Exchange Rate Pass-Through to Domestic Prices in Uganda: Evidence from a Structural Vector Auto-Regression (SVAR)

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Abstract

This paper examines the degree of exchange rate pass through to inflation in Uganda with quarterly data over the period 1999Q3 to 2012Q2 using a triangulation of well specified Vector Error Correction (VEC) and Structural Vector Auto-Regression (SVAR) models. The findings show strong and significant association between the exchange rate movements and inflation in Uganda, and that the pass-through to domestic inflation, although incomplete, is modest and persistent with a dynamic exchange rate pass-through elasticity of 0.48. This suggests that exchange rate movements remain a potentially important source of inflation in Uganda. Using variance decomposition, it is found that exchange rate shocks have a modest contribution to inflation variance, although inflation is mainly driven by own shocks especially at shorter horizons. The policy implication arising from these findings is that the monetary authority must be vigilant at exchange rate movements and focus on exchange rate interventions which stem inflation pressure from the external sector.

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* The views expressed in this paper are those of the author and do not in any way
  represent the official position of the Bank of Uganda.

1 Introduction

One of the most challenging problems in the conduct of monetary policy in developing countries, especially small open economies like Uganda is exchange rate pass-through (ERPT hereafter). The expression 'ERPT' is generally used to refer to the change in local currency domestic prices resulting from 1 percent change in the exchange rate. Due to its importance in international finance and monetary policy, the extent and timing of ERPT is an issue of interest for monetary policymakers as ERPT is a key ingredient of monetary policy and forecasting models of prices. In its conduct of monetary policy, the central bank requires an understanding of the transmission mechanism of monetary policy to be able to respond adequately to different shocks. In a small open economy such as Uganda, one of such shocks is the exchange rate, which, in addition to the standard aggregate demand channel, provides an important transmission channel for monetary policy. Moreover, as noted in IMF (2006), ERPT also has implications for external adjustment, i.e. the larger the ERPT, the larger will be the response of trade balance to nominal exchange rate volatilities.

As the empirical knowledge of the ERPT matters for the conduct of monetary policy, the subject has spawned many studies through the years. A number of empirical studies suggest the response of consumer prices to exchange rate changes in Sub-Saharan Africa (SSA) is low, and in some cases even zero. For instance, Anguyo (2008) using vector error correction model (VECM) found that the ERPT to inflation in Uganda is low, a finding that is consistent with a number of other studies. For example, both Mwase (2006) and Nkunde (2006) using a structural vector auto-regression (SVAR hereafter) find low ERPT for Tanzania. Similarly, for Ghana, two studies, Frimpong and Adam (2010), and Devereux and Yetman (2003), the former based on vector auto-regression (VAR) models and the latter a single equation approach find low ERPT. Chaoudhri and Hakura (2001) also report low pass-through for a number of SSA countries (Ghana, South Africa, Zimbabwe), while for Tunisia and Ethiopia, the pass-through is zero. According to this side of the story, it is now dangerously close to being elevated to a stylized fact that the ERPT for developing countries is low or where evidence of 'disconnect' between exchange rates and prices exist, it either implies a greater degree of insulation or greater effectiveness of monetary policy.

On the other hand Sanusi (2010), using SVAR model finds substantially large, although incomplete pass-through for Ghana, with a dynamic pass-through elasticity of 0.79. Chaoudhri and Hakura (2001) also find modest pass-through elasticity for Kenya, Cameroon and Zambia. The conflicting findings of empirical studies on the size of ERPT call for further studies (especially considering the recent change in the macroeconomic environment following the 2008 financial and economic crisis) in developing countries. This paper estimates the ERPT in Uganda using a triangulation of VECM and a SVAR approaches to track pass-through from exchange rate fluctuations to each stage of the distribution chain in a simple integrated framework.

The empirical results show robust evidence of a positive long-run correlation between the degree of ERPT and inflation. The paper also finds that the ERPT, measured by impulse response functions (IRF), is modest and persistent, although incomplete, with a dynamic exchange rate pass-through elasticity of 0.48. Consistent with the IRFs, variance decomposition reveals that exchange rate shocks have a modest contribution to inflation variance, but inflation 'own' shocks dominate the volatility of inflation. Our finding is broadly consistent with the findings for some SSA countries, but substantially larger than the earlier estimates on Uganda. We argue that the ERPT is modest if computed rightly, although the recent fluctuations resulting from the recent economic turmoil and the nascent recovery from it may have contributed to the measured ERPT to inflation in Uganda.

The rest of the paper is structured as follows. Section 2 gives an over view of nominal exchange rate and inflation developments in Uganda while the literature review is drawn in Section 3. The econometric methodology is discussed in Section 4. Empirical results are given in Section 5 while the conclusions and policy implications are drawn in Section 6.

2 Preliminary Analysis of Exchange and inflation

For open-economies, inflation comes from both domestic factors (internal pressure) and also overseas factors (external pressure). The sources of external factors are the increase in the world commodity prices or exchange rate fluctuation.

In a floating exchange rate regime, one of the channels through which exchange rate movements affect the inflation rate, is in terms of the interplay of the aggregate demand (AD) and aggregate supply (AS). For instance, in terms of aggregate supply, depreciation (devaluation) of domestic currency can affect the price level directly through imported goods that domestic consumers buy. However, this condition occurs if the country is the recipient countries of international prices (international price taker). Non direct influence from the depreciation (devaluation) of currency against the price level of a country can be seen from the price of capital goods (intermediate goods) imported by the manufacturer as an input. The weakening of exchange rate will make the price of inputs to be more expensive, thus contributing to a higher cost of production. Manufacturers will certainly increase the cost to the price of goods that will be paid by consumers. As a result, the price level aggregate in the country increases or if it continues it will cause inflation.
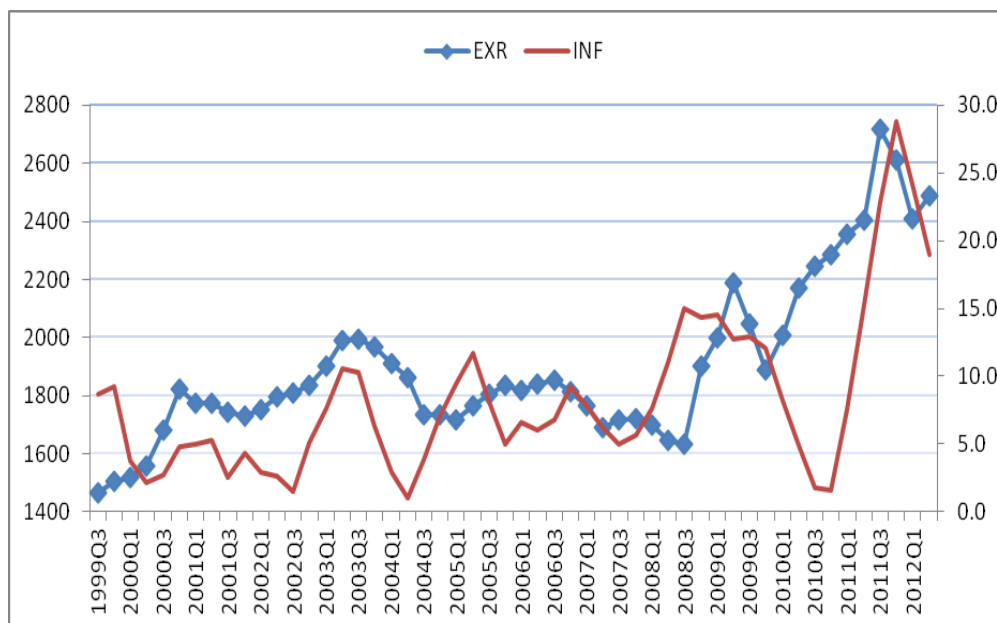
Notes: United States Dollar (USD): Uganda shillings (Ushs) nominal exchange rate
        and inflation (in percentage) on the left and right vertical axis respectively.

Figure 1: Nominal Exchange Rate (Ugandan Shillings per US Dollar) and
           Inflation Developments

        Figure 1 depicts the relationship between the nominal exchange rate (USD:
Ushs) (where an increase in the exchange rate means depreciation) on the primary
axis and inflation on the secondary axis. From the figure, on the whole the
Ugandan shilling has depreciated, albeit at different rates. For instance, the
shillings on average depreciated at a rate of about 5.2 percent annually between
1996 and 2007, but this rose to about 8.0 percent during and after the 2008
financial crisis. Similarly, the evolution of the inflation rate has closely mimicked
the exchange rate developments. For instance, the inflation rate averaged about 5
percent between 1996 and 2007, corresponding to the period with relatively stable
exchange rate. However, the average inflation rate picked up thereafter, reaching
highs of 28% in 2011.

3 Literature Review

ERPT is generally regarded as the change in local currency domestic prices resulting from 1 percent change in the exchange rate. As in Sanusi (2010), the general literature distinguishes between direct and indirect channels through which changes in the exchange rate may be transmitted to consumer prices. The direct channel of movements in the exchange rates on domestic prices is through prices of imported consumer goods or through domestically produced goods priced in foreign currency. While the indirect channel is through prices of imported intermediate goods as changes in the exchange rate may influence costs of production (see Sahminan, 2002 cited in Sanusi, 2010, p.28).

Empirical studies investigating the magnitude of the exchange rate pass-through are abound, albeit with much focus on industrialised countries, i.e. the Euro area, the United States and Japan. Surveys and discussions of the literature on the exchange rate pass-through are provided in Goldberg and Knetter (1997), Menon (1995) and many others, including empirical studies such as McCarthy (2000), Gagnon and Ihrig (2001), Campa and Goldberg (2001), Choudhri and Hakura (2006) and Ito and Sato, (2007) among others. In terms of estimation approaches, both the popular ordinary least squares (OLS) and vector autoregressive (VAR) approaches are used. The collective evidence can be summarized as follows. First, the degree and dynamics of ERPT is incomplete, and the pass-through to import prices tends to be higher in both magnitude and speed than that to consumer prices. Secondly, estimates across countries and estimates across studies for a particular country are significantly different and at times conflicting, i.e. the evidence is inconclusive. Thirdly, there is a general decline in the degree of pass-through in the 1990s, majorly attributed to the low inflation environment achieved in most industrialized countries.

Studies on the pass-through in developing economies are somewhat limited, although the few existing works tend to show similar results to those of industrialized countries. For example, Chaoudhri and Hakura (2001) found zero elasticity of exchange rate pass–through to inflation in Bahrain, Singapore, Canada and Finland. With regard to sub-Saharan Africa (SSA) countries, Kiptui *et al*. (2005) using a vector error correction approach find incomplete pass-through in Kenya during the period 1972-2002. In particular, their results show that an exchange rate shock leads to a sharp increase in inflation that evens out after four quarters, with exchange rate accounting for 46 percent of inflation variance. Similarly, Chaoudhri and Hakura (2001) found exchange rate pass–through of 0.09 for Kenya, including 0.14 for Ghana, 0.02 for South Africa, 0.06 for Zimbabwe, 0.16 for Burkina Faso and zero for Tunisia and Ethiopia.

Using quarterly data for the period 1990-2006 and an SVAR model, Mwase (2006) quantifies the exchange rate pass-through for Tanzania: first for the full sample; and second for two sub-periods, i.e. periods prior to and after 1995. He finds pass-through elasticity of 0.028 in the full-sample, and 0.087 in the period

before 1995, which however declines to 0.023 after 1995. Overall, he finds that the exchange rate pass-through has declined despite depreciation of the currency. In another study on Tanzania, Nkunde (2006) uses the same SVAR framework for the period 1990-2005. He finds an incomplete exchange rate pass-through to inflation, where a 10% depreciation was associated with a 0.05% increase in inflation after a two quarter lag. Sanusi (2010) also uses an SVAR model applied on quarterly observations for the period 1983Q3 to 2006Q3 to estimate the pass-through effects of exchange rate changes to consumer prices for the Ghanaian economy. He finds that the pass-through, although incomplete, is substantially large, with a dynamic pass-through elasticity of 0.79. For the same country (Ghana), Frimpong and Adam (2010) uses vector auto-regression (VAR) models applied on monthly data for the period 1990-2009. They find incomplete, decreasing and low exchange rate pass-through. To be exact, they find that a 1% depreciation is associated with a 0.025% increase in inflation after a quarter after initial impact, increasing to 0.09% after eight quarters and deceasing sluggishly to 0.07% after twelve quarters of its initial impact.

The only study to our knowledge, Anguyo (2008), uses vector error correction model (VECM) to examine the effect of exchange rate changes on consumer prices in Uganda. Using monthly data for the period 1996M7 – 2007M5, he finds low, significant and persistent exchange rate pass-through to inflation. Specifically, he found that a 1% exchange rate depreciation results in a 0.056% increase in inflation, in the second month (*ibid*: 91). In common with the findings of low exchange rate pass-through in the literature (see for example, Stulz, 2006; Devereux and Yetman, 2002; Taylor, 2001; Chaoudhri and Hakura, 2001; Devereux and Engel, 2001; among others), the author attributes the results to among other factors, low inflation environment and fairly stable exchange rate.

Whereas as in figure (1), Uganda could be characterized as having low inflation environment and fairly stable exchange rate up to 2007, the trend appears to have changed following the 2008 financial and economic crisis in several emerging markets. While the Uganda shilling depreciated on average at a rate of 5.2 percent (annualized) between 1996 and 2007, the rate of depreciation rose to 8.0 percent during and after the 2008 financial crisis. Over the same period, up to 2007, inflation averaged 5%, but picked up thereafter, reaching highs of 28% in 2011. This recent change in the macroeconomic environment calls for renewed investigation of the ERPT in Uganda. Even more, we argue that there is something profoundly wrong with the way the exchange rate pass-through has been computed. This study, as in Sanusi (2010), re-examines the estimation of exchange rate pass-through using a triangulation of VECM and a SVAR approaches. VECM is a good starting point because then, the assumption of weak exogeneity and endogeneity of variables, essential for ordering of variables in SVAR is not assumed (as in most studies) but tested.

4 Econometric Model

4.1   Vector Autoregressive Framework

In a Vector Autoregressive framework, all variables are endogenous. As a reduced form representation of a large class of dynamic structural models (Hamilton 1994: 326-7), VAR offers both empirical tractability and a link between data and theory using minimal assumptions about the underlying structure of the economy. In the current application where the macro variables are likely to be non-stationary, it is convenient to couch the empirical analysis in a VAR framework in its unrestricted error correction representation of the form

$$\Delta \mathbf{z_t} = \Pi \mathbf{z_{t-1}} + \sum_{i=1}^{k-1} \Gamma_1 \Delta \mathbf{z_{t-i}} + \Phi \mathbf{d_t} + \varepsilon_t \tag{1}$$

Where $\mathbf{z_t}$ is a vector of endogenous variables, each of the $(n \times n)$ matrices Γ_i and Π comprise coefficients to be estimated by Johansens's (1988) maximum likelihood procedure using a $(t = 1,...,T)$ sample of data, $i = 1,...,k-1$ is the number of lags included in the system, $\mathbf{d_t}$ is a vector of deterministic terms (constants, linear trends, 'spike' and/or intervention dummies), Δ is a first difference operator and ε_t is a vector of structural innovations, with zero mean, i.e. $E(\varepsilon_t) = 0$, a time-invariant positive definite covariance matrix Σ, and are serially uncorrelated, i.e. $E(\varepsilon_t \varepsilon'_{t-k}) = 0$ for $k \neq 0$. Of paramount interest in the VAR analysis is the Π vector which represents a matrix of long-run coefficients, defined as a multiple of two $(n \times r)$ vectors, α and β' i.e. $\Pi = \alpha \beta'$.

The β' vector represents the co-integrating vectors that quantify the 'long-run' (or equilibrium) relation(s) amongst the variables in the system while α is a vector of loadings of the cointegrating vectors, denoting the speed of adjustment from disequilibrium. In addition, zero restrictions on α reveals weak exogeneity status of the corresponding variable in the system. Finding the existence of cointegration is the same as finding the rank (r) of the Π matrix. If it has full rank, the rank $r = n$, and we have n cointegrating relationships, that is, all variables are potentially I(0).

The first step to estimating the above model is the choice of variables that should be included. Following McCarthy (2000), Hahn (2003), Ito and Sato (2006) and Anguyo (2008) among others, we set up a 5-variable VAR model, constituting the output gap (y_gap_t), nominal exchange rate (exr_t), core CPI ($CoreCPI_t$), oil price index (oil_t) and the 91 day Treasury bill rate (r_t). These include two indicators of aggregate demand, i.e. the core CPI ($CoreCPI_t$) and the output gap (y_gap). The y_gap is defined as the percentage deviation of actual output from the trend or equilibrium level of output, where trend output is generated from quarterly GDP using the Hodrick-Prescott (HP) filter. Accordingly, a positive number indicates positive 'excess demand' position with output above its trend level.

Thus, we consider the 5 x 1 vector $\mathbf{z_t} = \left(\mathbf{y\_gap_t}, \mathbf{exr_t}, \mathbf{oil_t}, \mathbf{corecpi_t}, \mathbf{r_t}\right)'$ with all variables in natural logs.[2] This choice of structural ordering of the variables in the VAR need not be interpreted as an attempt to provide a strict identification of structural shocks, but to analyze the long-run impact of exchange rates on domestic prices. In addition to these five endogenous variables, we control for financial crisis in the last quarter of 2008, and increases in world commodity prices beginning the third quarter of 2011. In what follows, we now turn to the identification of exchange rate shocks and impulse responses.

4.2   Structural VAR (SVAR)

One of the main shortcomings of the unrestricted VAR (UVAR) approach is the difficulty of interpreting the impulse responses. This is because the choice of the Choleski decomposition in the UVAR is not unique given the number of alternative sets of orthogonalised impulse responses which can be obtained from any estimated VAR model. Sim's (1980) own approach of circumventing this problem by choosing an orthogonalisation – typically imposing causal ordering on the VAR has not been fully accepted in the literature. In the absence of such restrictions, the orthogonalised impulse responses are difficult to interpret, so that the estimated model gives few meaningful insights into the economic system that it represents. The SVAR approach builds on Sims' approach but attempts to identify the impulse responses by imposing a priori restrictions on the covariance matrix of the structural errors and/or on long-run impulse responses themselves.

The SVAR permits contemporaneous relationships between the elements of a vector of endogenous variables. In this way, we can model dynamic and contemporaneous endogenity between variables. In matrix form following Hamilton (1994), the SVAR can be written as:

$$\beta_0 x_t = k + \beta_1 x_{t-1} + \beta_2 x_{t-2} + \cdots + \beta_p x_{t-p} + \mu_t \tag{2}$$

Where x_t is an endogenous variable, ε_t is a white noise error term. The white noise errors means that the structural disturbances are serially uncorrelated such that $E|\mu_t \mu_t'| = D$, where D is a diagonal matrix. Pre-multipying equation (2) by β_0^{-1} gives us the reduced form (VAR) of the dynamic structural model:

$$x_t = \beta_0^{-1} \left(k + \beta_1 x_{t-1} + \beta_2 x_{t-2} + \cdots + \beta_p x_{t-p} + \mu_t \right)$$

$$x_t = c + \varphi_1 x_{t-1} + \varphi_2 x_{t-2} + \cdots + \varphi_p x_{t-p} + \varepsilon_t \tag{3}$$

where $\varphi_s = \beta_0^{-1} \beta_s$, $(s = 1, 2, \ldots, p)$, $c = \beta_0^{-1} k$ and $\varepsilon_t = \beta_0^{-1} \mu_t$.

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[2] Note that natural log transformation was applied on actual GDP before generating the trend and so is the y_gap.

The variance-covariance matrix is given by:

$$E\left|\varepsilon_t\varepsilon_t'\right| = \beta_0^{-1}E\left|\mu_t\mu_t'\right|\left(\beta_0^{-1}\right)' = \beta_0^{-1}D\left(\beta_0^{-1}\right)' = \Omega$$

To generate the structural shocks, we use a Cholesky decomposition of the variance-covariance matrix of the reduced form VAR residuals $\hat{\Omega}$. Since the estimation of the SVAR model has k^2 more parameters than the VAR, in order to find a unique solution we require both the order condition and the rank condition to be satisfied. The order condition requires that the number of parameters in the matrices β_0 and D should be less than the number of free parameters in the matrix Ω. Since Ω is a symmetric matrix, then the number of free parameters of the matrix Ω is defined by $(k(k+1)/2$, where k is the number of endogenous variables included in the system.

Assuming that D is a diagonal matrix, then β_0 can have no more free parameters than: $k(k-1)/2$. We can impose two different restrictions on matrix β_0. The first is the normalisation restriction that aims to assign the value of 1 to variables $x_{t,i}$ in each of the ith equation. And the second is the exclusion restriction that aims to assign zero to some variables in the equation (especially contemporaneous relations). These restrictions are defined by the theoretical model.

The rank condition for identification of a structural VAR is more complex. This requires that the columns of the matrix J be linearly independent; which Hamilton (1994), defines as:

$$J = \left| \frac{\partial vech(\Omega)}{\partial \theta_B'} \quad \frac{\partial vech(\Omega)}{\partial \theta_D'} \right|$$

The operator $vech()$ picks out the distinct element Ω, making the condition sufficient for local identification. Imposing the restrictions suggested by the theoretical model, we construct the matrix B_0 and find the relationship between the error terms of the reduced form and the structural disturbances: $\varepsilon_t = \beta_0^{-1}\mu_t$

$$\begin{bmatrix} \varepsilon_t^{oil} \\ \varepsilon_t^{y-gap} \\ \varepsilon_t^{r} \\ \varepsilon_t^{exr} \\ \varepsilon_t^{corecpi} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 & 0 \\ a_{31} & a_{32} & 1 & 0 & 0 \\ a_{41} & a_{42} & a_{43} & 1 & 0 \\ a_{51} & a_{52} & a_{53} & a_{54} & 1 \end{bmatrix} \begin{bmatrix} u_t^{oil} \\ u_t^{y-gap} \\ u_t^{r} \\ u_t^{exr} \\ u_t^{corecpi} \end{bmatrix} \tag{4}$$

Where u_t^{oil} denotes the oil price (supply) shock, u_t^{gdp} is the output gap (demand) shock; u_t^{r} is the monetary shock; u_t^{exc} is the nominal exchange rate (external)

shock; and u_t^{cpi} is the inflation shock. Thus, according to our theoretical model, the matrix B has 10 free parameters to estimate, which are exactly the same we require for the order condition to be satisfied. This recursive identification scheme is based on Ito and Sato (2007), Hahn (2003) and McCarthy (2000), and implies that the identified shocks contemporaneously affect their corresponding variables and those variables that are ordered at a later stage, but have no impact on those that are ordered before.

In the resulting B matrix in the system in (4), it is reasonable to order the most exogenous variables first. As we will show in Table 2, the pump price of crude oil is exogenous to the domestic economy, so oil price shocks are modelled as independent of shocks to other variables in the system. Given the set up in (4), this amounts to a set of four restrictions, since the restriction imposes zero on the 2nd, 3rd, 4th and 5th elements of its first row. In the second row, we have three additional restrictions based on the assumption that shocks to the output gap are influenced by shocks to the pump price of crude oil and are independent of shocks to all other variables in the system. That is, the monetary policy rate, exchange rate and inflation are assumed to have no contemporaneous effects on the output gap. This is equivalent to imposing zero restrictions on the 3rd, 4th and 5th elements of the second row in the matrix. Shocks to the monetary policy rate are assumed to be influenced by shocks to the pump price of crude oil and shocks to the output gap. Exchange rate and inflation are assumed to have no contemporaneous effects on the monetary policy rate. This assumption adds two more restrictions, and is equivalent to imposing zero restrictions on the 4th and 5th elements of the third row in the matrix. We also assume that shocks to the exchange rate are influenced by shocks to the pump price of crude oil, shocks to the output gap and shocks to the monetary policy rate. In this case, inflation is assumed to have no contemporaneous effect on exchange rate. This assumption forms the 10th restriction, and therefore meets the minimum requirement for the order condition to be satisfied. Finally, shocks to domestic inflation are ordered last because they are assumed to be influenced by shocks to all the variables in the system.

Under this structure, the model is estimated as a SVAR using a Cholesky decomposition. The impulse response of CPI inflation to the orthogonalised shocks of exchange rate movements then provide estimates of the effect of exchange rate on domestic inflation. In addition, variance decompositions of CPI inflation enable us to determine the importance of each of the variables in the system in domestic price fluctuations.

5 Data and Empirical Results

5.1 The Data

All data on the five endogenous variables to be used in the analysis are from Bank of Uganda data base, and cover the period 1999Q3-2012Q2. Graphs of the

level and differences of the data are given in the Appendix **1**. All of them seem to follow the same pattern. Save for the *y_ gap*, they are definitely not stationary as they are not mean-reverting in levels. However, in differences they all seem to be mean-reverting and therefore stationary. Hence, all variables seem to be I(1), except *y_ gap*, which seems to be I(0). After visual inspection, the series are formally tested for the order of integration or non-stationarity using the Augmented Dickey Fuller (ADF) unit root test (Dickey and Fuller, 1979). Results of the unit root test are provided in Appendix 2, and indicate that only *y_ gap* is I(0) or stationary, while the rest of the variables are non-stationary. All I(1) variables are I(0) in first differences. From the graphs of the levels, it appears the data contains a linear trend. However, it is highly unlikely that there is a linear trend in the data because it would contradict the assumption of rational behaviour in the financial markets. If it was the case, then it would, to some extent be possible to predict the future interest rates. Hence, the constant term μ_0 will be restricted to lie in the cointegrating space. Furthermore, because our data is high frequency (i.e. quarterly observations), all series, save for nominal exchange rate, are adjusted for seasonal effects.

The unrestricted 5-dimensional model is estimated with a restricted constant. The choice of the lag- length was determined as the minimum number of lags that meets the crucial assumption of time independence of the residuals, based on a Lagrange Multiplier (LM) test. We began with 4 lags. Although both Schwarz and Hannan-Quinn information criteria suggested four lags, with three lags, the LM test could not reject the null hypothesis of no serial correlation in the residuals. Thus, all subsequent VAR analysis was implemented with three lags.

Having determined the appropriate specification of the data generating process, cointegration rank was determined using Johansen's (1988) *trace statistic*.[3] However, this has been shown to have finite sample bias with the implication that it often indicates too many cointegrating relations, i.e. the test is over-sized (Juselius, 2006: 140-2; Cheung and Lai, 1993b; Reimers, 1992). Hence, as shown in Table 1, we also report results for small sample Bartlett correction, which ensures a correct test size (Johansen, 2002). Based on the results, the presence of two equilibrium (stationary) relations, even when corrected for small sample bias among the variables at the conventional 5 percent level of significance cannot be rejected. Furthermore, even the recursive graphs of the Trace-Test Statistics in Appendix 3 also suggest 2 stationary relationships.

However, given the fact that the cointegration rank classifies the eigenvectors into *r* stationary and *p-r* non-stationary directions, every stationary variable included in the model correspondingly increases the number of cointegration equations (Harris and Sollis, 2005, p. 112). In Appendix **2**, *y_gap*

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[3] In the test, the determination of the cointegrating rank, *r* relies on a top-to-bottom sequential procedure. This is asymptotically more correct than the bottom-to-top alternative (i.e. Max-Eigen statistic) [Juselius, 2006: 131-134].

variable is I(0) while the rest of the variables are I(1), implying that potentially, the I(0) variable may have added an additional cointegrating relation in the model. Thus, adjusting the number of cointegration equations for the one I(0) variable leaves only one long-run relation. Even more, economic interpretability of the alpha coefficients (not reported here) supports the possibility of only one cointegrating relationship.

Table 1: I(1)-Analysis

p-r	r	Eig.Value	Trace	Trace*	Frac95	P-Value	P-Value*
5	0	0.700	137.564	106.451	88.554	0.000	0.001
4	1	0.500	79.718	67.920	63.659	0.001	0.020
3	2	0.459	46.430	33.300	42.770	0.020	0.327
2	3	0.216	16.957	9.851	25.731	0.426	0.924
1	4	0.104	5.291	4.550	12.448	0.564	0.666

Notes: Constant/Trend assumption: Restricted constant; Frac95: the 5% critical value of the test of H(r) against H(p). The critical values as well as the *p*-values are approximated using the Γ distribution (Doornix, 1998); '*' is the small sample Bartlett correction.

Source: Authors computation using CATS in RATs by J.G. Dennis, H. Hansen, S. Johansen and K. Juselius, Estima 2012.

5.2   Estimation of the long-run Exchange rate Pass- Through

With a unique relationship among the macrovariables, identification of the long-run relation becomes relatively direct. We normalize the only existing cointegration relation on core CPI to identify a cointegrated relation among the 5-variables in the system. Table 2 reports the long-run parameters, error correction and weak exogeneity tests (which then guide our re-specification of the VECM, conditioning on weakly exogenous variables). Based on the results in the table, weak exogeneity of oil_t cannot be rejected, amounting to the finding that oil_t does not enter the short-run equation determining the rest of the variables. The resulting conditional VECM model are represented by BETA (transposed)* for the long-run parameters and ALPHA (transposed)* for the error correction parameters, and is interpreted below.

Consistent with the Taylor (2000) hypothesis cited in Ca'Zorzi *et al.* (2007: 5), estimates suggest, ceteris paribus, a positive long run correlation of nominal exchange rate with core CPI. Estimates also show positive long-run correlation of y_*gap* and oil_price index with core CPI and a negative association with the monetary policy rate. In the long-run, estimates show that a unit percentage point q-o-q depreciation in the nominal exchange rate leads to a more than proportional

Table 2: Estimates of the Long-run Exchange rate Pass- Through

The Matrices based on 1 cointegrating vector				
BETA (transposed):				
y_gap_t	Exr_t	Oil_t	$CoreCPI_t$	R_t
2.072	1.226	0.108	-1.000	-0.022
(2.401)	(11.036)	(4.476)		(-5.905)
ALPHA (transposed)				
-0.076	-0.323	-0.140	-0.122	24.854
(-2.675)	(-2.954)	(-0.324)	(-2.997)	(3.444)
Test of Weak exogeneity: LR-Test, $\chi^2(1)$				
17.868	5.259	0.154	7.295	4.676
(0.012)	(0.022)	(0.694)	(0.007)	(0.0031)
BETA (transposed)*				
1.965	1.233	0.105	-1.000	**-0.023**
(2.240)	(10.918)	(4.286)		**(-6.054)**
ALPHA (transposed)*				
-0.068	-0.330	0.000	-0.119	24.674
(-1.604)	(-3.687)	(0.000)	(-2.965)	(3.586)

Notes:
i) In parenthesis are *t-ratios* for BETA and ALPHA (transposed) and *P-values* for weak exogeneity tests;
ii) Exclusion of a constant term (executed in CATS in RATs by J.G. Dennis, H. Hansen, S. Johansen and K. Juselius, *Estima* 2012) could not be rejected, so this is excluded from the long-run estimates.
iii) Whilst by construction core CPI inflation ironically excludes fuel prices, implying a zero restriction on oil_t in the model, this operation is not statistically accepted: Chi-square(1) = 2.989 (0.084). This rejection mimics a possible second round effect associated with oil prices on core CPI inflation – thus, oil_t variable is retained in the model
iv) The weakly exogenous variable remain in the long-run model, i.e. the cointegration vector (BETA (transposed)*, although its short-run behaviour is not modelled (For details, see Harris and Sollis, 2005, p.135-7).

increase in domestic inflation (about 1.2%). This is a huge effect, highlighting the vulnerability of the Ugandan economy to changes in the exchange rate. The deviation of output from trend has a significant positive long-run relationship with domestic prices, implying that inflation is very sensitive to aggregate demand. Furthermore, a measure of monetary policy rate that determines changes in short-

term interest rates (the 91-days Treasury bill rate) in response to changes in economic conditions has a significant negative association with domestic inflation. This is in line with the conjecture that the higher the Treasury bill rate, the lower the inflation and vice versa. The speed-of-adjustment is about 12 percent.

Note however that these are partial derivatives (by construction) predicated on the *ceteris paribus* clause (Lütkepohl and Reimers 1992), and have been interpreted in this light. As argued in Lloyd et *al*. (2006), where variables in an economic system are characterised by potentially rich dynamic interaction (as is the case here), inference based on 'everything else held constant' is both of limited value and may give a misleading impression of the short- and long-run estimates. Therefore, since what we want is to actually estimate what might happen to all variables in the system following a perturbation of known size in the exchange rate equation, impulse response analysis, which describes the resulting chain reaction of knock-on and feedback effects as it permeates through the system, provides a tractable and potentially attractive value of the exchange rate pass-through providing no other shocks hit the system thereafter (see Johnston and DiNardo, 1997). This is discussed in the next section.

5.3   Impulse Responses and Variance Decomposition

The results of the exchange rate shock under the identification scheme described in Section 4.2 above are shown in Figure 2. Specifically, Figure 2 shows the impact of a one standard deviation shock, defined as an exogenous, unexpected, temporary depreciation in the exchange rate with a 95 percent confidence level on domestic price inflation[4], output gap, oil price and 91-day Treasury bill rate in period 0. The solid line in each graph is the estimated response while the dashed lines denote a two standard error confidence band around the estimate. Since the data are in first differences of logarithms, the IRFs need to be regarded as measuring a proportional change in the rest of the macrovariables due to one standard innovation (at the initial period) in the exchange rate.

It is clear from the figure that the effect of an exchange rate shock on domestic price is fairly gradual (taking about 4 quarters to reach the full impact) and persistent. Based on the numbers in Table 3, the immediate effect of a structural one standard deviation shock to the exchange rate of 0.041 (or 4.1%) depreciation is about 0.007 (or 0.7%) increase in the domestic price level. This suggests an impact exchange rate pass-through elasticity of 0.16.

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[4] The rest of the IRF of the entire system are shown in Figure 4 in the Appendix.

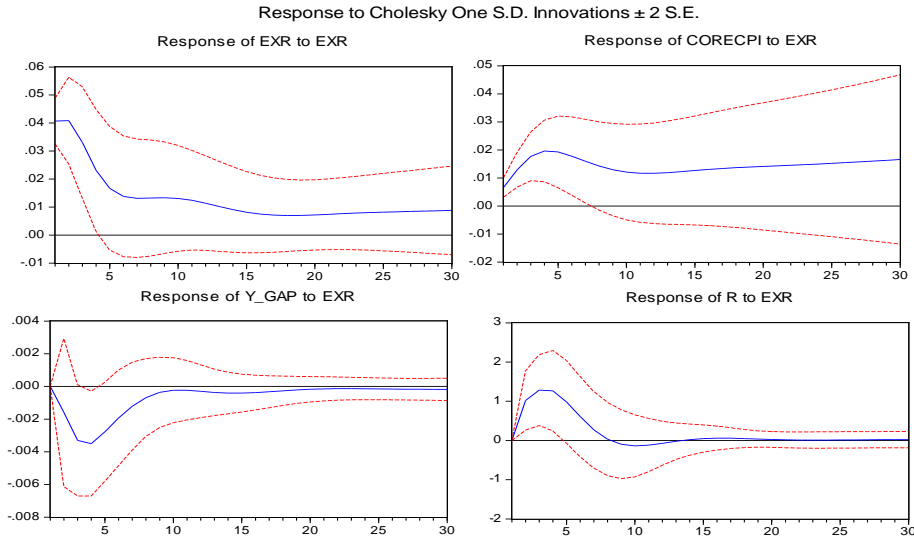Response to Cholesky One S.D. Innovations ± 2 S.E.

Figure 2: Responses to Exchange Rate Shocks

The full effect of this shock, which is realized after about 4 quarters, is about 0.0196 (or 1.96%) increase in the price level, implying a dynamic exchange rate pass-through elasticity of 0.48. As shown in Figure 3, exchange rate shock leads to a sharp increase in inflation that eases after reaching a full effect in 10 quarters, but remains persistent over the long-run. This suggests a strong second round effect (i.e. a prolonged impact of exchange rate movements on inflation) and the fact that prices are generally sticky downwards. With respect to the central bank reaction, the figure indicates that unexpected, temporary depreciation in the exchange rate is followed by a lagged monetary policy tightening (with the impact peaking within the second quarter). Thereafter the central bank eases its reaction (probably informed by corresponding easing of inflation), and eventually settles to its long-run equilibrium path. The output gap responds to exchange rate shock by declining, with the volatility dying out after the ninth quarter.

Table 3: Effect of Cholesky (d.f. adjusted) One S.D. Exr Innovation

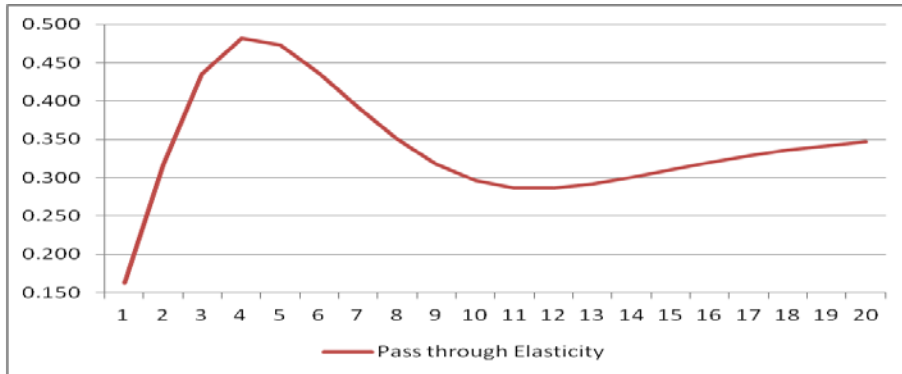| Period	Oil	y_gap	Monetary policy rate	Exchange rate	CORECPI
T=1	0	0	0	0.040695	0.006603
T=2	-0.00113	-0.0016	1.018728	0.040856	0.012896
T=3	0.012943	-0.00331	1.285803	0.033012	0.017706
T=4	0.019442	-0.0035	1.26761	0.023095	0.019605

Figure 3: Dynamic Elasticity of Exchange rate pass-through

Notes: Pass-through elasticity = $\dfrac{\%\,\Delta corecpi_t}{\%\,\Delta exr_0}$ where the denominator is the initial exchange rate shock (assuming no other shock hits the system).

　　　To summarise, the IRFs indicate that the exchange rate pass-through in Uganda is fairly modest, persistent and incomplete. In comparison, our estimate of the pass-through elasticity of 0.48 is consistent with those in Chaoudhri and Hakura (2001), who found the pass-through elasticity of 0.39 for Kenya and Cameroon, and 0.46 for Zambia. Other comparable developing countries studies are Sanusi (2010) who found the pass-through elasticity of 0.79 for Ghana (although fairly large), Anguyo (2008) estimates it at 0.056 for Uganda and Mwase (2006) who estimates the pass-through of 0.028 for Tanzania (the latter two being fairly low). Also, Kiptui et *al*., (2005) found an incomplete pass-through in Kenya that dies out after 4 quarters, with the exchange rate explaining 46 percent of inflation variability.
　　　As noted earlier, one may want to use variance decompositions to gain insights into the relative contribution of the structural shocks in explaining volatilities in inflation. The variance decomposition is shown in Table 4. Consistent with the IRFs discussed above, variance decomposition reveals that exchange rate shocks have a modest contribution to inflation variance, but inflation is mainly driven by own shocks especially at shorter horizons. Specifically, exchange rate shocks account for 25 to 40 percent (at 1 to 10 quarters horizon respectively), while own shocks account for about 74 to 26 percent over the same horizon, suggesting as in (Choudhri and Hakura, 2001) that the level of inflation dominates the volatility of inflation. As before, the contribution of exchange rate own shocks suggests high inflation persistence, underscoring the importance of other factors, other than those explicitly accounted for here in Uganda's inflationary process.

Table 4: Variance decomposition of CoreCPI

				Shocks		
Period	S.E.	oil	y_gap	Monetary policy rate	Exchange rate	CORECPI
T=1	0.01	0.05	0.92	0.09	25.40	73.55
T=2	0.02	0.03	1.31	1.25	36.96	60.44
T=4	0.04	3.06	6.17	1.58	48.52	40.67
T=6	0.06	9.47	8.00	1.15	48.75	32.62
T=8	0.07	15.12	9.08	2.93	44.49	28.39
T=10	0.08	18.83	9.63	5.69	39.86	25.99

Source: Author's computation using E-views 7.2

Finally, our model is checked for robustness, and as results in Appendix 4 suggest, the estimated model produces Gaussian errors, i.e. are normally distributed, serially uncorrelated and have a constant variance.

6 Conclusions and Policy Implications

This paper undertakes an extensive analysis of exchange rate pass-through in Uganda with quarterly data over the period 1999Q3 to 2012Q2 using a triangulation of well specified VECM and SVAR models. A summary of the key results is as follows:

Output gap, nominal exchange rate, oil prices, CPI inflation and short-term interest rates form a long-run stationary relation, which is a statistical analogue of the theoretical link between the inflationary environment and the pass-through. Normalizing the only relation on CPI inflation reveals, as expected a strong and significant association between the exchange rate movements and inflation in Uganda. The impulse responses indicate a fairly modest, persistent and incomplete exchange rate pass-through in Uganda, with a dynamic exchange rate pass-through elasticity of 0.48. This contrasts slightly with the findings in a comparable study on Uganda in Anguyo (2008), where the pass-through in Uganda is found to be as low as 0.056, although persistent and incomplete. We argue that the modest pass-through found here could be attributed to analytical choices and recent fluctuations resulting from the recent economic turmoil and the nascent recovery from it. In addition, variance decomposition reveals that exchange rate shocks have a modest contribution to inflation variance, although it is mainly driven by own shocks at shorter horizons and is persistent over the long-run. We also find that unexpected, temporary depreciation in the exchange rate is followed by a lagged monetary policy tightening (with the impact peaking within the second quarter).

In conclusion, dynamic elasticity of exchange rate pass-through and consequently inflation is persistent. This suggests that exchange rate movements remain a potentially important source of inflation in Uganda. The policy implication arising from these findings is that the monetary authority must be vigilant at exchange rate movements so as to take prompt monetary policy action and focus on exchange rate interventions which stem inflation pressure from the external sector.

Appendix 1: Level and differences Series plots
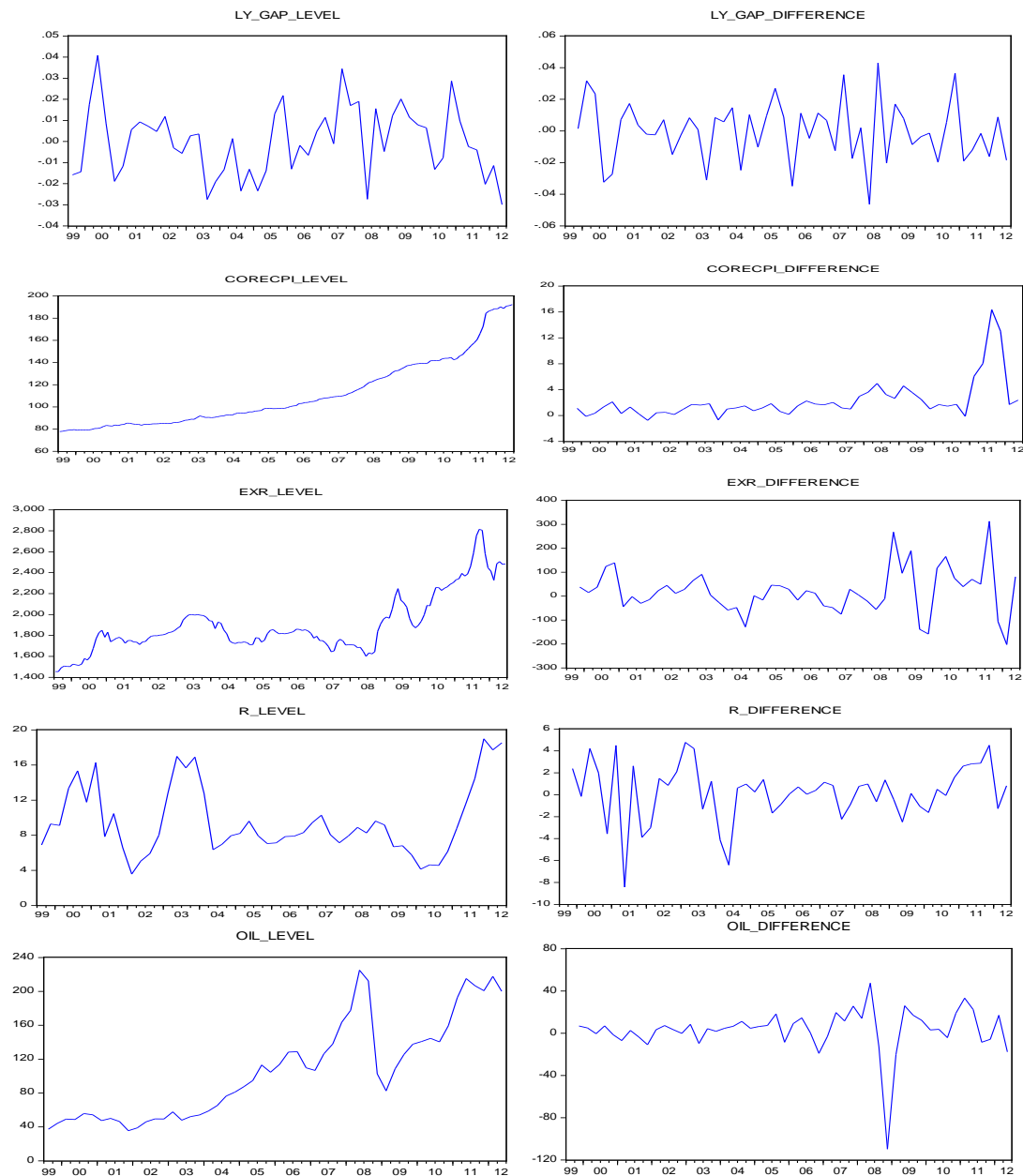
Appendix 2: ADF model framework

In theory, a vector $\mathbf{z}_t$ is said to be integrated of order d (i.e. $z_t \sim I(d)$) if variables in $\mathbf{z}_t$ can be differenced d times to induce stationarity. We employed the commonly used Augmented Dickey Fuller (ADF) unit root test (Dickey and Fuller, 1979) which takes the following specification:

$$\Delta z_t = c_0 + c_2 t + \gamma z_{t-1} + \sum_{i=1}^{\rho} \delta_i \Delta z_{t-i} + \varepsilon_t \tag{1}$$

Where, c_0 is the intercept term, c_2 and γ are coefficients of time trend and level of lagged dependent variable respectively, Δ is the first difference operator and ε_t are white noise residuals. ρ is the lag-length introduced to account for autocorrelation and is chosen using the minimum of the information criteria: Akaike Information criterion [AIC], Schwarz Bayesian criterion [SC] or the Hannan-Quinn Criterion [HQ].

To evaluate whether the sequence $\{\mathbf{z}_t\}$ contains a unit root, we estimated (1) and tested the significance of the parameter of interest, i.e. γ. If $\gamma = 0$, the sequence $\{\mathbf{z}_t\}$ contains a unit root or is otherwise stationary. In the equation, the null hypothesis that $\gamma = 0$ is rejected if the t-statistic is less than the critical value reported by Dickey and Fuller (DF) (1981), as this is a lower tailed test. Furthermore, mindful of the fact that critical values of the t-statistic do depend on whether an intercept (c_0) and/or time trend (t) is included in the regression equation and on the sample size (Enders 2010: 206), the τ_τ - statistic, scaled by the 5 per cent critical value is used for $n = 50$ usable observations. Critical values for the τ_τ - statistic are obtained from Table A in Enders (2010: 488).

The Augmented Dickey-Fuller (ADF) Unit root test					
	ADF test in Level			ADF test in First difference	
Macrovariables	$H_0: \gamma = 0$	Lag-Length	Inference	$H_0: \gamma = 0$	Inference
Oil_t	-3.460 (-3.502)	3	I(1)	-6.440 (-3.502)	I(0)
$CCPI_t$	1.123 (-3.502)	2	I(1)	-4.642 (-3.502)	I(0)
Exr_t	-1.588 (-3.502)	1	I(1)	-5.834 (-3.502)	I(0)
R_t	-3.441 (-3.502)	2	I(1)	-7.376 (-3.502)	I(0)
y_gap	-4.906 (-3.502)	5	I(0)		

Notes: L= log; Akaike Information criterion [AIC], Schwarz Bayesian criterion [SC] and Hannan-Quinn Criterion [HQ] were used (maximum set at 6 lags). An unrestricted intercept and restricted linear trend were included in the ADF equation when conducting unit root test of all the series in levels. Numbers in parenthesis are the 5 per cent critical values, unless otherwise stated. All unit-root non-stationary variables are stationary in first differences.

Source: Author's Computations using E-Views 7.2

Appendix 3: Recursive graphs of the Trace-Test Statistics

The recursive graphs of the trace statistics can also be used to determine rank (the base period was chosen to be 2000q2-2008q1). The graphs are of two versions: the X- form (i.e. the full model) and the R-form (i.e. the concentrated model which is cleaned for short run effects). For the purpose at hand, the R version is used. The graphs should grow linearly for i=1,…,r because they are functions of non-zero eigenvalues and be constant for i=r+1…p because they are functions of zero eigenvalues. About three of them seem to grow over the period, however the lowest one not very much. Although it is not clear from this test which rank to choose, the evidence is not necessarily inconsistent with r=2.
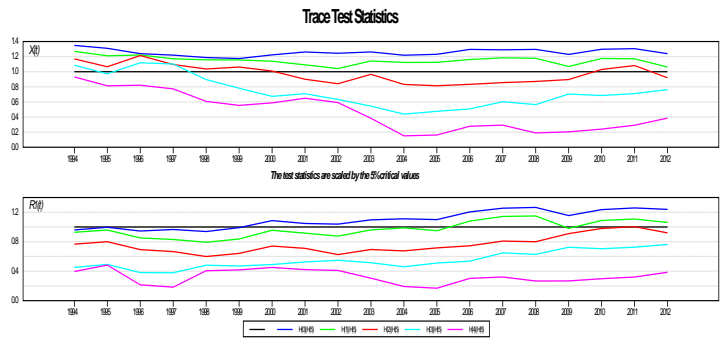
Notes: Due to the difficulties implementing routines in CATS in RATS, periods are annualized, but actually denote the following (1995: 2008q1; 1996: 2008q2; ....)

Source: Authors computation using CATS in RATs by J.G. Dennis, H. Hansen, S. Johansen and K. Juselius, Estima 2012
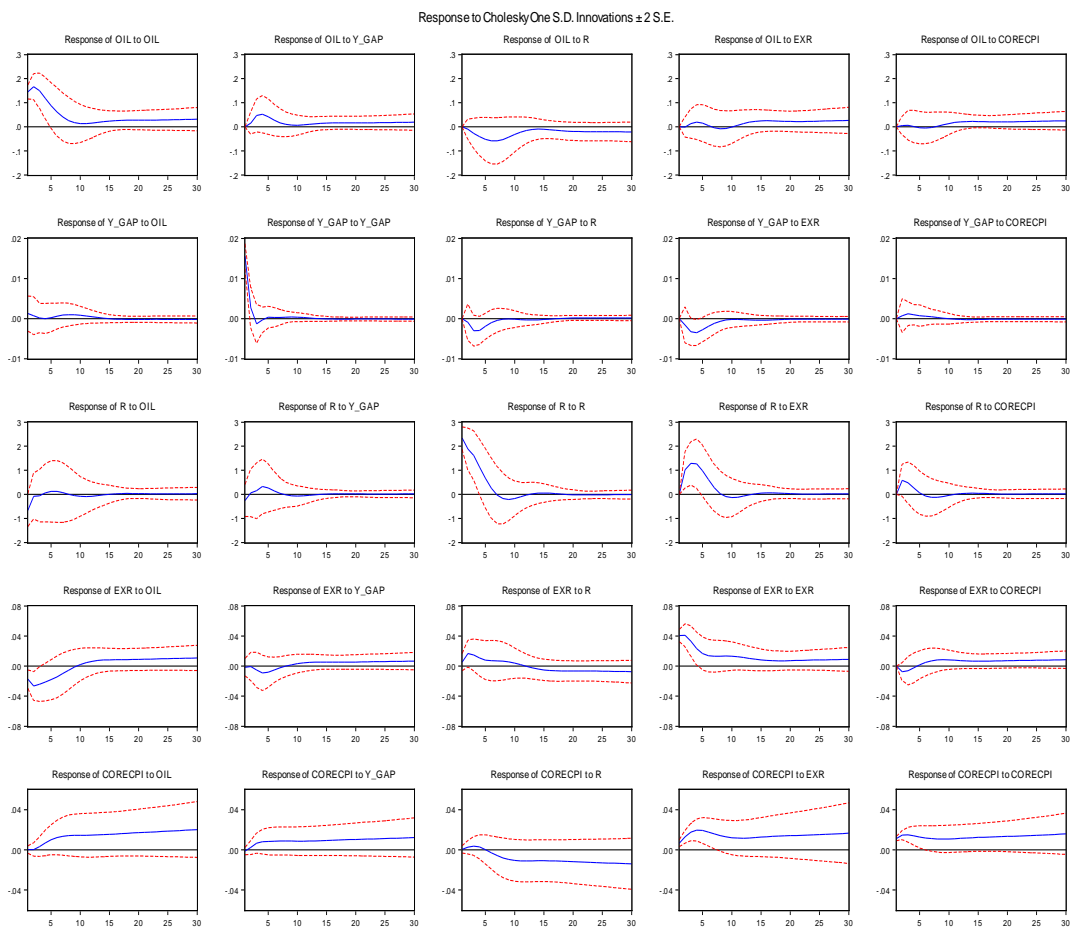

Figure 4: Responses to Exchange rate shocks

Appendix 4: Robustness checks

VAR Residual Serial Correlation LM Tests
Null Hypothesis: no serial correlation at lag order h
Date: 01/25/13   Time: 16:15
Sample: 1999Q3 2012Q2
Included observations: 49

Lags	LM-Stat	Prob
1	28.78860	0.2728
2	30.29016	0.2136
3	35.99333	0.0717
4	28.13275	0.3018
5	24.30735	0.5017
6	34.63295	0.0951
7	21.51861	0.6634
8	16.92205	0.8846
9	13.56134	0.9689
10	28.47485	0.2865
11	24.09715	0.5138
12	16.55196	0.8974

Probs from chi-square with 25 df.

VAR Residual Normality Tests
Orthogonalization: Cholesky (Lutkepohl)
Null Hypothesis: residuals are multivariate normal
Date: 01/25/13   Time: 16:16
Sample: 1999Q3 2012Q2
Included observations: 49

Component	Skewness	Chi-sq	df	Prob.
1	0.307565	0.772538	1	0.3794
2	-0.325808	0.866900	1	0.3518
3	-0.120731	0.119036	1	0.7301
4	0.159392	0.207482	1	0.6487
5	-0.314636	0.808464	1	0.3686
Joint		2.774420	5	0.7347

Component	Kurtosis	Chi-sq	df	Prob.

1	3.163547	0.054610	1	0.8152
2	3.015359	0.000482	1	0.9825
3	2.526892	0.456988	1	0.4990
4	2.641754	0.262028	1	0.6087
5	3.067639	0.009341	1	0.9230
Joint		0.783448	5	0.9781

Component Jarque-Bera	df		Prob.
1	0.827148	2	0.6613
2	0.867382	2	0.6481
3	0.576024	2	0.7498
4	0.469509	2	0.7908
5	0.817805	2	0.6644
Joint	3.557868	10	0.9651

VAR Residual Heteroskedasticity Tests: No Cross Terms (only levels and squares)
Date: 01/25/13   Time: 16:16
Sample: 1999Q3 2012Q2
Included observations: 49

Joint test:

Chi-sq	df	Prob.
493.5807	495	0.5096

Individual components:

Dependent	R-squared	F(33,15)	Prob.	Chi-sq(33)	Prob.
res1*res1	0.830950	2.234282	0.0497	40.71657	0.1672
res2*res2	0.529872	0.512309	0.9462	25.96373	0.8032
res3*res3	0.821433	2.090973	0.0649	40.25022	0.1801
res4*res4	0.576814	0.619558	0.8767	28.26389	0.7020
res5*res5	0.639453	0.806166	0.7073	31.33321	0.5502
res2*res1	0.574407	0.613484	0.8814	28.14597	0.7076

res3*res1	0.908629	4.520193	0.0016	44.52283	0.0868
res3*res2	0.554818	0.566489	0.9147	27.18611	0.7515
res4*res1	0.782932	1.639474	0.1545	38.36365	0.2393
res4*res2	0.690836	1.015696	0.5082	33.85097	0.4263
res4*res3	0.661039	0.886449	0.6286	32.39089	0.4973
res5*res1	0.706431	1.093796	0.4424	34.61511	0.3907
res5*res2	0.777985	1.592817	0.1693	38.12126	0.2477
res5*res3	0.619906	0.741331	0.7701	30.37539	0.5984
res5*res4	0.674403	0.941490	0.5759	33.04573	0.4650

Source: Authors computation using E-views 7.2

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