# Public investment in transportation and communication and growth: A dynamic panel approach*

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Abstract

This paper investigates the relationship between public investment in transportation and communication and economic growth using traditional instrumental estimation and a mixed fixed and random coefficient approach in the context of a dynamic panel framework. We find that there is a dynamic effect of public investment in transportation and communication on economic growth and its impact seems to be positive and to be a long-run effect. However, for the reverse causal relationship proposed by the investment acceleration hypothesis, we find that there is significant heterogeneity across countries and thus the reverse causality does not seem to run. These results suggest that greater heterogeneity in one causal direction than in the other can be instructive point in terms of more traditional analyses.

# 1 Introduction

In the neoclassical models of Solow (1956) and Swan (1956), because the return to government expenditures or good governance is viewed not to be sufficiently large enough to prevent diminishing returns to capital and the sources of long-run growth are exogenous factors – e.g., the dynamics of population and of technological progress – any policy aimed at achieving long run persistent growth can affect the rate of growth only during the transition to steady state (i.e., only *temporary* increase in growth rate).[1] On the other hand, in endogenous growth models such as Romer (1986, 1990), Lucas (1988), Barro (1990), and Rebelo (1991), not only direct investment in physical capital and labour but also investments in knowledge and human capital, in research and development and in public infrastructure play an important role in production and with this extended concept of capital, these models have constant or increasing returns in the factors that can be accumulated and the long-run growth is determined by the parameters of the model not by technological innovations or population growth. From this point of view, the endogenous models suggest that any temporary change in economic environment is capable of generating permanent effects and this opens up the possibility of fiscal policy to have long-run effect on growth.

This view has sparked a flurry of empirical studies on the relationship between the size of government (both at aggregate and disaggregate levels) and economic growth.[2] There is, however, conflicting evidence in the literature regarding the question as to how the composition of government expenditure affects economic growth.[3] In particu-

[1]A similar result is obtained by the Arrow and Kurz (1970) model of the influence of public investment on growth, where, the public capital stock enters into the private production function but the marginal returns to public capital is not bounded away from zero.

[2]Examples include Barro (1990, 1991), Barro and Sala-i-Martin (1995, 1999), Easterly and Rebelo (1993), Cashin (1995), Deverajan, Swaroop and Zou (1996), Kocherlakota and Yi (1996, 1997), Agell, Lindh and Ohlsson (1997), Mendoza, Milesi-Ferretti, and Aseaet (1997), Miller and Ressek (1997), Kneller, Bleany and Gemmell (1999), and Gupta, Clements, Baldacci, and Mulas-Granados (2002), and Bose, Haque and Osborn (2007).

[3]Consider, for example, the association between government size (as measured either by the level of total public expenditure or by the level of public consumption expenditure) and economic growth. According to some studies, such association is significant and positive (Ram, 1986). The same association has been found to be significant and negative in other studies (e.g. Landau 1983; Grier and Tullock 1989; Alexander 1990; Barro 1990, 1991). Yet, some studies have found this associaation to be insignificant or fragile (e.g. Kormendi and Meguire 1985; Levine and Renelt 1992). A similar variation in results can also be observed among studies, which look for the growth effects of public expenditures at a much-disaggregated level.

lar, on the relationship between public investment in transportation and communication (infrastructure) and economic growth, there has been a mixed picture. In the endogenous model, the public investment in infrastructure such as roads, highways, water supply, and airports, provides private producers with inputs that are complementary to more directly productive investments and raises the private sector productivity and thus has positive effects on long-run growth rates. In this setting, private returns to scale may be diminishing, but social returns–which reflect spillovers of knowledge or other externalities–can be constant or increasing.[4] Aschauer (1989) finds that core infrastructure – streets, highways, airports, mass transit, and other public capital – has the most explanatory power for private-sector productivity in the United States over the period 1949 - 1985. In a cross-country study, Easterly and Rebelo (1993) find, using the pooled regressions, that only public investment in transportation and communication (hereafter T&C) among the sectorial components of government investment, is consistently positively correlated with growth with a very high coefficient (between 0.59 and 0.66). On the other hand, Deverajan *et al.* (1996) find, from the study of 43 developing countries over 20 years, that transport and communication expenditures have a negative correlation with per-capita real GDP growth. Miller and Russek (1997) report that the estimated coefficient for the ratio of transportation and communication expenditure to GDP is positive but not statistically significant for 23 developing countries. Yet, another recent study by Bose et al. (2007) uses 30 developing countries over two decades of 1970s and 1980s and reports that both investment and total expenditure in T&C sector have insignificant effects on growth, after controlling for government budget constraint.

Why does this previous literature provide conflicting results? A theoretical perspective and recent econometric literature on the panel data analysis for developing countries direct our attention to addressing this question. First of all, it takes time for public investment in T&C to affect growth and thus, a consideration of time is important for investigating the effect of infrastructure on growth. From this point of view, a

[4] For example, as in Barro and Sala-i-Martin (1992, 1995), all productive government expenditures are complementary to private capital and therefore can be viewed as additional inputs to firm's production and ones predict that the public investment in infrastructure raises the long-run growth by enhancing the productivity of the private sector.

dynamic model might be more desirable than a contemporaneous cross-section analysis. Secondly, as the lagged dependent variable is correlated with the error term in the dynamic panel, there is a clear simultaneity problem and the usual approach for dealing with this problem is to first-difference the data and to use instrumental approach. In practice, however, it is often difficult to find good instruments for the first-differenced lagged dependent variable, which can itself create problems for the estimation. Kiviet (1995) shows that panel data models that use instrumental variable estimation (GMM) often lead to poor finite sample efficiency and bias.[5] Thirdly, as pointed out in Pesaran and Smith (1995), Weinhold (1999), and Nair-Reichert and Weinhold (2001), when we impose an (erroneous) assumption that the coefficients on the explanatory variables (lagged dependent variables) are equal across units, there could be significant bias on all the estimated coefficients introduced if in fact the coefficients on the lagged dependent variable are not constant across the cross section and this bias cannot be addressed with instrumental variable estimation (GMM). From these points of view, a more careful investigation is warranted for addressing the relationship.

The purpose of this paper is to revisit and examine the causal relationship between public investment in T&C and economic growth, bearing the time consideration and the econometric issues in mind. In exploring this end, our study is different from the previous literature on several grounds. First of all, we consider a dynamic panel model using a new and much richer data set for 15 developing countries over 1970 to 1987. Secondly, we not only employ an instrumental variable approach, but also apply the mixed fixed and random coefficients model (hereafter MFR) of Weinhold (1999) and Nair-Reichert and Weinhold (2001) to avoid biased parameter estimates resulting from the situation where there is an assumption of homogeneity of the coefficient on the lagged dependent variable(s) but there is actually cross-sectional heterogeneity in the

[5] Kiviet (1995) exploits a formula for the bias of the Least-Squares Dummy Variable (LSDV) estimator in the case of dynamic panel data models, and shows that the bias approximation is very accurate and the bias corrected LSDV estimator is consistent not just for $T \to \infty$, but also for finite $T$ and $N \to \infty$. In the Monte Carlo simulation study, he finds that the bias corrected LSDV estimator shows hardly any bias and is not outclassed by the standard Instrumental Variables (IV) estimators (Anderson and Hsiao (1981, 1982)) and various Generalized Method of Moments (GMM) estimators (Arellano and Bond (1991)) and Ahn and Schmidt (1995), whereas most of the standard IV or GMM estimation techniques produce poor results in particular customary finite sample situations.

dynamic components . Thirdly, following the accelerating effect of output on investment as in Clark (1979) and Wagner's law (the tendency for government expenditure to be higher at higher level of per capita GDP), we examine reverse causality in which public investments in T&C follow growth and thus rapid growth leads to higher investments in this sector.[6]

Parallel to the literature on public investment (fixed capital formation) on infrastructure, a large body of empirical literature has found a robust positive relationship between fixed capital formation, as measured by investment shares of GDP, and long run growth (e.g., Levine and Renelt 1992; Mankiw et al. 1992). Some authors have gone a step further to show the direction of the causality. For example, De Long and Summers (1991, 1992) assert that faster growth is triggered by higher investment rates or higher capital formation in the form of equipment investment. In contrast with this view, more recent empirical evidence (e.g., Carroll and Weil 1994; Blomstrom et al. 1996) shows that rapid growth leads to higher fixed investment.[7]

While the literature on fixed investment includes the fixed capital formation both by private and public sector, our study focuses on the fixed capital formation in public sector only. The study on the causes and consequences of public sector investment in infrastructure (such as transport and communication) is distinctively different from the above literature and bears the significance at least for three reasons.[8] First, private and public investments differ in their characteristics that cause them to affect output (or, growth of output) of firms in a different manner. While private capital is a direct input to production, public capital acts as an indirect input that enhances the productivity of private inputs. Second, due to the public goods nature of the infrastructures such as roads, highways, airports, telecommunication, etc., private sector fails – or is unable – to provide these goods. This causes the government intervention in infrastructure. Third, while private investment is financed from the private sector savings, public sector

[6] Musgrave (1969, p.74) mentions that the most plausible formulation of Wagner's hypothesis appears to be in terms of a positive correlation between the share of government expenditure in GDP and income per capita. Also, there is a large body of literature on Wagner's law, which includes Gandhi (1971), Abizadeh and Gray (1985) and Ram (1987), Park (1996), Al-Faris (2002), among others.

[7] There are some studies that show tht causality running from both the directions (e.g., Podrecca and Carmeci 2001)

[8] See Stern (1991) for an overview of th role of public policy in economic development.

investment needs to be financed through taxation that may crowd out private investment by reducing private sector savings. Hence, as long as the productivity enhancing effect of public investment dominates its negative (distortionary) effect through taxation, public investment should raise growth.[9]

Our results confirm and extend the conclusion of earlier study that public investment in T&C Granger causes economic growth. The estimated coefficients in our study are somewhat lower, whereas the sizes of the coefficients are disturbingly high in Easterly and Rebelo (1993).[10] Furthermore, when we allow heterogeneity of dynamics in the developing country panel, the estimation results of the MFR model support a causal relation from public investment in T&C to growth more strongly than in the instrumental variable estimation. However, from both approaches, instrumental estimation and the MFR model, we do not find evidence of reverse causality. In particular, the estimation results of MFR model indicate that there is a great deal of heterogeneity across countries in the reverse causal relationship.

The plan of the paper is as follows. Section 2 review the empirical evidence on the relationship between T&C and growth. Section 3 describes a methodology used in this panel study and Section 4 describes the data and reports empirical results. Concluding remarks are offered in Section 5.

## 2 Transportation and communication and growth: a discussion on the empirical evidence

In terms of empirical literature, the relationship between public investment in infrastructure and growth do not seem to prove causality. Aschauer (1989) shows that movements in public investment bring forth movements in private-sector output which are as much as four to seven times as large as the public-sector outlays, while changes in government

[9] See Barro (1990), barro and Sala-I-Martin (1992), Glomm and Ravikumar (1994, 1997), Cashin (1995), Deverajan et al. (1996), among others for the effect of public investment on economic growth.

[10] Easterly and Rebelo (1993) use instrumental variable estimation to investigate whether reverse causation is responsible for significant effect of T&C investment on growth. The instruments used to instrument public investment variables including T&C investment are initial income, population size, share of agriculture in GDP and continent dummies.

consumption have, at best, a small positive influence on production, somewhat smaller than unity.

Easterly and Rebelo (1993) try to provide a comprehensive summary of the statistical association between measures of fiscal policy, the level of development, and the rate of growth using standard data sources–Summers and Heston (1991), the Government Financial Statistics (GFS), the International Financial Statistics (IFS)–combined with newly created data–information contained in World Bank reports–for public investment. They construct decade-average public investment ratios by sector from the available data in each decade and enter them into pooled regressions of decade-average per capita growth. In these regressions, they use a similar set of conditioning variables as Barro (1991) and extend this regression to include one public investment variable at a time. They find that transportation and communication investment seem to be consistently positively correlated with growth with a high coefficient (between 0.59 and 0.66) and it is still significant in the growth regression when they control for private investment. Since this result cannot exclude the possibility that the association between public investment and growth is due to reverse causation: public investment may simply be higher in periods of fast expansion, further they investigate whether reverse causation is responsible for the result and perform a standard instrumental estimation with instruments of initial income, population size, share of agriculture in GDP, and continent dummies for Africa and Latin America. They report that the effect of transport and communications on growth is robustly significant with instrumental variables, but the size of the coefficients is high: they obtain a coefficient of 2 for transport and communication investment and a coefficient of 0.7 for general government investment.

Devarajan et al. (1996) investigate the relationship between the composition of public expenditure and economic growth. They derive conditions under which a change in the mix of public spending could lead to a higher steady-state growth rate for the economy. The conditions depend not just on the physical productivity of different components of public spending but also on the shares of government expenditure allocated to them. In their empirical study for 43 developing countries from 1970 through 1990, they use pooled regressions with the choice of a five-year forward moving average of per-capita real GDP growth to reflect the fact that public expenditures often take time before their effects register on output growth, to eliminate short-term fluctuations induced by shifts in public expenditure, and to increase the number of time series observations in the panel data. They find that all of standard candidates for productive expenditure–capital, transport and communication, health, and education–had either a negative or insignificant relationship with economic growth and the current expenditure was associated with higher economic growth. Their results suggest that expenditures which are normally considered productive could become unproductive if there is an excessive amount of them and imply that developing-country governments have been misallocating public expenditures in favor of capital expenditures at the expense of current expenditures.

Miller and Russek (1997) examine the effects of fiscal structure on economic growth by considering a systematic way that controls for the mode of financing and by incorporating the government budget constraint into the growth regressions so that one can identify how a particular change in fiscal policy is financed. In their empirical study for 39 countries with annual data for 1975 to 1984, they estimate two sets of regression equations where the one does not disaggregate total revenue and expenditure and the other does. They find that for developing countries, debt-financed increases in government expenditure retard economic growth and tax-financed increases lead to higher growth, while for developed countries, debt-financed increases in government expenditure do not affect economic growth and tax-financed increases lead to lower growth and that debt-financed increases in transportation and communication expenditure does not affect growth, while debt-financed increases in defense, health, and social security and welfare expenditures retard growth in developing countries.

In a more recent study, Bose et al. (2007) examine the growth effects of government expenditure using the data for two decades, 1970s and 1980s, with a particular focus on disaggregated government expenditures by explicitly recognizing the role of the government budget constraint and the possible biases arising from omitted variables. They show at the disaggregated level that government investment in education and total expenditures in education are the only outlays that are significantly associated with growth once the budget constraint and omitted variables are taken into consideration, while both investment and total expenditures in T&C sector turn out to have no significant effect on growth.

Overall, the findings of these studies paint a mixed picture; the relationship between T&C and growth is sometimes significant and positive, sometimes significant and negative, and sometimes not significant. Although these studies consider many different econometric techniques, they do not examine the effect of public investment in T&C in a systematic way that incorporate a dynamic behavior of infrastructure and avoid potential bias from the situation where it is assumed that the coefficients on the lagged dependent variables are equal across units but in fact the coefficients are not constant across the cross section in the dynamic panel. Since their analysis of decade averages implies only two data points per country, the results of Easterly and Rebelo (1993) and Bose et al. (2007) on causality from infrastructure to growth might cast doubt on the validity of the procedure in terms of dynamic behavior of infrastructure. Deverajan et al. (1996) and Miller and Russek (1997) do consider some dynamic structure of infrastructure but do not reflect on econometric issue resulting from an assumption of homogeneity of the coefficient on the lagged dependent variable(s) when the coefficient is heterogenous. From this point of view, a dynamic panel framework in which we reflect the effect of public investment in T&C on growth over time and incorporate heterogeneous behavior of cross-units into model structure might address the relationship more appropriately than previous studies.

## 3 Methodology

The review on previous literature in Section 2 raises a potential weakness in their econometric approaches and suggests a dynamic panel framework. In a dynamic panel data model, we can not use the pooling regression or the Least Squares Dummy Variable (LSDV) estimation method due to the bias resulting from the correlation between the lagged dependent variables and the error term as shown in Nickell (1981), Anderson and Hsiao (1981, 1982), Hsiao (1986) and Kiviet (1995), among others.[11] The usual approach for dealing with this problem is to first-difference the data to remove the fixed effects and then estimate the model using instruments. Holtz-Eakin *et al.* (1988) adopt this approach in a framework for testing Granger causality in panels and suggest using a time-varying set of instruments that includes both differences and levels. Following Holtz-Eakin *et al.* (1988), we consider a bivariate dynamic panel model:

$$y_{it} = \alpha_0 + \sum_{j=1}^{m} \alpha_j y_{it-j} + \sum_{j=1}^{m} \beta_j x_{it-j} + f_i + \varepsilon_{it}, i = 1, 2, ..., N, \tag{1}$$

where $y_{it}$ and $x_{it}$ are the dependent variable and the causal variable at time $t$ for country $i$ respectively, $f_i$ is the fixed effect, the lag length $m$ is sufficiently large to ensure that $\varepsilon_{it}$ is a white noise error term and the $\alpha'$s and $\beta'$s are the coefficients of the linear projection of $y_{it}$ on a constant, past values of $y_{it}$ and $x_{it}$ and the individual effect $f_i$. Taking differences in equation (1) to eliminate the fixed effects leads to the model:

$$\Delta y_{it} = \sum_{j=1}^{m} \alpha_j \Delta y_{it-j} + \sum_{j=1}^{m} \beta_j \Delta x_{it-j} + u_{it}, i = 1, 2, ..., N, \tag{2}$$

where $\Delta y_{it-j} = y_{it-j} - y_{it-j-1}$ for $j = 0, 1, .., m$, $\Delta x_{it-j} = x_{it-j} - x_{it-j-1}$ for $j = 1, 2, , ..., m$ and $u_{it} = \varepsilon_{it} - \varepsilon_{it-1}$. Because $\Delta y_{it-1}$ is correlated with the first difference error term, $u_{it}(= \varepsilon_{it} - \varepsilon_{it-1})$, it is necessary to use instrumental variable procedures. Following Holtz-Eakin *et al.* (1988, 1989), we can estimate the equation (2) by using 2SLS with a time-varying set of instruments. Holtz-Eakin *et al.* suggest that the vector of instrumental variables, $\mathbf{Z}_{it}$, that is available to identify the parameters of equation (2), is

$$\mathbf{Z}_{it} = [1, y_{it-2}, y_{it-3}, ..., y_{i1}, x_{it-2}, ..., x_{i1}]'.$$

The authors address the question of whether $x$ Granger causes $y$ or not by testing the joint hypothesis:

$$\beta_1 = \beta_2 = ... = \beta_m = 0. \tag{3}$$

[11] See Baltagi (2001) for a useful overview of this issue.

In our study, we start with this procedure to address the question of whether public investment in T&C Granger causes economic growth.

However, there might be some potential problems for this instrumental approach. First of all, in practice it is often difficult to find good instruments for the first-differenced lagged dependent variable, which can itself create problems for the estimation. Kiviet (1995) shows that dynamic panel data models that use instrumental variable estimation methods (in particular, GMM) often lead to poor finite sample efficiency. Secondly, when this instrumental approach imposes the assumption that the coefficients of the lagged dependent variables are homogenous across the cross section, Pesaran and Smith (1995) point out that the homogenous assumption of the coefficient on the lagged dependent variables could introduce significant bias if in fact the coefficients are not constant across the cross section and this bias cannot be addressed with instrumental variables estimation (GMM).

To avoid the potential problem resulting from an assumption of homogeneity of the coefficient on the lagged dependent variable(s) when it is heterogeneous, Weinhold (1999) and Nair-Reichert and Weinhold (2001) propose a mixed fixed and random coefficients model (MFR) in which the intercepts and the coefficients on the lagged dependent variables are specific to the cross section units, while the coefficients on the exogenous variables are assumed to be normally distributed across the cross section. Thus, the MFR model allows for greater heterogeneity in the parameters on exogenous variables as well as lagged dependent variables than do the traditional models and accurately models the true relationship to avoid a misspecification bias. This model is originally developed by Hsiao *et al.* (1989) in a non-dynamic, non-fixed-effects panel data model of regional electricity demand and adapted in Weinhold (1996, 1999) as an alternative specification for panel data causality testing and of estimating panel data models with heterogeneous dynamics. In her Monte Carlo experiments, Weinhold (1999) shows, that the MFR model performs very well compared to instrumental variable approaches and the bias on the exogenous variable's parameter estimate of the MFR model is relatively small when $T$ is between 10 and 25 and $N$ is between 20 and 40[12]. Following Weinhold

[12] The bias ranges from 0.002 to 0.003 when the true value of the coefficient is 0.2. For further details

(1999) and Nair-Reichert and Weinhold (2001), we consider an alternative specification for dynamic panel data model:

$$y_{it} = \alpha_i + \sum_{j=1}^{m} \gamma_{ij} y_{it-j} + \sum_{j=1}^{m} \beta_{ij} x_{it-j} + \varepsilon_{it}, \tag{4}$$

where the coefficients on the lagged dependent variables, $\gamma_{ij}$, are country-specific, the coefficient on the exogenous explanatory variable $x$, $\beta_{ij}$, is drawn from a random distribution with mean $\overline{\beta}_j, \beta_{ij} = \overline{\beta}_j + v_i$ and $v_i$ is a random disturbance. In essence, this approach uses information on the distribution of the estimates on the lagged exogenous variable to extract the required information and to address the question of the direction of causality or possible joint determination between economic variables in a panel data set. Weinhold (1999) suggests that the estimated variance of the random coefficients can be used as a diagnostic tool to determine the extent of heterogeneity in the relationship in question and thus, if the estimated variance is quite large relative to the coefficient estimates, this is a signal of significant heterogeneity in the panel. In our study, we employ this approach for further investigation of the causality between public investment in T&C and growth.

# 4 Estimation results

## 4.1 Data

Existing studies aiming at evaluating growth effects of public investment at a disaggregated level frequently suffer from the 'sparseness of data' problem.[13] For us, however, this problem poses a greater challenge due to the fact that a formal test for causality requires usage of lags and leads of the variables in question and such an analysis needs to be based on data sets containing relatively large number of observations per country. To overcome these shortcomings, we aimed at collecting a large and balanced data

on the MFR model, see Weinhold (1999).

[13] For example, due to shortage of data, Easterly and Rebelo (1993) have based their analysis on the decade averages implying two data points per country. On the other hand, *Government Financial Statistics* (GFS), the primary source for data on government expenditures, does not provide any data for sectoral public investment.

set on central government (consolidated accounts) capital expenditure in the T&C sector for developing countries after consulting a large collection of World Bank Country Economic Reports and Public Expenditure Reviews.[14] But we ended up with a panel of 15 developing countries with annual data from 1970 to 1987 without any missing observations.[15]

In an earlier exercise, Easterly and Rebelo (1993) collected data on public investment by sectors. We differ from this existing data set on two grounds. First, our data set provides a time series for public investment expenditures in transportation and communication sector, which is useful for a formal causality test. Second, the measure of public investment used by Easterly and Rebelo (1993) has got a tendency (as acknowledged by the authors themselves) to overstate public investment by including investments by public firms (enterprise) that have activities and goals similar to those of the private sector. In contrast, we strictly follow the *GFS* guidelines and exclude public enterprise investment. Our data for the growth rate of GDP is taken from World Bank CDROM and our study uses the bivariate estimation.[16]

Even though the time period is only 18 years, the ADF test (Augmented Dickey-Fuller test) for a unit root in public investment in T&C indicates that all countries have such a unit root. To avoid a specification with non-stationary explanatory variables leading to spurious results in a panel, we take the growth rate of our variables as adopted in Nair-Reichert and Weinhold (2001). Thus, we ask whether an increase in the growth rate of public investment in T&C helps forecast an increase in economic growth. In other words, we try to address the question whether a relatively high growth rate of public investment in T&C will lead to relatively high GDP growth rate.

[14] Our data set and further details about the data sources are available on request from the authors. We wish to thank the World Bank for allowing us to use their archive at Washington D.C.

[15] Countries in our sample, are Bahamas, Congo, Ethiopia, Guatemala, Indonesia, Kenya, Malaysia, Morocco, Pakistan, Rwanda, Sierra Leone, Sri Lanka, Tanzania, Thailand, and Zambia.

[16] There may be other important contemporaneous control variables for the problem and thus our approach might be open to the criticism that the omitted variable bias could occur. Nevertheless, we use the bivariate model based on two grounds. First of all, it is data availability. We could not collect annual data for some basic control variables in our sample. Secondly, we implicitly assume that the lagged dependent variable provides a good proxy variable for many omitted variables.

### 4.2 Causality from public investment in T&C to economic growth

As outlined in section 3, we start with a traditional panel causality test proposed by Holtz-Eakin *et al.* (1988) for dynamic panel models. From the equations (1) and (2), we have:

$$\Delta GY_{it} = \sum_{j=1}^{m} \alpha_j \Delta GY_{it-j} + \sum_{j=1}^{m} \beta_j \Delta GTCI_{it-j} + u_{it}, i = 1, 2, ..., N, \quad (5)$$

where $GY_{it}$ and $GTCI_{it}$ are the growth rates of GDP and public investment in T&C for the country $i$ at time $t$ and $\Delta GY_{it}$ and $\Delta GTCI_{it}$ are the first differences. How can we choose the correct lag length, $m$? Holtz-Eakin *et al.* (1988, 1989) discuss how to find the "best" value of $m$. First of all, we choose a relatively large value of $m$ to be sure to avoid truncating the lag structure inappropriately. Denote by $\tilde{m}$ the relatively large value of $m$ used for initial estimation of the model. Re-estimate the system (5) with $m = \tilde{m} - 1$. If the increase in the sum of squared residuals is "large", then $m = \tilde{m}$ is accepted. If the increase is "small", then try $m = \tilde{m} - 2$. Continue testing successively smaller lag lengths until one is rejected by the data, or $m = 0$. This procedure is consistent with the "general to specific" methodology. Following this procedure, we estimate equation (5) with initial lag length 5 (i.e., $\tilde{m} = 5$) and instruments, constant, $GY_{t-j}$, and $GTC_{t-j}$, $j = 2, ..., 6$. For the null hypothesis that $m = \tilde{m} - 1$ $(i.e., m = 4)$, the test statistic $(\chi^2_{(2)})$ is 35.12 and we reject the null at the conventional level, implying that the increase in the sum of squared residuals with smaller lag length is large, implying that $m = \tilde{m} = 5$.[17]

Table 1 presents the results from 2SLS estimation of equation (5).[18] While the estimated coefficients on $\Delta GTCI_{it-j}$, $j = 1, 2, 3, 4$, are not statistically significant at the

[17] Holtz-eakin et al. (1988) state that while in principle it is desirable to begin by specifying an arbitrarily long initial lag length, the additional lag structure can affect the size of weighting matrix and for such large matrices of weighting matrix, standard numerical procedures for inversion may yield unsatisfactory results. In our previous version of the paper, following Holtz-Eakin et al. (1988), we initially assume a lag length $m = 3$. However, as we are looking for long-run effect, we tried to take sufficient lag-length. But as we are constrained by our time series of 18 years, we cannot go beyond 5 year lag-length as opposed to 10-year effect investigated by Easterly and Rebelo (1993). We feel that this can be considered as a convenient compromise between Deverajan et al. (1996) who use 5 year forward moving average and Easterly and Rebelo (1993) who use 10 year average.

[18] Here, our 2SLS estimation is different from Holtz-Eakin et al. (1989) described in Section 3 in terms of a time-varying set of instruments. We have constant instruments under assumption that only constant lagged $GY_{t-j}$ and $GTC_{t-j}$ are valid instruments at each point in time.

conventional level, that on $\Delta GTCI_{it-5}$ is statistically significant. In contrast to Easterly and Rebelo (1993), in which the effect of public investment in T&C on growth is robustly significant with instrumental variables but the size of the coefficients is disturbingly high, the values of the estimated coefficients in our case are -0.002 and 0.003 and thus somewhat lower. The Wald test of the null hypothesis that $\beta_1 = ... = \beta_5 = 0$ indicates that we reject the null at the 5% level and thus Holtz-Eakin *et al.* dynamic panel causality test suggests that growth in public investment in T&C Granger causes GDP growth.[19]

However, even though the net value of the estimated coefficient on $\Delta GTCI_{it-j}$ is positive, the value of only statistically significant estimated coefficient, $\hat{\beta}_5$, is 0.0006, which is close to zero. As in the previous literature, our result based on the dynamic panel instrumental variable estimation might not indicate clearly that infrastructure has a positive impact on economic growth.

One possible reason might be attributed to heterogeneity both in the dynamic structure as well as the relationship between the public investment in T&C and growth. The econometric analysis presented in table 1 is based on underlying assumptions about the homogeneity of the relationships in question across countries in the panel. However, it is reasonable to expect a bit of heterogeneity both in the dynamic structure as well as in the relationship between economic growth and public investment in T&C, especially, in a panel of developing countries. To investigate whether this result can be attributed to heterogeneity in the cross-country units, we employ the MFR model described in Section 3. Following Nair-Reichert and Weinhold (2001), we use orthogonalization which is necessary to ensure that the coefficients are independent which in turn allows their estimated variances to be appropriately interpreted. That is, we have:

$$GY_{it} = \alpha_i + \sum_{j=1}^{m} \gamma_{ij} GY_{it-j} + \sum_{j=1}^{m} \beta_{ij} GTCI^{o}_{it-j} + \varepsilon_{it}, \tag{6}$$

[19] Looking at the insample t-statistic of the coefficients on past causal variables is not technically what is called Granger causality test but since the in-sample approach is more commonly used (as in Nair-Reichert and Weinhold 2001), it is being done here in the interest of keeping things consistent with the literature.

where $GTCI^o_{it-j}$ denotes the orthogonalized growth rate of public investment in T&C after the linear influences of the other right-hand side variables have been removed (including any other lags of this variable if multiple lag lengths are used). As in the 2SLS estimation, we chose the lag length, $m = 5$.[20] The estimated mean and variance of the indicated causal variables over countries are reported in table 2, as are the standard errors of the estimated means.[21]

Weinhold (1999) and Nair-Reichert and Weinhold (2001) point out that if the estimated variance of the coefficients on $GTCI^o_{it-j}$ is quite large relative to the mean of the estimated coefficients, this is a signal of significant heterogeneity in the panel. The estimated variances of the random coefficients are not large, implying that there might not be a great deal of the heterogeneity across this panel. In contrast to 2SLS estimation, the estimated means of the coefficients on $GTCI_{it-1}$ and $GTCI_{it-3}$ are positive and statistically significant but the value of the estimated coefficients on $GTCI_{it-2,}$ $GTCI_{it-4,}$ and $GTCI_{it-5,}$ are negative but not statistically significant. The Wald test of the null hypothesis that $\beta_1 = ... = \beta_5 = 0$ shows that we reject the null. Thus, the statistically significant positive values and the Wald test imply that the public investment in T&C has a positive impact on economic growth and seem to support that there is a dynamic effect of public investment in T&C on growth.[22,23] Overall, there is not a great deal of heterogeneity in this relationship. Nevertheless, the MFR model seems to be an appropriate methodology for explaining previous controversial results and taking heterogeneous behavior in developing countries into account. In addition, the magnitudes of the values on the estimated coefficients of both estimations, Holtz-Eakin *et al.*'s

[20] In the MFR model framework, we investigate initial lag length $m = 5$. The $\chi^2_{(2)}$ for the wald test of the null hypothesis that $\alpha_5 = \beta_5 = 0$ is 8.52 and we reject the null at the 5% level, indicating that the lag length $m = 5$ is valid.

[21] We thank Diana Weinhold for graciously sharing her code for MFR estimation.

[22] We test the null hypothesis that the sum of the coefficients on the lags of the T&C investment is zero to investigate the long-run effect (or net effect) and we do not reject the null in the instrumental variable estimation (p-value 0.65). Nevetheless, we do not have significant implication on this result because all estimated coefficients on lagged T&C investments are not statistically significant except the coefficient on the fifth lagged T&C investment. In the MFR estimation, we do not reject the null hypothesis that the sum of all coefficients is zero but we do reject the null hypothesis that the sum of the first lagged and the third lagged coefficients, which are statistically significant, are zero (p-value 0.00003).

[23] This paper follows the popularized version of Granger causality tests and thus we look at the in-sample t-statistic of the coefficients on causal variables. Since the in-sample approach is more commonly used, this seems to be justifiable in the interest of keeping things consistent with the literature.

instrumental estimation and the MFR model, are not quite different.

In sum, our results from Holtz-Eakin *et al.*'s instrumental estimation and the MFR model for the dynamic panel suggest that public investment in T&C Granger causes economic growth and supports that infrastructure such as transportation and communication matters for economic growth in developing countries. Furthermore, the values of the estimated coefficients on public investment in T&C are considerably lower in contrast to previous literature as in Aschauer (1989), and Easterly and Rebelo (1993).[24]

### 4.3 Reverse causality

In terms of the accelerating effect of output on investment, as in Clark (1979) and Wagner's law in Abizadeh and Gray (1985) and Ram (1987), there might be reverse causality, which means that public investments in T&C follow growth and thus rapid growth leads to higher investments in this sector. To investigate this issue, we employ the same methodology. First of all, we consider Holtz-Eakin *et. al.*'s (1988) instrumental variable estimation for the dynamic panel as follows:

$$\Delta GTCI_{it} = \sum_{j=1}^{m} \alpha_j \Delta GTCI_{it-j} + \sum_{j=1}^{m} \beta_j \Delta GY_{it-j} + u_{it}, i = 1, 2, ..., N; t = 1, .., T. \quad (7)$$

Table 3 shows the 2SLS estimation results of equation (7). As in the causality test of equation (5), we choose the lag length at $m = 5$. The estimated coefficients on all $\Delta GY_{it-j}s$ are positive but not statistically significant and high, except on $\Delta GY_{it-2}$ which is statistically significant at the 10% level, suggesting that there might be a great deal of heterogeneity in the reverse relationship . The Wald test of the null that all coefficients on $\Delta GY_{it-j}s$ are zero, indicates that we can not reject the null at the conventional level.

[24] One interpretation of why the coefficients found in this study are smaller than those of Easterly & Rebelo (1993) might be because that study allows the effect of T&C investment to last up to 10 years, whereas this paper allows only up to 5 years. We have tried to investigate this issue further but as described in the Footnote 16, unfortunately, the examination is limited by data availability. In addition, the small coefficients in Table 1 following Holtz-Eakin *et al.* (1988) do not give us too much information other than indicating its economically negligible effect on growth. These small coefficients may also be attributed to potential heterogeneity that we addressed in by MFR model. Hence, we feel that more interesting result lies in the estimates of MFR model, where we have regressed the growth of GDP per capita on growth of T&C investment as the estimated coefficient can be readily interpreted as the heteroschedasticity-corrected lagged elasticities of T&C investment on GDP per capita (for example, these are 1.8% for one-year lagged T&C and 3.1% for 3-year lagged T&C).

This result implies that the reverse causality does not apply and thus economic growth does not Granger cause public investment in transportation and communication.

To examine this reverse causality taking heterogeneity in cross-country units into account, following Weinhold (1999) and Nair-Reichert and Weinhold (2001), we also estimate the MFR model:

$$GTCI_{it} = \alpha_i + \sum_{j=1}^{m} \gamma_{ij} GTCI_{it-j} + \sum_{j=1}^{m} \beta_{ij} GY^{o}_{it-j} + \varepsilon_{it}, \quad (8)$$

where we choose $m = 5$. The estimated mean and variance of the indicated causal variables are reported in table 4 as are the standard errors of the estimated means. While the estimated means of estimated coefficients on $GY_{it-1}$ and $GY_{it-2}$ are positive but not statistically significant, those of $GY_{it-3,}$ and $GY_{it-4}$ are negative and not statistically significant. The estimated coefficient on $GY_{it-5}$ is statistically significant but the value of the estimated coefficient is disturbingly high and negative. In particular, the variances of the estimated mean on the random coefficients are much larger relative to the mean, implying that there are a great deal of heterogeneity across this panel in the reverse causal relationship.

This finding of the MFR causality test is particularly interesting in the context of traditional causality test and underlying economic mechanism. First of all, the result indicates that heterogeneity in one causal direction can be significantly different from that in the other, which has not been considered in the traditional causality test. Secondly, while the economic mechanism from the growth-enhanced effect of the infrastructure works systematically over various countries, the mechanism based on Wagner's law may significantly vary across countries. One of possible reasons is that while the investment on infrastructure is closely tied to economic growth in most countries, high economic growth does not necessarily encourages constructive infrastructure over various countries.

Overall, from the estimation results in the MFR model and Holtz-Eakin *et al.*'s instrumental variable estimation for the dynamic panel causality test, we do not find significant evidence that there is a reverse causal relationship between growth and public investment in T&C and our empirical study shows little evidence on reverse causality that the investment acceleration hypothesis works in the case of public investment in transportation and communication and economic growth for developing countries.

## 5 Concluding remarks

Empirical literature on the relationship between public investment in transportation and communication and economic growth has reported a mixed picture; sometimes significant and positive, sometimes significant and negative, and sometimes not significant. In addition, the size of the estimated coefficient on public investment in T&C is somehow high, implying a result which naturally casts doubt on the validity of the procedure.

This paper re-examined this issue by considering the dynamic effect of public investment in T&C on growth over time and allowing for heterogeneity in developing countries. For this end, we started with Holtz-Eakin *et. al.*'s (1988) instrumental estimation which is a benchmark model for a dynamic panel causality test. Our results in the instrumental variable estimation show that public investment in T&C matters for economic growth and the values of the estimated coefficients on lagged public investments in T&C are relatively lower than in previous literature. However, most values of estimated coefficients are not statistically significant and the value of statistical significant coefficient is close to zero, indicating that it is not clear for the public investment on T&C to have an positive impact on growth.

To investigate further whether these results are attributed to heterogeneity in developing countries, we employ the mixed fixed and random coefficient model (MFR) of Weinhold (1999) and Nair-Reichert and Weinhold (2001). The estimation results indicate that the net effect of public investment in T&C on growth is positive and statistically significant. From the MFR estimation in which we allow heterogeneity across the countries, we find more significant evidence that public investment in T&C has a positive impact on economic growth over time. Overall, both estimations suggest that public investment in T&C takes time to affect growth and thus a dynamic panel model is desirable for studying the relationship between infrastructure such as T&C and economic growth.

However, from both approaches, we do not find significant evidence on the reverse causality which is suggested by the acceleration effect of output on investment and Wagner's law. In particular, the MFR model estimation suggests that there is a great deal of heterogeneity across developing countries. These results indicate that heterogeneity in on causal direction can be significantly different from that in the other and it is necessary to be considered in the traditional causality test. Overall, we feel that the effect of public investment in transportation and communication on economic growth is generally significant and considerable, while the other way around is questionable for developing countries.

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Table 1

Holtz-Eakin *et al.* (1988) Causality test

$$\Delta GY_{it} = \sum_{j=1}^{m} \alpha_j \Delta GY_{it-j} + \sum_{j=1}^{m} \beta_j \Delta GTCI_{it-j} + u_{it},\ j = 1, ..., 5$$

| variable | Est. coeff. |
|---|---|
| $\Delta GY_1$ | $-0.494^{***}$ (0.070) |
| $\Delta GY_2$ | $-0.273^{***}$ (0.076) |
| $\Delta GY_3$ | $-0.300^{***}$ (0.073) |
| $\Delta GY_4$ | $-0.367^{***}$ (0.070) |
| $\Delta GY_5$ | $-0.362^{***}$ (0.070) |
| $\Delta GTCI_1$ | $-0.002$ (0.002) |
| $\Delta GTCI_2$ | 0.003 (0.004) |
| $\Delta GTCI_3$ | 0.003 (0.004) |
| $\Delta GTCI_4$ | 0.0001 (0.003) |
| $\Delta GTCI_5$ | $0.0006^{***}$ (0.0002) |
| $H_0 : \beta_1 = ... = \beta_5 = 0$ | $16.87^{***}$ |
| $N$ | 165 |
| $R^2$ | 0.132 |

Note: a. Instruments are constant, $GY_{it-j}, GTCI_{it-j}, j = 2, .., 6$.

b. *** denotes the statistical significance at the 1% level. Est. coeff. denotes the estimated coefficient on the explanatory variables.

Table 2

MFR Causality Test

$$GY_{it} = \alpha_i + \sum_{j=1}^{m} \alpha_{ij} GY_{it-j} + \sum_{j=1}^{m} \beta_{ij} GTCI_{it-j} + u_{it},\ j = 1, ..., 5$$

| variable | Est. coeff. | Coeff. variance |
|---|---|---|
| $GTCI_1$ | $0.018^{**}$ (0.008) | 0.002 |
| $GTCI_2$ | $-0.007$ (0.010) | 0.004 |
| $GTCI_3$ | $0.031^{***}$ (0.009) | 0.004 |
| $GTCI_4$ | $-0.010$ (0.008) | 0.016 |
| $GTCI_5$ | $-0.013$ (0.008) | 0.037 |
| $H_0 : \beta_1 = ... = \beta_5 = 0$ | $22.41^{***}$ | |
| $N$ | 180 | |
| $R^2$ | 0.279 | |

Note: a. *** and ** denote a statistical significance at the 1% level and at the 5% level respectively.
b. Est. coeff., and Coeff. variance denote the estimated mean and the estimated variance of the of random coefficients and $N$ is the number of observations. Figures in parentheses are the standard errors of the estimated mean of the random coefficients.

Table 3

Holtz-Eakin *et al.* (1988) *Reverse* Causality test

$$\Delta GTCI_{it} = \sum_{j=1}^{m} \alpha_j \Delta GTCI_{it-j} + \sum_{j=1}^{m} \beta_j \Delta GY_{it-j} + u_{it},\ j = 1, ..., 5$$

| variable | Est. coeff. |
|---|---|
| $\Delta GTCI_1$ | $-0.998^{***}$ (0.080) |
| $\Delta GTCI_2$ | $-0.830^{***}$ (0.185) |
| $\Delta GTCI_3$ | $-0.519^{***}$ (0.197) |
| $\Delta GTCI_4$ | $-0.203$ (0.162) |
| $\Delta GTCI_5$ | 0.010 (0.011) |
| $\Delta GY_1$ | 4.312 (3.619) |
| $\Delta GY_2$ | $6.661^{*}$ (3.921) |
| $\Delta GY_3$ | 3.902 (3.767) |
| $\Delta GY_4$ | 0.976 (3.724) |
| $\Delta GY_5$ | 3.079 (3.579) |
| $H_0 : \beta_1 = ... = \beta_5 = 0$ | 3.73 |
| $N$ | 165 |
| $R^2$ | 0.134 |

Note: a. Instruments are constant, $GY_{it-j}, GTCI_{it-j}, j = 2, .., 6$.

b. *** and * denotes the statistical significance at the 1% level and the 10% level respectively.

Table 4

MFR Reverse Causality Test

$$GTCI_{it} = \alpha_i + \sum_{j=1}^{m} \alpha_{ij} GTCI_{it-j} + \sum_{j=1}^{m} \beta_{ij} GY_{it-j} + u_{it},\ j = 1, ..., 5$$

| variable | Est. coeff. | Coeff. variance |
|---|---|---|
| $GY_1$ | 2.582 (3.115) | 106.0 |
| $GY_2$ | 0.645 (3.792) | 74.12 |
| $GY_3$ | −0.964 (3.600) | 605.6 |
| $GY_4$ | −0.868 (3.948) | 693.5 |
| $GY_5$ | −31.00*** (7.062) | 187,708 |
| $N$ | 180 | |
| $R^2$ | 0.02 | |

Note: a. *** and ** denote a statistical significance at the 1% level and at the 5% level respectively.

b. Est. coeff., and Coeff. variance denote the estimated mean and the estimated variance of the of random coefficients and $N$ is the number of observations. Figures in parentheses are the standard errors of the estimated mean of the random coefficients.