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Abstract

This paper offers an evaluation of the output contribution of infrastructure. Drawing from a large data set of infrastructure stocks covering 88 countries and spanning the years 1960-2000, and using a panel time-series approach, the paper estimates a long-run aggregate production function relating GDP to human capital, physical capital, and a synthetic measure of infrastructure given by the first principal component of infrastructure endowments in transport, power and telecommunications. Tests of the cointegration rank allowing it to vary across countries reveal a common rank with a single cointegrating vector, which we interpret as the long-run production function. Estimation of its parameters is performed using the pooled mean group (PMG) estimator, which allows for unrestricted short-run parameter heterogeneity across countries while imposing the (testable) restriction of long-run parameter homogeneity. The long-run elasticity of output with respect to the synthetic infrastructure index ranges between 0.07 and 0.10. The estimates are highly significant, both statistically and economically, and robust to alternative dynamic specifications and infrastructure measures. There is little evidence of long-run parameter heterogeneity across countries, whether heterogeneity is unconditional, or conditional on their level of development, population size, or infrastructure endowments.

JEL Classification: H54, E23, O40.

Keywords: Infrastructure, panel cointegration, parameter heterogeneity.

Resumen

Este artículo ofrece una evaluación de la contribución al PIB de la infraestructura. A partir de un conjunto de datos de infraestructura que abarca 88 países durante los años 1960-2000, y utilizando un enfoque de panel de series de tiempo, el trabajo estima una función de producción agregada de largo plazo con el PIB, capital humano, capital físico, y un medida sintética de la infraestructura que viene dada por el primer componente principal de la dotación de infraestructura en transporte, energía y telecomunicaciones. Contrastes de cointegración (que permiten rangos de cointegración heterogéneos entre países) revelan un rango común con un solo vector de cointegración, que interpretamos como la función de producción a largo plazo. La estimación de sus parámetros se realiza con el estimador "Pooled Mean Group (PMG)", que permite heterogeneidad en los parámetros de corto plazo al mismo tiempo que impone la restricción (contrastable) de homogeneidad en los parámetros de largo plazo. La elasticidad de largo plazo de la producción con respecto al índice sintético de infraestructura varía entre 0,07 y 0,10. Las estimaciones son muy significativas, tanto estadística como económicamente, y robustas alternativas especificaciones dinámicas y medidas de infraestructura. Hay poca evidencia de la heterogeneidad de los parámetros a largo plazo entre los países, tanto si la heterogeneidad es incondicional como condicional al nivel de desarrollo, el tamaño de la población, o las dotaciones de infraestructura.

Códigos JEL: H54, E23, O40.

Palabras clave: Infraestructura, cointegración en paneles, heterogeneidad en parámetros.

1 Introduction

The macroeconomic literature has long been interested in the contribution of infrastructure capital to aggregate productivity and output. Numerous theoretical papers have approached it using an aggregate production function including public capital as an additional input, first in the context of Ramsey-type exogenous growth models [e.g., Arrow and Kurz (1970)] and later in endogenous growth models [Barro (1990), Futagami, Morita and Shibata (1993)]. This analytical literature has grown enormously in the last fifteen years, exploring a multitude of variants of the basic models, such as alternative financing schemes, simultaneous consideration of public capital and productive current spending flows, utility-yielding public capital, or public infrastructure congestion.[1]

Quantitative assessments of the contribution of infrastructure are critical for many policy questions —such as the output effects of fiscal policy shocks instrumented through public investment changes [e.g., Leeper, Walker and Yang (2010); Ilzetzki, Mendoza and Végh (2010)], or the extent to which public infrastructure investment can be self-financing [Perotti (2004)]. The empirical literature offering such quantitative assessments took off with the seminal work of Aschauer (1989) on the effects of public infrastructure capital on U.S. total factor productivity. The literature has boomed over the last two decades, with dozens of papers using a large variety of data and empirical methodologies, and with widely contrasting empirical results.[2] For example, Bom and Ligthart (2008) report that in a large set of empirical studies using industrial-country data in a production function setting, estimates of the output elasticity of public capital range from -0.175 to +0.917.

However, much of the empirical literature on the contribution of infrastructure to aggregate output is subject to major caveats. Studies based on time-series have often ignored the non-stationarity of aggregate output and infrastructure capital, which typically display stochastic trends. This has sometimes led to implausibly high estimates of the productivity of infrastructure, owing to spurious correlation between both variables [Gramlich (1994)].[3] In addition, empirical studies also have to deal with potential simultaneity between infrastructure and income levels. For example, richer or faster-growing countries are likely to devote increased resources to infrastructure development. Failing to control for these and similar forms of reverse causality implies that estimates of the output elasticity of infrastructure may be confounded with the income elasticity of the demand for infrastructure services, and hence may suffer from upward biases.[4] Finally, studies using cross-section or panel macroeconomic data typically fail to account for the potential heterogeneity in the output elasticity of infrastructure across countries or states, which could arise from

1. See for example Turnovsky (1997); Glomm and Ravikumar (1997); Baier and Glomm (2001), and Ghosh and Roy (2004).

2. See for example Sánchez-Robles (1998); Canning (1999); Demetriades and Mamuneas (2000); Röller and Waverman (2001); Esfahani and Ramírez (2003); Calderón and Servén (2004). A recent overview of relevant empirical literature is provided by Romp and de Haan (2007).

3. One example is Aschauer's original estimate of the output elasticity of public capital, which was so high that the implied marginal product of infrastructure capital was close to 100% per year.

4. A way out of this problem is to use a full structural model in the empirical estimation. In this vein, some empirical studies have used stripped-down versions of Barro's (1990) framework [e.g., Canning and Pedroni (2008)]. An alternative is to use some kind of instrumental variable approach, ideally featuring outside instruments for infrastructure. For example, Calderón and Servén (2004) employ demographic variables as instruments —alone or in combination with internal instruments— in a GMM panel framework. Roller and Waverman (2001) follow a similar approach.

technological features such as network effects, scale economies and other factors that may affect the output elasticity of public capital.[5]

This paper estimates the contribution of infrastructure to aggregate output using a production function framework including as inputs infrastructure assets, human capital, and non-infrastructure physical capital. We use a large cross-panel dataset comprising 88 countries and over 3,500 country-year observations, drawn from countries with very different levels of income and infrastructure endowments.

One distinguishing feature of our approach is that, in contrast with much of the earlier literature, we use physical measures of infrastructure rather than monetary ones —such as a public investment flow or its accumulation into a public capital stock. We do this for two reasons. First, as an abundant literature has argued, public expenditure can offer a very misleading proxy for the trends in the public capital stock, as the link between spending and capital is mediated by the extent of inefficiency and corruption surrounding project selection and government procurement practices, which can vary greatly across countries and over time [e.g., Pritchett (2000); Keefer and Knack (2007)]. Second, our interest here is infrastructure capital, rather than broader public capital, and in many countries the two can be very different owing to the involvement of the public sector in non-infrastructure industrial and commercial activities (a common occurrence in virtually all countries over a good part of our sample period), and due also to the increasing participation of the private sector in infrastructure industries worldwide, especially since the 1990s.

The paper's approach allows us to tackle some of the main methodological problems of earlier literature and extend it in several dimensions. First, we take account of the multidimensionality of infrastructure[6] and, in contrast with the abundant literature that measures infrastructure in terms of an investment flow or its cumulative stock, or in terms of a single physical asset (such as telephone density), we consider three different types of core infrastructure assets —in power, transport and telecommunications— summarized into a synthetic infrastructure index, constructed through a principal-component procedure. Second, we use a panel cointegration approach to deal with the non-stationarity of the variables of interest and avoid the 'spurious regression' problem of much of the earlier time-series literature. Third, to address concerns with identification and reverse causality, we establish that only one cointegrating relation exists among the variables, and that this applies to all the countries in the panel. We interpret this relationship as the aggregate production function, and verify that our infrastructure index and the other productive inputs are exogenous with respect to its parameters —the parameters of interest in our context, in the terminology of Engle, Hendry, and Richard (1983). We estimate these parameters using the Pooled Mean Group estimator of Pesaran, Shin and Smith (1999), which allows for unrestricted cross-sectional heterogeneity of the short-run dynamics while imposing homogeneity of the long-run parameters. Fourth, we deal explicitly with potential cross-country heterogeneity of the (long-run) parameters of the production function through individual and joint Hausman tests of parameter homogeneity, as well as through additional experiments that let the output elasticity of infrastructure vary with selected country characteristics.

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5. In this vein, Gregoriou and Ghosh (2009) estimate the growth effects of public expenditure in a panel setting, and find that they exhibit considerable heterogeneity across countries.

6. Canning (1999) also considers the multidimensionality of infrastructure using three different physical measures; however, reverse causality issues are not addressed and the single cointegration rank hypothesis is imposed and not tested.

Our estimates of the output elasticity of infrastructure lie in the range of 0.07 to 0.10.[7] Moreover, our estimates are very precise, and robust to the use of alternative econometric specifications and alternative synthetic measures of infrastructure. Likewise, the estimated elasticities of the other inputs— human and non-infrastructure physical capital —are in line with those found in the empirical macroeconomic literature [Bernanke and Gurkaynak (2001); Gollin (2002)]. They are also highly significant and robust to the various experiments we perform.

We also find little evidence of heterogeneity across countries in the output elasticities of the inputs of the aggregate production function. Specifically, the output elasticity of infrastructure does not seem to vary with countries' level of per capita income, their infrastructure endowment, or the size of their population.

The rest of the paper is organized as follows: Section 2 describes the dataset. Section 3 lays out the methodological approach. Section 4 describes the empirical results. Finally, section 5 concludes.

7. This is very close to the value that emerges from the meta-study by Bom and Ligthart (2008) of the output elasticity of public capital. After adjusting for publication bias, they place the output elasticity of public capital at 0.086.

2 Data

Our goal is to estimate the contribution of infrastructure capital to output in a large panel data set using an infrastructure-augmented aggregate production function framework, in which aggregate output is produced using non-infrastructure physical capital, human capital, and infrastructure. The data set is a balanced panel comprising annual information on output, physical capital, human capital, and infrastructure capital for 88 industrial and developing countries over the period 1960-2000, thus, totaling 3,520 observations. The Appendix lists the sample countries used in the analysis.[8]

Real output is measured by real GDP in 2000 PPP US dollars from the Penn World Tables 6.2 [Heston, Summers and Aten (2006)]. The data on physical capital was constructed using the perpetual inventory method. To implement it, the initial level of the capital stock was estimated using data on the capital stock and real output from PWT 5.6 for those countries for which such data is available. We extrapolate the data for countries without capital stock information in PWT 5.6 by running a cross-sectional regression of the initial capital-output ratio on (log) real GDP per worker.[9]

As a robustness check, we also constructed an alternative 'back-cast' projection of capital per worker. Specifically, to construct an initial capital stock for the year 1960, we assume a zero capital stock in the distant past. Using the average in-sample growth rate of real investment (1960-2000), we project real investment back to 1930. Next, ignoring the capital stock that may have existed in that year, we accumulate the projected real investment forward into a capital stock series. The resulting level of the latter in 1960 is then taken as the initial capital stock, and the in-sample capital stock series is constructed accumulating observed investment.[10]

Our preferred measure is the capital stock series obtained from the PWT data. Assuming, as we do for the back-casting, that the pre-sample growth rate of real investment was equal to the average of the 1960-2000 period could be misleading, in view of the severe global shocks of the 1930s and 1940s (e.g., the Great Depression and World War II), which likely had a non-negligible reflection on the rates of growth capital stocks around the world. Nevertheless, we also report estimates using the capital stock series constructed through back-casting.

In turn, the stock of human capital is proxied by the average years of secondary schooling of the population, taken from Barro and Lee (2001). Finally, the labor input is proxied by the total labor force as reported by the World Bank's World Development Indicators.

Measuring physical infrastructure poses a challenge. Typically, the empirical literature on the output effects of infrastructure has focused on a single infrastructure sector.

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8. Country coverage is dictated by the availability of information. In particular, time coverage is limited by the human capital indicator, which is not available after 2000.

9. The regression used for extrapolation is: $K/Y = -1.1257+0.2727*\ln(Y/L)$, where K is the capital stock, Y is real GDP, and L is the labor force. The depreciation rate employed in the perpetual inventory calculations is 6%.

10. The rationale behind this calculation is that the assumed level of the capital stock in 1930 has only a very minor effect on the capital stock that results for 1960, and hence it is immaterial whether we set the level of capital stock at zero or some other arbitrary level in 1930.

Some papers do this by design,[11] while others take a broad view of infrastructure but still employ for their empirical analysis an indicator from a single infrastructure sector.[12] In reality, 'physical infrastructure' is a multi-dimensional concept that refers to the combined availability of several individual ingredients —e.g., telecommunications, transport and energy. In general, none of these individual ingredients is likely to provide by itself an adequate measure of the overall availability of infrastructure. For instance, a country may have a very good telecommunications network and a very poor road system or a highly unreliable power supply. In such situation, the availability of telecommunications services alone would provide a misleading indicator of the status of overall physical infrastructure.

However, attempting to capture the multi-dimensionality of infrastructure by introducing a variety of infrastructure indicators as inputs in the production function also poses empirical difficulties. It could lead to an over-parameterized specification, and hence to imprecise and unreliable estimates of the contribution of the individual infrastructure indicators. In our framework this is a concern not only for the usual reasons of multicollinearity —indeed, several of the infrastructure indicators we shall use are fairly highly correlated—[13] but also because, as described below, we shall use a nonlinear procedure to estimate the parameters of the production function. In these conditions, a parsimonious specification with relatively few regressors is much more likely to result in stable estimates robust to alternative choices of initial values.

For these reasons, we follow a different strategy. We use a principal component procedure to build a synthetic index summarizing different dimensions of infrastructure.[14] We focus on three key infrastructure sectors: telecommunications, power and road transport. This choice is consistent with previous literature on the output impact of infrastructure, which has typically focused on one of these individual sectors, most often telecommunications. The synthetic infrastructure index is the first principal component of three variables measuring the availability of infrastructure services in these three sectors. Specifically, the variables underlying the index are:

(a) *Telecommunications*: Number of main telephone lines, taken from the International Telecommunications Union's World Telecommunications Development Report CD-ROM. As a robustness check, we also experiment with an alternative measure, namely the total number of lines (main lines and mobile phones), from the same source.

11. For example, Röller and Waverman (2001) evaluate the growth impact of telecommunications infrastructure, and Fernald (1999) analyzes the productivity effects of changes in road infrastructure.

12. In the empirical growth literature, for example, the number of telephone lines per capita is usually taken as the preferred indicator of overall infrastructure availability; see for example Easterly (2001) and Loayza, Fajnzylber and Calderón (2005).

13. For instance, in our panel data set the full-sample correlation between the total number of phone lines (main and mobile) and overall power generation capacity is 0.92, while the correlation between total road length and overall power generation capacity is 0.65, and that between road length and main telephone lines is 0.61. In turn, the correlation between paved (as opposed to total) road length and power generation capacity is 0.83, while that between paved road length and main telephone lines is 0.84.

14. A similar approach is employed by Alesina and Perotti (1996) in their analysis of investment determinants, and by Sánchez-Robles (1998) to assess the growth effects of infrastructure.

(b) *Electric Power*: Power generation capacity (in Megawatts), collected from the United Nations' *Energy Statistics*, the United Nations' *Statistical Yearbook*, and the U.S. Energy Information Agency's *International Energy Annual*.[15]

(c) *Roads*: Total length of the road network (in kilometers), obtained from the International Road Federation's World Road Statistics, and complemented with information from national statistical agencies and corresponding national ministries.[16] To conduct robustness checks, we use two alternative measures of transport infrastructure: the length of the paved road network, collected from the same sources, and the combined total length of the road and railway network. The railway information is obtained from the World Bank's Railways Database and complemented with data from national sources.[17]

As we shall impose constant returns to scale in the estimations (see below), the three variables underlying the index (phone lines, power generation capacity and the length of the road network) are measured in per-worker terms, and expressed in logs.[18] Their first principal component accounts for 82 percent of their overall variance and, as expected, it is highly correlated with each of the three individual variables.[19] More specifically, the correlation between the first principal component and main telephone lines per worker is 0.96, its correlation with power generation capacity is 0.97, and its correlation with the total length of the road network is 0.74. In addition, all three (log-standardized) variables enter the first principal component with approximately similar weights:

$$z_{it} = 0.364 \cdot \ln\left(\frac{Z_1}{L}\right)_{it} + 0.354 \cdot \ln\left(\frac{Z_2}{L}\right)_{it} + 0.282 \cdot \ln\left(\frac{Z_3}{L}\right)_{it}$$

where z is the synthetic infrastructure index, (Z_1/L) is the number of main telephone lines (per 1,000 workers), (Z_2/L) is the power generation capacity (in GW per 1,000 workers), and (Z_3/L) represents the total length of the road network (in km per 1,000 workers).

Table 1 shows descriptive statistics for output, physical and human capital, and the various infrastructure indicators, for the cross-section corresponding to the year 2000. Output and the capital stock are expressed in PPP US dollars at international 2000 prices, while infrastructure variables are expressed in per worker terms.

15. The International Energy Annual (IEA) is the Energy Information Administration's main report of international energy statistics, with annual information on petroleum, natural gas, coal and electricity beginning in the year 1980. See webpage: http://www.eia.doe.gov/iea/.

16. One caveat regarding these data, as noted by Canning (1999), is that they may exhibit significant variations in quality. In particular, they do not reflect the width of the roads nor their condition.

17. The railways database can be found at http://go.worldbank.org/13EP3YJVV0.

18. Before applying principal component analysis, the underlying variables are standardized in order to abstract from units of measurement.

19. As shown below, the individual infrastructure measures display stochastic trends, and hence we obtain the first principal component by computing the weights from the (stationary) first-differenced series.

3 Econometric Methodology

The core of our empirical analysis consists in estimating the following production function:

$$y_{it} = \mu_i + \omega_t + \alpha k_{it} + \beta h_{it} + \gamma z_{it} + \varepsilon_{it} \qquad (1)$$

where *y* denotes real output, *k* and *h* represent physical capital and human capital, respectively, and *z* denotes the infrastructure capital. All variables (except human capital) are expressed in log per worker terms [e.g., k_{it} = ln(K_{it}/L_{it}) where L_{it} represents the workforce] and, in keeping with the majority of earlier literature, constant returns to scale have been imposed. The subscripts *i* and *t* index countries and years, respectively; μ_i and ω_t capture country-specific and time-specific productivity factors, and ε_{it} is a random disturbance that will be assumed uncorrelated across countries and over time. [20]

3.1 Panel unit root testing

Empirical assessments of the output contribution of infrastructure using time-series data have often failed to deal adequately with the non-stationarity of the variables. Here we address the issue using panel unit root and cointegration tests developed in the recent panel time-series literature. Unlike the traditional panel literature, which deals with samples in which the cross-sectional dimension *N* is large but the time dimension *T* is small, the panel time-series literature is concerned with situations in which both *N* and *T* are of the same order of magnitude; see Breitung and Pesaran (2008) for an overview. As a considerable literature has shown, the panel time-series (or multivariate) approach to integration and cointegration testing yields higher test power than separate, conventional tests for each unit in the panel; see, for example, Levin, Lin and Chu (2002).

The first step is to test for the stationarity of the variables under consideration, namely, output, the stocks of physical and human capital, and the composite index of infrastructure capital, all (except for human capital) measured in logs per worker. As a preliminary step, we remove the cross-sectional means from the data, to render the disturbances cross-sectionally independent. To test for the presence of a unit root in each panel series, we employ the unit root test of Im, Pesaran and Shin (2003) *(IPS),* which allows for heterogeneous short-run dynamics for different cross-sectional units. Specifically, the testing procedure averages the individual unit root test statistics.[21] The basic regression framework is the following:

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20. As noted by Canning (1999), infrastructure appears twice in (1): first as *z*, and then as part of overall physical capital *k*. Hence the total elasticity of output with respect to infrastructure capital can be approximated as $\eta \approx \gamma + \lambda\alpha$, where λ is the share of infrastructure in the overall physical capital stock. Evaluation of λ requires data on the price of infrastructure, which are not widely available. Nevertheless, calculations based on the data of Canning and Bennathan (2000) suggest that it is a small number. For the countries in our sample with available data, its median is 0.08, and its standard deviation is 0.05.

21. If the data are statistically independent across countries, under the null we can regard the average t-value as the average of independent random draws from a distribution with known expected value and variance (that is, those for a non-stationary series). This provides a much more powerful test of the unit root hypothesis than the usual single time series test. In particular, this panel unit root test can have high power even when a small fraction of the individual series is stationary. In this context, Karlsson and Lothgren (2000) find that the power of the IPS test increases monotonically with: (i) the number *N* of cross-sectional units in the panel; (ii) the time dimension *T* of each individual cross-sectional unit, and (iii) the proportion of stationary series in the panel.

$$y_{it} = \rho_i y_{i,t-1} + \sum_{k=1}^{p_i} \phi_{ik} \Delta y_{i,t-k} + \varsigma_i + \xi_{it} \tag{2}$$

with the null hypothesis of non-stationarity H$_0$: ρ_i=1, for all *i*, and the alternative H$_1$: ρ_i<1, for some *i*. The test is based on a $\bar{t}$ -statistic, defined as the average of the individual *ADF-t* statistics, $\bar{t} = \frac{1}{N}\sum_{i=1}^{N} t(\rho_i)$, where $t(\rho_i)$ is the individual *t*-statistic for testing the null hypothesis in equation (2). The critical values are tabulated by Im, Pesaran and Shin (2003).[22]

3.2 Panel cointegration testing

If the null of a unit root fails to be rejected, we next proceed to test for cointegration among the variables of interest. Several tests have been proposed in the literature for this purpose —e.g., McCoskey and Kao (1998), Kao (1999), and Pedroni (2004). However, all these tests simply evaluate the presence of cointegration and do not account for the potential existence of more than one cointegrating relationship. To assess the cointegration rank, we follow the approach of Larsson and Lyhagen (2000). Assume that the *p*-dimensional vector $Y_{it} = (y_{it}, k_{it}, h_{it}, z_{it})'$ for country *I* (where *p*=4 in our case and *I = 1,…,N)* has an error correction model (ECM) representation (if the Granger representation theorem holds). We first test the hypothesis that each of the *N* countries in the panel has at most *r* cointegrating relationships among the *p* variables. In other words, we test the null $H_0: \Upsilon_i = r_i \leq r$ for all *i=1,…,N*, against the alternative $H_a: \Upsilon_i = p$ for all *i=1,…,N*, where Υ_i is the number of cointegrating relationships present in the data for country *i*. To conduct this test, Larsson and Lyhagen (2000) define the *LR-bar* statistic as the average of the *N* individual trace statistics of Johansen (1995),

$$\Phi_{\overline{LR}}\left(H(r)/H(p)\right) = \frac{\sqrt{N}\left(\overline{LR}_{NT}(H(r)/H(p)) - E(W)\right)}{\sqrt{Var(W)}}$$

where $\overline{LR}_{NT}(H(r) \mid H(p))$ is the average of the individual trace statistics and *E(W)* and *Var(W)* are the mean and the variance of the variable *W*, whose asymptotic distribution is the same as that of the individual trace statistic. The cointegrating rank suggested by the testing procedure based on the standardized *LR-bar* statistic equals the maximum of the N individual ranks. Further, under fairly general conditions, Larsson et al. (2001) show that the standardized LR-bar statistic for the panel cointegration rank test is asymptotically distributed as a standard normal.

Since the time dimension of our sample is too short for the asymptotic properties of the individual trace statistics to hold, in our empirical application we use Reimers' (1992) small-sample correction —*i.e.* we multiply the individual trace statistics by $\left(T - Lp\right)/T$,

22. It has been shown that the empirical size of the IPS test is fairly close to its nominal size when *N* is small, and that is has the most stable size among the various panel unit root tests available [Choi (2001)]. However, when linear time trends are included in the model, the power of the test declines considerably [Breitung (2000); Choi (2001)].

where L is the lag length used to construct the underlying VAR, and p is the total number of variables.[23]

Next, we follow Larsson and Lyhagen (2000) and test for the smallest cointegration rank in the panel using the panel version of the principal component test developed by Harris (1997). Specifically, we test the null hypothesis $H_0 : \Upsilon_i = r_i = r$ against the alternative $H_a : \Upsilon_i = r_i < r$. Thus, this hypothesis is the opposite of that used for the *LR-bar* test in the sense that the alternative is that there are more than r cointegrating vectors. For this purpose, Larsson and Lyhagen (2000) developed the standardized *PC-bar* statistic:

$$\overline{PC_r} = \frac{\sqrt{N}\left(\bar{c}_r - E(W)\right)}{\sqrt{Var(W)}}$$

where $\bar{c}_r$ is the mean of the individual test statistics $\hat{c}_r = T^{-2}\sum_{t=1}^{T}\hat{S}_t'\hat{\Omega}_{nn}^{*-1}\hat{S}_t$, *E(W)* and *Var(W)* are the mean and the variance respectively of the variable *W,* whose asymptotic distribution is the same as that of $\hat{c}_r$, and we have defined $\hat{S}_t = \sum_{j=1}^{t}\hat{n}_j$ and $\hat{\Omega}_{nn}^{*} = \sum_{j=-T+1}^{T-1}k\left(\frac{j}{m}\right)\hat{\Gamma}_{ab}(j)$, where a and b are any time series, *k(.)* is a lag window, m is the bandwidth parameter and $\hat{\Gamma}_{ab}(j) = T^{-1}\sum_{t=1}^{T}a_{t-j}b_t$. In large samples, the PC-bar test follows a standard normal distribution. If at least one of the individual ranks is less than the hypothesized value, the test asymptotically rejects the null. Hence, the *PC-bar* statistic gives the minimum cointegration rank amongst all the cross-sectional units.

In short, we first use the *LR-bar* test to estimate the maximum number of cointegration relations, and then we use the *PC-bar* test to assess if for any country the number of cointegrating relations is less than the maximum given by the LR-bar test. If in the second step the null hypothesis cannot be rejected, the conclusion is that the number of cointegrating relations is the same for all cross-sectional units.

3.3 Heterogeneous panel data techniques

As we report below, the panel cointegration tests indicate a common unit cointegration rank among GDP, physical capital, human capital and the composite index of infrastructure. We interpret the single cointegration vector (whose parameters may vary across countries) as a long-run production function. To estimate its coefficients, we adopt a single-equation approach. If there were more than one cointegrating relation, single-equation estimation would only determine a suitable combination of the various cointegrating relations. However, in the presence of a single cointegrating vector, Johansen (1992) shows that if the equations of the marginal model have no cointegration, the single-equation estimator is equivalent to the estimator resulting from system estimation of all the equations.[24]

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23. As done for the panel unit root tests, we remove the cross-sectional means from the data prior to implementing the panel cointegration tests.

24. The single-equation analysis could be inefficient under certain circumstances [see Johansen (1992)].

To estimate the coefficients in equation (1), we use the pooled mean group *(PMG)* estimator developed by Pesaran, Smith, and Shin (1999).[25] In practical terms, we embed the production function equation (1) into an ARDL*(p,q)* model:

$$\Delta y_i = \phi_i \left(y_{i,-1} - F_{i,-1}\theta \right) + \sum_{h=1}^{p-1} \lambda_{i,h} \Delta y_{i,-h} + \sum_{h=1}^{q-1} \delta'_{i,h} \Delta F_{i,-h} + \mu_i \iota + \varepsilon_i \tag{3}$$

where $i = 1,...,N$ denotes the cross section units, and we impose homogeneity of the long-run coefficients $\theta_i = \theta \ \forall i$. Here $y_i = \left(y_{i1},..., y_{iT} \right)'$ is the *T* x *1* vector that contains the *T* observations of GDP for unit *I* in the panel, $F_i = (k_i, h_i, z_i)$ is the *T* x *3* matrix of inputs (physical capital, human capital and infrastructure), and ϕ_i are the coefficients that measure the speed of adjustment towards the long-run equilibrium. Also, ι is a *T* x *1* vector of ones and μ_i represents a country-specific fixed effect. The disturbances $\varepsilon_i = \left(\varepsilon_{i1},..., \varepsilon_{iT} \right)'$ are assumed to be independently distributed across *I* and *t*, with zero means and country-specific variances $\sigma_i^2 > 0$. As before, all the variables are cross-sectionally de-meaned prior to estimation in order to remove common factors, as required by the assumption of cross-sectional independence.

As equation (3) makes explicit, the *PMG* estimator restricts the long-run coefficients to be equal over the cross-section, but allows for the short-run coefficients, speed of adjustment and error variances to differ across cross-sectional units.[26] We therefore obtain pooled long-run coefficients and heterogeneous short-run dynamics. Thus, the *PMG* estimator provides an intermediate case between full parameter homogeneity, as imposed by the dynamic fixed effects estimator, and unrestricted heterogeneity, as allowed by the mean group *(MG)* estimator of Pesaran and Smith (1995), based on separate time-series estimation for each cross-sectional unit.

Estimation of the long-run coefficients in (4) is based on the concentrated log-likelihood function under normality. The pooled maximum-likelihood estimator of the long-run parameters is computed using an iterative non-linear procedure. Once the long run parameters have been computed, both the short-run and the error-correction coefficients can be consistently estimated running individual OLS regressions of Δy_i on $y_{i,-1} - F_{i,-1}\theta$.

To test the validity of the long-run parameter homogeneity restrictions, we use Hausman tests of the difference between MG and PMG estimates of the long-run coefficients. These are preferable to likelihood ratio tests owing to the 'large N' setting, which would cause the number of parameter restrictions to be tested by the likelihood ratio test to rise with sample size.

As described, the empirical strategy adopted in the paper is based on the single-equation estimation of the only cointegrating vector present in the data, which we interpreted as the aggregate production function. Our estimates of the output elasticity of physical capital, human capital and infrastructure, and the associated inference,

25. This estimator has been previously implemented in different contexts, for example Cameron and Muellbauer (2001) analyze the relationship between earnings, unemployment and housing in a panel of UK regions, while Égert et al. (2006) consider exchange rates, productivity and net foreign assets in a panel of countries.

26. In the context of country-level production functions, it seems reasonable to allow for heterogeneity of the short-run-dynamics due to, for instance, differences in adjustment costs across countries.

are obtained from an equation describing the time path of GDP per worker, with the time path of the three inputs determined by some unspecified marginal model. For this approach to be valid, the inputs have to be weakly exogenous —in the sense of Engle, Hendry, and Richard (1983)— for the parameters of the cointegrating relation. In this context, weak exogeneity means that changes in the inputs (e.g., infrastructure) do not react to deviations from the long-run equilibrium, although each input may still react to lagged changes of both GDP per worker and the other inputs of the production function. If changes in the inputs did react to deviations from the estimated long-run equation, the implication is that the single equation used in the analysis could be capturing the demand for physical capital, human capital or infrastructure rather than the production function —or a combination of both.

The requirement that the inputs be weakly exogenous can be verified through a standard variable-addition test. Specifically, as shown by Johansen (1992) and Boswijk (1995), weak exogeneity for the long-run parameters can be checked by testing the significance of the cointegrating vector in a reduced-form regression of each input on its own past and those of output and the other inputs of the production function.

Formally, weak exogeneity of the inputs amounts to the requirement that the δ coefficients not be significantly different from zero in the following system of equations [Johansen (1992)]:

$$\begin{aligned}
\Delta k_{it} &= \beta_{ki}\Delta f_{it-1} + \gamma_{ki}\Delta f_{it-2} + \lambda_{ki}^{1}\Delta y_{it-1} + \lambda_{ki}^{2}\Delta y_{it-2} + \delta_{ki}\xi_{it}(\hat{\theta}) + v_{it}^{k} \\
\Delta h_{it} &= \beta_{hi}\Delta f_{it-1} + \gamma_{hi}\Delta f_{it-2} + \lambda_{hi}^{1}\Delta y_{it-1} + \lambda_{hi}^{2}\Delta y_{it-2} + \delta_{hi}\xi_{it}(\hat{\theta}) + v_{it}^{h} \\
\Delta z_{it} &= \beta_{zi}\Delta f_{it-1} + \gamma_{zi}\Delta f_{it-2} + \lambda_{zi}^{1}\Delta y_{it-1} + \lambda_{zi}^{2}\Delta y_{it-2} + \delta_{zi}\xi_{it}(\hat{\theta}) + v_{it}^{z}
\end{aligned} \tag{4}$$

where $\xi_{it}(\hat{\theta}) = y_{i,t-1} - f_{i,t-1}\hat{\theta}$ is the estimated long-run equilibrium error term from equation (4) above, and $\Delta f_{it-1} = (\Delta k_{it-1}, \Delta h_{it-1}, \Delta z_{it-1})$. We estimate the system of equations country by country using the SURE estimator proposed by Zellner (1962). Once we have all the country-specific SURE estimates we compute the Mean Group (MG) estimator [Pesaran and Smith (1995)] and we carry out a Wald test under the null that the three coefficients on the added error terms are jointly zero. If soothe null is not rejected, we can conclude that the three inputs are weakly exogenous with respect to the parameters of the cointegrating relation.

4 Empirical results

Our empirical implementation starts by checking the order of integration of the different variables and testing for the existence of cointegration among them. Then we turn to the estimation of the parameters of the cointegrating relation(s).

4.1 Integration and cointegration

Table 2 reports the panel integration and cointegration tests. Panel A in the table shows the results of applying the panel unit root test of Im, Pesaran and Shin (2003) to each of the model's variables. In every case, the test statistic lies well below the 5% critical level, thus failing to reject the null of a unit root. Individual tests for each country (not shown) yield a similar verdict – they fail to reject the null in the overwhelming majority of cases. After taking first differences, however, the panel test (not reported) rejects the null of nostationarity for each of the variables. From this we conclude that all the variables are I(1).

We next test for cointegration. A battery of residual-based panel tests (not reported to save space), whose alternative hypotheses variously include homogeneous and heterogeneous cointegration [Kao (1999); Pedroni (1995 and 1999)], strongly support the view that the model's variables are cointegrated. However, as already noted, these tests are uninformative about the number of cointegrating relations, and with more than two I(1) variables under consideration, the possibility of multiple cointegration vectors cannot be ruled out. Further, in a panel context the possibility that different cross-sectional units may have different orders of cointegration cannot be dismissed either.

To assess the cointegration rank, we turn to the LR-bar test of Larsson, Lyhagen and Lothgren (2001) and the panel version of the PC-bar test of Harris (1997) proposed by Larsson and Lyhagen (2000). As already mentioned, we proceed in two stages. We first use the LR-bar test to establish the maximum cointegrating rank —i.e., the maximum number of cointegrating relations present in any of the panel's cross-sectional units (countries). We then use the panel version of the PC-bar test to establish the minimum cointegrating rank.

As Panel B of Table 2 reports, the LR-bar test overwhelmingly rejects the null that the maximum rank is zero (the test statistic of 9.03 is far above the 5% critical value of 1.96), but cannot reject a maximum rank of one. In turn, panel C shows that the PC-bar test cannot reject a minimum cointegrating rank of one – the computed test statistic of 1.21 is well below the critical 5% value of 1.96.

Since the maximum cointegration rank from the LR-bar test and the minimum cointegration rank from the PC-bar test coincide, the null hypothesis of a common cointegrating rank for all countries in the panel cannot be rejected. Hence, taken together the test results imply that, for each of the sample countries, there exists one single cointegrating vector among the four variables in our study, which we shall interpret as the infrastructure-augmented production function.

4.2 Estimation results

The next step is to estimate the parameters of the single cointegrating vector, which in principle might differ across countries. As discussed, we opt for the PMG estimator of Pesaran, Shin and Smith (1999), which estimates the coefficients of the long-run relation along with those characterizing the short-term dynamics. We use Hausman specification tests to assess the validity of the homogeneity restrictions imposed by PMG on the long-run parameters.

Table 3 reports a variety of PMG estimates of the long-run parameters, using alternative dynamic specifications —*i.e.*, different orders of the ARDL formulation of the equation of interest— and including time dummies to account for common factors (columns 1-5) or excluding them (column 6). The first thing to note is that, with the exception of the last column in the table, the parameter estimates in the different columns are very similar to each other. They are also very precisely estimated, which is unsurprising given the large number of observations (over 3,500) and the relatively parsimonious model employed.

In the first column, the order of the ARDL specification is determined (separately for each country) using the Schwarz criterion, subject to a maximum of two lags for both the dependent and independent variables. The estimated coefficient of the capital stock is 0.34, very close to the values commonly encountered in the empirical macroeconomic literature. The coefficient of the human capital variable is 0.10, likewise in the range of previous estimates in the literature, while that of the synthetic infrastructure index equals 0.08.[27] All three estimates are significantly different from zero at the 1 percent level. Further, the Hausman tests of parameter homogeneity, reported also in the table, show little evidence of cross-country heterogeneity of any of the individual parameters (all the p-values exceed 0.20). The same applies to the Hausman test of the joint null of homogeneity of all parameters, reported at the bottom of the table, whose p-value equals 0.44. The second column of Table 3 uses the Akaike information criterion rather than the Schwarz criterion to determine lag length, still subject to a 2-lag maximum. As is customary, this choice leads to somewhat more generous lag specifications, with a majority of countries selecting longer lag lengths than under the Schwarz criterion. However, it causes little change in the size or significance of the parameter estimates in the table, and it does not affect the Hausman tests of parameter homogeneity.

In turn, column 3 of the table imposes equal lag length (two lags) for all variables and countries, instead of allowing it to be determined by information criteria. Relative to columns 1 and 2, this results in a further loss of degrees of freedom, to an extent that depends on the country-specific number of lags that were being selected by the information criteria, and a slight deterioration of the precision of the estimates. However, it is of little consequence for the values of the estimates, their overall significance, or the verdict of the Hausman tests.

We next assess the effect of alternative choices of maximum lag length, using again the Schwarz criterion. Column 4 restricts the maximum lag length to 1. Relative to column 1, this adds 88 observations to the estimation sample, but is otherwise of little consequence for the coefficient estimates, their precision, and the Hausman tests. Column 5 summarizes the opposite exercise, raising maximum lag length to 4. This leads to the loss of 176 observations relative to column 1, but again there is no material change in any of the results.

27. The estimated coefficient of human capital ranges from 0.08-0.085 in Bloom et al. (2004), to 0.11-0.13 in Temple (1998), and 0.06-0.11 in Miller et al. (2002). On the other hand, our estimated coefficient of infrastructure capital is similar to those reported by Le and Suruga (2005) (0.076), Eisner (1991) (0.077), Duffy-Deno et al. (1991) (0.081) and Mas et al. (1996) (0.086).

Lastly, column 6 in Table 3 examines the role of common factors by re-estimating the specification in the first column omitting the time dummies. This does cause major changes in the parameter estimates: the coefficient of the capital stock rises above 0.40, and that of the infrastructure synthetic index becomes negative and insignificant. This confirms the importance of taking into account common factors (i.e., GDP and productivity shocks correlated across countries) in the estimation.

In Table 4 we explore the robustness of the results to the use of alternative measures of infrastructure and the capital stock. In all cases we employ the Schwarz criterion with a maximum lag length of 2 to select the dynamic specification. For ease of comparison, column 1 just reproduces the results from the first column of Table 3. In column 2, we replace the indicator of telephone density underlying the synthetic infrastructure index, using total phone lines (fixed plus mobile) instead of main lines, which is the variable conventionally employed in the growth literature. We recalculate the synthetic index as the first principal component of total phone lines, roads, and power generation capacity, all expressed in log per worker terms. This causes fairly modest changes in the estimates: the human capital parameter falls from 0.10 to 0.07, and overall precision declines somewhat, but there is little change in the infrastructure coefficient estimate and the results of the homogeneity tests.

Column 3 replaces road density with the density of land transport lines, including both roads and railways. As before, this leads to a new synthetic infrastructure index, but the estimation results obtained with it are virtually identical to those in the first column. In column 4, we use a narrower measure of roads, namely paved roads. The only noticeable change concerns the estimated coefficient of the human capital variable, which declines by half, while its standard error doubles. However, there is virtually no change in the other estimates. In turn, the Hausman tests now show some borderline evidence against the cross-country homogeneity of the coefficient of the human capital stock. Column 5 presents the results obtained replacing the principal-component index with an average index of infrastructure in which all the three infrastructure variables (roads, phone lines and electricity generating capacity) receive the same weight. Compared with our baseline specification in Column 1, the use of the average index causes practically no changes in the estimates.

Lastly, column 6 assesses the robustness of the results to the use of an alternative capital stock series, constructed through the back-casting method described earlier. Once again, the estimation results —including remarkably the coefficient on the capital stock itself— are virtually indistinguishable from those in the first column of the table, although now there is some indication of cross-country heterogeneity of the capital stock coefficient.

Overall, these experiments suggest that the parameter estimates of the infrastructure-augmented production function are fairly robust to alternative specifications concerning the short-run dynamics as well as the precise choice of explanatory variables. Moreover, the experiments also reveal little evidence of cross-country heterogeneity in the output elasticity of infrastructure.

However, the tests reported so far are concerned with unconditional heterogeneity, and it might be possible to gain power by testing for more specific forms of parameter heterogeneity. For example, it could be argued that, owing to network effects, the elasticity of output with respect to infrastructure should be higher in countries with larger infrastructure endowments than the rest.[28] Alternatively, the elasticity could vary with the level of development —as captured for example by GDP per worker— reflecting the fact that poorer countries are less able to use infrastructure effectively. As another hypothesis, the output elasticity of infrastructure could depend negatively on the size of the overall population, owing to congestion effects.

To verify this, we re-estimate the model in column 1 of Table 3 without imposing homogeneity across countries of the long-run parameter of the infrastructure synthetic index, and then look for patterns of heterogeneity in the individual-country estimates of that parameter along the three dimensions just mentioned —GDP, infrastructure endowment (both in per worker terms), and population size.

Figure 1 plots the resulting country-specific estimates of the infrastructure long-run coefficient against each of the three variables just mentioned. While there are some obvious outliers, the conclusion from all three graphs is clear: there is no relationship between the country-specific coefficient estimates and the three variables considered. This points to the absence of cross-country heterogeneity of the output elasticity of infrastructure along any of these dimensions.

Table 5 presents the results of more formal tests of parameter heterogeneity along these dimensions, using the country-specific estimates of the output elasticity of infrastructure obtained above. The first two columns of the table test if the output contribution of infrastructure varies across countries with their respective level of income per worker. We divide the country-specific estimates into two groups, one consisting of countries with high income and the other of countries with low income. In column 1, the groups are drawn using the World Bank's list of 'high income' countries; the low-income group is made up by all other countries in the sample. In column 2, the grouping is based instead on the sample median income per worker in the year 2000. In each case, the table reports the simple average of the parameter estimates of each of the two groups, along with the *p*-value of the test of difference in group means. In both columns, the mean estimate is slightly larger in the low-income group, but the difference is small and statistically insignificant.

Column 3 of Table 5 reports a similar experiment distinguishing between countries with high and low infrastructure endowments, again defined by the sample median. The mean estimates of the infrastructure elasticity in the two groups are numerically and statistically very similar. Thus, there is little evidence that the output elasticity of infrastructure varies systematically with the degree of infrastructure development. Lastly, column 4 defines the two country groups according to country size, as given by population, with the sample median as the relevant dividing line. The output contribution of infrastructure might be expected to be larger in countries with smaller population, owing to congestion effects. The pattern of the mean estimates of the two groups seems to accord with this view: the mean estimate is much higher for small countries than for large ones (where it is actually negative and close to zero). However, the difference between the two falls well short of statistical significance.

In summary, our results indicate that the elasticity of GDP per worker with respect to the synthetic infrastructure index is around 0.08. This finding is robust to alternative econometric specifications and alternative definitions of the synthetic infrastructure index,

28. This is similar in spirit to Gregoriou and Ghosh (2009), who let the growth contribution of public expenditure vary with countries' average level of expenditure.

as well as alternative measures of the capital stock. In addition, we find little evidence of heterogeneity of such elasticity across countries. Further experiments also suggest that the output contribution of infrastructure does not vary with countries' population, their level of income, or their infrastructure endowment. In light of our empirical specification, this suggests that cross-country variation in the marginal productivity of infrastructure is solely driven by variation in the ratio of infrastructure to output. In other words, the marginal product of infrastructure is higher wherever the (relative) infrastructure stock is lower.

Finally, we turn to the test of weak exogeneity of the production inputs described in section 3. The Wald test statistic computed from estimation of the system of equations (4) equals 5.97. Under the null of weak exogeneity of the three inputs, it follows a chi-square distribution with 3 degrees of freedom, and hence the test yields a p-value of 0.12, failing to reject the null.[29] Therefore we conclude that physical capital, human capital and infrastructure are weakly exogenous for the parameters of the cointegrating vector. This supports our interpretation that we are in fact estimating the production function instead of, for instance, an infrastructure demand equation, or a combination of both relations.

Our estimates of the output contribution of infrastructure are significant not only statistically, but also economically. To illustrate this, consider an increase in the level of infrastructure provision from the cross-country median in the year 2000 (an index of -4.65, roughly similar to the value observed in Tunisia in 2000) to the 75th sample percentile in the same year (an index of -3.69). This would translate in a 7.7 percent [=0.08*(-3.69+4.65)] increase in output per worker.[30] Similar calculations show that: (a) an increase in infrastructure provision from the median level observed among lower-middle income countries (-4.67, roughly equivalent to Bolivia in the year 2000) to that of the median upper-middle income country (-4.02, Uruguay) would yield an increase in output per worker of 5.2 percent, and (b) raising the level of infrastructure provision from the value observed in the median upper-middle income country to that of the median high-income country (-2.93, corresponding to Ireland) would raise output per worker by 8.7 percent.

29. This corresponds to a baseline specification including two lags of all the variables. However, similar results were obtained with different lag specifications.

30. Note that 0.08 is the estimated coefficient of the infrastructure index in regression [1] of Table 3.

5 Conclusions

This paper adds to the empirical literature on the contribution of infrastructure to aggregate output. Using an infrastructure-augmented production function approach, the paper estimates the output elasticity of infrastructure on a large cross-country panel dataset comprising over 3,500 annual observations. The paper addresses several limitations of earlier literature. It uses a multi-dimensional concept of infrastructure, combining power, transport and telecommunications infrastructure into a synthetic index constructed through a principal component procedure. The econometric approach deals explicitly with the non-stationarity of infrastructure and other productive inputs, reverse causality from output to infrastructure, and potential cross-country heterogeneity in the contribution of infrastructure (or any other input) to aggregate output.

The empirical strategy involves the estimation of a production function relating output per worker to non-infrastructure physical capital, human capital, and infrastructure inputs. Our estimates, based on heterogeneous panel time-series techniques, place the output elasticity of infrastructure in a range between 0.07 and 0.10, depending on the precise specification employed. The estimates are highly significant and robust to a variety of experiments involving alternative econometric specifications and different synthetic measures of infrastructure. Some illustrative calculations show that the output contribution of infrastructure implied by these results is also economically significant. Moreover, our estimates of the output contribution of human capital and non-infrastructure physical capital are likewise significant and broadly in line with those reported by earlier literature.

Lastly, tests of parameter homogeneity reveal little evidence that the output elasticity of infrastructure varies across countries. This is so regardless of whether heterogeneity is unconditional, or conditional on the level of development, the level of infrastructure endowments, or the size of the overall population. The implication is that, across countries, observed differences in the ratio of aggregate infrastructure to output offer a useful guide to the differences in the marginal productivity of infrastructure.

REFERENCES

ALESINA, A., and R. PEROTTI (1996). "Income Distribution, Political Instability, and Investment", *European Economic Review, 40,* pp. 1202-1229.

ARROW, K. J., and M. KURZ (1970). *Public investment, the rate of return and optimal fiscal policy,* Baltimore, MD: The Johns Hopkins University Press.

ASCHAUER, D. (1989). "Is Public Expenditure Productive?", *Journal of Monetary Economics,* 23, pp. 177-200.

BAIER, S. L., and G. GLOMM (2001). "Long-run growth and welfare effects of public policies with distortionary taxation", *Journal of Economic Dynamics and Control,* 25, pp. 2007-2042.

BARRO, R. J. (1990). "Government spending in a simple model of exogenous growth", *Journal of Political Economy,* 98, pp. 103-125.

BARRO, R. J., and J. W. LEE (2001). "International Data on Educational Attainment: Updates and Implications", *Oxford Economic Papers,* 53 (3), pp. 541-563.

BERNANKE, B. S., and R. S. GÜRKAYNAK (2002). "Is Growth Exogenous? Taking Mankiw, Romer, and Weil Seriously", in B. S. Bernanke and K. S. Rogoff (eds.), *NBER Macroeconomics Annual 2001,* Cambridge, MA: National Bureau of Economic Research, pp. 11-72.

BLOOM, D. E., D. CANNING and J. SEVILLA (2004). "The Effect of Health on Economic Growth: A Production Function Approach", *World Development,* 32 (1), pp. 1-13.

BOM, P. R. D., and J. E. LIGTHART (2008). *How productive is public capital? A meta-analysis,* CESIfo Working Paper 2206, January.

BOSWIJK, H. (1995). "Efficient Inference on Cointegration Parameters in Structural Error Correction Models", *Journal of Econometrics,* 69, pp. 133-158.

BREITUNG, J. (2000). "The Local Power of Some Unit Root Tests for Panel Data", in B. Baltagi (ed.), *Nonstationary Panels, Panel Cointegration, and Dynamic Panels,* Advances in Econometrics, Vol. 15, JAI: Amsterdam, pp. 161-178.

BREITUNG, J., and M. H. PESARAN (2008). "Unit Roots and Cointegration in Panels", in L. Matyas and P. Sevestre (eds.), *The Econometrics of Panel Data: Fundamentals and Recent Developments in Theory and Practice,* Third Edition, Springer Publishers, Ch. 9, pp. 279-322.

CALDERÓN, C., and L. SERVÉN (2003). "The Output Cost of Latin America's Infrastructure Gap", in W. Easterly and L. Servén (eds.), *The Limits of Stabilization: Infrastructure, Public Deficits, and Growth in Latin America,* Stanford University Press and the World Bank, pp. 95-117.

— (2004). *Trends in infrastructure in Latin America,* World Bank Policy Research Working Paper 3401.

CAMERON, G., and J. MUELLBAUER (2001). "Earnings, Unemployment, and Housing in Britain", *Journal of Applied Econometrics,* 16, pp. 203-220.

CANNING, D. (1999). *The Contribution of Infrastructure to Aggregate Output,* The World Bank Policy Research Working Paper 2246.

CANNING, D., and E. BENNATHAN (2000). *The social rate of return on infrastructure investments,* The World Bank Policy Research Working Paper Series 2390, July

CANNING, D., and P. PEDRONI (2008). "Infrastructure, long-run economic growth and causality tests for cointegrated panels", *The Manchester School,* 76 (5), pp. 504-527.

CHOI, I. (2001). "Unit Root Tests for Panel Data", *Journal of International Money and Finance,* 20, pp. 249-272.

DEMETRIADES, P., and T. MAMUNEAS (2000). "Intertemporal Output and Employment Effects of Public Infrastructure Capital: Evidence from 12 OECD Economies", *The Economic Journal,* 110, pp. 687-712.

DUFFY-DENO, K. T., and R. W. EBERTS (1991). "Public Infrastructure and Regional Economic Development: A Simultaneous Equations Approach", *Journal of Urban Economics,* 30, pp. 329-343.

EASTERLY, W. (2001). "The Lost Decades: Explaining Developing Countries' Stagnation in spite of policy reform 1980-1998", *Journal of Economic Growth,* 6 (2), pp. 135-157.

ÉGERT, B., K. LOMMATZSCH and A. LAHRÉCHE-RÉVIL (2006). "Real Exchange Rate in Small Open OECD and Transition Economies: Comparing Apples with Oranges?", *Journal of Banking & Finance,* 30, pp. 3393-3406.

EISNER, R. (1991). "Infrastructure and Regional Economic Performance: Comment", *New England Economic Review,* September/October, pp. 47-58.

ENGLE, R. F., and C. W. J. GRANGER (1987). "Co-integration and Error Correction: Representation, Estimation, and Testing", *Econometrica,* 55, pp. 251-276.

ENGLE, R. F., D. F. HENDRY and J. F. RICHARD (1983). "Exogeneity", *Econometrica,* 51, pp. 277-304.

ESFAHANI, H., and M. RAMÍREZ (2003). "Institutions, Infrastructure and Economic Growth", *Journal of Development Economics,* 70 (2), pp. 443-477.

FERNALD, J. G. (1999). "Roads to Prosperity? Assessing the Link between Public Capital and Productivity", *The American Economic Review,* 89, pp. 619-638.

FUTAGAMI, K., Y. MORITA and A. SHIBATA (1993). "Dynamic analysis of an endogenous growth model with public capital", *Scandinavian Journal of Economics,* 95, pp. 607-625.

GHOSH, S., and U. ROY (2004). "Fiscal policy, long-run growth, and welfare in a stock-flow model of public goods", *Canadian Journal of Economics,* 37 (3), pp. 742-756.

GLOMM, G., and B. RAVIKUMAR (1997). "Productive government expenditures and long-run growth", *Journal of Economic Dynamics and Control,* 21 (1), pp. 183-204.

GOLLIN, D. (2002). "Getting Income Shares Right", *Journal of Political Economy,* 110 (2), pp. 458-474.

GRAMLICH, E. M. (1994). "Infrastructure Investment: A Review Essay", *Journal of Economic Literature,* 32 (3), pp. 1176-1196.

GREGORIOU, A., and S. GHOSH (2009). "On the heterogeneous impact of public capital and current spending on growth across nations", *Economics Letters,* 105 (1), pp. 32-35.

HARRIS, D. (1997). "Principal Components Analysis of Cointegrated Time Series", *Econometric Theory,* 13, pp. 529-557.

HAUSMAN, J. (1978). "Specification Tests in Econometrics", *Econometrica,* 46, pp. 1251-1271.

HESTON, A., R. SUMMERS and B. ATEN (2006). *Penn World Table Version 6.3,* Philadelphia, PA: University of Pennsylvania, Center for International Comparisons of Production, Income and Prices, September.

ILZETZKI, E., E. MENDOZA and C. VÉGH (2010). *How big (small) are fiscal multipliers?,* NBER Working Paper 16479.

IM, K. S., M. H. PESARAN and Y. SHIN (2003). "Testing for Unit Roots in Heterogenous Panels", *Journal of Econometrics,* 115, pp. 53-74.

JOHANSEN, S. (1988). "Statistical Analysis of Cointegration Vectors," *Journal of Economic Dynamics and Control*, 12, pp. 231-254.

— (1992). "Cointegration in Partial Systems and the Efficiency of single-equation Analysis", *Journal of Econometrics,* 52, pp. 389-402.

— (1995). *Likelihood-based Inference in Cointegrated Vector Autoregressive Models,* Oxford: Oxford University Press.

KAO, C. (1999). "Spurious Regression and Residual-based Tests for Cointegration in Panel Data", *Journal of Econometrics,* 90, pp. 1-44.

KARLSSON, S., and M. LÖTHGREN (2000). "On the Power and Interpretation of Panel Unit Root Tests", *Economic Letters,* 66, pp. 249-255.

KEEFER, P., and S. KNACK (2007). "Boondoggles, Rent-Seeking, and Political Checks and Balances: Public Investment under Unaccountable Governments", *Review of Economics and Statistics,* 89, 566-572.

LARSSON, R., and J. LYHAGEN (2000). *Testing for common cointegrating rank in dynamic panels, Working Paper Series in Economics and Finance,* No. 378, Stockholm School of Economics.

LARSSON, R., J. LYHAGEN and M. LOTHGREN (2001). "Likelihood-based Cointegration Tests in Heterogenous Panels", *Econometrics Journal*, 4, pp. 109-142.

LE, M. V., and T. SURUGA (2005). "Foreign Direct Investment, Public Expenditure and Economic Growth: The Empirical Evidence for the Period 1970-2001", *Applied Economics Letters,* 12, pp. 45-49.

LEEPER, E. M., T. B. WALKER and S. YANG (2010). "Government Investment and Fiscal Stimulus in the Short and Long Runs", *Journal of Monetary Economics,* forthcoming.

LEVIN, A., C. F. LIN and C. S. CHU (2002). "Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties", *Journal of Econometrics,* 108, pp. 1-24.

LOAYZA, N. V., P. FAJNZYLBER and C. CALDERÓN (2005). *Economic Growth in Latin America and the Caribbean: Stylized Facts, Explanations and Forecasts,* Washington, DC: The World Bank, Latin America and the Caribbean Studies, 156 pp.

MAS, M., J. MAUDOS, F. PÉREZ and E. URIEL (1996). "Infrastructure and Productivity in the Spanish Regions", *Regional Studies,* 30, pp. pp. 641-650.

MCCOSKEY, S., and C. KAO (1998). "A Residual-Based Test of the Null of Cointegration in Panel Data", *Econometric Reviews,* 17, pp. 57-84.

MILLER, S. M., and M. P. UPADHYAY (2002). "Total factor productivity and the convergence hypothesis", *Journal of Macroeconomics,* 24, pp. 267-286.

PEDRONI, P. L. (1995). *Panel Cointegration: Asymptotic and Finite Sample Properties of Pooled Time Series Tests, With an Application to the PPP Hypothesis,* Indiana University Working Papers in Economics No. 95-013.

— (1999). "Critical values for cointegration tests in heterogeneous panels with multiple regressors", *Oxford Bulletin of Economics and Statistics,* 61 (4), pp. 653-690.

— (2004). "Panel Cointegration: Asymptotic and Finite Sample Properties of Pooled Time Series Tests with an Application to the Purchasing Power Parity Hypothesis", *Econometric Theory,* 20, pp. 597-625.

PEROTTI, R. (2004). *Public investment: another (different) view,* IGIER Working Paper 277, December.

PESARAN, H., Y. SHIN and R. SMITH (1999). "Pooled Mean Group Estimation of Dynamic Heterogeneous Panels", *Journal of the American Statistical Association,* 94, pp. 621-634.

PESARAN, H., and R. SMITH (1995). "Estimating long-run Relationships from Dynamic Heterogeneous Panels", *Journal of Econometrics,* 68, pp. 79-113.

PRITCHETT, L. (2000). "The Tyranny of Concepts: CUDIE (Cumulated, Depreciated, Investment Effort) Is Not Capital", *Journal of Economic Growth,* 5 (4), pp. 361-384.

REIMERS, H. E. (1992). "Comparison of Tests for Multivariate Cointegration", *Statistical Papers*, 33, pp. 335-359.

RÖLLER, L.-H., and L. WAVERMAN (2001). "Telecommunications Infrastructure and Economic Development: A Simultaneous Approach", *American Economic Review,* 91, pp. 909-923.

ROMP, W., and J. DE HAAN (2007). "Public Capital and Economic Growth: A Critical Survey", *Perspektiven der Wirtschaftspolitik,* 8 (s1), pp. 6-52.

SÁNCHEZ-ROBLES, B. (1998). "Infrastructure Investment and Growth: Some Empirical Evidence", *Contemporary Economic Policy,* 16, pp. 98-108.

TEMPLE, J. (1998). "Robustness tests of the augmented Solow model", *Journal of Applied Economics,* 13 (4), pp. 361-375.

TURNOVSKY, S. J. (1997). "Fiscal policy in a growing economy with public capital", *Macroeconomic Dynamics,* 1, pp. 615-639.

ZELLNER, A. (1962). "An efficient method of estimating seemingly unrelated regressions and tests for aggregation bias", *Journal of the American Statistical Association,* 57, pp. 348-368.

Table 1.
Descriptive Statistics
Output and Inputs for the year 2000

	Mean	Std. Dev.	Min.	Max.	Unit
GDP	21536	20048	1603	75288	2000 US Dollars
Physical Capital	48539	58035	600	248032	2000 US Dollars
Physical Capital (BC)	48644	58153	597	247570	2000 US Dollars
Secondary Education	1.5882	1.1113	0.0712	4.4438	Years
Electricity	0.0017	0.0022	0.0000	0.0118	Gigawatts
Main Phone Lines	0.4561	0.4713	0.0028	1.4051	Number of lines
Cell Phones	0.4479	0.5453	0.0004	1.6927	Number of lines
Roads	0.0141	0.0163	0.0011	0.0827	Kilometers
Paved Roads	0.0079	0.0116	0.0001	0.0540	Kilometers
Rails	0.0006	0.0008	0.0000	0.0049	Kilometers

All variables are expressed in per worker terms. The basic descriptive statistics were computed over a sample of 88 countries in the year 2000. BC refers to the back-casting method of construction of the capital stock series, where the initial capital stock is computed by projecting the level of real investment into the past and assuming a negligible level of capital stock in 1930.

Table 2.
Panel Unit Root and Cointegration Tests

PANEL A: Panel Unit Root Test

Variable	Test Statistic
GDP	-6.20
Physical Capital	-7.08
Secondary Education	-1.77
Infrastructure	-3.35

5% critical value for the null hypothesis of unit root is 1.96 in all cases.

The sample covers 88 countries and the years 1960-2000

Test employed: Im, Pesaran and Shin (2003)

PANEL B: Panel LR-bar Test

Maximum rank	Test Statistic
0	9.03
1	0.85

The null hypothesis of maximum cointegration rank is sequentially tested against the alternative of maximum rank equal to p (i.e. the number of variables considered). The 5% critical value is 1.96.

The sample covers 88 countries and the years 1960-2000

Test employed: Larsson, Lyhagen and Lothgren (2001)

PANEL C: Panel PC-bar Test

Minimum rank	Test Statistic
1	1.21

Given the maximum cointegration rank tested in Panel B, the null hypothesis of minimum cointegration rank is sequentially tested against the alternative of smaller minimum cointegration rank. The 5% critical value is 1.96.

The sample covers 88 countries and the years 1960-2000

Test employed: Larsson and Lyhagen (2000)

In all tests variables are expressed in log per worker terms and common factors in the series are removed.

Table 3
Estimation of the Production Function
Alternative Dynamic Specifications

Column	(1)	(2)	(3)	(4)	(5)	(6)
Max# of lags	2	2	2	1	4	2
Information criterion	SBC	AIC	Imposed	SBC	SBC	SBC
Common factors	Yes	Yes	Yes	Yes	Yes	No
Physical Capital	**0.34**	**0.33**	**0.36**	**0.35**	**0.34**	**0.41**
t-ratio	35.2	30.5	22.7	31.4	32.4	33.4
hausman p-value	0.54	0.95	0.43	0.78	0.44	0.52
Secondary Education	**0.10**	**0.12**	**0.10**	**0.12**	**0.11**	**0.12**
t-ratio	15.6	14.8	8.09	18.7	17.1	16.0
hausman p-value	0.24	0.19	0.21	0.19	0.20	0.64
Infrastructure	**0.08**	**0.07**	**0.10**	**0.08**	**0.08**	**-0.02**
t-ratio	7.45	6.73	6.58	8.33	8.77	-1.49
hausman p-value	0.21	0.18	0.16	0.40	0.88	0.40
joint hausman p-value	0.44	0.38	0.24	0.25	0.45	0.85
Average R^2	0.36	0.40	0.48	0.28	0.42	0.35
Observations	3432	3432	3432	3520	3256	3432

Dependent variable is log GDP. All variables are expresed in log per worker terms. Infrastructure is an aggregate index of electricity generating capacity, main phone lines and roads. Country specific short run dynamics are either imposed or determined by information criteria (Schwarz (SBC) or Akaike (AIC)). For each regressor, the p-value from the test of the null of cross-country homogeneity is reported under the t-statistic of its respective coefficient estimate; the p-value from the joint test is reported at the bottom of the table.

Table 4
Estimation of the Production Function
Alternative measures of infrastructure and the capital stock

Column	(1)	(2)	(3)	(4)	(5)	(6)
Variable Changed	Base	Total Phone Lines	Roads plus Rails	Paved Roads	Average Infrastructure Index	BC Physical Capital
Physical Capital	**0.34**	**0.35**	**0.34**	**0.34**	**0.34**	**0.33**
t-ratio	35.2	32.8	35.2	26.6	35.5	18.0
hausman p-value	0.54	0.80	0.48	0.58	0.92	0.05
Secondary Education	**0.10**	**0.07**	**0.10**	**0.05**	**0.11**	**0.10**
t-ratio	15.6	6.84	15.8	3.98	16.2	10.3
hausman p-value	0.24	0.55	0.24	0.11	0.26	0.20
Infrastructure	**0.08**	**0.07**	**0.08**	**0.07**	**0.08**	**0.09**
t-ratio	7.45	5.45	7.51	5.20	7.80	5.53
hausman p-value	0.21	0.37	0.23	0.41	0.36	0.14
joint hausman p-value	0.44	0.69	0.43	0.33	0.63	0.20
Average R^2	0.36	0.37	0.36	0.35	0.37	0.35

Dependent variable is log GDP. All variables are expresed in log per worker terms. Infrastructure is an aggregate index of electricity generating capacity, main (or main plus cells) phone lines and roads (or roads plus rails or paved roads). BC refers to the Back Cast construction method of the physical capital stock. Sample size is 3432 in all columns. Country specific short run dynamics are determined by the Schwarz information criterion with a maximum # of lags of 2. For each regressor, the p-value from the test of the null of cross-country homogeneity is reported under the t-statistic of its respective coefficient estimate; the p-value from the joint test is reported at the bottom of the table.

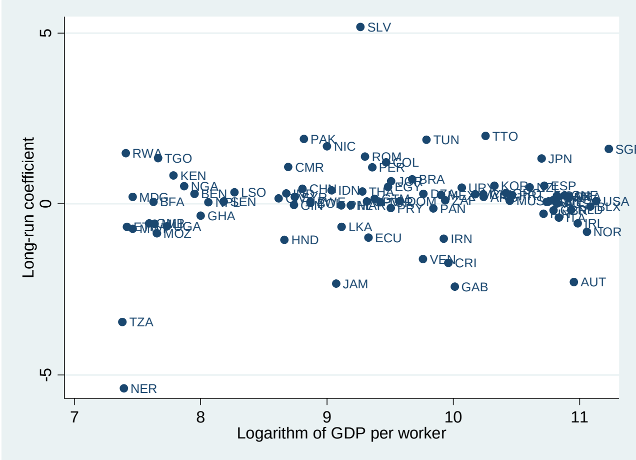
Table 5
Additional Homogeneity Tests

Column	(1)	(2)	(3)	(4)
	Per Capita Income (A)	Per Capita Income (B)	Infrastructure Endowment	Total Population
High	0.054	0.044	0.059	-0.016
Low	0.059	0.062	0.055	0.131
p-value	0.985	0.940	0.988	0.576

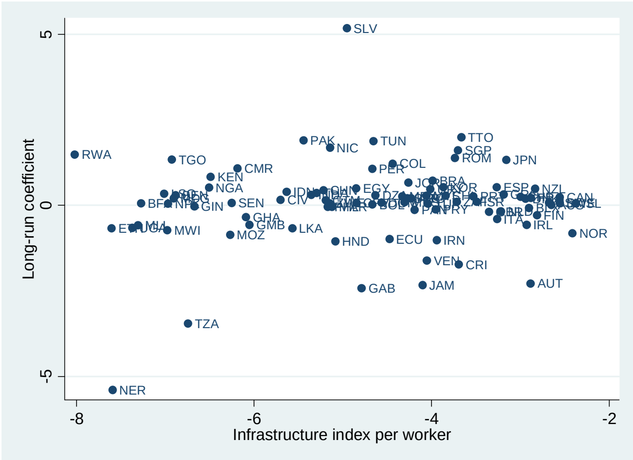
This table reports the results of tests of difference in means with unequal variances carried out by sub-groups in which heterogeneity of the effects of infrastructure might be a concern. For this purpose, country specific PMG infrastructure coefficients are estimated and group specific means are computed. The null hypothesis of equality of group specific means is tested and p-values are reported. Countries are grouped into 'high' and 'low' categories, where high refers to countries with high levels of per capita income (above the sample median per capita income (A) or according to the World Bank's definition of high income (B)), infrastructure (above the sample median) or population (above the sample median).

Figure 1
Output Elasticity of Infrastructure across countries

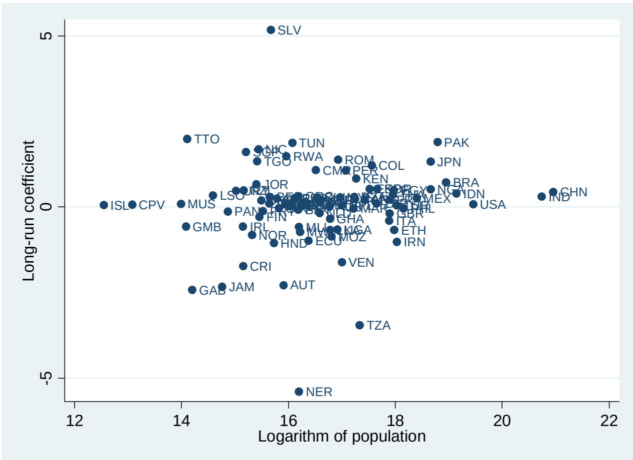
1.1 Output elasticity of infrastructure vs. the level of output



1.2 Output elasticity of infrastructure vs. aggregate infrastructure



1.3 Output elasticity of infrastructure vs. population



Appendix

Table A1: List of Countries

Panel A: Rich countries		
Australia	Iceland	Portugal
Austria	Ireland	Singapore
Belgium-Luxemburg	Israel	Spain
Canada	Italy	Sweden
Denmark	Japan	Switzerland
Finland	Netherlands	United Kingdom
France	New Zealand	United States
Greece	Norway	

Panel B: Developing Countries		
Argentina	Egypt, Arab Rep.	Ghana
Brazil	El Salvador	Guinea
Chile	Guatemala	India
Costa Rica	Honduras	Indonesia
Gabon	Iran, Islamic Rep.	Kenya
Korea, Rep.	Jamaica	Lesotho
Malaysia	Jordan	Madagascar
Mauritius	Morocco	Malawi
Mexico	Paraguay	Mali
Panama	Peru	Mozambique
South Africa	Philippines	Nepal
Trinidad and Tobago	Romania	Nicaragua
Turkey	Sri Lanka	Niger
Uruguay	Syrian Arab Republic	Nigeria
Venezuela	Thailand	Pakistan
Algeria	Tunisia	Rwanda
Bolivia	Benin	Senegal
Cape Verde	Burkina Faso	Tanzania
China	Cameroon	Togo
Colombia	Cote d'Ivoire	Uganda
Dominican Republic	Ethiopia	Zimbabwe
Ecuador	Gambia	

Rich countries are those defined as high income countries by The World Bank

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