TIME-SERIES PROPERTIES AND EMPIRICAL EVIDENCE OF GROWTH AND INFRASTRUCTURE*

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Abstract

After more than forty years studying growth, there are two classes of growth models that have emerged: exogenous and endogenous growth models. Since both try to mimic the same set of long-run stylized facts, they are observationally equivalent in some respects. Our goals in this paper are twofold. First, we discuss the time-series properties of growth models in a way that is useful for assessing their fit to the data. Second, we investigate whether these two models successfully conform to *post-war* U.S. data. We use cointegration techniques to estimate and test long-run capital elasticities, exogeneity tests to investigate the exogeneity status of TFP, and Granger-causality tests to examine temporal precedence of TFP with respect to infrastructure expenditures. The empirical evidence is robust in confirming the existence of a unity long-run capital elasticity. The analysis of TFP reveals that it is not weakly exogenous in the exogenous growth model. Granger-causality test results show unequivocally that there is no evidence that TFP for both models precede infrastructure expenditures not being preceded by it. On the contrary, we find some evidence that infrastructure investment precedes TFP. Our estimated impact of infrastructure on TFP lay roughly in the interval (0.19, 0.27).

*We gratefully acknowledge the comments and suggestions Heather Anderson, Steve de Castro, Wouter de Haan, Rob Engle, Alvaro Escribano, Jesus Gonzalo, Clive Granger, Paul Lau, Ajax Moreira, Marcelo Portugal, Valerie Ramey, Chor-Yiu Sin, Pedro Valls Pereira, two anonymous referees and seminar participants at UC San Diego, Getulio Vargas Foundation, IPEA-Rio, and PUC-Rio, and conference participants at ESEM 1994 and SBE 1994. All remaining errors are ours. We thank Laura Almeida, Alessandra Gouvêa, and Ana Carolina Letichevsky for excellent research assistance, and CNPq and Pronex for financial support.

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Resumo

Após mais de quarenta anos estudando-se modelos de crescimento econômico, dois modelos básicos dominam a literatura: os de crescimento endógeno e de crescimento exógeno. Dado que ambos tentam reproduzir os mesmos “fatos estilizados”, eles são similares do ponto de vista observacional em alguns aspectos. Aqui, discute-se as propriedades de séries de tempo de ambos esses modelos de forma a poder-se examinar sua aderência aos dados, que é investigada usando-se agregados macroeconômicos americanos no período desde o pós-guerra. Testes de cointegração são usados para investigar se a produtividade total dos fatores (TFP) é exógena, e testes de causalidade de Granger são usados para investigar se os gastos de infra-estrutura precedem temporalmente a TFP. Confirma-se inequivocamente a existência de uma elasticidade unitária de longo prazo para o estoque de capital na função de produção, o que confirma a aderência de ambos os modelos aos dados americanos. A análise usando-se a TFP confirma que esta não é exógena para o modelo de crescimento exógeno. Ademais, em testes de causalidade, concluiu-se que a TFP não precede os gastos de infra-estrutura, sendo possivelmente precedida por aqueles. O impacto (elasticidade) de longo prazo da infra-estrutura na TFP está aproximadamente no intervalo (0.19, 0.27).

Key words: Growth models; cointegration.

JEL Code: O40 and C52.

1. Introduction.

After more than forty years studying growth, there are two stylized classes of growth models that have emerged. The first is that of exogenous growth models, based on Solow (1957), Cass (1965) and Koopmans (1965), and the second is that of endogenous growth models based on Romer (1986, 1987, 1989), among others. The first basic difference among these two classes is the behavior of productivity. For the exogenous class, productivity is assumed to grow at an exogenous rate. This implies a balanced-growth path for macroeconomic aggregates, and thus a unity long-run capital elasticity in production. Because models in this tradition use the condition of decreasing returns to capital for the existence of equilibrium, the only

way to assure growth in steady-state is to require exogenous technological change. In the endogenous class of models this condition is dropped since it is thought to be too *ad hoc*. To assure growth, external effects due to social capital are introduced, although decreasing returns to private capital are still maintained. The externality hypothesis by itself does not guarantee growth, since it is still possible that the marginal product of capital (private plus social) will converge to zero. For example, with Cobb-Douglas technology, unless the long-run elasticity of capital is at least one there will be no growth in steady-state. Of course, if the long-run elasticity is above unity there will be explosive growth. Since this is not a stylized fact of modern economies, testing the externality hypothesis has focused on the case of unity elasticity. A consequence of a unity elasticity is that macroeconomic aggregates will follow a balanced-growth path. Regarding productivity, little or nothing is imposed in the class of endogenous growth models.

From the discussion above, it is clear that these two models are observationally equivalent regarding (i) a balanced-growth path for macroeconomic aggregates, and specifically (ii) a unity long-run capital elasticity; see Lau and Sin (1997a). Thus, estimating long-run elasticities cannot be a strategy in distinguishing which of these models best fits the data. Despite that, short-run capital elasticities differ for both models. For the endogenous class it is also one, but for the exogenous class it will usually be less than unity, reflecting the fact that there are decreasing marginal returns in equilibrium.

Since the observational equivalence result is relatively new, before it became known several authors have tried to confirm the theoretical predictions of a given model class by examining long-run elasticities; see Neusser (1991), who finds support for the exogenous-growth model using cointegration, and the cross-country study in Romer (1987), who finds a unity capital elasticity supporting the

endogenous-growth model. Another problem of econometric testing of growth models is that not all studies were careful in distinguishing between the short- and the long-run capital elasticity. For example, the time-series evidence in Romer (1987) and Benhabib and Jovanovic (1991) *inter-alia* rejects a unity capital elasticity. The data used in these regressions are either first- or quasi-differenced, which raises the issue of what coefficient is being estimated. On the one hand, if there is no cointegration in the data, one can estimate consistently the short-run capital elasticity from first-differenced regressions. Regarding quasi-differenced regressions, consistency will depend on common factors being correctly imposed; see Hendry and Mizon (1978). On the other hand, if there are long-run relationships, ignoring them yields inconsistent estimates of short-run elasticities in both cases.

There are several aspects of growth models that have to be taken into account for successfully assessing their fit. These issues have not been thoroughly discussed by the literature. Moreover, they have not been used appropriately when confronting theoretical models with the data. In this paper, we first set out what are the main time-series properties of growth models regarding: the order of integration of macroeconomic aggregates, the cointegration relationships implied by each class of models, and the order of integration and the exogeneity status of their respective Total Factor Productivity (TFP). Given these theoretical implications, we then use *post-war* U.S. data on output, and capital and labor inputs to assess the fit of long-run elasticities of a broader class of models, comprising endogenous and exogenous growth models. We later test TFP exogeneity, considering the potential influence of infrastructure investment; see Aschauer (1989). With that regard, we first investigate whether or not the Solow residual of the exogenous growth model is (weakly) exogenous. For the endogenous growth model, we investigate the

properties of a new Solow residual, taking into account a unity long-run capital elasticity. Finally, for the Solow residuals of both models, we estimate what is the long-run impact of infrastructure expenditures following Aschauer.

To estimate and test long-run elasticities we use cointegration techniques as proposed by Johansen (1988, 1991), where estimation and testing are all likelihood-based. TFP exogeneity tests are based on the typology of Engle, Hendry and Richard (1983), and performed according to Johansen (1995, 122–123). Granger-causality tests for TFP and infrastructure investment take the form of exclusion tests, where we consider the criticism of Toda and Phillips (1993) of such tests.

Section 2 presents a discussion of a stylized version of both theoretical models, showing the observational equivalence result. It also discusses the properties of TFP for both models. Section 3 presents a discussion of the recent applied literature using our previous theoretical results. Section 4 presents the econometric techniques used on the empirical section. Empirical results are presented in Section 5, and Section 6 concludes.

2. Theory.

2.1. Stylized growth models.

This section discusses two distinct classes of economic models: endogenous and exogenous growth models. We present here stylized versions of these models under very restrictive parameterizations of preferences and technology, which capture the essence of the long-run co-movement in the data we want to discuss. Although these stylized facts could be discussed under a less restrictive setting, we chose their restricted version for expositional reasons.

Under Cobb-Douglas technology, Romer's (1986, 1987, 1989) endogenous growth model can be characterized using a representative consumer who maximizes his/her flow of utility, choosing a sequence of consumption $\{C_t\}$ subject to a resource constraint:

$$\begin{aligned} & \text{Max}_{\{C_t\}} E_0 \sum_{t=0}^{\infty} \beta^t U(C_t), \\ & \text{s.t.} \\ & C_t + I_t = Y_t = A K_t^\alpha L_t^{(1-\alpha)} \bar{K}_t^\theta \exp(z_t) \\ & K_{t+1} = I_t + (1 - \delta) K_t \\ & 0 < \beta < 1, \end{aligned} \tag{2.1}$$

where Y_t is output, K_t is the stock of physical capital, L_t is hours worked, I_t is investment, δ is the depreciation rate of the capital stock, β is the one-period discount factor for future utility, $E_t(\cdot)$ is the conditional expectation operator, and $\{z_t\}$ is the stochastic process representing productivity. Aggregate capital per capita ($\bar{K}_t$) is included to represent the externalities due to the fact that knowledge is non-appropriable.

In equilibrium we have $\bar{K}_t = n \cdot (K_t/L_t)$, where n is the number of firms.¹ In per-worker terms the production function is given by $(Y_t/L_t) = A(K_t/L_t)^\alpha [n \cdot (K_t/L_t)]^\theta \exp(z_t)$. Hence, the (private) marginal productivity of capital is given by:

$$F_{K/L}(\cdot) = \alpha A n^\theta (K_t/L_t)^{(\alpha+\theta-1)} \exp(z_t).$$

After solving the representative consumer's problem² the conditional expectation of the growth rate of the economy (γ) is given

¹Romer (1989) assumes that the number of firms and consumers is the same, so that K_t/L_t stands both for capital per labor and capital per firm.

²Note that equilibrium is guaranteed because the technology still displays constant returns-to-scale to private factors and because the whole sequence $\{\bar{K}_t\}_{t=0}^{\infty}$ is taken as given when the consumer solves his/her problem. Only then is the equilibrium condition $\bar{K}_t = n \cdot (K_t/L_t)$ imposed.

by:

$$\begin{aligned} 1 + \gamma_t &= E_t [F_{K/L}(\cdot) + (1 - \delta)] \beta \\ &= E_t \left[\alpha A n^\theta (K_t/L_t)^{(\alpha+\theta-1)} \exp(z_t) + (1 - \delta) \right] \beta. \end{aligned} \quad (2.2)$$

If $\alpha + \theta < 1$ the model displays no growth. If $\alpha + \theta = 1$ the model displays sustainable growth as long as $\alpha A n^\theta E_t [\exp(z_t)] + (1 - \delta)$ is greater than $1/\beta$. If $\alpha + \theta$ exceeds unity there will be explosive growth, which is not a stylized fact of modern economies. Under this conjecture, testing this class of endogenous growth models amounts to checking if $\alpha + \theta = 1$. Under this condition, the growth rate of the economy is stationary, although (log) output itself will be a non-stationary process growing at rate γ .

Exogenous growth models inspired in Solow (1957), Cass (1965) and Koopmans (1965), and updated by King, Plosser and Rebelo (1988) and King, Plosser, Stock, and Watson (1991), will display growth of output per-capita as long as productivity grows exogenously. Otherwise, decreasing marginal returns will ultimately imply no input growth, and thus, under appropriate conditions, per-capita output will reach a steady-state value. Using Cobb-Douglas technology and a random walk productivity process growing at rate μ on average, $\ln(w_t) = \mu + \ln(w_{t-1}) + \epsilon_t^p$, the problem faced by the consumer is the following:

$$\begin{aligned} &\text{Max}_{\{c_t\}} E_0 \sum_{t=0}^{\infty} \beta^t U(C_t), \\ &s.t. \\ &C_t + I_t = Y_t = A K_t^\alpha L_t^{(1-\alpha)} w_t \\ &K_{t+1} = I_t + (1 - \delta) K_t \\ &0 < \beta < 1 \\ &0 < \alpha < 1. \end{aligned} \quad (2.3)$$

After solving the consumer's problem the conditional growth rate of the economy (μ) is given by:

$$\begin{aligned} 1 + \mu_t &= E_t [F_{K/L}(\cdot) + (1 - \delta)] \beta \\ &= E_t [\alpha A (K_t/L_t)^{(\alpha-1)} w_t + (1 - \delta)] \beta. \end{aligned} \quad (2.4)$$

We now discuss some revealing time-series properties of these two classes of growth models.

Proposition 1. *In the problem (2.1), under $\alpha + \theta = 1$, log-utility, and full-depreciation of the capital stock, the closed-form solution for $\ln(K_{t+1})$ implies it has a unit root even with $\{z_t\}$ stationary. Since the closed-form solutions for $\ln(Y_t)$, $\ln(C_t)$ and $\ln(I_t)$ are linear functions of $\ln(K_t)$, they all have common unit roots. Hence, they cointegrate in such a way, that,*

$$\begin{aligned} \ln(Y_t) - \ln(K_t), \\ \ln(I_t) - \ln(Y_t), \text{ and,} \\ \ln(C_t) - \ln(Y_t), \end{aligned} \quad (2.5)$$

form a basis for the cointegrating space. This is true, regardless of the persistency properties of $\{z_t\}$, i.e., even if $\{z_t\}$ is a stationary process; for example, stationary i.i.d. process, stationary ARMA(p, q) process.

Proof. When solving this problem, the consumer takes the sequence $\{\bar{K}_t\}_{t=0}^{\infty}$ parametrically. Thus, the problem reduces to a standard log Cobb-Douglas dynamic optimization one. Under these assumptions, it is a well known result (Sargent, 1987, chapter 1) that there is a closed form solution to this problem, it being straightforward to show that:

$$K_{t+1}/L_{t+1} = \alpha \beta (K_t/L_t)^{\alpha} \bar{K}_t^{\theta} \exp(z_t). \quad (2.6)$$

In equilibrium, $\bar{K}_t = n \cdot K_t/L_t$. Thus, the law of motion of the log of capital per hour can be written as:

$$\ln(K_{t+1}/L_{t+1}) = \mu_1 + (\alpha + \theta) \ln(K_t/L_t) + z_t. \quad (2.7)$$

Where μ_1 includes fixed parameters from preferences and technology. When $\alpha + \theta = 1$, capital per hour has a unit root and (lagged) equation (2.7) is:

$$\ln(K_t/L_t) = \mu_1 + \ln(K_{t-1}/L_{t-1}) + z_{t-1}. \quad (2.8)$$

Using this result in the decision rules and output equation we obtain:

$$\begin{aligned} \ln(Y_t/L_t) &= \mu_2 + \ln(K_t/L_t) + z_t, \\ \ln(C_t/L_t) &= \mu_3 + \ln(K_t/L_t) + z_t, \\ \ln(I_t/L_t) &= \mu_1 + \ln(K_t/L_t) + z_t, \end{aligned} \quad (2.9)$$

Since $\ln(K_t/L_t)$ is $I(1)$, and output per hour, consumption per hour and investment per hour contain $\ln(K_t/L_t)$, they are all $I(1)$ as well. Cointegration as postulated in Proposition 1 above follows immediately from subtracting respectively $\ln(K_t/L_t)$ from $\ln(Y_t/L_t)$, $\ln(C_t/L_t)$ from $\ln(Y_t/L_t)$, and $\ln(I_t/L_t)$ from $\ln(Y_t/L_t)$, using their expressions in (2.9). $\square$

Proposition 1 offers an empirical test for the necessary condition of the increasing returns hypothesis.³ If $\alpha + \theta = 1$, the variables in the model will have unit roots. Moreover, they will cointegrate with known coefficients given by (2.5) above. For example, there is a unity long run elasticity between $\ln(Y_t/L_t)$ and $\ln(K_t/L_t)$, or between $\ln(Y_t)$ and $\ln(K_t)$.

³We discuss below why this is only a necessary condition.

Although unit roots and restrictions (2.5) on cointegration coefficients are a necessary condition for the increasing-returns hypothesis, they are not sufficient for it. Benhabib and Jovanovic (1991) note that when the data have unit roots, it does not necessarily follow that $\alpha + \theta = 1$. As discussed by King, Plosser and Rebelo (1988), and by King et al. (1991), restrictions (2.5) on cointegration coefficients are also present in the standard neoclassical growth model with a random walk (log) productivity process presented above.⁴ We thus present a second result.

Proposition 2. *In problem (2.3), under log-utility, full-depreciation of the capital stock, and a random-walk (log) productivity process, $\ln(w_t) = \mu + \ln(w_{t-1}) + \epsilon_t^p$, the closed-form solutions for $\ln(Y_t)$, $\ln(C_t)$, $\ln(K_t)$ and $\ln(I_t)$ ⁵ are linear functions of the same forcing variable ($\ln(w_t)$). Hence, these variables share a common unit root, and cointegrate in such a way, that,*

$$\begin{aligned} &\ln(Y_t) - \ln(K_t), \\ &\ln(I_t) - \ln(Y_t), \text{ and,} \\ &\ln(C_t) - \ln(Y_t), \end{aligned}$$

form the basis of the cointegrating space.

Proof. See King, Plosser and Rebelo (1988) and King et al. (1991). $\square$

In Proposition 2 the variables of interest still cointegrate with the same coefficients of (2.5) despite the fact that $0 < \alpha < 1$, i.e., despite decreasing marginal returns. Comparing this result to that of

⁴This results generalizes for any $I(1)$ productivity process.

⁵Notice that full depreciation of the capital stock implies $\ln(K_t) = \ln(I_t)$.

the endogenous growth model illustrates a major difference: for the latter, growth, integration, and cointegration follows because agents do not consider their external effect when buying an additional unit of capital. When the external effect has the “right impact on output”, i.e., $\theta = 1 - \alpha$, this extra unit of capital is just enough to make the aggregate capital stock (K_t) grow in equilibrium. Hence, it is the externality of social capital that generates growth. This leads to identical steady-state growth for all variables of interest, i.e., cointegration. For the exogenous growth model, cointegration is obtained forcefully *via* the productivity process, since it is a consequence of the existence of this unique integrated forcing variable ($\ln(w_t)$), which dominates all the long-run co-movement properties of the aggregates in the system.

Despite these conceptual differences, Propositions 1 and 2 show that endogenous and exogenous growth models are observationally equivalent with respect to cointegration restrictions. This was proved by Lau and Sin (1997a), although the initial idea can be traced back at least to Benhabib and Jovanovic (1991).⁶

2.2. Solow residuals of growth models.

Despite the fact that the endogenous and exogenous growth models are observationally equivalent with respect to cointegration restrictions, their Total Factor Productivity (TFP) processes have quite different properties. The first difference is about their order of integration.

Proposition 3. *Given the assumptions in Propositions 1 and 2, if $\ln(Y_t/L_t)$, $\ln(C_t/L_t)$, $\ln(K_t/L_t)$ and $\ln(I_t/L_t)$ are all integrated processes of order one ($I(1)$), then the TFP processes of*

⁶See their footnote 16.

the endogenous growth model is $I(0)$ whereas the TFP process of the exogenous growth model is $I(1)$. More generally, if $\ln(Y_t/L_t)$, $\ln(C_t/L_t)$, $\ln(K_t/L_t)$ and $\ln(I_t/L_t)$ are all integrated processes of order $d(I(d))$, it follows that the TFP process of the endogenous growth model is $I(d-1)$ whereas the TFP process of the exogenous growth model is $I(d)$.

Proof. From equation (2.7), under $\alpha + \theta = 1$, it follows that:

$$z_t = -\mu_1 + \Delta \ln(K_{t+1}/L_{t+1}), \quad (2.10)$$

which establishes that the TFP process of the endogenous growth model is $I(0)$. Notice that the only way to have z_t being $I(1)$ is having $\Delta \ln(K_{t+1}/L_{t+1})$ being $I(1)$ as well. But this makes $\ln(K_{t+1}/L_{t+1})$ and all other variables of interest $I(2)$ processes. For the exogenous growth model, the order of integration of $\ln(Y_t/L_t)$, $\ln(C_t/L_t)$, $\ln(K_t/L_t)$ and $\ln(I_t/L_t)$ must be the same of that of $\ln(w_t)$. Then it follows that $\ln(w_t)$ is $I(1)$. Generalizing the proof for the $I(d)$ case is straightforward. From (2.10) it is obvious that $\ln(K_{t+1}/L_{t+1})$ (and all other variables of interest) is of one order of integration higher than z_t , and thus the result follows. $\square$

Another key difference about TFP processes is the exogeneity assumption on the process:

$$\ln(w_t) = \mu + \ln(w_{t-1}) + \epsilon_t^p.$$

Since the growth rate of productivity is given by:

$$\Delta \ln(w_t) = \mu + \epsilon_t^p,$$

it will grow exogenously as long as there are no external factors influencing $\{\epsilon_t^p\}$. On the class of endogenous growth models little

is imposed on $\{z_t\}$. In particular, the exogeneity assumption used on the other class of growth models is either dropped or omitted. Hence, we refrain from testing its exogeneity status.

Measuring TFP for the endogenous-growth model is straightforward. Labelling p_t^n its Total Factor Productivity, and recalling that in equilibrium $Y_t/L_t = A \exp(z_t) (K_t/L_t)^\alpha (n \cdot K_t/L_t)^\theta$, we have, after a logarithm transformation, and under $\alpha + \theta = 1$:

$$\begin{aligned} p_t^n &= (y_t - l_t) - (k_t - l_t) = y_t - k_t \\ &= a + z_t, \end{aligned} \tag{2.11}$$

where lower-case characters denote logarithms, and $a = \ln(A n^\theta)$. TFP is equal to the Solow residual (z_t) (up to a constant). Thus, TFP, or the (demeaned) Solow residual, can be easily calculated using $y_t - k_t$.

For the class of exogenous-growth models, calculating its TFP requires a consistent estimate of α . Labelling p_t^x as its TFP, we obtain after similar manipulations:

$$\begin{aligned} p_t^x &= (y_t - l_t) - \alpha(k_t - l_t) \\ &= \ln(A) + \ln(w_t). \end{aligned} \tag{2.12}$$

Once again, TFP, or the (demeaned) Solow residual, can be easily calculated using $(y_t - l_t) - \alpha(k_t - l_t)$.

2.3. Infrastructure expenditures, growth, and Solow residuals.

Solow's (1957) famous result – that economic growth cannot be understood solely in terms of the accumulation of physical capital and labor – spurred a vast literature which sought to find economic factors that could explain the variation of output not captured by

input variation. In a typical paper, researchers investigate whether the behavior of “Solow residuals” could be explained by the behavior of a set of economic variables. A vintage of this literature has focused on the role of infrastructure. Aschauer (1989), Nadiri and Manuneas (1992) and Munnell (1990), suggested that the behavior of the Solow residual, and thus productivity, can be partially explained by the evolution of public infrastructure. For example, public infrastructure investment per hour (G_t/L_t) could be a separate argument of the production function. In this case, TFP could be decomposed into two parts, a non-modelled one, and a modelled one - capturing the effect of (G_t/L_t).

An alternative way of modelling the infrastructure effect is to include infrastructure capital as an input of the economy’s production function. This has been tried by Lau and Sin (1997b), but the approach has been criticized by Bougheas and Demetriades (1997) for lacking a fundamental explanation of the infrastructure role in production. Bougheas and Demetriades, on the other hand, propose “introducing infrastructure as a technology which reduces the fixed cost of producing intermediate goods”, i.e., specialization. A by-product of their specification is a non-linear relationship between infrastructure capital and growth.

Regardless of how the role of infrastructure is modelled it is interesting to investigate the following issues for the two classes of growth models considered above: Is TFP of the exogenous growth model (weakly) exogenous when we consider infrastructure expenditures? For the two models considered, do movements in infrastructure precede or are preceded by movements in TFP (Granger-causality)? For the two models considered, what are the long-run elasticities of TFP with respect to Infrastructure? Is there a non-linear relationship between TFP and infrastructure expenditures for the two models considered here (Bougheas and Demetriades)?

3. Previous Empirical Results.

The fact that the long-run capital elasticities of the endogenous and the exogenous growth model are the same creates a problem for econometric tests based on them. Examples are capital elasticity estimates from cross-section regressions using several countries, where cross-section data are a time average with a long time span, and estimates from time-series regressions using integrated data in levels. In these cases, what is being tested is the fit of either endogenous or exogenous growth models, and not the fit of a single one of them. Since only recently the observational equivalence of these models became a well known fact (Lau and Sin (1997a)), there have been several studies in the past that used long-run capital elasticities to confirm or dismiss the fit of the stylized versions of the exogenous or the endogenous growth model presented above; see Neusser (1991), Romer (1987), and Ferreira (1993).

The evidence that the long-run capital elasticity is unity is abundant:

1. Cross-country studies: using data from Maddison (1983) on cross-country rates of growth, Romer's (1987) physical capital estimate is 0.87 and not significantly different from one. With the Summers and Heston (1991) data base, his estimated coefficient is 0.75. Ferreira's (1993) cross-country studies reach a similar conclusion using data from Summers and Heston (1991), the IMF, and Benhabib and Spiegel (1994). Physical capital coefficient estimates are usually close to one and very robust to changes in the specification of the model or to changes of the data used.
2. Time-series evidence: using data in levels and cointegration techniques, Neusser (1991) confirms the theoretical long-run elastic-

ities in either Propositions 1 and 2 above. Benhabib and Jovanovic (1991) running a regression in levels with annual data, find a capital-elasticity estimate of 1.06; see Table 7, p. 91. Using data from Maddison (1983) on three G-7 countries, Lau and Sin (1997a) find evidence supporting cointegration between output and capital, although the capital elasticity is far from unity in some cases.

Despite the favorable evidence above some time-series studies have found a capital elasticity estimate far from unity. In Romer's (1987) article, when annual or decade data in first differences are used, results show a large and significant coefficient for the labor input and an insignificant coefficient for the capital input. The same pattern is found by Benhabib and Jovanovic, who applied maximum likelihood to U.S. data with the error term assumed to be an $ARMA(1, 2)$ process. For quarterly data, they find a labor coefficient of about 0.65 and estimated $\alpha + \theta$ close to 0.23. The evidence against a unity capital elasticity is hard to interpret, simply because what is being estimated is not the long-run capital elasticity. The major problem lies on ignoring the stochastic properties of the data, which are probably non-stationary due to unit roots. If first differences are applied, inconsistent estimates of long-run elasticities may result if long-run relationships are ignored. The same problem occurs if quasi-differencing is applied. In this case, even if there are no long-run relationships in the data, inconsistencies will arise due to imposing common factors.

The relationship between public infrastructure and productivity was discussed and tested originally by Aschauer (1989). Using time series data, Aschauer's estimate of the effect of public capital on TFP is about 0.50, which is relatively high. Munnell (1990) using regional data concurs with Aschauer's estimates. Cost function duality estimates at the industry level by Nadiri and Manuneas (1992) and

Morrison and Schwartz (1992) are relatively smaller (roughly 0.15) but confirm Aschauer's initial premise. The same applies to the cross-country evidence in Ferreira (1993) and Easterly and Rebelo (1993). More recently, Bougheas and Demetriades (1997) estimate a non-linear relationship between infrastructure capital and growth, finding an inverted U-shaped relationship between them.

In most cases, the empirical evidence tends to confirm a supply side role for government (or private infrastructure in some cases) affecting productivity. Exceptions are Hulten and Schwab (1992), using a growth accounting framework, and Holtz-Eakin (1992), using panel data at the state level. Although causality between productivity and public infrastructure is crucial for this discussion, it has not been thoroughly studied.⁷

4. Econometric Tests.

As discussed in the previous section, both endogenous and exogenous growth models are observationally equivalent with respect to cointegration tests. Therefore, the initial step in our empirical investigation is to examine whether U.S. aggregate data conform to these theoretical long-run restrictions. In doing so, we are not testing any of these two models, but a broader class of models that include both. Before cointegration tests are performed, we investigate the order of integration of the data using the Augmented Dickey-Fuller test; see Dickey and Fuller (1979, 1981). There are several cointegration tests in the literature, which is now well known. Gonzalo (1994) compares their properties and concludes that Johansen's (1988, 1991) likelihood-based method is the most adequate overall; see Johansen

⁷ One exception is Ho and Sorensen (1993), where reverse Granger-causality is found from two-digit manufacturing TFP to public infrastructure. Mixed evidence is found on Granger-causality from infrastructure to TFP.

and Juselius (1990) for an application. The test consists of estimating r – the number of linearly independent cointegrating vectors (or cointegrating rank) – and the corresponding cointegrating vectors by maximum likelihood. Two statistics can be used for estimating r : the *Trace* and $\lambda_{\max}$ statistics, which have non-standard asymptotic distributions. Moreover, the technique has the advantage of allowing testing hypotheses on cointegrating vectors, conditioned on knowledge of r . The latter are the usual likelihood-ratio tests and have the nice property that their limiting distribution is χ^2 .

The relationship between growth and infrastructure will be investigated in several ways. Regarding Solow residuals of exogenous growth models we will perform exogeneity tests using the typology of Engle, Hendry and Richard (1983). Suppose that we are interested in conducting inference on a given set of parameters β , and have two series r_t , and s_t . Then, r_t is said to be weakly exogenous for β if we can factor the joint density of $(r_t, s_t)'$, $f(r_t, s_t)$, into the conditional and marginal distributions, $g(s_t|r_t)$ and $h(r_t)$ respectively, where the parameters of the joint and of the marginal densities are separable in such a way that, to learn β , one does not need to conduct inference on the parameters of the marginal density $h(r_t)$. This “separation” property allows doing conditional inference on β using r_t . Strong exogeneity for β requires, in addition to weak exogeneity, that r_t is not *Granger-caused* by s_t . If this is not the case, it is impossible to do conditional forecasting for s_t , since one needs to learn s_t in order to learn the future of r_t .

Weak exogeneity tests will use the results in Johansen (1995, pp. 122-123). There, it is shown that, under cointegration, weak exogeneity can be tested *via* the significance of the error-correction term(s). Granger non-causality tests will take the usual format of testing for exclusion restrictions. Toda and Phillips (1993) show that these test results can lead to wrong inferences when the num-

ber cointegrating vectors in the system is unknown.⁸ We thus present test results under a variety of assumptions regarding cointegration between Solow residuals of the exogenous growth model and infrastructure expenditures.

5. Empirical Results.

5.1. The data set.

The data set used consists of U.S. Post-war quarterly log of private GNP y_t , log of hours worked l_t , log of the capital stock k_t , log of public investment on infrastructure g_t . Private output was obtained subtracting government GNP from total GNP.⁹ Hours worked used the Household Survey monthly series accumulated to generate quarterly figures. Public investments in infrastructure represent the sum of federal, state, and local non-military expenditures on structure and durables. Capital stock data were constructed using the gross private domestic investment series, accumulated using the perpetual inventory method at three different annual depreciation rates: 6%, 8% and 10%. The initial capital stock was calculated following Young(1995). All these series were extracted from Citibase and are available from 1947:1 through 1994:3 in most cases. The exceptions are l_t , for which data are available from 1947:1 through 1993:4, and g_t , available from 1959:1 through 1994:3. We also used an additional version of the capital stock series, based on the annual figures

⁸Wald-exclusion tests should preferably be performed using the vector error-correction model, i.e., conditioning in cointegration after testing. Wrong inference is still possible but less likely in this case.

⁹This is an important difference between our data set and the one used by Lau and Sin (1997a). Using private GNP may be critical in investigating a unity elasticity for the two models considered. With the government sector included, there is the potential to depart from unity capital elasticity, since fiscal variables may be used to do systematic counter-cyclical policies. This is exactly their finding.

provided by the Survey of Current Business. These data were interpolated by the FED to yield quarterly observations in order to run the MPS model, and are available from 1958:1 through 1988:4. Figure 1 displays the production function data. Notice that although all series show signs of non-stationarity, $y_t - l_t$ and $k_t - l_t$ follow each other very closely in the sample period.

5.2. Unit roots and cointegration.

The first step is to test the data for unit roots, since it only makes sense to test for cointegration if they have a long run component. The Augmented Dickey-Fuller test was used including a time trend and a lag structure necessary to obtain white noise residuals.¹⁰ Results are presented in Table 1. At usual significance levels, we cannot reject the hypothesis that y_t , l_t , k_t , g_t , $y_t - l_t$, and $g_t - l_t$ contain one unit root. However, for $y_t - k_t$, we do reject the presence of a unit root, regardless of the version of the capital stock series used. This result is promising, since it is an indication that y_t and k_t , may be proportional in the long run.

Cointegration tests were performed using two distinct data sets in forming the Vector Autoregression (VAR). The first uses the variables in the aggregate production function in per hour terms, i.e., $y_t - l_t$ and $k_t - l_t$, and the second uses all three variables in the production function, i.e., y_t , l_t and k_t .

For the bi-variate VAR, the existence of a long run relationship requires $y_t - l_t$ and $k_t - l_t$, to cointegrate with the cointegrating vector given by $(1, -1)'$. The first step in applying Johansen's cointegration test is to determine the order of the VAR, and decide how to model its deterministic components. The lag-order search was accomplished

¹⁰We started the lag search with 8 lags going down by one lag up to lag zero. The number of lags chosen corresponds to the first significant lag using the t -statistics.

Figure 1
GNP per hour and Capital per hour

Table 1
Augmented Dickey-Fuller test

Variable	Number of Lags	t-ratio
y_t	1	-2.4559
(6% depreciation)	5	-0.31829
(8% depreciation)	5	-0.47550
(10% depreciation)	5	-0.62713
(Survey Series)	6	-1.9083
l_t	5	-2.4153
$y_t - l_t$	6	-0.9976
$k_t - l_t$ (6% depreciation)	5	-1.5688
$k_t - l_t$ (8% depreciation)	5	-1.4288
$k_t - l_t$ (10% depreciation)	5	-1.2816
$k_t - l_t$ (Survey Series)	3	-1.4398
$y_t - k_t$ (6% depreciation)	1	-4.0232**
$y_t - k_t$ (8% depreciation)	1	-4.1077**
$y_t - k_t$ (10% depreciation)	1	-4.1792**
$y_t - k_t$ (Survey Series)	2	-3.8211*
g_t	0	-0.4732
$g_t - l_t$	0	-0.1846
$(y_t - l_t) - 1/3(k_t - l_t)$ (6% depreciation)	2	-2.8352
$(y_t - l_t) - 1/3(k_t - l_t)$ (8% depreciation)	2	-2.8905
$(y_t - l_t) - 1/3(k_t - l_t)$ (10% depreciation)	2	-2.9419
$(y_t - l_t) - 1/3(k_t - l_t)$ (Survey Series)	2	-3.3600

Notes: A constant, seasonal dummies, and a time trend are included in each regression. The lag length was chosen using the first significant t statistic of the highest possible lag, starting at lag 8. If a statistic is significant at 5%, it is labeled with (*), and if it is significant at 1% with (**).

using the Schwarz criterium in conjunction with diagnostics tests.¹¹ The VAR was estimated with a constant, seasonal dummies, and a linear trend term, since the data in levels, and $y_t - k_t$, show signs of

¹¹The first is helpful to decide on the trade-off between degrees of freedom and residual sum of squares, but has little information on whether the model passes specification tests.

containing a linear trend and possibly a quadratic term. The VAR order chosen for the different versions of the constructed $k_t - l_t$, was three, but five lags is also a possibility. For the Survey of Current Business version of $k_t - l_t$, the lag order chosen is three. Cointegration tests were performed using these lag structures. There is overwhelming evidence that there is one cointegration vector. Conditioned on the cointegrating rank being one, we tested the restriction that the VAR does not contain a time trend. For all cases this restriction is strongly rejected, supporting the initial guess that the VAR should be modelled with the trend present.

For three lags, the results of cointegration tests reject that there are zero cointegrating vectors for all versions of the calculated $k_t - l_t$ as shown in Table 2. Moreover, the hypothesis that there is one cointegrating vector cannot be rejected for all cases at usual confidence levels. Estimates of the long run elasticity of capital per hour are all marginally higher than one. To test if they are equal to one, we used the likelihood ratio test proposed in Johansen (1991), conditioned on the cointegrating rank being one. The Results show that a unity coefficient cannot be rejected in all cases with very high confidence. For the Survey of Current Business version of $k_t - l_t$, at three lags, we reject the null of zero cointegrating vectors with 95% confidence using the λ_{max} statistic.¹² Moreover, we cannot reject that the cointegrating rank is one using the trace and the λ_{max} statistic. Capital per hour long run elasticity is estimated to be 0.82. At usual confidence levels, conditioned on one cointegrating vector, we cannot reject that it is unity.

¹² With 90% confidence, both the λ_{max} and the *Trace* statistic reject that the cointegrating rank is zero.

Table 2
Bi-Variate Cointegration Tests

k_t version	Trend χ^2	N° of lags in the VAR	H_0	$\lambda_{\max}$	Trace	Capital elasticity	P-Value
6%	15.7**	3	$r = 0$	17.86*	19.56*	1.042	0.7515
			$r \leq 1$	1.698	1.698		
Depreciation	11.6*	5	$r = 0$	17.34*	18.32*	1.119	0.3435
			$r \leq 1$	0.979	0.979		
8%	15.8**	3	$r = 0$	18.64*	20.69*	1.038	0.7588
			$r \leq 1$	1.452	1.452		
Depreciation	11.1**	5	$r = 0$	16.24	17.35	1.136	0.3137
			$r \leq 1$	1.109	1.109		
10%	15.7**	3	$r = 0$	19.63*	20.86*	1.033	0.7818
			$r \leq 1$	1.231	1.231		
Depreciation	10.6**	5	$r = 0$	18.51*	19.38*	1.101	0.3825
			$r \leq 1$	1.868	1.868		
Survey Series	13.4**	3	$r = 0$	16.95*	17.47	0.821	0.5005
			$r \leq 1$	0.515	0.515		

Notes: Cointegration tests are based on Johansen (1988, 1991) method. Critical values are extracted from Osterwald-Lenum (1992), Table 2, which represents the case where the VAR and the Error Correction term have a constant, seasonal dummies and an unconstrained time trend. These deterministic components were chosen using the Trend χ^2 test above of the null hypothesis that the trend coefficient is zero in the system. The cointegration rank is denoted by r . Conditioned on $r = 1$, the P-Value presented above is associated with the test statistic of the following restriction on the cointegrating space:

$$H_0 \beta = \begin{pmatrix} y & 1 \\ k & -1 \end{pmatrix}.$$

If a test statistic is significant at 5%, it is labeled with (*), and it is significant at 1% with (**).

For five lags, we found very similar results for $k_t - l_t$. One cointegrating vector is found for all but one version of capital per hour (8% depreciation) at 5% significance. However, at 10% significance, there is one cointegrating vector for all versions of the capital stock. Conditioned on rank one, we cannot reject a unity long run elasticity for $k_t - l_t$.

Thus, for the bi-variate VAR, there is overwhelming evidence that output per hour and capital per hour have a long run relationship with unity elasticity, conforming to the two models described above. This evidence is robust to variations of the capital series used or to variations in the lag length of the VAR.

We now turn to the evidence of the tri-variate VAR using y_t , l_t and k_t . Based on statistical tests in Table 3, we found again that all VAR's should include a constant term, seasonal dummies, and a time trend term. For the constructed version of k_t , the preferred lag order is three, but four lags is also a possibility. For the Survey of Current Business version of k_t , we choose a VAR of order two, but order three is also possible.

Using the VAR with 3 lags, for the constructed k_t , we found vast evidence that there is one cointegrating vector. In the long run relationship, point estimates for the capital coefficient are all around 0.9, and those for hours are all near zero. We thus tested the joint hypothesis that the latter is zero and the former is one. For all three versions of k_t we could not reject it, i.e., $y_t - k_t$ is stationary.¹³ These results are robust to changes in the lag order, as can be seen for the results using four lags presented in Table 3.

The analysis for the Survey of Current Business version of k_t yields slightly different results, in which finding a long run unity

¹³We also performed the test restricting the capital coefficient to unity and letting the hours coefficient free. Again, with high confidence, the null cannot be rejected.

Table 3
Tri-Variate Cointegration Tests

k_t version	Trend χ^2	N° of lags in the VAR	H_0	$\lambda_{\max}$	Trace	Capital elasticity	P-Value
6% Depreciation	15.16**	3	$r = 0$	35.84*	43.82**	0.970	0.3668
			$r \leq 1$	7.62	7.98		
			$r \leq 2$	0.357	0.357		
	16.50**	4	$r = 0$	40.01**	47.54**	0.965	0.3726
			$r \leq 1$	7.06	7.53		
			$r \leq 2$	0.470	0.47		
8% Depreciation	15.26**	3	$r = 0$	35.22**	43.34**	0.930	0.3283
			$r \leq 1$	7.68	8.12		
			$r \leq 2$	0.44	0.44		
	15.62**	4	$r = 0$	39.11**	46.85**	0.924	0.297
			$r \leq 1$	7.16	7.74		
			$r \leq 2$	0.60	0.60		
10% Depreciation	15.62**	3	$r = 0$	34.60**	42.98**	0.898	0.251
			$r \leq 1$	7.76	8.28		
			$r \leq 2$	0.52	0.52		
	17.02**	4	$r = 0$	38.34**	46.30**	0.893	0.1943
			$r \leq 1$	7.28	7.95		
			$r \leq 2$	0.68	0.68		
Survey Series	10.02**	3	$r = 0$	24.38	35.19*	0.848	0.5313
			$r \leq 1$	10.28	10.81		
			$r \leq 2$	0.52	0.52		
	13.40**	2	$r = 0$	28.56*	41.73**	0.717	0.052
			$r \leq 1$	12.57	13.18		
			$r \leq 2$	0.61	0.61		

Notes: Cointegration tests are based on Johansen (1988, 1991) method. Critical values are extracted from Osterwald-Lenum (1992), Table 2, which represents the case where the VAR and the Error Correction term have a constant, seasonal dummies and an unconstrained time trend. These deterministic components were chosen using the Trend χ^2 test above of the null hypothesis that the trend coefficient is zero in the system. The cointegration rank is denoted by r . Conditioned on $r = 1$, the P-Value presented above is associated with the test statistic of the following restriction on the cointegrating space:

$$H_0 \beta = \begin{pmatrix} y & 1 \\ k & -1 \\ \ell & 0 \end{pmatrix}.$$

If a test statistic is significant at 5%, it is labeled with (*), and it is significant at 1% with (**).

elasticity depends on the lag length used. For three lags, conditioned on one cointegrating vector, we cannot reject unity elasticity. However, with two lags, we marginally reject it. The results of the Monte-Carlo exercise in Gonzalo (1994) suggest preferring the results with a higher lag order, since omitting the dynamics in cointegrating analysis may lead to inconsistent estimates of cointegrating vectors.

Overall, the outcome of unit-root and cointegration tests confirm the adequacy of the two growth models discussed above.

5.3. Is exogenous-growth TFP exogenous?

For the exogenous growth model we impose $\alpha = \frac{1}{3}$ in constructing its Solow residual. This estimate is by far the most widely used for that purpose; see Cooley and Prescott (1995) and McGrattan (1994). The ADF integration test suggests that the Solow residual is an integrated process of order one for the exogenous growth model under $\alpha = \frac{1}{3}$. Next, we investigate the exogeneity of TFP by setting up a VAR including the exogenous-growth Solow residual and (log) infrastructure expenditure per-hour $g_t - l_t$. Plots of $(y_t - l_t) - \frac{1}{3}(k_t - l_t)$ and $g_t - l_t$ are shown in Figure 2. Exogeneity-test results are presented in Table 4. The lag length chosen is four and all VAR's include a constant, a time trend, and seasonal dummies. Cointegration tests reveal that TFP and infrastructure expenditures have a long-run relationship. At 10% significance and using the *Trace* test statistic, there is one cointegrating vector for all versions of the capital stock used. At 5% there is one cointegrating vector for all capital stock series but the Survey series. Conditioning on the existence of one cointegrating vector, we reject the hypothesis that TFP is weakly exogenous for the parameters of interest in the conditional model using infrastructure expenditures. Note, however, that infrastructure expenditures are not Weakly Exogenous either for the parameters of interest in the conditional model using TFP.

Figure 2
Infrastructure Expenditures per hour and Exogenous-Growth Model Solow Residual

Table 4
Cointegration and Weak Exogeneity Tests for TFP of
the Exogenous-Growth Model and Infrastructure Expenditures

k_t version	Trace-test cointegration rank at 5% (10%) significance	Weak exogeneity test for infra- structure exp. (P-Value)	Weak exogeneity test for TFP (P-Value)
6% Depreciation	1 (1)	0.001	0.007
8% Depreciation	1 (1)	0.001	0.006
10% Depreciation	1 (1)	0.001	0.006
Survey Series	0 (1)	0.016	0.026

Notes: 1) Uses a VAR with four lags, a constant, a time trend, and seasonal dummies; 2) Weak Exogeneity tests conditioning on the existence of one cointegrating vector.

5.4. TFP and infrastructure in endogenous and exogenous growth models.

In this section, given our previous empirical results, it is natural to label $(y_t - k_t)$ as the Solow residual of the endogenous-growth model. As discussed above, for the exogenous-growth model we use $\alpha = \frac{1}{3}$ in constructing the Solow residual.

For the endogenous growth model, Granger non-causality test results between TFP and $g_t - l_t$ are presented in Table 5; see Figure 3 for plots of the data. Since there is an imbalance between the order of integration of the Solow residual (which is $I(0)$) and $g_t - l_t$ (which is $I(1)$), we use the former in levels and the infrastructure expenditures series in first differences. It is well known that causality-test results depend on the lag structure of the VAR. Thus, we present them for lag orders four, three, and two. At the 5% significance level, in all

cases but one, there is no Granger-causality from TFP to $g_t - l_t$.¹⁴ On the other hand, we find that $g_t - l_t$ Granger-causes TFP for almost every lag structure.

Table 5
Granger causality tests for TFP of the
endogenous-growth model and infrastructure expenditures

k_t version	TFP $\rightarrow g_t - l_t$	$g_t - l_t \rightarrow$ TFP	Lags in VAR
6% Depreciation	0.053	0.061	4
	0.472	0.025*	3
	0.413	0.045*	2
8% Depreciation	0.055	0.075	4
	0.560	0.030*	3
	0.552	0.046*	2
10% Depreciation	0.050*	0.091	4
	0.707	0.036*	3
	0.697	0.047*	2
Survey Series	0.257	0.198	4
	0.849	0.081	3
	0.911	0.054	2

Notes: Results presented are P-values testing the null of no Granger-causality. Tests were based on estimated bi-variate VAR's at different lag lengths. Variables in bi-variate VAR are: $(y_t - k_t)$ and $(g_t - l_t)$, a constant, a time trend, and seasonal dummies. Since $(y_t - k_t)$ is $I(0)$ and $(g_t - l_t)$ and are $I(1)$, the former is used in levels and the latter in first differences. If a test statistic is significant at 5%, it is labeled with (*), and if it is significant at 1% with (**).

¹⁴Notice that with a VAR of order 4, Granger-causality p-values are very close to 0.05, a feature that is not observed for VAR's of lower orders. This supports the hypothesis of a seasonal factor showing up in a causality test, since it only happened with the VAR of order 4, and quarterly data are used.

Figure 3
Infrastructure Expenditures per hour and Endogenous-Growth Model Solow Residual

For the exogenous growth model, Granger non-causality test results between TFP and $g_t - l_t$ are presented in Table 6. Since both series are $I(1)$, the distribution of the test statistics depend on whether or not the two series are cointegrated; see Toda and Phillips (1993). Thus, we chose to present test results for the case of no cointegration and of one cointegrating vector, ruling out the case where both series are $I(0)$. If there is one cointegrating vector, test results show that there is feedback for all accumulated capital series at 5% significance, i.e., Granger causality in both directions. For the Survey series there is weak evidence that TFP Granger-causes infrastructure expenditures: no-causality from infrastructure expenditures at 5% but causality at 10%. Thus, we conclude that $g_t - l_t$ Granger-causes TFP and that the reverse is also true. If there is no cointegration test results are very different: there is no Granger causality in either way at 5%. At 10% there is weak evidence of causality from $g_t - l_t$ to TFP.

Putting all these results together, it is clear that we can state unequivocally that there is no evidence that TFP precedes infrastructure expenditure and that it is not preceded by it. Moreover, there is some evidence that $g_t - l_t$ Granger-causes TFP, not being Granger-caused by it. The evidence is strong for the endogenous growth model, but, for the exogenous growth model, these results depend on whether or not the two series are cointegrated. If we take into account cointegration-test results (see table 4, where one cointegration vector is found in most cases), there is weak evidence that $g_t - l_t$ precedes TFP (10% significance). However, as pointed-out by Toda and Phillips, these test results can be misleading.

Table 6
Granger causality tests for TFP of the
exogenous-growth model and infrastructure expenditures

a) assuming one cointegrating vector

k_t version	TFP $\rightarrow g_t - l_t$	$g_t - l_t \rightarrow$ TFP	Lags in VAR levels
6% Depreciation	0.002**	0.018*	4
	0.002**	0.028*	3
	0.007**	0.029*	2
8% Depreciation	0.002**	0.016*	4
	0.002**	0.025*	3
	0.007**	0.026*	2
10% Depreciation	0.002**	0.014*	4
	0.002**	0.022*	3
	0.008**	0.023*	2
Survey Series	0.021*	0.060	4
	0.017*	0.061	3
	0.031*	0.078	2

b) assuming no cointegration

k_t version	TFP $\rightarrow g_t - l_t$	$g_t - l_t \rightarrow$ TFP	Lags in VAR levels
6% Depreciation	0.179	0.085	4
	0.423	0.088	3
	0.958	0.164	2
8% Depreciation	0.186	0.085	4
	0.426	0.089	3
	0.958	0.162	2
10% Depreciation	0.172	0.084	4
	0.419	0.086	3
	0.919	0.142	2
Survey Series	0.303	0.159	4
	0.523	0.158	3
	0.872	0.226	2

Notes: Results presented are P-values testing the null of no Granger-causality. Tests were based on estimated bi-variate VAR's at different lag lengths. Variables in bi-variate VAR are: $(y_t - l_t) - \frac{1}{2}(k_t - l_t)$ and $(g_t - l_t)$, a constant a time trend, and seasonal dummies. Since both variables are $I(1)$ they are used in levels. If a test statistic is significant at 5%, it is labeled with (*), and if it is significant at 1% with (**).

The next step in our empirical investigation is to calculate the infrastructure elasticity of Solow residuals for the two growth models. Results are presented in Table 7. For the endogenous growth model, due to the imbalance in orders of integration between $y_t - k_t$ and $g_t - l_t$, usual statistical inference is invalid, but consistent estimates can still be obtained. Elasticities are calculated from the static (long run) solutions of the VAR in levels using different versions of $y_t - k_t$ and $g_t - l_t$. In all VAR's the lag length used is four. Our estimates are roughly between 0.19 and 0.27. For the exogenous growth model, long run solutions are calculated using $(y_t - l_t) - \frac{1}{3}(k_t - l_t)$ and $g_t - l_t$. Elasticity results are very similar to those of the endogenous growth model for all accumulated series. For the Survey series we obtained an odd negative elasticity. Since when the Survey series is used there is evidence that TFP and infrastructure do not cointegrate (see Table 4) it is hard to attach an economic interpretation to this estimate. With that caveat, regardless of the growth model considered, infrastructure elasticity of TFP remains in the range 0.19-0.27. This is undoubtedly a sizable effect, although much smaller than the initial estimates reported in Aschauer(1989) (about 0.50).

The final step of our empirical investigation of TFP and infrastructure expenditures is to perform diagnostic tests on the estimated linear models (Table 7). We use the RESET test for non-linearity (Ramsey (1969)), including initially a quadratic projection term, and then a quadratic and a cubic term jointly. The results are clear: for the endogenous growth model, there is almost no sign of misspecification of the estimated linear relationship. The exception being the regression run with the Survey series. For the exogenous growth model, using the 10% significance level, all models show signs of misspecification, although at the 5% level they do not.

Table 7
Infrastructure expenditures elasticity and non-linearity tests

k_t version	Endogenous growth model elasticity	Endogenous growth model: Reset test (P-Values)	Exogenous growth model elasticity	Exogenous growth model: Reset test (P-Values)
6% Depreciation	0.266	0.89 (0.61)	0.211	0.08 (0.06)
8% Depreciation	0.228	0.75 (0.46)	0.220	0.09 (0.07)
10% Depreciation	0.201	0.48 (0.34)	0.226	0.09 (0.08)
Survey Series	0.190	0.09 (0.19)	-0.076	0.09 (0.13)

Notes: Estimates are based on long run solutions estimated of bi-variate VAR's in levels, using lag length 4, a constant, and seasonal dummies. For the endogenous-growth model, variables in the bi-variate VAR are $(y_t - k_t)$ and $(g_t - l_t)$. For the exogenous-growth model, variables in the bi-variate VAR are $(y_t - l_t) - \frac{1}{3}(k_t - l_t)$ and $(g_t - l_t)$. The second and last column of the Table present the P-values of the Reset (Ramsey (1969)) test for non-linearity for the endogenous and exogenous growth model respectively. The Reset test includes, for both models, up to the 2nd and up to the 3rd powers of the linear regression projection.

For the endogenous growth model, these results confirm the existence of a supply side role for government infrastructure expenditures. It indicates that an increase in public infrastructure affects productivity positively and thus the marginal returns to capital and labor. On average, if there are better roads, ports, communication systems, etc., the same flow of private capital and labor services is able to produce the same output faster or a higher output in the same amount of time. According to our results, a 10% increase in public infrastructure outlays boosts TFP in the long run by approximately 2%. Since U.S. public expenditures in infrastructure decreased in the 1970's, only reverting this downward trend in the mid 1980's, this can partially explain the observed productivity slowdown for this period. Although with different magnitudes, this result agrees

with Aschauer (1989) and Morrison and Schwartz (1992), who used a completely different methodology in investigating the role of public infrastructure.

For the exogenous growth model the results are more subtle to interpret. On the one hand, it is hard to reconcile the data with an exogenous TFP. Granger-causality results, if anything, corroborate the initial evidence that the Solow residual is not exogenous (strongly exogenous in this case). On the other hand, infrastructure expenditure is not exogenous either, and there is only weak evidence that it Granger-causes TFP. Moreover, there is some evidence that the relationship between TFP and infrastructure investment is non-linear.

6. Conclusions and Further Research.

The basic goals of this paper were: (i) to discuss the time-series properties of growth models in a way that it would be useful for assessing their fit to the data. We focus on the class of endogenous and exogenous growth models (Romer (1986, 1987, and 1989), and Solow (1957), Cass (1965) and Koopmans (1965) respectively), and (ii) to investigate whether a broader class of models, including both of them, successfully conforms to U.S. *post-war* data. We use cointegration techniques to estimate and test long-run capital elasticities, exogeneity tests to investigate the exogeneity status of TFP, and Granger-causality tests to examine temporal precedence. Exogeneity and causality tests are conducted with respect to infrastructure expenditures.

The empirical evidence confirms the existence of a unity long-run capital elasticity, which validates both growth models. It is robust to using several measures of the capital stock and different specifications for the estimated dynamic model. Based on the estimated unity long-run capital elasticity, we construct a new measure of the

Solow residual for the endogenous growth model. Regarding the exogenous growth model, we construct its Solow residual using $\frac{1}{3}$ as the capital share. The TFP of the exogenous growth model is not weakly-exogenous for the parameters of interest when we consider infrastructure investment. Causality tests show unequivocally that there is no evidence that TFP for both models precede infrastructure expenditures not being preceded by it. On the contrary, we find some evidence that infrastructure investment precedes TFP, which is stronger for the endogenous growth model. This may be interpreted as a supply side role for public investments with respect to TFP. Our estimated impact of infrastructure on TFP lay roughly in the interval (0.19, 0.27) and is not very different across the class of growth models considered. Although these estimates show a sizable impact of infrastructure on TFP, they are considerably smaller than the initial estimates in Aschauer (1989) (about 0.50), which were thought to be too large. Finally, if TFP can be regarded as the measure of our ignorance, this last result represents a reduction of it.

Submitted in July of 1997. Revised in September of 1998.

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