Bulletin of Applied Economics, 2020, 7(2), 49-76

Public and Foreign Investment Spending in the Argentine Case. A Cointegration Analysis with Structural Breaks, 1960-2015.

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Abstract

This paper examines whether public investment spending and inward foreign direct investment (FDI) enhance labor productivity growth in Argentina. Using annual data, it estimates a dynamic labor productivity function for the 1960-2015 period that incorporates the impact of public and private investment spending, education expenditures, the labor force, and export growth. It tests for both single and two-break unit root tests, as well as performing cointegration tests with an endogenously determined regime shift over the 1960-2015 period. Cointegration analysis suggests that a long-term relationship exists among the relevant variables. The error correction (EC) models suggest that (lagged) increases in public investment spending and education have a positive and significant effect on the rate of labor productivity growth Also, the model is estimated for a shorter period (1970-2015) to capture the impact of inward FDI flows. The estimates suggest that (lagged) FDI flows have a positive and significant impact on labor productivity growth, while increases in the labor force have a negative effect. From a policy standpoint, the findings call into question the politically expedient policy in many Latin American countries, including Argentina during the 1990s and 2000s, of disproportionately reducing public capital expenditures on education and infrastructure to meet reductions in the fiscal deficit as a proportion of GDP. The results give further support to pro-investment and pro-growth policies designed to promote public investment spending and attract inward FDI flows.

JEL classification numbers: C22, O10, O54
Keywords: Complementarity Hypothesis, Education expenditures, Endogenous growth, Foreign Direct Investment (FDI), Gregory-Hansen cointegration single-break test, Lee-Strazicich two-break unit root test, Johansen Cointegration Test, Public Investment, Vector Error Correction model (VECM).

1 Introduction

Since the mid-eighties, major Latin American countries such as Chile, Brazil, and Mexico have adopted an outward-oriented, market-based strategy of economic growth by liberalizing their trade and financial sectors, as well as dismantling and privatizing their state-owned enterprises. Argentina began this process of economic stabilization and structural reform in earnest following the country's adoption of a currency board system in 1991—the so-called "Convertibility Plan"-- under the administration of Carlos Saul Menem.[2] The Plan set the Argentinean peso to the dollar on a one-to-one basis, thus eliminating the

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[2] Argentina's privatization, liberalization, and deregulation program is discussed and analyzed in Baer et. al. (2002) and Weisbrot et. al. (2002).

Article Info: *Received*: May 1, 2020. *Published online*: June 15, 2020

ability of the government to finance budget deficits via direct money. A key accomplishment of the currency board was to reduce dramatically the rate of inflation from 2,314 percent in 1990 to 4.1 percent in 1994, and less than 1 percent in 1998! Another accomplishment of the plan was a dramatic increase in FDI net inflows: from US$ 1.84 billion in 1991 to an all-time high of US$23.9 billion in 1999, before falling to US$11.7 billion in 2000. These capital flows fell to US$1.6 billion in 2003 as a result of the economic and financial debacle the economy experienced following the collapse of the Convertibility Plan in 2002 (see World Investment Report, 2012).

Unfortunately, the stabilization of the Argentine economy during the nineties was not sustainable because of the impact of several external shocks and ill-conceived policies that paved the way for the economic and financial debacle associated with the collapse of the Convertibility Plan in 2001-2002. First, the highly adverse effects of the Tequila crisis in 1995-96 generated massive capital flight, a liquidity crisis, and high real interest rates with their knock-on effects on the balance sheets of the banks and the real economy. Second, the Asian and Russian crises led to a significant flight of capital and, once again, a substantial rise in real interest rates and their adverse effects on investment and consumption spending. Third, the devaluation of the Brazilian currency (the real) in 1999 had a severe effect on the Argentine economy because close to 30 percent of its exports were destined for that country (see Weisbrot et. al., 2002). The devaluation of the Real had a disproportionate effect because the promotion of outward-oriented policies had significantly increased the relative importance of its exports (and foreign capital) in fueling economic growth. For example, exports of goods and services averaged 9.5 percent of GDP during the 1990-2001 period as compared to 8.6 percent during the 1980-90 period (see ECLAC, 2011). Finally, the economic situation was further exacerbated by the fact that the dollar continued to appreciate in real terms relative to the Euro and the Yen, thus further undermining the competitiveness of the Argentine economy given its hard peg to the dollar and its policy of unrestricted capital mobility (see Baer et. al., 2002).

Moreover, IMF-sponsored stabilization policies adopted by the Argentine government in the wake of these shocks led to short- and long-term economic (negative) effects on consumption, investment, and economic growth (see Calva, 1997; Maia and Kweitel, 2003; Stiglitz, 2003; Taylor 1997; Baer et. al., 2002; Weisbrot and Sandoval, 2007; and Weisbrot, 2011). These programs often call for across-the-board cuts in public spending and tight restrictions on credit creation to meet stringent fiscal deficit targets, reduce the rate of inflation, and free resources to service the external debt.[3] Nowhere is this more evident than in the disappointing and erratic behavior of Argentine private capital formation during the past two and a half decades. Table 1 below shows that Argentina's private investment as a proportion of GDP fell dramatically during the "lost decade" of the 1980s, reaching a low 9.4 percent in 1990 which amounted to less than half its level in 1980. Following the adoption of the Convertibility Plan it rose to a high of 19.1 percent in 1994, from which it fell again to 9.2 percent in 2002 and a dismal 7.6 percent in 2003 as a result of the country's economic crisis following the collapse of the currency board.[4] In this connection, most economists believe that it is essential for Argentina (and other Latin economies) to improve and sustain its investment performance if it is going to lay the groundwork for rapid and sustained economic growth and employment opportunities for its rapidly expanding labor force (see Bose, et al., 2009; Calva, 1997; Moguillansky, 1996; Baer et al., 2002; and Weisbrot and Sandoval, 2007).

Several prominent investigators have cited the dramatic fall in public investment in economic and social infrastructure, brought about by the need to meet the stringent fiscal deficit targets of the stabilization program, as one possible factor in explaining the poor investment performance of Argentina and other Latin American countries. Table 1 below shows that public investment spending in economic

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[3] Weisbrot (2002) observes that in 2002 the IMF demanded that the Argentine government enact spending cuts of 10 percent across-the-board, in addition to a 30 percent reduction in outlays for goods and services and a 13 percent cut in salary and pensions for government employees (p. 13).

[4].It should be mentioned that during the 1990s only a handful of countries, notably Chile and Costa Rica, have managed to record investment ratios comparable to those before the onset of the debt crisis in the early 1980s.

and social infrastructure as a proportion of GDP fell precipitously from 4.6 percent in 1990 to barely 1 percent in 1994, only to rise to 2 percent during the 1995-99 period before falling again in the 2001-2003 period to less than 1 percent. Moreover, the average public investment spending on economic infrastructure in the 1990s and early 2000s is only a third of that of the 1980s and barely one fifth of the average levels recorded during the 1960s and 1970s. However, Table 1 also reveals that under the pro-investment and pro-growth policies of both Kirschner administrations (2003-2015), public investment as a percentage of GDP has risen dramatically since 2006, recording, consistently, levels around 3 percent. It is also notable that private investment spending as a proportion of GDP has risen by one full percentage point during the 2011-2015 period as compared to 2003-2010.

The basic idea is that public investments in highways, bridges, sewerage systems, water supplies, and education and health services often generate substantial positive spillover benefits for the private sector by reducing the direct (and indirect) costs of producing, transporting, and delivering goods and services to consumers. (See Albala-Bertrand and Mamatzakis, 2001; Aschauer, 1989; Cardoso, 1993; Devarajan and Zou 1994; Devarajan et al., 1996; Khan and Reinhart, 1990; Bose, et al., 2009; Ramirez, 2002 and 2010; Serven and Solimano, 1993; and Weisbrot and Merling, 2018).[5] Moreover, capital expenditure cuts may undermine some or all of the long-term efficiency gains expected from the implementation of market-based, outward-oriented reforms such as privatization of state-owned firms and the liberalization of trade and finance (see Killick, 1995; and Stiglitz, 2012). After all, the newly privatized firms in liberalized (open) markets will need adequate and reliable economic infrastructure in order to produce, transport, and market their goods and services at home and abroad in a cost-effective manner.[6] Given the importance and controversial nature of this topic, this paper analyzes the impact of public investment spending on infrastructure and education, inward FDI flows, and export growth on the economic growth and labor productivity of the Argentine economy during the 1960-2015 period. The choice of Argentina is warranted for several reasons. First, Argentina is a large and strategically important country in Latin America and the second largest economy in South America. This is a situation that promises to continue as a result of the country's participation in the important regional trade agreement named Mercosur and its status as a major regional exporter of maize and soybeans. Second, Argentina is one of the few countries in Latin America that has reliable and disaggregated annual time-series data on public investment spending on economic and social infrastructure going as far back as the decade of the sixties. This data set thus enables researchers to test whether increases in government *investment* spending on economic infrastructure *per se*, rather than *overall* public investment expenditures, displace or promote private investment spending, economic growth, and (labor) productivity. Third, beginning with the Menem administration (1989-1999) and continuing under the ill-fated administrations of Fernando De La Rua and Duhalde (2000-2002), Argentina pursued a far-reaching neoliberal strategy of economic growth that was heralded by pundits and multilateral organizations as a prototype for the region—at least before the onset of the crisis. However, under both Kirschner administrations (2003-2015), but particularly that of Cristina Fernandez de Kirschner (2007-2015), the Argentine government reversed itself and pursued a more activist set of Keynesian-style policies designed to promote broad-based economic growth with an emphasis on the promotion of public investment and education.[7]

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[5] This paper only addresses the direct output effects of increasing public investment spending. It ignores the impact of public investment spending on the relative prices that private firms face for key inputs and services.

[6] Of course, the public sector need not provide these public goods directly; the goods can be contracted out to the private sector in accordance with government regulations and guidelines. In fact, many governments in Latin America (including Argentina) have awarded concessions to private firms to produce and provide quasi-public goods and services. However, if the monitoring or supervision cost of outsourcing public works projects is high, then the bias in favor of privatizing these types of expenditures is removed (see Ramirez, 2002).

[7] Nestor Kirschner, a member of the Justicialist (Peronist) Party, served as President of Argentina during the 2003-2007 period, and his wife, Cristina Fernandez de Kirschner, succeeded him as President of Argentina in 2007. Under both Kirschner administrations, the economy grew at average annual rates exceeding 8 percent and levels of poverty and unemployment experienced a dramatic fall from their crisis levels in 2001-02; there also was a

Finally, in 2015 the pendulum swung back under the administration of the right-center president Mauricio Macri (2015-2019). The government returned to the neoliberal policies of the past, including the adoption of a three-year draconian IMF-sponsored program in June 2018 to combat a severe economic and financial crisis that erupted in 2018. The crisis was largely the result of ill-conceived export tax cuts early on in the administration and a large and rapid run-up in foreign borrowing. To make matters worse, there was an equally rapid and ill-timed removal of currency controls amid rising interest rates in the U.S. which, in turn, fueled massive capital flight, and last but not least, a severe drought ravaged the country's all-important agro-export sector. The ongoing crisis has sent the peso reeling more than 50 percent since 2018 and industrial output has plummeted, unemployment and poverty have risen sharply, real wages have nose-dived, and interest have soared to more than 70 percent in 2019, risking a repeat of the devastating economic crisis of 2001-02 (see ECLAC, 2018; and Weisbrot and Merling, 2018).[8]

The paper is organized as follows. Section II provides a conceptual framework for incorporating the public or FDI capital stock in a modified neoclassical production function. The model presented in this section is intended solely to motivate the ensuing discussion and although the relevant parameters cannot be estimated directly given the inherent data limitations present in the Argentine case, the discussion highlights how researchers might proceed if the relevant data becomes available. Next, the paper introduces a rough empirical counterpart to the model presented in the previous section and discusses the nature and limitations of the data used in this study. Section IV presents both single-break (Zivot-Andrews) and two-break (Lee-Strazicich) unit root tests for the variables included in the dynamic production relationship. Using cointegration analysis with structural breaks (Gregory-Hansen), this section also tests whether there is a stable long-term relationship among the relevant regressors of the modified production function. In so doing, this paper goes beyond other empirical studies of the complementarity hypothesis by addressing the important question of spurious correlation among the model variables. The section is brought to a close by generating several error-correction (EC) models that are used to track the historical data on the growth rate of labor productivity for the period under review. Section V presents estimates for a small-scale Vector Error Correction Model (VECM) that treats all the included variables as endogenous. In general, the estimates are consistent with those of the larger (EC) models. The last section summarizes the paper's major findings.

2 Conceptual Framework

On the supply side, the positive externalities generated by additions to the public (or FDI) capital stock can be formalized by incorporating them in an augmented Cobb-Douglas production function of the following form (see Barro and Sala-I-Martin 1995; and De Mello 1997):

$$Y = A\,\varphi\,[L, K_p, E] = A\,L^{\alpha}\,K^{\beta}\,E^{(1-\alpha-\beta)} \qquad (1)$$

where Y is real output, K_p is the private capital stock, L is labor, and E denotes the externality generated by additions of the public capital stock or FDI capital stock (α and β are the shares of domestic labor and private capital respectively, and A captures the efficiency of production. Initially, it is assumed that α and

significant increase in government spending on housing, health and economic infrastructure, as well as a generous extension of social security coverage and a substantial rise in real wages (see Weisbrot, 2011, pp. 8-12).

[8] Preliminary estimates by ECLAC (2018) indicate that the growth rate in real GDP fell by 2.6 percent in 2018 after expanding in 2017 by 2.9 percent. The data also indicate that during 2018 inflation surged to an estimated 45.5 percent from 25 percent the previous year, and that the unemployment rate rose by a full percentage point to 9.4 percent compared to the previous year. Economic activity is expected to continue to fall in 2019 because the US56.3bn stand-by agreement with the IMF calls for substantial real cuts in in public expenditures on subsidies (energy and transportation), public sector wages, and public investment on economic infrastructure. For further details, see ECLAC (2018), pp. 1-4 and Statistical Annex, pp. 108-138; and Weisbrot and Merling, 2018, pp. 1-19.

β are less than one, such that there are diminishing returns to the labor and capital inputs.
The externality, E, can be represented by a Cobb -Douglas function of the type:

$$E = [L, K_p, K_g^{\gamma}]^{\theta} \quad (2)$$

where γ and θ are, respectively, the marginal and the intertemporal elasticities of substitution between private and public (FDI) capital. Let $\gamma > 0$, such that a larger stock of public (or FDI) capital generates a positive externality to the economy; i.e., knowledge or technological progress is an accidental by-product of capital investment by relatively small firms in the form of FDI or public investment. If $\theta > 0$, intertemporal complementarity prevails and, if $\theta < 0$, additions to stock of public (FDI) capital crowd out private capital over time [see Jones, 2011].

Combining equations (1) and (2), we obtained,

$$Y = A\, L^{\alpha + \theta(1 - \alpha - \beta)}\, K_p^{\beta + \theta(1 - \alpha - \beta)}\, K_g^{\gamma\theta(1 - \alpha - \beta)} \quad (3)$$

A standard growth accounting equation can be derived by taking logarithms and time derivatives of equation (3) to generate the following dynamic production function:

$$g_y = g_A + [\alpha + \theta(1-\alpha-\beta)]g_L + [\beta + \theta(1-\alpha-\beta)]g_{Kp} + [\gamma\theta(1-\alpha-\beta)]g_{Kg} \quad (4)$$

where g_i is the growth rate of i = Y, A, L, K_p, and K_g. Equation (4) states that (provided γ and $\theta > 0$) additions to the stock of public (FDI) capital will augment the elasticities of output for labor and capital by a factor $\theta(1-\alpha - \beta)$.

<u>Preferences</u>.

The demand side of the economy can be included into the model via the following intertemporal utility maximization framework:

$$\text{Max } u(t) = \int_o^{\infty} u(c(t))\, e^{-\rho t} L(t)\, dt \quad (5)$$
$$\text{s. t. } \dot{k}_p = A\, k_p^{\beta + \theta(1-\beta)}\, k_g^{\gamma\theta(1-\beta)} - c - \delta k_p, \text{ and } k_p(0) \geq 0.$$

where, for convenience, lower-case letters are defined in per capita terms and ρ is the discount rate, L(t) is the size of the family, c(t) is per capita consumption, and δ represents the rate of depreciation. For convenience, the initial population is normalized to 1 so that the analysis in aggregate and per capita terms is the same. The instantaneous utility function of the representative consumer is assumed to exhibit constant relative risk and can be written in the following general form:

$$u(c(t)) = (c(t)^{1-\sigma} - 1) / (1- \sigma) \quad (6)$$

σ denotes the relative risk aversion coefficient or the inverse of the elasticity of substitution between current and future consumption; i.e., σ is an index of the representative consumer's willingness to exchange current consumption for future consumption. Letting u(c) = lnc, for simplicity, and solving the standard optimal control problem in equation (5), we obtained the following equation:

$$\dot{c}/c = 1/\sigma\{A[\beta + \theta(1-\beta)]k_p^{\beta + \theta(1-\beta) -1}\, k_g^{\gamma\theta(1-\beta)} - \rho\} \quad (7)$$

Equation (7) can be interpreted as follows in the presence of intertemporal complementarity between public and private capital (i.e., $\theta > 0$): the economy grows at a positive rate whenever the marginal product of capital, net of depreciation, can be kept above the rate of time preference (discount) [the dot over c denotes the rate of change of per capita consumption]. The marginal productivity of private capital, in turn, is augmented and kept above the discount rate by additions to the stock of public capital. Finally, the larger the intertemporal elasticity of substitution of current consumption for future consumption, as captured by the inverse of the relative risk coefficient, σ, the higher the rate of growth of the economy. Put differently, the sacrifice of current consumption is less costly to the representative consumer when present and future consumption are good substitutes (see Jones, 2011).

3 Empirical Model

It is often not possible to generate estimates of equations (3) and (4) directly because of the poor quality of existing data for key variables over a long time. Instead, investigators have used proxies for variables such as the labor force and/or the stocks of private and public capital such as population data rather than labor force data, or substituted investment data (as a proportion of GDP) for capital stock data (see Aschauer, 1989; Cardoso, 1993; Lin, 1994; and Ramirez, 2002; 2010). However, models that use these proxies have to impose unduly restrictive assumptions (e.g., such as a fixed capital-output ratio) or unrealistic assumptions (a constant labor force participation rate) that can generate both misspecified relationships and significant measurement errors.

In the case of Argentina we are fortunate to have annual labor force data going as far back as 1960, but we do not have consistent estimates of the public and private capital stock series, or for that matter, reliable estimates of the rate of depreciation from which such a series could be generated. Researchers in the field of economic development have circumvented this problem by estimating a dynamic production, which defines the relevant variables in terms of percentage growth rates, thus permitting them to generate proxies for the percentage growth rates in the respective capital stocks. Following their lead, this study includes the annual ratio of public and private investment spending to the gross domestic product as alternative proxies. Finally, for reasons explained in Section IV, the empirical model was estimated with *changes* in the investment ratios because these ratios were determined to be non-stationary in level form. Using annual data, this study extends previous empirical work by estimating an empirical counterpart of the dynamic production function in equation (4) for the 1960-2015 period for a total of 56 initial observations without the FDI variable; it also esimtates eq. (4) between 1970 and 2015 with the FDI variable for a total of 46 initial observations.[9]

The most general formulation of the growth equation is given below,

$$\Delta y = \alpha + \beta_1 \Delta l + \beta_2 \Delta(i_p) + \beta_3 \Delta(i_g) + \beta_4 \Delta(i_f) + \beta_5 \Delta(e_g) + \beta_6 \Delta(x) + \beta_7 D_1 + \beta_8 D2 + \varepsilon \tag{8}$$

lower case letters denote natural logarithms, and Δ denotes the change in the variable in question; y is real GDP (1993 pesos); l, as indicated above, refers to the labor force (thousands occupied); i_p denotes the ratio of private investment to GDP, while i_g represents public investment spending on economic infrastructure as a proportion of GDP, viz., roads, bridges, and ports[10]--it therefore excludes investment expenditures by state-owned enterprises which are more likely to crowd out private investment spending and output; i_f the ratio of foreign direct investment to GDP and it is expected to have a positive effect because increased FDI flows are associated with a greater transfer of technology and managerial

[9] Annual data for the FDI ratio were not available for Argentina prior to 1970 (see Maitan and Keitel, 2003; and World Investment Reports, 2011 and 2012).

[10] It should be noted that the reported government investment data contains a portion that is devoted to health and education expenditures, and should be treated separately as public (human) capital investment. At the risk of some overlap, I have included in this study includes a separate proxy for human capital, viz., government expenditures on education as a proportion of GDP. The study also made use of enrollment ratios for primary, secondary and tertiary education for the period under review, but the data was more limited and sporadic in nature and had to be interpolated for some years. See World Development indicators, published by World Bank (2018).

knowhow, learning-by-doing, and greater market discipline. However, FDI flows may also have a negative effect on the growth rate of a country if they give rise to substantial reverse flows in the form of remittances of profits and dividends and/or if the TNCs obtain substantial tax and other concessions from the host country (see Agbloyor et al., 2014; DeMello, Jr., 1997; Ram and Zhang, 2002; and Ramirez, 2019). e_g is real government expenditures on education (at all levels) as a proportion of GDP, and it is a proxy for human capital. This variable is expected to directly or indirectly crowd in private investment expenditures, and thus affect output (and labor productivity) positively. The x variable denotes exports of goods and services. The export promotion hypothesis suggests that its growth rate is expected to not only have a direct effect on economic growth, but also indirectly via the increased investment and realization of economies of scale by the exporting firms, and the concomitant diffusion of technological and managerial knowhow throughout the economy generated by the export sector (see Ram and Zhang, 2002). D_1 is a dummy variable that takes a value of one for the crisis years, and 0 otherwise, while D_2 equals 1 for the impact of the currency board, and 0 otherwise. Finally, ε is a normally distributed error term.

Data

The annual data used in this study were obtained from official government sources such as the *Direccion Nacional de Politicas Macroeconomica, Ministerio de Economia y Produccion (Ministry of Economy and Production, various issues)* and the *Instituto Nacional De Estadistica y Censos de la Republica Argentina (National Institute of Statistics and Census of Argentina)*. Other relevant economic data have been obtained from the *World Development Indicators for Argentina,* World Bank (2018); ECLAC, *Statistical Yearbook for Latin America and the Caribbean, 2018,* and the *International Finance Corporation* (Everhart and Sumlinski, 2001).

The dependent variable was estimated as the growth rate in labor productivity by subtracting the growth rate in the labor force from the percentage change in GDP in eq (4). Defining the dependent variable in this manner reverses the expected sign of the labor variable because of diminishing returns to the labor input. The sign of β_1 is anticipated to be positive in the GDP formulation while, as indicated above, it is expected to be negative in the labor productivity specification. β_2 is expected to be positive, while the sign of β_3 can be positive or negative depending on whether increases in public in public investment complement or substitute for private capital formation. Lags were included for this variable because of the delayed impact of government investment spending on economic infrastructure, private investment spending and private output growth.[11] B_4 is expected to have a positive sign, but for reasons alluded to above, its sign could also be negative (see Ram and Zhang, 2002). The sign of β_5 is also expected to be positive because government expenditures on education (a proxy for human capital) may directly or indirectly crowd in private investment expenditures and thus affect output (and labor productivity) positively. β_6 is expected to be positive for reasons alluded to above, while β_7 is anticipated to be negative for obvious reasons. Finally, β_8 is expected to be positive.

4 Unit Roots, Structural Breaks, and Cointegration Analysis

Initially, conventional unit root tests (without a structural break) were undertaken for the variables in question given that it is well-known that macro time series data tend to exhibit a deterministic and/or stochastic trend that renders them non-stationary; i.e., the variables have means, variances, and covariances that are not time invariant (see Granger and Newbold 1986). This study tested the variables

[11] Another important reason for including lags is that it reduces, to some extent, the criticism of reverse causation from the rate of growth in GDP to the growth rate in public investment spending, i.e., the economic argument which suggests that public investment is a normal good whose rate of growth will decline when the rate of output growth (or productivity) declines and tax revenues fall and increase during periods of rapid economic activity and rising tax revenues. The order of the lag length was determined by applying the AIC and SIC criteria.

in question for a unit root (non-stationarity) by using an Augmented Dickey-Fuller test (ADF) with a lag length automatically determined by the Schwarz Information Criterion (SIC).

Before reporting the unit root tests, it is important to acknowledge that when dealing with historical time series data for developing countries such as Argentina or Chile investigators are often constrained by the relatively small number of time series observations (usually annual data). This is the case in this study where the sample size is just at the threshold level of 50 annual observations recommended by Granger and Newbold (1986), which may compromise the power of the unit root (and cointegration) tests—not to mention distorting the size or significance of the tests as well (see Charemza and Deadman, 1997). However, a growing literature contends that the power of unit root (and cointegration) tests depends on the length or *time span* of the data rather than the number of observations per se. That is, for a given sample size n, the power of the test is greater when the time span is longer. Thus, unit root or cointegration tests based on 56 observations over 56 years have considerably more power than those based on 100 observations over 100 days (see Bahnam-Oskooee 1996; and Shiller and Perron, 1985).[12]

Table 2A presents the results of running an ADF test (one lag) for the variables in both level and differenced form under the assumption of a stochastic trend only, i.e., the test is run with a constant term and no time trend.[13] It can be readily seen that all the variables in the level form are nonstationary; i.e., they appear to follow a random walk with (positive) drift (see Nelson and Plosser, 1982). In the case of first differences, however, the null hypothesis of non-stationarity is rejected for all variables (except one) at least at the 5 percent level.[14] Thus, the evidence presented suggests that the variables in question follow primarily a stochastic trend as opposed to a deterministic one, although there is the possibility that for given sub-periods they follow a mixed process.[15]

Given the relatively low power of the ADF unit root tests when the data generating process is stationary but with a root close to the unit root, Table 2B presents the results of running a confirmatory KPSS stationarity test (Kwaitkowski et al., 1992). This test has a no unit root (stationary) null hypothesis, thus reversing the null and alternative hypotheses under the Dickey Fuller test. The reported results in both level and differenced form under the assumption of a constant are, by and large, consistent with those reported in Table 2A. For example, except for the labor productivity variable, the null hypothesis of no unit root can be rejected for all the variables in level form at the 5 percent level of significance; i.e., they appear to follow a random walk with (positive) drift. In the case of first differences, however, the null hypothesis of stationarity cannot be rejected for all variables at least at the 5 percent level. Thus, the evidence presented suggests that the variables in question follow primarily a stochastic trend as opposed

[12] For example, Hakkio and Rush (1991) contend that in nearly non-stationary time series "the frequency of observation plays a very minor role" in cointegration [and unit root] analysis because "cointegraton is a long-run property, and thus we often need long spans of data to properly test it" (p. 579). Similarly, Bahmani-Oskooee (1996) observes that in cointegration (and unit root) analysis using annual data over 30 years "is as good as using quarterly data over the same period" (p. 481). To some degree, this addresses the strong analytical and policy inferences drawn from a relatively small sample size.

[13] A stochastic trend is one where the random component of the series itself, say variable xt, contributes directly to the long run pattern of the series, either upward or downward. However, in the case of a deterministic trend the deviations from the non-stationary mean over time are quickly corrected. It is also possible for the variable in question to display both a stochastic and deterministic trend process over time. For further details see Charemza and Deadman, (1997, pp. 84-92).

[14] The order of the lag length was determined by applying both the Akaike Information Criterion (AIC) and the Schwarz Bayesian Information Criterion (SIC). For all the variables in this study, the ADF tests with one lag showed the lowest value for both the AIC and SIC criteria.

[15] This study also performed an ADF test (one lag) on the variables in logarithmic form with a deterministic trend. The results indicate that the null hypothesis of non-stationarity cannot be rejected for any of the variables in level form with a deterministic trend, suggesting that the variables in question do not exhibit a deterministic time trend throughout the period under review.

to a deterministic one, although there is a possibility that for given sub-periods they follow a mixed process.

A. Single-Break Unit Root Test.

The conventional results reported in Table 2 may be misleading because the power of the ADF test may be significantly reduced when the stationary alternative is true and a structural break is ignored (see Perron, 1988); that is, the investigator may erroneously conclude that there is a unit root in the relevant series. To test for an unknown one-time break in the data, Zivot and Andrews (1992) developed a data dependent algorithm that regards each data point as a potential break-date and runs a regression for every possible break-date sequentially. The test involves running three regressions (models): model A which allows for a one-time change in the intercept of the series; model B which permits a one-time change in the slope of the trend function; and model C which combines a one-time structural break in the intercept and trend (Waheed et. al., 2006). Following the lead of Perron, most investigators report estimates for either models A and C, but another relatively recent study Sen (2003) has shown that the loss in test power ($1-\beta$) is considerable when the correct model is C and researchers erroneously assume that the break-point occurs according to model A. On the other hand, the loss of power is minimal if the break date is correctly characterized by model A but investigators erroneously use model C. Because of this, Table 3 reports the Zivot-Andrews (ZA) one-break unit root test results for model C in level form along with the endogenously determined one-time break date for each time series.

The estimates reported in Table 3 for the series in level form are consistent with those in Table 2. For all of the series in question, Table 3 shows that the null hypothesis with a structural break in both the intercept and the trend cannot be rejected at the 5 percent level of significance.[16] Also, the Z-A test identifies endogenously the single most significant structural break in every time series. In view of space constraints, Figures 1 and 2 below show, respectively, the endogenously determined break-date for the labor productivity (lprod) and public investment (lginv) series.

B. Two-Break Unit Root Test.

So far, only tests for the presence of a single endogenously determined structural break have been undertaken. Lee and Strazicich (2003) have developed a two-break minimum Lagrange Multiplier (LM) unit root test that shows that assuming erroneously that there is one structural break in the data when, in fact, there are two leads to a further loss of power. Moreover, the LM unit root test developed by Lee and Strazicich (LS) enables the investigator to properly test for structural breaks under *both* the null and alternative hypotheses, thus eliminating size distortions that lead to the over-rejection of the null hypothesis of a unit root. This study, therefore, performed two-break unit root tests for all the variables in level form utilizing the LSUNIT (lags =1, model=crash, breaks=2) command in Rats 9.2. It was determined that the null hypothesis of a unit root under two endogenously determined structural (intercept) breaks could not be rejected at either the 1 or 5 percent level of significance (with the possible exception of the education variable which is marginally significant at the 5 percent level, see Table 4 below). The (LS) test was also performed with two endogenously determined structural intercept *and* trend breaks (lags =1, model = break, breaks=2) and the results are consistent with those reported for the intercept breaks (the education variable is no longer significant at the 5 percent level). (These results are available upon request.) Thus, the more powerful (LS) unit root test strongly suggests that all the included variables are I(1), by and large, consistent with the reported estimates for the (ZA) procedure. To save space, the (LS) Two-break (Bs) unit root results for the log of real GDP (Y), the log of productivity (y), the log of the ratio of private capital formation to GDP (i_p), the log of the ratio of public investment to GDP, and the log of the ratio of educational expenditures to GDP are reported in Table 4 below.

C. Cointegration Analysis.

[16] The Z-A one-break point unit root test was also performed for the relevant time series in differenced form under the assumption of model C and the null hypothesis was rejected at the 5 percent level or lower in all cases.

Given that the variables are integrated of order one, I(1), it is necessary to determine whether there is at least one linear combination of these variables that is I(0). Does there exist a stable and non-spurious (cointegrated) relationship among the regressors in each of the relevant specifications? This was done by first using the cointegration method proposed by Johansen and Juselius (1990). The Johansen method was chosen over the one originally proposed by Engle and Granger (1987) because it is capable of determining the number of cointegrating vectors for any given number of non-stationary series (of the same order); it also has a well-defined limiting distribution (see Harris, 1995).

Table 5 below shows that the Johansen test for both the output and labor productivity equations show that the null hypothesis of no cointegrating vector can be rejected at least at the five percent level; i.e., there exists a unique linear combination of the I(1) variables that links them in a stable and long-run relationship.[17] The signs of the cointegrating equation are reversed because of the normalization process and they suggest that, in the long run, the private and government investment variables have a positive and highly significant effect on Argentine labor productivity. The relatively high private capital (investment) elasticity reported in Table 4 is consistent with the extant empirical literature for developing (and developed) countries, and may be explained by FDI-induced or educational externalities in the form of better managerial know-how and the transfer of superior technology that "inflate" the private investment elasticity estimate by a positive factor θ (see DeMello, Jr., 1997). For example, a *ceteris paribus* 10 percent increase in the ratio of private investment to GDP raises output per worker by an estimated 5.6 percent in the long run. Admittedly, the relatively high coefficient for the private investment variable may be due to measurement error, omitted variables such as human capital, and/or simultaneity bias.

D. Cointegration Analysis with Structural Breaks.

Before turning to the EC models, it should be noted that the cointegrating test performed in this study does not allow for structural breaks in the sample period, whether level (intercept) shifts or regime (intercept and slope) shifts. However, Gregory and Hansen (1996) have shown that ignoring these breaks reduces the power of conventional cointegration tests similar to conventional unit root tests and, if anything, should lead to a failure to reject the null hypothesis of no cointegrating vector, which is not the case in the present study. This study, therefore, undertook a confirmatory G-H cointegration test with level *and* regime shifts and the results are consistent with the Johansen test. The G-H test with endogenously determined level (intercept) shift generated a minimum ADF* stat. = -6.027 [break point=2005] which is smaller than the tabulated 5 % critical value [-6.05 (1%); -5.57(5%); -5.33(10%)] reported by Gregory and Hansen. Thus, the null hypothesis of no cointegration with endogenously determined level break is rejected at the 5 percent level of significance. A regime test (intercept and slope) was also performed and the minimum ADF* stat. = -6.312 [break point= 2004] was essentially the same as the tabulated 5% critical value [-6.890 (1%); -6.320 (5%); -6.16 (10%)], thus leading to rejection of the null hypothesis once again. It should be noted that the break dates are found by estimating the cointegrating relationship for all possible break dates in the sample period. The Rats program selects the break date where the modified [trimmed] ADF* = inf ADF (τ) test statistic is at its minimum. The optimal lags for the G-H test were determined via the AIC and general–to-specific criteria.

E. Error Correction Models.

The lagged residual (error correction (EC) term) from the cointegrating equation, measuring the deviation between the current level of output (labor productivity) and the level based on the long-run relationship, was included in a set of EC models. For simplicity, consider the EC model without lags (and dummy variables) given in equation (6) below:

[17] The dummy variables were treated as exogenous variables in the cointegration test. The variables in question are also cointegrated with the inclusion of the export variable. The trace statistic in this case is equal to 88.56 and the null hypothesis of no cointegration is rejected at the 5 percent level (critical value= 76.972). The Max-eigen value test also indicates one cointegrating vector [39.977 > 34.80 (5% critical value)].

$$\Delta y = \alpha + \beta_1 \Delta l + \beta_2 \Delta(i_p) + \beta_3 \Delta(i_g) + \beta_4 \Delta(e_g) + \beta_5 \Delta(i_f) + \beta_6 \Delta(x) + \delta EC_{t-1} + \varepsilon \qquad (9)$$

The coefficients (β=s) of the changes in the relevant variables represent short-run elasticities, while the coefficient, δ (< 0), on the lagged EC term obtained from the cointegrating equation in level form denotes the speed of adjustment back to the long-run relationship among the variables. To conserve space, Table 6 below presents results only for the labor productivity growth rate relationship.[18]

The results for eqs. (1)-(4) (for the longer period without the FDI variable but with the inclusion of the export variable) suggest that changes in employment growth have an (expected) negative impact on the growth rate in labor productivity. The estimates for the private and public investment ratio variables suggest that they have a positive and statistically significant effect when lagged two to four periods.[19] This result is not altogether surprising because the positive externalities generated from additions to the stock of roads, bridges and ports are likely to affect labor productivity with a lag. The estimate for the lagged total education ratio variable also has an anticipated positive and statistically significant effect on the rate of labor productivity growth (see eq. 2). This is particularly the case for secondary and tertiary education (see the economically significant estimates in eqs. (3) - (4)). The estimate for primary education was positive as well but not significant at the 5 percent level (available upon request). Turning to the export variable, it too has a positive and consistently significant lagged impact in eqs. (1) – (4), which is consistent with the export promotion hypothesis. The estimates for the dummy variables in eqs. (2) - (5) suggest that the economic and financial crises that have hit Argentina have had a highly adverse effect on labor productivity growth, while the implementation of the currency board had a highly positive and significant impact.

The lagged EC terms are negative and statistically significant, suggesting, as in equation (3), that a deviation from long-run labor productivity growth this period is corrected by 13 percent in the next year. The results in Table 6 are also robust to the exclusion and inclusion of the dummy variables. Finally, the EC regressions were tested for specification error such as omitted variables and/or functional form via Ramsey's Regressions Specification Error Test (RESET) and we were unable to reject the null hypothesis of no specification error at the 5 percent level of significance (results available upon request).

Turning to the results with the FDI variable in eq. (5), they suggest that inflows of FDI have a positive (lagged) and significant effect on labor productivity growth. (The export variable was excluded from these regressions because it is highly correlated with inward FDI flows, with a simple correlation coefficient of 0.854, thus essentially capturing the same effect.) The other variables in eq. (5) retained their economic and statistical significance. Again, these results have to be taken with caution given the shortened period of estimation.

The EC models were also used to track the historical data on labor productivity growth in Argentina. Table 7 below reports selected Theil inequality coefficients obtained from historical simulations of the productivity growth equations (2) and (5). In general, the predictive power of the model is considered to be relatively good if the coefficient is at or below 0.3 [Theil, 1966]. The results reported in Table 7 meet this performance criterion, particularly for eq. (2) (the root mean squared errors (RMS) are relatively low as well). The sensitivity analysis on the coefficients shows that changes in the initial or ending period did not alter appreciably the predictive power of eq. (2) (it was not possible to conduct a similar analysis for eq. (5) because of insufficient data points). Figures 3 and 4 corresponding to equations (2) and (5), respectively, provide visual evidence of the models' ability to track the turning points in the actual series. (DLPROD refers to the actual data and DLPRODF denotes the forecast.) They show that the rate of labor productivity growth was, in general, positive during the decade of the nineties, highly erratic in the seventies, and mostly negative during the "lost decade" of the eighties. In fact, during the first half of the nineties there was a sharp upward turn in output (labor productivity) growth, punctuated by a sharp drop

[18] The EC regressions for the output equation are essentially the same as those for the labor productivity regressions (except for the reversal in sign of the labor force variable) because they are a parametric transformation of one another

[19] The order of the lag length was determined by applying both the AIC and SIC criteria.

in 1995 as a result of the "tequila effect" associated with the Mexican peso crisis of 1994-95, followed, in turn, by three years of positive growth, only to culminate in a sharp contraction during the economic crisis years of 1999-2002.

Figures 3 and 4 also show that since 2003 there has been an upward surge in labor productivity growth (except for the recession years of 2009 and 2014) associated with both the administrations of Nestor Kirschner (2003-2007) and Cristina Fernandez de Kirschner (2007-2015). Weisbrot and Sandoval (2007) attribute this favorable turn of events to some factors, not the least of which is the abandonment of the currency board, which had become a "strait-jacket with regard to monetary policy," and the adoption of a stable and competitive real exchange rate which has stimulated both the growth of exports and import-competing industries. Also, they contend that the government's adoption of unorthodox (pro-growth) policies, in the form of an accommodating monetary policy and a boost in public investment spending, have stimulated both internal demand and private capital formation (see Table 1).

Finally, Weisbrot and Sandoval emphasize the Kirschner administration's firm stance vis-à-vis the IMF in negotiating and restructuring Argentina's defaulted external debt in 2005, which significantly reduced the country's debt-service ratio from 52.2 percent of GDP in 2005 to 36.9 percent in 2008, thus freeing up scarce resources for its pro-growth policies (including public investment in economic and social infrastructure. In this connection, Table 1 shows a significant rise in this variable as a proportion of GDP during the 2006-2015 period) (see Weisbrot, 2011, pp. 9-11; and ECLAC 2018). Unfortunately, under the current neoliberal administration of Mauricio Macri (2015-19), the hard-won gains of the previous administrations were at first eroded, and then reversed during the current crisis. For example, there has been a dramatic rise in Argentina's gross public debt as a proportion of GDP (up sharply 20.3 percentage points in 2018 to 77.4 percent of GDP!), private investment has plummeted as a proportion of GDP and is forecast at just 8.5 percent of GDP in 2019—the lowest level since the crisis years of 2001-02. Last but not least, public capital expenditures as a percent of GDP have plummeted from 2.7 percent in 2015 to an estimated 1.5 percent in 2018 (see ECLAC, 2018, p.136; and Weisbrot and Merling, 2018, pp. 9-10).

5 Vector Error Correction Model

In this section we utilized a VECM framework because it allows investigators to model the short-run correction mechanism of a system of variables to their long-run equilibrium without deciding, a priori, about the endogeneity or exogeneity of the included variables. The vector autoregressive framework treats all variables as endogenous and determines the direction of causality between them based on econometric tests instead of assuming exogeneity based on preconceived assumptions. The flexibility of this specification represents a distinct advantage in modeling the system-wide error correction mechanisms of the relevant variables. VECM specifications can only be estimated if there is a cointegrating relationship among the variables which, according to the Johansen test, is indeed the case in our study.

First, an unrestricted small-scale VECM is estimated using ln(Y/L), lnL, $\ln I_p$, and $\ln I_g$. variables to examine whether a system-wide error correction mechanism exists for these variables. The VECM is estimated based on Model 4 with 2-3 lags (based on AIC and SBC criteria) to allow for the impact of real private and public investment expenditures which, a priori, are likely to have a delayed effect on labor productivity.[20] The export, education, and foreign investment variables could not be included in the system due to the limited number of observations and insufficient degrees of freedom to generate the relevant estimates. The unrestricted VECM finds a negative and highly significant adjustment coefficient for one of the four equations, viz., D(lnY/L)[LPROD]; and it finds a positive and significant coefficient for D(lnL)[LLABOR]. This means that for the labor productivity equation there is a short-run adjustment mechanism back to the equilibrium relationship when shocks to the system are sustained; however, this is

[20] Model 4 (with an intercept and trend) is chosen based on the Schwarz Criteria by Rank (rows) and Model (columns). The VECM was also estimated with dummy variables D1 and D2 as defined in the text. These variables are assumed to be exogenous.

not the case for the labor variable which is positive and significant. The labor productivity specification is the best, based on the adjusted R-squared values and the AIC/SBC criteria, that is, when treating D(lnY/L) as the "dependent variable". In this specification, the long-run cointegration equation has significant coefficients for all the variables and is consistent with the results obtained for the Johansen test in the previous section.

Turning to the short-run EC specification for D(lnY/L) [LPROD], it can be readily seen that the error correction term is negative and highly significant, implying that a 10 percent shock (deviation) away from the long-run equilibrium in the current quarter is corrected by 1.1 percent in the subsequent period, which is consistent with the results reported in Table 8 below. The short-run estimates for this equation also suggest that percentage changes in lnI_g [LGINV]"crowd-in" labor productivity when lagged between one and three years. Insofar as percentage changes in labor productivity are concerned, the short-run estimates suggest that it has a positive and highly significant effect on itself and private investment expenditures when lagged three years. Finally, percentage changes in the labor force have a positive effect on private investment when lagged two to three years.

On the other hand, treating D(lnL)[LLABOR] as the "dependent variable," it can be readily ascertained from Table 8 that the error correction term is positive and significant at the 5 percent level, implying that a 10 percent deviation away from equilibrium during the current year generates a subsequent movement away from equilibrium of 0.49 percent. Changes in D(lnI_p) [LPRIV] have a positive and significant effect on its current values and the labor variable when lagged two to three periods, while percentage changes in labor have a negative and significant effect on labor productivity when lagged one year, but no statistically significant effect on the other variables, including its values.

A very useful property of the VECM framework is that it enables the investigator to impose zero restrictions on the adjustment coefficients of each equation, thus determining which variables can be treated as weakly exogenous in the system, thereby omitting them from the interdependent system of variables. Based on this weak exogeneity test, Table 9 below indicates that the private investment and public investment variables can be omitted from the system (treated as weakly exogenous) because the null hypothesis of a zero restriction is not rejected for these variables at least at the 5 percent level. In other words, in this simple four equation system, we can treat lnY/L and lnL as endogenous variables, while lnI_p and lnI_g are weakly exogenous.

To investigate further the "causal" relationship among these variables, we performed a Granger Block Causality test without the restrictions and three lags. This test examines all four equations and tries to determine whether the presumed exogenous variables can be omitted from each equation individually or as a group. This test finds "causality" or precedence in the D(lnY/L) equation at the 5% level for all variables taken together (p-value: 0.0267), and for D(lnI_p) and D(lnI_g) individually (p-value$_{DlnIp}$:0.006, p-value $_{DlnIg}$: 0.007). This finding is consistent with the results reported above. The test finds the presence of weak "causality" or precedence in the D(lnL) equation for the percentage change in private investment (p-value $_{DlnIp}$: 0.110). It also finds

some evidence of reverse causality or weak precedence in the D(lnI_p) equation from percentage changes in the labor productivity variable (p-value $_{DlnY/L}$: 0.112). Finally, the remaining relationships in the block are not statistically significant.

Based on this important information, we assessed visually the dynamic interactions among the variables in the system by appealing to impulse response functions for the variables in our simple system. This study employed a generalized decomposition process first proposed by Pesaran and Shin (1998) because it constructs an orthogonal set of innovations that do not depend on the VAR ordering. Figure 5 below shows the generalized impulse responses of the four variables in question to both a unitary shock in their values and the rest of the variables over 10 years period. It can be readily ascertained that the response of lnY/L [LPROD] to one standard deviation (SD) innovation in both private [LPRIV] and public investment [LGINV] spending is significant and sustained (particularly after two to three periods); on the other hand, the reverse line of "causation" is not as strong. Figure 5 also shows that the lagged response of private investment to a one (SD) innovation in public investment is positive and sustained (again, after two to three periods), while the reverse is not as strong. Interestingly, and somewhat unexpectedly, the response lnL [LLABOR] is relatively strong to both one standard innovation in public investment and one standard innovation in private investment, but this effect is not captured in the Granger Block causality test reported above. Finally, the response of each variable to its one SD innovation is relatively strong and sustained over the period under review.[21]

Overall, based on the analysis of the unrestricted VECM system, we can conclude that short-run deviations of the included variables from their long-run cointegrated relationship are corrected in subsequent years for the labor productivity equation in the system. The restricted VEC models suggest that there are two weakly exogenous variables in the system, that is, not all variables can be treated as endogenous. In line with this finding, the Block Granger causality test finds weak two-way "causality" between D(lnY/L) and D($\ln I_p$), and one-way causality or precedence from public investment spending and the labor variable. In general, the short-run impulse response functions are somewhat consistent with the VECM and Granger Block exogeneity tests, except for the labor variable equation.

6 Conclusion

This paper presented a simple model that explicitly includes the impact of the public (or FDI) capital stock on both the supply and demand sides of a hypothetical economy. The discussion showed that if significant complementarities are present between public (or FDI) and private capital (i.e., if a positive externality is present), then diminishing returns to the private inputs can be prevented or postponed indefinitely. The conceptual model laid the groundwork for the empirical analysis of labor productivity growth in the Argentine case for the 1960-2015 period in Sections III and IV. Several key findings were obtained.

First, Zivot-Andrews and Lee-Stratizicich unit root tests in the presence of single and two-time breaks indicate, respectively, that the null hypothesis of non-stationarity cannot be rejected for the relevant series in level form, but can be rejected in first differences. This represents a significant contribution to the extant literature which does not address the low power of conventional unit root tests in the presence of structural breaks. Second, the Johansen and Gregory-Hansen cointegration tests revealed that the null hypothesis of no cointegration can be rejected at the five percent level, thus suggesting that the I(1) variables have a unique and stable relationship that keeps them in proportion to one another in the long run even in the presence of an endogenously determined structural regime (intercept and trend) break. This is a highly important finding because previous empirical studies have applied the OLS method directly to nonstationary variables in level form, thus generating spurious or miss-specified regressions. Third, the cointegrating equations were used to generate a set of EC models that reconcile both the short and long-run properties of the variables included in the output and labor productivity relationships. As the theory predicts, the EC models have negative and statistically significant error correction terms, suggesting that deviations from long-run labor productivity (output) growth are corrected in subsequent periods.

Fourth, the individual EC estimates indicated that the growth rate of private and public investment as a proportion of GDP, as well as the growth rate in exports, education ratio, and the FDI ratio, have a positive and statistically significant effect on the growth rate of labor productivity, while the growth rate

[21] Accumulated responses to generalized one S.D. innovations were also estimated for the included variables and are available upon request.

in the labor force has a negative impact. The VECM results for the small-scale model that treats all variables as endogenous also report~~s~~ estimates that are consistent with the EC estimates in Table 6. Fifth, the reported Theil inequality coefficients for the selected EC models suggested that they were able to track and simulate the turning points of the historical series in labor productivity relatively well.

Finally, the EC model estimates showed that during the decade of the nineties the rate of labor productivity growth was mostly positive, while during the decade of the seventies the annual estimated rate of output growth became erratic, culminating in a marked decrease (often negative rates) during the decade of the eighties--the so-called “lost decade of development.” The labor productivity growth estimates for the first half of the nineties did reveal a robust increase, thereby suggesting that the currency board’s taming of inflationary pressures and the opening of the economy to foreign direct investment had a positive effect. During the second half of the 2000s, there was an upsurge in labor productivity growth which coincided with the promotion by both Kirschner administrations of broad-based pro-investment and pro-growth policies. For example, there was a significant increase in public investment as a proportion of GDP, averaging 3 percent during the 2007-2015 period—more than triple its average during the 2000-2006 period. Private investment as a proportion of GDP also rose during the 2007-2015 period. Unfortunately, as indicated above, many of the hard-won economic and social gains of both Kirschner administrations were at first eroded and then reversed under the neoliberal administration of Mauricio Macri (2015-2019).

From a policy standpoint, the findings in this paper are important because they suggest that cash-strapped governments of Latin America, such as the Argentine one, can maximize the growth potential of their economies by directing scarce resources to investments in economic and social infrastructure, particularly education at the secondary and tertiary levels. The findings also suggest that attracting “bolted down” capital in the form of FDI inflows, as well as promoting exports, are likely to have a beneficial effect on labor productivity growth. These investments, through a positive externality effect, are likely to increase the marginal productivity of the private inputs directly (as well as indirectly), thereby increasing private investment, output, and labor productivity.

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Table 1: Argentina: Investment as a Share of GDP (in percent), 1960-2015

Year	Private Investment	Public Investment
1960	11.3	9.7
1970	13.1	8.1
1980	19.2	6.1
1990	9.4	4.6
1992	14.9	1.8
1994	19.1	0.8
1996	16.1	2.0
1998	17.9	2.0
2000	15.4	1.0
2002	9.2	0.7
2004	10.5	1.3
2006	13.2	2.5
2008	15.1	3.3
2010	14.9	3.5
2011	14.1	2.8
2012	13.8	3.0
2013	13.7	3.0
2014	13.5	3.1
2015	13.6	3.1
Simple Average		
1960-1969	11.0	8.9
1970-1979	13.6	9.1
1980-1989	15.0	4.9
1990-1999	15.7	1.6
2000-2010	12.7	2.1
2011-2015	13.7	3.0

Source: IFC, Trends in Private Investment in Developing Countries, Statistics for 1970-2000. Washington, D.C., The World Bank, 2016; M.E.P., Argentina: Sustainable Output Growth After the Collapse. Buenos Aires, Ministerio De Economia Argentina

Table 2: PART A. Unit Root Tests for Stationarity, Sample Period 1960-2015

Variables	Levels	First Difference	5% Critical Value[1]	1% Critical Value
ln(Y)	-0.48	-5.81**	-2.92	-3.56
ln(Y/L)	-2.21	-5.53**	-2.92	-3.56
lnL	0.72	-6.46**	-2.92	-3.56
$\ln I_p$	-0.43	-6.47**	-2.92	-3.56
$\ln I_g$	-1.55	-7.10**	-2.92	-3.56
$\ln e_g$	-1.07	-6.53**	-2.92	-3.56
$\ln I_f$ [2]	-1.79	-8.30**	-2.92	-3.56
lnX	-0.94	-6.49**	-2.92	-3.56

PART B. KPSS (LM) No Unit Root Tests for Stationarity with constant,1960-2015.

Variables	Levels	First Difference	5% Critical Value
lnY	0.869*	0.083	0.463
ln(Y/L)	0.451	0.098	0.463
lnL	0.892*	0.205	0.463
$\ln I_p$	0.854*	0.152	0.463
$\ln I_g$	0.654*	0.094	0.463
$\ln e_g$	0.699*	0.066	0.463
$\ln I_f$	0.739*	0.055	0.463
lnX	0.889*	0.081	0.463

[1]MacKinnon critical values for rejection of hypothesis of a unit root. [2]Unit root test for the FDI variable undertaken for the 1969-2015 period. *Significant at the 5 percent level; **Significant at the 1 percent level. Asymptotic critical values for rejection (LM-Stat.)

Table 3: Zivot-Andrews One-break Unit Root Test, Sample Period 1960-2015

Variables	Levels	Break Year	5% Critical Value[a]	1% Critical Value
Ln(Y)	-3.27	1981	-5.08	-5.57
In(Y/L)	-3.81	1985	-5.08	-5.57
lnL	-4.32	1995	-5.08	-5.57
$\ln I_p$	-4.49	1981	-5.08	-5.57
$\ln I_g$	-4.62	1991	-5.08	-5.57
$\ln e_g$	-4.02	1992	-5.08	-5.57
$\ln I_f$	-5.70**	1969	-5.08	-5.57
lnX	-4.03	1973	-5.08	-5.57

Estimations undertaken with Eviews 11.0. One-break unit root test for the FDI variable undertaken for the 1969-2015 period.

Table 4: Lee-Strazicich Two-Break Unit Root Test, 1960-2015

Variable	Coefficients	T-ratios	1% cv	5% cv
SY(-1)	-0.179	-2.66	-4.073	-3.563
Constant	0.032	5.12	---	---
B1: 1977	-0.055	-1.22	---	---
B2: 1984	-0.086	-1.92	---	---
Sy(-1)	-0.139	-2.37	-4.073	-3.563
Constant	0.019	2.96	---	---
B1: 1977	-0.046	-1.17	---	---
B2: 1984	-0.073	-1.91	---	---
Si_p (-1)	-0.343	-3.24	-4.073	-3.563
Constant	0.012	0.05	---	---
B1: 1968	0.359	2.14	---	---
B2: 2002	-0.384	-2.07	---	---
Si_g (-1)	-0.251	-2.70	-4.073	-3.563
Constant	-0.058	-1.71	---	---
B1: 1990	-0.887	-3.99	---	---
B2: 2002	-0.455	-1.95	---	---
Se_g(-1)	-0.463	-3.52	-4.073	-3.563
Constant	0.013	0.61	---	---
B1: 1984	-0.618	-4.26	---	---
B2: 1991	0.480	3.28	---	---

Notes: The coefficients on the SY(-1),Sy(-1),Si_p(-1),Si_g(-1), and Se_g are lagged de-trended variables test for the presence of a unit root. B1 and B2 are the endogenously determined breaks in the intercept for the sample period. The estimations were undertaken with Rats 9.2.

Table 5: Johansen Cointegration Rank Test (Trace), 1960-2015

A. Series: lnY, lnL, LnI_g , and lnI_p.

Test assumption: No Linear deterministic trend in the data.

Eigenvalue	Likelihood Ratio	5% Critical Value	No. of CE(s)
0.404	57.359	54.08	None
0.305	29.378	35.19	At most 1
0.123	9.706	20.26	At most 2
0.047	2.614	9.17	At most 3

B. Series: ln(Y/L), lnL, lnI_g , and lnI_p.

Test assumption: no linear deterministic trend in the data.

Eigenvalue	Likelihood Ratio	5% Critical Value	No. of CE(s)
0.404	57.359	54.08	None
0.305	29.378	35.19	At most 1
0.123	9.706	20.26	At most 2
0.047	2.614	9.17	At most 3

Normalized Cointegrating Vector; coefficients normalized on ln(Y/L) in parenthesis.

Vector	ln(Y/L)	lnL	lnI_g	lnI_p	Constant
1.	1.000	0.147	-0.055	-0.271	-1.247
		(0.078)	(0.019)	(0.045)	(0.440)

Note: Standard errors are in parenthesis. Signs in cointegrating vector are reversed due to normalization.

Table 6: Dependent Variable is: (ΔlnY_t - ΔlnL_t), 1960-2015, OLS Regressions

Variables	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**
Constant	0.01 (1.39)*	0.003 (0.42)	0.003 (0.41)	-0.001 (-0.14)	0.006 (0.98)
ΔlnL_t	-0.33 (-1.95)**	-0.35 (-3.36)**	-0.49 (-4.17)**	-0.42 (-3.87)	-0.27 (-1.64)*
ΔlnI_{pt-3}	0.19 (5.69)**	0.05 (3.33)**	0.05 (2.54)**	0.05 (2.75)**	0.04 (2.34)**
ΔlnI_{gt-2}	0.02 (2.01)**	0.03 (2.86)**	0.03 (4.19)**	0.03 (3.10)**	0.03 (2.92)**
ΔlnI_{ft-3}	---	---	---	---	0.02 (2.29)**
ΔlnX_{t-3}	0.08 (3.04)**	0.07 (2.51)**	0.08 (2.29)**	0.08 (3.28)**	---
Δlne_{gt-2}	0.04 (2.59)**	0.04 (3.87)**	---	---	0.04 (2.68)**
$\Delta lnse_{gt-2}$	---	---	0.30 (2.08)**	---	---
$\Delta lnte_{gt-2}$	---	---	---	0.15 (2.06)**	---
ECT_{t-1}	-0.15 (-1.90)**	-0.11 (-2.36)**	-0.13 (-2.20)**	-0.12 (-2.19)**	-0.11 (-2.36)**
DUM1	---	-0.05 (-2.38)**	-0.04 (-5.23)**	-0.04 (-5.79)**	-0.04 (-5.97)**
DUM2	---	0.05 (7.15)**	0.05 (6.32)**	0.06 (7.49)**	0.04 (5.07)**
Adj R^2	.50	.75	.74	.73	.70
S.E.	.032	.023	.025	.025	.024
D.W.	1.99	1.95	1.99	1.96	2.08
AIC	-3.55	-4.44	-4.29	-4.31	-3.73
SIC	-3.74	-4.03	-3.84	-3.89	-3.99

Note: Figures in parentheses are t-ratios. AIC denotes Akaike Information Criterion, SIC is the Schwarz Information Criterion, and S.E. is the standard error of regression.*10%; **5%.

Table 7: In-Sample Forecast Evaluation for Error Correction Models

	Equation (2)	**Equation (5)**
	Sample: 1960-2015	**Sample: 1970-2015**
Root Mean Squared Error (RMS)	0.022	0.025
Mean Absolute Error (MAE)	0.016	0.019
Theil Inequality Coefficient (TIC)	0.281	0.310
Bias Proportion (BP)	0.011	0.000
Variance Proportion (VP)	0.041	0.081
Covariance Proportion (CP)	0.947	0.911
	Sample: 1960 1999	
RMS	0.022	---
MAE	0.018	---
TIC	0.286	---
BP	0.000	---
VP	0.071	---
CP	0.928	---
	Sample: 1970 2015	
RMS	0.023	---
MAE	0.018	---
TIC	0.294	---
BP	0.021	---
VP	0.094	---
CP	0.884	---

Note: In-sample forecast evaluation estimates generated with EVIEWS 11.0

Table 8: Vector Error Correction Results

	D(In(Y/L)	D(lnL)	D(lnI_p)	D(lnI_g)
Error Correction Coefficient	-0.109	0.0497	-0.0137	0.0347
Standard Error	(0.022)	(0.016)	(0.117)	(0.272)
t-statistics	[-4.858]	[3.067]	[-0.117]	[0.127]
R-squared	0.748	0.547	0.585	0.100
Adj. R-squared	0.635	0.358	0.412	-0.307
Akaike AIC	-4.195	-4.836	-0.873	0.808
Schwarz SBC	-3.954	-4.236	-0.272	1.409

Table 9: Exogeneity Test

Ho: weakly exogenous variable	Chi-Square statistics	Probability
D[ln(Y/L)], A(1,1)=0	20.100	0.000
D(lnL), A(2,1)=0	9.522	0.002
D(LnI_p), A(3,1)=0	0.018	0.894
D(lnI_g), A(4,1)=0	0.015	0.902

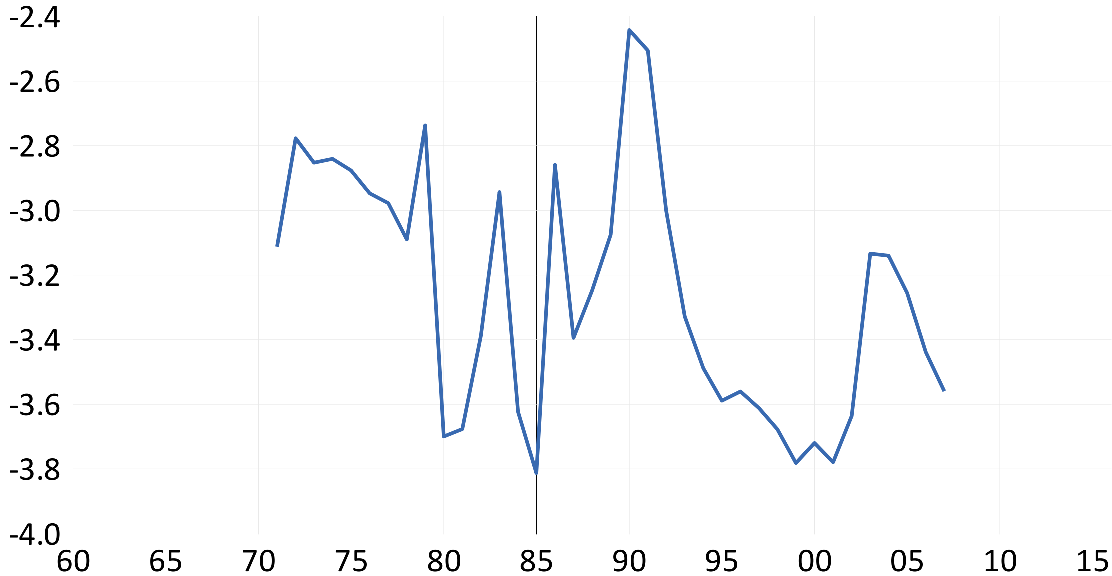

Figure 1: Zivot-Andrew Breakpoints

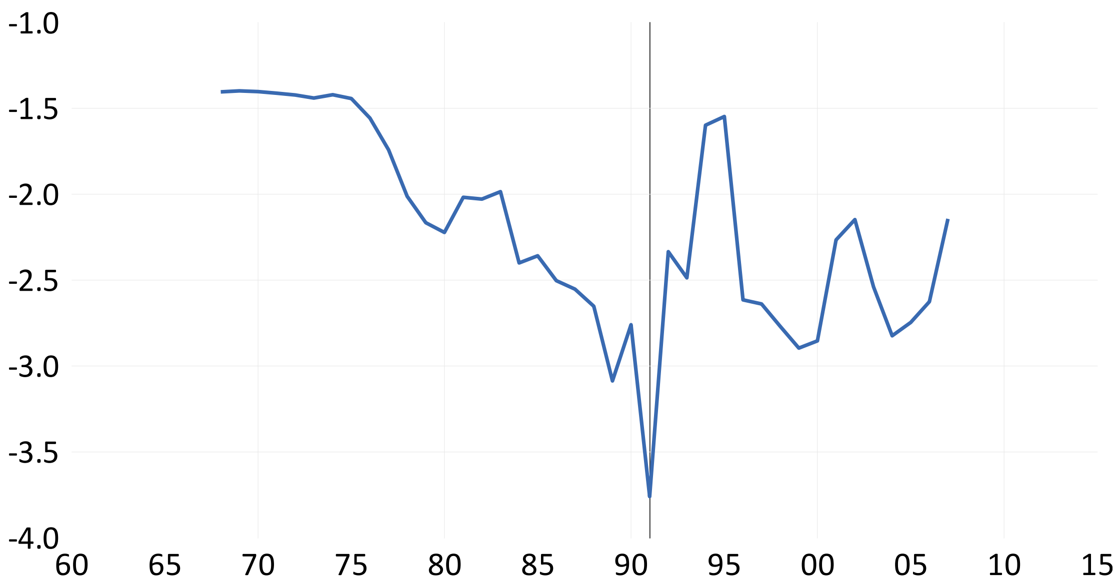

Figure 2: Zivot-Andrew Breakpoints

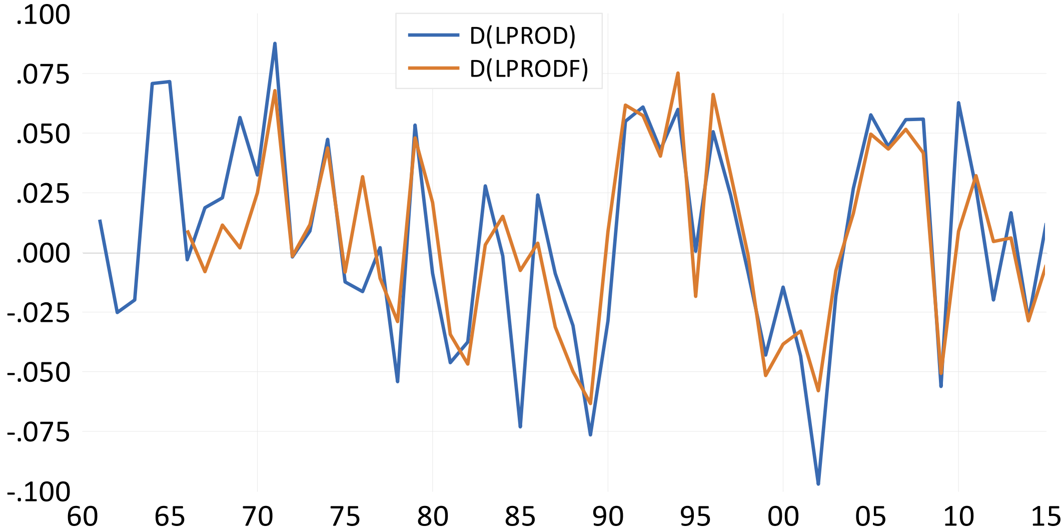

Figure 3: Labor Productivity Growth: Actual (DLPROD) and Forecast (DLPRODF), 1960-2015.

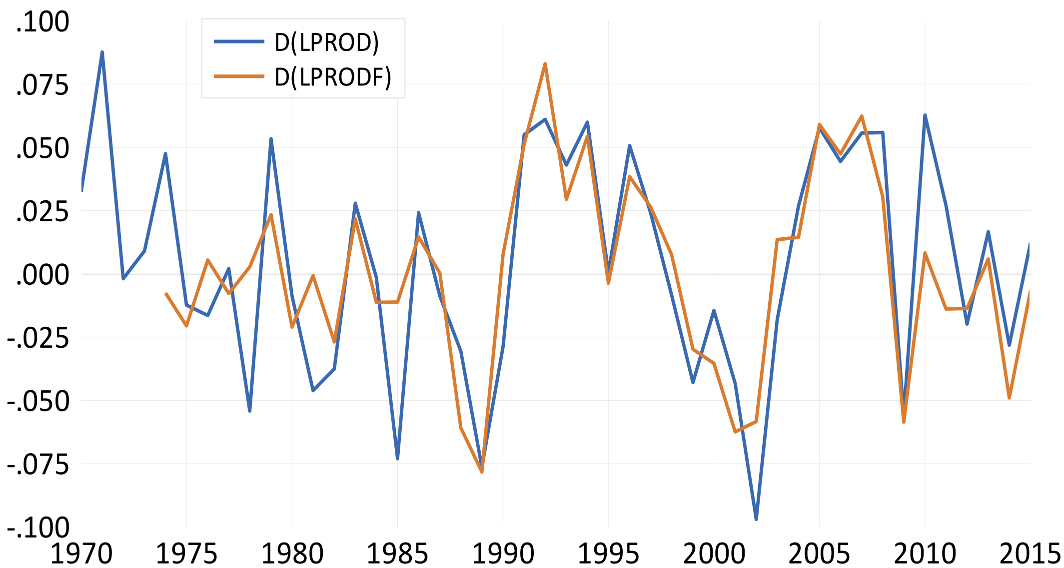

Figure 4: Labor Productivity Growth: Actual (DLPROD) and Forecast (DLPRODF), 1970-2015.

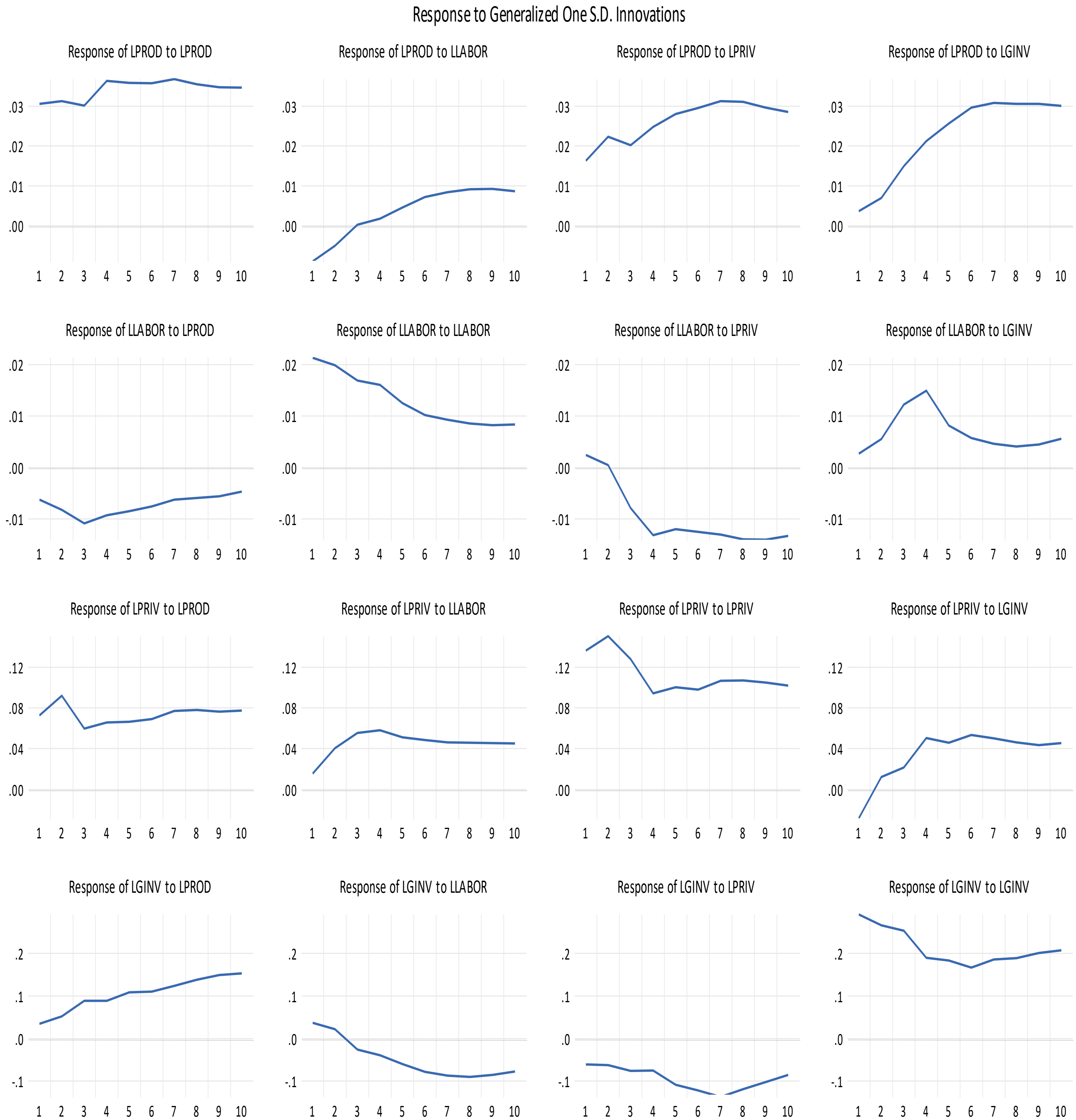

Figure 5: Impulse Response Functions