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Identifying Dornbusch's Exchange Rate Overshooting with Structural VECs: Evidence from Mexico*

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Abstract

In this paper we use data from Mexico to identify Dornbusch's (1976) exchange rate overshooting hypothesis. We specify and estimate a structural cointegrated VAR that considers explicitly the presence of a set of long-run theoretical relations on macroeconomic variables (a purchasing power parity, an uncovered interest parity, a money demand, and a relation between domestic and U.S. output levels). We then impose a recursiveness assumption to identify the response of domestic variables to a monetary policy shock. The long-run restrictions embedded in the model are themselves identified, estimated, and tested using an ARDL methodology that is robust to the degree of persistence of the time series and, in particular, to whether they are trend- or first-difference stationary. With this approach, we are able to find that the response of the exchange rate to monetary policy shocks is consistent with Dornbusch's model.

Keywords: Cointegration; vector error correction models; exchange rate overshooting; money demand; monetary policy shock; purchasing power parity; uncovered interest rate parity.

J.E.L. classification number. C32, C51, E10, E17.

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1 Introduction

Although Dornbusch's (1976) overshooting hypothesis has become a central component of international macroeconomic theory and of monetary policy modeling, the empirical evidence from small open economies has not generally supported it. In fact, in many cases the evidence has found puzzling results suggesting that the exchange rate depreciates after a positive shock to domestic interest rates (see e.g. Sims, 1992; Racette and Raynauld, 1992; Grilli and Roubini, 1995; Mojon and Peersman, 2003). This exchange rate puzzle not only contradicts the overshooting hypothesis, but is also inconsistent with the theoretical response that the value of the local currency should exhibit after a monetary policy shock. In other cases, a long and persistent hump-shaped appreciation of the currency after a contractionary monetary policy shock has been found, the so-called "delayed overshooting" (see Eichenbaum and Evans, 1995; Grilli and Roubini, 1995; Lindé, 2003). This is again inconsistent with Dornbusch's model and, in particular, with the expected depreciation that would make uncovered interest parity hold after the initial appreciation of the currency.

Other papers have argued that these results reflect an inappropriate identification strategy and, in particular, may be responding to bi-directional causality between exchange rates and interest rates (see Cushman and Zha, 1997; Kim and Roubini, 2000; Bjørnland, 2009). Indeed, if monetary policy responds to exchange rate shocks by increasing interest rates to avoid the inflationary consequences of the shock, this may lead to a positive correlation between interest rates and exchange rates that may make it difficult to identify the response of exchange rates to exogenous monetary policy shocks. In this context, the combined evidence from these papers shows for the cases of non-U.S. G-7 countries, Australia, New Zealand and Sweden that an appropriate identification strategy that both explicitly takes into account the features of a small open economy and its interactions with the rest of the world, and that identifies the monetary policy response to exchange rate shocks, is helpful to identify theoretically consistent exchange rate responses.

In this paper, we present evidence from Mexico supporting this view. The identification strategy we follow is based on: i) imposing a set of long-run theoretical restrictions on a structural cointegrated VAR model; and, ii) using a recursiveness assumption to identify the response of domestic variables to a monetary policy shock. While Cushman and Zha (1997) and Bjørnland (2009) also achieve identification through the estimation of structural VARs, our strategy differs from theirs.[1] In particular, we impose a much

[1] Cushman and Zha (1997) estimate a structural VAR that explicitly identifies a monetary policy function. However, they don't restrict the model to satisfy long-run theoretical relations. Bjørnland

richer long-run structure to the model, which implies a large set of over-identifying restrictions. Thus, in contrast with the approach followed previously by this literature, we are able to test the validity of our identification strategy.

Formally, our strategy starts with the specification and estimation of a small quarterly macroeconometric model for the Mexican economy. We follow the Garratt et al. (2006) approach based on the estimation of a structural cointegrated VAR that considers explicitly the presence of a set of long-run theoretical restrictions on macroeconomic variables, but that otherwise leaves the short-run dynamics of the system unrestricted.[2] The long-run relations are themselves identified, estimated, and tested using an autoregressive distributed lag (ARDL) methodology that is robust to the degree of persistence of the regressors (Pesaran, 1999; Pesaran et al., 2001). This may be especially relevant in countries such as Mexico, where diverse structural breaks may affect the degree of persistence of macroeconomic series, making it difficult to determine unambiguously the stochastic properties of the variables as trend or first-difference stationary. The long-run relations consider explicitly the fact that Mexico is a small open economy and, therefore, account for the relation that the core domestic macroeconomic aggregates would be expected to have with foreign variables. In particular, the model contains four long-run relations: (i) a purchasing power parity; (ii) an uncovered interest parity; (iii) a money demand; and (iv) a relation between domestic and U.S. output levels. The dynamic properties of the model are illustrated with the use of Generalized Impulse Response Functions (Koop et al., 1996; Pesaran and Shin, 1998). We then use the estimated model to recursively identify a monetary policy shock and to analyze the response of macroeconomic variables to this shock.

This paper differs from previous attempts to identify monetary policy shocks in the Mexican economy. In contrast to the regime-switching VARs with recursive identifying restrictions estimated by Del Negro and Obiols-Homs (2001) and Gaytán and García-González (2006), we find significant effects on output and prices. Furthermore, we also find that, consistent with Dornbusch's (1976) overshooting hypothesis, a contractionary monetary policy shock induces a quick and strong appreciation of the exchange rate, followed by a gradual depreciation. In particular, we show that, once we follow our identification strategy, both the so-called price and exchange rate puzzles disappear. Moreover, our results do not exhibit the "delayed overshooting" anomaly found in pre-

(2009), in contrast, assumes a long-run neutrality restriction on the exchange rate, which allows her to impose a contemporaneous interaction between the interest rate and the exchange rate without sacrificing identification. No further long-run restrictions are imposed.

[2] See Garratt et al. (2003) for an application of this methodology to the UK economy and Assenmacher-Weschea and Pesaran (2009) for an application to the Swiss economy.

vious attempts to identify the exchange rate overshooting mechanism. These results are consistent with economic theory and are broadly similar to those found by Carrillo and Elizondo (2015) who estimate a VAR model with variables expressed as gaps and both recursive and sign identifying restrictions. Our approach however, differs from theirs by providing a unified account of trends in real, financial and foreign variables. Since the variance of many macroeconomic series seems to be dominated by low frequency components, directly modelling the trends and their inter-relations may be particularly helpful for the study of the effects of monetary policy in the long-run. An additional distinguishing feature from these papers is that, by considering over-identifying restrictions, this approach imposes a foundation based on economic theory that is testable and that increases the degrees of freedom for estimation. These are useful features in an environment with small samples and possible regime changes.

The paper proceeds as follows. In Section 2, we present the econometric formulation of the structural cointegrated model used in this paper. The data and the analysis of the individual long-run relations is found in Section 3. Section 4 contains the estimation and testing of the model, while section 5 presents the impulse-response analysis, including the response of the endogenous variables to a monetary policy shock. Some concluding remarks are contained in Section 6.

2 Econometric Formulation of the Model

The model we consider is a structural cointegrated vector autoregressive (VAR) model that relates the core macroeconomic variables of the Mexican economy to current and lagged values of key foreign variables. Our modelling strategy begins with the explicit statement of the long-run relation between the variables of the model obtained from macroeconomic theory. The deviations from these long-run relations (i.e. the error correcting terms), are then embedded within an otherwise unrestricted VAR model in the variables of interest to obtain a cointegrating VAR model which has the structural long-run relation as its steady-state solution. This allows us to test for the presence of the cointegrating relations and the validity of the over-identifying restrictions implied by the long-run economic theory.

A very general structural model for the determination of an $m \times 1$ vector of variables of interest, $\mathbf{z}_t$, is given by the Structural Vector Error Correction Model (SVECM)

$$\mathbf{A}\Delta\mathbf{z}_t = \widetilde{\mathbf{a}} + \widetilde{\mathbf{b}}t - \widetilde{\mathbf{\Pi}}\mathbf{z}_{t-1} + \sum_{i=1}^{p-1} \widetilde{\mathbf{\Gamma}}_i \Delta\mathbf{z}_{t-i} + \boldsymbol{\varepsilon}_t. \tag{2.1}$$

Here, $\mathbf{A}$, $\widetilde{\mathbf{\Pi}}$, and $\widetilde{\mathbf{\Gamma}}_i$ are $m \times m$ matrices and $\widetilde{\mathbf{a}}$ and $\widetilde{\mathbf{b}}$ are $m \times 1$ vectors, all containing structural parameters. The $m \times 1$ vector of errors in (2.1), $\boldsymbol{\varepsilon}_t$, is assumed to have zero mean and a diagonal positive definite covariance matrix $\boldsymbol{\Omega}$ (i.e. $\boldsymbol{\varepsilon}_t$ contains the structural shocks). If there are r cointegrating vectors, $0 < r < m$, then $\widetilde{\boldsymbol{\Pi}}$ will be of rank r. In this case, we may decompose it as $\widetilde{\boldsymbol{\Pi}} = \widetilde{\boldsymbol{\alpha}}\boldsymbol{\beta}'$ where $\widetilde{\boldsymbol{\alpha}}$ is an $m \times r$ matrix of adjustment coefficients and $\boldsymbol{\beta}$ is an $m \times r$ matrix of long-run coefficients.

Within this framework, $\boldsymbol{\beta}'\mathbf{z}_{t-1}$ is interpreted as the error correction term. In particular, as explained in Davidson et al. (1978), $\boldsymbol{\beta}'\mathbf{z}_t = \mathbf{0}$ can be interpreted as the long-run equilibrium of the dynamic system so that, whenever $\boldsymbol{\beta}'\mathbf{z}_{t-1}$ differs from zero, there are deviations from the long-run equilibrium that tend to be corrected through changes in $\mathbf{z}_t$. In some cases, these relations can be supplemented with deterministic intercepts and trends. Thus, according to Garratt et al. (2003), the long-run relations can be expressed as

$$\boldsymbol{\xi}_t = \boldsymbol{\beta}'\mathbf{z}_t - \mathbf{b}_0 - \mathbf{b}_1 t, \tag{2.2}$$

where $\boldsymbol{\xi}_t$ is therefore interpreted as the deviation from the long-run equilibrium defined by the condition $\boldsymbol{\beta}'\mathbf{z}_t - \mathbf{b}_0 - \mathbf{b}_1 t = \mathbf{0}$.

Hence, the long-run relations are embedded in the matrix $\boldsymbol{\beta}$. However, as is well known, $\boldsymbol{\beta}$ can not be identified from $\widetilde{\boldsymbol{\alpha}}$ without extra restrictions. In order to identify it, at least r^2 restrictions are needed. Normalization readily yields r restrictions, so that $r^2 - r$ additional restrictions are required.

The most common approach to obtain identification is probably the one proposed by Johansen (1991), which is based on statistical procedures. However, an interesting alternative to obtain the missing $r^2 - r$ restrictions relies on economic theory, as proposed by Garratt et al. (2003). Since the restrictions are imposed on the cointegrating vectors, the relevant theory is that of the long-run. In this context, the modelling strategy is to embed $\boldsymbol{\xi}_t$ in an otherwise unrestricted VAR(p). This strategy allows us to avoid restricting the short-run dynamics, by assuming that changes in $\mathbf{z}_t$ can be well approximated by a linear function of a finite number of past changes in $\mathbf{z}_t$ (Sims, 1980). In this regard, we consider the VECM(p-1)

$$\Delta\mathbf{z}_t = \mathbf{a}_0 - \boldsymbol{\alpha}\boldsymbol{\xi}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta\mathbf{z}_{t-i} + \mathbf{u}_t,$$

that, given equation (2.2), can be written as

$$\Delta\mathbf{z}_t = \mathbf{a} + \mathbf{b}t - \boldsymbol{\Pi}\mathbf{z}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta\mathbf{z}_{t-i} + \mathbf{u}_t, \tag{2.3}$$

which is a reduced form VECM where $\mathbf{a} = \mathbf{A}^{-1}\widetilde{\mathbf{a}}$, $\mathbf{b} = \mathbf{A}^{-1}\widetilde{\mathbf{b}}$, $\boldsymbol{\Pi} = \mathbf{A}^{-1}\widetilde{\boldsymbol{\Pi}} = \mathbf{A}^{-1}\widetilde{\boldsymbol{\alpha}}\boldsymbol{\beta}' = \boldsymbol{\alpha}\boldsymbol{\beta}'$, $\boldsymbol{\Gamma}_i = \mathbf{A}^{-1}\widetilde{\boldsymbol{\Gamma}}_i$, $\mathbf{u}_t = \mathbf{A}^{-1}\boldsymbol{\varepsilon}_t$, and $\mathbf{u}_t$ has covariance matrix $\boldsymbol{\Sigma}$ with $\boldsymbol{\Omega} = \mathbf{A}\boldsymbol{\Sigma}\mathbf{A}'$. The estimation of (2.3) can be carried out using the long-run structural modelling approach proposed by Pesaran et al. (2000) and Pesaran and Shin (2002).

Within this framework, it is important to distinguish between endogenous and weakly exogenous variables, since this could lead to a more efficient estimation approach. We treat foreign variables as weakly exogenous, this is, they affect the domestic variables contemporaneously but, while they could be affected by lagged changes of domestic and foreign variables, they are not affected by disequilibria in the domestic economy. This implies that none of the error correction terms enter the equations for the change in the foreign variables.

3 Long-Run Level Relations

Our choice of variables is determined by the purpose of the model, namely modelling the monetary transmission mechanism. In this context, the model should include those relations from economic theory that may be associated to the mechanisms through which the economy responds to monetary policy shocks. It should also include the variables that are expected to influence domestic output and inflation in a small open economy such as Mexico. We consider: (i) a purchasing power parity, which links the domestic price level, the exchange rate and the foreign (U.S.) price level; (ii) an uncovered interest rate parity, which relates the domestic nominal interest rate to the foreign (U.S.) nominal interest rate and to the depreciation rate; (iii) a condition for long-run solvency requirements, the money demand, which relates the real money stock to the real output and the interest rate; and, (iv) a condition for long-run output determination, based on a relation between domestic and foreign real output. In addition, once we estimate the full model, we consider the price of oil, the foreign (U.S.) producer price index and the TED spread as exogenous variables in the system, given their importance as determinants of Mexico's terms of trade, prices for imported intermediate goods and global financial liquidity shocks, respectively.[3] These variables' only role in the VECM will be to affect the short-run dynamics.

[3]The TED spread is a commonly used measure of the funding conditions and is defined as the spread between 3-Month LIBOR on U.S. dollars and the 3-Month Treasury Bill.

3.1 Data

We use quarterly, seasonally adjusted data for the period 1990q1-2015q4.[4] Table 1 describes the variables. Domestic variables are real gross domestic product (GDP), y_t; Mexican consumer price index, p_t; 91-day CETES interest rate, r_t; money aggregate M1 in real terms, m_t^d; and the nominal exchange rate expressed as Mexican pesos per U.S. dollar, e_t. Foreign variables are U.S. real GDP, y_t^*; U.S. consumer price index, p_t^*; three-month t-bill interest rate, r_t^*; WTI oil price, p_t^o; U.S. producer price index pp_t^*; and the TED spread, $\widetilde{r}_t^*$. All variables, except the interest rates and the spread, are transformed to their natural logarithms.[5]

To see the extent to which the long-run relations under investigation have held historically, the data is presented in Figures 1 to 4. Figure 1 shows the variables in the purchasing power parity relation and suggests that it may hold in the long-run. The large devaluation associated with the crisis of the mid-90s is evident, as is the response of prices to this devaluation. The recent financial crisis resulted in the observed sharp depreciation in the last two quarters of 2008, the latter matched with a price differential response. Two further events are noticed, the financial stress steamed from the Greek debt crisis towards the end of 2011 and, the sharp depreciation of the Mexican currency observed from the last quarter of 2014 to the end of the sample, which was partially a response to the observed decrease in oil prices.

Figure 2 shows the variables in the uncovered interest parity relation, the domestic and the foreign (U.S.) interest rates in panel (a), the depreciation of the exchange rate, in panel (b), and all the relevant variables simultaneously in panel (c). There appears to be a positive relation between the interest rates, although the foreign rate is smoother. When both rates grow apart, the gap seems to be filled, at least in part, by depreciations of the Mexican currency. This is evident during the crisis of the mid of the nineties, the financial crisis starting in 2008q3, the Greek debt stress episode in 2011q4 and at the end of the sample throughout 2015.

Figure 3 shows the variables in the money demand equation. The levels of real M1 and real GDP appear in panel (a).[6] After the crises of both the mid-90s and the end

[4] To guarantee that all regressions start at 1990q1 and are therefore comparable, irrespective of the lag order chosen, data before 1990q1 was used.

[5] Interest rates are expressed as $0.25\ln(1 + R/100)$, where R is the interest rate in percent per annum, to make units compatible with other rates of change used (e.g., the depreciation rate) which are computed as the first (log) difference of the quarterly levels. The Mexican Consumer Price Index and M1 are seasonally adjusted using the TRAMO-SEATS methodology (Gomez and Maravall, 1996).

[6] M1 at constant prices increased 84% between 1990 and 1991. This level shift, which became apparent in the last quarter of 1991, is explained by the increase in checking accounts (151.7%) that occurred as a result of a change in regulation. The new regulation transferred funds in *fideicomisos*

of the first decade of the 2000's, both series have similar trend properties, hence, an income elasticity of unity seems plausible. Panel (b) illustrates the relation between (inverse-) velocity and the interest rate. The Figure is supportive of a negative interest rate semi-elasticity.

Figure 4 shows the domestic and foreign real GDP. After the crisis in the mid-90s and up to the first half of 2008, these series display similar trends. This may reflect the production-side links formed between the manufacturing sectors of both economies, especially after the North American Free Trade Agreement (NAFTA) was enacted (see Graham and Wada, 2000). After the 2008 crisis, however, the U.S. GDP growth rate seems to have been somewhat lower than that of Mexico -which experienced a sharper decrease in economic activity in 2009.

The exogenous variables that we use in the estimation of the full system are also displayed in Figures 5 to 7. Figure 5 shows oil prices with three steep declines, one related to the 1998 Asian debt crisis, a second decline during the 2009 financial crisis and the last decline beginning the second half of 2014. The behavior of U.S.Producer Price Index is contained in Figure 6 while, that of the behavior of the TED Spread is displayed in Figure 7.

3.2 Testing for the Existence of Level Relation

Before conducting the full system estimation, we undertake a single-equation analysis of each of the long-run relations mentioned above. We first test for the existence of a long-run relation between the levels of the series in each equation using the bounds testing approach of Pesaran et al. (2001). Then, given that the data does not seem to reject the existence of a levels relation, we use the ARDL modeling approach of Pesaran (1999) to estimate its coefficients. The estimates from the ARDL long-run relations will indicate which parameter restrictions are likely to be valid. Importantly, the ARDL approach is robust to the unit root properties of the data, and hence knowledge of the integration order of the variables is not necessary. This is particularly useful for the case of nominal variables in Mexico, given that a change in their persistence may have occurred in 2001 (Chiquiar et al., 2010).

The individual models that we estimate correspond to the four long-run relations that we expect to find: purchasing power parity (PPP), uncovered interest rate parity (UIP), money demand (MD) and output determination (OUT). We estimate each

abiertos de inversion de valores and in master accounts to (interest paying) checking accounts. To account for this shift, we include in the money equation below a dummy variable that takes the value of one from 1991q4 on.

ARDL in error correction form. To determine the appropriate lag length and whether a deterministic linear trend is required, each model was estimated by Ordinary Least Squares with and without a linear time trend, using one to eight lags. We selected the best specification using the Schwarz information criterion (BIC). Pesaran et al. (2001) provide bounds for the critical values for a *t-test* of the significance of the coefficient associated with the error correction term (speed of adjustment) and for a joint test of the exclusion of the lagged levels. These bounds provide a lower and an upper critical value for the null of no level relation (i.e., no cointegration if the variables contain an unit root). When the test statistic lies outside the bounds (in absolute value), the null is rejected in favour of the alternative hypothesis of the existence of a level relation among the variables included in the equation, irrespectively of the persistence properties of the variables involved.[7]

The tests for the existence of long-run level relations are summarized in Table 2. The second column shows the coefficient of the error correction term, and the next shows the *t-statistics* with their respective lower and upper bound critical values at the 10% level. The following three columns show the *F-statistic* for the exclusion of the levels of the variables and the corresponding lower and upper 10% critical values. Finally, we present the adjusted R^2 and the final specification of each ARDL model.[8] In the four models we get negative and significant error correction coefficients, as can be seen by the fact that the *t-statistic* exceeds the bounds (in absolute value).[9] In addition, the *F-statistic* also exceeds the upper bound in the four cases. Thus, there is evidence to reject the null hypothesis of no level relations in all the long-run relations. Hence, we find evidence of the existence of four stable long-run relations. The long-run relations estimated with the ARDL models are shown below, with delta method standard errors in parenthesis. The order is: PPP; UIP; MD; and OUT

$$
\begin{aligned}
p_t &= \underset{(0.818)}{0.019} + \underset{(0.196)}{0.491} p_t^* + \underset{(0.055)}{1.015} e_t, \\
\Delta e_t &= \underset{(0.007)}{0.001} + \underset{(0.274)}{1.154} r_t - \underset{(1.159)}{2.329} r_t^*, \\
m_t^d &= \underset{(3.553)}{-2.867} + \underset{(0.176)}{0.019} D914_t + \underset{(0.383)}{1.113} y_t - \underset{(3.482)}{13.978} r_t, \\
y_t &= \underset{(0.471)}{0.437} + \underset{(0.028)}{0.047} D091_t + \underset{(0.032)}{0.042} D121_t + \underset{(0.044)}{0.010} D151_t + \underset{(0.050)}{0.897} y_t^*.
\end{aligned}
$$

[7]If the test statistic lies within the interval formed by the upper and lower bounds, we would need to know the order of integration of the variables to make conclusive inference as suggested by Pesaran et al. (2001).

[8]In the last column, p corresponds to the number of lags of the dependent variable, and m_k are the lags of the *k-th* regressor.

[9]In the case of the output relation, the test statistic is slightly higher than the critical value at the 10% significance level upper bound.

Notice that all relations were estimated with a constant and without a trend, and that we include the following shift dummies: $D914_t$ accounts for the level change in real money balances after 1991q4; $D091_t$ captures the structural change observed in both the Mexican and U.S. GDP starting in 2009q1; $D121_t$ controls for the steep deceleration the U.S. economy suffered in 2012q1 discussed previously; and $D151_t$ aims to capture the aforementioned divergence on the slope of the GDP series. We did not find evidence in favour of a linear trend in any of the long-run relations at conventional levels. All the relevant coefficient estimates are significant at the 5% significance level and have the expected signs.[10]

For the PPP relation, the long-run specification shows that it is not unreasonable to impose a unit elasticity in both the foreign price level and in the exchange rate.[11] Concerning the UIP relation, and according to the theory, we find that the Wald tests are unable to reject the individual null hypotheses of unit coefficients with opposite signs for r_t and r_t^* at the 5% significance level. This is also true for the Wald tests of their joint equality to one.[12] With regard to the demand for money, a unit income elasticity cannot be rejected. In particular, the p-value of this test is 0.786. Finally, we claim that, with the available evidence, the relation between the real domestic and foreign output has a unit elasticity in the long-run.[13] These long-run relations are consistent with a broad class of open-economy macroeconomic models (e.g., Garratt et al., 2003; Clarida et al., 2001).

Given the results described above, it seems reasonable to assume that PPP and UIP hold in the long-run, and that money demand exhibits a unit elasticity with respect to output. It also seems suitable to restrict the cointegration space so that domestic output has a unit elasticity with respect to foreign output. In this context, abstracting from unimportant constants, once we estimate the full system in the following section we will impose the following (oveidentifying) restrictions into the cointegrating space,

[10] We note that all the dummy variables included in OUT are significant when estimated with maximum likelihood within the VAR system below.

[11] The Wald test for the estimate related to e_t is unable to reject the null that the coefficient equals one. The corresponding test for the estimate related to p_t^* rejects with a p-value of 0.01, but we note that a joint Wald test rejects the null that both coefficients are equal with a p-value of 0.034. This may reflect the fact that both p_t and p_t^* include non-tradable goods in their definitions, as opposed to the theoretical derivation of this relation. The likelihood ratio test for over-identifying restrictions within the VECM estimated below will confirm that these restrictions are valid in the system estimation.

[12] The p-value of the Wald test for the null that the coefficient of the domestic interest rate is one is 0.577. On the other hand, the test for the null that the coefficient related to the foreign interest rate is one is 0.255. Finally, the p-value for the joint test that both coefficients are one is 0.434.

[13] The p-value for this test is 0.043. The discussed changes in the behavior of the output series, particularly in the post-2009 period seem to be weakening their long-run relation, the latter situation, however, may be temporary. Indeed, if we estimate the relation up to 2008q2, prior to the financial crisis, we find that the p-value is 0.555.

which correspond respectively to the PPP, UIP, MD and OUT long-run relations[14]

$$\begin{aligned}
p_t &= p_t^* + e_t + \xi_{1t}, \\
\Delta e_t &= r_t - r_t^* + \xi_{2t}, \\
m_t^d &= \beta_{D914} D914_t + y_t + \beta_r r_t + \xi_{3t}, \\
y_t &= \beta_{D091} D091_t + \beta_{D121} D121_t + \beta_{D151} D9151_t + y_t^* + \xi_{4t}.
\end{aligned}$$

Notice that these four long-run relations can be expressed compactly as

$$\boldsymbol{\xi}_t = \boldsymbol{\beta}' \begin{pmatrix} \mathbf{z}_t \\ \mathbf{D}_t \end{pmatrix} - \mathbf{b}_0,$$

where $\mathbf{z}_t = \left(y_t, p_t, r_t, \Delta e_t, m_t^d, p_t^* + e_t, y_t^*, r_t^*\right)'$ and $\mathbf{D}_t = (D914_t, D091_t, D121_t, D151_t)'$, $\mathbf{b}_0$ is a vector of unimportant constants, $\boldsymbol{\xi}_t$ are the deviations from long-run equilibrium, $p_t^* + e_t$ are foreign prices in domestic currency and

$$\boldsymbol{\beta}' = \begin{pmatrix} 0 & 1 & 0 & 0 & 0 & -1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & -1 & 0 & 0 & 0 & -1 & 0 & 0 & 0 & 0 \\ -1 & 0 & \beta_r & 0 & 1 & 0 & 0 & 0 & \beta_{D914} & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 & 0 & -1 & 0 & 0 & \beta_{D091} & \beta_{D121} & \beta_{D151} \end{pmatrix}. \tag{3.1}$$

The only long-run coefficients that we leave unrestricted, to be estimated from the data, are the interest semi-elasticity of the money demand and the coefficients for the shift dummies. All other parameters are restricted to their long-run theoretical values. We now define $\widetilde{\mathbf{z}}_t = (\mathbf{z}_t', \mathbf{D}_t')'$ and separate $\mathbf{z}_t$ into endogenous and weakly exogenous variables. The endogenous variables are $\mathbf{x}_t = \left(y_t, p_t, r_t, \Delta e_t, m_t^d, p_t^* + e_t\right)'$, and the weakly exogenous are $\mathbf{x}_t^* = (y_t^*, r_t^*)'$. From now on, we will treat $p_t^* + e_t$ as a single variable, which we will refer to as foreign prices in domestic currency.[15] Also notice that the error correction terms for PPP, UIP, MD and OUT will be denoted by $\xi_{1t}, \xi_{2t}, \xi_{3t}$ and ξ_{4t}, respectively.

4 Estimation and Testing of the Model

4.1 Unit Root Tests

Before we describe the estimation of the full system, we summarize in this subsection the results of the unit root tests on the variables included in the model. The Augmented Dickey-Fuller (ADF) test from Dickey and Fuller (1979) and Said and Dickey (1984)

[14]The validity of these restrictions will be tested jointly in the full system estimation below and, as will be seen, is not rejected in that context.

[15]In the unit root tests shown in the following section, however, we will make the tests individually for the foreign price level and for the exchange rate.

statistics and the DF-GLS test from Elliott et al. (1996) for the variables in levels and in first differences are presented in Table 3.

The results of both the ADF and the DF-GLS tests suggest that we can treat y_t, r_t, m_t^d, e_t, y_t^*, p_t^*, r_t^*, p_t^o, $\widetilde{r}_t^*$ and pp_t^* as I(1) variables. There is certain ambiguity regarding the order of integration of the domestic price level, p_t. Given the results in Table 3 and a further ADF test on $\Delta^2 p_t$, the evidence suggests that the domestic price level seems to behave like an I(2) variable. However, specific tests for inflation in Mexico have shown that there is a change in the persistence of it around the beginning of 2001 (Chiquiar et al., 2010). Hence, the price level would seem to behave like an I(2) series only for the first part of the sample, and as an I(1) variable for the second part. In the analysis we will treat p_t as being an I(1) variable.[16]

4.2 Long-Run Relations

We now undertake the system estimation of the structural cointegrating VAR describing the full dynamic behavior of the variables included in the macroeconomic model. The cointegrating VAR embeds the structural long-run relations described before as the steady-state solution, while the short-run dynamics are estimated from the data without restrictions. Formally, the long-run relations we identified previously are incorporated into an otherwise unrestricted VAR in differences of the variables. In particular, the VAR includes lags of the differenced series of Mexico's macroeconomic variables, as well as the first lag, $\boldsymbol{\xi}_{t-1}$, of the deviations from the long-run equilibrium conditions (i.e. the error correction terms). The U.S. interest rate and GDP also enter the VAR, although we assume these variables are weakly exogenous and, thus, are not supposed to respond to the lagged error-correcting terms. In error-correction form, the model to be estimated can be written as

$$\Delta \mathbf{z}_t = \mathbf{a} - \boldsymbol{\alpha}\boldsymbol{\beta}' \widetilde{\mathbf{z}}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \mathbf{z}_{t-i} + \mathbf{u}_t. \tag{4.1}$$

We first determine the specification of (4.1) by choosing its appropriate lag length and assessing whether it is appropriate to assume that the long-run solution may be represented by a set of four cointegrating relation; this is, we test whether it is reasonable to set the rank of the $\boldsymbol{\Pi}$ matrix to four. In this context, we first determine the lag order of the underlying VAR. In order to do so, we estimate a VAR in levels including

[16] Treating inflation as a variable without a stochastic trend is consistent with our sample corresponding to a time period when fiscal dominance was absent (see Capistrán and Ramos-Francia, 2009). For instance, Mexico started the public debt renegotiations with the so-called Brady Plan on July 1989.

$y_t, p_t, \Delta e_t, r_t, m_t^d, y_t^*, r_t^*$ and $p_t^* + e_t$. We also include the set of dummy variables and Δp_t^o, $\Delta \widetilde{r}_t^*$ and Δpp_t^* with one lag as exogenous variables. All computations are carried out over the period 1990q1 to 2015q4. Table 4 shows the results from the application of different lag order selection criteria: the Akaike information criterion (AIC), the Schwarz information criterion (BIC), the Hannan-Quinn criterion (HQ) and Lütkepohl's final prediction error (FPE). Notice that the AIC points to a lag order of 4, whereas the HQ and FPE criteria favor a lag of order 2 and the BIC a lag order of 1. We proceed with a lag length of p = 2 because, given our sample size, considering a higher number of lags did not seem appropriate as the number of coefficients to be estimated would rise quickly. Moreover, a lag order of 1, as suggested by the BIC, would not be able to fully capture the short run dynamics, as discussed by Garratt et al. (2006).

Having decided on the lag order of the VAR, we now need to determine the appropriate number of cointegrating relations that should be included (or cointegrating rank, R). To do so, we compute the corresponding trace test statistics. Note that the usual critical values for these may not be applicable in our model, given the inclusion of the dummy variables and of a set of weakly exogenous variables. We therefore simulate the critical values for the test.[17] Table 5 shows the eigenvalues and the trace statistic, together with their simulated 5% critical values and the corresponding tests' p-values. Note that the test indicates the presence of four cointegrating vectors, that is $R = 4$. This seems a sensible choice, considering the four long-run theoretical restrictions described in the previous section.

Given the results above, we now restrict the cointegrating space to include the four long-run relations identified with the ARDL models described previously. In particular, we impose oveidentifying restrictions on $\boldsymbol{\beta}'$ as in (3.1). To test the validity of these restrictions, we conduct a likelihood ratio (LR) test as outlined in Garratt et al. (2006). We first define a set of $R^2 = 16$ exactly identifying restrictions on $\boldsymbol{\beta}'$ as a subset of those contained in (3.1) as follows

$$\boldsymbol{\beta}'_{EI} = \begin{pmatrix} \beta_{1,1} & 1 & \beta_{1,2} & \beta_{1,3} & \beta_{1,4} & \beta_{1,5} & \beta_{1,6} & \beta_{1,7} & 0 & 0 & 0 & 0 \\ 0 & \beta_{2,2} & \beta_{2,3} & 1 & \beta_{2,5} & \beta_{2,6} & 0 & \beta_{2,8} & 0 & \beta_{2,10} & 0 & \beta_{2,12} \\ \beta_{3,1} & 0 & \beta_r & \beta_{3,4} & 1 & 0 & 0 & \beta_{3,8} & \beta_{D914} & \beta_{3,10} & 0 & \beta_{3,12} \\ 1 & 0 & \beta_{4,3} & \beta_{4,4} & 0 & 0 & \beta_{4,7} & \beta_{4,8} & 0 & \beta_{D091} & \beta_{D121} & \beta_{D151} \end{pmatrix} \quad (4.2)$$

The overidentified model contains 47 restrictions, 39 directly seen in (3.1) and 8 that stem from the fact that r_t^* and y_t^* are assumed to be weakly exogenous. Hence there are 23 over-identifying restrictions and the LR test statistic should distribute as

[17] See the details of the CATS procedure to simulate these critical values in Dennis (2006).

a chi-square with 23 degrees of freedom (Pesaran and Shin, 2002). However, given the relatively large dimension of the VAR model and the available degrees of freedom, we proceed to test the significance of the LR statistic using critical values computed using bootstraping techniques.[18] We obtain the critical values from a parametric bootstrap with 1000 replications. For each replication, an artificial data set is generated assuming first that (4.2) is the cointegrating matrix of the data generating process and then that (3.1) is so. This is done using the observed initial values of each variable, the estimated model under both sets of restrictions, a set of random innovations, and taking the (weakly) exogenous variables as fixed across replications.[19]

The LR test on the validity of the over-identifying restrictions is carried out on each of the simulated data sets and the empirical distribution of the test statistic is derived across all the replications. The LR test statistic is 93.56, while the simulated critical values are 148.95 for the 5% and 141.11 for the 10% significance levels. Interestingly, our estimates from the full VAR model fail to reject the validity of the imposed long-run over-identifying restrictions.

We now turn to a brief description of the results of the estimation of $\boldsymbol{\beta}'$ in (3.1). The estimation method employed is quasi maximum likelihood.[20] Apart from unimportant constants, only five long-run coefficients are estimated freely (the semi-elasticity of money demand with respect to the interest rate and the coefficients for the shift dummies). The rest of the long-run coefficients are fixed at their theoretical values. The estimate of the semi-elasticity of money demand with respect to the interest rate is -55.434, with a t-statistic of -18.891. This result means, as expected, that money demand responds negatively to the interest rate. On the other hand, the $D914_t$ shift dummy enters the money demand equation with a coefficient of -0.015 which is estimated imprecisely (its t-statistic is -0.112). The shift dummies included in the OUT relation are all positive and significant at the 5% level, $D091_t$ has as an associated parameter 0.065 with a t-statistic of 5.761; $D121_t$ has as an associated parameter 0.028 with a t-statistic of 2.058; and $D151_t$ has as an associated parameter 0.120 with a t-statistic of 5.321.

The time plots of the four cointegrating relations are shown in Figure 8. In particu-

[18]The asymptotic approximation is only valid when relatively large samples of data are available (see, Abadir et al., 1999). However, even if the asymptotic distribution were valid, the bootstrap often greatly reduces the error in the rejection probability typically found when using small sample sizes (Horowitz, 2001).

[19]The random innovations are obtained from a normal multivariate distribution with mean equal to zero and a covariance matrix equal to the covariance matrix estimated from the residuals of each model.

[20]See Pesaran and Shin (2002) for the properties of quasi maximum likelihood in this context.

lar, we plot each equilibrium error, corrected for short-run dynamics. We should expect these series to be stationary. We note that, although deviations from PPP seem to be relatively persistent, there indeed seems to be an error-correcting behavior that leads this equilibrium error to be stationary. The UIP condition and the money demand function seem to exhibit an even clearer error-correcting behavior. Finally, apart from the large shocks during the Tequila crisis (1994-1995), the financial crisis in 2009 and the last part of the sample in 2015, deviations from the output long-run relation seem to behave in a stationary fashion, although they are also quite persistent.

As a last robustness check, we test the recursive stability of the long-run estimates by means of the approach in Hansen and Johansen (1999). In particular, starting from a base sample going from 1990q1 to 2006q2, we re-estimate recursively the model, adding an additional observation at a time. Let $\boldsymbol{\beta}^{(T)}$ denote the estimate of the long-run parameters using the full sample, and $\boldsymbol{\beta}^{(t)}$ denote the estimate using a sample going from 1990q1 to t, for $t = 2006q3....2015q4$. Then, we compute two tests for the constancy of $\boldsymbol{\beta}$. The first is a max test on the difference between $\boldsymbol{\beta}^{(T)}$ and $\boldsymbol{\beta}^{(t)}$ (max test of beta constancy) and the second is a recursive test of $\boldsymbol{\beta}^{(T)} \in span\{\boldsymbol{\beta}^{(t)}\}$ (test of beta = "Known beta", see Dennis, 2006). The results are plotted in Figure 9. The test statistics are scaled by the corresponding 5% critical value, so that a test statistic larger than unity would lead to a rejection of the null of stability. We include the test statistics when all the parameters are reestimated in each recursion (X form) and when the short run dynamics are concentrated out with the full sample estimates (R form). As may be noted, we find no evidence of structural instability in the long-run cointegration relations with the max test, and only one period of instability about 2006q2 when looking at the known beta test. Had we scaled the test statistics in this case with the 10% critical value, we would not be able to reject the null of stability.[21] We should note that, even if most of the long-run parameters were fixed from the start, an important misspecification would lead to instability of the parameters that were indeed estimated.

4.3 Error Correction Equations

We now turn to a brief analysis of the estimates of the reduced-form error correction equations. Table 6 summarizes these estimates and includes some diagnostic statistics. The deviations from the long-run relations are statistically significant, at conventional levels, in several equations. In particular, deviations from PPP, $\xi_{1,t-1}$, are statistically significant in the equation for the domestic interest rate and domestic prices, suggest-

[21] When we test with an estimation sample that runs from 1990q1 to 2008q2, that is, prior to the financial crisis, we find no evidence of instability.

ing that the latter tend to adjust to attain PPP in the long-run through short-run changes in the level of inflation. Domestic prices also respond in a statistically significant manner, with a positive sign, to positive deviations of GDP from its long-run relation with U.S. output, $\xi_{4,t-1}$. In turn, as expected, real money balances adjust to correct short-run deviations from money demand, $\xi_{3,t-1}$. Similarly, domestic output exhibits a significant error-correcting behavior with respect to deviations from its long-run equilibrium relation with the U.S. output level. These deviations, in turn, also enter significantly into the money and interest rate equations. Finally, it is noteworthy to mention that interest rate changes seem to lead to an appreciation of the currency and to an error-correcting behavior of the interest rate. The first of these results would seem to be consistent with a Dornbusch (1976) type of mechanism in which a monetary policy shock (here assumed to correspond to a rise in the interest rate), would lead to an appreciation of the currency that overshoots the exchange rate's long-run response, until its value is consistent with devaluation expectations that, given the increase in the domestic interest rate, satisfy the UIP condition. We will see further results supporting the presence of this mechanism below.

Concerning the results of the diagnostic statistics, Doornik and Hansen (1994) test for normality suggest that the residuals of all equations corresponding to domestic variables exhibit non-normality. A visual analysis of the residuals suggests that this reflects large outliers associated with the first quarter of 1995, when the Tequila Crisis kicked in. By re-estimating the model with a dummy accounting for these outliers, we found that the departure from normality derived from this shock does not seem to have significant effects on our main findings, particularly, since we are estimating through quasi maximum likelihood. In contrast with the strong rejection of normality, ARCH tests cannot reject the null of homoskedasticity of errors in any of the equations, while serial correlation LM tests reject the null of no serial correlation at a 10% significance level only in the case of the U.S. GDP. Finally, a Ramsey RESET test rejects the null of no miss-specification for all cases except for the U.S. GDP and U.S. interest rate. However, trying to correct for this problem by including further lags would possibly imply a significant loss of degrees of freedom.

5 Impulse-Response Analysis

In this section, we illustrate the dynamic properties of the model estimated previously by means of the relevant impulse-response functions. While the long-run response of each endogenous variable is ultimately driven by the long-run relations that are embedded in the model, the study of the short-run dynamic responses to different shocks, as summarized by these impulse-response functions, may be relevant on its own right for forecasting and policy making. This analysis may give a more complete picture of the short-run dynamics of the system than the partial, equation-by-equation, results presented above.

We first analyze the response of each variable to non-structural shocks. In particular, we illustrate the effect of a one standard error shock to each variable on the time path of all the endogenous variables in the model. The analysis is based on the computation of Generalized Impulse Response Functions (GIRFs) (see Koop et al., 1996 and Pesaran and Shin, 1998). The GIRFs are invariant to the ordering of the variables in the VAR and do not require specific identifying assumptions. Instead, their computation relies on the estimated reduced-form errors covariances that consider the contemporaneous linkages between shocks that have prevailed historically.

Having described the main dynamic properties of the model through the use of GIRFs, we conduct an exercise intended to identify the dynamic response of the main endogenous variables of the model to a monetary policy shock. We do not attempt to describe the effects of other structural shocks. To identify the monetary policy shock, in this particular exercise we rely on a Cholesky decomposition based on a specific ordering of the variables in the VAR, as suggested by Sims (1980). In particular, we follow an identification approach that satisfies a recursiveness assumption, wherein we suppose that monetary policy shocks are orthogonal to the information set that the monetary authority is assumed to have when setting its policy instrument (see Christiano et al., 1999 and Garratt et al., 2003).

Specifically, we assume that, when Mexico's central bank makes its monetary policy decisions, reflected in the short-run interest rate, it observes the current values of the U.S. interest rates and output levels, the foreign prices in domestic currency and the domestic price level. This assumption allows us to identify a monetary policy feedback rule that sets the interest rate as a function of current values of these variables and of lags of all variables in the system.[22] This, along with the recursiveness assumption, allows us to recover the series of monetary policy shocks which, in this context, are

[22]Garratt et al. (2003) show that this identification scheme may be rationalized by a decision-based approach in which an interest rate policy rule is derived by assuming that the central bank minimizes a quadratic loss function in inflation and output growth. The latter is subject to a structural model that links the interest rate to the target variables, conditioning on the available information. Christiano et al. (1999) choose not to include current GDP after the interest rate solely on the basis that a restrictive shock to the latter implies a non-negative response of the former. We, however, do not encounter such a puzzle and follow the more sensible assumption that the central bank does not know the current level of GDP at the time of the shock as argued by Rudebusch (1998).

orthogonal to the current shocks in the U.S. interest rates and output levels, the foreign prices in domestic currency, the domestic prices and the lags of all variables in the system.[23]

To make this identification scheme operational, we order the variables in the VAR as follows

$$\mathbf{z}_t = \left(r_t^*, y_t^*, e_t + p_t^*, p_t, \vdots\, r_t, \vdots\, \Delta e_t, y_t, m_t^d \right)', \tag{5.1}$$

that is, we order the equations in the VAR so that all variables whose current values are assumed to be known by the central bank locate prior to the interest rate equation, while those that are assumed to be observed with a lag locate at the end of the $\mathbf{z}_t$ vector. With the ordering (5.1), the Cholesky decomposition-based impulse-responses of each variable to shocks in the interest rate equation may be interpreted as the dynamic responses of each variable to the monetary policy shock.[24] As shown in Christiano et al. (1999) and in Garratt et al. (2003), these dynamic responses will be invariant to the particular ordering of the equations within the blocks of variables that appear before and after the interest rate in the VAR.[25]

5.1 Response to Non-Structural Shocks

Figures 10-12 illustrate the response of each endogenous variable to a unit standard error shock in each one of the variables in the model.[26] As discussed before, these are responses to observable, non-structural shocks. Thus, it is hard to identify the underlying mechanisms that lead to the dynamic correlations that are illustrated in these GIRFs. In particular, while these impulse-response functions summarize the

[23]Results are invariant to the inclusion of the peso dollar exchange rate before or after the interest rate in the recursive order of the VAR. We choose the latter ordering since our main concern is to assess the effect of monetary policy on the exchange rate. We thank an anonymous referee for this suggestion. Note that this identification scheme in turn implies that, in the VAR, the foreign prices in domestic currency, the domestic prices, the U.S. interest rates and output levels can only respond with a lag to monetary policy shocks.

[24]For this exercise, we are considering the short-term interest rate as the monetary policy instrument. However, during our sample Banco de México has used different monetary policy instruments. Before the Tequila crisis, Mexico had a target zone for the exchange rate. With the crisis, Mexico was forced to float the currency, and between 1995 and January 2008, Banco de México used a non-borrowed reserve requirements target, called the “Corto”, to affect interest rates. Nonetheless, since April 2004, Banco the México had started to send signals to the market about its desired level of interest rates, although the *de jure* switch to the use of the overnight interest rate as the monetary policy instrument was made in January 2008 (Banco de México, 1996; Banco de México, 2007).

[25]If we were interested in identifying the dynamic effects of other structural shocks, then the ordering of the variables in the blocks that appear before and after the interest rate would become relevant.

[26]We used the response functions of the depreciation rate and of the price level to compute the implied responses of the exchange rate and the inflation rate.

dynamic interrelation observed historically among the variables in the model, they may not be directly linked to the response to any particular structural shock. This limits the economic interpretation of these functions. It is important to emphasize from the start, however, that the identification of these responses may be especially reliant on the large macroeconomic shock that led to the crisis of the Mexican economy at the mid nineties, in which a large devaluation led to high inflation and a decrease in money demand, in a context of a deep recession and high interest rates. As may be seen in Figures 10-12, the historical (reduced-form) correlation observed among the variables in the model suggests the presence of positive links between the dynamics of exchange rate, prices and interest rates, as well as a negative correlation of these variables with money demand and output.

Concerning the responses of output to diverse observable shocks displayed in Figure 10 a), there are several findings that deserve attention. As expected, in the long-run Mexico's GDP is fundamentally determined by shocks in U.S. output levels. In the short run, however, output also responds in a non-negligible way to other shocks as well. In particular, GDP seems to respond negatively to exchange rate, prices and interest rate shocks. The response to exchange rate shocks is larger in absolute value than the response to prices in the short-run. The results seem to suggest that real devaluations may have an initially negative impact on domestic demand that offsets their positive effect on exports, so that their net effect on output is negative in the short-run.[27] On the other hand, while the short-run negative response of output to interest rate shocks could in principle reflect a monetary transmission mechanism, we will see in the identification exercise below that the apparent negative response of output to monetary policy shocks seems to be much more delayed than what the results described here suggest. Thus, the apparent negative short-run response of output to interest rate shocks we find here seems to be driven mainly by some other unobservable shock that leads simultaneously to increases in interest rates and decreases in output levels.

According to the GIRFs analysis, inflation increases temporarily after an exchange

[27]This result may be linked to the literature on the contractionary effects of devaluations. In particular, while a real devaluation may stimulate output by inducing an increase in the production and exports of traded goods, this result could be offset by balance sheet effects, by a contraction of supply when imported inputs and capital are used in production or by a redistribution of resources towards groups with higher propensities to save. In this context, the net effect on output of a real devaluation could become negative (see Diaz-Alejandro, 1963; Krugman and Taylor, 1978; Agénor, 1991; Gavin, 1992; Céspedes et al., 2004; Frankel, 2005). The empirical evidence from developing countries tends to suggest that these mechanisms may indeed be sufficiently strong to induce a negative effect of devaluations on output (see Edwards, 1989; Moreno, 1999). For the particular case of Mexico, Kamin and Rogers (2000) find evidence that real devaluations are contractionary. The negative response of output to real exchange rate shocks that we uncover in our analysis is consistent with their findings.

rate shock is observed (see Figure 11 a)). This is the mechanism through which the domestic price level adjusts to satisfy the purchasing power parity relation in the long-run (see Figure 10 b)). In contrast, as expected, a money demand shock seems to lead to a reduction in inflationary pressures. Also as expected, after approximately one year, demand pressures as measured by shocks to Mexico or to U.S. GDP levels seem to increase inflation temporarily.

A result that may seem odd, and that resembles the findings in Sims (1992), Racette and Raynauld (1992) and Grilli and Roubini (1995), is the fact that an increase of the domestic interest rate seems to lead to a temporary increase in inflation and to a depreciation of the currency as shown in Figures 11 a) and b), respectively. That is, these impulse-response functions exhibit both the price and the exchange rate puzzles. If we considered a monetary transmission mechanism, given the UIP condition, we would expect a negative response of inflation and a quick appreciation of the currency, followed by a gradual depreciation, after an interest rate shock (Dornbusch, 1976). As mentioned before, however, in this GIRFs analysis we do not claim to be identifying the response to structural shocks in general, nor to monetary policy shocks in particular. Indeed, the results we find here could be affected by two-way causality between the interest rate and the exchange rate. In particular, the monetary authority could be reacting endogenously to exchange rate shocks by raising interest rates to avoid their inflationary consequences, consistently with this possibility, we find a strong positive response of the interest rate to shocks in the exchange rate (see Figure 12 a)). This response would make it difficult to uncover the causal relation going from the interest rate to the exchange rate, unless we use a specific identification strategy to isolate exogenous monetary policy shocks. This is the main insight of Cushman and Zha (1997) and Bjørnland (2009). We pursue this in the next subsection.

We observe that money demand shocks tend to lead to an appreciation of the currency, while U.S. interest rate shocks tend to depreciate it, both phenomena displayed in Figure 11 b). These results seem to be driven by the effects of investors' portfolio decisions on the foreign exchange market. In particular, an increase in the demand for pesos would tend to reduce demand pressures in the foreign exchange market, while an increase in the foreign interest rate would lead to a depreciation of the exchange rate.

The responses of domestic interest rates are shown in Figure 12 a). We already noted that they seem to show an especially strong short-run positive response to exchange rate and price shocks. We also note a positive response of domestic interest rates to shocks in the U.S. interest rates. For a given depreciation rate, this seems to be associated with the UIP relation identified previously. The estimated response of interest rates to domestic and foreign output shocks is positive for horizons beyond four quarters. This could suggest that we are either identifying the positive effect on interest rates from demand shocks (i.e., shifts of the IS schedule) or a monetary policy rule responding to demand shocks.

Money demand responds positively to output shocks in the short run but, in contrast with the long-run relation identified previously, the response seems to turn negative after several quarters (see Figure 12 b)). This holds even when the money demand long-run relation is satisfied by the GIRFs illustrated here. The reason why this odd result may be appearing is that, as already seen, an output shock also leads to an increase in interest rates after several quarters, which itself reduces money demand directly and indirectly through an eventual negative effect on output itself. Finally, an increase in interest rates, prices and the exchange rate seem to be negatively correlated with money demand.

5.2 Responses to a Monetary Policy Shock

We first discuss the main features of the monetary policy rule implied by the recursive identification given by the Cholesky ordering in (5.1). Figure 13 contains the orthogonalized shocks. To illustrate the statistical significance of the estimated responses, we approximated their distribution with a Monte Carlo integration exercise.[28] Following Sims and Zha (1999), we show the 16% and 84% quantiles from these distributions as confidence bands. These quantiles would correspond to approximately one standard deviation if we were doing symmetrical error bands based upon estimates of the variance. As may be noted, the estimated responses of the short-term interest rate to diverse shocks are sensible and consistent with a standard monetary policy rule.[29] In particular, the monetary authority seems to increase interest rates after demand shocks or perceived inflationary pressures. This is evident by noting that interest rates rise after both Mexico and U.S. output shocks, or after an exchange rate shock (which, as noted before, tends to lead to higher inflation). This policy-induced response of the interest rate to exchange rate shocks is precisely the mechanism that we claim may have made it difficult to identify a Dornbusch type of response of the exchange rate in the previous section. We note that interest rates also rise after a shock to U.S. interest rates. This

[28]The exercise was conducted using the RATS software. The procedure computes the first moments and variances of the posterior distribution of the orthogonalized impulse responses, using a Monte Carlo integration exercise based on Zellner (1971). We conducted 5,000 replications of the estimates for the impulse response functions. For further details, see Example 13.3 in Doan (2004).

[29]We restrict the scale of the vertical axis in all the graphs that appear in Figure 13 to be the same, given that all show the effect on the interest rate.

may reflect the fact that, to avoid an unwanted depreciation of the currency that could lead to higher inflation, the monetary authority would need to increase domestic rates after an increase in foreign interest rates.

We now illustrate the response of the economy to a monetary policy shock, as identified through the recursive ordering in (5.1) described previously. Figure 14 summarizes the results. The impulse-responses include the 16% and 84% quantiles from the distribution approximated with the Monte Carlo integration exercise. In general, the results are consistent with the prior expectations concerning the response the economy should exhibit to a monetary contraction. An important result, in contrast with those of the previous section, is that our identification strategy leads to responses that do not exhibit the price and the exchange rate puzzles any more.

In particular, the results suggest that the main mechanism through which monetary policy affects inflation in Mexico seems to be fundamentally linked to its short-run effects on the exchange rate which, in turn, is a significant determinant of the price level. The mechanism that leads to these results seems to be consistent with a Dornbusch (1976) type model: a contractionary monetary policy shock induces a strong immediate appreciation of the exchange rate, followed by a gradual depreciation.[30] The timing and the direction of the exchange rate response to a monetary policy shock we uncover here is similar to those found by Bjørnland (2009) for other small open economies.[31]

This result, in turn, is reflected in the response of prices to the monetary policy shock. Driven by the quick effect that the monetary policy shock has on the nominal exchange rate, a monetary contraction seems to lead to a temporary reduction in inflation.[32] In particular, a tightening of monetary policy tends to temporarily appreciate the currency and, through this channel, to reduce inflationary pressures. This transmission mechanism seems to be fairly fast as the largest impact on inflation happens within the first year after the shock. Moreover, both on impact and in the medium term, the monetary policy shock seems to have the expected negative effect on output.

Our results differ significantly from previous attempts to identify the exchange rate monetary transmission mechanism in Mexico in different dimensions. First, the results presented in Figure 14 do not show a delayed, hump-shaped response of the exchange

[30] The support we find for Dornbusch (1976) overshooting result may in part reflect the fact that we imposed UIP as one of the long-run restrictions in our model. It is relevant to recall, however, that we tested the validity of this restriction before imposing it and within the full system estimation.

[31] Using an identification strategy that allows for contemporaneous effects, Bjørnland (2009) finds that the maximum impact on the exchange rate occurs within 1-2 quarters. The response we find is also very sharp, and more in line with the Dornbusch effect than with delayed overshooting.

[32] The initial response of inflation to an increase of the interest rate may be rationalised as a one-and-for-all increase in the credit costs for producers. We note nonetheless that said response is particularly small.

rate to monetary policy shocks that violates uncovered interest parity (see Martínez et al., 2001; Gaytán and García-González, 2006). Second, previous studies have been unable to uncover significantly large effects on output and prices (e.g. Del Negro and Obiols-Homs, 2001) or have found overall positive responses of output to monetary policy shocks, that are hard to rationalize theoretically (e.g., Gaytán and García-González, 2006). Just as in the case of the identification of an exchange rate overshooting mechanism, the reason why we are able to identify stronger and theoretically consistent output and price responses as opposed to the previous literature may be linked to our identification strategy and with our explicit use of long-run equilibrium conditions in our estimation procedure. We note that our results are broadly similar with those found by Carrillo and Elizondo (2015) but we use a model that considers explicitly long run relations and trends. At the same time, we use a more extended sample, which indicates that results are robust across different regimes of exchange rate and inflation.[33]

6 Concluding Remarks

In this paper we find that four long-run relations suggested by economic theory are satisfied by data in our sample: a purchasing power parity, an uncovered interest parity, a money demand and a relation between Mexico's and U.S. real outputs. We use these relations as the foundation of a structural cointegrated VAR model for the Mexican economy. This approach allows us to directly model the trends in the data, while leaving the short-run dynamics unrestricted. The model reflects Mexico's condition of a small open-economy, with both real and financial links to the U.S. economy. In addition, the model includes money, a variable that has been missing in some models recently used in central banks, but that plays an important role as suggested by our data.

The estimated model can be used to analyze the dynamics of the Mexican economy and to produce forecasts. In this paper we have advanced in the former. First, we use the model to analyze the implied (unconditional) dynamics of the economy via the Generalized Impulse Responses. We find positive links between the dynamics of the exchange rate, prices and interest rates, as well as negative correlations of these variables with money and output. Second, we use the model to analyze the transmission mechanism of monetary policy, using a recursive strategy for identification. We find

[33] We include in Figure 15 the impulse response functions that result from re-estimating the short-run dynamics of the VAR within the period 2001-2015. The results are qualitatively the same as those included in Figure 14.

that a contractionary monetary policy shock seems to have a temporary negative effect on output and prices, and that it induces a sharp appreciation followed by a gradual depreciation of the exchange rate. The timing of the effects point to the exchange rate channel as one of the main channels through which monetary policy could be affecting prices in the Mexican economy.

Future research should look at the forecasting ability of the model. For this purpose, the long-run restrictions may help in terms of the typical trade-off between bias and variance, since they may reduce the variance of the forecasting errors by reducing parameter uncertainty, without a large effect on the bias. The model can also be used as the core model to help specific, satellite, models, as advanced by Garratt et al. (2006). For instance, a sectoral model to analyze output in different sectors of the economy, or a satellite model to analyze different components of inflation, such as the inflation of tradables and non-tradables. Under this approach, a block recursive structure could allow variables and error correction terms in the core model to influence the dynamics of the satellite models, but not vice-versa.

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Variable	Source
y_t:= natural log of the Mexican real GDP s.a. (2003=100)	INEGI
p_t:= natural log of the Mexican Consumer Price Index, s.a. (2002=100)	INEGI
r_t:= $0.25\ln(1+R_t/100)$, R_t = 91 days CETES interest rate per annum	Banco de México
m_t:= natural log of the Mexican real M1 money stock s.a. ($M1/P$)	Banco de México
e_t:= natural log of the nominal peso-dollar effective exchange rate	Banco de México
y_t^*:= natural log of the U.S. real GDP s.a. (2009=100)	BEA
p_t^*:= natural log of the U.S. Consumer Price Index s.a. (1982-1984=100)	BLS
r_t^*:=$0.25\ln(1+R_t^*/100)$, R_t^* = 3 month T-Bill interest rate per annum	Federal Reserve
p_t^o:= natural log of the West Texas Intermediate oil price	IMF
pp_t^*:= natural log of Producer Price Index for All Commodities, Index (1982=100)	Federal Reserve
$\widetilde{r}_t^*$:= TED spread is the spread between 3-Month LIBOR based on U.S. dollars and 3-Month Treasury Bill	Federal Reserve

Note: s.a. means that the variables are seasonally adjusted using the TRAMO-SEATS methodology, Gomez and Maravall (1996).

Table 1: List of Variables in the Model

	EC	*t-stat*	CV Bounds		*F-stat*	CV Bounds		$\bar{R}^2$	$ARDL(p, m_1, \ldots, m_k)$	
PPP	-0.07	-6.38	-3.21	-2.57	9.87	2.63	3.35	0.89	(2,2,2)	
UIP	-1.10	-14.95	-3.21	-2.57	53.26	2.63	3.35	0.75	(1,1,0)	
MD	-0.10	-8.10	-3.46	-2.57	6.43	2.63	3.35	0.78	(1,1,0)	$D914$
OUT	-0.19	-3.61	-3.66	-2.57	3.70	3.02	3.51	0.28	(1,2)	$D091, D121, D151$

Note: PPP denotes purchasing power parity (p, p^*, e), UIP the interest rate parity $(\Delta e, r, r^*)$, and MD money demand (m, y, r). Estimates of the long-run coefficientes are shown in the text. The columns 2 to 5 show the error-correction term (EC), its *t-ratio* and the lower and upper bound of the associated 5% critical values. The next two columns show the *F-statistic* for exclusion of the levels variables and the respective 10% lower and upper critical value bounds. The adjusted $\bar{R}^2$ refers to that of the regression in first differences. The sample period is 1990q1 to 2015q4. The specification shows the number of lags: p corresponds to the number of lags of the dependent variable, m_k is the number of lags for the *k-th* regressor (e.g. for PPP the dependent variable is p, the first regressor is p^* and the second regressor is e), in each case a (restricted) constant was included. $D914, D091, D121$ and $D151$ are *dummy* variables defined in the text. The lag length was chosen according to the BIC criterion with a maximum length of 8.

Table 2: Autoregressive Distributed Lag Models

Variable	ADF	Lag structure	DF-GLS	Lag structure
Tests with a constant and a trend				
y_t	-2.81	0	-2.28	0
p_t	-1.55	5	-0.91	3
m_t^d	-3.24	1	-1.78	1
e_t	-1.63	0	-1.52	3
y_t^*	-1.08	2	-1.14	2
p_t^*	-1.95	2	-0.70	1
p_t^o	-1.81	2	-1.76	5
pp_t^*	-1.69	2	-1.64	2
Tests with a constant				
r_t	-2.48	4	-0.63	4
r_t^*	-2.07	4	-0.69	4
$\widetilde{r}_t^*$	-2.72	3	-0.99	3
Δy_t	-8.91	0	-2.24	6
Δp_t	-2.34	2	-2.03	2
Δr_t	-4.72	2	-2.96	10
Δm_t^d	-3.56	3	-2.36	5
Δe_t	-4.50	2	-0.66	6
Δy_t^*	-2.35	8	-2.00	8
Δp_t^*	-1.88	11	-1.89*	11
Δr_t^*	-3.24	2	-1.73*	2
Δp_t^o	-8.47	0	-7.84	0
$\Delta \widetilde{r}_t^*$	-11.86	0	-2.01	12
Δpp_t^*	-6.82	0	-1.52	12

Note: The 5% CV for the Augmented Dickey Fuller (ADF) tests are -3.45 (constant and a trend) and -2.88 (constant), Dickey and Fuller (1979) and Said and Dickey (1984). The 5% CV for the Elliott et al. (1996) DF-GLS tests are -3.02 (constant and a trend) and -1.94 (constant). The lag structure is selected using the Modified Akaike Criterion from Ng and Perron (2001). $*p - value \leq 10\%$. The sample period runs from 1990q1 to 2015q4.

Table 3: Unit Root Tests

Lags	AIC	BIC	HQ	FPE$\times 10^{-35}$
1	-79.00	-75.34*	-77.51	5.93
2	-79.76	-74.47	-77.61*	2.96*
3	-79.88	-72.96	-77.07	2.97
4	-80.05*	-71.51	-76.59	3.04

Note: AIC is the Akaike information criterion, BIC the Schwartz information criterion, HQ is the Hannan-Quinn criterion and FPE is the final prediction error. *Denotes the suggested lag order for the VAR estimated from 1990q1 to 2015q4.

Table 4: Lag Length Selection Criteria

Rank	Eig.Value	Trace	95% crit.	P-Value
0	0.83	393.35	170.23	0.00
1	0.54	213.67	134.28	0.00
2	0.43	134.28	101.28	0.00
3	0.37	77.55	72.46	0.02
4	0.21	29.95	46.75	0.56
5	0.06	6.41	24.34	0.93

Note: The sample period is 1990q1 to 2015q4. Critical values are simulated with 5,000 replications.

Table 5: Trace Test and Simulated Critical Values

	Δp_t	$\Delta(p_t^* + e_t)$	Δm_t^d	Δy_t	Δr_t	$\Delta\Delta e_t$	Δy_t^*	Δr_t^*
$\xi_{1,t-1}$	-0.078***	0.059	0.055	0.004	-0.043**	0.051	-	-
	(0.02)	(0.144)	(0.066)	(0.023)	(0.02)	(0.145)	-	-
$\xi_{2,t-1}$	0.044	0.563	-2.504***	-0.134	0.338	0.165	-	-
	(0.216)	(1.599)	(0.737)	(0.25)	(0.217)	(1.603)	-	-
$\xi_{3,t-1}$	0.005	0.007	-0.049***	-0.003	0.002	0.000	-	-
	(0.004)	(0.032)	(0.015)	(0.005)	(0.004)	(0.032)	-	-
$\xi_{4,t-1}$	0.148***	0.255	-0.303***	-0.092**	0.116***	0.267	-	-
	(0.031)	(0.232)	(0.107)	(0.036)	(0.031)	(0.233)	-	-
Δp_{t-1}	0.238**	-0.132	0.536	0.217*	-0.154	-0.211	0.041	-0.007
	(0.104)	(0.768)	(0.354)	(0.120)	(0.104)	(0.770)	(0.071)	(0.008)
$\Delta(p_{t-1}^* + e_{t-1})$	0.078	0.086	2.170***	0.024	-0.294	-0.502	0.003	0.003
	(0.214)	(1.582)	(0.729)	(0.247)	(0.214)	(1.586)	(0.146)	(0.017)
Δm_{t-1}^d	-0.012	-0.172	0.074	0.022	-0.031*	-0.178	0.014	0.000
	(0.016)	(0.119)	(0.055)	(0.019)	(0.016)	(0.120)	(0.011)	(0.001)
Δy_{t-1}	-0.129	0.577	0.522	0.237**	-0.106	0.580	0.158**	0.013
	(0.097)	(0.716)	(0.330)	(0.112)	(0.097)	(0.718)	(0.066)	(0.008)
Δr_{t-1}	-0.241	-2.812**	-0.125	0.170	-0.440***	-2.849**	0.153	-0.003
	(0.146)	(1.080)	(0.497)	(0.168)	(0.146)	(1.082)	(0.099)	(0.012)
$\Delta\Delta e_{t-1}$	0.031	-0.145	0.074	0.024	0.018	-0.159	-0.011	-0.003**
	(0.020)	(0.147)	(0.068)	(0.023)	(0.020)	(0.148)	(0.014)	(0.002)
Δy_{t-1}^*	0.034	-0.429	-0.431	0.185	0.110	-0.458	0.383***	0.056***
	(0.171)	(1.263)	(0.582)	(0.197)	(0.171)	(1.266)	(0.116)	(0.014)
Δr_{t-1}^*	0.506	15.311**	-5.820*	0.162	2.031**	15.501**	0.257	0.578***
	(0.944)	(6.975)	(3.213)	(1.089)	(0.944)	(6.991)	(0.643)	(0.075)
R^2	0.909	0.303	0.821	0.630	0.411	0.612	0.398	0.662
Ser. Corr. test	1.287	0.055	1.87	1.582	1.511	0.039	4.23**	1.475
p-value	0.257	0.815	0.171	0.209	0.219	0.843	0.040	0.225
Normality test	36.246***	248.134***	61.739***	63.94***	132.254***	255.467***	7.339**	24.287***
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.025	0.000
ARCH(2) test	1.493	0.049	0.168	0.132	0.009	0.047	5.253*	1.578
p-value	0.474	0.976	0.919	0.936	0.995	0.977	0.072	0.454
RESET test	86.415***	8.956***	5.142**	11.963***	49.262***	6.948***	0.235	0.238
p-value	0.000	0.004	0.026	0.001	0.000	0.010	0.629	0.627

Note: Standard errors are included in parenthesis. The error correction terms $\xi_{i,t-1}$ are defined in the main text. The estimates related to the dummy variables are: $\beta_{D914} = 0.015(0.136)$, $\beta_{D091} = -0.065(0.011)$, $\beta_{D121} = -0.028(0.014)$ and $\beta_{D151} = -0.120(0.023)$. Constants are not shown. Asterisks denote: *significance at the 10 percent level; **significance at the 5 percent level; ***significance at the 1 percent level. Serial Correlation test refers to a Breusch-Godfrey Serial Correlation LM test, Normality test refers to a Doornik-Hansen test for Normality.

Table 6: Reduced-form Error Correction Equations

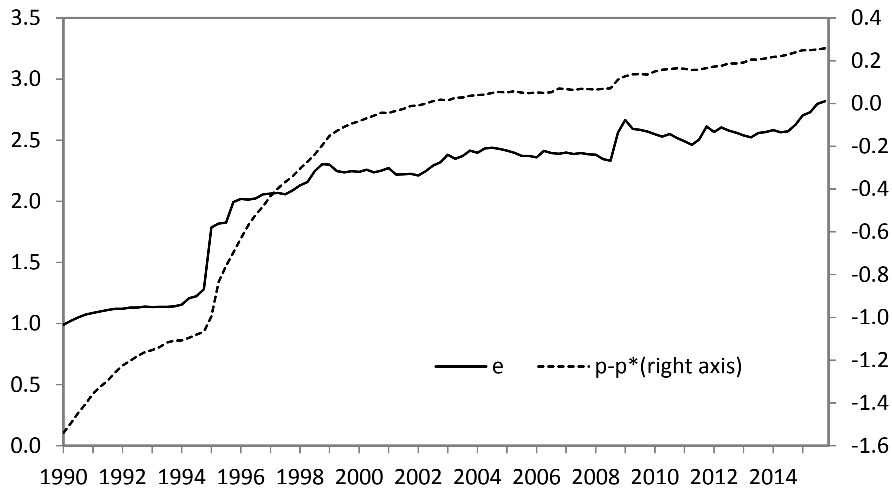
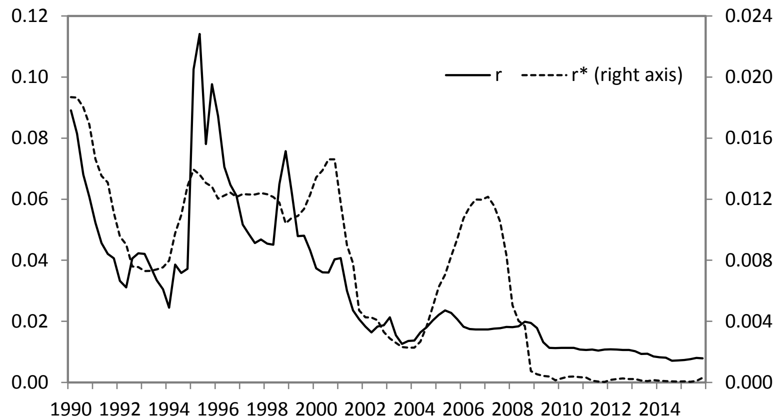
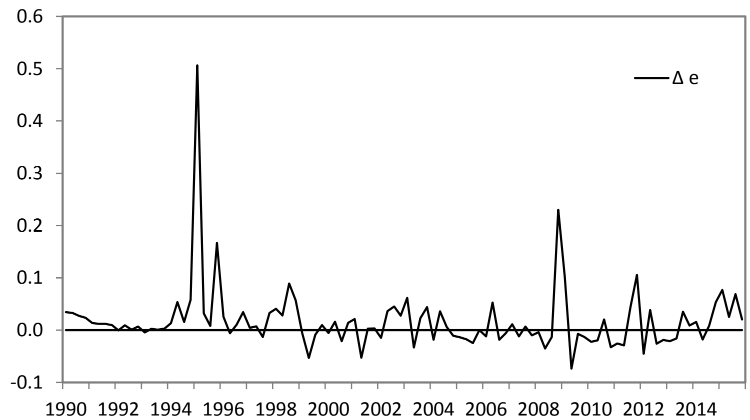


Figure 1: Exchange Rate and the Ratio of Domestic to Foreign Prices

Panel a) Domestic and foreign interest rates



Panel b) Depreciation rate



Panel c)

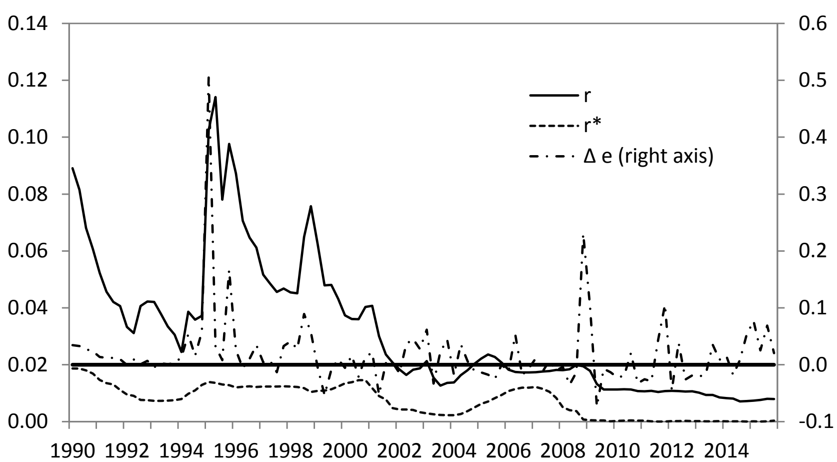
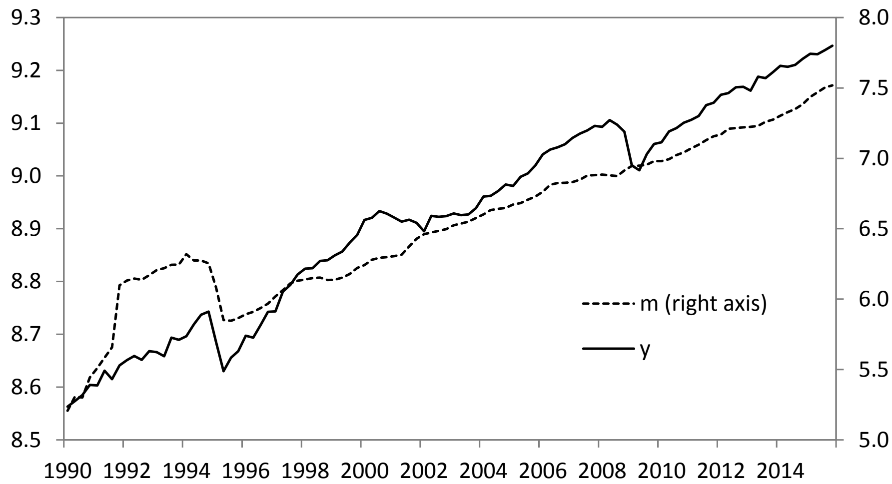


Figure 2: Domestic and Foreign Interest Rates and Depreciation Rate

Panel a) Real M1 and Real GDP



Panel b)

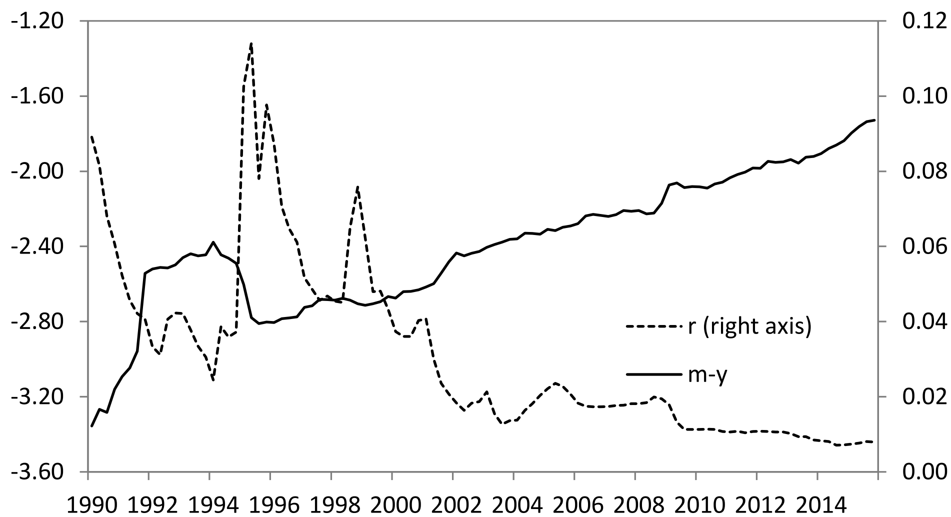


Figure 3: Real M1, Real GDP and Interest Rate

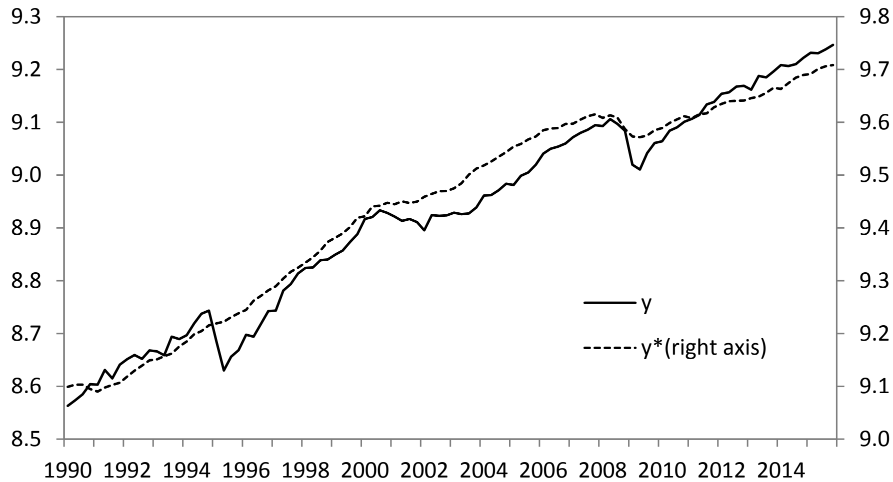
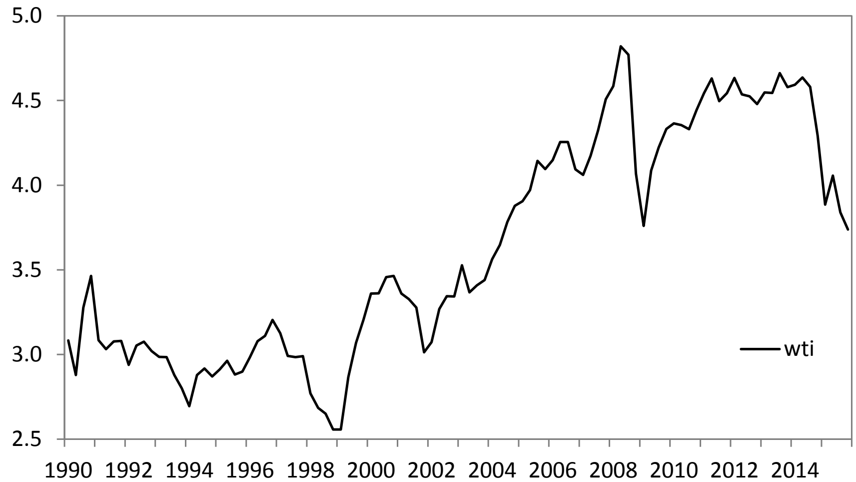


Figure 4: Domestic and Foreign Real GDP

Panel a) Oil Price



Panel b) Oil price, first difference

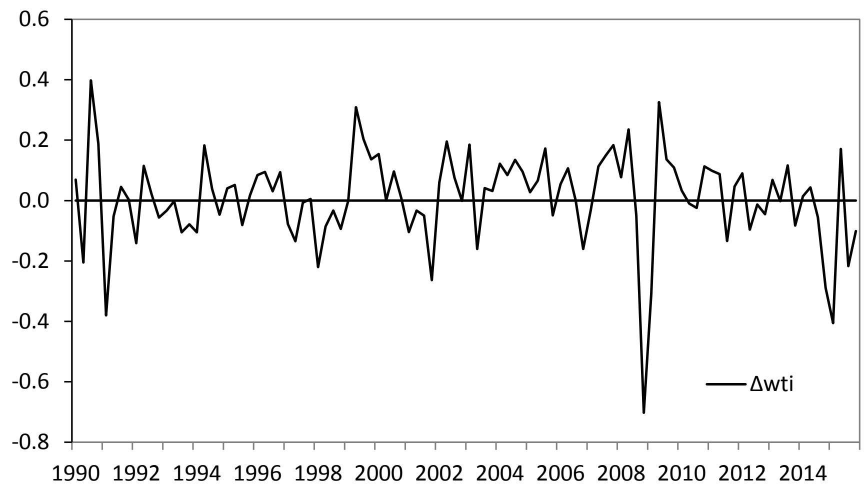
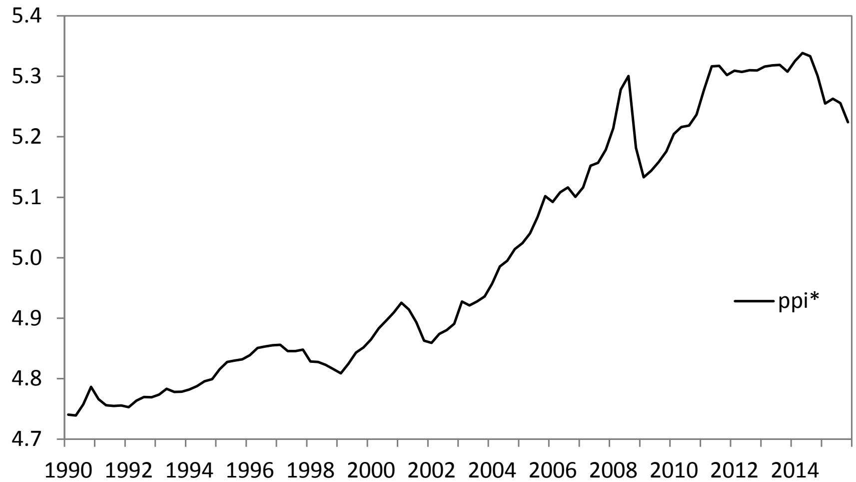


Figure 5: Oil price

Panel a) U.S. Producer Price Index



Panel b) U.S. Producer Price Index, first difference

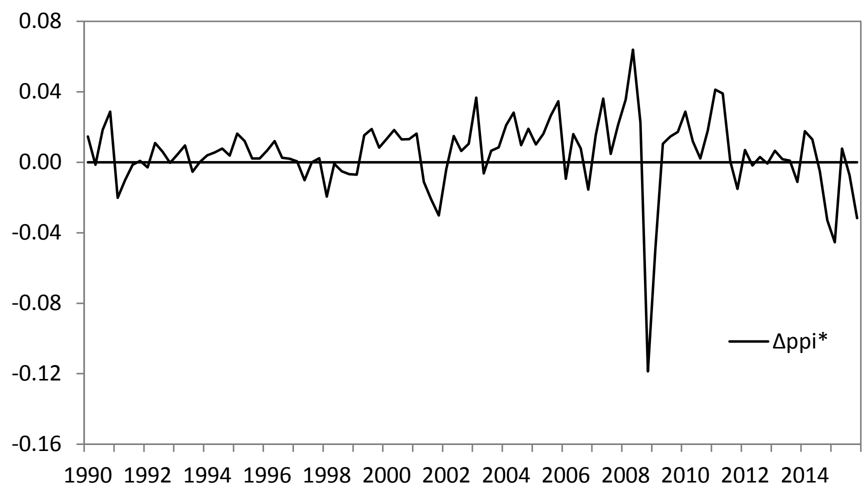
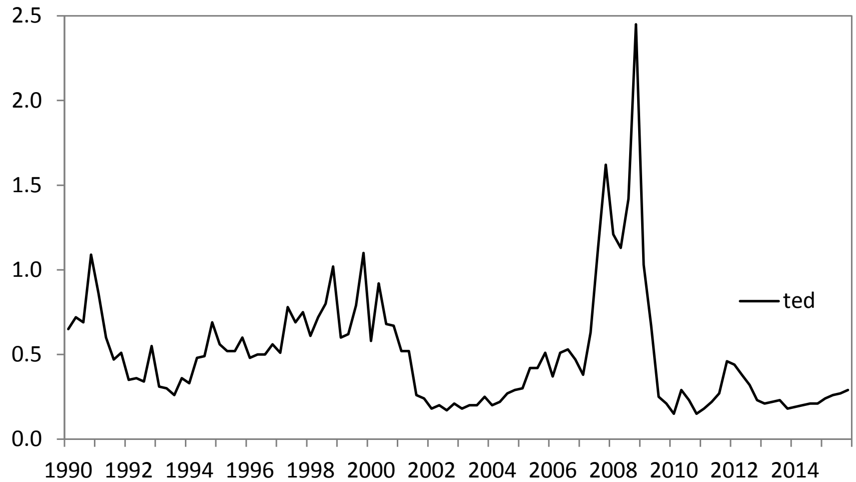


Figure 6: U.S. Producer Price Index

Panel a) TED Spread



Panel b) TED Spread, first difference

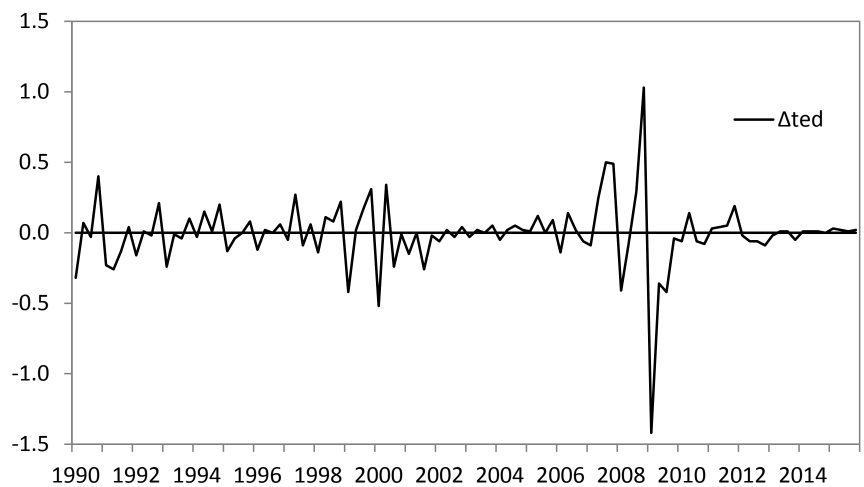


Figure 7: TED Spread

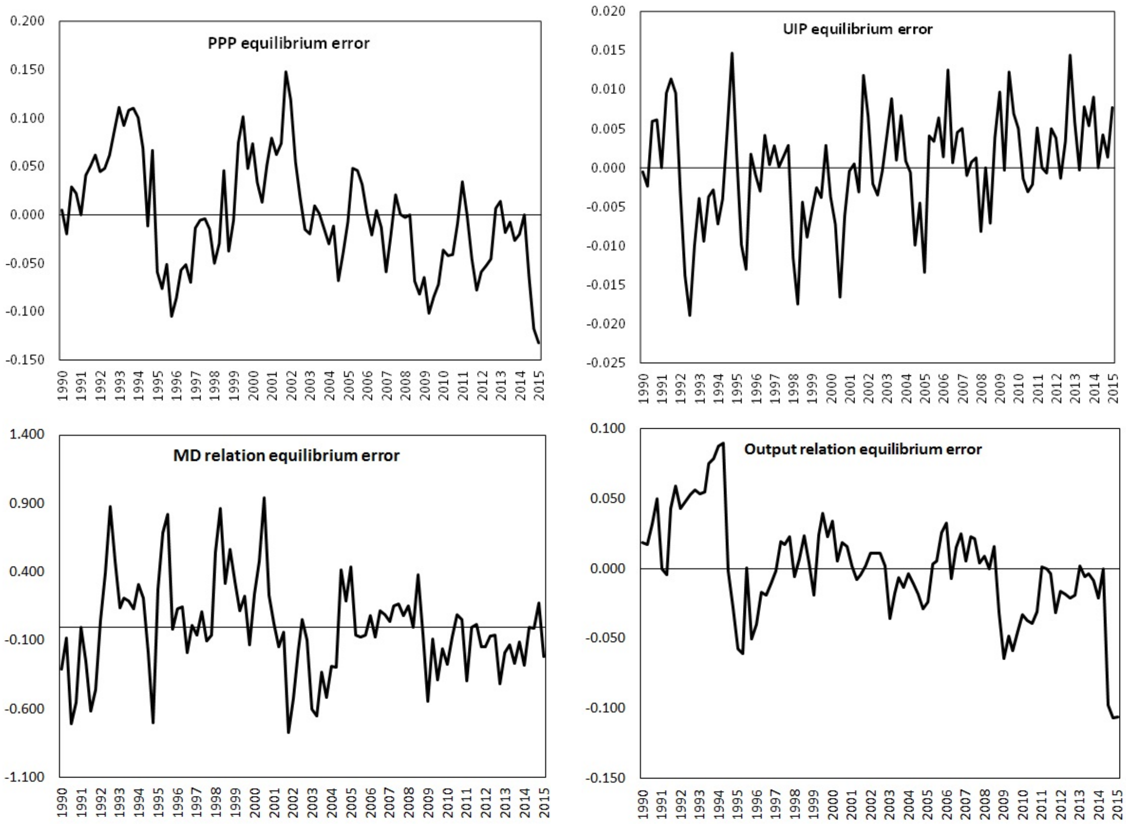
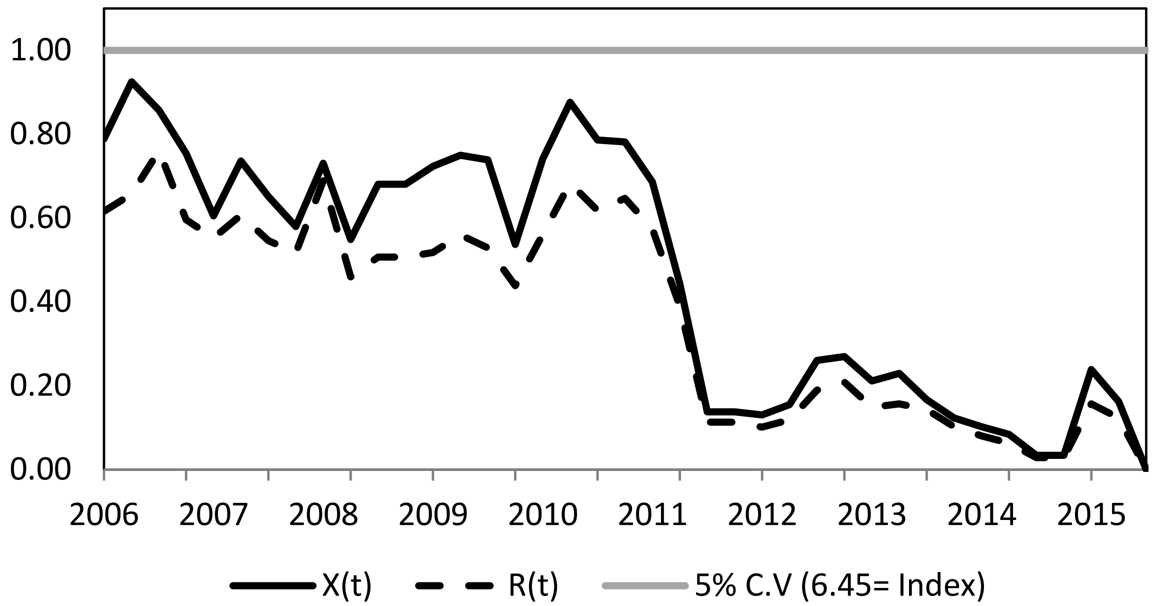


Figure 8: Cointegrating Relations

Panel a) Max test of beta constancy



Panel b) Test of beta = "Known beta"

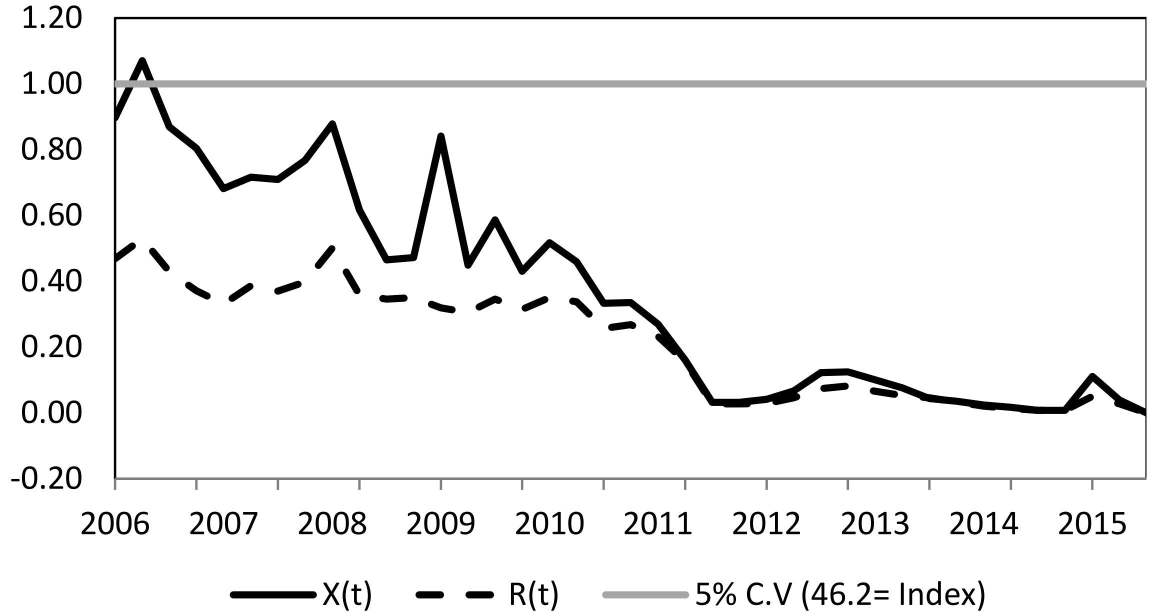
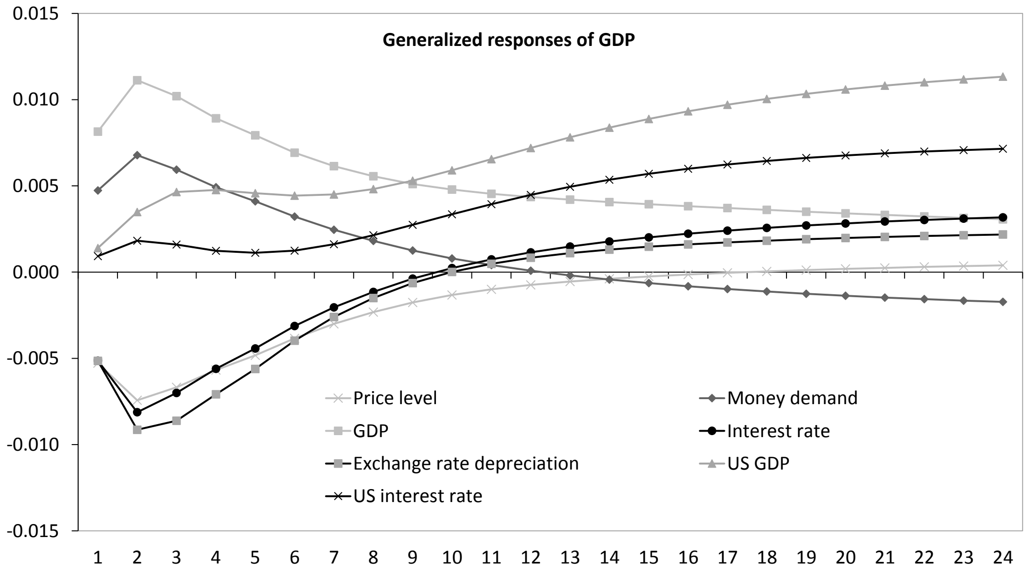


Figure 9: Test of Beta

Panel a)



Panel b)

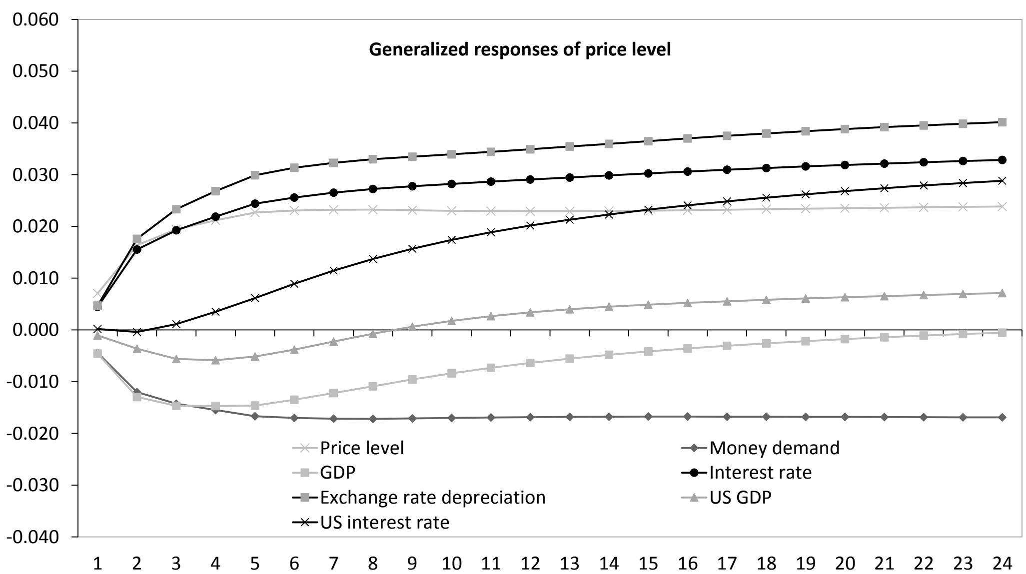
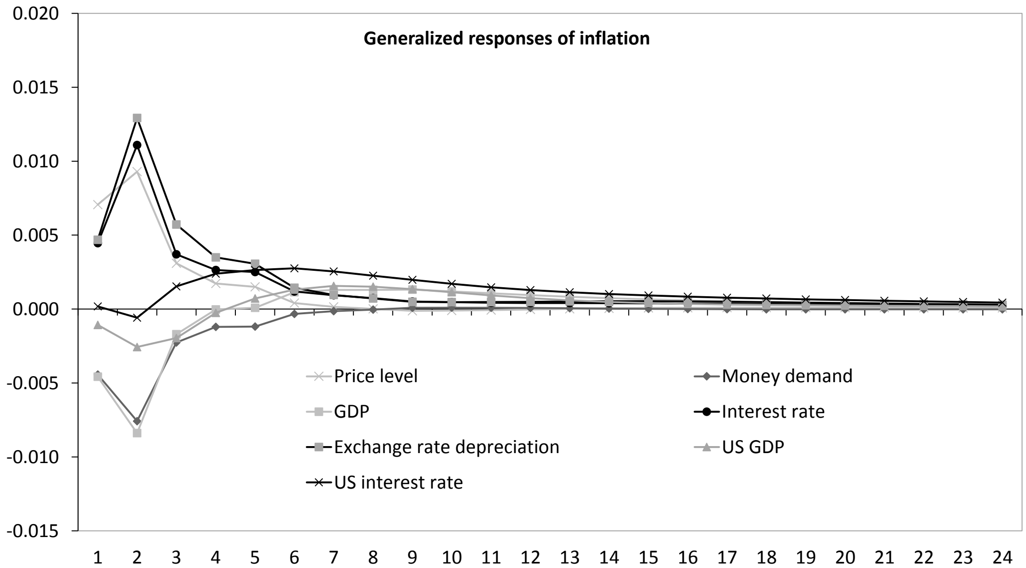


Figure 10: Generalized Impulse Response Functions of GDP and domestic price level

Panel a)



Panel b)

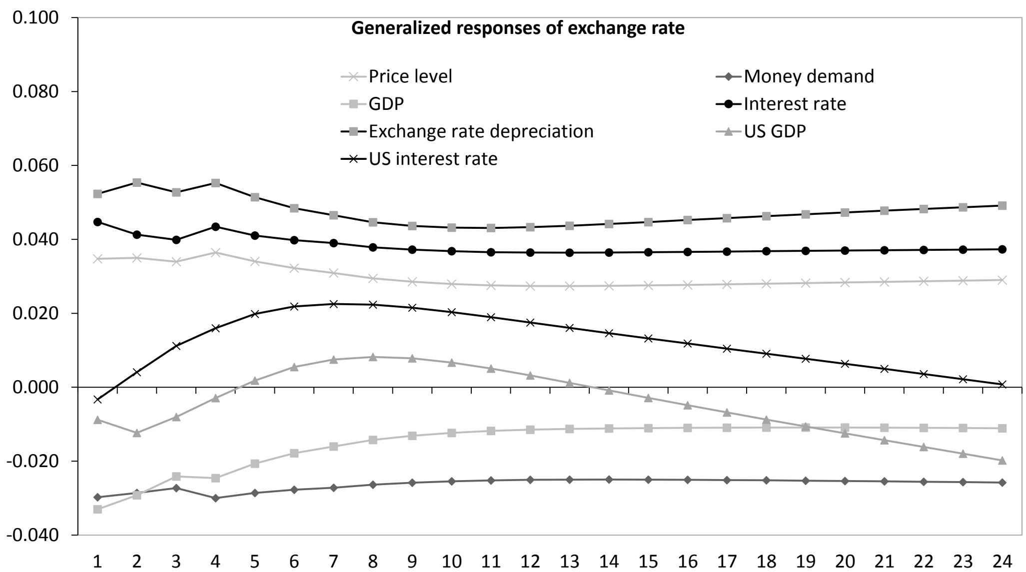
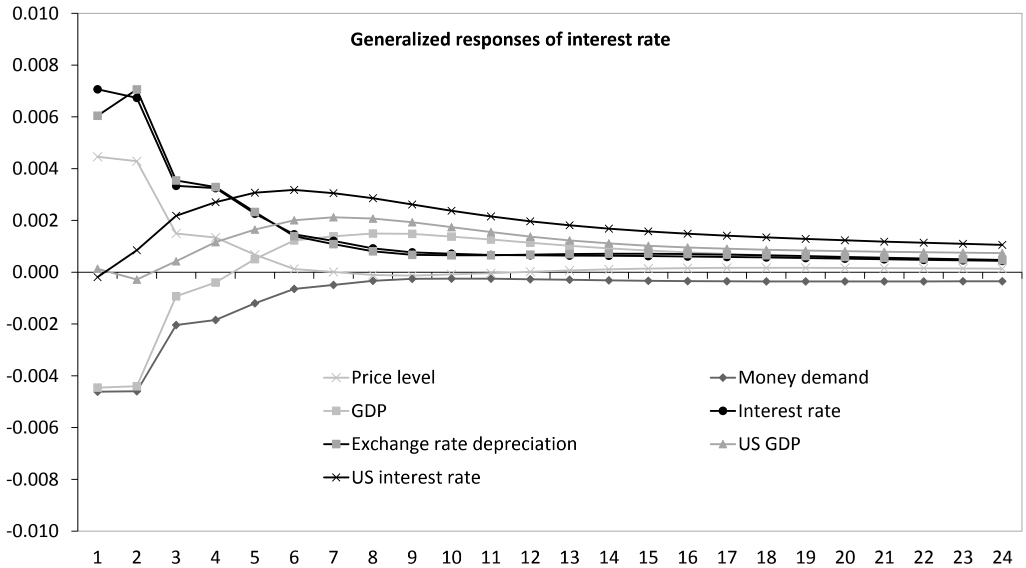


Figure 11: Generalized Impulse Response Functions of inflation and the exchange rate

Panel a)



Panel b)

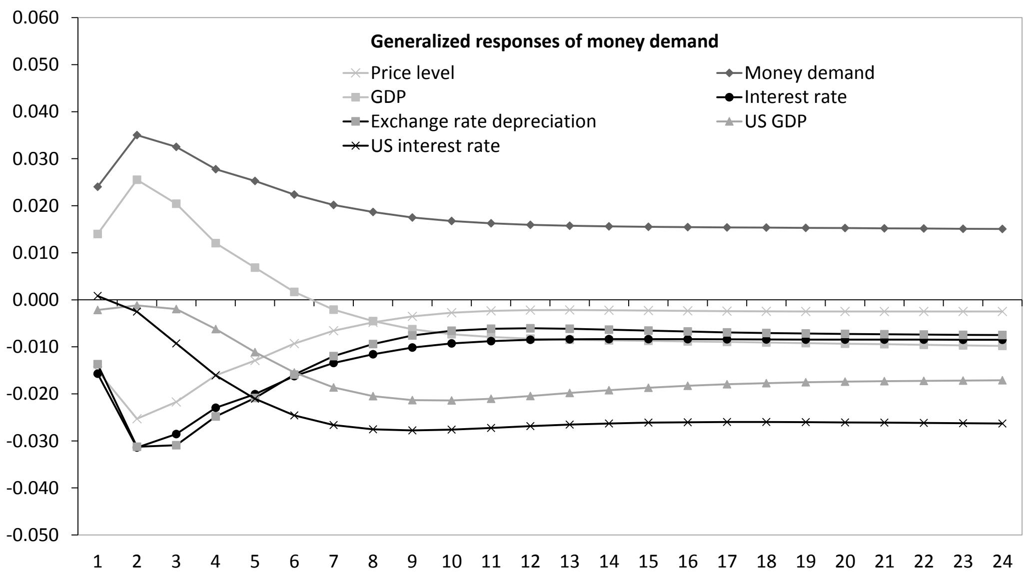


Figure 12: Generalized Impulse Response Functions of the interest rate and the money demand

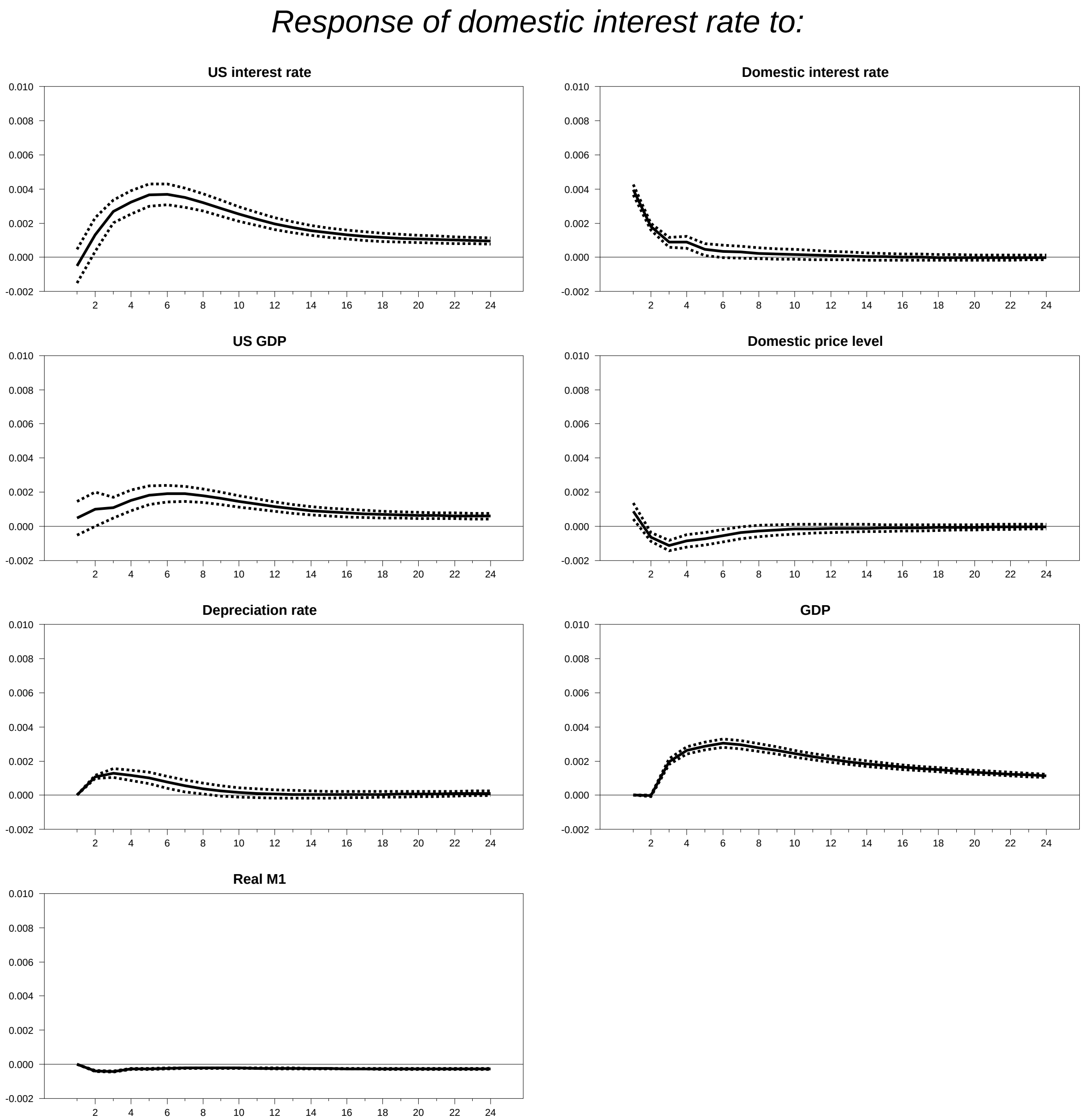


Figure 13: Impulse Response Functions from a Monetary Policy Rule

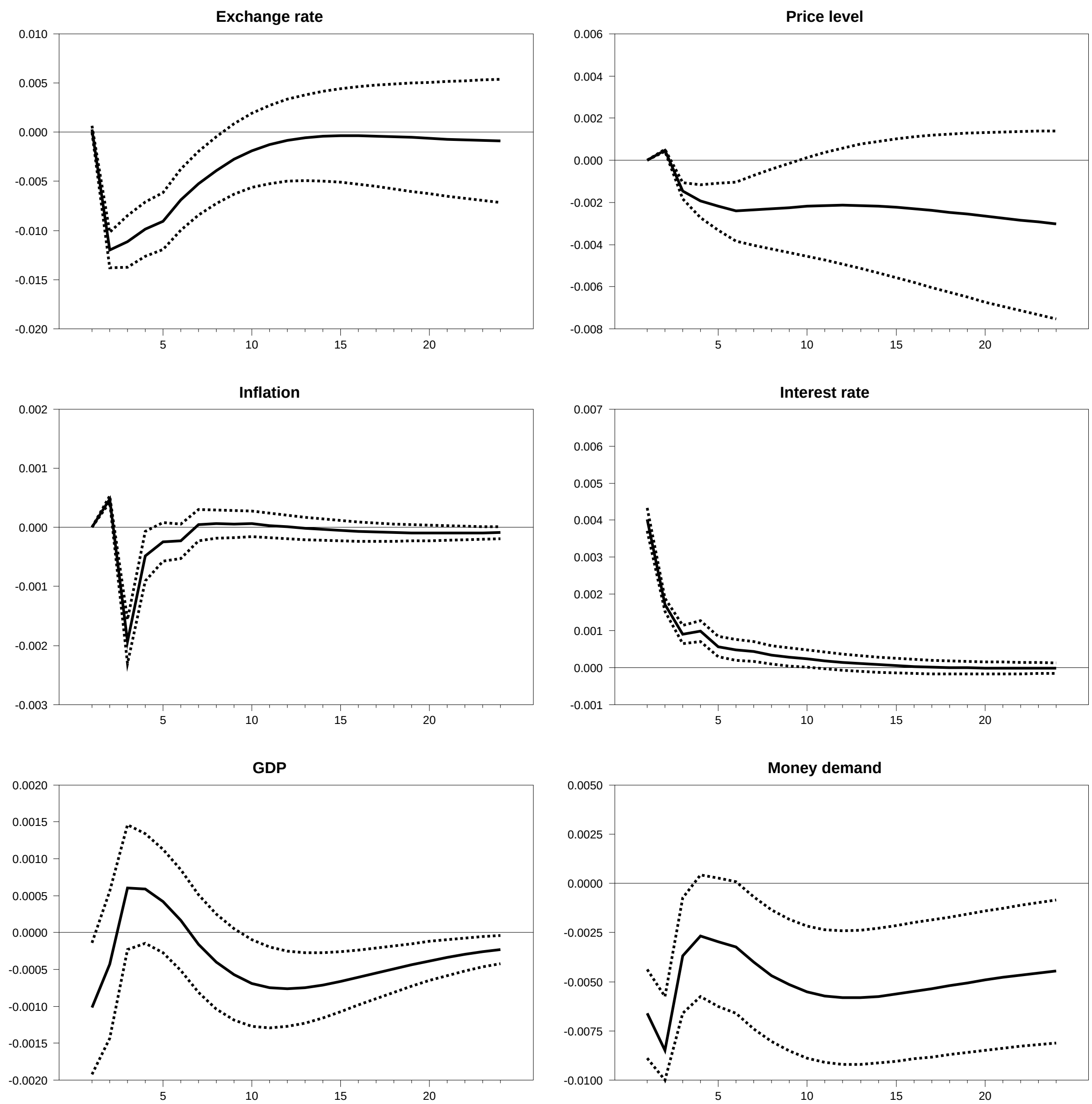


Figure 14: Impulse Response Functions to a Monetary Policy Shock

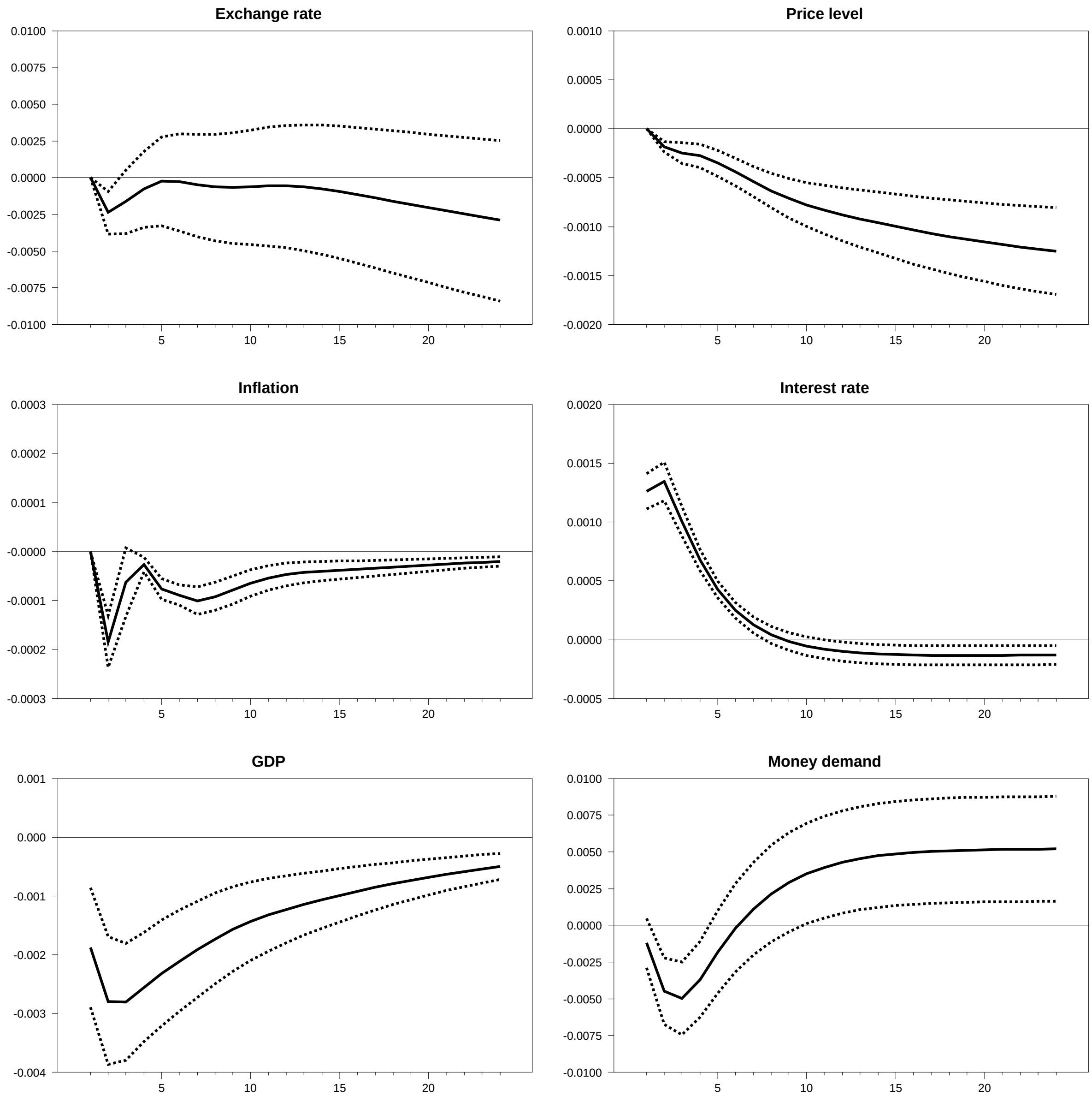


Figure 15: Impulse Response Functions to a Monetary Policy Shock. Sample 2001-2015.