Abstract

This paper examines the question of whether foreign direct investment (FDI) enhances labor productivity growth in Mexico. Using cointegration analysis, a dynamic labor productivity function for the 1958-2010 period is estimated that includes, inter alia, the impact of changes in the stock of private and foreign capital per worker. The vector error correction model (VECM) estimates suggest that increases in both private (lagged) and foreign (lagged) investment per worker have a positive and economically significant effect on the rate of labor productivity growth. However, after taking into account the growing remittances of profits and dividends, there is a marked decrease in the economic effect of foreign capital per worker on the rate of labor productivity growth. The paper assesses the short-term interactions of the relevant variables via impulse response functions (IRFs) and variance decompositions (VDCs) based on a decomposition process that does not depend on the ordering of the variables.

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Keywords: vector error correction model, foreign direct investment, Gregory-Hansen cointegration single-break test, impulse response functions, Johansen cointegration test, KPSS (no unit root) stationarity test, labor productivity growth, variance decompositions, Zivot-Andrews single-break unit root analysis.
1. INTRODUCTION.

During the last decade of the 20th century and the first decade of the 21st century, Mexico experienced a sharp and sustained increase in inward foreign direct investment flows (FDI) that can, in part, be traced to the country’s relatively successful implementation of macroeconomic stabilization measures and structural reform programs; it is part and parcel of the general surge in FDI inflows to developing countries associated with the rapid globalization of production and distribution (Chakraborty and Basu, 2002; Rodrik, 2011). FDI inflows to the country jumped from just $2.5 billion in 1989 to $13.2 billion in 2000 and all-time high of $25.7 billion in 2008, before sharply falling to only $6.6 billion in 2010 and an estimated $10.6 billion in 2011 as result of the worldwide recession and financial crisis (ECLAC, 2012). Foreign Investors have been attracted to Mexico as a result of privatization and debt conversion programs, the liberalization of the tradeable sector associated with the passage of NAFTA, and the wholesale removal of restrictive FDI legislation concerning the repatriation of profits, prior authorization of investments, and sectorial restrictions such as local content and export requirements (Agosin, 1995; Blomstrom & Wolff, 1994; Carrada-Bravo, 1998; De Mello, 1997; Ramirez, 2010).

This paper tests whether changes in the accumulated stock of gross foreign direct investment (FDI) and, for the first time in the economic literature on Mexico, the stock of net FDI exert a positive and significant effect on Mexican labor productivity. The focus on Mexico is based both on its strategic economic and geopolitical importance to the United States, and its adoption of comprehensive market based and outward-oriented policies beginning with the administration of Miguel de La Madrid (1982-88) and followed in earnest by the neoliberal administrations of Salinas de Gortari (1988-1994), Ernesto Zedillo (1994-2000), and Vicente Fox (2000-2006).
Mexico is also chosen because the recent surge in FDI flows has been channeled primarily to newly created manufacturing firms in the export-oriented *maquiladora* sector along the U.S.-Mexico border. Although an in-depth examination of the country’s experience with the sharp increase in FDI flows to this sector is beyond the scope of this paper, it is evident that these flows have, and will continue to play, an integral role in the transfer of up-to-date technology, long-term capital, and managerial knowhow to the Mexican nation (Blomstrom and Wolff, 1994).

The layout of the paper is as follows: First, the paper gives an overview of FDI flows to Mexico during the 1990s and 2000s in absolute terms, relative to gross fixed capital formation, and in terms of their sectorial destination. Second, FDI’s potential impact on economic and labor productivity growth is rationalized by incorporating the foreign capital stock as an additional argument in a modified neoclassical production function. The third section presents an empirical counterpart to the simple model developed in Section III and, using cointegration analysis, proceeds for the first time in the extant literature to assess the impact of changes in the stock of FDI--both with and without the inclusion of profit and capital remittances--on Mexican labor productivity growth. The short-term dynamic interactions of the model are addressed via error correction models (ECMs), as well as impulse response functions (IRFs) and variance decompositions (VDCs) of the endogenous variables. The concluding section summarizes the major results and offers some policy prescriptions for attracting FDI into the region and enhancing its positive direct and indirect effects.
2. OVERVIEW OF FDI FLOWS TO MEXICO.

Figure 1 below shows that the relaxation of FDI rules in 1989 and the commencement of NAFTA negotiations in 1991 explains the rapid increase in FDI flows after 1992. The dramatic increase in FDI flows to the country, however, takes place after the passage of NAFTA in 1993. The strength of FDI flows is further revealed by the fact that despite the serious economic downturn in Mexico in 1995, and the associated “Tequila effect” which reduced FDI inflows in 1995 and 1996, they staged a remarkable recovery during the rest of the decade, easily surpassing the pre-crisis levels. The chart shows that, on average, FDI flows more than quintupled, from US$3.3 billion in the 1985-93 period to US$11.3 billion during 1994-2000, and US$16.8 billion during the 2001-2010 period (ECLAC, 2012). The major supplier of FDI flows to Mexico during the decade of the nineties (and historically) has been the United States followed, in order of importance, by Great Britain, Japan, Germany, and France (ECLAC, 2012).
In relative terms, net FDI inflows to Mexico as a percentage of GDP rose from 1.2 percent in 1991 to 3.0 percent in 1999, after registering a NAFTA-induced high of 3.9 percent in 1995, and then averaging 2.82 percent during the 2000-2008 period, before falling sharply to only 1 percent in 2010 as a result of the U.S.’s Great Recession (ECLAC, 2012). From an economic standpoint, the importance of these inflows is more fully appreciated by focusing on their evolution relative to the country’s gross fixed capital formation (ECLAC 2012). Table 1, part A, shows that throughout the decade of the 1990s, and particularly after 1997, FDI flows are averaging 12.9 percent of Mexico’s gross fixed capital formation, but well below those of Chile—the region’s stellar performer.

Critics of FDI, however, contend that these flows, rather than contributing to Mexico’s financing of capital formation, are, in fact, a drain on the country resources because they generate substantial reverse flows in the form of remittances of profits and dividends to the parent companies (Cypher & Dietz, 2004). For example, Figure 1 above shows that despite the fact that since 1985 FDI inflows to Mexico have consistently exceeded profit and dividend remittances, the latter have more than quadrupled between 1990 and 2008 (from $2.3 to $10 billion). Relative to the inflows of FDI during the 1990-2010 period, Mexico’s remittances of profits and dividends averaged almost 46.8 percent (ECLAC, 2012).

One way to gauge the net contribution of FDI to private capital formation is by subtracting from these gross inflows the repatriation of profits and dividends to the parent companies. If we subtract profits and dividends from FDI flows and express the net figure as a proportion of fixed capital formation, it is evident from Table 1 A that the net contribution of FDI inflows to gross
fixed capital formation in Mexico is far less than that advertised by the unadjusted FDI flows. It is also evident that during 2005-2010 period FDI’s contribution has declined relative to the 1994-2004 period. Finally, it is important to note that the net contribution of FDI would be further reduced if we could accurately measure the amount of capital that leaves the region as a result of the widespread practice of intra-firm transfer pricing to avoid taxes and restrictions on the repatriation of profits (see Cypher & Dietz, 2004).

(Table 1 here)

From an economic standpoint, it is preferable to concentrate on the accumulated stock of FDI, rather than the flow variable, because it is the former that ultimately determines the marginal productivity of private capital (and labor). For Latin American as a whole the stock of FDI in constant 1990 dollars rose from $175.6 billion in 1990 to $507.4 billion in 2010 (World Investment Report, 2012). This represents almost a tripling in the stock of FDI of these countries, an increase which is far greater than that of the entire “lost decade” of the 1980s. Focusing on Mexico, its stock of FDI rose from $37.1 billion 1990 and accelerated after 1993 (following the passage of NAFTA) reaching an impressive level of $102.0 billion by year-end 2010 (World Investment Report, 2012). From a relative standpoint, Table 1, Part B shows that as a percentage of GDP the stock of FDI rose from 20.9 percent in 1990 to one third of GDP in 2000, and a high of 48.7 percent in 2009. Even when the stock is adjusted for the remittances of profits and dividends, Table 1 B shows that the net stock of FDI rose from 16.2 percent in 2000 to almost 25 percent in 2010. Endogenous growth theory suggests that if this accumulation of capital in the form of FDI has generated substantial spillover benefits in terms of innovation and managerial knowhow, both of an indirect and direct nature, then the long-term positive
contribution of this surge in FDI during the decade of the 1990s and 2000s cannot be adequately measured by focusing solely on flow variables.

From a sectorial standpoint, net FDI flows into Mexico during the decade of the nineties have been primarily channeled to “greenfield” investments in the manufacturing sector, particularly in branches with a strong participation by TNC’s and with investments oriented towards exports such as those of the maquiladora industry. For example, during the 1990s and 2000s sixty percent of FDI flowed into industrial sectors producing small automobiles (e.g., Ford escorts and VW bugs) and auto parts (engines), industrial machinery and computers, electronic equipment, transportation equipment, food processing, and basic petrochemicals. The banking and financial services industry absorbed twenty five percent, and the remainder was channeled to commerce (12%), mining (1.8%), and agriculture (1%). This evolution in FDI flows is consistent with Dunning’s (1988) locational advantage hypothesis, particularly now that NAFTA has “locked in” many of the neoliberal reforms initiated by the De la Madrid (1982-88), Salinas (1988-94) and Zedillo (1994-2000) administrations (see World Investment Report, 2011).

3. FDI FLOWS AND ECONOMIC GROWTH.

Economic theory suggests that FDI inflows to developing countries increase their stock of capital which, in turn, raises the host country’s labor productivity and incomes, a process that eventually translates into higher levels of output, employment creation, and potential tax revenues. In addition to the direct effects of FDI, De Mello (1997), Huang (2004) and Kehal (2005) observe that indirect positive spillover effects on overall efficiency may arise from the
enhanced competition generated by foreign firms, the transfer of needed technology and managerial knowhow to local firms, and trade-induced learning-by-doing effects as local firms attempt to overcome competition in the global market.

The voluminous empirical literature on the hypothesis of FDI-Led Growth (FLG) by no means speaks with one voice but, in general, finds in cross-country studies that the “nexus between FDI and the host country’s growth seems generally positive” when host countries adopt a more liberalized trading regime (see De Mello, 1997; Huang, 2004; Kehal, 2005; Ram & Zhang, 2002; Zhang, 2001). Recent country studies also suggest that export-oriented FDI, such as that undertaken by TNCs in Chile, Mexico and China, may promote exports (and economic growth) by establishing assembling plants and helping host firms access international markets for exports (see Aitken et al., 1997; De Mello 1997; Zhang, 2001). Other country studies, notably by Chakraborty and Basu (2002) for India and De Mello (1997) for Brazil, have found no support for the FLG hypothesis; in fact, their empirical work suggests that the line of causality runs from GDP growth to FDI—a finding consistent with the market-seeking FDI hypothesis of Dunning (1988) and Mortimore (2003), viz., that FDI is attracted to growing internal markets for services such as telecommunications, gas and electricity, retail commerce, and financial services.

FDI flows may also have a negative effect on the growth prospects 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 or other concessions from the host country. These negative effects would be further compounded if the expected positive spillover effects from the transfer of technology are minimized or eliminated altogether because the technology transferred is inappropriate for the host country’s factor proportions (e.g., too capital intensive); or, when this
is not the case, as a result of overly restrictive intellectual property rights and/or prohibitive royalty payments and leasing fees charged by the TNCs for the use of these “intangibles” (see Ram & Zhang, 2002; and Kehal, 2005).

After Ram and Zhang (2002) and Ramirez (2000) it is possible to analyze the potential impact of inward FDI on private investment and growth in Mexico by appealing to the modified neoclassical production function given in equation (1) below.

\[
Y = AF [L, K_p, K_f, \alpha_i]
\]

where: \( Y \) is the level of real output; \( L \) denotes labor; \( K_p \) is the stock real private capital; \( K_f \) refers to the stock of real FDI capital; \( A \) captures the efficiency of production; and \( \alpha_i \) represents other relevant variables that researchers have incorporated into the modified production function to explain the level of GDP such as real exports of goods and services, the provision of real credit to the private sector, and real government consumption expenditures (see De Mello, Jr., 1997; Feder, 1983; Ram 1986; and Ramirez, 2000).

By treating the stock of FDI as a separate input in the production function, a \textit{ceteris paribus} increase in inward FDI gives rise to three conceptually distinct effects. First, if the stock of FDI capital is productive and complements the private capital stock by augmenting domestic savings and investment, a \textit{ceteris paribus} increase in the stock of FDI capital will increase output directly in the same way that an increase in any other factor of production raises output (\( F_3 > 0 \)).
Secondly, if it generates (unmeasured) positive spillover effects via the transfer of advanced technology and managerial skills to local firms, then it will indirectly increase private investment and output by further raising the marginal productivity of domestic capital \((F_{23} > 0)\) relative to the real interest rate. Third, it will increase output via its positive impact on the marginal productivity of labor; i.e., by increasing the amounts of both domestic private and foreign capital per worker \((F_{12} \text{ and } F_{13} > 0)\).

Of course, as discussed above, an increase in inward FDI that leads to substantial reverse flows of profits and dividends, the transfer of inappropriate technology, and/or eliminates domestic enterprises through intense competition (as a result of TNCs’ monopoly over state-of-the-art technology) could lead to a situation where foreign and private capital are direct substitutes \((F_{23} < 0)\). Under these conditions there would still be a positive direct effect, but a negative indirect effect that could more than offset it; i.e., when the following condition arises:

\[
\left[ (F_3 + F_{13}) + F_{23} - F_{12} \right] < 0.
\]

Finally, in the case where foreign and public capital are independent \((F_{23} = 0)\), perhaps in the case of a foreign copper mining enclave with little forward and backward linkages to the rest of the economy, a *ceteris paribus* increase in foreign direct investment will generate a direct positive effect on output.

4. **EMPIRICAL MODEL.**

Mexico has a sufficiently long (and official) time series data set (extending back to the decade of the sixties) for both private investment and FDI flows so that, using a perpetual inventory method, capital stock data can be generated for both types of capital. In addition, there
is annual data on profit remittances going back to the decade of the fifties, so we can compute a “net” capital stock variable for FDI, viz., a capital variable that takes into account the outflow of profits from the gross inflows of FDI. This permits us to test empirically the hypothesis put forth by critics on the left which contends that FDI inflows do not contribute to capital formation and actually divert resources away from the financing of private investment spending, thereby slowing down economic growth. We are also fortunate to have official data on the economically active population (EAP) for the period under review—as opposed to the common but problematic practice of using the total population variable as a proxy (e.g., see Ram and Zhang, 2002).

For estimation purposes, a modified Cobb-Douglas (CD) production function under the assumption of constant returns to scale is estimated for the 1958-2010 period. Taking total derivatives of the CD production function in log form (including time trend), and adding dummy variables and a random disturbance term, this study extends previous empirical work on Mexico by estimating a labor productivity growth model of the following form:

\[
\Delta y_i = \alpha + \beta_1 \Delta k_{pi,i} + \beta_2 \Delta k_{fg,i} \left[ \Delta k_{fn,i} \right] + \beta_3 \Delta C_{g,i} + \beta_4 \Delta X_{i,i} + \beta_5 D_1 + \beta_6 D_2 + \beta_7 T + \epsilon \tag{2}
\]

where lower case letters denote natural logarithms of the relevant variables in per worker (EAP) terms, and \( \Delta \) is the difference operator; \( y \) denotes the natural log of output per worker (1970 pesos); \( \Delta y_i \) is the labor productivity growth rate and it is defined as, \((\Delta \ln Y - \Delta \ln L)\); \( k_p \) denotes the natural log of the stock of private capital per worker (1970 pesos); \( k_{fg} \) and \( k_{fn} \) denote, respectively, the natural logs of gross and net stocks of FDI capital per worker (1970 pesos); \( \Delta k_{pi} \) is the growth rate in the private capital stock per worker and is defined as \((\Delta \ln K_p - \Delta \ln L)\), and likewise for the other variables in factor-intensive form; \( C_g \) is the log of real government
consumption expenditures (1970 pesos); X refers to the log of real exports (1970 pesos); \( D_1 \) is a dummy variable that equals 1 for the economic and political crises years of 1976, 1982-83, 1987, 1995, and 2009 and 0 otherwise, while \( D_2 \) equals 1 for the petroleum led expansion of 1978-81; \( T \) is a time trend designed to capture any secular trend in labor productivity growth during the period in question. Finally, \( \varepsilon \) is a normally distributed error term.

The economic rationale for the inclusion of the variables in equation (2) and the interpretation of their respective coefficients is given below. Following the lead of Aschauer (1989) and Blomstrom and Wolff (1994), the dependent variable in this study was estimated as a labor productivity growth variable. The coefficients represent the annual percentage change in labor productivity growth associated with a respective percentage change in the growth rate of the variables in question. \( \beta_1 \) is expected to be positive because the more private capital workers have to work with, the more productive they are, *ceteris paribus*. \( \beta_2 \), on the other hand, can be positive or negative depending on whether the accretions of gross (net) foreign capital stock per worker (the flow of FDI per worker) complements or substitutes for private capital formation.

The government consumption variable is added to the model because government expenditures on collective consumption goods such as food, housing, and salaries of public employees are a significant component of aggregate expenditures and may directly or indirectly (via output taxes and subsidies) crowd out private consumption expenditures and thus affect output in a negative fashion (see Aschauer, 1990; Barro, 1990; Barro & Sala-I-Martin, 1995). However, it may also be the case that part of these expenditures go to financing primary and secondary education (as they do in several developed and developing countries, including Mexico). To the extent that they do, they may generate in the long run a positive spillover effect to the private sector in the
form of a better educated workforce that can efficiently seize the market opportunities offered by
the transfer of technology and managerial knowhow associated with FDI, thus affecting labor
productivity in a positive manner. In view of these ambiguous effects, the sign of $\beta_3$ is
indeterminate. $\beta_4$ is expected to have a positive effect because the export promotion hypothesis
suggests that export growth contributes not only directly to labor productivity growth but also
indirectly via increased domestic and foreign investment undertaken by TNCs seeking to exploit
scale economies realized by the exporting firms to a larger market (see Feder, 1983; Sheehy,
1990; Zhang, 2001). $\beta_5$ is anticipated to have a negative sign for obvious reasons. $\beta_6$ is
expected to be positive because of the high rates of economic growth associated with the short-
lived petroleum boom of 1978-81. Finally, $\beta_7$ is expected to have a negative sign because the
annual average growth rate in Mexico has decelerated over the period under review.

5. EMPIRICAL RESULTS

Time series data such as the ones used in this study tend to exhibit either a deterministic
and/or stochastic time trend and are therefore non-stationary; i.e., the variables in question have
means, variances, and covariances that are not time invariant. According to Engle and Granger
(1987), the direct application of OLS or GLS to non-stationary data produces regressions that are
misspecified or spurious in nature. Consequently, this study tested the variables in question for
a unit root (non-stationarity) by using an Augmented Dickey-Fuller test (ADF) (Dickey-Fuller,

Table 2, Part A 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, i.e., the test is run with a
constant term and no time trend.\textsuperscript{11} It can be readily seen that all the variables in level form are nonstationary; i.e., they appear to follow a random walk with (positive) drift (Nelson & Plosser, 1982). In the case of first differences, however, the null hypothesis of non-stationarity is rejected for all variables at least at the 5 percent level. Thus, the evidence presented above suggests that the variables in question follow primarily a stochastic trend; i.e., they are integrated of order one, I(1).

(Table 2 here)

In view of 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 2 (part B) presents the results of running a KPSS stationarity test (Kwiatkowski 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. It is used as a confirmatory test because in the presence of insufficient information, due to a relatively small sample size, it defaults to the stationary data generating process. The reported results in both level and differenced form under the assumption of a constant and time trend are, by and large, consistent with those reported in Table 2 (part A). For example, the null hypothesis of no unit root (stationarity) 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, with the exception of the private capital-labor ratio variable, 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 to a deterministic one, although the possibility that for given subperiods they follow a mixed process cannot be rejected.

A. Single-Break (Zivot-Andrews) Unit Root test.

Although suggestive, the results reported in Table 2 may still be misleading because the power of conventional unit root tests may be significantly reduced when the stationary alternative is true and a structural break is ignored (see Perron, 1989); that is, the investigator may erroneously conclude that there is a unit root in the relevant series. In order 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 break date is selected where the t-statistic from the ADF test of unit root is at its most negative; i.e., a break date will be chosen where the evidence is least favorable for the null hypothesis of a unit root. The test involves running the following 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 (with a constant only) and C (with both a constant and trend), but in a relatively recent study Seton (2003) has shown that the loss in test power (1-β) 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. Table 3 below reports the Zivot-Andrews (Z-A) 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. With the possible exception of the real GDP variable (which is close to the 5% threshold), the unit root null hypothesis with a structural break in both the intercept and the trend cannot be rejected at the 5 percent level of significance. As can be seen from the table, the Z-A test also identifies endogenously the single most significant structural break in every time series.

[Table 3 about here]

B. Cointegration Analysis.

In view of the above, it is necessary to determine whether there is at least one linear combination of these non-stationary variables (in level form) that is I(0). In words, does there exist a stable and non-spurious (cointegrated) relationship among the relevant variables over the period in question? The Johansen and Juselius (1990) method was used first to determine whether there is a stable and unique long-run relationship among the relevant variables in logarithmic form, viz., the natural log of average labor productivity, the log of private capital per worker, and the log of foreign capital per worker. (After all, one of the major advantages of the Johansen method over the Engle-Granger two-step approach is that it is able, in the presence of more than two variables, to determine whether there is more than one cointegrating vector.)

Table 4 (Part A) below shows that the likelihood ratio (L.R.) tests suggest that the null hypothesis of no cointegrating relationship can be rejected at the 5 percent (and 1 percent) levels (trace statistic = 43.961 > critical value (5 percent) = 42.19 ; and Max-Eigen statistic = 31.5 >
critical value (5 percent) = 25.8), thereby suggesting the presence of one unique cointegrating equation from which residuals (EC terms) can be obtained to measure the respective deviations between the current level of labor productivity and the level based on the long-run relationship. The cointegrating vector (normalized on y) is also displayed in Part B of Table 4. The signs are reversed because of the normalization process and they clearly show that, in the long run, $k_p$ and $k_i$ have a positive and highly significant effect on Mexican labor productivity. For example, a ceteris paribus 10 percent increase in the foreign capital stock per worker raises output per worker by an estimated 2.24 percent in the long run. Alarmingly, the negative (and highly significant) trend term in the cointegrating equation suggests that labor productivity growth attributable to technological or residual change in Mexico has been negative over the 1958-2010 period.

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 clearly not the case in the present study. Nevertheless, as a confirmatory test, I performed a Gregory and Hansen (1996) cointegration test with endogenously determined level (intercept) shift (CC) and obtained a minimum ADF* stat. = -5.179 [break point=1972] which is smaller than the tabulated 1% critical threshold level [-5.13 (1%); -4.61(5%); -4.34(10%)] reported by Gregory and Hansen. Thus, the null hypothesis of no cointegration with endogenously determined break is rejected at the 1 percent level of
significance. Thus, even when a structural break is allowed for, the null of no cointegration is rejected. It should be noted that the break date is 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.

(Table 4 here)

C. Error Correction Models and In-Sample Forecasting.

The information provided by the L.R. tests can now be used to generate a set of EC models that capture the short- and long-run behavior of the labor productivity relationship. The changes in the relevant variables represent short-run elasticities, while the coefficient on the EC term represents the speed of adjustment back to the long-run relationship among the variables. Table 4 below displays a partial list of the EC models estimated in this study. Equations (1)-(3) include the gross foreign capital stock (without profit and dividend remittances), while equations (4)-(6) are estimated with the net foreign capital stock (with profit and dividend remittances taken into account). EC_{t-1} represents the lagged residual from the cointegrating equation, and (**) denotes significance at the 5 percent level or less.

The estimates in eqs. (1) - (6) suggest that the lagged impact of changes in the growth rate of the private capital stock per worker are positive and statistically (and economically) significant. The relatively high capital elasticities reported in Table 5 are consistent with the extant empirical literature for developing (and developed) nations, and may be explained by FDI-induced externalities in the form of better managerial know-how and the transfer of superior technology that “inflate” the private capital elasticity estimate by a positive factor θ (see De Mello, 1997). In other words, unless one can measure (the spread of) superior managerial know-how and
control for it, then the growth of private capital will increase output by more than its private marginal product. In fact, Paul Romer (1987) claims in his seminal article that the size of the externality associated with the accumulation of broadly defined capital may be so large as to cause the social marginal product of capital to exceed its private marginal product by a factor of two or even three times (see Benhabib & Jovanovic, 1991, p. 83).\textsuperscript{xiv}

(Table 5 here)

The growth rate of gross foreign private capital stock per worker in eqs. (1)- (3) has a positive and statistically (and economically) significant effect when lagged four periods.\textsuperscript{xv} This result is not altogether surprising because the positive externalities generated in the form of a greater transfer of technology and managerial know-how are likely to impact labor productivity with a considerable lag. The estimates in all equations also suggest that lagged changes in the growth rate of exports have a positive and marginally significant effect on labor productivity growth, thereby providing some support for the export promotion hypothesis. Insofar as the dummy variables are concerned, the estimates indicate that they have the anticipated signs and are highly significant in all EC regressions. The estimates in eqs. (1) and (4) of Table 5 also suggest that they are robust to the exclusion and inclusion of the dummy variables. Finally, the coefficients on both the time trend and consumption variables (not reported) were negative for the reported equations and only marginally significant over the 1963-2010 period.\textsuperscript{xvi}

Turning to the estimates in equation (6), they suggest that the impact of the net FDI capital per worker variable (which takes into account profit and capital remittances) on labor productivity growth is considerably smaller in economic magnitude (by a factor of almost four). This is, perhaps, not altogether surprising in view of the fact that the net capital stock is smaller than the
gross stock, but what is surprising is that despite this reduced scale effect, the foreign capital variable retains its statistical significance. (In the EC model where a 5 percent rate of depreciation is assumed for the stock of FDI capital, the net FDI variable has a positive but marginally significant effect.) The other variables also maintain their economic and statistical significance. The relative fit and efficiency of both EC regressions is quite good and, as the theory predicts, the EC terms are negative and statistically significant, suggesting, as in equation (6), that a deviation from long-run labor productivity growth this period is corrected by about 55 percent in the next year. Stability tests were also undertaken to determine whether the null hypothesis of no structural break could be rejected at the 5 percent level. The Chow breakpoint tests suggested that the hypothesis could not be rejected for the economic crises year 1976 (p-value= 0.4761), 1982 (p-value= 0.1828), 1987, and 1995 (p-value= 0.1736). A Ramsey RESET specification test was undertaken for the reported regressions and, in the case of eq. (6), the null hypothesis of no misspecification bias could not be rejected (p-value= 0.5464).

Figure 2 below, corresponding to eq. (6), provides further visual evidence of the ability of the EC model to track the turning points in the actual series. It also serves to highlight the highly variable and downward trend in Mexican average labor productivity growth over the period in question, particularly after 1980. Figure 3, in turn, reports the Theil inequality coefficient of 0.213 for eq. (6), and it is well below the threshold level of 0.3; its variance, bias, and covariance statistics are also close to their optimal values (see Theil, 1966). Sensitivity analysis on the coefficients suggested that changes in the initial or ending period did not alter the predictive power of the selected EC models.
Figure 2

Figure 3.
Theil Inequality Coefficient, 1963-2010.
D. VECM Analysis and Impulse Response Functions.

In order to assess the dynamic interactions of the variables in the labor productivity function, this study also estimated a four period multivariate error correction model (VECM), treating initially labor productivity, domestic private capital per worker, and foreign capital per worker as endogenous, and the rest of the variables as exogenous (see Sims 1980; Johansen, 1990). In view of the fact that the Cholesky decomposition is arbitrary and sensitive to the ordering of the variables, this study employed a generalized decomposition process first proposed by Pesaran and Shin (1998). Basically it constructs an orthogonal set of innovations that does not depend on the VAR ordering.

Figure 4 below shows the generalized impulse responses of the three variables in question to both a unitary shock in their own values and the rest of the variables over a 10 year period. It can be readily ascertained that the response of Mexican labor productivity to a one standard deviation (SD) innovation in foreign capital per worker is significant and sustained (particularly after three to four periods); the reverse line of “causation” is also strong, particularly after 3 periods. Figure 4 also shows that the lagged response of private capital per worker to a one (SD) innovation in foreign capital per worker is positive and sustained (again, after three periods), while the reverse is not supported (and in fact becomes negative eventually). Finally, the response of labor productivity to private capital per worker is declines after one to two periods and then increases dramatically in magnitude after three periods which is not altogether surprising because it is capturing the reinforced (indirect) positive effect from a unitary innovation in foreign capital per worker. The reported statistics thus seem to support the complementarity hypothesis outlined
in Section III above; i.e., increases in foreign capital per worker appear to have a positive lagged (direct) effect on Mexican labor productivity and (after a three-period lag) begin to crowd in private capital per worker which, in turn, raises labor productivity again. These dynamic interactions (and reinforcement effects) may also explain, in part, the high estimate for the private capital elasticities reported in Table 5.

Next, this study traced the variance decomposition (VDC) of each variable over a ten year period. The VDC gives information about the relative importance of each random (one-
standard deviation) shock to the endogenous variables in the VECM. The results (available upon request) suggest that, after 10 years, a unitary shock in foreign capital per worker explains 16 percent of the accumulated forecast error variance of labor productivity and 17 percent of that of private capital spending per worker. Again, both proportions are significant while the reverse proportions (viz., the accumulated percentage variance of kf due to y and the accumulated percentage variation of kf due to kp) are 37 percent and 6 percent, respectively. 

Finally, a unitary shock in private capital spending per worker explains 24 percent of the accumulated variance in labor productivity, while the reverse proportion is 13 percent.

Finally, 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 variable can be treated as weakly exogenous in the system, thereby omitting it from the interdependent system of variables. Based on this weak exogeneity test, one of the variables, viz., foreign capital per worker, can be omitted from the system (treated as weakly exogenous) because the null hypothesis of a zero restriction is not rejected for this variable at least at the 5 percent level. (See Appendix).

Table 6. Exogeneity Test

<table>
<thead>
<tr>
<th>H0: weakly exogenous variable</th>
<th>Chi-Square statistics (53 obs)</th>
<th>Probability</th>
</tr>
</thead>
<tbody>
<tr>
<td>y, A(1,1)=0</td>
<td>15.393</td>
<td>0.0000</td>
</tr>
<tr>
<td>kp, A(2,1)=0</td>
<td>10.929</td>
<td>0.0009</td>
</tr>
<tr>
<td>kf, A(3,1)=0</td>
<td>2.768</td>
<td>0.0961</td>
</tr>
</tbody>
</table>
6. CONCLUSIONS

This paper has analyzed the impact of FDI flows in a leading country of Latin America: Mexico. Several major findings were presented. First, the descriptive evidence presented in Section II suggests that gross FDI flows to Mexico have been substantial since the second half of the nineties, particularly in relation to domestic capital formation and GDP, and that a large proportion of these funds have been directed to innovative “greenfield” investments in the manufacturing sector, particularly the export-oriented maquiladoras, where the international trade links are strong and the positive spillover effects associated with the transfer of technology and trade-induced learning-by-doing among local firms—as they attempt to overcome competition in the global market—are considered to be high.

Second, Section II also presented evidence which casts doubt on the contribution of FDI to private capital financing and growth. It was revealed that Mexico’s reverse flows of profits and capital have not been inconsequential over the period in question, and that once they are subtracted from the gross inflows of FDI into the country, the contribution of “net” FDI to the financing of private capital formation is reduced accordingly.

Third, the paper presented a simple growth model that explicitly incorporates the positive or negative externalities associated with changes in the stock of FDI per worker. The econometric evidence presented for Mexico suggested that the variables included in the underlying per capita production function have a stable and unique (cointegrating) relationship that keeps them in proportion to one another in the long run, even though each variable in level form has a stochastic trend. The individual EC estimates indicated that the growth rate of the private and (gross) foreign capital stock per worker variables have a positive and statistically significant effect on the growth rate in labor productivity. In this connection, the positive (lagged) impact
of gross FDI per worker on labor productivity growth in the Mexican case is generally in line with the results obtained by other investigators who have analyzed the impact of FDI spillovers on productivity growth (see Blomstrom and Wolff, *op. cit*).

However, contrary to the work of these investigators, this paper presented evidence which suggests that when profit remittances are deducted from gross FDI flows, the economic impact of the resulting “net”FDI capital per worker variable is reduced in magnitude and statistical significance. The reported estimates also provide some support for the export promotion hypothesis. This is not altogether surprising because a significant share of inward FDI flows to Mexico since 1986 (Mexico’s accession to the GATT) has been directed to establishing export assembly operations along the U.S.-Mexico border. Both the government consumption per worker variable and the time trend were found to have a negative and marginally significant effect on Mexican labor productivity growth. The Theil inequality coefficients and Figure 3 suggested that the reported EC models were able to track and simulate the turning points of the historical series quite well. Finally, the reported IRFs and VDCs suggest that increases in foreign capital spending per worker have both a positive and significant effect on labor productivity and private capital spending per worker which, in turn, raises labor productivity again. The estimates also suggest that increases in labor productivity seem to induce higher levels of foreign and private (domestic) capital spending per worker.

From a policy perspective, no strong conclusions can be derived from the results given that they are not generated from a sectoral and/or ownership analysis where issues such as whether foreign-owned capital is more productive than domestically owed capital would have to be addressed. However, the evidence for Mexico suggests that, in the short run, the country may gain by making major institutional-legal changes such as the relaxation or (conditional)
elimination of restrictions on profit and capital remittances, the selective elimination of prior authorization of investments, and the gradual opening of formerly “priority” sectors to foreign investors. However, in the long run, these “market-friendly” measures need to be accompanied by improvements in worker productivity, the upgrading of economic infrastructure, and the introduction of innovative managerial practices so that the receiving country is able to graduate into higher-valued activities. Otherwise, the potential positive spillover effects associated with FDI inflows may be partially (or fully) offset by the costs associated with growing profit remittances, extravagant tax concessions, and the potential compromise of the national patrimony. The econometric evidence in this paper casts some doubt on the “positive” effect of FDI flows on labor productivity growth in the Mexican case, i.e., once the reverse flows of profits and dividends are accounted for in the calculation of the net FDI capital variable. Moreover, the benefits may be further reduced if the major countries of the region engage in this form of “incentive competition” because it may lead to a “beggar-thy-neighbor” effect as TNCs redistribute a fixed volume of FDI among the competing countries, and in so doing, capture for themselves the “lion’s share” of the social benefits.

NOTES
1. Blomstrom and Persson (1983) and, more recently, Blomstrom and Wolff (op. cit) present empirical evidence for Mexican manufacturing industries in the 1970-75 period that suggest that the presence of foreign capital--as measured by the share of employment accounted by foreign firms--has a positive impact on the rate of labor productivity growth of domestic firms within the industry in question (1994, Table 10.5, p. 270). The authors also find (labor) productivity growth convergence between local and foreign firms in industries with a greater presence of multinational corporations (see also Table 10.9, eqs. 4 and 5, p. 275). A study for Uruguay by Kokko, Tansini and Zejan (1996) finds that foreign capital, as measured by the foreign plants’ share of total output, has a positive impact only in those firms that have a relatively small technology gap with foreign firms (see Table 1, p. 608). The results suggest that firm-specific differences in the ability to capture and apply foreign technologies and managerial knowhow are crucial in determining whether FDI flows will have a positive or negative effect.
2. Critics such as Cypher (1997) also emphasize that substantial flows of capital leave LDCs such as Mexico via the subsidiaries’ widespread practice of intra-transfer pricing, viz., their over-invoicing of imports from, and under-invoicing exports to, the parent company.


4. The interested reader is referred to a recent empirical study by Bosworth and Collins (1999) which presents pooled data for developing nations during the 1979-95 period. The estimates suggest that FDI has a positive and statistically significant effect on private investment spending and savings, while portfolio investment has-- in both cases--a statistically insignificant effect. For further details see Table 4, p.162.

5. The share of exports in GDP as well as real exports per worker were also used (separately) as proxies for openness and their inclusion generated results very similar to those obtained from real exports.

6. The rationale for focusing on the rate of labor productivity growth, rather than output growth, resides in the notion that the accumulation of FDI over time generates a spillover efficiency to the workforce via the transfer of better machinery and equipment as well as managerial knowhow. In fact, even if the TNCs do not share the technology with the local firms, they may still have a positive effect on overall labor productivity as local firms are forced to train their labor and management to catch up with their foreign competitors. For further detail see Blomstrom and Wolf (op. cit) who find that labor productivity growth in Mexico is positively associated with the degree of foreign concentration in a given industry (pp. 263-284).

7. Following the neoclassical approach, the coefficients on labor, private capital, and foreign capital are constrained to sum to unity so that the growth rate elasticities of the dynamic production function equal their factor shares, provided, of course, that factor and product markets are perfectly competitive (see Solow, 1956). However, in view of the fact that this may be an untenable assumption for a developing country such as Mexico, I tested the restriction via the Wald test and failed to reject the null hypothesis of constant returns to scale (se note 15 below). For further detail see Coe and Moghadam (1993).

8. It should be noted that the government consumption data contains a portion that is devoted to health and education expenditures, and it should be treated as public (human) capital investment. However, I searched various standard government publications such as INEGI’s Anuario Estadistico de los Estados Unidos Mexicanos; and Nafinsa’s La Economia Mexicana en Cifras, and was not able to find disaggregated government consumption data from which I could generate a human capital variable. I was also not able to obtain data for R&D expenditures in the Mexican case because they are unavailable for the period under review.

9. This study estimated the labor productivity (and output) functions in level form and the results displayed very high R²s and F statistics, as well as low D.W. statistics. These results are typical of spurious regressions.

10. 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. The null hypothesis was also rejected at the 5 percent with a constant and a
time trend. For further details see Charemza and Deadman (1997, pp. 84-92).
11. The order of the lag length was determined by applying both the Akaike Information
Criterion (AIC) and the Schwartz Bayesian Criterion (SBC). Lower values for these performance
statistics indicate a better fit to the data. Between the two criteria, the SBC is more reliable
because the AIC is known to be biased towards choosing an over-specified model.
12. The EC regressions for the output growth equation are qualitatively the same as those for the
labor productivity growth regressions, and they are available upon request. I tested the restriction
that the sum of the growth elasticities for private capital and labor is equal to 1. The Wald test (p-
value= 0.2316) suggested that the assumption of constant returns to scale cannot be rejected. A
similar was obtained when I ran the test with the growth elasticity of the foreign capital stock
included as an additional variable (p-value=.2001), thus suggesting the presence of constant
returns to scale. However, in view of the well-known imperfections present in Mexican labor and
product markets, it is not clear how one should interpret these estimates.

13. At this juncture, it is important to note that although the growth rate in private domestic
capital per worker is lagged in eqs. (1)-(6), it is possible for the line of causation to run the other
way. To test for this possibility I ran a multi-variate VECM Granger Causality/Block Exogeneity
Wald Test with four lags. The results show that the null hypothesis that \( \Delta k \) does not “Granger
cause” \( \Delta y \) can be rejected at the 5 percent level (p-value: 0.011), as well as in the other direction
(p-value: 0.004). Thus, there seems to be causation in both directions in the short run. Of course,
this test says nothing about “causation” per se; it only provides information about whether
changes in one variable precede changes in another. The estimates also suggest that changes in \( k \)
precede changes in \( y \) (p-value=0.01), but not the other way around (p-value= 0.06). Finally, the
results for the pairs \( \Delta k \) and \( \Delta k \) suggest that changes in \( k \) precede those in \( k \) at the 5 percent
level (p=0.010) , as well as the other way around (p=0.012). (At this juncture, it is important to
note that the nature of the lag for the relevant variables was determined via the Akaike-Schwarz
criteria.)
14. The high capital elasticity may also be due to simultaneity bias and omitted variables such as
public capital per worker. For example, it is likely that over time the capital stock is an
endogenous variable that is correlated with the error term (which may be capturing technological
change), thus generating an elasticity estimate which is above capital’s share in output (see De
Mello, 1997).
15. The gross FDI capital per worker variable was also positive and statistically significant when
lagged three and five periods.
16. In most of the estimated EC models, the (lagged) government consumption variable per
worker was found to have a negative but statistically marginal effect on labor productivity
growth. Similar results were obtained when I included a time trend.
17. The sample size for the EC models is between 50 and 53 data points after adjusting for
endpoints which is at or above the threshold level of 50 observations recommended by Granger
and Newbold (1986)–cited in Charemza and Deadman(1997). The IRFs and VDCs were
generated from a four period VECM model for the 1958-2010 period (treating initially labor
productivity, domestic private capital per worker, and foreign capital per worker as endogenous,
and the rest of the variables as exogenous). In general, the results are consistent with those of the
Engle-Granger two-step method. Both the domestic capital stock per worker (when lagged one and three periods) the foreign capital stock per worker (when lagged four periods) were positive and statistically significant, with t-ratios of 2.33 and 1.90, respectively. The labor productivity lag became positive and significant when lagged three to four periods. The error correction term was negative (coefficient: -0.332) and highly significant (t-ratio:-3.873), suggesting that deviations from the long-run labor productivity relationship are corrected in subsequent periods. The export variable was positive (coefficient: 0.12) and significant (t-ratio: 1.90), while the dummy variables had their expected signs and were statistically significant.

18. To conserve space the VDC table is not reported, but is available upon written request.

<table>
<thead>
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</thead>
<tbody>
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<td>16.4</td>
<td>15.7</td>
<td>12.8</td>
<td>12.9</td>
<td>19.1</td>
<td>15.3</td>
<td>11.0</td>
<td>13.5</td>
<td>8.8</td>
<td>9.3</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Source: Computed from Banco de Mexico, [http://www.banxico.org.mx](http://www.banxico.org.mx); ECLAC [1996-97, Table VIII. 4, p. 126]; ECLAC [2002, Table A-5, p. 40; and Tables 12 and 15, pp. 760 and 763]; and ECLAC [2012].


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</thead>
<tbody>
<tr>
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<td>11.2</td>
<td>8.0</td>
<td>7.2</td>
<td>8.0</td>
<td>15.7</td>
<td>12.2</td>
<td>6.0</td>
<td>2.1</td>
<td>4.2</td>
<td>7.1</td>
<td>1.0</td>
<td>1.1</td>
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</table>

Source: Same as above.

Part B


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</thead>
<tbody>
<tr>
<td>Mexico</td>
<td>20.9</td>
<td>20.7</td>
<td>21.8</td>
<td>29.3</td>
<td>30.7</td>
<td>33.0</td>
<td>39.4</td>
<td>42.5</td>
<td>42.9</td>
<td>42.4</td>
<td>43.1</td>
<td>45.2</td>
<td>47.7</td>
<td>47.0</td>
</tr>
</tbody>
</table>

Source: Author’s estimations based on data obtained from Nacional Financiera, S.A. Banco de Mexico, and ECLAC[2002; 2012].


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<thead>
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<td>Mexico</td>
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<td>19.0</td>
<td>21.6</td>
<td>21.6</td>
<td>24.5</td>
<td>24.7</td>
<td>23.8</td>
<td>23.7</td>
<td>24.9</td>
<td>26.4</td>
<td>24.9</td>
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</tbody>
</table>

Source: Same as above.
TABLE 2, Part A. Mexico: Unit Root Tests for Stationarity with constant only, Sample Period 1958-2010.

<table>
<thead>
<tr>
<th>Variables</th>
<th>Levels</th>
<th>First Difference</th>
<th>5% Critical Value(^1)</th>
<th>1% Critical Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>lnY</td>
<td>-1.69</td>
<td>-4.27(^*)</td>
<td>-2.91</td>
<td>-3.59</td>
</tr>
<tr>
<td>lnL</td>
<td>0.71</td>
<td>-3.59(^*)</td>
<td>-2.91</td>
<td>-3.59</td>
</tr>
<tr>
<td>lnK(_r)</td>
<td>0.32</td>
<td>-3.62(^*)</td>
<td>-2.91</td>
<td>-3.59</td>
</tr>
<tr>
<td>lnK(_p)</td>
<td>-1.13</td>
<td>-3.28(^*)</td>
<td>-2.91</td>
<td>-3.59</td>
</tr>
<tr>
<td>lnC(_g)</td>
<td>-2.56</td>
<td>-3.93(^*)</td>
<td>-2.91</td>
<td>-3.59</td>
</tr>
<tr>
<td>lnX</td>
<td>-0.32</td>
<td>-6.22(^*)</td>
<td>-2.91</td>
<td>-3.59</td>
</tr>
<tr>
<td>y</td>
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<td>-7.87(^*)</td>
<td>-2.91</td>
<td>-3.59</td>
</tr>
<tr>
<td>k(_p)</td>
<td>-0.85</td>
<td>-4.17(^*)</td>
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<td>-3.59</td>
</tr>
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<td>k(_r)</td>
<td>-0.02</td>
<td>-4.25(^*)</td>
<td>-2.91</td>
<td>-3.59</td>
</tr>
</tbody>
</table>

\(^1\)MacKinnon critical values for rejection of hypothesis of a unit root. *Denotes significance at the 5 percent level; **denotes significance at the 1 percent level


<table>
<thead>
<tr>
<th>Variables</th>
<th>Levels</th>
<th>First Difference</th>
<th>5% Critical Value(^1)</th>
<th>1% Critical Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>lnY</td>
<td>0.25(^*)</td>
<td>0.03</td>
<td>0.146</td>
<td>0.216</td>
</tr>
<tr>
<td>lnL</td>
<td>0.15(^*)</td>
<td>0.14</td>
<td>0.146</td>
<td>0.216</td>
</tr>
<tr>
<td>lnK(_r)</td>
<td>0.17(^*)</td>
<td>0.13</td>
<td>0.146</td>
<td>0.216</td>
</tr>
<tr>
<td>lnK(_p)</td>
<td>0.22(^*)</td>
<td>0.13</td>
<td>0.146</td>
<td>0.216</td>
</tr>
<tr>
<td>lnC(_g)</td>
<td>0.22(^*)</td>
<td>0.07</td>
<td>0.146</td>
<td>0.216</td>
</tr>
<tr>
<td>lnX</td>
<td>0.84(^*)</td>
<td>0.09</td>
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<td>0.216</td>
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<td>y</td>
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<td>0.146</td>
<td>0.216</td>
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<td>k(_p)</td>
<td>0.21(^*)</td>
<td>0.16(^*)</td>
<td>0.146</td>
<td>0.216</td>
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<tr>
<td>k(_r)</td>
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<td>0.09</td>
<td>0.146</td>
<td>0.216</td>
</tr>
</tbody>
</table>

\(^1\)*Denotes significance at the 5 percent level; **denotes significance at the 1 percent level
### TABLE 3. Part B. Zivot-Andrews One-break Unit Root Test, Sample Period 1958-2010

<table>
<thead>
<tr>
<th>Variables</th>
<th>Levels</th>
<th>Break Year</th>
<th>5% Critical Value¹</th>
</tr>
</thead>
<tbody>
<tr>
<td>y</td>
<td>-4.83</td>
<td>1975</td>
<td>-5.08</td>
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<tr>
<td>kp</td>
<td>-4.32</td>
<td>1985</td>
<td>-5.08</td>
</tr>
<tr>
<td>kf</td>
<td>-4.46</td>
<td>1984</td>
<td>-5.08</td>
</tr>
<tr>
<td>lnY</td>
<td>-5.00</td>
<td>1978</td>
<td>-5.08</td>
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<tr>
<td>lnL</td>
<td>-2.32</td>
<td>1992</td>
<td>-5.08</td>
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<tr>
<td>lnX</td>
<td>-3.68</td>
<td>1976</td>
<td>-5.08</td>
</tr>
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</table>

¹Critical values for rejection of null hypothesis of a unit root with a structural break in both the intercept and trend (Model C). *Denotes significance at the 5 percent level.
<table>
<thead>
<tr>
<th>Eigenvalue</th>
<th>Trace Statistic</th>
<th>5% Critical Value</th>
<th>Probability</th>
<th>No. of CE(s)</th>
</tr>
</thead>
<tbody>
<tr>
<td>0.448</td>
<td>43.961*</td>
<td>42.915</td>
<td>0.039</td>
<td>None</td>
</tr>
<tr>
<td>0.172</td>
<td>12.457</td>
<td>25.872</td>
<td>0.779</td>
<td>At most 1</td>
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<tr>
<td>0.087</td>
<td>3.461</td>
<td>9.16</td>
<td>0.498</td>
<td>At most 2</td>
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</tbody>
</table>

Test assumption: Linear deterministic trend.

<table>
<thead>
<tr>
<th>Eigenvalue</th>
<th>Max-Eigen Stat.</th>
<th>5% Critical Value</th>
<th>Probability</th>
<th>No. of CE(s)</th>
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</thead>
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<td>12.15</td>
<td>0.934</td>
<td>At most 2</td>
</tr>
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</table>

*Denotes significance at the 5 percent level.

Part B: Normalized Cointegrating Vector; coefficients normalized on y.

<table>
<thead>
<tr>
<th>Vector</th>
<th>y</th>
<th>k_p</th>
<th>k_r</th>
<th>constant</th>
<th>Trend</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>1.000</td>
<td>-0.788</td>
<td>-0.224</td>
<td>1.207</td>
<td>0.287</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(-5.96)**</td>
<td>(-2.00)**</td>
<td>(1.29)</td>
<td>(4.57)**</td>
</tr>
</tbody>
</table>

**Note:** T-ratios are in parenthesis. Signs are reversed in the cointegrating vector because of the normalization process.
TABLE 5. Mexico: Error Correction Model; Dependent Variable is: (ΔlnY_t – ΔlnL_t), 1963-2010.

<table>
<thead>
<tr>
<th>Variables</th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
<th>(5)</th>
<th>(6)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>-0.02</td>
<td>-0.02</td>
<td>-0.03</td>
<td>-0.01</td>
<td>-0.01</td>
<td>-0.02</td>
</tr>
<tr>
<td></td>
<td>(-1.51)*</td>
<td>(-1.09)</td>
<td>(-2.70)**</td>
<td>(-0.92)**</td>
<td>(-1.07)</td>
<td>(-1.31)*</td>
</tr>
<tr>
<td>Δk_{t-4}</td>
<td>0.18</td>
<td>0.17</td>
<td>0.19</td>
<td>---</td>
<td>---</td>
<td>---</td>
</tr>
<tr>
<td></td>
<td>(2.30)**</td>
<td>(2.04)**</td>
<td>(2.84)**</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>ΔlnX_{t-1}</td>
<td>---</td>
<td>0.06</td>
<td>0.04</td>
<td>---</td>
<td>0.04</td>
<td>0.05</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(2.22)**</td>
<td>(1.73)*</td>
<td></td>
<td>(1.63)*</td>
<td>(3.11)**</td>
</tr>
<tr>
<td>ECT_{t-1}</td>
<td>-0.95</td>
<td>-0.76</td>
<td>-0.71</td>
<td>-0.97</td>
<td>-0.86</td>
<td>-0.55</td>
</tr>
<tr>
<td></td>
<td>(-5.28)**</td>
<td>(-5.09)**</td>
<td>(-5.00)**</td>
<td>(-6.19)**</td>
<td>(-5.99)**</td>
<td>(-4.12)**</td>
</tr>
<tr>
<td>D_1</td>
<td>---</td>
<td>---</td>
<td>-0.04</td>
<td>---</td>
<td>---</td>
<td>-0.05</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>(-2.91)**</td>
<td></td>
<td></td>
<td>(-4.52)**</td>
</tr>
<tr>
<td>D_2</td>
<td>---</td>
<td>---</td>
<td>0.07</td>
<td>---</td>
<td>---</td>
<td>0.07</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>(5.84)**</td>
<td></td>
<td></td>
<td>(3.42)**</td>
</tr>
</tbody>
</table>

Adj R²           | .65     | .71     | .81     | .61     | .65     | .81     |
S.E.             | 0.04    | 0.03    | 0.02    | 0.04    | 0.03    | 0.02    |
A.C.             | -3.77   | -3.82   | -4.35   | -3.57   | -3.64   | -4.41   |
S.C.             | -3.56   | -3.56   | -3.95   | -3.35   | -3.38   | -3.95   |
D.W.             | 1.94    | 1.91    | 2.03    | 2.00    | 1.98    | 1.99    |
F-Stat.          | 15.23** | 15.44** | 20.03** | 12.81** | 11.79** | 17.57** |

Note: same as in previous table.
REFERENCES


*Anuario Estadistico de los Estados Unidos Mexicanos*. Aguascalientes, Mexico (1998): INEGI


