

The Mexican Peso and the Korean Won Real Exchange Rates: Evidence from Productivity Models.

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Abstract: Using the U.S. as benchmark, Korean data starting from 1970:1 to 2000:4 and Mexican data from 1983:1 to 2000:4 are decomposed into traded (manufacturing) and non-traded sectors (mostly services). All three countries experienced growing gaps of productivity in favor of traded goods production over time. Likewise, the Won and the Peso both depreciated over time with respect to the U.S. dollar (USD). We find that the traditional purchasing power parity (PPP) model performs remarkably well for the Mexican Peso, while the standard productivity model is strongly rejected for both countries. However, a version of the model shows that as Mexican traded goods productivity rises, the Peso nominal exchange rate appreciates (coefficients between -2.03 and -2.16), while as U.S. traded goods productivity rises, the Peso depreciates (coefficients between 2.06 and 2.48), and the effect of price differentials on the exchange rate is either 1.08 or 1.09 . The results are less supportive for Korea, most likely reflecting the failure of a real model to take into account financial considerations (randomness in spot rates) that drive the won real exchange rate.

Keywords: cointegration, non-traded goods, traded goods, PPP, productivity models.

JEL Classification Numbers: F11, F31.

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1. Introduction

The econometric approach to productivity differentials and real exchange rates is based on the following conjecture: if a long-run equilibrium exists, productivity innovations explain permanent changes in real exchange rates. A representative study of this line of research is Strauss (1996) who finds that: i) increases in the domestic productivity of traded goods (relative to non-traded) appreciates the real exchange rate in all six industrial economies studied, and ii) increases in the foreign productivity of traded goods depreciates the real exchange rate in four of the economies. The theoretical idea is of course much dated, going back at least to works of the sixties [Balassa (1964) and Samuelson (1964)]. Hsieh (1982), in one of the first empirical implementations of the productivity-based model, estimates by ordinary least squares and instrumental variables the basic equation under German and Japanese annual data from the period 1954-1976 against their major trade partners.

The results in Hsieh (1982) are now viewed with skepticism, especially because time series methods have been extended in many ways during the last 20 years. The results in Strauss (1996) lack power as well, due to its approach under annual data covering the period 1960-1990. Moreover, as emphasized by Froot and Rogoff (1995), the field of real exchange rates is especially vulnerable to survivorship bias. This is so because, due to data limitations, evidence is almost exclusively established for industrial economies. Studies in this field that apply original econometric methods under data from industrial economies include: Mark (1990), Cheung and Lai (1993), Chen (1995), and Costa and Crato (2001). More narrowly, on the Balassa-Samuelson effect, Canzoneri et al. (1999) and Faria and León-Ledesma (2003) apply techniques, such as: panel cointegration and the bounds test approach, respectively, to industrial economies.

Less developed countries (LDCs) or newly industrialized countries (NICs) are less frequently explored. It is true that empirical studies of purchasing power parity (PPP) are now often found under the various time series methods, including cointegration, for such countries. Examples include: McKnown and Wallace (1989) on four high inflation economies, Mollick (1999) and Alves et al. (2001) on Brazil in the very long run, Salehizadeh and Taylor (1999) and Holmes (2002) for several emerging economies, and Choudhry (2000) for Eastern European countries. A related set of studies on LDC countries is based on the econometrics of trend breaks, of which Mahdavi and Zhou (1997) and Zhou (1997) are typical examples.

None of these studies, however, has focused on the productivity-model of the real exchange rate as put forward by Hsieh (1982). The most likely reason for the scarcity of research is data availability. Despite advances in data gathering and technology, LDCs still suffer from severe data limitations. This study attempts to remedy this problem, employing a carefully trimmed dataset for Korea and Mexico, using the U.S. as benchmark. Using country-specific data sources in the three countries, we define the *traded goods sector* as manufacturing and the *nontraded goods sector* as the sum of seven broad services areas: electricity, gas and water; construction; wholesale and retail sale, restaurants and hotels; transport, storage and communication; financial services and insurance; community, social and personal services; and government services. This division is backed by OECD guidelines and has been used before for industrial economies by Strauss (1996).

The empirical exploration of the productivity model for the Korean and Mexican economies is straightforward. Suppose that both countries experience higher productivity growth in the traded sector than in the non-traded sector. This is an assumption entirely consistent with the data as will be shown below. The price of traded goods in these countries

remains constant due to the law of one price on traded goods. Higher productivity in this sector will lead to higher wages paid to its workers, which will imply higher wages in the non-traded good sector due to labor mobility. However, because the productivity gains in the non-traded good sector are smaller than those in the traded good sector, firms in the non-traded good sector will have to raise prices to absorb the wage increase. This will bring about a rise in the overall price level domestically. If the foreign (say, U.S.) price level is unchanged, the rise in domestic prices will translate into a real exchange rate appreciation. Conversely, slower productivity growth in the traded good sector leads to a real exchange rate depreciation.

In order to reconsider the sector productivity-based model for Korea and Mexico, we borrow from the methodologies of Strauss (1996), Canzoneri et al. (1999) and Chinn and Alquist (2002). This research route has the advantage of presenting, first, the *traditional PPP model* with the spot rate adjusting one by one with home and foreign price levels. It then moves to the *standard productivity model*, in which the real exchange rate is driven by productivity differentials in each country. Rejections of the former are widespread in the literature. Strauss (1996), in particular, attribute them for the six economies studied to “measurement errors” or innovations in productivity differentials. The standard productivity model, on the other hand, has found mixed support across quarterly data studies [Canzoneri et al. (1999), Chinn and Alquist (2002)] and broad support across annual data studies in Strauss (1996) and in the non-parametric approach of Tille et al. (2001). Combining the definition of the real exchange rate and the standard productivity model leads to what may be called the *augmented productivity sector model*. Evidence in Strauss (1996) based on such model is focused on likelihood tests of linear combinations of series, rather than estimation of the vector coefficients themselves.

This paper aims to remedy the lack of cointegration evidence, on a quarterly data basis, of the productivity sector model for economies such as Korea and Mexico. We find that the traditional PPP model performs remarkably well for the Mexican peso. Not only cointegration is found under both Johansen and Stock and Watson methodologies, but also the coefficient of the nominal exchange rate with respect to price differentials in Mexico varies in a narrow range from 0.94 to 1.04. The standard productivity model, however, is strongly rejected for both countries, while the performance of the augmented productivity model is extremely good for the Mexican Peso against the USD: as Mexican traded goods productivity rises, the nominal exchange rate appreciates, while as U.S. traded goods productivity rises, the nominal exchange rate depreciates, and the effect of price differentials on the exchange rate is close to 1.

In Korea the relative traded goods productivity gap grows overall but suffers a slowdown in the mid to late 80s, which contrasts to the monotonically increasing traded goods productivity gap in Mexico. These contrasting patterns between domestic productivities in Korea and Mexico may lie behind the failure of the model in Korea. Another possibility is that the strong co-movement between nominal and real exchange rates suggests a better model is one more grounded on financial variables [e.g., Chinn and Meese (1995)], rather than on real variables. Finally, relying on Canzoneri et al. (1999), we replicate the model for Japan as the benchmark country against Korea and find unstable and poor estimates. This, together with the fact that the Korean economy is now very much related to the U.S., rules out the benchmark country hypothesis for the failure of the productivity model.

The remainder of the paper is organized as follows. Section 2 presents the derivation of the models to be estimated and section 3 discusses the data. Section 4 summarizes the major findings and section 5 reviews the results and suggests extensions for further work.

2. The Models

According to the purchasing power parity (PPP) doctrine, the nominal exchange rate (S , expressed as the domestic currency price of one unit of foreign exchange) adjusts to movements in domestic (P) and foreign (P^*) price levels. The real exchange rate (Q) is defined by the nominal exchange rate corrected to movements in domestic and foreign price levels, which results, in logarithmic form:

$$s = k + p - p^* \quad (1)$$

$$q \equiv s + p^* - p \quad (2),$$

where k is a constant equal to zero when absolute PPP holds and different from zero when some form of relative PPP holds. A large set of studies surveyed by Froot and Rogoff (1995) explores that temporary deviations from PPP imply q follows a covariance stationary process. If there is no long run relationship, there is no trend for prices to return to levels dictated by PPP. If PPP does not hold, q has a unit root and deviations from PPP are permanent.

Due to impossibility of arbitrage-based transactions, PPP is imposed on the traded goods (T) sector:

$$p_T = p^*_T + s \quad (3)$$

The general price level, however, is composed of traded and non-traded (NT) goods as follows, with respective shares given by α and α^* :

$$p = (1 - \alpha)p_T + \alpha p_{NT} \quad (4a)$$

$$p^* = (1 - \alpha^*)p^*_T + \alpha^* p^*_{NT} \quad (4b)$$

Substitution of these equations into (2) yields the equivalent expressions:

$$q = -\alpha(p_{NT} - p_T) + \alpha^*(p^*_{NT} - p^*_T) \quad (5a)$$

$$q = -\alpha \ln(P_{NT}/P_T) + \alpha^* \ln(P^*_{NT}/P^*_T) \quad (5b)$$

The pair of equations given by (5) suggests that q depends on the relative price of nontradables in the two economies. A rise in the relative domestic price of nontradables implies an appreciation of the real exchange rate, while a rise in the foreign relative price of nontradables causes a depreciation of the real exchange rate. The differential productivity model proposed by Hsieh (1982), based on the Balassa-Samuelson paradigm, is a 2-country model with one factor (labor), in which its supply is fixed at the aggregate. Labor, however, may move within the two sectors, which leads to wage equalization across sectors.

Suppose now that the marginal product per worker (MPL) in the 2 sectors is given by A_T and A_{NT} . Price competition leads to:

$$P_T = W/A_T, P_{NT} = W/A_{NT}, P^*_T = W^*/A^*_T, \text{ and } P^*_{NT} = W^*/A^*_{NT} \quad (6)$$

Due to measurement difficulties associated with MPL, we replace it by the average product per worker as in Hsieh (1982). Among the justifications provided in Canzoneri et al. (1999) for this shortcut, the most important is perhaps that this precludes data on (sector)

capital stock, requiring information on (sector) employment or value added. Replacing (6) into (5a) leads to the **standard productivity model** and to the **augmented productivity model**, respectively:

$$q = -\alpha (a_T - a_{NT}) + \alpha^* (a^*_T - a^*_{NT}) \quad (7)$$

$$s = (p - p^*) - \alpha (a_T - a_{NT}) + \alpha^* (a^*_T - a^*_{NT}) \quad (8)$$

When $(\tilde{a}_T - a_{NT})$ grows, the relative price of nontradables $(p_{NT} - p_T)$ grows. By (5a), q falls, ensuing a real exchange rate appreciation. If the data generation process is non-stationary, cointegration analysis searches a long-run equilibrium relationship within the vectors below:

$$Z1 = [s, p - p^*] \quad (9a)$$

$$Z2 = [q, a_T - a_{NT}, a^*_T - a^*_{NT}] \quad (9b)$$

$$Z3 = [s, p - p^*, a_T - a_{NT}, a^*_T - a^*_{NT}] \quad (9c)$$

The vector $Z1$ is the **traditional PPP model** with the spot rate adjusting one by one with price levels at home and abroad. The vector $Z2$ contains the **standard productivity model**, in which the real exchange rate is ultimately driven by productivity differentials in each country, and vector $Z3$ is another representation of the theory, in what we call the **augmented productivity model**, since now both PPP and relative productivities are considered. Positive evidence on $Z2$ and $Z3$ suggests the Balassa-Samuelson framework is operative. If, in particular, a linear combination of (9c) is stationary, there is a cointegration relationship among the spot rate, relative prices, and productivity differentials in the two countries.

3. The Data

The data consists of annual and quarterly series of South Korea and the United States for the period 1970-2000 and Mexico for the period of 1983-2000. The data set is created with several series taken from country-specific data sources as follows. For Mexico, from the Instituto Nacional de Estadística Geografía e Informática (INEGI: <http://www.inegi.gob.mx>). For South Korea, from the National Statistic Office of Korea (NSOK: <http://www.nso.go.kr>). For the U.S., from the Bureau of Economic Analysis (BEA: <http://www.bea.doc.gov>), and from the Bureau of Labor and Statistics (BLS: <http://stats.bls.gov>).

In order to calculate traded and nontraded average labor productivities, we classify sector product using a methodology similar to Strauss (1996), who used the Organization for Economic Cooperation and Development (OECD) classification. The traded goods sector includes all of manufacturing and the nontraded goods sector includes all of the following services: electricity, gas and water; construction; wholesale and retail sale, restaurants and hotels; transport, storage and communication; financial services and insurance; community, social and personal services; and government services.

Figures 1 and 2 contain the most important series for Korea and Mexico, respectively. The traded and nontraded average labor productivities (A_{NT} and A_T of section 2) are calculated by dividing each of the total production by sector in constant prices by the total (full and part time) employment of the respective sector. The sector productivities in a given country appear in the figures as $lnprodkr$, $lnprodmx$, and $lnprodus$. It can be seen a modest slowdown in Korean relative sector productivity during the 80s and that the 90s are years of steady growth in productivity. This pattern contrasts to Mexican relative traded goods sector productivity

growth, which is monotonically increasing. In the U.S., relative productivity across sectors displays a fairly well behaved growing trend.

The nominal Won and Mexican Peso exchange rates with respect to the U.S. dollar (USD) are called, in logarithms, Innerkr and Innermx in figures 1 and 2. The logarithm of the real exchange rate with USD as the foreign currency is simply called q . The consumer price index (P), and the price indexes of each sector in Mexico and in South Korea (P_{NT} and P_{T}) are obtained from the same databases of each country mentioned above. The U.S. CPI is obtained from the BLS. The difference, in logarithms, between each country's price level and the U.S.'s is called Incps . For all three countries, CPI and price indexes of nontradable and tradable sectors are on a 1995 basis. Employment is in thousands of people for all three countries. For the three countries, GDP-nontradable and GDP-tradable are in (billions of won, thousands of pesos, or millions of USD) 1995 constant prices.

The quarterly data for Mexico and Korea are originally in non-seasonally adjusted (nsa) form, while the U.S. series are seasonally adjusted (sa). To use them together, we convert the nsa series to sa, employing the estimation method X-12 ARIMA with holiday and trading day adjustment, the same used by the U.S Bureau of the Census.

For the construction of sector-based productivities, the U.S. GDP by industry does not exist on a quarterly basis as defined above. It is possible, however, to see that U.S. National Income by industry is appreciably different than GDP by industry in levels, but the growth rates are very close. Therefore, in order to create quarterly U.S. GDP by industry (real and nominal), an interpolation routine is adopted converting annual data to quarterly data with the growth rates of the seasonally adjusted quarterly National Income (without capital consumption adjustment). This is the method recommended by BEA.

4. Results

4.1. Preliminary Results

Consider first the unit root evidence in table 1 for the period 1970:01 to 2000:04 for Korea and for the period 1983:01 to 2000:04 for Mexico. Three types of unit root tests are used. The first is the usual Augmented Dickey-Fuller (ADF) test. We employ the sequential approach suggested by Ng and Perron (1995) for choosing the number of lags in the estimating equations that appear in parenthesis close to the reported ADF figure. We also employ the Elliott, Rothenberg and Stock (ERS) (1996) generalized least squared (DF-GLS) test, which is also computed under the null that the series is non-stationary. Finally, we report the results of the Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) (1992) test, in which the null hypothesis becomes the stationary series.¹

The ADF and DF-GLS tests in the upper part of table 1 show that, except for the productivity differentials between sectors in the U.S. (Inprod) for the DF-GLS, the unit roots null is not rejected in levels at the 5% level. Similarly, the KPSS rejects the null of stationarity in all cases, except for the Inpnt in Mexico when it does so only at 10%. Thus, according to DF-GLS tests only, productivity differentials in the U.S. appear to be stationary in levels. In the lower part of table 1 the ADF and DF-GLS tests reject the null in all cases, except for inflation differentials between Mexico and the U.S. and the DF-GLS test for Inprod in Korea.

¹ For the KPSS test, table 1 contains values calculated from lag-length truncation parameter equal to 4. The Bartlett kernel is used for spectral estimation method. As a robustness check, we choose different choices of lag-length selection in the KPSS test using lag-lengths (L) of 0, 4 and 13 as computed from the formula $\text{lag} = 0$, or $\text{lag} = \text{integer} [L(T/100)^{.25}]$. These three choices match precisely the original choices of KPSS for lag-length, in which Monte Carlo simulations show that size distortions appear as L increases [KPSS, 1992, p.170]. The results using $\text{lag} = 0$ and the ones reported in table 1 are broadly consistent. Using $\text{lag} = 4$, the results are very similar to $\text{lag} = 0$. If we did not correct for error autocorrelation (and had taken $\text{lag} = 0$), we would be assuming that iid errors are plausible under the null (of stationarity around a deterministic trend), which is not justified easily. Results, including the I(2) for inflation differentials between Mexico and the U.S., are very much in tandem with the ones reported in table 1, across different values of L.

The latter, however, is not rejected under the KPSS test, which implies Korean Inprod can be reasonably inferred to be I(1). The price differentials between Mexico and the U.S. follow probably a I(2) process since ADF for Incpi in second differences, not reported, according to ADF is -7.89 (0) and -6.99 (0) according to DF-GLS.

In order to address the low power of unit root tests, some mixed results are reinforced by complementary KPSS tests. For example, in first differences, while in Korea all series except price differentials appear to be I(1) at 5%, Mexican data reject the null of stationarity for exchange rates (real and nominal) and price differentials at 5%, which would suggest the two series are I(2) in Mexico according to KPSS tests. Note, however, that by both ADF and DF-GLS tests, these series are I(1). With respect to inflation differentials between Mexico and the U.S., there is consistency across the three tests. In fact, the I(2) judgment on Mexico-U.S. inflation differentials is clear from ADF, DF-GLS and KPSS tests alike.

4.2. Main Results and Discussion

Tables 2 and 3 contain Johansen cointegration tests for Korea and Mexico under the assumption of linear deterministic trend in the data, and all series with stochastic trends. The VAR lag length is chosen by a search on a maximal order of $k = 8$ based on the highest degree of coincidence of all statistic criteria involved (likelihood ratio, final prediction error, Akaike information criterion, Schwarz information criterion, and Hannan-Quinn information criterion). We conduct Lagrange Multiplier tests of serial correlation under the null that there is no serial correlation. It was never possible to reject the null at 5% for the lag-length chosen by the criteria above, suggesting the VARs are devoid of serial correlation problems.

Tables 2 and 3 collect Johansen cointegration tests for Korea and Mexico, respectively. In table 2 for Korea, there is cointegration only for the PPP model at the 5% level, although the

estimated coefficient (0.30) is statistically insignificant and far from the theoretical value of 1. The other two models fail miserably for Korea: there is no cointegration and the coefficients do not match the expected values. In table 3 for Mexico there is rejection of the null of no vectors at 5% according to the trace test for the PPP model. This implies there exists for Mexico a long-run positive relationship between price differentials and the nominal exchange rate, as postulated by PPP in equation (9a). In Korea, for all models, the null hypothesis of no cointegration is never rejected. Table 2 suggests the results are very negative for all models under Korean data.

The most interesting result is, however, the augmented productivity model in table 3 for Mexico, representing vector (9c) above. In this case, both maximal eigenvalue and trace tests reject the null of zero cointegration at the 1% level. More importantly, the coefficients make economic sense in all cases. First, higher productivity differentials towards the tradable sector in Mexico yields an exchange rate appreciation (-1.96), although the standard error is large. Second, higher productivity differentials towards the tradable sector in the U.S. leads to a statistically significant exchange rate depreciation (+4.12). This magnitude appears to be large but is not entirely unheard of. A recent study by Chinn and Alquist (2002), for example, reports robust elasticities of the Euro-USD real exchange rate with respect to log productivity differentials between 4 and 5. Finally, price differentials move positively with the nominal Mexican Peso exchange rate: a statistically significant +0.76 coefficient, away from 1.0, but still positive and significant.

Given the evidence in table 1 that uncovered a probable I(2) for price differentials between Mexico and the U.S. (Incp_i), we perform next Stock and Watson (1993) dynamic ordinary least squares (DOLS) estimations in table 4. All model specifications are listed at the

top of table 4 and the lagged and lead terms on $(p - p^*)$ are differenced twice to achieve stationarity in the case of Mexico.² For each country, we estimate two sets of equations with 2 leads and lags for all terms as suggested by Stock and Watson (1993) and also eliminate the insignificant differenced terms based on a 10% or less confidence level.

The “PPP only” model supports a unit estimate of $(p - p^*)$ in Korea and close to 1 in Mexico. In both cases, the coefficient is statistically significant and in agreement with theory. Higher Mexican price levels with respect to the U.S. require a higher Mexican peso spot rate. The standard productivity model implies statistically significant 1.65 or 1.50 levels for the domestic productivity differential in Korea and much higher 5.89 or 7.91 β_1 coefficients in Mexico. The positive level found for β_1 does not match one’s priors since higher productivity of tradables implies a depreciation of the real exchange rate, in violation of equation (7). The U.S. productivity level is found positive and statistically significant only for Mexico in the specification with all leads and lags (11.32). This matches equation (7) since a higher U.S. productivity level would contribute to a stronger dollar and a lower peso, which would imply a positive correlation with the real exchange rate. However, the results on β_1 seriously cast doubt (again) on the standard productivity model for both countries.

Observe next the evidence from the augmented productivity model at the lower part of table 4. Consider the results for Mexico first at the right side of the table. The negative β_1 coefficient at a little over -2 suggests that higher domestic productivity of tradables implies an appreciation of the real exchange rate as formulated in equation (7). Also, the β_2 coefficient of over 2 implies that the U.S. productivity level affects positively the nominal exchange rate in

² Recall that in the demand for money estimations in Stock and Watson (1993), the net national product price deflator (p , in logarithms) is judged to be either $I(1)$ or $I(2)$. This motivates their specification of p as $I(2)$ and the lower triangular representation for an $I(d)$ process satisfying a set of conditions supporting the DOLS estimations. The $I(2)$ feature of prices is found in other economies as well. See, for example, Muscatelli and Spinelli (2000) for Italian money demand in the very long run.

Mexico according to both specifications with all leads and lags and the more parsimonious one. Finally, the β_3 coefficient is very close to its theoretical value of one in Mexico, implying that rises in the Mexican inflation with respect to U.S. inflation are accompanied by depreciations of the Mexican Peso exchange rate. For Korea, however, the augmented productivity model does not fit so well. Indeed, the only theoretical “correct” coefficient is the one for inflation differentials (β_3), varying between 0.60 and 0.64. The other two are statistically significant but with the contrary sign to theoretical predictions.

We offer three plausible explanations for the complete failure of the augmented productivity model in Korea. First, the benchmark adopted is the U.S. economy, which at first glance appears much more important in Mexico than in the Korean case.³ However, the U.S. is the largest South Korean trade partner today and South Korea is the United States' 7th largest trading partner, 6th largest export market and 4th largest market for agricultural goods, according to the Korea International Trade Association (<http://www.kita.org>). This suggests the benchmark route does not sound particularly convincing. Although skeptic, we estimate once more the models above under Japanese and Korean annual data, since monthly data for sector breakdown in Japan was unavailable. Of course, the estimates in this case suffer from low power of the cointegration tests but the results were never satisfactory for the won/yen nominal and real exchange rates as function of relative sector productivities in Japan and Korea.⁴

The second possible explanation lies in the trend of the explanatory variables in the augmented productivity model. Under Korean data, the relative traded goods productivity gap

³ Canzoneri et al. (1999), for example, report the productivity-based model is much more successful when the German mark (DEM) is used instead of the USD. Indeed, the benchmark argument appears in other studies. Due to abnormally high fluctuations in the USD during the 80s, Strauss (1996) estimates the productivity model under the DEM and the French Franc as benchmarks with more supportive findings.

⁴ Results of unit roots, Johansen, and Stock and Watson cointegration tests for the won/yen estimations are available upon request, as well as a new data description. There was considerable more variation in estimates than the estimates above under quarterly data. Results were very sensitive to lag-length and, especially, to elimination of insignificant terms in the Stock and Watson procedure.

is growing overall but suffers a slowdown in the mid to late 80s. This contrasts to the monotonically increasing traded goods productivity gap in Mexico. Since higher domestic traded goods productivity gap implies an appreciation of the currency, the fact that the series fails to behave uniformly may contribute to a coefficient contrary to the expected in the Korean case. In fact, the β_1 coefficient is positive and significant in Korea (varying from 0.75 to 0.79).

The third possibility is the institutional framework associated with the exchange rate regime. Inspection of figures 1 and 2 indicates that nominal and real exchange rates move closely together in the two countries. In other words, the source of the fluctuation in q is solely the randomness in the spot rate. A co-movement between nominal and real exchange rates can only happen when inflation differentials are relatively stable, which occurs only in the more recent past. The Mexican Peso adopted various pegged exchange regimes during the eighties, until December of 1994 when the financial crisis erupted. Similarly, The Korean Won was effectively pegged to the USD until March 1990, when the authorities adopted the so-called market average rate (MAR) system, in which supply and demand determined exchange rates subject to a daily price level [Takagi (1996)]. In 1997, the Won started to float in the wake of Asian currency crisis turmoil. Before the nineties, however, there was room for fluctuations in the two exchange rates, which could explain the failure of the standard productivity model and the good fit of the augmented model for Mexico, for example. The implication of this standpoint is that any model of the exchange rate is better built with financial (monetary aggregates, interest rates, etc.) rather than real (productivities) series, along the lines of Chinn and Meese (1995) and Chinn (1998).

5. Concluding Remarks

This paper finds that the traditional PPP model performs remarkably well for the Mexican peso under quarterly data from 1983 to 2000. Cointegration is found under both Johansen and Stock and Watson methodologies and the coefficient of the nominal exchange rate with respect to price differentials in Mexico varies narrowly from 0.94 to 1.04 across methods. The standard productivity model, however, linking the real exchange rate to domestic and foreign productivity differentials between traded and non-traded sectors is strongly rejected for both countries, in contrast to Strauss (1996) for annual data (1960-1990) of industrial economies and to Tille et. al. (2001), who also use annual data for around 30 years.

Finally, the performance of the augmented productivity model is extremely good for the Mexican peso against the USD: as Mexican traded goods productivity rises, the nominal exchange rate appreciates, while as U.S. traded goods productivity rises, the nominal exchange rate depreciates, and the effect of price differentials on the exchange rate is either 1.08 or 1.09.

The results of the productivity-based model are much less supportive for Korea. Borrowing from Canzoneri et al. (1999), we replicate the model for Japan as the benchmark country against Korea and find poor and unstable estimates. Another possible reason for the failure of the model in Korea is that the relative traded goods productivity gap grows overall but suffers a slowdown in the mid to late 80s. This pattern contrasts to the monotonically increasing traded goods productivity gap in Mexico. Finally, the strong co-movement between nominal and real exchange rates may suggest a model grounded on financial variables [e.g., Chinn and Meese (1995)], rather than on real ones. Though appealing, this interpretation leaves unexplained why the standard productivity model for Korea and Mexico is so much worse *economically* than those for industrial countries [Strauss (1996) and Tille et. al. (2001)].

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Figure 1. Korean series from 1970:1 to 2000:4: q (the log of the real exchange rate), \lnnerkr (the log of the nominal exchange rate), $\ln cps$ (overall price differentials between Korea and the U.S.), $\ln prodkr$ (productivity differentials between tradable and non-tradable sectors in Korea), and $\ln produs$ (sector productivity differentials in the U.S.)

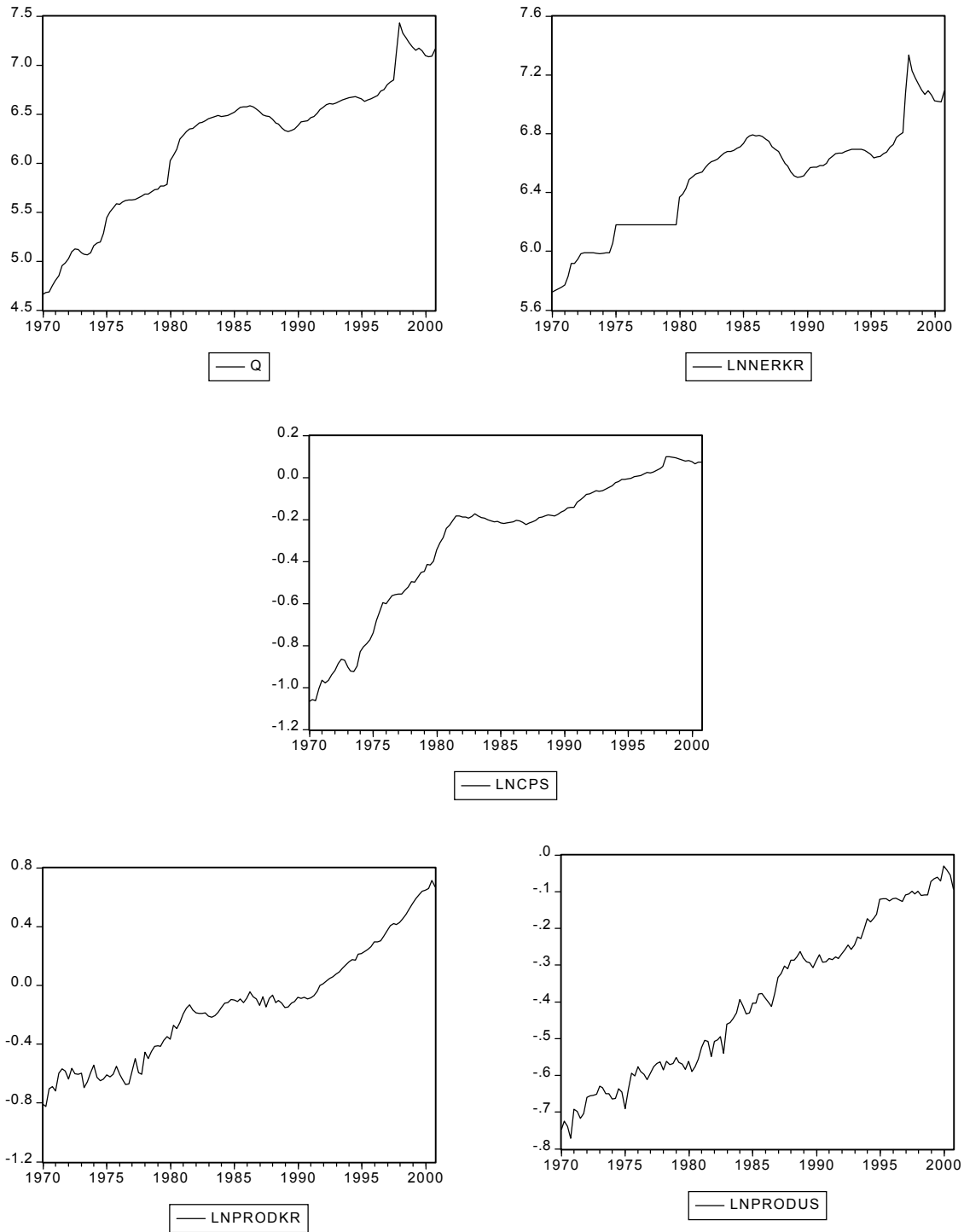


Figure 2. Mexican series; 1983:1 to 2000:4: q (the log of the real exchange rate), \lnnermx (the log of the nominal exchange rate), \lnpcps (overall price differentials between Mexico and the U.S.), \lnprodmx (productivity differentials between tradable and non-tradable sectors in Mexico), and \lnprodus (sector productivity differentials in the U.S.)

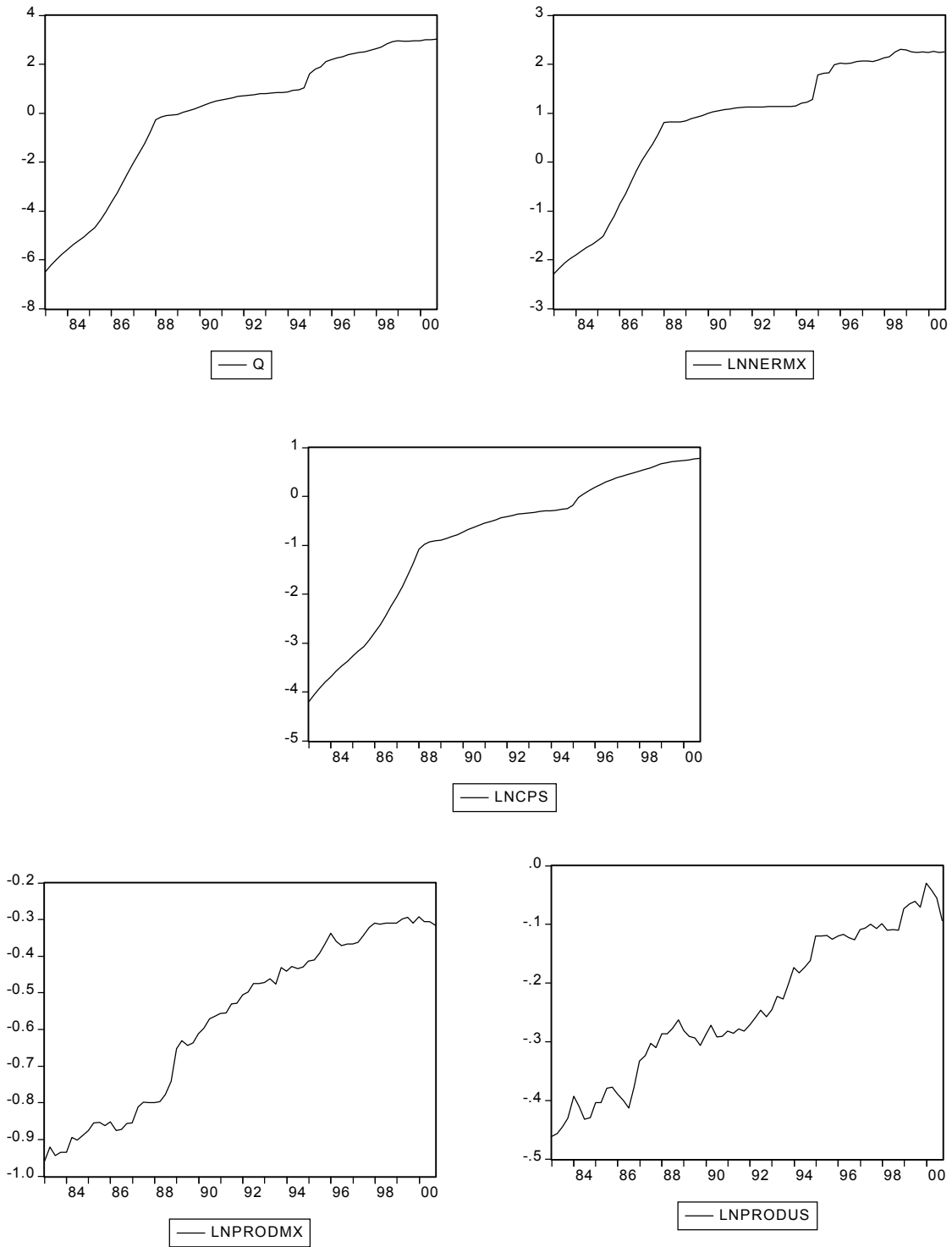


Table 1. Unit Root Tests (Quarterly):
Korea and USA: 1970.1-2000.4; México: 1983.1-2000.4.

Unit Root Test	q	lnner	lnpnt	lnprod	lncps
Series in Levels [statistic (k)] for tests with constant and trend					
ADF					
Korea	-2.38 (3)	-2.63(3)	-2.92(1)	-1.90(0)	-1.76(1)
USA			-2.37(1)	-3.18(4)*	
Mexico	-1.65(1)	-2.27(3)	-2.68(1)	-0.57(0)	-1.83(1)
DFgls					
Korea	-1.51(1)	-2.15(1)	-1.41(1)	-2.03(0)	-0.65(1)
USA			-1.51(4)	-3.15(4)**	
Mexico	-0.99(1)	-1.65(3)	-1.60(1)	-0.94(0)	-1.32(1)
KPSS					
Korea	0.44***	0.30***	0.50***	0.23***	0.52***
USA			0.24***	0.17***	
Mexico	0.31***	0.28***	0.12*	0.28***	0.33***
Series in First Differences [statistic (k)] for tests with constant only					
ADF					
Korea	-7.18(0)***	-7.43(1)***	-15.79(0)***	-10.31(1)***	-6.65(0)***
USA			-8.07(3)***	-9.99(1)***	
Mexico	-3.07(0)**	-4.67(0)***	-6.23(0)***	-8.11(0)***	-2.14(0)
DFgls					
Korea	-7.21(0)***	-7.59(0)***	-4.88(2)***	-1.38(8)	-6.61(0)***
USA			-14.00(0)***	-10.55(0)***	
Mexico	-2.68(0)***	-1.90(2)*	-5.97(0)***	-3.16(1)***	-1.55(0)
KPSS					
Korea	0.250	0.10	0.26	0.15	0.74***
USA			0.12	0.06	
Mexico	0.70**	0.57**	0.22	0.33	0.78***

Notes: The variables are defined as follows: **q** stands for the log of the real exchange rate; **lnner** stands for the log of the nominal exchange rate; **lnpnt** stands for the log of the relative price of nontradables; **lnprod** stands for the log of traded to nontraded sectors productivity; **lndpi** stands for the log of the relative general price level. ADF (k) refers to the Augmented Dickey-Fuller t-tests for unit roots, k is the selected lag length, DF gls (k) refers to Elliott-Rothenberg-Stock Dickey Fuller GLS test statistic for unit root, and KPSS refers to the Kwiatkowski-Phillips-Schmidt-Shin test. For the series in levels, the ADF (k), DF gls (k), and KPSS of each entry are estimated with a constant and trend as suggested by visual plots in figure 1. For the unit root tests in first-differences the test has only a constant. In the ADF and DF-gls tests, k is determined by the Campbell-Perron's lag length selection procedure developed formally in Ng and Perron (1995). The method starts with an upper bound, $k_{\max}=8$, on k. If the last included lag is significant, choose $k = k_{\max}$. If not, reduce k by one until the coefficient of the last lag becomes significant (we use the 5% value of the asymptotic normal distribution to assess significance of the last lag). If no lags are significant, set $k = 0$. In the KPSS test the spectral estimation method is the Bartlett kernel and the bandwidth is be verified for different values of the lag truncation parameter. Reported in the table are the statistics for lag truncation = 4; see the text for explanation. The symbols *, **, and *** attached to the figure indicate rejection of the null of no-stationarity at the 10%, 5%, and 1% levels, respectively.

Table 2. Cointegration Tests among Key Variables in Korea (1970:1-2000:4).

Model Specification	Max. Eigenv. Test	5% Crit. Val.	Trace Test	5% Crit. Val.	Null Hyp. on Coint. Vectors (C.V.)	VAR Lag Length Criteria
PPP MODEL						
[s, (p-p*)]	15.02*	14.07	19.19*	15.41	No C.V.s	2
Vector: q= 0.30(p - p*) (0.20)	4.17*	3.76	4.17*	3.76	At most 1	[LR, FPE, AIC, SC HQ]
PRODUCTIVITY MODEL						
[q, (a _T - a _{NT}), (a* _T - a* _{NT})]	9.63	20.97	14.71	29.68	No C.V.s	2
Vector: q= 2.53(a _T - a _{NT}) (1.10) -2.58 (a* _T - a* _{NT}) (1.99)	4.53	14.07	5.08	15.41	At most 1	[LR, FPE AIC]
	0.56	3.76	0.56	3.76	At most 2	
AUGMENTED MODEL						
[s, (p-p*), (a _T - a _{NT}), (a* _T - a* _{NT})]	21.46	27.07	43.34	47.21	No C.V.s	2
Vector: s= +1.95(a _T - a _{NT}) (0.65) -2.40(a* _T - a* _{NT}) (1.13) +1.15(p - p*) (0.48)	16.87	20.97	21.88	29.68	At most 1	[FPE,
	4.89	14.07	5.01	15.41	At most 2	AIC,
	0.12	3.76	0.12	3.76	At most 3	HQ]

Notes: The variables are defined as in table 1. Intercept and trend are included in cointegrating equation and test VAR. The data in levels have linear trends but cointegrating equation has only intercepts. The symbols ** indicates significance (rejection of the null hypothesis) at the 1% level and * indicates significance at the 5% level. Below the calculated values of the estimated coefficients by the Johansen cointegration method are marginal significance levels (p-values) with the exact probability. From the k-order VAR model, the ΔX_t and X_{t-k} are regressed on a constant and $\Delta X_{t-1}, \dots, \Delta X_{t-k+1}$. The obtained residuals R_{0t} and R_{kt} are used in the construction of the residual product matrices matrix S_{ij} . The matrix of cointegrating vectors is then estimated as the eigenvectors associated with the eigenvalues $\lambda_1 > \lambda_2 > \dots > \lambda_r > 0$ found as the solution to $|\lambda S_{kk} - S_{k0}S_{00}^{-1}S_{0k}|$. The test statistics (for n=2,3,4 variables in the system) are based on the maximal eigenvalue test and the trace test. The maximal eigenvalue test (null is r cointegration vectors against the alternative of r+1 cointegration vectors) is based on the statistic: $\lambda_{\max} = -T \ln(1 - \lambda_{r+1})$, where T is the sample size, r is the number of cointegrating vectors, and λ_i are the eigenvalues above. The trace test (null is at most r cointegration vectors against the alternative of more than r cointegration vectors) is based on the statistic:

$$\lambda_{\text{trace}} = -T \sum_{i=r+1}^p \ln(1 - \lambda_i).$$

The lag-length selection criteria used were given by combination of the sequential modified likelihood ratio (LR) test, the final prediction error (FPE), and the Akaike (AIC), Schwarz (SIC) and Hannan-Quinn (HQIC) information criteria. Below the used lag-length for each specification are the criteria adopted.

Table 3. Cointegration Tests among Key Variables in Mexico (1983:1-2000:4).

Model Specification	Max. Eigenv. Test	5% Critical Value	Trace Test	5% Critical Value	Null Hyp. on Coint. Vectors (C.V.)	VAR Lag Length Criteria
PPP MODEL						
[s, (p-p*)]	11.78	14.07	16.72*	15.41	No C.V.s	2
Vector:	4.94*	3.76	4.94*	3.76	At most 1	[LR, FPE, AIC, SC HQ]
s= 1.04(p - p*) (0.02)						
PRODUCTIVITY MODEL						
[q, (a _T - a _{NT}), (a* _T - a* _{NT})]	21.04*	20.97	34.89*	29.68	No C.V.s	2
Vector:	7.93	14.07	13.85	15.41	At most 1	[LR, FPE, AIC, HQ]
q= 28.30(a_T - a_{NT}) (6.13)	5.92*	3.76	5.92*	3.76	At most 2	
-16.79 (a* _T - a* _{NT})						
(9.98)						
AUGMENTED PRODUCTIVITY MODEL						
[s, (p-p*), (a _T - a _{NT}), (a* _T - a* _{NT})]	39.06**	27.07	89.34**	47.21	No C.V.s	8
Vector:	34.37**	20.97	50.28**	29.68	At most 1	[LR, FPE, AIC]
s= -1.96(a_T - a_{NT}) (1.35)	15.65*	14.07	15.91*	15.41	At most 2	
+4.12 (a* _T - a* _{NT})	0.27	3.76	0.27	3.76	At most 3	
(1.42)						
+0.76(p - p*)						
(0.11)						

Notes: The variables are defined as in table 1. Intercept and trend is included in cointegrating equation and test VAR. The data in levels have linear trends but cointegrating equation has only intercepts. The symbols ** indicates significance (rejection of the null hypothesis) at the 1% level and * indicates significance at the 5% level. Below the calculated values of the estimated coefficients by the Johansen cointegration method are marginal significance levels (p-values) with the exact probability. From the k-order VAR model, the ΔX_t and X_{t-k} are regressed on a constant and $\Delta X_{t-1}, \dots, \Delta X_{t-k+1}$. The obtained residuals R_{0t} and R_{kt} are used in the construction of the residual product matrices S_{ij} . The matrix of cointegrating vectors is then estimated as the eigenvectors associated with the eigenvalues $\lambda_1 > \lambda_2 > \dots > \lambda_r > 0$ found as the solution to $|\lambda S_{kk} - S_{k0} S_{00}^{-1} S_{0k}|$. The test statistics (for n=2,3,4 variables in the system) are based on the maximal eigenvalue test and the trace test. The maximal eigenvalue test (null is r cointegration vectors against the alternative of r+1 cointegration vectors) is based on the statistic: $\lambda_{\max} = -T \ln(1 - \lambda_{r+1})$, where T is the sample size, r is the number of cointegrating vectors, and λ_i are the eigenvalues above. The trace test (null is at most r cointegration vectors against the alternative of more than r cointegration vectors) is based on the statistic:

P

$\lambda_{\text{trace}} = -T \sum_{i=r+1}^p \ln(1 - \lambda_i)$. The lag-length selection criteria used were given by combination of the sequential modified

likelihood ratio (LR) test, the final prediction error (FPE), and the Akaike (AIC), Schwarz (SIC) and Hannan-Quinn (HQIC) information criteria. Below the used lag-length for each specification are the criteria adopted.

Tabla 4. Stock and Watson (1993) DOLS Cointegration Tests of the Models.

PPP only: Korea: $s_t = \beta_0 + \beta_1(p - p^*)_t + d_p(L) \Delta(p - p^*)_t + \varepsilon_t$

Mexico: $s_t = \beta_0 + \beta_1(p - p^*)_t + d_p(L) \Delta^2(p - p^*)_t + \varepsilon_t$

Standard Productivity: $q_t = \beta_0 + \beta_1(a_T - a_{NT})_t + d_a(L) \Delta(a_T - a_{NT})_t +$

$\beta_2(a^*_T - a^*_{NT})_t + d_{a^*}(L) \Delta(a^*_T - a^*_{NT})_t + \varepsilon_t$

Augmented: Korea: $s_t = \beta_0 + \beta_1(a_T - a_{NT})_t + d_a(L) \Delta(a_T - a_{NT})_t + \beta_2(a^*_T - a^*_{NT})_t + d_{a^*}(L) \Delta(a^*_T - a^*_{NT})_t +$

$\beta_3(p - p^*)_t + d_p(L) \Delta(p - p^*)_t + \varepsilon_t$

Mexico: $s_t = \beta_0 + \beta_1(a_T - a_{NT})_t + d_a(L) \Delta(a_T - a_{NT})_t + \beta_2(a^*_T - a^*_{NT})_t + d_{a^*}(L) \Delta(a^*_T - a^*_{NT})_t +$

$\beta_3(p - p^*)_t + d_p(L) \Delta^2(p - p^*)_t + \varepsilon_t$

Estimated Model	Korea (all leads and lags)	Korea (only significant leads and lags)	Mexico (all leads and lags)	Mexico (only significant leads and lags)
PPP Only				
β_0	6.83*** (0.04)	6.81*** (0.04)	1.65*** (0.03)	1.65*** (0.03)
β_1	0.94*** (0.07)	0.98*** (0.06)	0.94*** (0.02)	0.94*** (0.02)
T (sample size)	119	122	66	70
Adj. R ²	0.873	0.883	0.985	0.989
Standard Prod.				
β_0	6.52*** (0.21)	6.58*** (0.12)	5.63*** (0.62)	5.95*** (0.55)
β_1	1.65*** (0.39)	1.50*** (0.21)	5.89* (3.19)	7.91*** (2.94)
β_2	0.08 (0.71)	0.26 (0.37)	11.32** (5.51)	7.60 (4.77)
T (sample size)	119	121	67	68
Adj. R ²	0.868	0.868	0.903	0.905
Augmented Prod.				
β_0	6.55*** (0.08)	6.56*** (0.08)	1.14*** (0.16)	1.11*** (0.15)
β_1	0.79*** (0.18)	0.75*** (0.14)	-2.16*** (0.37)	-2.03*** (0.35)
β_2	-0.75*** (0.33)	-0.70*** (0.30)	2.48*** (0.56)	2.06*** (0.59)
β_3	0.61*** (0.11)	0.64*** (0.09)	1.08*** (0.06)	1.09*** (0.05)
T (sample size)	119	121	66	67
Adj. R ²	0.910	0.915	0.992	0.992

Notes: The method of estimation is the DOLS developed by Stock and Watson (1993). The models are: the PPP only (s as function of p-p*), the standard productivity model (the real exchange rate explained by sector productivity differentials in the two economies) and the augmented productivity model (the exchange rate explained by sector productivity and inflation differentials in the two economies) The estimates include 2 leads and 2 lags of the first differences in the regressions. In the equations above, d(L) represents the polynomial in terms of the lag operator. Below the reported coefficients are the standard errors computed by Newey-West correction for heteroskedascity and autocorrelation.