). Similarly, the risk of social conflict may not linearly increase in population size, as suggested by Esteban and Roy (2008).

(iii) Finally, we consider the role of *per capita income*. For one, richer countries are less vulnerable to social conflict (Blattman and Miguel, 2010). Thus, for richer countries the social conflict channel (through which population size would fuel government growth) may be less important. For another, richer countries tend to be more open to international trade, e.g., as found in Ram (2009); increased exposure to trade may in turn create demand for higher government spending to insure against the risks of trade (Rodrik, 1998). Finally, Wagner’s law postulates that richer countries are more prone to government expansion (e.g., Shelton, 2007), where richer countries are also expected to fund public goods (e.g., culture) for which scale effects may be less important. In sum, differences in the scope and type of government activity may differ with a country’s level of economic development, potentially also influencing how population size influences government size.

We run a series of common correlated effects mean-group estimations as specified in equation (4) for a number of sub-samples, where the sub-samples differ with respect to the characteristics outlined above. To create these sub-samples, we use the interquartile mean of the ethnic fractionalization, population size and per capita income data series. Relying on this mean provides

[12] For instance, Alesina et al. (2003) report a fractionalization index for China of 0.154, while the index for many smaller African countries (e.g., Angola, Burkina Faso or Cote d’Ivoire) lies between 0.7 and 0.8.

some protection against outliers; at the same time, it allows us to split the sample into two sub-samples of roughly equal size.

The empirical results are reported in Table 5. First, the results suggest that in countries with an above-average level of ethnic fractionalization there is a positive long-run effect of population size on government size, while the same is not true for the sub-sample where ethnic fractionalization is smaller. Consistent with our expectations, this finding suggests that that government size is especially responsive to population pressures under such circumstances. For instance, this may be due to social conflict being more pronounced in highly fractionalized societies, leading to, e.g., more public spending on security or other public goods compared to less fractionalized countries. Second, 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 after a certain population threshold has been reached. Finally, we find that population size has a positive long-run effect on government size both for a sub-sample of comparatively rich and poor countries. Thus, a country's level of economic development does not seem to play an obvious role in moderating the government size-population size nexus.

—Table 5 here—

4. Conclusion

There are conflicting schools of thought regarding the effect of population size on government size. One school argues that larger countries benefit from scale economies and reduced exposure to the risks of international wars and trade and can thus afford smaller governments. Another school of thought argues that larger countries necessitate larger governments to counter congestion costs and the dangers of conflict, both of which increase with country size. Accordingly, a review of existing empirical studies shows that the evidence on the government size-population size nexus is mixed.

We re-examine this nexus using panel data for 130 countries for the 1970-2014 period. We argue that previous analyses of the effect of population size on government size are incorrectly specified models by not properly accounting for cross-sectional dependence, non-stationarity and cointegration as well as parameter heterogeneity. Using a panel time-series approach that adequately considers and models these issues, we find that population size has a positive long-run effect on government size, suggesting that the detrimental effects of population size (more congestion, greater risk of social conflict) dominate its beneficial ones. We also show that this effect is especially important to countries that are vulnerable to social conflict due to ethnic heterogeneity and that exhibit a country size exceeding 10 million inhabitants.

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, these countries cannot expect to see their government size shrink relative to their GDP as population size increases; rather, the opposite appears to be true. Given that many empirical studies suggest that (too) large governments may produce undesirable socio-economic outcomes (reduced economic growth, crowding-out of private investment etc.), the role of population size and growth in determining government size should consequently not be disregarded especially in developing countries and emerging markets.

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Tables

Variable	N*T	Mean	Standard Deviation	Minimum	Maximum
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 Income	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.

Table 1: Summary Statistics

Panel A: 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 test 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 B: 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).

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

	(1)	(2)	(3)	(4)	(5)	(6)
Method →	POLS	POLS	POLS	FE	FE	FE
ln(Population Size)	-0.058	-0.056	-0.064	-0.253	-0.404	-0.556
	(0.016)***	(0.016)***	(0.016)***	(0.077)***	(0.117)***	(0.148)***
Year-Fixed Effects	No	Yes	Yes	No	Yes	Yes
Additional Controls †	No	No	Yes	No	No	Yes
Number of Observations	5,850	5,850	5,850	5,850	5,850	5,850
Root MSE	0.488	0.484	0.479	0.336	0.329	0.322
CADF-statistic	-1.43	-1.27	-1.20	-1.27	-1.25	-1.30
(*p*-value)	(1.00)	(1.00)	(1.00)	(1.00)	(1.00)	(1.00)
CD-statistic	32.16	-0.10	0.12	40.43	-2.62	-1.90
(*p*-value)	(0.00)***	(0.92)	(0.91)	(0.00)***	(0.01)**	(0.06)*
Roy-Zellner test-statistic				4,091.02	4,223.99	24,234.23
(*p*-value)				(0.00)***	(0.00)***	(0.00)***

Notes: Dependent variable=ln(Government Size). Constant not reported. POLS=Pooled OLS estimation. FE=Fixed-effects estimation. Cluster-robust standard errors in parentheses. *p<0.1, **p<0.5, ***p<0.01.

†: Additional controls are per capita income and age dependency ratio (both logged).

Table 3: Pooled OLS and Fixed-Effects Estimates

	(1)	(2)	(2)	(3)	(4)	(5)	(6)
Method →	Pooled-CCE	MG-CCE	MG-CCE	Dynamic MG-CCE	Dynamic MG-CCE	Dynamic MG-CCE	Dynamic MG-CCE
Short-Run Estimates							
Δ ln(Population Size)	-0.568	2.310	4.172	14.198	15.680	17.031	11.456
	(0.581)	(2.641)	(4.212)	(8.190)*	(7.452)**	(31.598)	(21.624)
Δ ln(Population Size) $_{t-1}$				-10.051	-7.455	31.996	-2.347
				(7.074)	(7.467)	(56.296)	(46.822)
Δ ln(Population Size) $_{t-2}$						31.523	-4.599
						(51.757)	(44.648)
Δ ln(Population Size) $_{t-3}$						-20.339	3.165
						(22.393)	(19.338)
Δ ln(Government Size) $_{t-1}$				0.172	0.185	0.338	0.268
				(0.021)***	(0.019)***	(0.044)***	(0.033)***
Δ ln(Government Size) $_{t-2}$						0.169	0.141
						(0.035)***	(0.027)***
Δ ln(Government Size) $_{t-3}$						0.132	0.110
						(0.024)***	(0.019)***
Long-Run Estimates							
ln(Population Size) $_{t-1}$	0.034	0.505	1.372	2.127	4.085	3.404	1.873
	(0.049)	(0.278)*	(0.527)***	(0.765)***	(1.141)***	(1.245)***	(1.127)*
Ln(Government Size) $_{t-1}$	-0.177	-0.381	-0.472	-0.630	-0.801	-0.979	-0.945
	(0.034)***	(0.021)***	(0.023)***	(0.032)***	(0.032)***	(0.061)***	(0.047)***
$\bar{t}$-Statistic		-2.832	-2.837	-3.107	-3.698	-2.905	-3.358
(p-value)		(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***
Long-Run Average Coefficient							
ln(Population Size)		1.324	2.904	3.376	5.099	3.475	1.983
		(0.726)*	(1.116)***	(1.208)***	(1.427)***	(1.274)***	(1.200)*
Additional Controls †	No	No	No	No	Yes	No	Yes
Number of Lags of Cross-Sectional Averages	0	0	3	3	3	3	1
Number of Observations	5,720	5,720	5,330	5,330	5,330	5,330	5,330
Root MSE	0.129	0.107	0.095	0.088	0.078	0.074	0.077
CADF-statistic	-2.75	-3.15	-3.38	-3.46	-3.64	-3.33	-3.558
(*p*-value)	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***
CD-statistic	1.03	1.75	1.62	1.33	1.54	-0.05	1.64
(*p*-value)	(0.30)	(0.08)*	(0.11)	(0.19)	(0.13)	(0.96)	(0.11)

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

†: Additional controls are per capita income and age dependency ratio (both logged).

Table 4: Common Correlated Effects Mean-Group Error-Correction Estimates

	(1)	(2)	(3)	(4)	(5)	(6)
Short-Run Estimates						
Δ ln(Population Size)	52.894	-37.333	17.229	1.938	-40.502	58.449
	(29.732)*	(41.413)	(10.715)	(22.924)	(54.674)	(34.177)*
Δ ln(Population Size) $_{t-1}$	-87.806	61.639	-34.617	8.780	62.836	-100.961
	(70.484)	(76.486)	(29.002)	(52.586)	(79.675)	(80.549)
Δ ln(Population Size) $_{t-2}$	74.383	-37.904	41.181	-12.873	-36.459	82.531
	(75.295)	(65.605)	(25.108)	(51.244)	(56.370)	(83.464)
Δ ln(Population Size) $_{t-3}$	-34.192	-2.591	-23.085	5.222	-10.515	-32.220
	(35.506)	(26.374)	(18.555)	(19.306)	(21.442)	(37.325)
Δ ln(Government Size) $_{t-1}$	0.319	0.268	0.167	0.152	0.267	0.369
	(0.051)***	(0.049)***	(0.039)***	(0.039)***	(0.063)***	(0.061)***
Δ ln(Government Size) $_{t-2}$	0.164	0.119	0.085	0.067	0.152	0.170
	(0.043)***	(0.038)***	(0.031)***	(0.027)**	(0.053)***	(0.042)***
Δ ln(Government Size) $_{t-3}$	0.112	0.093	0.053	0.035	0.114	0.135
	(0.029)***	(0.029)***	(0.021)**	(0.021)*	(0.038)***	(0.028)***
Long-Run Estimates						
ln(Population Size) $_{t-1}$	2.743	1.461	2.245	0.658	3.527	2.515
	(1.198)**	(1.672)	(0.970)**	(1.047)	(1.957)*	(1.215)**
Ln(Government Size) $_{t-1}$	-0.921	-0.842	-0.601	-0.726	-0.914	-0.971
	(0.089)***	(0.063)***	(0.059)***	(0.059)***	(0.080)***	(0.089)***
$\bar{t}$-Statistic	-2.986	-2.800	-3.168	-3.647	-2.862	-3.017
(p-value)	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***
Long-Run Average Coefficient						
ln(Population Size)	2.977	1.735	3.735	0.906	3.856	2.590
	(1.261)**	(1.997)	(1.559)**	(1.448)	(2.140)*	(1.241)**
Number of Lags of Cross-Sectional Averages	2	2	3	3	2	3
Sub-Sample	EF >0.46	EF <0.46	POP >10 mill.	POP <10 mill.	GDP > 6,000	GDP < 6,000
Number of Countries	63	67	62	68	64	66
Number of Observations	2,747	2,583	2,542	2,788	2,624	2,706
Root MSE	0.086	0.071	0.085	0.086	0.057	0.092
CADF-statistic	-3.18	-3.33	-3.56	-3.47	-3.50	-3.51
(*p*-value)	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***
CD-statistic	1.13	0.76	1.47	-0.92	0.95	0.34
(*p*-value)	(0.26)	(0.45)	(0.14)	(0.36)	(0.34)	(0.74)

Notes: Dependent variable=Δ ln(Government Size). EF=Ethnic Fractionalization. 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 5: Sub-Sample Analysis

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

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