

Explaining Apparent deviations from Covered Interest Parity: Evidence from Mexico

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Abstract

This paper tests and quantifies the effects of reduced funding liquidity conditions on the covered interest parity (CIP) relating the U.S. Dollar-Mexican Peso market. To this end, a vector error-correction model is estimated. Results suggest, first, that apparent deviations from the CIP disappear when measures of funding liquidity for market participants are considered. Second, the exchange rate forward premium and the U.S. interest rate adjust towards the CIP cointegrating relationship. Finally, a structural analysis shows that deviations from CIP are mostly determined by shocks on the funding liquidity in the U.S. while funding liquidity conditions in Europe also have a non-negligible role. From the policy perspective, the paper underlies the relevance of funding liquidity measures when assessing whether the foreign exchange market works efficiently. As ever, there are some caveats in the analysis to consider. First, funding liquidity measures may shift from non- to stationary regimes. Second, market participants may not be able to fund their liquidity at reference rates. The financial series present considerable ARCH-like behaviour, this may be a source of information to explore in further work.

JEL Classification: C51, G12, F31, F37.

Keywords: Covered interest parity; vector autoregression models; vector error-correction models; forward premium; exchange rate; funding liquidity.

Sobre la prelación de las restricciones al crecimiento económico: abogando por la perspectiva basada en la huella ecológica

Resumen

Este artículo se prueba y cuantifica los efectos de la reducción en el acceso al fondeo sobre la paridad de tasas cubierta (CIP) entre el dólar de Estados Unidos y el peso mexicano. Para ello se estima un modelo de vectores con corrección de error. Los resultados sugieren que, primero, las desviaciones aparentes de la CIP desaparecen cuando se incluyen medidas de la liquidez. Segundo, la prima forward en el mercado del tipo de cambio y la tasa de interés externa se ajustan hacia la relación de cointegración. Finalmente, choques sobre la liquidez en Estados Unidos y en Europa determinan las desviaciones de la CIP. En términos de política, se subraya la relevancia de las medidas de liquidez al evaluar el funcionamiento del mercado cambiario. Existen algunas limitaciones en el análisis que merecen consideración. Primero, las medidas de liquidez podrían cambiar entre regímenes de no-a estacionario. Segundo, los participantes de mercado podrían enfrentar condiciones de liquidez distintas a las que las tasas de interés sugieren. Las series financieras muestran un comportamiento ARCH, ello será objeto de trabajo futuro.

Clasificación JEL: C51, G12, F31, F37.

Palabras clave: Paridad de tasas cubierta; modelos de vectores autorregresivos; modelos de vectores con corrección de error; prima forward; tipo de cambio; liquidez de fondeo.

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* No source of funding for research development



1 Introduction

Deviations from the covered interest parity (CIP) across different foreign exchange markets, have been the norm in the aftermath of the financial crisis that erupted in the summer of 2007. These have been observed in the U.S. Dollar(USD)-Sterling, USD-Euro, and USD-Yen parities (Du et al. 2018). Deviations from the CIP in the USD-Mexican Peso (MXN) parity have also been observed. Notwithstanding the MXN is among the most traded currencies in the world (BIS, 2016), these deviations reveal apparent risk-less arbitrage opportunities.

The computation of deviations from CIP using USD-MXN future contracts, the yields on U.S. Treasury Bills and Mexico's CETE, all for 1-month, displayed in Figure 1 shows the dates of the major shocks from the start of the financial crisis. In particular, the behaviour of these deviations is consistent with economic theory for the period January 2005-August 2007, since it is around zero. Major disruptions generated by severe financial stress events that unfolded in 2007 (e.g. the Bear Stearns buyout, the Lehman Brothers failure, the Troubled Asset Relief Program (TARP) approval within the U.S. Congress and the AIG's bailout) are represented by extremely large deviations from the zero level. After the second half of 2009, the main sources of financial stress were the fragility of Europe's banking sector and regulation introduced in the aftermath of the financial crisis.

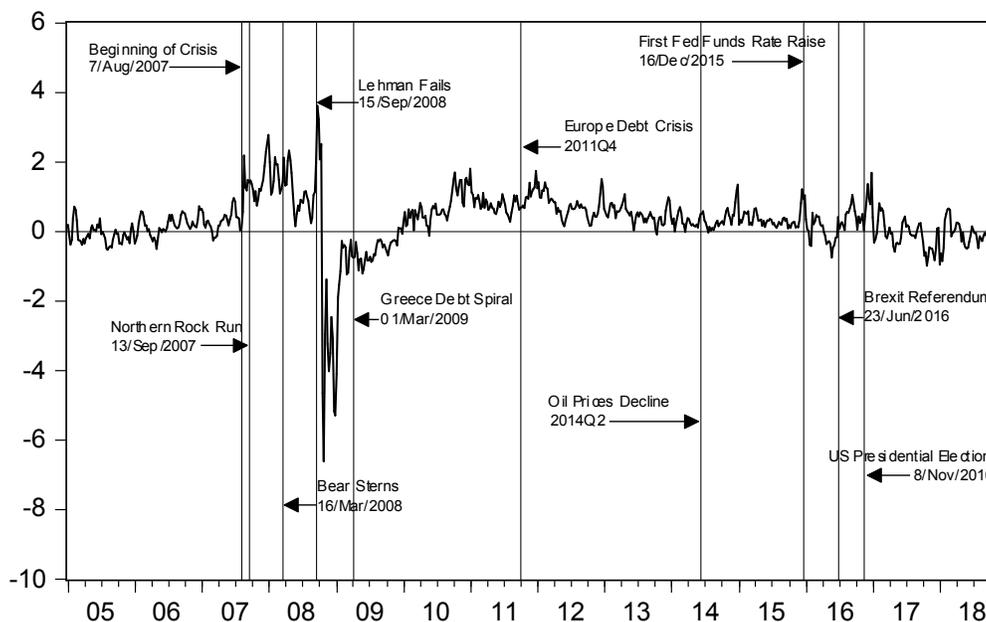


Figure 1: Deviations from the CIP in percentage points for the period January 2005 to October 2018 computed using expression (2.4) below. Weekly averages of daily data. Source: Bloomberg.

A priori, it seems plausible the described events and regulation would cause funding liquidity in financial markets to fall. Reduced liquidity is defined as (i) the reduced availability of funds to undertake financial transactions in assets that are perceived riskier than those with the highest rate by rating agencies, and (ii) higher quality or quantity collateral demands. Given that underlying the CIP is a forward contract on the exchange rate, a further source of frictions might be found in the risk for the counterpart fulfilling their part of the contract. In this paper, the relative importance of reduced liquidity conditions in explaining deviations

from CIP is tested and quantified.

The sequence of financial shocks from the summer of 2007 through the summer of 2012 and new regulation introduced since, suggests that there were two different sources of funding liquidity shocks: (1) U.S.-born, and (2) Europe-born. Their role in explaining deviations from the CIP is analysed accordingly. Shocks born in the U.S. are straightforward to relate to deviations from the CIP provided, among other reasons, that: (i) the U.S. economy is Mexican exports' main destination, and so changes in the demand for goods made in Mexico affect the risk premium firms must pay when they are looking for sources of funding; (ii) the largest Mexican firms have access to credit valued in USD, which may be provided by U.S. based financial firms; and (iii) there is relatively high mobility of labour between the two economies, which changes the relative prices of production inputs, among them the interest rate.

Europe-born shocks are not as directly related to deviations from the CIP but they are still relevant for the analysis. The lack of a clear relationship between deviations in the CIP and Europe-born shocks, may bias analysts and economists' judgement towards the claim that these shocks have only a short-run effect. An econometric-based answer to the latter statement is, however, absent in the literature.

To determine whether shocks in funding liquidity generate deviations from the CIP, and whether these deviations are persistent, a vector error-correction model (VECM) is used which includes the following variables: (i) the 1-month exchange rate forward premium; (ii) the yield on 1-month U.S. Treasuries; (iii) the yield on 1-month CETE; (iv) the LIBOR-OIS spread in USD; and (v) the analogous measure for Europe.¹

The main findings are, first, apparent deviations from the CIP are persistent. Once the effects of funding liquidity shocks in the U.S. and Europe are included, however, an error-correcting relationship emerges (i.e. the CIP is satisfied). This means that arbitrage opportunities do not exist, once a measure to the costs of funding liquidity for the agents is included. Second, both the exchange rate forward premium and the U.S. interest rates adjust towards the error-correcting relationship that explains deviations from the CIP. Interestingly, the yield on 1-month Mexican CETE has its own stochastic trend, notwithstanding the strong relationship between U.S. and Mexico's economies and the small open economy status of the latter. Third, a structural analysis confirms that the estimated model is consistent with intuition relating the variables included in the VECM. Finally, it confirms that deviations from CIP are explained by changes in funding liquidity conditions. In particular, over time the relative importance of U.S. funding liquidity conditions is similar to that of deviations from CIP themselves. Moreover, contrary to what may be expected, the relationship between liquidity conditions in Europe and deviations from CIP is non-negligible.

To produce a survey of the vast literature around the CIP would require a great amount of space, but among the pioneers are Aliber (2), Frenkel and Levich (18), and Frenkel and Levich (19). Indeed, up until the 2007 financial crisis, tests confirmed the validity of the CIP. These were made using high-frequency data sets (Akram et al., 2008, and Fong et al., 2010), panel data and single equation econometric techniques (Mancini-Griffoli and Ranaldo, 2011, Ivashina et al., 2015, Sushko et al., 2016, and Avdjiev et al., 2017), and time series multiple equation techniques (Clarida and Taylor, 1997, Shigeru and Shu, 2006, Roll et al., 2007, and Eichengreen et al., 2012). Cointegration techniques have also been used to determine if the CIP is satisfied (Peel and Taylor, 2002, and Sushko et al., 2016).

Since deviations from CIP in major foreign exchange markets have been the norm in the post 2007 crisis period, literature has turned its attention to find possible explanations. This has taken the form of omitted variables analysis, much in the same way the present paper does for an Emerging Market economy (Mancini-

¹As pointed out by a referee, theoretically the funding liquidity measure for Mexico is also associated with deviations from CIP. Including such measure does not change the main results. In particular, the liquidity measure is not able to account for deviations from CIP. These estimates are shown in Appendix A.

Griffoli and Rinaldo, 2011, Levich, 2012, Ivashina et al., 2015, Sushko et al., 2016, Liao, 2016, Borio et al., 2016, Avdjiev et al., 2017, Cenedese et al., 2017, and Du et al., 2018). More recently, attention has turned to bands-about the CIP and one-way arbitrage, as suggested by Levich (28).

The CIP related to the USD-MXN parity is an interesting case to analyse since the financial crisis may be considered as an exogenous shock (i.e. not caused by Mexican macroeconomic outlook which at the time enjoyed sound macroeconomic indicators, and a sustainable current account deficit, IMF 2007). The references for this foreign exchange market include Carstens (9), Khor and Rojas-Suarez (25), and Bush (8). In particular, Carstens analyses the determination of the forward exchange rate for delivery after three months. In that period Mexico had a fixed (spot) exchange rate and observed considerable external imbalances concerning both its current account and outstanding external debt. This led to a generalised belief that a devaluation of the MXN with respect to the USD was about to take place. The present paper updates Carstens' analysis of the CIP for the USD-MXN market, and is the first to use a VECM framework. A further important difference relative to Carstens' work is the floating exchange rate regime under which the MXN is determined in the period analysed here (2005 to 2018).

Khor and Rojas-Suarez (25) test the CIP for Mexican internal debt denominated in USD but payable in MXN and external debt. They use cointegration analysis to assess the sovereign risk indirectly through a long-run relationship between the aforementioned debt instruments. Their work includes three main findings. First, the CIP is satisfied for most of the period 1987-1991. Second, there is indeed a long-run relationship between the interest rate on USD denominated bonds and the yield to maturity of the bonds issued in the external market. Finally, a policy implication in which they suggest that by improving internal economic conditions the interest rate of the Mexican debt will decline sensibly. Khor and Suarez-Rojas's work sheds light on the history of the CIP for Mexico in the period studied which, as in Carstens' work, has a fixed exchange rate regime and the econometric methods are single-equation based.

Bush (8) exploits a detailed data set that allows to shed light on the prevailing conditions in the market for hedging transactions that are linked to the CIP. Using regression analysis, she finds that the funding gap observed in the hedging market has a prominent role in explaining said deviations. Data used for this study covers the 2013-2018 period where the MXN is determined in floating regime. Bush's empirical model is also single-based, hence silent on possible dynamic relations among determinants of the CIP.

The measures of funding liquidity conditions used in the present paper to explain deviations from the CIP aim to satisfy the definition advanced by Brunnermeier and Pedersen (7): *funding liquidity* is concerned with how easily a trader can get funding to trade. Brunnermeier and Pedersen's model shows how under particular conditions market liquidity "dries up", is common across assets, relates to volatility, suffers flight-to-quality, and co-moves with the market.

Of particular relevance for the present paper, is the work of Peel and Taylor (34) who use a VECM to address the validity of the Keynes-Einzig conjecture: Given deviations from the CIP, arbitrage opportunities will only be realised if the premium is high enough (5% in the USD-Sterling exchange rate during the 1920's). Peel and Taylor's work inspired the econometric model and notation used in the present paper. Since the CIP may be regarded as an equilibrium condition, the VECM is a suitable tool to analyse deviations about it and to tests its validity. The VECM also allows economic interpretation and testing the validity of the parameters in the cointegrating relationship, the loading matrix and the short-run dynamics. A key difference is that bands are not estimated *a la* Keynes-Einzig, which would require non-linear estimation techniques, and are beyond the scope of the present paper.

The rest of the paper is divided in 5 further sections. Section 2 presents briefly the theory behind the

CIP. In section 3 details of the data set are provided. The econometric model and the empirical results are outlined in section 4. The concluding remarks are found in section 5.

2 Theory

2.1 Covered Interest Parity

Economic theory predicts a number of stable relationships among variables through time. Of particular interest in the field of international finance is the CIP. This relationship predicts that under market efficiency, the forward premium (given by the difference between the forward exchange rate and the spot exchange rate) must be equal to the interest rate differential among the assets in two countries (plus some transaction costs or minimum risk premium). Formally, let S_t be the spot exchange rate of USD per 1 MXN and let F_t^k be the related forward contract agreed at period t with maturity of k months. Throughout the paper, the followed convention is that the investor resides in the U.S. More generally, he can always borrow in USD and is looking to invest in Mexico. Thus, the U.S. and Mexico are the domestic and the foreign country, respectively. Also, let $i_{us,t}^k$ and $i_{mex,t}^k$ be the interest rate on the sovereign bond with maturity of k months issued in domestic currency by the U.S. and Mexico, respectively. Note that by taking $i_{us,t}^k$ to compute the CIP, it is implicitly assumed that the traders have access to loans at this rate. This may be true for those at large banks. Assuming individuals are rational and absence of financial frictions and costs, the CIP may be stated as

$$\frac{F_t^k}{S_t} = \frac{1 + i_{us,t}^k}{1 + i_{mex,t}^k}. \quad (2.1)$$

When deviations from the CIP occur, the economic theory predicts that demand and supply forces will re-establish (2.1) through arbitrage. This is, in theory inequalities provide risk-less profitable opportunities in a friction-less set-up. An example of the latter is as follows. Assume

$$\frac{F_t^k}{S_t} (1 + i_{mex,t}^k) > 1 + i_{us,t}^k. \quad (2.2)$$

If an investor in the home country borrows B USD for k months at rate $i_{us,t}^k$, he can buy B/S_t MXN. Then, he can lend his B/S_t MXN at a rate $i_{mex,t}^k$. At the same time, he signs a forward contract where he promises to deliver $(1 + i_{mex,t}^k)B/S_t$ MXN at the end of k months in exchange for $F_t^k B(1 + i_{mex,t}^k)/S_t$ USD. After this sequence of transactions, the investor has in his hand $F_t^k B(1 + i_{mex,t}^k)/S_t$ USD and he owns $B(1 + i_{us,t}^k)$ USD. If (2.2) holds, then he can pocket a profit of $F_t^k B(1 + i_{mex,t}^k)/S_t - B(1 + i_{us,t}^k)$ USD.

The arbitrage process ensures that (2.1) holds, even if there are deviations in the short run. Empirical tests of the CIP advanced by Aliber (2) or Peel and Taylor (34) are based on the following approximation to expression (2.1)

$$P_k \frac{F_t^k - S_t}{S_t} = i_{us,t}^k - i_{mex,t}^k - c, \quad (2.3)$$

where P_k is a factor that adjusts to annual terms the forward premium and c accounts for possible transaction costs and minimum risk premium. Since interest here lies in *deviations* from the CIP, these are denoted by

δ_t^k , and the forward premium by $\Phi_t^k = P_k \frac{F_t^k - S_t}{S_t}$. That is

$$\delta_t^k = \Phi_t^k - i_{us,t}^k + i_{mex,t}^k + c. \quad (2.4)$$

Figure 1 above shows relationship (2.4) for $k = 1$.

2.2 Funding Liquidity Measures

Among the key money-market indicators for "liquidity events" is the LIBOR-OIS spread, LOIS. It has been used as a market indicator of interbank liquidity conditions by central bankers and market participants as explained by Bernanke (4), Hull (20), Paulson (33), and Thornton (39). In the international finance literature, LOIS has served as a proxy for credit risk. It has a relatively simple interpretation: when LOIS increases, the funding liquidity decreases. A description of how both the LIBOR and the OIS are constructed is found in chapter 3 of Smith (37). Thus, in the absence of negative financial events, the spread between the two rates is small. In times of financial distress, however, loans signed in a LIBOR contract are riskier than overnight loans, causing LOIS to increase. Since the aim of this paper is to obtain a measure of the effects from liquidity shocks on the CIP, LOIS for the U.S. banking sector is used. To compare the effects of said shocks with those born in Europe, LOIS for the European banking sector which is the analogue measure is included in the analysis.²

3 Data

All the analysis is conducted transforming daily data provided by Bloomberg into weekly averages. This transformation helps to avoid "day effects" (e.g. Friday is more prone to suffer a sell-off if an important is due to occur during the weekend), it also reduces volatility.³ The sample period considered is January 2005-October 2018. The sample starts 36 months after the end of the U.S. recession dated at November 2001 by the NBER's Business Cycle Committee. This allows undertaking the analysis of the CIP during a period of "normal" conditions.

The focus in this paper is on deviations from the CIP in 1-month sovereign bonds given by δ_t^1 as defined in (2.3). This tenor is chosen on the basis that this market has a larger volume and turnover than other maturities in possession of non-Mexican investors. To ease notation, the super-index $k = 1$ in the variables is omitted from now on. F_t is the inverse of the exchange rate agreed on the forward contract. The inverse is taken because USD per 1 MXN is used in the analysis, whereas contracts are quoted in MXN per 1 USD. Analogously, S_t is the inverse of the spot exchange rate at date t . The variable $i_{us,t}$ is the daily yield on the 1-month U.S. Treasury Bill in annual terms. The latter is obtained from the "constant Maturity Treasury". Finally, $i_{mex,t}$ is the weekly average of the daily yield of the 28-day (1-month) Treasury Certificate known as CETE in annual terms and $P_k = 12 \times 100$, as explained above.

The use of weekly averages rather than the end-of-period data, allows to deal with the autocorrelation induced by the fact that the frequency of the original data is daily whereas F_t matures in 30 days. This is why, contrary to what is typically observed, taking weekly averages reduces autocorrelation. Figures 2 to 5 display the time series of the variables. It is worth noting that all series display a more volatile behaviour

²As mentioned above, theoretically the funding liquidity measure for Mexico is also associated with deviations from CIP. Results are not affected by excluding said measure. As shown in Appendix A such liquidity measure is not able to account for deviations from CIP.

³Appendix B contains the Bloomberg ticker id of each variable.

after the first half of 2007. In particular, they show large deviations from the trend at the beginning of the last quarter of 2008. This coincides with financial distress caused by events like the Lehman failure and the AIG's bail-out.

Regarding Figure 4 (left) two main reasons account for the near zero yield of the 1-month Treasury Bills: (1) The monetary stimulus provided by the Federal Reserve; and (2) the flight-to-quality phenomenon where the U.S. received a large capital inflow despite being the center of the financial crisis as detailed in Bernanke (4) and Paulson (33). A second period of financial turbulence, caused among other factors by the end of the zero-level of the Federal Funds rate in December 2015, is noticeable across the Figures. Moreover, particular shocks to Mexico are observed. In addition the Fed's rates increase, Figure 2 reflects the oil-prices decline in mid 2014, and the uncertainty associated with the U.S. Presidential election in November 2016.

Figure 3 displays how the forward premium has evolved within the sample. From definition if Φ_t decreases away from zero, it reflects either an appreciation of the spot MXN (or a USD depreciation) or a depreciation of the MXN (or a USD appreciation) embedded in the future contract. When Φ_t increases, it reflects an MXN depreciation (or a USD appreciation).

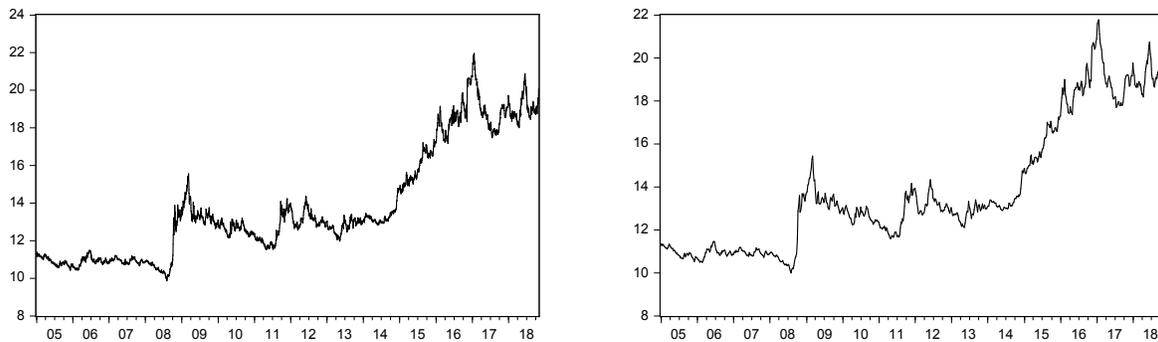


Figure 2: (Left) Spot exchange rate MXN per 1 USD, S_t^{-1} . (Right) Forward exchange rate MXN per 1 USD, F_t^{-1} . Weekly averages of daily data. Source: Bloomberg.

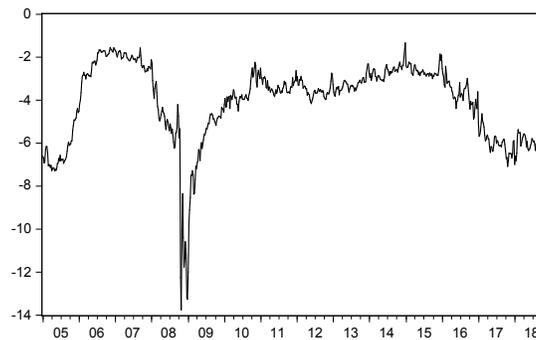


Figure 3: Forward exchange rate premium in percentage points, Φ_t . Weekly averages of daily data. Source: Bloomberg.

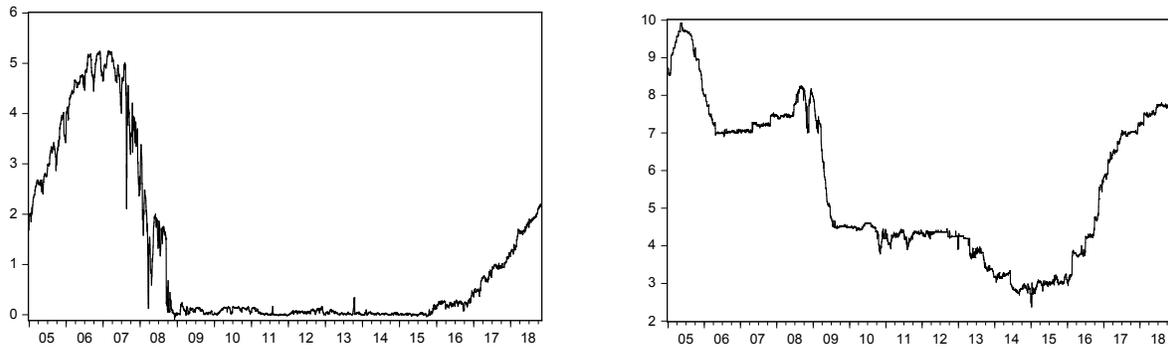


Figure 4: (Left) U.S. Treasury Bill yields for 1 month in percentage points, $i_{us,t}$. (Right) Mexican 28-day CETE in percentage points, $i_{mex,t}$. Weekly averages of daily data. Source: Bloomberg.

LOIS for the U.S. banking system, $LOIS_{us,t}$, is computed as the weekly average of the difference between the daily 3-month LIBOR in USD and the Overnight Index Swap. The variable $LOIS_{eur,t}$ stands for LOIS for Europe. This is obtained by taking the weekly average of the daily difference between the 3-month Euribor in Euros and the Overnight Index Swap, both published on the Euro Overnight Index Average (EONIA) website. Behaviour of these two measures of funding liquidity shown in Figure 5 is stable, up to the summer of 2007. In particular, $LOIS_{us,t}$ (left) displays several “hump-shape” of financial distress. In 2011 associated with sovereign debt problems in southern Europe. In 2016 associated, possibly to marked changes in regulation (see Du et al., 2018), the “Brexit” Referendum, and sharp volatility observed in markets, mainly reflecting a sharp decline in stock prices. $LOIS_{eur,t}$ (right) also displays several “hump-shaped”. None, however, after those associated with the European sovereign debt.

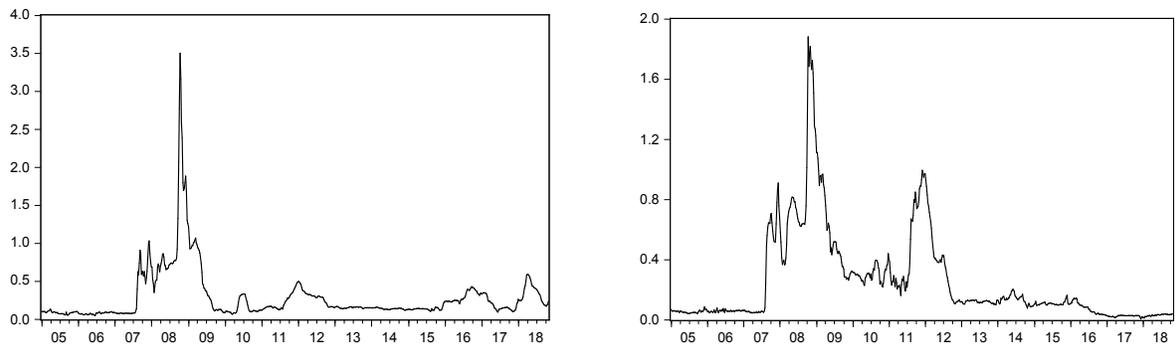


Figure 5: (Left) Libor-OIS spread in percentage points, $LOIS_{us,t}$. (Right) Euribor-EONIA rate spread in percentage points, $LOIS_{eur,t}$. Weekly averages of daily data. Source: Bloomberg.

It is worth noting that there is evidence to treat all variables as $I(1)$. In general, a battery of tests fail to reject the null-hypothesis for the presence of a unit root as Table 1 shows. A borderline case arise for $LOIS_{us,t}$. In particular, the $DF - GLS$ test proposed by Elliott, Rothenberg, and Stock (16), shown to possess the highest power, when applied to $LOIS_{us,t}$ only rejects the null-hypothesis at the 1% level.

Variable	ADF^μ	ERS	ERS^+	PP^μ
Φ_t	-2.260588	-1.536808	-0.921692	-2.870119*
$i_{us,t}$	-1.423635	-1.418677	-0.564666	-1.036137
$i_{mex,t}$	-2.798172	-1.310672	-0.906296	-1.247117
$LOIS_{us,t}$	-2.385599	-2.352037*	-1.211621	-3.089500*
$LOIS_{eur,t}$	-2.227629	-1.865794	-1.294154	-2.666757
5% Critical Value	-2.866976	-1.941444	-1.941444	-2.866868

Table 1: Unit root tests. Columns show *test-statistic* for the Augmented Dickey-Fuller test with a mean (ADF^μ), the Elliot-Rothenberg-Stock (ERS) $DF - GLS$ test with a mean, and the Phillips-Perron test with a mean (PP^μ). All tests were carried out in levels with a constant. Null Hypothesis: there is a unit root. The lag-length is chosen by E-Views according to the Modified Akaike Information Criterion with a maximum of 19. *Reject the null hypothesis at the 5% level. +Sub-sample January 2008 to October 2018.

Note, however, that if these tests are conducted from the period of the peak of the crisis onwards, from September 2008 to October 2018, all variables are unambiguously $I(1)$ as shown in Table ???. This is fairly direct to explain. $LOIS_{.t}$ observations were very stable before the crisis, given that there were no major liquidity strains between 2005 and the summer of 2008. Thus, all variables are treated as $I(1)$ in the rest of the paper.

4 Econometric Model and Estimates

Details on the econometric specification of the VECM are now provided. There are several reasons why this model is a suitable tool for the analysis. First, it allows to analyse $I(1)$ variables, and to test the existence of an equilibrium stationary relationship among them (in this case the CIP). Second, it can be transformed into its “common trends representation”, which allows writing each equation of the system as a linear combination of residuals from the VECM (i.e. a Moving Average form), and testing for the number of variables adjusting towards the CIP. Third, it allows the estimation of Impulse Response Functions (IRFs) through the corresponding orthogonalization of the residuals. Finally, it allows to obtain variance and historical decompositions under appropriate conditions of stationarity of the system.

Let $n = 5$ and define the $n \times 1$ variable vector⁴

$$X_t = (\Phi_t, i_{us,t}, i_{mex,t}, LOIS_{us,t}, LOIS_{eur,t})'$$

Let the vector autoregressive model (VAR) in its vector error-correction form of order p , VECM(p),

$$\Delta X_t = \alpha\beta'X_{t-1} + \sum_{s=1}^{p-1} \Gamma_s \Delta X_{t-s} + \varepsilon_t, \tag{4.1}$$

$$\varepsilon_t = (\varepsilon_{\Phi,t}, \varepsilon_{i_{us,t}}, \varepsilon_{i_{mex,t}}, \varepsilon_{LOIS_{us,t}}, \varepsilon_{LOIS_{eur,t}})'. \tag{4.2}$$

$$\varepsilon_t \sim N(0, \Sigma), \tag{4.3}$$

where $\Pi = \alpha\beta'$ is a $n \times n$ matrix, α and β are $n \times r$ matrices, Γ_s are $n \times n$ short-run effect matrices, ε_t is the vector containing the reduced form errors, and the $n \times n$ matrix Σ is symmetric and is allowed to be

⁴As mentioned above, inclusion of a funding liquidity for Mexico does not change the main results.

non-diagonal. Note that all elements of X_t are $I(1)$ as shown in Table 1. The latter means that for the model (4.1) to yield appropriate inference, the presence of $r(=rank(\Pi))$ cointegrating relations is needed.

4.1 Specification

Lag-Length

The first step in specifying the VECM is to choose the number of lags, p , used in estimation. Table 2 displays the Bayesian Information Criterion (BIC), the Hannan-Quinn Criterion (H-Q), and the Lagrange Multiplier Test (LM) for autocorrelation for values of $p = 1, \dots, 15$. From Table 2 it is possible to conclude that $p = 11$ will imply that residuals are free from autocorrelation, since the LM test fail to reject the null hypothesis of no-autocorrelation at the 5% level. Even though $p = 11$ may seem large, Kilian and Lütkepohl (26) document that if the true lag-length parameter is unknown, an over-parameterised VAR may yield more accurate IRFs. They also discuss why the Ljung-Box Q test may be biased when $I(1)$ series are involved, hence it is not reported here. In this paper, over-parametrization should not be a concern, provided the sample size is 722 observations.

The canonical VAR analysis no longer requires the residuals to satisfy Normality and Homoskedasticity. On the one hand, Normality is no longer a necessary condition to estimate through Maximum Likelihood, since reliance on Quasi-Maximum Likelihood techniques is enough. On the other hand, Heteroskedasticity is no longer an issue in estimating a VAR according to Juselius (24).

p	Reg	BIC	H-Q	LM(1)	LM(p)
15	76	-22.944	-24.449	0.000	0.004
14	71	-23.093	-24.500	0.001	0.001
13	66	-23.206	-24.514	0.000	0.000
12	61	-23.306	-24.515	0.000	0.000
11	56	-23.468	-24.578	0.069*	0.025
10	51	-23.541	-24.552	0.000	0.000
9	46	-23.656	-24.567	0.000	0.001
8	41	-23.767	-24.580	0.000	0.015
7	36	-23.861	-24.574	0.000	0.000
6	31	-23.969	-24.583*	0.000	0.000
5	26	-24.028	-24.543	0.000	0.000
4	21	-24.067*	-24.483	0.000	0.000
3	16	-24.054	-24.371	0.000	0.000
2	11	-23.788	-24.006	0.000	0.000
1	6	-23.294	-23.413	0.000	0.000

Table 2: p is the lag-length. Reg is the number of regressors in each equation of the VAR. BIC is the Bayesian Information Criterion on lag length determination. H-Q is the Hannan and Quinn Criterion. LM(p) is the Lagrange Multiplier Test of autocorrelation of order p , for $p = 1, \dots, 16$. * Is the suggested lag according to each criterion. Estimations were made in Dennis, Hansen, Johansen, and Juselius (12) CATS 2 in RATS.

Rank Test

One of the appealing features of model (4.1) is that it allows to test whether the equilibrium relationships predicted by economic theory are satisfied by the data. In particular, when each of the equations $(\beta' X_{t-1})_j$ is $I(0)$, the long-run relationships are satisfied. This is, a subset of elements of X_t are cointegrated. Note that

$j = 1, 2, \dots, r < n$. Another appealing feature of the model (4.1) is that it allows to impose restrictions on the matrix β . Moreover, if the number of restrictions in each row β'_j is larger than $r - 1$, these are testable. The latter means that determining r , the cointegrating rank, is a necessary condition for undertaking inference on the cointegrating relationships and the $n - r$ common trends. Johansen (23) tests is a popular way to select r .⁵

$n - r_0$	r_0	Trace-Stat	Trace-Stat*	Crit-5%	$p - value$	$p - value^*$	BIC
5	0	130.13	121.06	76.81	0.000	0.000	
4	1	70.14	61.27	53.94	0.001	0.009	-9.428
3	2	27.00	22.10	35.07	0.293	0.592	-9.414
2	3	11.56	9.49	20.16	0.497	0.693	-9.381
1	4	3.33	2.94	9.14	0.531	0.600	-9.356

Table 3: $n - r$ is the number of Common Trends. r is the cointegrating rank (i.e. the number of cointegrating relations). Trace-Stat is Johansen's Trace Statistic. Crit 5% is the critical value for the size of 5%. Trace-Stat* is Johansen's small sample corrected Trace Statistic. The null hypothesis is: cointegrating rank = r_0 . BIC is the Bayesian Information Criterion.

Table 3 contains the tests for several possible values of r denoted r_0 . In this case, Johansen's Trace Test suggests $r = 2$. Estimation of $\hat{\beta}$ yields:⁶

$$\hat{\beta}' X_t = \begin{bmatrix} \Phi_t - 0.79i_{us,t} + 0.80i_{mex,t} + 3.07LOIS_{us,t} - 0.45LOIS_{eur,t} - 0.19 \\ \Phi_t - 1.30i_{us,t} + 1.25i_{mex,t} + 1.28LOIS_{us,t} - 3.32LOIS_{eur,t} - 0.88 \end{bmatrix}. \quad (4.4)$$

Inspection of (4.4) reveals that estimates for the main variables of CIP are similar across cointegration relationships, hence imposing $r = 2$ may be redundant. A further criterion to choose r is discussed in Martin et al. (32). In particular, they suggest that a valid alternative to Johansen's test is selecting the restricted model with the least BIC. The last column in Table 3 provides said criterion. Indeed, it suggests $r = 1$. The estimation of $\hat{\beta}$ with $r = 1$ yields:

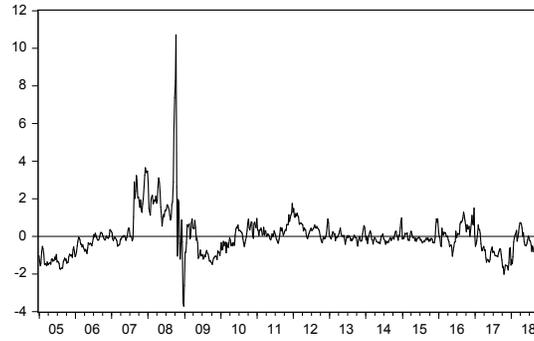
$$\Phi_t - \underset{(0.084)}{0.791}i_{us,t} + \underset{(0.074)}{0.798}i_{mex,t} + \underset{(0.570)}{3.070}LOIS_{us,t} - \underset{(0.547)}{0.448}LOIS_{eur,t} - \underset{(0.307)}{0.189}. \quad (4.5)$$

Expression (4.5) presents the cointegrating relationship.⁷ At this point it is not possible to interpret the estimates. It is possible, however, to discuss the stationarity of the cointegrating relationship. Figure 6 presents $\hat{\beta}' X_t$ and suggests that it is stationary, as desired. Unit root tests are shown in Table 4 confirming the latter. In particular, the Augmented Dickey Fuller and Phillips-Perron tests are included, without a mean or a trend. The Table also shows a $t - test$ statistic for the sample mean of $\hat{\beta}' X_t$ being zero, justifying the exclusion of any deterministic terms in the unit root tests. Results show unambiguous support for the absence of a unit root, confirming the existence of one cointegrating relationship.

⁵Estimations in the rest of the paper were carried out in Doornik and Juselius (13) CATS 3 in OxMetrics, unless stated otherwise.

⁶Full estimates of the long-run matrix $\hat{\Pi}$, loading $\hat{\alpha}$, and short-run matrices $\hat{\Gamma}_j$ for $j = 1, \dots, 10$ are not included for brevity, but are available upon request.

⁷Throughout the paper standard errors are presented in brackets unless stated otherwise.

Figure 6: Cointegrating Relationship $\hat{\beta}' X_t$.

Variable	<i>ADF</i>	<i>PP</i>	<i>t - statistic</i>
$\hat{\beta}' X_t$	-3.16	-6.08	
Sample Mean of $\hat{\beta}' X_t$			0.24
5% Critical Value	-1.94	-1.94	-1.96

Table 4: Unit root tests for the cointegrating relationship $\hat{\beta}' X_t$. Columns show *test-statistic* for the Augmented Dickey-Fuller test (*ADF*) and the Phillips-Perron test (*PP*) with no mean or trend. Null Hypothesis: there is a unit root. The lag-length for the *ADF* test is chosen according to the Modified Akaike Information Criteria with a maximum of 17. No constant was included since the *t - statistic* for the mean of $\hat{\beta}' X_t$ suggests it has zero mean, as reported in the fourth column. Estimation was carried out in EViews 10.

4.2 Testing the Theoretical Relationship

Assuming that a Data Generating Process can be modelled by a VECM(p) entails the *a priori* conjecture that level relationships among the elements of X_t are stationary. Hence, economic theory outlined in Section 2 is used to justify the only cointegration relationship: the CIP. Economic theory predicts a stable relationship among its components, which in econometric terms means that δ_t as defined in expression (2.4) is stationary. Let the parameters to be estimated of the CIP, be given by $\beta = (\beta_1, \beta_2, \beta_3, \beta_4, \beta_5, c)'$. If the data obey the CIP in equilibrium, then the estimate $\hat{\beta}$ will satisfy:

1. $\hat{\beta}_1 \approx -\hat{\beta}_2$.
2. $\hat{\beta}_1 \approx \hat{\beta}_3$.
3. $\hat{\beta}_4$ and $\hat{\beta}_5$ not statistically different from zero.

Results displayed in expression (4.5) somewhat satisfy conditions 1 and 2, however, condition 3 is far from being observed. The VECM(p) allows to test whether 3 can be imposed. Let $\tilde{\beta}^R$ be the restricted estimate of β , for the set of restrictions R . The first set of restrictions to test, $R = A$, is defined as

$$\tilde{\beta}_1^A = -\tilde{\beta}_2^A, \quad (4.6a)$$

$$\tilde{\beta}_1^A = \tilde{\beta}_3^A, \quad (4.6b)$$

$$\tilde{\beta}_4^A = 0, \quad (4.6c)$$

$$\tilde{\beta}_5^A = 0. \quad (4.6d)$$

Imposing the restriction set (4.6a)-(4.6d), the following estimates are obtained:

$$\tilde{\beta}^A X_{t-1} = \Phi_{t-1} - i_{us,t-1} + i_{mex,t-1} - \underset{(0.094)}{0.304} = \delta_{t-1}. \quad (4.7)$$

The conducted test to assess whether the restrictions imposed in (4.7) are valid suggest they are not. Table 5 presents the results of the test which rejects the null hypothesis of $\tilde{\beta}^A$ being valid.

Test of Restrictions	χ^2	<i>p</i> - value
Test set A	25.169	0.000
Test set B	4.3757	0.1122

Table 5: Test for the validity of the Restricted Model versus the unrestricted model. The test for set A assumes a $\chi^2(4)$ distribution. The test assumes a $\chi^2(2)$ distribution. Null hypothesis: Restrictions are valid.

Since the aim is to test whether changes in funding liquidity are affecting the CIP, define the restrictions set *B* as a subset of *A* where *B* is composed only by (4.6a) and (4.6b). That is, define $\tilde{\beta}^B$ as the restricted estimate of β , where $\tilde{\beta}_1^B = -\tilde{\beta}_2^B$ and $\tilde{\beta}_1^B = \tilde{\beta}_3^B$ are imposed and both $\tilde{\beta}_4^B$ and $\tilde{\beta}_5^B$ are left unrestricted. Estimating the model under the restriction set *B* yields:

$$\begin{aligned} \tilde{\beta}^{B'} X_{t-1} &= \Phi_{t-1} - i_{us,t-1} + i_{mex,t-1} + \underset{(0.402)}{2.300} LOIS_{us,t-1} - \underset{(0.430)}{0.960} LOIS_{eur,t-1} - \underset{(0.098)}{0.717} \\ &= \delta_{t-1} + \underset{(0.402)}{2.300} LOIS_{us,t-1} - \underset{(0.430)}{0.960} LOIS_{eur,t-1} - \underset{(0.098)}{0.717} \\ &= \delta_{t-1}^B. \end{aligned} \quad (4.8)$$

As shown in Table 5 the Null Hypothesis of the restriction set *B* being valid is not rejected at a 10% level. This means that there is a stationary equilibrium relationship that includes the funding liquidity measures and has a mean different from zero.

Once the validity of the relationship is established, a test to determine whether δ_{t-1}^B is stationary ensues. Figure 7 suggests that δ_{t-1}^B is stationary. The latter is confirmed by the results of the unit root tests included in Table 6. This is, δ_t^B , as defined in expression (4.8), is indeed stationary, hence cointegrated. With δ_t^B in hand, an economic interpretation for the econometric results obtained so far can be provided.

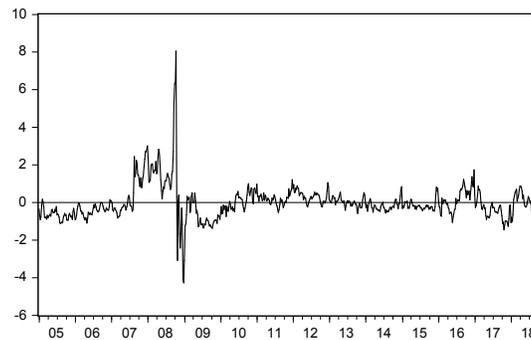


Figure 7: Cointegrating Relationship $\hat{\beta}^{B'} X_t = \delta_t^B$.

Variable	<i>ADF</i>	<i>PP</i>	<i>t – statistic</i>
$\widehat{\beta}^{B'} X_t$	-3.49	-7.28	
Sample Mean of $\widehat{\beta}^{B'} X_t$			-0.28
5% Critical Value	-1.94	-1.94	-1.96

Table 6: Unit root tests for the cointegrating relationship $\widehat{\beta}^{B'} X_t$. Columns show *test-statistic* for the Augmented Dickey-Fuller test (*ADF*) and the Phillips-Perron test (*PP*) with no mean or trend. Null Hypothesis: there is a unit root. The lag-length for the *ADF* test is chosen according to the Modified Akaike Information Criteria with a maximum of 17. No constant was included since the *t – statistic* for the mean of $\widehat{\beta}^{B'} X_t$ suggests it has zero mean, as reported in the fourth column. Estimation was carried out in EViews 10.

4.3 Economics of the Estimation

Expression (4.8) should be interpreted as an equilibrium relationship among the 5 elements of X_t . In particular, said expression may be written as

$$\delta_t = -\underset{(0.402)}{2.300} LOIS_{us,t-1} + \underset{(0.430)}{0.960} LOIS_{eur,t-1} + \underset{(0.098)}{0.717} \quad (4.9)$$

where ν_t is a mean zero stationary random variable as implied by Table 6.

The relationship between δ_t and $LOIS_{us,t}$ is negative, which can be rationalised as follows: (1) An increase in the $LOIS_{us,t}$ spread signals an increase of the perception of credit risk, thus a decrease in funding liquidity in the U.S. (2) The increase in $LOIS_{us,t}$ will cause δ_t to decrease in one of three ways: (a) USD appreciates if there is a flight to quality, diminishing the forward premium Φ_t ; or (b) $i_{us,t}$ increases since individuals demand less treasuries due to mature in the short run; or (c) both.

The relationship between δ_t and $LOIS_{eur,t}$ is positive since: (1) An increase in $LOIS_{eur,t}$ signals a decrease in funding liquidity in Europe. (2) There is a flight-to-quality to the U.S. assets. This causes: (a) an appreciation of the USD, thus Φ_t increases (closer to zero); or (b) $i_{us,t}$ decreases; or (c) both.

Finally the constant term represents the minimum profit investors demand to trade MXN and Mexican Treasuries. It is worth noting that despite the similar nature of changes in funding liquidity conditions, whether these are originated in the U.S. or Europe matters. In particular, the two sources of funding liquidity shocks have opposite sign on δ_t .

4.4 Weak Exogeneity and the Common Trends Representation

Weak Exogeneity

A further advantage provided by the model (4.1) is that it allows to test for “weakly exogenous variables”. As defined by Juselius (24) pp. 193, a variable $x_{j,t} \in \{X_t\}$ is said to be weakly exogenous if it affects other variables in the system X_t while it is not affected by them. This is, if the row j of matrix α contains only zeros. Since α is a $n \times r$ matrix, and $r = 1$ in this paper, then a weak exogeneity test is equivalent to testing for each of the 5 rows of α for being different from zero. To this end, a simple *t – test* statistic is not suitable since the comparison is made across 5 models. In particular, a Likelihood-Ratio statistic is required. Table 7 displays the weak exogeneity test for the variables in X_t .

r	5% Crit-val	Φ_t	$i_{us,t}$	$i_{mex,t}$	$LOIS_{us,t}$	$LOIS_{eur,t}$
1	3.841	9.674 [0.002]	12.984 [0.000]	0.093 [0.760]	0.038 [0.846]	1.889 [0.169]

Table 7: Weak Exogeneity Test. LR-Test statistic, $\chi^2(r)$, p – values in brackets. Null Hypothesis: The variable is weakly exogenous.

Results from the test allow to conclude that $i_{mex,t}$, $LOIS_{us,t}$ and $LOIS_{eur,t}$ are weakly exogenous. Note, however, that the p -value for $i_{us,t}$ implies that this variable adjusts towards the cointegrating relationship. There is, however, a plausible rationalisation for this seemingly counter-intuitive result. In particular, note that (4.9) contains $LOIS_{us,t}$ and note that funding liquidity shortages may be strongly related to an increase of demand of the safest asset in the market (i.e. the U.S. treasuries). The latter, in turn, should cause yields to change. A similar argument may apply when the change in funding liquidity conditions is born in Europe. This may explain why within the estimated system, $i_{us,t}$ is not weakly exogenous. The latter is confirmed by the restricted estimation and a test for the validity of the restrictions embedded in $\tilde{\alpha}$, contained in Table 8:

$$\tilde{\alpha} = \begin{pmatrix} -0.314, & -0.129, & 0, & 0, & 0 \end{pmatrix}' \quad (4.10)$$

(0.059) (0.024)

Test restricted α	χ^2	p – value
Test of Restrictions	5.2608	0.1537

Table 8: Test for the validity of the restricted estimate $\tilde{\alpha}$ versus the unrestricted model. The test assumes a $\chi^2(3)$ distribution. Null hypothesis: Restrictions are valid. β is left unrestricted.

The restricted estimate for α in equation (4.10), implies that Φ_t and $i_{us,t}$ are the only variables of the system that adjust in response to deviations from the cointegration relationship $\tilde{\beta}'X_{t-1}$. Moreover, the estimates imply that it takes slightly more than 3 and 7.5 weeks for deviations from the previous period equilibrium relationship to dissipate for Φ_t and $i_{us,t}$, respectively. This is, suppose there is a *stationary* one-time shock to δ_t^B , seemingly creating arbitrage opportunities. Their effect on Φ_t will take 3 or 7.5 weeks to dissipate entirely.

Common Trends Representation

The VECM also allows writing (4.1) in its common trends representation or Moving Average (MA) form, as proved in the Granger Representation Theorem. This allows both the estimation of the $n - r$ common trends in the system X_t and a very intuitive interpretation. The MA form is written as

$$X_t = C \sum_{s=1}^t \varepsilon_s + \sum_{j=0}^{\infty} C_j^* \varepsilon_{t-j} + X_0, \quad (4.11)$$

$$C = \beta_{\perp} \left[\alpha'_{\perp} \left(I_n - \sum_{s=1}^{p-1} \Gamma_s \right) \beta_{\perp} \right]^{-1} \alpha'_{\perp} = \tilde{\beta}_{\perp} \alpha'_{\perp} \quad (4.12)$$

where β_{\perp} is the orthogonal complement of β , defined as the $(n - r) \times r$ matrix satisfying $\beta'_{\perp} \beta = 0$ and similarly for α_{\perp} .⁸ The first component of equation (4.11) deserves special attention. In particular, note that the product $C \sum_{s=1}^t \varepsilon_s$ is $\tilde{\beta}_{\perp} \alpha'_{\perp} \sum_{s=1}^t \varepsilon_s$ using (4.12). It is possible to interpret $\tilde{\beta}_{\perp}$ as a measure of the importance (weight) that each of the common trends $\alpha'_{\perp} \sum_{s=1}^t \varepsilon_s$ has on X_t .

⁸Proof of (4.11) and (4.12) is contained in Lütkepohl (30).

The $n \times n$ matrix C defined in (4.12) is known as the "long-run impact matrix" and it is interpreted in two possible ways. Column-wise, the element c_{ij} , if statistically significant, means the cumulated effect $\sum_{s=1}^t \varepsilon_{j,s}$ has a relevant effect of c_{ij} on x_i . Row-wise, the element c_{ij} means that x_i may be permanently influenced by $\sum_{s=1}^t \varepsilon_{j,s}$ in some measure c_{ij} . The matrices C_j^* are the current and previous "one-time" effects of each element of ε_t on the system X_t and X_0 is its initial value.

For the present model, restricted estimates of the long-run impact matrix yield the following common trend representation.⁹ Recall that Σ is a full matrix as described in (4.3), Table 9 shows the corresponding correlations.

	ε_{Φ}	$\varepsilon_{i,us}$	$\varepsilon_{i,mex}$	$\varepsilon_{LOIS,us}$	$\varepsilon_{LOIS,eur}$
S.E.	0.355	0.166	0.186	0.0667	0.0572
ε_{Φ}	1.000				
$\varepsilon_{i,us}$	0.634	1.000			
$\varepsilon_{i,mex}$	-0.788	-0.217	1.000		
$\varepsilon_{LOIS,us}$	-0.877	-0.403	0.561	1.000	
$\varepsilon_{LOIS,eur}$	-0.746	-0.516	0.426	0.856	1.000

Table 9: Residual Standard Errors and Cross-Correlations.

The MA form is the representation of each variable as the sum of the history of previous shocks. Although correlated, these shocks are useful in explaining changes in each element of system X_t . Here, the restricted estimate \tilde{C} is used to construct the estimated MA expressions. Let $C_{[k,\cdot]j}^*$ be the k th row of matrix C_j^* . Omitting stationary and terms not different from zero statistically, the MA representation for Φ_t is given by

$$\Phi_t = -3.98 \sum_{j=1}^t \varepsilon_{i,mex,j} - 5.53 \sum_{j=1}^t \varepsilon_{LOIS,us,j} + \sum_{j=0}^{\infty} C_{[1,\cdot]j}^* \varepsilon_{t-j}. \quad (4.13)$$

In line with the weak exogeneity test, equation (4.13) shows all the stochastic trends have a non-negligible effect on Φ_t . $LOIS_{eur,t}$ is present through the correlation between $\varepsilon_{LOIS,us}$ and $\varepsilon_{LOIS,eur}$ which is relatively high, as shown in Table 9.

For $i_{us,t}$, equation (4.14) shows its non weakly exogenous status. In particular, it is affected by shocks on the forward premium

$$i_{us,t} = -0.364 \sum_{j=1}^t \varepsilon_{\Phi,j} + 0.89 \sum_{j=1}^t \varepsilon_{i,us,j} + \sum_{j=0}^{\infty} C_{[2,\cdot]j}^* \varepsilon_{t-j}. \quad (4.14)$$

The results of the estimation for $i_{mex,t}$ displayed in (4.15) suggests some de-coupling of the monetary policy for the period. A case for the presence of at least a second stochastic trend could be made easily in order to explain changes in the Mexican Treasury yield. This issue deserves a deeper study, since the correlations between $i_{mex,t}$ and the rest of the variables is not high enough to justify some second-order effect from a different variable in the system. These second-order effects may be present since the reduced form

⁹Full estimation output is available upon request.

errors ε_t are contemporaneously related

$$i_{mex,t} = \frac{2.69}{(0.601)} \sum_{j=1}^t \varepsilon_{i,mex,j} + \sum_{j=0}^{\infty} C_{[3,\cdot]j}^* \varepsilon_{t-j}. \quad (4.15)$$

Finally, the MA forms of $LOIS_{us,t}$ and $LOIS_{eur,t}$ in (4.16) and (4.17) show these variables behaving as two weakly exogenous elements, with a high correlation among them

$$LOIS_{us,t} = \frac{1.58}{(0.239)} \sum_{j=1}^t \varepsilon_{LOIS,us,j} + \sum_{j=0}^{\infty} C_{[4,\cdot]j}^* \varepsilon_{t-j}, \quad (4.16)$$

$$LOIS_{eur,t} = \frac{0.799}{(0.205)} \sum_{j=1}^t \varepsilon_{LOIS,us,j} + \frac{0.722}{(0.236)} \sum_{j=1}^t \varepsilon_{LOIS,eur,j} + \sum_{j=0}^{\infty} C_{[5,\cdot]j}^* \varepsilon_{t-j}. \quad (4.17)$$

So far, evidence suggests unambiguous support for the relevance of both $LOIS_{us,t}$ and $LOIS_{eur,t}$ in explaining the behaviour of the variables that define δ_t . The previous analysis comes a long way in separating effects from each shock into the behaviour of each element of X_t . There is a shortcoming, however, and that is related to the covariance matrix of the reduced form errors, Σ contained in Table 9, not being diagonal. Shocks between $LOIS_{us,t}$ and $LOIS_{eur,t}$ cannot be distinguished clearly. To accomplish this, a structural analysis is undertaken in the next section.

4.5 Structural Analysis

The Structural VECM(p) can be readily derived from expressions (4.1) or (4.11). In order to relate the reduced-form errors vector ε_t to the "structural" errors vector u_t , a Cholesky decomposition is used. This identification strategy relies on the assumption of some causal direction among the variables. Thus, it is assumed that $i_{us,t}$ is independent from the rest of the variables in X_t and affects $LOIS_{us,t}$. In turn, $LOIS_{eur,t}$ is affected by shocks to $LOIS_{us,t}$. Also, $i_{mex,t}$ is subject to changes in all non-domestic variables. Finally, Φ_t responds to shocks to all previous variables. The selected order obeys the economic logic behind the variables in the system. That is, structurally, the $i_{us,t}$ follows its own shocks; the banking system of the U.S. is larger than the European one, however interconnected; the European banking sector is independent of the structural shocks to $i_{mex,t}$, given that Mexico is a small open economy.¹⁰

Impulse Response Analysis

The model VECM(11) in (4.1) is used to estimate the IRFs, imposing the identified restriction set (4.8). The responses of both Φ_t and $i_{mex,t}$ are displayed in Figure 8. An increase in $i_{mex,t}$ yields a negative response of Φ_t , reflecting an MXN appreciation. Also, more stringent funding liquidity conditions given by increases in either $LOIS_{us,t}$ or $LOIS_{eur,t}$ yield a positive response of Φ_t . This is, an appreciation of the USD. The right column shows that increases in Φ_t , given by an appreciation of the MXN, yield a negative response of $i_{mex,t}$. Moreover, increases in either $i_{us,t}$ or $LOIS_{eur,t}$ imply an increase in $i_{mex,t}$. Note that this is the expected monetary policy reaction to more stringent funding conditions, in a broad sense.

The response of the rest of the variables to the shocks is shown in Figure 9. As suggested by the weak exogeneity analysis, $i_{us,t}$ responds positively to increases in Φ_t and to no other shock. Moreover, $LOIS_{us,t}$

¹⁰Results do not change if an ordering obeying the common-trends analysis from above is followed instead. Moreover, results are also similar if a liquidity measure for Mexico is included. See Appendix A

responds negatively to shocks to $i_{us,t}$ and $LOIS_{eur,t}$. Finally, $LOIS_{eur,t}$ responds negatively to increases in both Φ_t and $i_{us,t}$, while responding positively to shocks to $LOIS_{us,t}$.

Forecast Variance Decomposition Analysis

A forecast variance decomposition analysis is undertaken to complete the analysis of the average effect of shocks to the variables. Having established that funding liquidity shocks have a role in explaining deviations from the CIP, the aim is to quantify the relative importance of changes in $LOIS_{us,t}$ and $LOIS_{eur,t}$. To this end, a VAR model in levels based on the VECM(11) is estimated. As discussed by Kilian and Lütkepohl (2002), the analysis may well be carried out on the VECM(11) even though all variables contained in X_t are $I(1)$. This, however, may entail some loss of information since the model is expressed in differences.

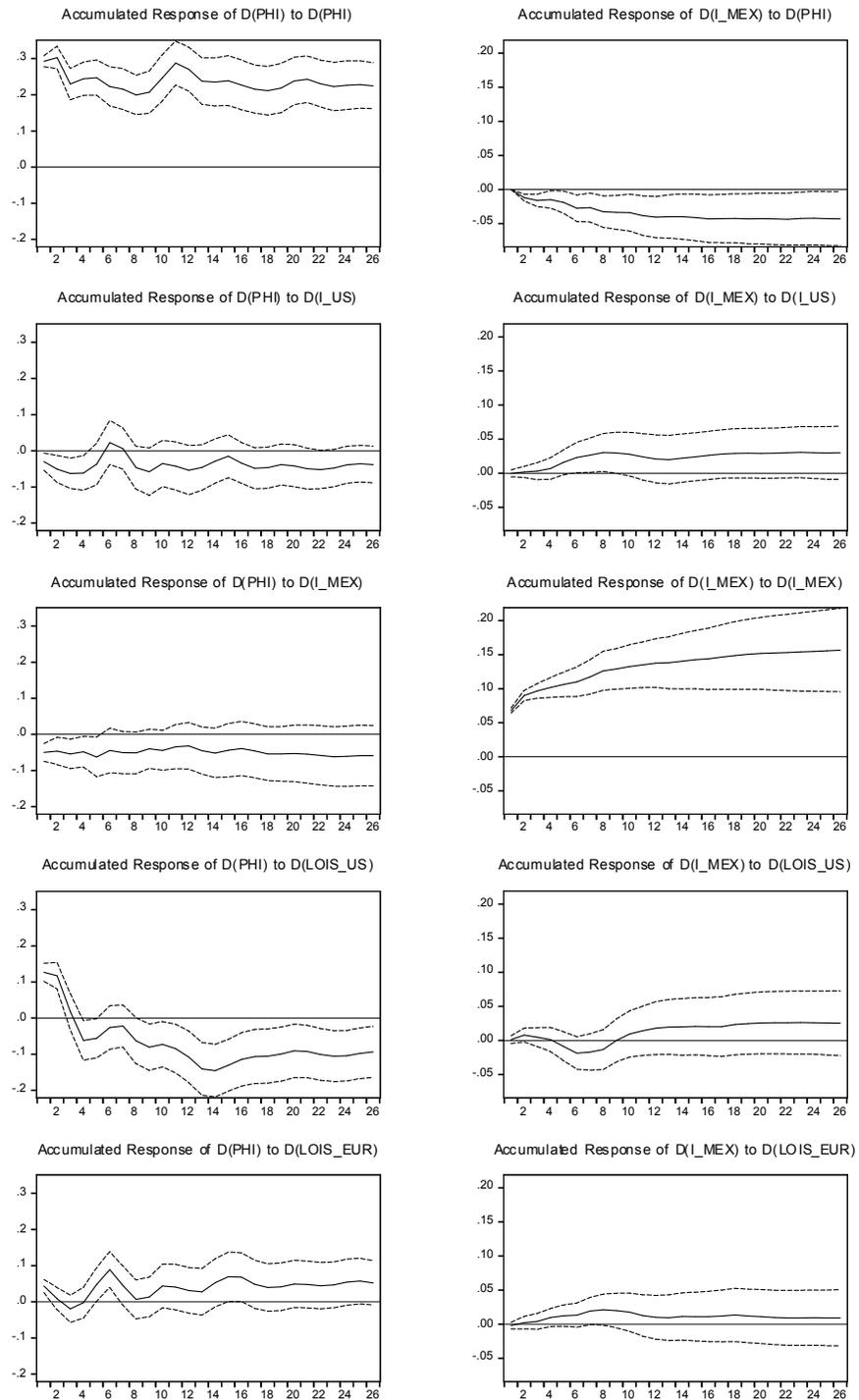


Figure 8: Impulse Response Functions and ± 2 Standard Errors computed by Monte Carlo Simulations. Estimation was carried out in EViews 10.

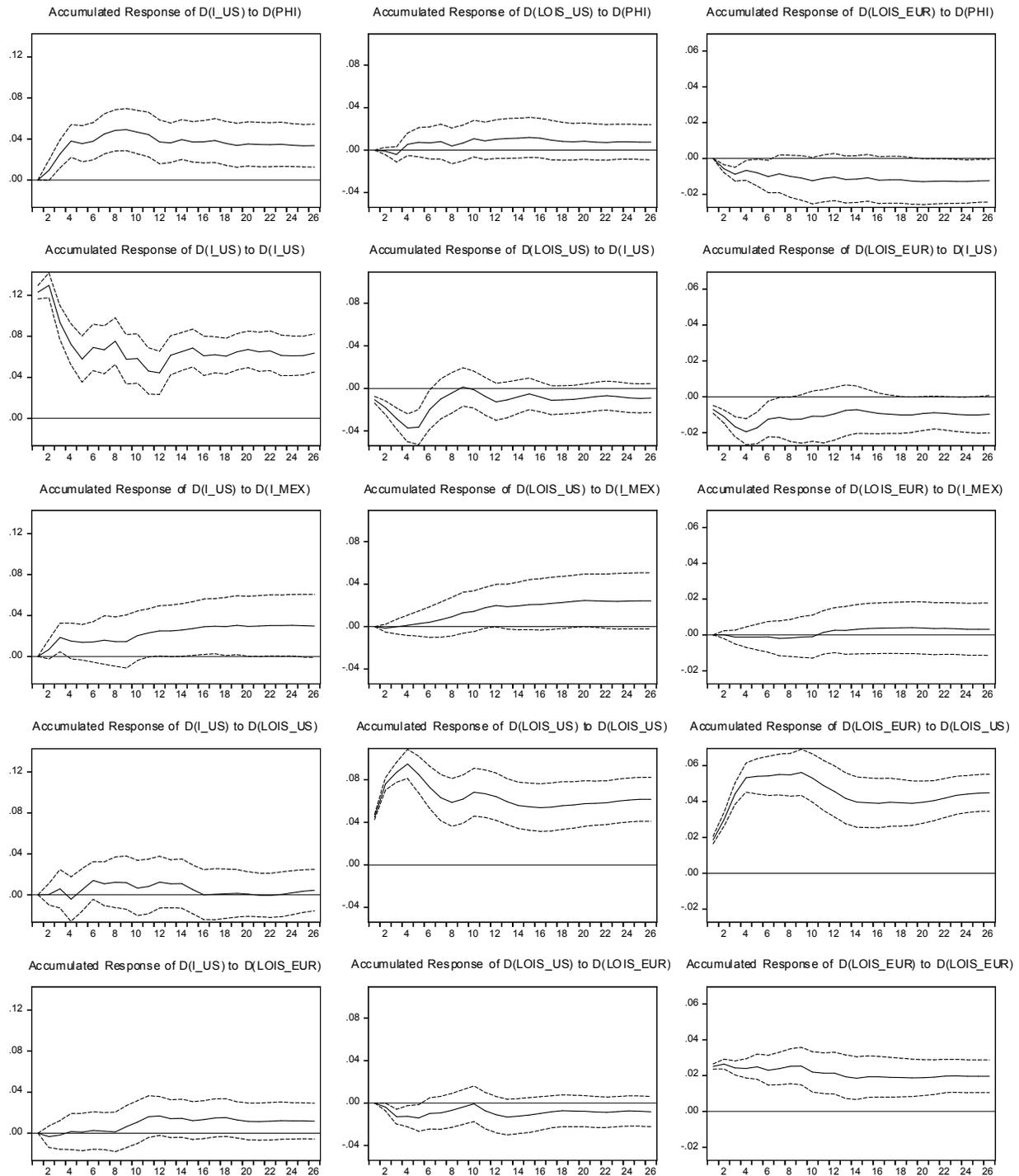


Figure 9: Impulse Response Functions and ± 2 Standard Errors computed by Monte Carlo Simulations. Estimation was carried out in EViews 10.

The estimated VAR model in levels is given by

$$\tilde{X}_t = c + \sum_{j=1}^{11} \Theta_j \tilde{X}_{t-j} + v_t, \quad (4.18)$$

where $\tilde{X}_t = (LOIS_{us,t}, LOIS_{eur,t}, \delta_t)'$, δ_t is defined in (2.4), Θ_j are $n \times n$ parameter matrices, c is a $n \times 1$ vector of constant terms, and v_t is the error term. The estimation of model (4.18) and inference obtained from it should pose no obstacle, as long as the system is stable. That is, all the roots of the system are within the unit circle. Indeed, the largest implied root from the VAR in levels is 0.974, well below 1.¹¹

Results from the forecast error variance decomposition for δ_t are summarised in Table 10. The 52 week (one year) horizon serves to find the convergence level of each variance decomposition percentage. The Cholesky ordering is that given in \tilde{X}_t . Results suggest that the relative importance of δ_t to explain its own forecast errors decreases within the horizon from 75.5% to 44.9%. The relative importance of $LOIS_{us,t}$ increases from 21.3% to 40.5% while that of $LOIS_{eur,t}$ increases from 3.3% to 14.6%.

Horizon	S.E.	δ_t	$LOIS_{us,t}$	$LOIS_{eur,t}$
1	0.364533	75.46490 (2.92715)	21.27821 (2.67605)	3.256888 (1.12950)
5	0.593349	76.85711 (2.82405)	20.53360 (2.58550)	2.609291 (1.37738)
10	0.674929	64.28828 (5.02050)	29.05565 (4.38689)	6.656070 (3.39428)
15	0.775626	53.96670 (5.53852)	38.88517 (5.29684)	7.148130 (4.13493)
25	0.824555	48.13689 (6.10063)	41.03941 (6.17755)	10.82371 (6.31722)
52	0.862945	44.93074 (6.25840)	40.46760 (6.83538)	14.60166 (8.61251)

Table 10: Forecast Error Variance Decomposition. S.E. stands for the forecast error at each horizon. The labelled columns labelled with the elements of \tilde{X}_t are the percentage of the forecast variance attributed to each variable. Standard Errors computed by 10^4 Monte Carlo Simulations are shown within brackets. Estimation was carried out in EViews 10.

Interestingly, the relative importance of $LOIS_{us,t}$ is statistically the same as that of δ_t in explaining the forecast error variance of δ_t after 15 weeks. The latter may be rationalised by, first, noting that the U.S. banking sector is the largest in the world and, second, the analysis of the CIP involves the USD, the currency in which $LOIS_{us,t}$ is determined.

A further result worth of attention is that, however small when compared with $LOIS_{us,t}$, changes in $LOIS_{eur,t}$ are important in explaining deviations from the CIP. Indeed, in a horizon of 52 weeks, 14.6% of the forecasts error variance may be attributed to this funding liquidity measure. Thus evidence supports that, notwithstanding the apparent weak relationship that may exist between the USD-MXN market and European financial markets, there is a non-negligible link.

Historical Decomposition Analysis

Up to this stage, the structural analysis suggests that *on average*, changes in funding liquidity conditions are relevant to determine deviations from CIP. The analysed period is characterised by a sequence of financial

¹¹ADF and PP unit root tests for de-mean δ_t reject the null hypothesis of the presence of a unit root. Moreover, residuals estimated in model (4.18) are free from autocorrelation according to the LM test.

events, not all taking place simultaneously. Hence, an Historical Decomposition Analysis may shed light on the *relative importance of each measure of funding liquidity at each date*. Figure 10 displays the results. In particular, the solid line is the relative importance of each variable on the behaviour of δ_t while the shade corresponds to de-mean δ_t itself.

Some conclusions emerge. Prior to the onset of the financial crisis in summer 2007, most of the dynamics of δ_t were explained by itself (i.e. changes in Φ_t , $i_{us,t}$, and $i_{mex,t}$). After August 2007, the funding liquidity measures are able to explain several episodes of stark deviations from CIP. In particular $LOIS_{us,t}$ is able to account for the observed large deviations from early 2008 to mid 2009, including of course those registered in the fourth quarter of 2008 when Lehman failed, AIG was bailed-out, and TARP was approved by the U.S.' Congress. It then explains a large share of δ_t from mid 2010 to mid 2011. Finally, beginning in 2012, possibly related to new regulation concerning funding of commercial banks, the funding liquidity measure re-gained importance.

The role of $LOIS_{eur,t}$ is reflected mainly in four episodes. First, from mid 2007 to mid 2008, a period where, as documented by Ivashina et al. (2012), European banks experienced marked USD liquidity constraints. Second, from mid 2010 to early 2011, when several southern European economies were on the brink of entering into a sovereign debt crisis. Third, throughout 2011 when the sovereign debt crisis was at its peak, particularly in Greece. The role of $LOIS_{eur,t}$ started to diminish after the European Central Bank (ECB) stated that "Within our mandate, the ECB is ready to do whatever it takes to preserve the euro." in mid 2012.¹² Finally, towards the end of the sample in 2018, this funding liquidity measure is able to account for deviations from the CIP.

5 Concluding Remarks

In this paper, econometric evidence is presented to support the hypothesis that funding liquidity constraints have a significant role in explaining deviations from CIP relating the USD-MXN. Results show that the LOIS spread in USD is positively related to deviations from CIP whereas the LOIS spread in EUR is negatively related, underscoring that the source of the shock matters. Moreover, it is found that both funding liquidity measures are able to explain the forecast error variance of deviations from CIP. Interestingly, LOIS in USD accounts for a similar share of said variance than deviations from CIP after 15 periods. An historical decomposition analysis suggests that both LOIS in USD and EUR are able to account for both positive and negative deviations from CIP. The analysis suggests that the the apparent arbitrage opportunities do not exist once funding liquidity conditions are included. This is, there should be no deviations from the CIP once funding liquidity conditions are considered. From the policy perspective, this paper underlies the relevance of funding liquidity measures when assessing whether the USD-MXN foreign exchange is functioning efficiently.

There are some caveats in the analysis to consider. First, the borderline conclusion of the funding liquidity measures being $I(1)$ should be present before making further conclusions. Second, many market participants are not able to fund their liquidity at LIBOR or OIS rates, but at higher ones. Future work should aim to find a funding liquidity measure for non-prime borrowers. Further research should focus on testing different models that account for the borderline stationarity of the $LOIS_{.,t}$. In the same vein, different measures associated with market liquidity, such as Credit Default Swaps or Volatility indices may be informative. Finally, the financial series present considerable ARCH-like behaviour, this may be a source of information to explore.

¹²Speech by Mario Draghi, President of the European Central Bank at the Global Investment Conference, London on July 26, 2012.

Acknowledgements: Valuable comments and suggestions from a anonymous referees and the participants of the seminar series at Directorate General of Economic Research at Banco de México are acknowledged. This work was greatly benefited from comments and discussion with Nicolás Amoroso, Marcus Chambers, Miguel Díaz, Raúl Ibarra, Carlos Lever, Antonio Noriega, Patrice Robitaille, Salvador Rodríguez, Joao Santos-Silva, and Lucio Sarno. Andrea Miranda and Ezequiel Piedras provided excellent research assistance.

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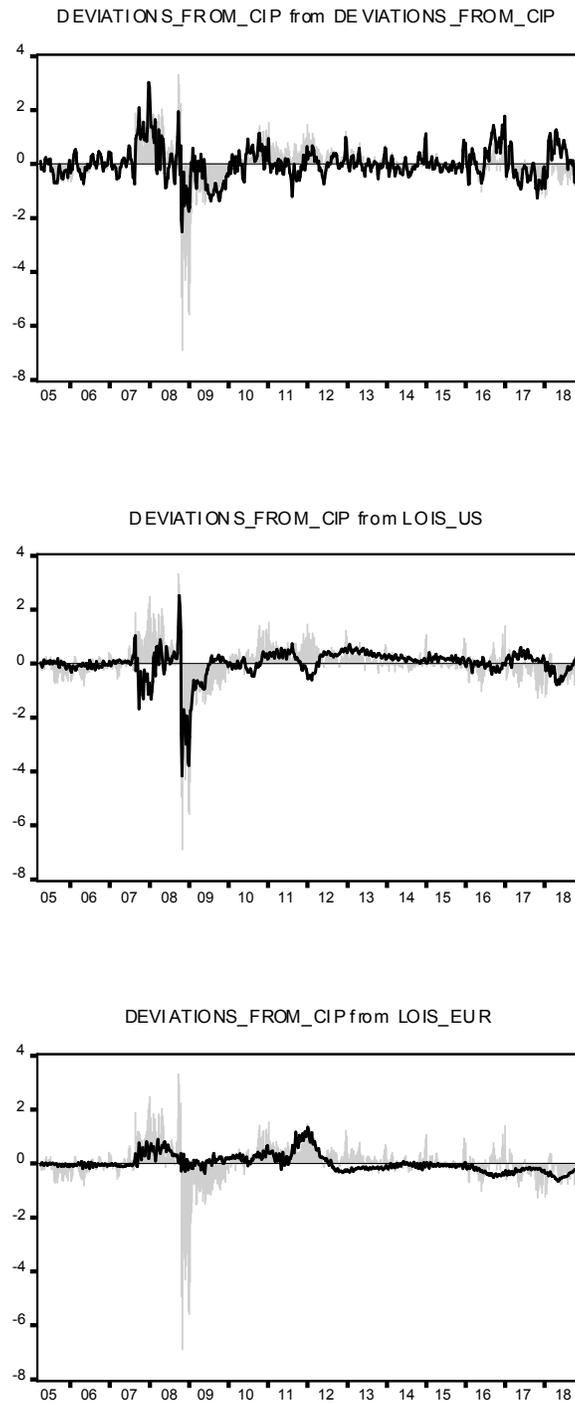


Figure 10: Historical Decomposition of δ_t computed by Monte Carlo Simulations. Shaded area corresponds to δ_t , solid line corresponds to the relative importance of each variable explaining δ_t . Estimation was carried out in EViews 10.

Appendices

A LOIS for Mexico

Horizon	S.E.	δ_t	$LOIS_{US}$	$LOIS_{EUR}$	$LOIS_{MEX}$
1	0.362699	75.91505 (2.72313)	20.53417 (2.68532)	3.220211 (1.15654)	0.330562 (0.34808)
5	0.594195	76.94096 (3.13446)	20.45739 (2.93466)	2.390065 (1.14156)	0.211591 (0.66572)
10	0.675832	64.08866 (5.06013)	29.44557 (4.66296)	6.083867 (3.22748)	0.381896 (1.12275)
15	0.774654	53.74312 (5.45571)	39.06096 (5.58296)	6.555362 (3.46195)	0.640556 (1.07696)
25	0.822831	47.89034 (5.87807)	40.83643 (6.08406)	10.06468 (4.57344)	1.208542 (1.65333)
52	0.864981	43.98718 (6.58314)	40.87865 (6.63079)	13.89360 (6.35129)	1.240569 (2.08445)
		(6.64223)	(6.66857)	(6.47433)	(2.13731)

Table 11: Forecast Error Variance Decomposition. S.E. stands for the forecast error at each horizon. The labelled columns labelled with the elements of \tilde{X}_t are the percentage of the forecast variance attributed to each variable. Standard Errors computed by 10^4 Monte Carlo Simulations are shown within brackets. Estimation was carried out in EViews 10.

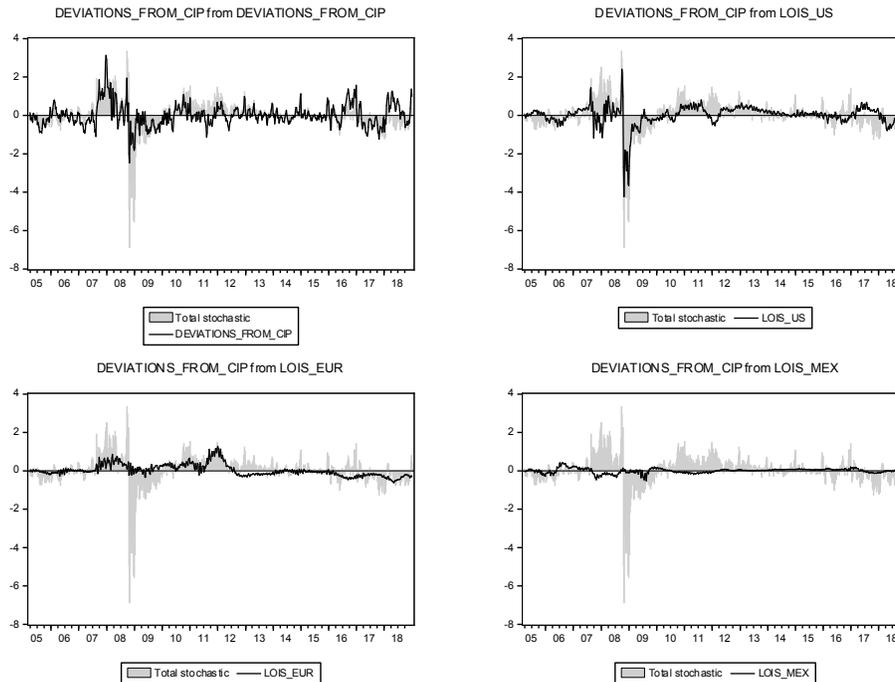


Figure 11: Historical Decomposition of δ_t computed by Monte Carlo Simulations. Shaded area corresponds to δ_t , solid line corresponds to the relative importance of each variable explaining δ_t . Estimation was carried out in EViews 10.

B Data

This appendix contains the Bloomberg ticker of each variable:

1. Main Variables:

- Spot Exchange Rate USD per 1 MXN: $(S_t)^{-1}$ =USDMXN Curncy.
- Forward Exchange Rate USD per 1 MXN at maturity 1 month: $(F_t)^{-1}$ =MXN1M Curncy.
- Interest Rate of U.S. Treasury Debt 1-month constant maturity: $i_{us,t}$ =USGG1M Index.
- Interest Rate of Mexican Debt in MXN of 1-month maturity: $i_{mex,t}$ =MPTBA Curncy.

2. LIBOR-OIS

- LIBOR-OIS Spread for the U.S.: $LOIS_{us,t}$ =US0003M Index - USSOC Curncy.
- LIBOR-OIS Spread for Europe: $LOIS_{eur,t}$ =EUR003M Index - EUSWEC Curncy.
- LIBOR-OIS Spread for Mexico: $LOIS_{mex,t}$ =MXIB91DT Index - MPSW28T Curncy.