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BUSINESS CYCLES IN MEXICO AND THE UNITED STATES: DO THEY SHARE COMMON MOVEMENTS?

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In this document I apply a recently developed econometric technique to prove the existence of common movements between time series. Said methodology is used to test and measure the existence of common cycles between the economies of Mexico and the United States for the 1993-2001 period. It is found that both economies share a common trend and a common cycle. Also, given the existence of one common cycle between these economies, it is found that transitory shocks affecting Mexico's GDP are more important than when a conventional trend-cycle decomposition methodology is applied. Finally, it is shown that there are efficiency gains in forecasting by considering the common cycle restriction in a bivariate vector error correction model that includes the Mexican and the U.S. GDPs.

JEL classification codes: C32, O51, O54

Key words: time series models, U.S. GDP, Mexican GDP

I. Introduction

The Mexican economy has opened up since the mid 1980's. Undoubtedly, this openness has led to an increase in the importance of the shocks coming from external sources. In fact, various authors have found that the economic growth of Mexico is not only conditioned on the behavior of its fundamental macroeconomic variables, but also on the dynamics of international markets.

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Mejía (1999, 2000) and Torres (2000), for example, showed that the Mexican economy responds significantly to developments in the US economy, especially since the North-American Free Trade Agreement (NAFTA) was signed. This result is in line with those obtained in studies that analyze economic interactions among several countries. Anderson et al. (1999), for instance, identified a significant relationship between trade openness and the synchronization of economic cycles in a set of 37 countries.

The economic relationship between Mexico and the U.S. is evident in the evolution of some of their economic indicators. For example, it is apparent that, since 1993, Mexico's GDP shares its trend behavior with the U.S. GDP (Figure 1). This fact can be seen more clearly in the annual growth rate of the series (Figure 2).

Nevertheless, during the 1980s and the beginning of the 1990s the synchronization of the real sectors of both economies was unclear. Perhaps because, during that period, the main economic links between Mexico and the U.S. were the large amounts of external debt incurred by the former, which were not reflected in common movements between the Mexican and the U.S. GDPs. However, the synchronization of GDPs in Mexico and the U.S. became evident with the implementation of the NAFTA.

Given these facts, one cannot help but to wonder if it is possible to support the hypothesis that the Mexican and the United States economies share common movements, both in the short run and in the long run. If that is the case, the obvious question is: what are the consequences of a synchronized behavior regarding the measurement of the business cycle in Mexico and the U.S.?

The purpose of this document is to answer the above questions. First, in order to test for the existence of co-movements between the Mexican and the U.S. economies, I apply an econometric technique recently developed. Particularly, I implement the methodology proposed by Vahid and Engle (1993), where the authors present a test for the presence of common cycles between macroeconomic time series, conditioned on the existence of cointegration relationships. And second, once the parameters that characterize this relationship are obtained, the estimated cyclical pattern is contrasted with that obtained by applying a commonly used trend-cycle decomposition

Figure 1. Mexican and U.S. Quarterly GDP, 1993-2001
(Logarithms, Seasonally Adjusted Series)

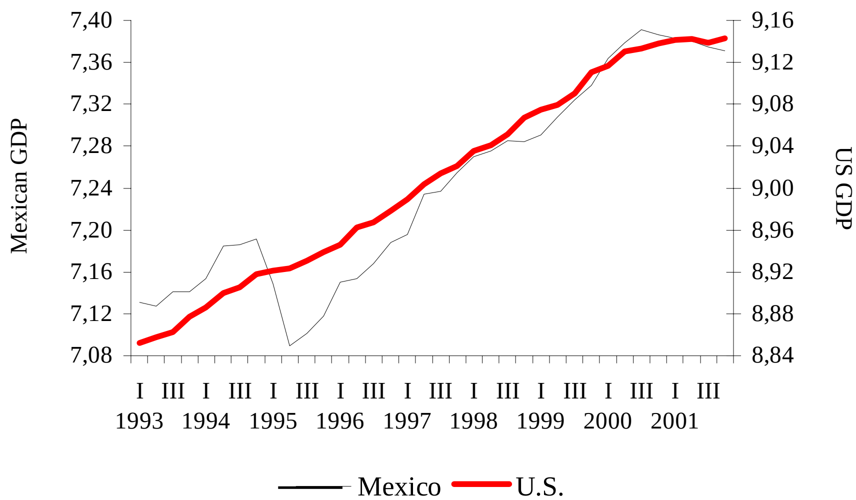
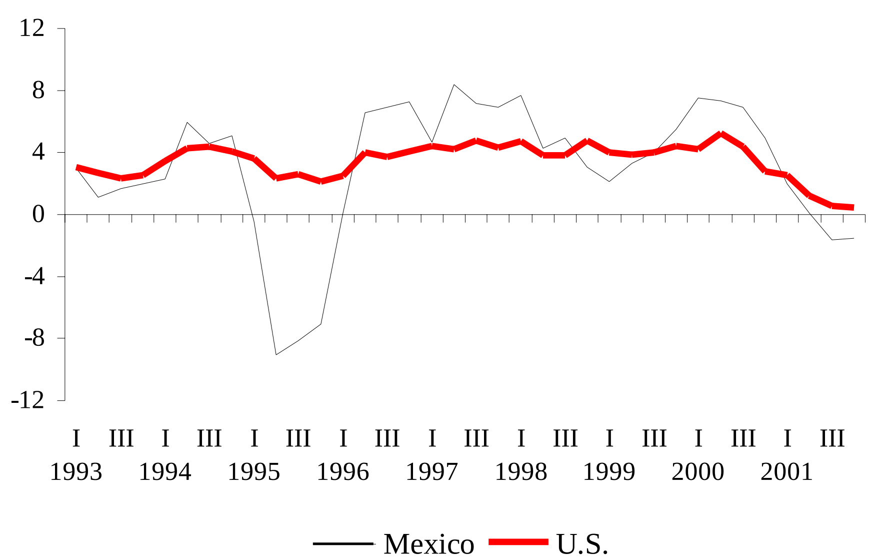


Figure 2. Mexican and U.S. Quarterly GDP, 1993-2001
(Yearly Percentage Change over the Same Quarter,
Seasonally Adjusted Series)



methodology. Attention is focused on evaluating the relative importance of transitory shocks and on the efficiency gains in forecasting obtained by including the estimated co-movements restrictions in a model that explains the behavior of the Mexican GDP.

The rest of the paper is organized as follows. The next section presents a review of the literature related to the existing theories and applications that deal with the topic of common movements between macroeconomic time series. In section III, the methodology employed in this analysis is described and the issues related to the implementation of said methodology are presented. In section IV, the results of the analysis are reported and commented on. Finally, section V presents the concluding remarks.

II. Literature Review

A. Theory

There is a wide range of statistical tools available to establish if a set of time series exhibit common movements. The theory of cointegration (Engle and Granger, 1987; Stock and Watson, 1988; Johansen, 1991) and the theory of common features (Engle and Kozicki, 1993) are a couple of them.

When a group of non-stationary time series that are integrated of order one, I(1), are cointegrated, there exist at least one linear combination of them that is stationary and integrated of order zero, I(0), so the number of stochastic trends in the system is reduced by a number equal to the existing combinations that produce cointegrating relations.

The theory of common features between macroeconomic time series is a relatively recent research topic, initiated by the seminal work in Engle and Kozicki (1993). The authors proposed a test to determine the number of common features, and hence the number of common cycles, within a group of stationary variables. The test consists in determining the significance of the relevant history of the variables in the system, by imposing overidentifying restrictions in an instrumental variables regression, where the set of instruments is defined by the past history of the variables within the system.

Similarly, Vahid and Engle (1993) derived a test to determine the existence of common cycles within a group of non-stationary series, which is conditioned on the presence of cointegration in the system. Essentially, they developed a test based on that proposed by Engle and Kozicki (1993). However, Vahid and Engle take into account long run restrictions as well. Additionally, they also show that when the number of cointegrating (r) and cofeature (s) vectors sum up to the number of variables in the system ($r + s = n$), it is possible to obtain a trend-cycle decomposition of the variables in the system *à la* Beveridge and Nelson (1981).[1] Beveridge and Nelson asserted that every macroeconomic time series integrated of order one has a representation equal to the sum of a trend and a cyclical component.

However, when the condition $r + s = n$ is not met, i.e. $r + s < n$, it is possible to use the state space representation of the cointegrated system in order to derive its trend-cycle decomposition, as proposed by Proietti (1997) and Hecq et al. (2000).

Also, even when the serial correlation common feature restrictions are not significant, it is possible to test and estimate models where the cycles of the series in the system are not exactly synchronized by using the generalized method of moments technique, developed by Vahid and Engle (1997). This kind of analysis is known as codependent cycles. A codependence vector is present when there exists at least a linear combination of a set of stationary variables that has a serial correlation order lower than the series included in the system.

As it was mentioned before, in this document I use the Vahid and Engle (1993) methodology for determining the existence of a common trend and a common cycle between the Mexican and the U.S. GDPs.

B. Empirical Analyses

The subject of identifying common cycles between macroeconomic time series has been approached from two main streams: one refers to the analysis

[1] Gonzalo and Granger (1995) derived a methodology to obtain the trend-cycle decomposition through the parameters estimated for the cointegration space without considering the common cycle analysis. However, for the case $r + s = n$, both the Vahid and Engle and the Gonzalo and Granger methodologies are equivalent. See Proietti (1997) for greater details.

of common cycles present in domestic macroeconomic variables (Vahid and Engle, 1993, U.S.; Issler and Vahid, 2001, U.S.; Herrera and Castillo, 2003, Mexico); the other branch tests the hypothesis of international common cycles between a set of countries (Engle and Kozicki, 1993; Mejía, 1999, 2000; Torres, 2000).

The presence of common movements between the economies of Latin American countries has been studied from a variety of perspectives, without finding significant evidence of it in a multivariate setting. However, the evidence of co-movements is significant at the bivariate level. Engle and Issler (1993), for example, found that while Argentina and Brazil share both long and short run co-movements, Mexico does not have similar trend or cyclical behavior with any of those countries. Also, Arnaudo and Jacobo (1997) considered four South-American countries (Argentina, Brazil, Paraguay and Uruguay) and found, just like Engle and Issler, significant synchronization only between Argentina and Brazil. Finally, Mejía (1999, 2000), using a wide sample of countries, found no evidence of a Latin American common cycle. However in a bivariate context the author found significant co-movements between several countries (Argentina-Brazil, Argentina-Peru, Bolivia-Venezuela, Brazil-Peru, Chile-United States, Argentina-Bolivia, Mexico-Venezuela and Brazil-United States).

The evidence found in studies analyzing common cycles between the economies of Mexico and the United States is mixed (Mejía, 1999, 2000; Torres, 2000). Mejía considered a sample period from 1950 through 1995, where he could not reject the null hypothesis of a common cycle between both economies. Nevertheless, by partitioning the sample in two periods (one from 1948 through 1979 and the other from 1980 through 1997), Torres found that in the first period there was, indeed, a common cycle but by the second partition of the sample that behavior was not present.

This work extends the above studies in two fashions: firstly, it includes more recent data (1993-2001); secondly, it considers the inclusion of cointegrating restrictions (common trend) in a test for the existence of common cycles. This approach pretends to explore the implications of the growing economic integration between Mexico and the United States for the economic cycle of both economies, especially for the Mexican cycle.

III. Econometric Approach

I follow the Vahid and Engle (1993) econometric methodology to test and estimate the co-movements, both in the long and in the short run, of a group of time series. This methodology is briefly described next.[2] Consider a VAR representation for a set X_t of n variables that are integrated of order one in levels:

$$X_t = A_1 X_{t-1} + A_2 X_{t-2} + ... + A_p X_{t-p} + \Phi D_t + \varepsilon_t \tag{1}$$

where $A_i, \forall\, i = 1, 2, \ldots, p$, are coefficient matrices associated to the autoregressive structure, D_t is a matrix of deterministic variables and Φ its associated matrix of coefficients, and ε_t is a white noise random disturbance.

From equation (1) it is possible to rewrite the VAR in error correction form:

$$\Delta X_t = A X_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + \Phi D_t + \varepsilon_t \tag{2}$$

where $A = \sum_{i=1}^{p} A_i - I$ and $\Gamma_i = -\sum_{j=i+1}^{p} A_j$.

When $A = 0$, the presence of cointegration is rejected and the VAR should be estimated in first differences, with no restrictions placed on the long run specification. However, if A has reduced rank ($r < n$) there will be r linear combinations of X_t that are I(0), which means that the series are cointegrated with r cointegrating vectors. In this case, the matrix A can be decomposed in two matrices of rank r, $A = \alpha\beta'$, where α is the matrix of speed of adjustment and β is the matrix containing the cointegrating vectors. There exist several techniques to test and estimate cointegrating relations. In this paper, as suggested by Vahid and Engle, I apply the methodology proposed by Johansen (1991).

[2] For technical details consult the original paper, or, alternatively, the reviews contained in the applications of this methodology in Issler and Vahid (2001) and Herrera and Castillo (2003).

Once the cointegrating vectors have been estimated, we proceed to test the existence of common cycles restricted to the inclusion of the long run parameters. This test is carried out by first computing squared canonical correlations between the differenced variables in the system and its relevant history, which is defined by the lagged error correction term and the lagged differenced elements in X_t. The squared canonical correlations $\left(\lambda_i^2 \ \forall i = 1, \ 2, \ldots, n\right)$ are obtained by solving an eigenvalue problem of a matrix constructed with the aforementioned variables.[3] At this stage, we are looking for a rotation of the system such that the right hand side variables (the relevant history) will have no correlation with each other, so the exclusion of one of them will have no effect on the coefficients of the rest. Formally, the canonical correlations can be represented as follows:

$$CanCor \left\{ \Delta X_t \, , \begin{pmatrix} \beta ' X_{t-1} \\ \Delta X_{t-2} \\ \vdots \\ \Delta X_{t-p+1} \end{pmatrix} \mid D_t \right\}.$$

The presence of common cycles means that there exist s linear combinations of the differenced variables characterized by the following properties: they cannot be forecasted, and they eliminate the serial correlation pattern present in the system. The number of such linear combinations equals the number of squared canonical correlations equal to zero, which has to be a number smaller than or equal to the number of variables in the system minus the cointegrating rank, i.e. $s \leq n - r$. Each of such linear combinations is a cofeature vector and they are stacked in a $n \times s$ matrix $\tilde{\alpha}$, which is orthogonal to the cointegration space. Vahid and Engle showed that the existence of $\tilde{\alpha}$ implies that $\tilde{\alpha}' \, \Gamma_i = 0$ and $\tilde{\alpha}' \, \alpha = 0$. This means that there are $n - s$ common cycles in the system.

The test statistic for the null hypothesis that the s squared canonical correlations equals zero $\left(\lambda_j^2 = 0 \, , \forall j = 1, \ 2, \ldots, s\right)$ and that the largest are greater than zero, amounts to compute the following statistic:

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[3] See Anderson (1958) for greater detail.

$$C(p,s) = -(T-p-1)\sum_{j=1}^{s}\log\left(1-\lambda_j^2\right) \tag{3}$$

Under the null hypothesis that the dimension of the cofeature space is at least s, this test statistic has a χ^2 distribution with degrees of freedom equal to: $s\,(n\,p\,+\,r) - s\,(\,n - s)$.

Regarding the trend-cycle decomposition in a Beveridge-Nelson fashion, given that we are interested in the case where there exists at least one cointegrating vector ($n - r$ common stochastic trends) and one cofeature vector ($n - s$ common cycles), I briefly illustrate the implications when both r and s are greater than zero. In that case, there are two possibilities: (i) $r + s < n$ or (ii) $r + s = n$. For case (i), the Proietti (1997) or the Hecq et al. (2000) techniques may be used to obtain the trend-cycle decomposition. While for case (ii), the technique proposed by Vahid and Engle (1993), or the one suggested by Gonzalo and Granger (1995), is appropriate to decompose the system into its permanent and transitory components.

As emphasized at the beginning of this section, I follow the Vahid and Engle methodology. That decomposition is based upon the cointegrating and the cofeature vectors and its formal representation is as follows:

$$X_t = \tilde{\alpha}\,(\tilde{\alpha}'\tilde{\alpha})^{-1}\tilde{\alpha}'X_t + \beta\,(\beta'\beta)^{-1}\,\beta'\,X_t \tag{4}$$

$$= \textit{trend component} + \textit{transitory component.}$$

Finally, it is worth noting that the efficiency gains in forecasting of including significant short and long run restrictions can also be measured. This is done by imposing the common cycle-common trend restrictions in a vector error correction model (VECM) estimated by a full information method (full information maximum likelihood, FIML, for instance), and comparing the forecasts of such a model with those of an unrestricted VECM.

The pseudo structural form of the system that includes the restrictions mentioned above is defined as follows: first rotate $\tilde{\alpha}$ to have enough exclusion-normalization restrictions to avoid the indeterminacies that arise because any linear combination of the columns contained in it will be a cofeature vector. Vahid and Engle recommend making a rotation such that $\tilde{\alpha}$ includes an s-dimensional identity matrix:

$$\tilde{\alpha}_{n\times s} = \begin{pmatrix} I_s \\ \tilde{\alpha}^*_{(n-s)\times s} \end{pmatrix}.$$

Then, the FIML model to be estimated has the following reduced from:

$$\Delta X_t = \begin{pmatrix} -\tilde{\alpha}^{*\,\prime} \\ I_{n-s} \end{pmatrix} \left(\Gamma^*_1 \Delta X_{t-1} + \ldots + \Gamma^*_{p-1} \Delta X_{t-p+1} + \alpha^* \beta' X_{t-1} \right) + \upsilon_t$$

where Γ^*_i and α^* represent the partitions of Γ_i and α, respectively, corresponding to the bottom $n - s$ reduced form (unrestricted) VECM equations.

IV. Empirical Evidence

A. Data

The analysis considers quarterly seasonally adjusted data for the sample period 1993:1-2001:4.[4, 5] Non-seasonally adjusted data for the Mexican GDP were obtained from the Banco de Información Económica, Instituto Nacional de Estadística, Geografía e Informática, Mexico, and its seasonally adjusted series was estimated by applying the routine X-12 of the U.S. Department of Commerce, U.S. Census Bureau. Data for the U.S. GDP were obtained from the U.S. Department of Commerce, Bureau of Economic Analysis.

B. Setting the Number of Lags

The number of lags p to be included in the analysis is set according to the

[4] According to Hecq (1998), seasonal adjustment of a data set considered for common feature analysis will raise both size and power distortions, which will ignore and spuriously signal (low *t* ratios for the cofeature coefficients) common cycles, respectively. Cubadda (1999) claimed that seasonal adjustment could make a common cycle turn to a codependent cycle. Nevertheless, the results reported in the analysis below point to a significant common cycle between the seasonally adjusted series for the Mexican and the U.S. GDPs, which also exhibits a significant coefficient to describe such relation (non-spurious).

[5] The qualitative results presented in this paper do not change if the sample considered is reduced to 1994:1-2001:4 (not reported), where 1994:1 is the starting date of the NAFTA.

multivariate Akaike Information Criterion (AIC), which is obtained from an unrestricted VAR in levels. It follows that the number of lags to be used in the analysis in first differences will be equal to the number of lags defined in the above VAR (in levels) minus one, p-1.

The AIC indicated that the optimal number of lags in levels is 2 (p = 2), hence, the number of lags to be included in the analysis in first differences should be one. Additionally, other forms of misspecification rejected with that lag structure are: serial correlation, heteroskedasticity, non-normality and ARCH effects.[6]

C. Cointegration Analysis

Currently, there exists a vast literature that analyzes the U.S. GDP data generating process (DGP). Many studies conclude that its DGP is characterized by a unit root, which means that it is an I(1) variable (Cochrane, 1994 surveys some of these studies). Also, in the case of Mexico's GDP the evidence points to a unit root present in its DGP (Castillo and Díaz-Bautista, 2002). For the purpose of this analysis, I take these results as evidence in favor of the non-stationary nature of both GDPs.

Given that both series are integrated of the same order, it is possible to apply Johansen's (1991) cointegration (trace) test to determine if there exists a common trend between the Mexican and the U.S. GDPs. The results of the test are presented in Table 1.

I am not able to reject the null hypothesis of one cointegrating relationship between these variables at the 95 percent significance level, which means that both series share a common trend. Table 1 also presents the estimated cointegrating space from which the long run (normalized) elasticity of Mexican GDP with respect to the U.S. GDP is inferred and equals 0.84, with a standard error of 0.30. This implies that the equilibrium response of the Mexican GDP to a permanent shock in the U.S. economy is less than unity and significant. For example, an increase in factor productivity that induces a one percent sustained raise in the level of the United States economic activity will

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[6] Results not reported for brevity.

Table 1. Johansen's Cointegration Trace Test: Mexican and U.S. GDPs

Hypothesized number of cointegrating relations	Eigenvalues	Trace statistic	95% critical values
None	0.141	16.12	15.41
At most 1	0.034	3.01	3.76

Normalized cointegrating vector: log (Mexican GDP) = 0.84 log (U.S. GDP)
[0.301]

Notes: Seasonally adjusted series, 1993:1-2001:4. Number in brackets is the standard error of the coefficient.

produce a permanent increase in the long run level of Mexico's GDP equal to 0.84 percentage points.[7]

By looking at Figure 1, and given the fact that this is a bivariate system, one might think that this analysis should consider a linear trend and, perhaps, also a structural break. However, by including a linear trend in the cointegrating analysis, the long run relationship between both economies still holds (though the long run elasticity diminishes to 0.78). Similarly, a break in the trend is not needed in the analysis, since the shock observed at the beginning of 1995 can be well captured by a discrete jump in the intercept, not affecting significantly the estimated stochastic trend.

D. Existence of a Common Cycle

Conditional on the existence of cointegration, the Vahid and Engle (1993) methodology for testing and estimating common cycles is applied to determine if there exists a common cycle between the GDPs of Mexico and the United

[7] Torres and Vela (2002), for example, show that this transmission mechanism is via the trading sector.

States. The estimated test statistics are reported in Table 2. According to these, it is not possible to reject the presence of a common cycle between the Mexican and the U.S. economies at the 1.3% significance level.

Table 2. Vahid and Engle's Cofeature Test: Mexican and U.S. GDPs

Null hypothesis	Squared correlations	Cofeature statistic test	DF	p-value
s > 0	0.023	0.85	2	0.654
s > 1	0.346	16.11	6	0.013

Normalized cofeature vector: Δlog (Mexican GDP) = 3.78 Δlog (U.S. GDP)
[0.808]

Notes: Seasonally adjusted series, 1993:1-2001:4. Number in brackets is the standard error of the coefficient.

In Table 2 the normalized cofeature vector is also shown. It can be seen that the fluctuations of the Mexican GDP around its trend are 3.78 times those corresponding to the U.S. GDP, with a standard error of 0.81. The parameters estimated for the common cycle have another interpretation, the effect of a non-permanent one percent shock to the U.S. economy is reflected in an immediate shock to the Mexican economy of 3.78%.

The findings presented up to this point show that the Mexican economy overreacts in cyclical frequencies to shocks coming from the U.S. economy, however, the equilibrium response of Mexico's GDP to a permanent one percent change in the growth rate of the United States is less than a percentage point.

E. Trend-cycle Decomposition

When the cointegrating and the cofeature vectors exist (which means that they are unique), and the cointegration space (r) and the cofeature space (s) sum up to the number of variables in the system ($r + s = n$), it is possible to span the space projected by the n variables in the system as a trend-cycle decomposition à la Beveridge-Nelson,[8] as argued in Vahid and Engle (1993). This is easily done by manipulating the cointegrating and the cofeature vectors. This means that when $r + s = n$, the cointegrating and the cofeature vectors form a base to express each variable in the system in its trend-cycle decomposition as a linear combination of the set of variables within the same system (Vahid and Engle, 1993).

The matrices that span the basis to decompose each series in the system in a trend-cycle form are presented in Table 3. In terms of equation 4, Table 3 reports two squared matrices of order two which describe the basis for the trend component ($\tilde{\alpha}(\tilde{\alpha}'\tilde{\alpha})^{-1}\tilde{\alpha}'$) and for the transitory part $\left(\beta(\beta'\beta)^{-1}\beta'\right)$. Given that they are reduced rank matrices (equal one, as a result of the cointegrating and the cofeature restrictions), the Mexican GDP trend can be expressed as a linear combination of the one corresponding to the U.S. GDP. The same applies for the cycle. In this simple case ($n = 2 = r + s = 1 + 1$) the base for the trend component is a linear combination of the cofeature vector (where, if we normalize each row of this matrix by its first column, we obtain the normalized cofeature vector), and the rows of the base for the cyclical part are linear combinations of the cointegrating vector.

Table 3. Trend-Cycle Decomposition as Linear Combinations of the Variables: Mexican and U.S. GDPs

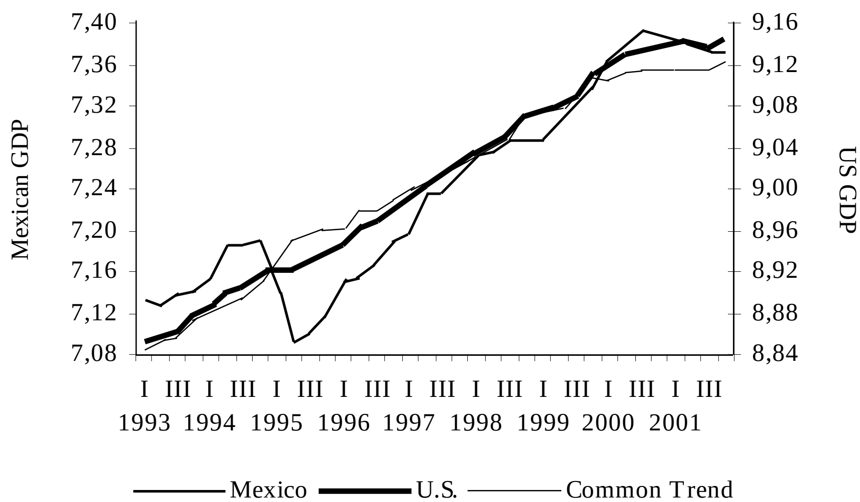
Variable	Mexican GDP	U.S. GDP
	Trends	
Mexican GDP	-0.28	1.07
U.S. GDP	-0.34	1.28
	Cycles	
Mexican GDP	1.28	-1.07
U.S. GDP	0.34	-0.28

Notes: Seasonally adjusted series, 1993:1-2001:4. The cycles are demeaned in order to obtain cycles averaging to zero.

[8] The details about this kind of decomposition can be found in Issler and Vahid (2001), Appendix B.

Figure 3 illustrates the estimated common trend for both economies. The variability observed in Mexico's GDP is greater than that corresponding to the U.S. GDP. As a consequence, the estimated common trend follows the U.S. variable closely, while the Mexican output exhibits well-pronounced fluctuations. This volatility is also reflected in the higher magnitude of the cyclical pattern of the Mexican economy with respect to the U.S. cycle (Figure 4).[9]

Figure 3. Mexican and U.S. GDPs and their Common Trend, 1993-2001 (Logarithms)

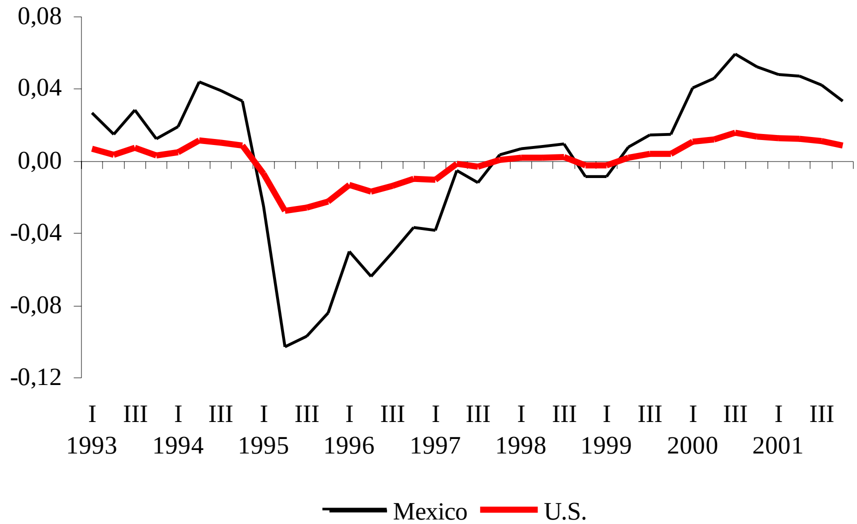


F. Comparison with the Hodrick-Prescott Filtering Technique

In order to contrast the dates and magnitudes of the cyclical fluctuations estimated above for the Mexican economy with those obtained by a commonly used trend-cycle decomposition methodology, a Hodrick-Prescott (HP)

[9] Another issue to consider is that the apparent non-cyclical behavior of the cyclical component for both economies could be due to the period considered for this analysis, which can be part of a larger fluctuation. With the availability of more data, this possibility could be explored in future research.

Figure 4. Cyclical Patterns in Mexican and U.S. GDPs, 1993-2001 (Gap between Observed and Trend Output in Logarithms)



univariate filter is estimated. The results are presented in Table 4. Turning points are defined by visual inspection of Figure 5.

The cyclical component estimated for the Mexican economy with the methodology applied in this document (first row of $\beta(\beta'\beta)^{-1}\beta' X_t$ in equation 4) displays comparable turning points (recession-expansion) at the same dates with respect to the cycle obtained with the HP filter. Despite the fact that both methodologies exhibit the same dates for the turning points,[10] notice that the transitory shocks estimated with the Vahid and Engle methodology are of greater magnitude (more importance) relative to those obtained by applying the commonly used methodology. Given that the technique used in this document considers long and short run restrictions, it is possible to argue that the previously mentioned fact follows from the efficiency gains derived with the implementation of the Vahid and Engle methodology in measuring economic cycles.[11]

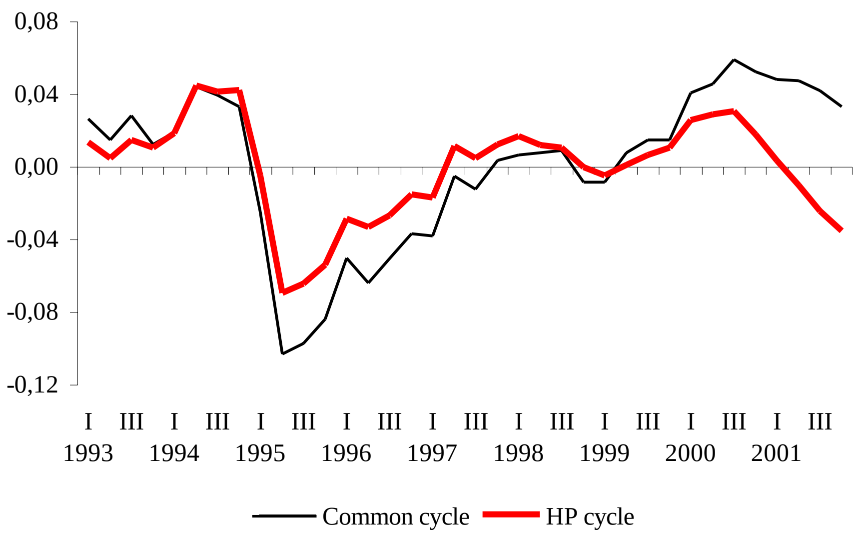
[10] Except in one case, which presents a mismatch of two quarters.

[11] Ideally, this should be demonstrated by means of a variance decomposition, however this is not possible due to the insufficient number of observations used in this analysis.

Table 4. Turning Points in Mexico: Shared Common Cycle with the U.S. and Hodrick-Prescott Cycle

	Common cycle	HP cycle
Peak	1994:II	1994:II
Trough	1995:II	1995:II
Peak	1998:III	1998:I
Trough	1999:I	1999:I
Peak	2000:III	2000:III

Figure 5. Cyclical Movements in the Mexican Economy: Common Cycle with the U.S. and Hodrick-Prescott Cycle, 1993-2001 (Gap between Observed and Trend Output in Logarithms)



G. Out of Sample Forecasts for Mexican and U.S. GDPs

Theoretically, it can be shown that when the predictions of two different models are compared, one unrestricted and the other including certain restrictions, there exist efficiency gains when the restrictions are imposed.

Empirically, the efficiency gains can be tested, first, through hypothesis tests to evaluate the validity of the restrictions imposed, and then by comparing the predictive ability of each specification. Vahid and Issler (2002) show that ignoring significant short run restrictions in a vector error correction model lead to the forecasts at business-cycle horizons to be less accurate than when they are considered.

In this section, I compare the results from computing out of sample forecasts for the period 2000:1-2001:4 from two representations of a bivariate system with the Mexican GDP and the U.S. GDP. The purpose of this exercise is to determine the existence of efficiency gains in forecasting when common-trend and common-cycle restrictions are imposed. Mean Squared Error (MSE) and the determinant of the MSEs matrix (|MSE|) are used as measures of precision of the forecasts. In this case, I compare two models: a vector error correction model without restrictions (UVECM); and a vector error correction including the common cycle restriction (RVECM), which is estimated by the Full Information Maximum Likelihood method (FIML). The results are presented in Table 5.

Table 5. Out of Sample Forecasts: Mean Squared Error (MSE) for Mexican and U.S. GDPs, UVECM vs. RVECM, 2000-2001

	UVECM	RVECM
Mexican GDP	0.0831	0.0250
U.S. GDP	0.0678	0.0225
| MSE |	0.0056	0.0006

The MSE results indicate that both Mexico's and the U.S. GDPs display a shorter forecast error in the restricted model than in the UVECM case. At the same time, the determinant of the MSEs matrix for the RVECM is almost one tenth of that obtained through the model without the common cycle restriction. Based on these results, it is evident that by imposing the restriction of a common cycle between the Mexican and the U.S. GDP there is a significant improvement in forecast ability.

V. Conclusions

It is evident that in recent years the Mexican and the U.S. economies have become more integrated due to the growing trade openness among them, particularly after the implementation of the NAFTA. The present analysis finds the existence of statistically significant common movements between both economies since 1993, both at trend and cyclical horizons.

A couple of stylized facts can be mentioned: first, transitory movements in the Mexican economy seem to be more important when the Vahid and Engle methodology is employed than when a Hodrick-Prescott filter, commonly used in this kind of analysis, is applied; second, the binding restrictions implied by a common cyclical pattern, between the Mexican and the U.S. economies, generates efficiency gains in forecasting, compared to the predictions obtained from a model that only includes long run (cointegrating) restrictions.

The efficiency gains obtained through this kind of analysis might be useful, for example, in forecasting the short-term growth of the Mexican economy conditioned to U.S. economic progress. Moreover, the methodology used in this document can be employed using monthly (instead of quarterly) economic indicators, such as industrial production, so one can take advantage of a structural analysis of this type. Some of these topics will be addressed in future research.

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