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Abstract

This paper gives an overview of remittance flows to Mexico during the 1980-2009 period in absolute terms, relative to GDP, in comparison to FDI inflows, and in terms of their regional destination. Next, the paper reviews the growing literature that assesses the impact of remittances on investment spending and economic growth. Third, it develops a simple endogenous growth model that explicitly incorporates the potential impact of remittance flows on economic and labor productivity growth. Fourth it presents an empirical counterpart to the conceptual model and, using single-break unit root and cointegration analysis, proceeds to determine the impact of changes in these flows on economic growth and labor productivity growth over the 1970-2009 period. The error-correction model estimates suggest that remittance flows to Mexico, along with other relevant variables, have a positive and significant effect on both economic growth and labor productivity growth. The concluding section summarizes the major results and discusses potential avenues for future research on this important topic.

J.E.L. Codes: C22, F24, 04, 015, 054

Keywords: Error-correction model, FDI inflows, Johansen Cointegration test, labor productivity growth, remittance flows, Theil inequality coefficient, Zivot-Andrews single-break unit root analysis
I. Introduction.

Over the past decade or so, remittance flows to Latin America and the Caribbean in general, and Mexico in particular, have increased dramatically, even surpassing their FDI inflows for selected years. Figure 1 below shows that remittance flows to Latin America and the Caribbean increased steadily from US$21.3 bn in 2001 to US$53 bn in 2005, before jumping to US$61.5 bn in 2006 and almost US$70 bn in 2008. The figure also reveals that the onset of the Great Recession in 2008-09 led to a significant reduction in these flows, but not as much as anticipated and considerably less than other private and official flows, such as FDI, portfolio investment, and ODA flows. Insofar as Mexico is concerned, it is the largest recipient of remittance flows in Latin America (and the third largest recipient in the world, after India and China) and, not surprisingly, it also recorded a dramatic increase in these flows for the period under review, from a level of US$10.2 bn in 2001 to US$19.3 bn in 2004 and US$26.3 bn in 2008—a figure that surpassed the country’s FDI inflows for that year (see ECLAC, 2010; and World Bank, 2009). In fact, remittance flows have become such an important source of foreign exchange earnings for the country over the last decade that they rank third, just behind Mexico’s earnings from maquiladoras (assembly-line industry) and oil (see Canas et. al., 2007).

Given the magnitude of these flows, both in absolute and relative terms, a growing literature has emerged that attempts to assess empirically the economic determinants of these flows to the region and individual countries, as well as their impact on economic growth, investment, savings, and poverty—to name a few. However, there are relatively few extant studies—and none for the case of Mexico—that try to assess over a sufficiently long time span the economic impact of these flows on a country’s economic and labor productivity growth rates. In this study
we attempt to overcome this lacuna in the extant literature by estimating a modified dynamic production function for Mexico over the 1970-2009 period. The layout of the paper is as follows: First, the paper gives an overview of remittance flows to Mexico during the 1980-2009 period in absolute terms, relative to GDP, and in terms of their regional destination. Second, it reviews the growing literature that attempts to assess empirically the impact of remittances on economic growth for selected developing countries, including several in Latin America and the Caribbean. Third, it develops a simple endogenous growth model that explicitly incorporates the potential impact of remittance flows on economic and labor productivity growth. The fourth
section presents an empirical counterpart to the theoretical model developed in Section III and, using single-break unit root and cointegration analysis, proceeds to determine the impact of changes in these flows on economic growth and labor productivity growth. The concluding section summarizes the major results and discusses potential avenues for future research on this important topic.

II. Overview of Remittance Flows to Mexico.

Although remittance flows to Mexico did not increase dramatically until the decade of the 2000s, they were by no means inconsequential during the decades of the 1980s and 1990s as shown in Table 1 below. Between 1980 and 1989 remittance flows almost tripled from US$1.04 billion to US$2.8 billion, and then more than doubled between 1990 and 1999, from US$3.1 to US$6.7 billion. Notably, for a number of years during the 1980s and early 1990s, remittance flows rivaled or exceeded Mexico’s inflows of FDI. From a relative standpoint, remittances increased as a share of GDP from a mere 0.5 percent of GDP in 1980 to a high of 2.3 percent of GDP in 1988, before falling somewhat during the decade of the 1990s to a stable and respectable annual average of 1.4 percent of GDP. More importantly, perhaps, given remittances’ potential role in financing private capital formation, remittance flows as a proportion of Mexico’s gross domestic capital formation rose from 4.3 percent in 1980 to 11.7 percent in 1988, and then stabilized at an annual average of about 9 percent of GDCF for the decade of the 1990s. The rapid growth in remittance flows to Mexico during the 1990s can be explained, in part, by the 1994-95 peso crisis which dramatically increased migratory flows to the U.S. in search of employment opportunities; the plentiful job opportunities associated with the rapid economic growth remittance flows to Mexico during the 1990s can be explained, in part, by the 1994-95 peso
crisis which dramatically increased migratory flows to the U.S. in search of employment opportunities; the plentiful job opportunities associated with the rapid economic growth experienced by the U.S. economy during the 1995-1999 period; and the credit-driven boom in U.S. construction activity where a disproportionate number of migrants find employment.

Table 1. Remittance Flows to Mexico in Absolute and Relative Terms, 1980-2009.

<table>
<thead>
<tr>
<th>Year</th>
<th>Remit.(US$, bn)</th>
<th>Remit.(% GDP)</th>
<th>Remit. (% of GDCF)</th>
<th>FDI Inflows (US$, bn)</th>
</tr>
</thead>
<tbody>
<tr>
<td>1980</td>
<td>1.04</td>
<td>0.54</td>
<td>4.3</td>
<td>1.62</td>
</tr>
<tr>
<td>1981</td>
<td>1.22</td>
<td>0.52</td>
<td>3.9</td>
<td>1.70</td>
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<tr>
<td>1982</td>
<td>1.23</td>
<td>1.85</td>
<td>16.0</td>
<td>0.63</td>
</tr>
<tr>
<td>1983</td>
<td>1.39</td>
<td>1.25</td>
<td>13.4</td>
<td>0.68</td>
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<tr>
<td>1984</td>
<td>1.56</td>
<td>1.11</td>
<td>10.9</td>
<td>1.43</td>
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<td>1985</td>
<td>1.62</td>
<td>1.53</td>
<td>19.8</td>
<td>1.73</td>
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<td>1986</td>
<td>1.77</td>
<td>2.04</td>
<td>27.1</td>
<td>2.42</td>
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<td>1.44</td>
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<td>1992</td>
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<td>6.8</td>
<td>4.90</td>
</tr>
<tr>
<td>1994</td>
<td>4.12</td>
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<td>1995</td>
<td>4.37</td>
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<td>1996</td>
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<td>1997</td>
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<tr>
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<tr>
<td>2009 a</td>
<td>21.12</td>
<td>3.20</td>
<td>14.0</td>
<td>3.82</td>
</tr>
</tbody>
</table>

Sources: World Bank (2009); Nacional Financiera, S.A. *La Economia Mexicana en
During the decade of the 2000s, up until the year 2007, i.e., before the adverse effects of the Great Recession of 2008-09 began to be felt, Table 1 shows that remittance flows to Mexico increased dramatically, both in absolute terms and relative to GDP and gross domestic capital formation. For example, in 2000 remittance flows amounted to US$7.53 bn and represented 1.3 percent of GDP; by 2007 they had shot up to US$27.1 bn or 3.4 percent of GDP, before falling to US$26.3 bn in 2008. In fact, Mexico was by far the largest beneficiary of remittance flows in all of Latin America and the Caribbean, and among the relatively larger economies of the region it was, with the exception of Peru, the biggest recipient in relation to its gross domestic product over the entire 2000-2009 period. The importance of these flows is further revealed by comparing them with FDI inflows to Mexico for the period under review. As can be seen from Table 1, remittance flows began to rival FDI inflows in 2000 and exceeded them by a significant margin after 2004, particularly in the recession year of 2009.

From a regional standpoint, Canas et al. report that the lion’s share of remittances were sent to the middle income (and poor) central western-states of Michoacan, Guanajuato, Morelos, Zacatecas, and Estado de Mexico (all at least 5% of gross state product). Several of the poorer southern states (Oaxaca, Guerrero and Chiapas) also received significant amounts of remittance flows (at least 5% of GSP). Only the wealthier border-states (Sonora, Chihuahua, Coahuila, and Nuevo Leon) received lower remittance flows (below 1% of GSP) because relatively few low-skilled workers emigrate to the United States from these states. Although the Banco de Mexico supplies information on the regional destination of remittance flows within Mexico, the United
States does not systematically track the origins of these flows within the United States. However, the IDB’s annual survey of remittance flows to Latin America gives us some indication of where these flows originated from because they ranked, not surprisingly, California first ($13.2 billion), Texas second ($5.2 billion), and New York third ($3.7 billion) (see Canas et. al., p. 4).

With the onset and aftermath of the Great Recession of 2008-09, remittance flows to Latin America in general, and Mexico in particular, have experienced an abrupt decline. For example, according to the World Bank (2010), between 2008 and 2009 Latin America and the Caribbean witnessed a decline in remittance flows of almost 10 percent, from $65 billion to $58 billion, while Mexico saw its remittance flows drop from $26.3 billion to $21.1 billion, or a 17.6 percent decline. The World Bank estimates that remittance flows to Mexico will rise slightly in 2010 to about US$22bn. The sharp decline in the case of Mexico can be partly explained by the steep drop in construction employment in the U.S. during 2009, where there is usually a lag of 4 to 6 months between a drop in economic activity (employment) and remittance flows to Mexico (see Mohapatra et al., 2011).

To make matters worse, the economic recovery in the U.S. has been lackluster and the prediction by many economists is for a “jobless recovery” through the end of 2011 and beyond. This means that the employment prospects and income levels of existing and prospective migrants will be adversely affected in years to come, thus undermining their willingness to migrate to the U.S. or, if already here, their ability to send remittances back home. Remittance flows to Mexico are also expected to remain weak in the coming years because tighter border controls have led to a significant reduction in emigration flows from Mexico beginning in the second quarter of 2008 and continuing through the first three quarters of 2009 (INEGI, 2010).
Interestingly enough, the tightening of immigration controls in border-states such as Arizona and California, has had the unintended effect of reducing return migration because of the unwillingness of migrants to try to re-enter the U.S. in view of its increased difficulty. Finally, the increased synchronization of the Mexican (Canadian) and U.S. business cycles since the passage of NAFTA means that when economic activity and employment prospects are poor in the U.S., they are downright dismal in Mexico, thereby acting as a further disincentive to any large-scale return migration (see Ramanarayanan, 2009).

III. The Impact of Remittance Flows on Economic Growth and Development.

There is now a significant and growing literature that attempts to assess empirically the impact of remittances on economic growth for selected developing countries, including several in Latin America and the Caribbean (but not necessarily Mexico). In general, remittances are expected to have a positive effect on the economic growth of the recipient countries when they complement national savings and augment the total pool of financial resources for investment projects. In this connection, Solimano (2003) reports that migrants in the United States, including Ecuadorans, Guatemalans, Mexicans, and Salvadorans, have formed permanent associations known as Home Town Associations (HTAs) which regularly send donations back to their communities to finance investments in small businesses and infrastructure projects such as water treatment plants, roads, bridges, and schools. To the extent that these flows become “institutionalized,” their positive effects on growth are likely to become more permanent.¹

In this connection, Ratha (2003) found that remittances had a positive and significant effect on investment in receiving countries such as Mexico, Egypt and Sub-Saharan Africa. Aitymbetov (2006) found that approximately 10 percent of remittances were used as some form of
investment in Kyrgyzstan, and thus had a positive impact on the economy. Similarly, Giuliano et al. (2006) argue that remittances help boost the growth rate of the economy in less financially developed countries by providing credit which would otherwise not be available. Insofar as Mexico is concerned, investigators have found that remittances are used to finance investment in micro-enterprises. For example, Woodruff (2006) found that there is, in general, a positive relation between investment spending and the growth of micro-enterprises. Woodruff determined that about 5 percent of remittance flows are invested in micro-enterprises and that this could have a significant impact on the long-term growth of these labor-intensive enterprises.

Other authors have found a positive and statistically significant relationship between remittance flows and economic growth. For example, Mundaca (2005), in a study assessing the impact of remittances on growth in selected countries in Central America, found a strong correlation between remittances and economic growth. Remittances had a significant impact on the growth of these economies, and the impact was stronger when the financial sector was included in the model. Mundaca carried out several estimations on the impact of remittances on growth using different variables to proxy for financial development. When domestic credit from banks was used as one of the explanatory variables, a 10 percent increase in remittances as a percentage of GDP increased GDP per capita by 3.49%. However, when no variables were included to proxy for financial development, a 10 percent increase in remittances as a percentage of GDP increased GDP per capita at a lower rate of 3.18% (see also Sharma and Ramirez, 2009).²

However, the positive findings are not readily accepted by other scholars working on this topic. Chami et al. (2005) found a negative correlation between remittances and growth. Basically remittances were found to be counter-cyclical in nature. They argue that remittances
act like compensatory transfers and, hence, do not aid in the process of economic growth. Their idea was that remittances are intended for consumption rather than investment. A similar finding is cited in Solimano (op. cit.) by Bendixen and Associates (2003). They reported that in the case of Ecuador around 60 percent of remittances are spent on food, medicines, house rents, and other basic commodities. Another possible negative effect on growth associated with remittances may result from the possibility of a “Dutch Disease” effect via an induced real appreciation of the domestic currency for countries with sizable remittance flows. To the degree that this happens, traditional and non-traditional exports (and import-competing industries) may be adversely affected, thus undermining investment spending and growth.

Finally, there are a several economic, institutional and social factors which have a potential effect on the size of remittance flows, and thus economic growth. The size of the migrant population, the length of stay away from their home country, the migrants’ income and that of family members back home, volatility of exchange rates, the economic freedom of the source country, the transfer costs, and the migrants’ motivation to go back. For example, Canas et al. (2007) raise the issue of falling money-transfer costs and the new measurement techniques adopted by the Banco de Mexico as the most important factors in determining the increased remittance flows to this country. They have determined that the size of the Mexican migrant population, income, and their attachment to the home country as being less important in determining the size of the flow. So, remittances have the potential to increase irrespective of the home-country situation if the costs of money transfer decrease. In the future, the cost of transferring funds will continue to fall, and thus remittance flows are likely to increase.
IV. Conceptual Model.

For reasons outlined above, remittance flows may have either positive or negative effect on the long-term growth prospects of a country. To the degree that they contribute to the financing of private capital formation, they augment both the stock of private capital and the productivity of labor, thus enhancing the country’s long-term economic growth. On the other hand, if remittances are primarily channeled to finance current consumption, then they reduce current investment spending, thereby reducing the stock of private capital and the country’s long-term growth prospects. Of course, it is possible that remittance flows are used by family members to finance expenditures on education and/or vocational training, and to the extent that they do, then they contribute to the formation of human capital, thus promoting future economic growth. Following the lead of De Mello (1997) and Vita (2004), remittance flows can be treated as a form of foreign capital that generates positive or negative externality effects to the domestic economy. It can be explicitly modeled via an augmented Cobb-Douglas production function of the following form:

\[ Y = A f [L, K_p , E] = A L^\alpha K_p^\beta E^{(1 - \alpha - \beta)} \]

where \( Y \) is real output, \( K_p \) is the private capital stock, \( L \) is labor, and \( E \) refers to the positive or negative externality generated by additions to the stock of foreign capital in the form of remittance flows. \( \alpha \) and \( \beta \) are the shares of domestic labor and private capital respectively, and \( A \) captures the efficiency of production. It is also assumed that \( \alpha \) and \( \beta \) are less than zero, such that there are diminishing returns to the labor and capital inputs.

The externality, \( E \), can be represented by a Cobb-Douglas function of the type:
\[ E = [L, K_p, K_r]^{\gamma} \]  \hspace{0.5cm} (2)

where \( \gamma \) and \( \theta \) are, respectively, the marginal and the intertemporal elasticities of substitution between private and foreign capital in the form of remittance flows. Let \( \gamma > 0 \), such that a larger stock of remittances generates a positive externality to the economy. If \( \theta > 0 \), intertemporal complementarity prevails and, if \( \theta < 0 \), additions to the stock of foreign capital in the form of remittances crowd out private capital over time and diminish the growth potential of the host country.

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

\[ Y = A L^{\alpha + \theta(1 - \alpha - \beta)} K_p^{\beta + \theta(1 - \alpha - \beta)} K_r^{\gamma \theta(1 - \alpha - \beta)} \]  \hspace{0.5cm} (3)

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

\[ g_Y = g_A + [\alpha + \theta(1 - \alpha - \beta)] g_L + [\beta + \theta(1 - \alpha - \beta)] g_{K_p} + [\gamma \theta(1 - \alpha - \beta)] g_{K_r} \]  \hspace{0.5cm} (4)

where \( g_i \) is the growth rate of \( i = Y, A, L, K_p, \) and \( K_r \). Equation (4) states that (provided \( \gamma \) and \( \theta > 0 \)) additions to the stock of foreign capital in the form of remittances will augment the elasticities of output with respect to labor and capital by a factor \( \theta(1-\alpha - \beta) \).

V. Empirical Model.

Mexico has a sufficiently long (and official) time series data set (extending back to the decade of the fifties and sixties) for a number of key variables, including private investment spending,
public capital formation, and FDI inflows, so that using a perpetual inventory method capital stock data can be generated for the different types of capital. Insofar as remittance flows are concerned, there is annual data going back only to the decade of the seventies, so it is not possible to generate a capital stock measure for this variable. Nevertheless, there are still a sufficient number of data points (40) to test empirically whether these flows have a beneficial or adverse impact on economic growth.

Official data on the economically active population (EAP), rather than just population data per se, are also available for the period under review. This study is thus the first, other than Woodruff’s (2006) at the micro level, to test whether remittance flows have a positive or negative effect on economic growth (and labor productivity growth) in Mexico during the 1970-2009 period. The most general formulation of the dynamic production function is given below,

$$\Delta Y = \alpha + \beta_1 \Delta L + \beta_2 \Delta K_p + \beta_3 \Delta K_g + \beta_4 \Delta K_f + \beta_5 \Delta R + \beta_6 D_1 + \beta_7 D_2 + \varepsilon_t \quad (5)$$

$\Delta$ is the difference operator; $Y$ represents the natural log of real GDP (1970 pesos); $L$, as indicated above, refers to the natural log of the EAP; $K_p$ denotes the natural log of the stock of private capital (1970 pesos); $K_g$ denotes the natural log of the stock of public capital; $K_f$ denotes the log of the stock of FDI capital (1970 pesos); $R$ is the natural log of remittance flows (1970 pesos); $D_1$ is a dummy variable that equals 1 for the crises years of 1976, 1982-83, 1987, 1995, 2001, 2008-09 and 0 otherwise, while $D_2$ equals 1 for the petroleum led expansion of 1978-81; Finally, $\varepsilon_t$ is a normally distributed error term.

The economic rationale for the inclusion of the additional variables in equation (5) and the interpretation of their respective coefficients is given below. The coefficients represent the
annual percentage change in real GDP associated with a respective percentage change in the variables in question. Following the lead of Aschauer (1989) and Blomstrom and Wolff (1994) the dependent variable in this study was estimated as a labor productivity growth equation by defining the variables in per capita terms using the economically active population. The sign of $\beta_1$ is anticipated to be positive in both the GDP growth rate formulation and labor productivity growth rate specification. $\beta_2$ is expected to be positive, while the sign of $\beta_3$ can be positive or negative depending on whether government investment spending “crowds in” or “crowds out” private investment spending. To the extent that public investment spending is directed to economic and social infrastructure in the form of roads, bridges ports, and primary education, it is likely to reduce the cost of doing businesses and thereby crowd in private investment spending. On the other hand, if public investment spending is channeled primarily to sectors that directly compete with the private sector and/or indirectly raise the cost of credit by competing for scarce funds, then it is likely to crowd out private investment spending (see Aschauer, 1989; Barro, 1990; and Ramirez, 2000).

Insofar as foreign capital is concerned, FDI inflows are likely to complement private capital formation if they bring needed financing and transfer managerial and technological knowhow. In this connection, the impact of remittance flows will be beneficial to long-term growth if, like FDI inflows, they contribute to financing private capital formation and are directed to investments in human capital and economic infrastructure. In view of these ambiguous effects, the signs of $\beta_4$ and $\beta_5$ are indeterminate. $B_6$ is anticipated to have a negative sign for obvious reasons. $B_7$ is expected to be positive because of the high rates of economic growth associated with the short-lived petroleum boom of 1978-81.
The economic data used in this study were obtained from official government sources such as INEGI (various issues), Nacional Financiera, S.A., *La Economia Mexicana en Cifras*, the Banco de Mexico, *Informe Anual* (various issues), and the World Bank’s *Migration and Development Brief* (various issues). Private and public investment data for Mexico have been obtained from International Finance Corporation, *Trends in Private Investment in Developing Countries: Statistics for 1970-2000* [2002]. The private, public, and foreign capital stock data were generated using a standard perpetual inventory model of the following form,

\[ K_t = K_{t-1} + I_t - \delta K_{t-1} \quad (6) \]

where \( K_{t-1} \) is the stock of capital at time \( t-1 \), \( I_{t-1} \) is the flow of gross investment during period \( t \), and \( \delta \) is the rate at which the capital stock depreciates in period \( t-1 \). In this study the initial stocks of private and foreign capital were estimated by aggregating over six years of gross investment (1955-1960), assuming an estimate of the rate of depreciation of 5 percent. To ensure the robustness of the econometric results, other estimates of the rate of depreciation were used (10 percent), as well as different estimates of the initial capital stock (e.g., summing over 4 and 6 years), but the results were not altered significantly.

VI. Empirical Results

Initially, conventional unit root tests (without a structural break) were undertaken for the variables in question given that it is well-known that macro time series data tend to exhibit a
deterministic and/or stochastic trend that renders them non-stationary; i.e., the variables in question have means, variances, and covariances that are not time invariant. In their seminal paper, Engle and Granger (1987) showed that 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, 1981) with a constant and deterministic trend given below,

$$\Delta X_t = a + (\rho - 1)X_{t-1} + \sum \theta_i \Delta X_{t-i} + bT + \mu_t$$  \hspace{1cm} (7)

where T denotes the time trend and failure to reject the null hypothesis of a unit root ($\rho = 1$) signifies the presence of a non-stationary process.

Table 2 (part A) below presents the results of running ADF test (one lag) on the variables in logarithmic form with a deterministic trend. The results indicate that the null hypothesis of non-stationarity cannot be rejected for any of the variables in level form with a deterministic trend, suggesting that the variables in question do not exhibit a deterministic time trend throughout the period under review. In other words, the common practice of detrending the data by a single trend line will not render the data in question stationary because the trend line itself may be shifting over time (see Charemza and Deadman, 1997). When the ADF test is applied to these variables in first differences under the assumption of a constant and deterministic time trend, all of the variables become stationary at the five percent level of significance.

Table 2 (part B) also presents the results of running an ADF test (one lag) for the variables in both level and differenced form under the assumption of a stochastic trend only, i.e., equation (11) above is run with a constant term and no time trend. 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 (see Nelson and 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 as opposed to a deterministic one, although the possibility that for given subperiods they follow a mixed process cannot be rejected.

A. Single-Break Unit Root Analysis.

Although suggestive, the conventional results reported in Table 2 may be misleading because the power of the ADF test may be significantly reduced when the stationary alternative is true and a structural break is ignored (see Perron, 1989); that is, the investigator may erroneously

<table>
<thead>
<tr>
<th>Variables</th>
<th>Level</th>
<th>First Difference</th>
<th>5% Critical Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Y</td>
<td>-1.74</td>
<td>-6.74*</td>
<td>-3.53</td>
</tr>
<tr>
<td>L</td>
<td>0.41</td>
<td>-6.31*</td>
<td>-3.53</td>
</tr>
<tr>
<td>Kp</td>
<td>-2.46</td>
<td>-4.21*</td>
<td>-3.53</td>
</tr>
<tr>
<td>Kr</td>
<td>-2.41</td>
<td>-3.62*</td>
<td>-3.53</td>
</tr>
<tr>
<td>Kg</td>
<td>-2.18</td>
<td>-6.76*</td>
<td>-3.53</td>
</tr>
<tr>
<td>R</td>
<td>-1.95</td>
<td>-7.60*</td>
<td>-3.53</td>
</tr>
</tbody>
</table>

MacKinnon critical values for rejection of hypothesis of a unit root. * denote significance at the 5 percent level.

<table>
<thead>
<tr>
<th>Variables</th>
<th>Levels</th>
<th>First Difference</th>
<th>5% Critical Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Y</td>
<td>-2.11</td>
<td>-6.28*</td>
<td>-2.93</td>
</tr>
<tr>
<td>L</td>
<td>-2.98</td>
<td>-5.15*</td>
<td>-2.93</td>
</tr>
<tr>
<td>Kp</td>
<td>-1.85</td>
<td>-2.97*</td>
<td>-2.93</td>
</tr>
<tr>
<td>Kf</td>
<td>0.05</td>
<td>-3.60*</td>
<td>-2.93</td>
</tr>
<tr>
<td>Kg</td>
<td>-2.44</td>
<td>-6.79*</td>
<td>-2.93</td>
</tr>
<tr>
<td>R</td>
<td>-1.22</td>
<td>-7.63*</td>
<td>-2.93</td>
</tr>
</tbody>
</table>

Same as above.

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). The models are given below:

\[ \Delta X_t = a + (\rho - 1)X_{t-1} + \sum \theta_i \Delta X_{t-i} + bT + cD_t + \mu_t \]  

(Model A)  

(8)
\[ \Delta X_t = a + (\rho - 1)X_{t-1} + \sum \theta_i \Delta X_{t-i} + bT + dDT_t + \mu_t \]  
(Model B) \hspace{1cm} (9)

\[ \Delta X_t = a + (\rho - 1)X_{t-1} + \sum \theta_i \Delta X_{t-i} + bT + cD_t + dDT_t + \mu_t \]  
(Model C) \hspace{1cm} (10)

Where \( D_t \) is a dummy variable to capture an intercept (mean) shift occurring at each possible break date (TB), and \( DT_t \) is a trend shift variable. Formally, \( D_t \) equals 1 if \( t > TB \), and 0 otherwise; and \( DT_t \) equals \( t - TB \) if \( t > TB \), and 0 otherwise.

Following the lead of Perron, most investigators report estimates for either models A and C, but in a relatively recent study Seton (2003) has shown that the loss in test power \((1 - \beta)\) is considerable when the correct model is C and researchers erroneously assume that the break-point occurs according to model A. On the other hand, the loss of power is minimal if the break date is correctly characterized by model A but investigators erroneously use model C. In view of this, Table 3 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 notable exception of the remittances variable, 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.\(^7\) The Z-A test identifies endogenously the single most significant structural break in every time series. The exception for the remittances variable is, in all likelihood, due to the abrupt break in 1970 due to the lack of reported data before that year. In view of space constraints, Figure 2 below shows visually the endogenously determined break-date for the real GDP series.\(^8\)

<table>
<thead>
<tr>
<th>Variables</th>
<th>Levels</th>
<th>Break Year</th>
<th>5% Critical Value$^1$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Y</td>
<td>-4.83</td>
<td>1978</td>
<td>-5.08</td>
</tr>
<tr>
<td>L</td>
<td>-2.29</td>
<td>1992</td>
<td>-5.08</td>
</tr>
<tr>
<td>$K_p$</td>
<td>-4.32</td>
<td>1985</td>
<td>-5.08</td>
</tr>
<tr>
<td>$K_f$</td>
<td>-4.46</td>
<td>1984</td>
<td>-5.08</td>
</tr>
<tr>
<td>$K_x$</td>
<td>-4.80</td>
<td>1978</td>
<td>-5.08</td>
</tr>
<tr>
<td>R</td>
<td>-6.97*</td>
<td>1970</td>
<td>-5.08</td>
</tr>
</tbody>
</table>

$^1$Critical 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.
Figure 2

Zivot-Andrew Breakpoints
B. Cointegration Analysis.

In order to determine whether there exists a stable and non-spurious (cointegrated) relationship among the regressors, this study employed the cointegration method first proposed by Johansen (1988). The Johansen method was chosen over the one originally proposed by Engle and Granger (1987) because it is capable of determining the number of cointegrating vectors for any given number of non-stationary series (of the same order), its application is appropriate in the presence of more than two variables, and more important, Johansen (1988) has shown that the likelihood ratio tests used in this procedure (unlike the DF and ADF tests) have well-defined limiting distributions. The maximum likelihood test involves estimating the following relationship:

$$\Delta X_t = \Sigma \Gamma_i \Delta X_{t-i} + \Pi X_{t-k} + \nu + \varepsilon_t, \ t=1, \ldots, T \quad (11)$$

where,

$$\Gamma = -I + \Pi_1 + \ldots + \Pi_i, \ I \ \text{(is a unit matrix)}, \ \forall i = 1, 2, \ldots t-k. \quad (12)$$

and

$$\Pi = -(I - \Pi_1 - \ldots - \Pi_k) \quad (13)$$

Where the $\Gamma_i$'s are estimable parameters that contain information on the short-run adjustments of changes in the vector $X$ ($\Delta X_t$), and the coefficient matrix $\Pi$ conveys information on a possible long-run equilibrium (cointegrated) relationship among the variables in question. The rank of $\Pi$ is equal to the dimension of the cointegrating space, and when it is equal to the number of variables, $p$, it has full rank; i.e., all included variables are stationary and a test for cointegration
becomes redundant. If, on the other hand, the rank of $\Pi$ is equal to the null matrix, then estimation of equation (11) above corresponds to estimating the production function in differenced form (i.e., without a stationary process). Finally, if the rank $\Pi$ is greater than 0 and equal to $r$, where $r < p$, then there exists $(p \times r)$ matrices, $Q$ and $R$, such that $\Pi = QR'$. $R$ is called the cointegrating matrix and it consists of $r$ cointegrating vectors $R = (R_1, \ldots, R_r)$, where $0 < r < p$. The linear combinations given by the vector $R'X$ are trend stationary even though $X$ itself is non-stationary, while $Q$ is a matrix of vector adjustment (error correction) coefficients.

In this last case, equation (12) can be estimated as an error correction (EC) model, viz., a model that combines both the short-run properties of economic relationships in first difference form and the long-run information provided by the data in level form. EC models thus enable the researcher to estimate the speed of adjustment back to the long-run (stable) condition among the variables. In this connection, Engle and Granger (1987) warn that failure to include the lagged residual of the cointegrating equation in a (short-run) model in difference form results in a misspecified relationship because the long-run properties of the model are ignored—properties which can only be extracted from the data in level form.

To save space, Table 4 below reports the Johansen maximum L.R. test for cointegration for only the output equation. The first column of the table gives the eigenvalues in descending order, while the second column reports the corresponding trace statistics generated from the maximum L.R. test statistic. The next two columns report, respectively, the 5 percent critical and p-values. Finally, the last column gives the null hypotheses, $p - r$, ranging from no cointegrating relationships up to at most four cointegrating vectors (given that $p = 5$). It can be ascertained from the L.R. ratio statistics that, in the presence of a constant in the cointegrating
and VAR equation, there exists a linear combination of the I(1) variables that links them in a
stable and long-run relationship. In fact, the p-value reported in the table shows that the null

**TABLE 4. Johansen Cointegration (Trace) Test, 1970-2009.**

A. Series: Y, L, Kp, Kf, and R.

Test assumptions: Intercept (no trend) in CE and VAR; Kg, D1, and D2 are treated as
exogenous variables.

<table>
<thead>
<tr>
<th>Eigenvalue</th>
<th>Likelihood Ratio</th>
<th>5%</th>
<th>P-value</th>
<th>No. of CE(s)</th>
</tr>
</thead>
<tbody>
<tr>
<td>0.957</td>
<td>170.027</td>
<td>69.82</td>
<td>0.000</td>
<td>None</td>
</tr>
<tr>
<td>0.452</td>
<td>46.866</td>
<td>47.856</td>
<td>0.062</td>
<td>At most 1</td>
</tr>
<tr>
<td>0.307</td>
<td>23.392</td>
<td>29.797</td>
<td>0.227</td>
<td>At most 2</td>
</tr>
<tr>
<td>0.204</td>
<td>9.093</td>
<td>15.495</td>
<td>0.357</td>
<td>At most 3</td>
</tr>
<tr>
<td>0.005</td>
<td>0.203</td>
<td>3.842</td>
<td>0.653</td>
<td>At most 4</td>
</tr>
</tbody>
</table>

B. Normalized Cointegrating Vector; coefficients normalized on Y.

<table>
<thead>
<tr>
<th>Vector</th>
<th>Y</th>
<th>L</th>
<th>Kp</th>
<th>Kf</th>
<th>R</th>
<th>Constant</th>
</tr>
</thead>
<tbody>
<tr>
<td>1.</td>
<td>1.000</td>
<td>-0.108</td>
<td>-0.129</td>
<td>0.29</td>
<td>-0.137</td>
<td>-3.154</td>
</tr>
<tr>
<td></td>
<td>(9.87)</td>
<td>(3.22)</td>
<td>(11.6)</td>
<td>(21.1)</td>
<td>(-7.03)</td>
<td></td>
</tr>
</tbody>
</table>

Note: t-ratios are in parenthesis. Signs are reversed in the cointegrating vector because of the
normalization process.

hypothesis of no cointegrating vector can be rejected at least at the one percent level, thereby
suggesting the presence of one cointegrating equation from which residuals (EC terms) can be
obtained to measure the respective deviations between the current level of output and the level
based on the long-run relationship. The long-run estimates reported in Part B of the Table
suggest that all variables, with the notable exception of the foreign capital variable, have positive and highly significant effect on the level of real GDP. Similar results were obtained for the labor productivity function and they are available upon request.

C. Error Correction Models.

The information provided by the L.R. tests is used to generate a set of EC models that capture the short- and long-run behavior of the labor productivity relationship. EC models enable the researcher to estimate the speed of adjustment back to the long-run (stable) condition among the variables. Engle and Granger (1987) warn that failure to include the lagged residual of the cointegrating equation in a (short-run) model in difference form results in a misspecified relationship because the long-run properties of the model are ignored. 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. For simplicity, consider the EC model with lags (and no dummy variables) given in equation (14) below:

\[
\Delta Y_t = \alpha + \beta_1 \Delta L_{t-i} + \beta_2 \Delta K_{pt-i} + \beta_3 \Delta K_{gt-i} + \beta_4 \Delta K_{ft-i} + \beta_5 \Delta R_{t-i} + \delta EC_{t-1} + \varepsilon_t \quad (14)
\]

The coefficients (\(\beta\)'s) of the changes in the relevant variables represent short-run elasticities, while the coefficient, \(\delta (< 0)\), on the lagged EC term obtained from the cointegrating equation in level form denotes the speed of adjustment back to the long-run relationship among the variables. To conserve space, the results for two of the EC models estimated in this study are given below, one for the growth rate in real GDP and the other with the growth rate in labor productivity (lower case letters denote per capita terms), respectively.

\[
\Delta(Y)_t = -0.03 + 0.4 \Delta L_t + 0.85 \Delta K_{pt-1} - 0.17 \Delta K_{gt-2} + 0.06 \Delta K_{ft-3} + 0.03 \Delta R - 0.56 EC_{t-1} - 0.04 D_1 + 0.09 D_2 \quad (15)
\]
\[
\begin{align*}
\Delta(y) & = -0.03 + 0.92\Delta(k_p)_{t-1} - 0.21\Delta(k_p)_{t-3} + 0.07\Delta(k_i)_{t-3} + 0.04\Delta(r)_{t-1} - 0.03\text{EC}_{t-1} - 0.03D_1 + 0.08D_2 \\
\Delta(y) & = -0.03 + 0.92\Delta(k_p)_{t-1} - 0.21\Delta(k_p)_{t-3} + 0.07\Delta(k_i)_{t-3} + 0.04\Delta(r)_{t-1} - 0.03\text{EC}_{t-1} - 0.03D_1 + 0.08D_2
\end{align*}
\]
EC estimates for the foreign capital variable are 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 economic growth (and labor productivity) with a lag. The public investment variable has a lagged negative and significant effect on real GDP growth and labor productivity growth which is somewhat unexpected. Perhaps it could be due to the fact that the variable does not measure economic infrastructure spending *per se*, but overall public investment spending which includes spending by state-owned enterprises and other public entities. During the 1970s and early 1980s most of the increase in public investment spending was associated with the growth of state-owned enterprises in sectors that were in direct competition with the private sector (see Ramirez, 1989). Insofar as the impact of remittance flows are concerned, the estimates for both equations suggest that they have a positive and highly significant effect when lagged one period. The reported estimates for the labor productivity growth in eq. (16) are consistent with those reported in equation (15). Turning to the dummy variables, the estimates suggest that they have the anticipated signs and are highly significant in both EC regressions. 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 (15), that a deviation from long-run labor productivity growth this period is corrected by about 56 percent in the next year. Finally, stability tests were 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 = 0.453), 1982 (p-value = 0.325), and 1995 (p-value = .773).
Before concluding, the EC models were used to track the historical data on the percentage growth rate labor productivity during the period under review. Figure 3 below, corresponding to equation (16) above, shows that the model was able to track the turning points in the actual series quite well. PGR refers to the actual series and (PGRF) denotes the in-sample forecast. In addition, Figure 3 shows that the Theil inequality coefficient for this model is 0.215, which is well below the threshold value of 0.3, and suggests that the predictive power of the model is quite good [see Theil, 1966]. The Theil coefficients can be decomposed into three major components: the bias, variance, and covariance terms. Ideally, the bias and variance components should equal zero, while the covariance proportion should equal one. The estimates reported in
Figure 4 suggest that all of these ratios are close to their optimum values (bias= 0.00, variance= 0.028, and covariance = 0.972). Sensitivity analysis on the coefficients also revealed that changes in the initial or ending period did not alter the predictive power of the selected models (results are available upon request).

VII. Conclusions.

Several major findings were presented in this paper. First, the evidence for Mexico suggests that remittance flows have been substantial since the second half of the eighties and during the decade of the nineties and beyond, particularly in relation to GDP, gross fixed domestic capital formation, and from a regional standpoint. It was also noted that these flows have, since the decade of the eighties, but particularly during the decade of the 2000s, rivaled or
exceeded FDI inflows. The extant literature suggests that to the extent that these flows are
cchanneled to finance private capital formation and economic infrastructure (broadly defined), it
portends well for the future growth prospects of the country. Moreover, remittance flows tend to
exhibit a greater degree of stability and less susceptibility to the business cycle than FDI inflows
which are often referred to as the “good cholesterol” of global (private) financial flows.

Second, unit root tests in the presence of one-time structural breaks in both the intercept and
trend indicate that the null hypothesis of non-stationarity cannot be rejected for the relevant
series in level form, with the notable exception of the remittances variable. This is probably due
to the lack of official (reported) data on remittance flows for Mexico before 1970 (the stipulated
break point). The rest of the estimates are consistent with those generated by conventional unit
root tests. This is an important finding because conventional unit root tests tend to exhibit low
power when structural breaks are ignored and the stationary alternative is true; i.e., investigators
are more likely to incorrectly conclude that the variable in question has a unit root. Third, the
Johansen cointegration test indicated that there is a stable relationship among the relevant
variables in level form which keeps them in proportion to one another over the long run. This is a
highly important contribution to the extant literature because previous econometric studies
relating to impact remittances in the Mexican case have failed to determine whether the
estimated relationships were spurious or not.

Fourth, the EC models reported above suggests that short-run deviations from the long-run
relationship for both real GDP and labor productivity are corrected in subsequent periods and,
equally as important, Figure 1 shows that the in-sample forecasts of the EC models are able to
track the turning points in the data relatively well. The short-run estimates suggest that
Remittance flows have a positive and lagged effect on both the growth rate in real GDP and labor productivity growth over the period in question. The other included quantitative variables also have their anticipated signs and they are statistically and economically significant. Finally, the qualitative variables also have their expected signs and they are statistically significant.

From a research standpoint, it would be highly important for future investigators to determine whether the massive inflows of remittances the country has received in recent decades have been directed to finance investment spending in physical and human capital or primarily to defray current consumption. As more disaggregated data becomes available on a regional or sectorial basis, it will be possible to conduct panel studies to determine whether remittance flows have greater positive (or negative) effect on investment spending and labor productivity growth in certain regions or sectors of the Mexican economy. This should help policymakers and the home town associations alluded to earlier channel remittances to where they can have their maximum effect in terms of financing investment, promoting economic growth, and alleviating poverty. Other avenues for future research will need to focus on what are some of the major economic and institutional determinants of remittance flows to countries such as Mexico. There is already a growing literature on this topic but relatively few time-series studies that identify the key economic and institutional variables in a rigorous fashion for the major recipients of remittance flows, viz., China, India and Mexico.
REFERENCES


ENDNOTES.

1 Ellerman (2003) reports that Mexican migrant associations send home between US $5000-$25,000 per year, while migrant associations from El Salvador send home donations of about US $10,000 per year. See also Solimano (2003).

2. Using panel cointegration analysis and the Fully-Modified OLS (FMOLS) methodology, Sharma and Ramirez (2009) report estimates for selected upper and lower income Latin American & Caribbean countries which suggest that remittances have a positive and significant effect on economic growth in both groups of countries. In addition, the interaction of remittances with a financial development variable reveal that these two variables act as substitutes and, similar to Mundaca’s findings, that the impact of remittances is more pronounced in the presence of the financial development variable.

3. 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).

4. Investment and FDI data published by the OECD was cross-checked with that found in INEGI and La Economia Mexicana en Cifras, and no significant differences were discerned.

5. 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.

6. A stochastic trend is one where the random component of the series itself, say variable $x_t$, 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 former process is referred to as a random walk with drift, either positive or negative, and it is captured by including a constant in the DF or ADF tests. It is also possible for the variable in question to display both a stochastic and deterministic trend process over time. For further details see Charemza and Deadman, (1997, pp. 84-92).

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

8. The analysis undertaken in this study only tests for the presence of a single endogenously determined structural break. In a recent paper, Lee and Stazicich (2003) show that when there are, in fact, two structural breaks in the data, assuming erroneously that there is only one can result in a loss of power of the test.
9. Estimates for the labor productivity function are consistent with those for the output function and are available upon request.

10. The public investment variable was assumed to be exogenous because Mexican policymakers abruptly reduced public investment spending in response to the demise of import-substitution industrialization policies as well as external pressure on the part of the IMF and other multilateral institutions (for further details, see Ramirez, 1997).