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Impact of monetary policy on exchange market pressure: The case of Nepal

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ABSTRACT

This paper uses a monetary model of exchange market pressure to examine the impact of monetary policy on the Nepalese exchange rate. Using a recently developed estimation technique, impulse indicator saturation, along with general-to-specific modeling, we find that a contractionary monetary policy results in easing of pressure on the exchange rate. The robustness of the results is confirmed using misspecification tests.

1. Introduction

One of the primary concerns of emerging market economies is that a rush of capital inflows will fuel consumption, which in turn will cause appreciation of the real exchange rate and erode its export competitiveness. In a similar vein, many other less-developed economies receive substantial remittance inflows, which are different from capital inflows but can nonetheless have similar implications on competitiveness. Nepal is one of the largest recipients of remittances, with its size reaching 23 percent of GDP in 2009. In the same year, Nepal was ranked the fifth-largest remittance receiving country in the world by the World Bank.[2] Fueled by remittances and other structural impediments in the economy, Nepal's trade deficit (goods and services) reached 21 percent of GDP in 2009. Some economic and non-economic factors have led to slowing of the economy in the last decade. Average GDP growth during 2000–2009 was 3.9 percent.

As a small open economy, Nepal has maintained a pegged exchange-rate regime with India as its monetary policy anchor. The authorities in Nepal contend that the pegged exchange-rate regime along with growing foreign-currency reserves based

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on remittances have been critical for macroeconomic stability. It is true that the country has not faced any serious balance of payments (BoP) crisis, but the inherent weaknesses in the system have kept unrelenting pressure on the INR-NPR exchange rate.[3] In order to study whether the monetary policy can be helpful in mitigating pressure on the exchange rate, we apply a variant of the monetary model applied in Girton and Roper (1977). The authors introduced the term "exchange market pressure" (EMP), which defines pressure on the domestic currency as a sum of the percentage change in the exchange rate and the percentage change in international reserves. As per the model, an expansionary monetary policy will increase pressure (to depreciate) on the exchange rate, while a contractionary policy will ease any such pressure.

In this study, we examine the interaction of monetary policy and EMP in Nepal for the period 1975–2009. In empirical estimation, we use a recently developed technique called impulse indicator saturation (IIS). In modeling the US expenditure on food, Hendry and Mizon (2011) note that empirical testing of a theory that is essentially "correct" may exhibit misspecification if the data generating process is not properly modeled. Many macroeconomic data often incorporate shocks to the system arising from structural changes in the economy following events, such as recessions, wars, and changes in the policy regime. The authors argue that a theory can perform well in a general framework, which incorporates dynamics as well as possible outliers, breaks, and shifts in the data. The IIS technique by Hendry, Johansen, and Santos (2008) allows for these possibilities in a special case of regression where a dummy variable is added for each observation. The application of the IIS technique is useful not only in detecting structural changes in the data but is also flexible enough to capture analytical insights when the model is complete and correct, reject the model if it is incorrect, and improve it when additional information is provided.[4]

The paper makes two contributions to the literature. This study is the first of its kind on Nepal, which is a low-income, land-locked country that has become increasingly dependent on remittances in recent years. It is a case where remittances are critical for macroeconomic stability in the short run, while there is evidence of adverse effects on the economy over long run through the Dutch-disease effects. By establishing a role for the monetary policy in a pegged exchange-rate regime, it can be expected that the policymakers will ameliorate the adverse effects. The second contribution is the use of a relatively new technique to estimate the EMP model. By applying the IIS technique, this study has attempted to test the empirical predictions of the theory while being robust to structural breaks and outliers, which are common in macroeconomic variables.

Our results show that using the IIS technique helped capture outliers in the data which turned out to be important in finding evidence that supported the theory. Accordingly, the results show that a contractionary monetary policy in Nepal can help ease pressure on the home currency. While the OLS estimates produced mostly insignificant effects and failed the misspecification tests, the results using the IIS technique showed better performance in both regards.

The remainder of the paper is organized as follows. Section 2 summarizes the EMP model and offers a brief review of the literature. Section 3 offers details on econometric estimation, including the IIS technique, general-to-specific (GETS) modeling, and the algorithm called Autometrics. This is followed by a discussion of the results in Section 4, which includes a comparison of the OLS and IIS estimates, results from a multiple structural breaks test, forecast evaluation, and an economic interpretation of the findings. Some concluding remarks are given in Section 5.

2. Exchange market pressure

Girton and Roper used the conventional money demand and supply functions as well as purchasing power parity (PPP) to derive a monetary model of EMP. A variant of the original model is adopted here, following Connolly and Da Silviera (1979), Kim (1985), Bahmani-Oskooee and Shiva (1998), Stavarek and Dohnal (2009), and many others.

The first component of the model is the money demand function, which is a stable function of real income (Y_t) and the price level (P_t):

$$M_t^d = kP_tY_t \tag{1}$$

where k is a constant and represents the fraction of income that firms and households desire to hold as money balance. The money supply, M_t^s, is a product of the money multiplier (m_t) and the monetary base ($B_t = R_t + D_t$):

$$\begin{aligned} M_t^s &= m_tB_t \\ &= m_t(R_t + D_t). \end{aligned} \tag{2}$$

The monetary base is backed by the aggregate stock of foreign assets, R_t, and domestic credit, D_t.

Two assumptions are needed at this point. First, the money market is assumed to be in a continuous equilibrium. Second, the PPP holds:[5]

$$M_t^d = M_t^s \tag{3}$$

$$P_t = E_tP_t^f. \tag{4}$$

[3] INR = Indian Rupee; NPR = Nepalese Rupee.

[4] See Johansen and Nielsen (2009), Santos and Oliveira (2010), and Castle et al. (2012).

[5] The nominal exchange rate (E_t) is expressed as units of the home currency per unit of the foreign currency.

Combining Eqs. (1)–(4) gives the following expression:

$$kE_tP_t^fY_t = m_t(R_t + D_t). \tag{5}$$

Taking the logarithm of both sides and then the time derivative yields

$$\frac{\dot{E}_t}{E_t} + \frac{\dot{P}_t^f}{P_t^f} + \frac{\dot{Y}_t}{Y_t} = \frac{\dot{m}_t}{m_t} + \frac{\dot{R}_t}{R_t + D_t} + \frac{\dot{D}_t}{R_t + D_t} \tag{6}$$

Redefine the variables as follows:

- $e_t = \dot{E}_t/E_t$ is the percentage change in the exchange rate;
- $rsv_t = \dot{R}_t/(R_t + D_t)$ is the change in foreign reserves as a proportion of the monetary base;
- $dc_t = \dot{D}_t/(R_t + D_t)$ is the change in domestic credit as a proportion of the monetary base;
- $mm_t = \dot{m}_t/m_t$ is the rate of change of the money multiplier;
- $p_t^f = \dot{P}_t^f/P_t^f$ is the percentage change in the foreign price level; and
- $y_t = \dot{Y}_t/Y_t$ is the percentage change in real income.

Rewriting the equilibrium in the growth rate form yields the following expressions

$$e_t + p_t^f + y_t = mm_t + rsv_t + dc_t, \tag{7}$$

$$rsv_t - e_t = -dc_t - mm_t + p_t^f + y_t. \tag{8}$$

To consistently apply the notation of EMP, Eq. (8) is rewritten as

$$e_t - rsv_t = dc_t + mm_t - p_t^f - y_t. \tag{9}$$

Here, rsv_t represents the change in foreign-currency reserves scaled by the monetary base and e_t is the percentage change in the INR-NPR exchange rate. The foreign price (p_t^f) represents inflation in India. The model suggests that an increase in domestic credit (dc_t), ceteris paribus, will result in loss of foreign-currency reserves when the exchange rate is fixed (i.e., $e_t = 0$), and in exchange-rate depreciation under a fully flexible system (i.e., $rsv_t = 0$). Clearly, the case of Nepal for the period of this study is that of a managed-peg regime, and therefore, an expansionary policy will result in either reserve losses, depreciation, or a combination of both. The monetary model predicts that both domestic credit and the money multiplier will have a coefficient of one. Similarly, any increase in domestic income or higher inflation in the foreign country will result in either appreciation of the home currency and/or increase in foreign reserves.

EMP has been used as an indicator of pressure on the currency by many researchers for different purposes. After the Girton–Roper paper, several authors have used their model, or variants thereof, to analyze the interaction of monetary policy and EMP. The model has been widely used in countries with either the pegged or managed-peg type exchange rate regimes. Even with some strong assumptions, such as the PPP and the quantity theory of money, there is general empirical support to the predictions of the Girton–Roper type monetary model. However, in the literature there are two key issues often discussed when using the EMP models.

The first one concerns the choice of a policy variable. As derived in the model, domestic credit is shown to reflect monetary conditions. In this regard, Tanner (2000, 2002) points out that although the interest rate reflects the ex-ante policy stance, domestic credit indicates if the monetary policy has been tight or loose ex-post. The author proposes using the interest-rate differential, in addition to domestic credit, to more fully incorporate policy changes and their impact. Therefore, domestic credit is included here to incorporate the broader impact of changes in the monetary policy. Similarly, the second issue pertains to the use of estimation technique. Initially, the OLS technique was frequently used. See, for example, Connolly and Da Silveria for Brazil and Kim for Korea. Mathur (1999) used a similar technique to study the EMP relationship in India. All these studies show the predicted effect of policy changes. However, there were concerns regarding the use of OLS as the estimation technique due to possible endogeneity of foreign reserves and domestic credit. In response, several recent papers report the use of VAR models (e.g., Tanner (2000, 2002)). Garcia and Malet (2007), Gochoco-Bautista and Bautista (2005), and Hegerty (2009) have also used VAR models. These studies show that the impact of policy variable is consistent with the model's prediction. Even though the VAR identification is a matter of concern, researchers have responded by applying either the alternative orderings of variables or used an order-invariant identification scheme.

3. Estimation

There are primarily two measures of EMP used in this study.[6] The first one uses the Girton–Roper's definition (*emp*1), while the second one (*emp*2) is constructed by adding an interest-rate differential term, $\Delta(i_t - i_t^f)$, to *emp*1. Net foreign assets

[6] In keeping with the literature, the two EMP measures are calculated as shown below, where all the variables are as defined in Section 2, including mb_{t-1}, which is the lagged value of the monetary base. $emp1 = \frac{e_t - e_{t-1}}{e_{t-1}} - \frac{nfa_t - nfa_{t-1}}{mb_{t-1}}$; $emp2 = \frac{e_t - e_{t-1}}{e_{t-1}} - \frac{nfa_t - nfa_{t-1}}{mb_{t-1}} + \Delta(i_t - i_t^f)$

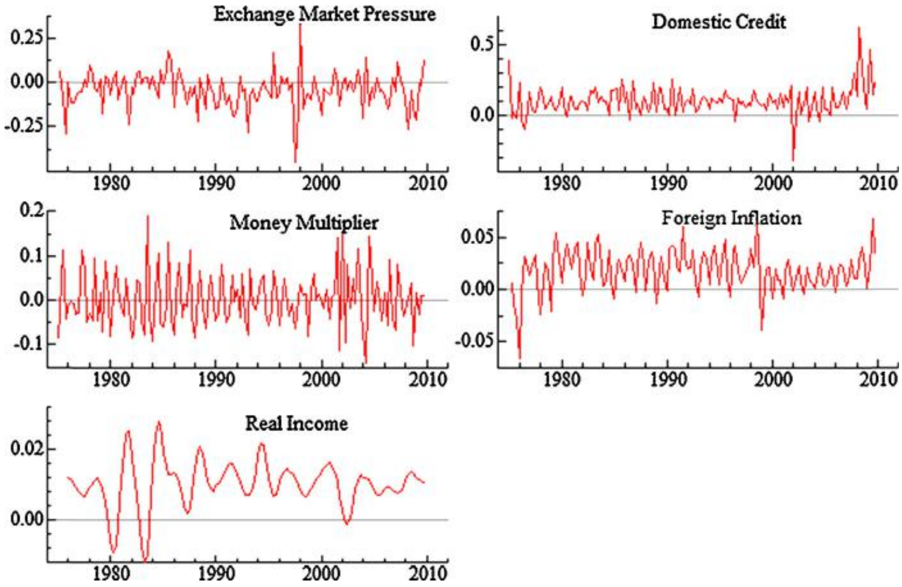

Fig. 1. Graph of all variables (first row left: exchange market pressure (*emp*1); first row right: domestic credit (dc_t); second row left: money multiplier (mm_t); second row right: foreign inflation (p_t^f); third row left: real income (y_t)).
Source: IMF, WDI and Central Bank of Nepal databases.

(nfa_t) are used to represent foreign-exchange reserves, while the central bank's discount rates (i_t is Nepal's and i_t^f is India's) are used as the relevant interest rates. The interest-rate differential augmented EMP (*emp*2) is relevant in Nepal as the monetary authorities may have tried to keep the interest rate above or close to India's, mainly to prevent capital flight. Even with capital account restrictions, the open border presents a challenge for avoiding the capital outflows, especially when significant arbitrage opportunities arise between the two countries.

We use quarterly data (1975Q1–2009Q4) for the empirical analysis.[7,8] Except for the data on output, all other variables are available on a quarterly basis. Real GDP is available only on an annual basis. Moreover, there is no industrial production data to proxy the output. Therefore, we apply a linear interpolation in the levels of real GDP to generate the quarterly output series.

The unit-root test results are shown in Table 1. Out of the three tests used (two unit-root tests and one stationarity test), we decide to follow the rejection of the unit-root hypothesis based on at least two tests.[9] This analysis shows that there is no firm evidence of a non-stationary series or to suggest otherwise. All variables have a high kurtosis and therefore reject the normality assumption.

As noted earlier, we apply the IIS technique for estimation. Particularly, we follow the estimation approach in Hendry and Mizon (2011). They use Autometrics to select a parsimonious model based on the GETS modeling. A similar estimation procedure is followed here. The dynamic models are simple autoregressive distributed lag (ARDL) single-equation estimates. As an alternative estimation technique, we also consider some vector autoregression (VAR) models. The motivation for using the VAR technique is to address the possible endogeneity between the variables. We begin with a brief discussion of the IIS technique.

3.1. Impulse indicator saturation

Hendry et al. (2008) propose selecting a regression model using the GETS approach when there are more variables than observations. It is a special case in which the regression equation is saturated by indicator variables, that is, the regression

[7] A graph of all variables is shown in Fig. 1.

[8] Nepal adopted a formal exchange-rate policy only in 1960, after the establishment of the central bank in the late 1950s. The data on all related variables are available only since 1975 and are collected from the IMF, WDI, and the central bank of Nepal databases.

In the literature, there is some variation in the use of domestic credit to represent monetary-policy stance. In Connolly and Da Silveria the author derive domestic credit by subtracting the rate of change in foreign exchange reserves from the rate of growth of the money supply. Some other studies (e.g., Kaminsky and Reinhart (1999)) use a broader measure, such as the monetary aggregate, while others use either the narrow money or the monetary base (e.g., Tanner (2000, 2002)). In this study, considering the importance of remittances in the economy, outstanding net domestic claims on all depositary institutions are used to derive a measure of domestic credit. This measure of domestic credit is expected to incorporate primarily the effect of policy changes, but it will also account for the effects of remittances received through the banking channel on credit flows.

[9] The unit root tests applied are DFGLS (Dickey–Fuller test with GLS trending) proposed by Elliott, Rothenberg, and Stock (1996) and Phillips and Perron (1988). The stationarity test of time series data is based on Kwiatkowski, Phillips, Schmidt, and Shin (1992).

Table 1
Various unit-root test results.

Variable	Test type	DFGLS Null: Uroot	PP Null: Uroot	KPSS Null: Stationary
$emp1$	c	−3.56***	−11.44***	0.17
dc_t	c	−1.14	−12.47***	0.22
mm_t	c	−0.55	−21.77***	0.08
p_t^f	c	−3.57***	−8.55***	0.16
y_t	c	−1.75*	−3.9***	0.09

Source: Author's calculation.
The lag selection for DFGLS is based on the Schwartz Information Criterion (SIC). A notation 'c' under the test type refers to a constant in the equation.$emp1 = \frac{e_t - e_{t-1}}{e_{t-1}} - \frac{nfa_t - nfa_{t-1}}{mb_{t-1}}$
* Significance level: 10 percent.
** Significance level: 5 percent.
*** Significance level: 1 percent.

equation is augmented by impulse dummies $d_{j,t} = 1_{(j=t)}$, one for every j. Under such a scenario, if one wishes to estimate μ and σ_ε^2 by regressing y_t on $\{\mu, d_{j,t}, j = 1, \ldots, T-1\}$, a perfect fit is achieved and nothing is learned. Therefore, in the first step, only half of the indicators is included in a general unrestricted model (GUM). As shown in Eq. (10), an i.i.d. random variable is regressed on a constant, augmented by a block of impulse indicators:

$$y_t = \mu + \sum_{j=1}^{T/2} \delta_j d_{j,t} + \varepsilon_t. \tag{10}$$

The equation contains $T/2$ parameters for $T/2$ impulse indicators for the first $T/2$ observations. A parsimonious model is then selected and stored from Eq. (10) subject to all mis-specification tests not being rejected and all retained variables being significant at the chosen significance level. Any indicator with an estimated t-value less than the critical value, $|t_{1,\hat{\delta}_t} < c_\alpha|$ (where α is the desired significance level such as 0.01 or 0.025 depending on T), is deleted. In the second step, the regression equation is estimated for the remaining half of the impulse indicators, and y_t is regressed upon $\{\mu, d_{j,t}, j = T/2+1, \ldots, T\}$. All the redundant indicators are eliminated subject to $|t_{2,\hat{\delta}_t} < c_\alpha|$ and the resulting parsimonious model is stored. The selected indicators from the terminal models are combined and re-estimated to give a final model wherein all indicators that did not truly belong to the model are dropped. When testing the null of no significant indicators in the regression, the authors find that the average retention rate of impulse indicators is αT, where α is the significance level (target size) of the test. This is true even when the sample is split into smaller sizes.

Hendry et al. further note that including additional regressors in the equation will add to the computational efforts, but do not affect the analysis. Johansen and Nielsen (2009) show that this approach can be generalized to a class of models including the dynamic ones. And, for a small α, the cost of retaining impulse indicators corresponds to omitting αT observations.[10] A more useful representation of Eq. (10) is

$$y_t = \beta' X_t + \sum_{i=1}^{T} \gamma_i 1_{(i=t)} + \varepsilon_t; \quad t = 1, \ldots, T \tag{11}$$

where X_t is an m-dimensional vector of regressors and γ_i represents the coefficient of significant impulse indicators where $1_{(t)} = 1$ for observation at time "t" and 0 otherwise.[11] Santos and Oliveira (2010) note that the IIS technique can detect breaks at unknown dates—both in the mean and variance. Hendry and Santos (2010, p. 9) argue that the IIS technique not only detects the outliers, "but may also reveal other shifts that are hidden by being 'picked up' incorrectly by other variables." By capturing the outliers, other locational shifts in the data can be detected. In general, the IIS technique is credited for detecting an unknown number of structural breaks, occurring at unknown times, with unknown duration, and occurring anywhere in the sample. Castle, Doornik, and Hendry (2012) sum up the usefulness of IIS by suggesting that it can tackle multiple breaks at unknown times and can handle fat-tailed distributions. Next, we discuss the GETS approach in Autometrics.

Autometrics is a computer-automated package proposed by Doornik (2009) that uses a tree-search method to detect and eliminate insignificant variables.[12] The tree-search method is an improvement over the multi-path search used in the GETS modeling. There are three main stages in selecting the variables beginning with an initial GUM. In the second stage, a variable is eliminated subject to the new model passing the encompassing test and maintaining congruency.[13] The encompassing test ensures that the new model is a valid reduction of the GUM at the chosen α. A path terminates when there is no variable to be

[10] For example, in a sample of 100 observations, with a target size of 0.01 (1 percent), only one observation is lost.
[11] Note that a constant is recommended in the regression. In the absence of a constant, the IIS technique in Autometrics is found to perform poorly in identifying breaks and shifts in the data. Autometrics allows retaining the constant and other variables in the regression equation.
[12] In a typical regression, there are, in total 2^N (N being the number of regressors) possible models from which to select a correct model.
[13] Bauwens and Sucarrat (2010) point out five aspects of model selection to be taken together as the definition of congruency. They are innovation errors, weak exogeneity, parameter constancy, theory-consistent identifiable structures, and data admissibility.

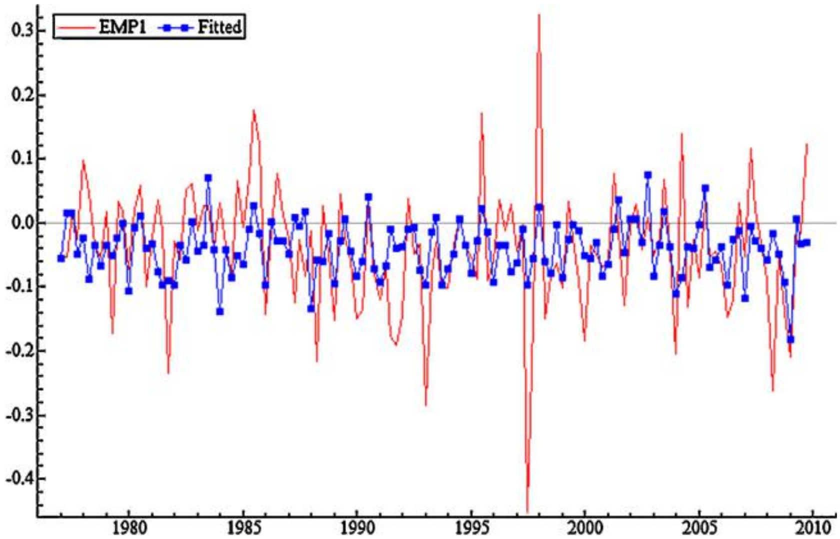

Fig. 2. Actual (straight line) and fitted *emp*1 (dependent variable; dotted line) using ordinary least squares.
Source: Author's calculation.

reduced following the criterion. In the final stage, there can be one or more terminal models, and the selection of the final model can be based on a tie breaker, such as the Schwarz Information Criterion (SIC).

In their estimation, Hendry and Mizon illustrate the use of the IIS technique in several steps. Accordingly, we first estimate an ARDL model with four lags using OLS. Then, we introduce the IIS technique and the GETS approach to select the impulse indicators in Autometrics and present the results.[14] In the final step, we allow Autometrics to select the impulse dummies, but we also allow Autometrics to pick the lags of all regressors. Hendry and Mizon find that retaining (contemporaneous) theory variables in the regression equation and allowing Autometrics to select the lags and impulses significantly improved the result.

The ARDL(p,q) model estimated in this study is shown in Eq. (12), where the selection of lags of the dependent variable (emp_t) and regressors (X_t) along with the selection of statistically significant impulse indicators are based on the GETS approach. All estimations were carried out using PCGive 13.1

$$emp_t = \beta_0 + \sum_{i=1}^{p} \alpha_i emp_{t-i} + \sum_{i=0}^{q} \beta_i X_{t-i} + \sum_{i=0}^{T} \gamma_i 1_{(i=t)} + \varepsilon_t. \tag{12}$$

4. Results

This section begins by briefly highlighting the difference in results due to different estimation techniques, followed by a summary discussion of VAR results. We then discuss IIS results. Finally, the usefulness of the IIS estimates is evaluated using a structural break test and forecast evaluation.

4.1. Overview of OLS and IIS estimates

As discussed in Section 3, the IIS technique is useful for detecting shifts, breaks and outliers in time-series data. This helps not only in identifying the underlying structure, but also in correcting the regression error and in attaining the right model specification. In this study, we find evidence that broadly concurs with previous observations on the use of the IIS technique. In general, the application of the technique improved results, mainly in getting the correct model specification which should be interpreted as the model being robust to structural change and/or outliers in the data. Moreover, by retaining the theory variables (contemporaneous value of all regressors), while allowing the IIS technique to select dummies and dynamic effects, the estimation resulted in economic effects that are consistent with the theory. As will be seen this is clearly an improvement over OLS, which resulted in incorrect specification and either wrongly signed or insignificant coefficients. In Figs. 2 and 3, different actual and fitted series are shown. Clearly, there is evidence of better fit using the IIS technique.[15]

[14] In applying the IIS technique, we retain all regressors including the constant in each stage of model selection.

[15] The estimates using *emp*2 as the dependent variable are not shown here considering brevity of the presentation. In a fuller version of this paper, we also use EMP based on one of the variance equalizing schemes. Additionally, we use annual data to complete the analysis using different definitions of EMP. The main conclusion in all these estimates is that the monetary policy has the predicted effect on the exchange rate.

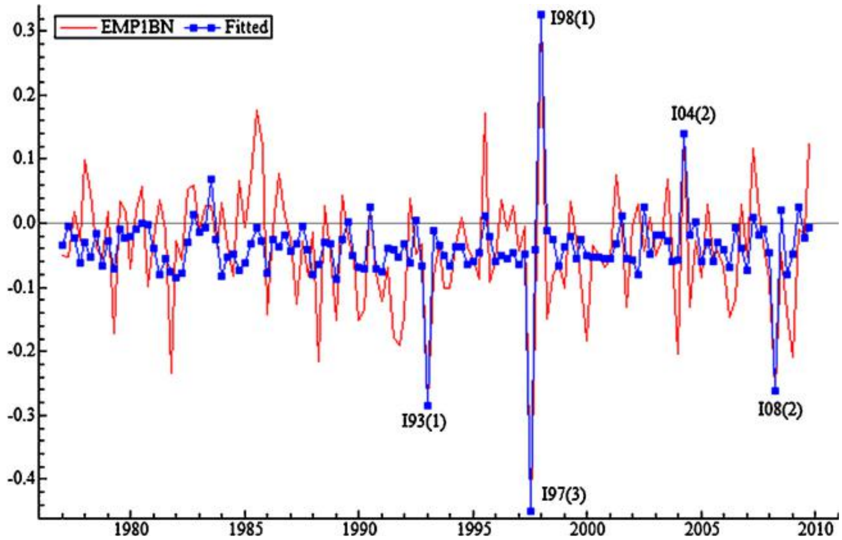

Fig. 3. Actual (straight line) and fitted *emp*1 (dependent variable; dotted line) using the impulse indicator saturation technique.
Source: Author's calculation.

A note of caution is required here. It is difficult to judge from the outset whether the wrongly signed coefficient or a weakly fitted model point to a weak estimation technique or is an outcome of a theory that is not supported by the data. It should also be noted that improving coefficients or the fitness of the model is not the main motivation for using the IIS technique. Rather as pointed above, we use the IIS technique to add a layer of robustness in our estimation such that it can handle problems often encountered in time series data.

Interestingly, however, the selection of impulse dummies did not reveal any regime shift, mainly because the dummies were sparsely located and transitory in nature, and appeared more like outliers. In modeling French inflation, Santos and Oliveira (2010) use eight consecutive periods of significant impulse dummies that have the same sign and a similar magnitude to characterize a regime. If such a regime existed, then a step indicator as proposed by Hendry, Doornik, and Pretis (2013) could be used to accommodate it. Castle, Doornik, and Hendry (2012) also suggest using a similar approach to identify regime shifts. We find no evidence of such a pattern in this study. A formal test of whether the selection of impulse dummies constituted any structural breaks in the data is discussed later.

Next, in terms of the VAR estimation, one of the key issues is model identification, commonly known as the causal ordering of variables. A detailed discussion on identification is given in the appendix. Various pretesting results, such as the Granger-causality test (Table 4) and the VAR-block exogeneity test (Table 5), failed to reveal a causal pattern. Therefore, we adopted a generalized impulse response analysis. Fig. 4 presents selected graphs to summarize the findings. First, neither a domestic credit shock nor a shock to real income has any statistically significant effect on exchange market pressure. Second, there is an instantaneous rise in exchange market pressure following a shock to the money multiplier. This effect, however, is temporary and soon the pressure on the currency starts declining. The rise and fall in EMP becomes insignificant before the second quarter.[16] The partial effect discerned from the VAR analysis is inadequate to guide us much since some of the mis-specification tests (Table 6), especially the normality test, do not comply. This remains a potential drawback of this study and a potential avenue for further research.

4.2. Estimates using the IIS technique

The main results are presented in Table 2. In the table, column (1) shows the results using OLS; in column (2), the IIS technique is introduced and general-to-specific modeling under Autometrics is allowed to select impulse indicators while retaining all theory variables and their lags. Finally, in columns (3a) and (3b) the IIS results based on selecting lags and impulse indicators are shown.[17] In the remainder of the table, various test results to check model specification are shown.

In column (1), except the money multiplier none of the variables is significant. The OLS estimates clearly show statistical inadequacy. This can be seen in the rejection of the null hypothesis of time-invariant volatility and non-normal residuals, respectively, under ARCH (autoregressive conditional heteroskedasticity) and normality tests.

[16] Foreign inflation is treated as an exogenous variable. See Appendix A for more detail.

[17] The difference between the two columns is in the use of target size (significance level), which is done to account for sensitivity associated with using a slightly tighter target size. Generally, a conservative strategy is suggested in applying different target sizes. Hendry and Mizon use 1 percent significance level in their model selection.

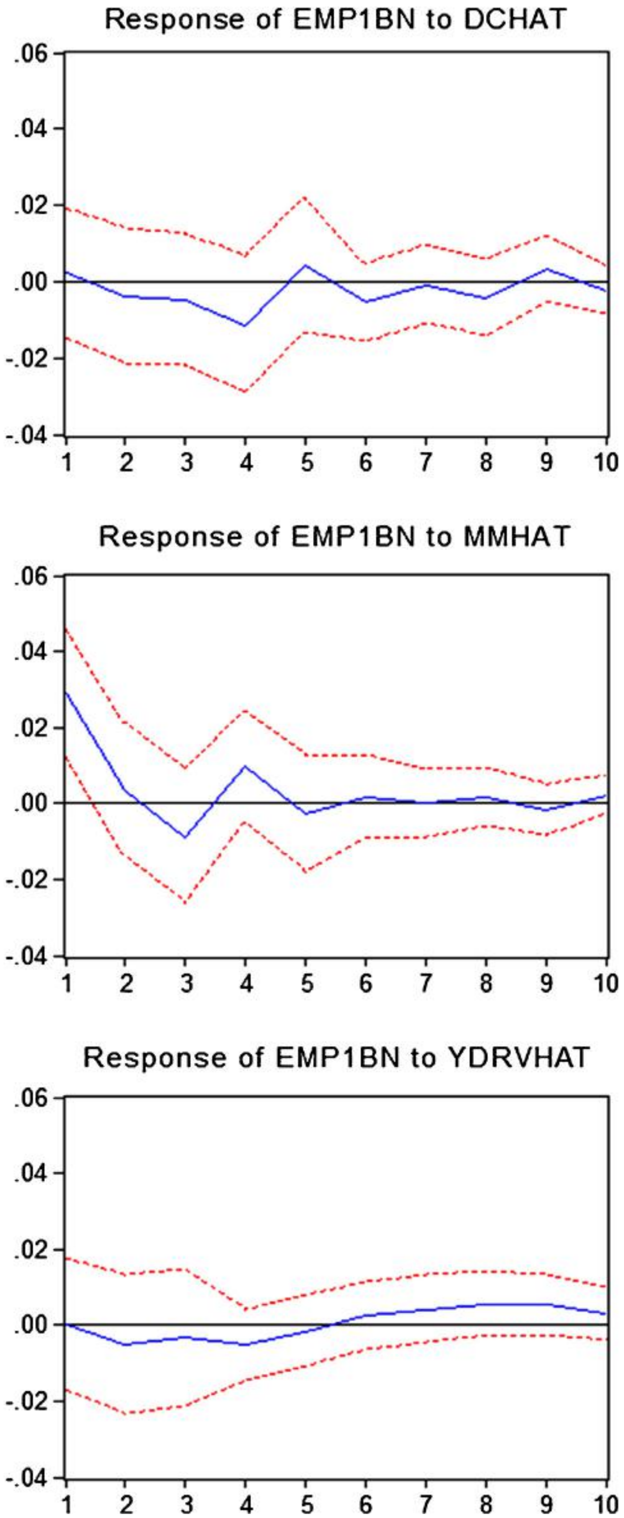

Fig. 4. Generalized impulse responses (top, middle and bottom graphs show the response of exchange market pressure (*emp*1) to a domestic credit shock (dc_t) a money multiplier shock (mm_t) and a real income shock (y_t), respectively).
Source: Author's calculation.

Table 2
Estimation results for 1975 Q1–2009Q4.

Variable	OLS	IIS and regressors "fixed"	IIS and selecting dynamics (TS 0.5%)	IIS and selecting dynamics (TS 1%)
	(1)	(2)	(3a)	(3b)
Constant	−0.05	−0.07**	−0.05***	−0.04***
	(0.04)	(0.03)	(0.02)	(0.02)
$emp1_{t-3}$	0.14	0.15*		
	(0.1)	(0.08)		
dc_t	0.09	0.3***	0.23***	0.21***
	(0.1)	(0.09)	(0.08)	(0.08)
dc_{t-3}	−0.15	−0.15*		
	(0.1)	(0.08)		
mm_t	0.63***	0.67***	0.42***	0.38***
	(0.21)	(0.18)	(0.13)	(0.13)
mm_{t-1}	0.34	0.43**		
	(0.23)	(0.18)		
p_t^f	0.01		0.14	0.05
	(0.67)		(0.41)	(0.4)
y_t	1.7		−1.6	−1.73*
	(5.9)		(0.98)	(0.96)
I93(1)				−0.21***
				(0.08)
I97(3)		−0.37***	−0.37***	−0.37***
		(0.08)	(0.08)	(0.08)
I98(1)		0.39***	0.35***	0.35***
		(0.09)	(0.08)	(0.08)
I04(2)		0.27***	0.25***	0.25***
		(0.09)	(0.08)	(0.08)
I08(2)		−0.38***	−0.34***	−0.33***
		(0.09)	(0.09)	(0.09)
σ	0.1	0.08	0.08	0.08
R^2	0.19	0.52	0.41	0.44
$\bar{R}^2$	0.01	0.39	0.37	0.4
AR test	1.1	0.5	0.63	0.79
	(0.36)	(0.78)	(0.67)	(0.56)
ARCH test	2.2*	0.25	0.16	0.13
	(0.07)	(0.91)	(0.96)	(0.97)
Normality test	25.26***	0.94	0.35	0.02
	(0.00)	(0.63)	(0.84)	(0.98)
Heteroskedasticity test	1.25	1.11	0.65	0.77
	(0.19)	(0.33)	(0.73)	(0.63)
Ramsey reset test	0.4	0.43	0.54	0.56
	(0.67)	(0.65)	(0.58)	(0.57)

Source: Author's calculation.
The dependent variable is the combination of the percentage change in the exchange rate and changes in the foreign asset scaled by the monetary base. $\left(emp1 = \frac{e_t - e_{t-1}}{e_{t-1}} - \frac{nfa_t - nfa_{t-1}}{mb_{t-1}}\right)$
The third column presents results based on the IIS technique where all regressors are retained in the model, that is, Autometrics does not eliminate these variables but only selects the impulse indicators. In the fourth and fifth columns, all (contemporaneous) theory variables are retained, but the selection of lags and impulse dummies is based on Autometrics.
Standard error for the coefficients is shown in parentheses, and diagnostic tests show the *p*-value in parentheses.
Target size (TS) of 0.5 percent is applied in column (3a) and 1 percent in columns (2) and (3b).
* *P*-value: 10 percent.
** *P*-value: 5 percent.
*** *P*-value: 1 percent.

Also, there is little explanatory power in this model as shown by adjusted R^2. It should be noted that the rejection of the normality test is likely driven by outliers in the data. These outliers are then picked up by IIS in later estimations as discussed below.

In column (2), domestic credit and the money multiplier are significant, although the third-period lag of domestic credit is wrongly signed. Foreign inflation and income growth are still insignificant. Moreover, there are four significant impulse indicators chosen in this step. The drastic improvement in model specification is evident. Not only do the residual tests confirm statistical adequacy, but also the model now has a decent explanatory power. The results in the next two columns (3a) and (3b) show that statistical inference is robust and the coefficients have the economic meaning as hypothesized by the theory.

In the two columns, the variables are correctly signed and almost identical. Domestic credit and the money multiplier are highly significant. Foreign inflation is insignificant in both, which may be due to the use of CPI-based inflation.[18] Income growth has a borderline significance in column (3a), with a p-value of 0.1052, but is significant at 10 percent in column (3b). The main difference between columns (3a) and (3b) is the inclusion of an additional impulse indicator, I93(1), in the latter.[19] In terms of model specification, all tests confirm randomness of the residuals. The *p*-values of these tests are given in parentheses. First, the AR test confirms the failure to reject the null of no autocorrelation in the residuals. Second, the ARCH and heteroskedasticity tests show that the null of no time-varying volatility in the residuals cannot be rejected. Third, the normality test confirms that the residuals are asymptotically normal. Finally, the Ramsey Reset test suggests there is no non-linear behavior in the residuals.

In addition to improved misspecification test results, some other parameters also showed improvement. For example, the adjusted R^2 is higher using IIS which indicates progress in model's explanatory power compared to the OLS estimate. Similarly, the σ estimate (standard deviation of the regression error) is also lower with IIS. However, it should be noted that these improvements are only used as reference measures and are not the motivation behind using the IIS technique. Taken together, using IIS helped to attain an empirically robust model in which the underlying data generating process was captured better compared to the OLS.

To offer economic interpretation of the findings, estimates from column (3b) are used. The coefficients reported for domestic credit and the money multiplier are well below one. The negative coefficient on income growth suggests that higher growth leads to appreciation of the exchange rate and/or greater reserve accumulation. Higher income growth means higher money demand, which ceteris paribus necessitates a fall in prices to maintain the real money balance, resulting in appreciation.

Several studies in the past have also failed to find full support for the monetary model's predictions. They commonly cite unrealistic assumptions regarding price flexibility, which is central in all monetary models. Moreover, there can be some restrictions in the economy, such as capital control and trade barriers, that can lessen the impact of monetary policy on EMP. A quick reference to previous studies on Nepal is relevant here. Ginting (2007) and Yelten (2004) suggest that prices in Nepal tend to be influenced by prices in India. Clearly, the pegged exchange rate with India is a big facilitator for co-movement in prices. These findings also indicate a limitation for policy effectiveness in Nepal. Unlike the one-to-one effect of the money supply on the exchange rate, it is therefore plausible to expect a lesser impact.

The partial validation of the predictions of the monetary model in this study is important from policymakers' point of view. The estimated effect of domestic credit on EMP suggests some role for monetary policy in influencing the exchange rate. Specifically, a contractionary monetary policy can help reduce pressure on the domestic currency.

Next, there are five significant impulse indicators in the chosen estimates. Except the first (I93(1)), all other indicators are in the post-liberalization period when there was a fixed nominal peg, so EMP includes the changes in foreign reserves only.[20] Moreover, the impulse indicators in the post-liberalization period (1993 onwards) are sparsely located and appear more like outliers. The I93(1) indicator captures the economically important transition period. This indicator is negatively signed, suggesting a reduction in EMP which is possibly an outcome of policy reforms. During the first quarter of 1993, the central bank had directly intervened to adjust the parity of the NPR/INR exchange rate for the last time. The exchange rate was revalued to 1.6 from 1.65, an appreciation of about 3 percent. This followed the announcement of current account liberalization, which was in sync with the liberal orientation of the economy at the time. It is important to note that the announcement of economic liberalization in India also occurred during this period.

The remaining impulse indicators capture fluctuations in the data that do not correspond to any known economically significant events in Nepal. Their inclusion can be better understood using macro data, especially the changes in the current account. The current account deficit increased to 7.89 percent of GDP in 1997. This trend continued until the third quarter. Gains in the financial account, however, moderated the impact on the overall balance of payments. The stock of net foreign assets grew in the third quarter of 1997, while it fell in the first quarter of 1998. These fluctuations in net foreign assets are captured by a negative coefficient in I97(3), signifying a fall in EMP, and a positive coefficient in I98(1), indicating a rise in EMP. Similarly, the positive coefficient in I04(2) and the negative coefficient in I08(2) may be explained by the fall and rise in the current account surplus, respectively.

Earlier it was argued that the selected impulse indicators, except I93(1), suggested outliers in the data. If instead the indicators had picked up structural changes in the data, that would require a different treatment. In order to confirm robustness of the estimates, a two-prong assessment approach is adopted. First, a formal structural break test is conducted. Second, a forecast evaluation is considered.

The structural break test used here is for multiple breaks in the data proposed by Bai and Perron (1998). The test is based on minimizing the sum of squared residuals from a linear regression, where the dates of the breaks are unknown variables to

[18] In this regard, a better indicator perhaps is the wholesale price index (WPI). As a check of robustness, we used the producers' price index (PPI) to represent foreign prices. There was a marginal improvement in the coefficient estimate, but it was not statistically significant. The WPI series for the whole period is not available.

[19] Note that the impulse indicator is read as, for instance I93(1), where 93 indicates year 1993 and number 1 in the parenthesis indicates the first quarter of the year.

[20] Nepal's economy until the early 1990s was more isolated from the global economy. There were several restrictions in the economy, including limited transactions on the current account. In a series of reform measures, several liberal economic policies were adopted.

be estimated. As a useful point of reference, we can consider five break points (as implied by the IIS estimates) in the data to begin with. However, as it turns out, owing to the large number of variables (four lags of each variable) in the regression equation and a total of 132 observations, a maximum of only four break points could be considered in the test. The results are presented in Table 3. As seen in the table, the number of break points with the lowest BIC is zero, that is, no break point. This confirms our earlier suspicion that the impulse indicators may have captured outliers in the data and that they do no indicate any structural breaks or existence of different regimes in the data. The model's performance is further tested based on forest accuracy. Our goal is to find an alternative way to test robustness of the IIS estimates. However, we do recognize that using forecast accuracy may not always be a reliable way to evaluate a model's performance.

Table 3
Bai and Perron (1998) multiple structural breaks test.

Breakpoints	BIC	Log-Lik	RSS	N. Coeff.
0	−143.77	135.36	0.99	26
1	−76.16	165.03	0.63	52
2	−38.5	209.68	0.32	78
3	−13.58	260.7	0.15	104
4	−120.58	377.68	0.03	130

Source: Author's calculation.
Chosen number of breaks: 0.
BIC is the Bayesian Information Criterion, Log-Lik is the log likelihood estimates, and RSS is the residual sum of squares. The last column shows the number of coefficients.

Table 4
Pair wise Granger-causality test with two lags using *emp*1 for period 1975:01–2009:04. (p_t^f is treated as an exogenous variable).

Null hypothesis	*F*-statistics	Probability
dc_t does not Granger cause *emp*1	0.51	0.85
*emp*1 does not Granger cause dc_t	0.45	0.89
mm_t does not Granger cause *emp*1	0.82	0.59
*emp*1 does not Granger cause mm_t	0.85	0.56
y_t does not Granger cause *emp*1	0.82	0.58
*emp*1 does not Granger cause y_t	0.46	0.88
mm_t does not Granger cause dc_t	0.86	0.55
dc_t does not Granger cause mm_t	1.66	0.12
y_t does not Granger cause dc_t	0.53	0.83
dc_t does not Granger cause y_t	0.45	0.89
y_t does not Granger cause mm_t	1.37	0.22
mm_t does not Granger cause y_t	0.32	0.96

Source: Author's calculation.

Table 5
VAR Granguer causality/Block exogeneity Wald test (1975:01–2009:04) (p_t^f is treated as an exogenous variable).

Dependent variable Excluded	χ^2	df	Probability
mm_t	4.23	4	0.38
dc_t	0.48	4	0.98
y_t	1.84	4	0.77
All	7.14	12	0.85
*emp*1	2.92	4	0.57
mm_t	1.18	4	0.88
y_t	0.88	4	0.93
All	5.96	12	0.92
*emp*1	2.24	4	0.7
dc_t	11.5	4	0.02
y_t	7.24	4	0.12
All	22.7	12	0.03
*emp*1	3.94	4	0.41
mm_t	2.08	4	0.72
dc_t	2.71	4	0.61
All	7.91	12	0.8

Source: Author's calculation.

Table 6
VAR residual test results.

Test type	Test statistics	*P*-Value
Autocorrelation test Null: no serial correlation	LM Stat(5): 7.51	0.96
Normality test Null: multivariate normal	Jarque–Bera: 244.89	0.00
Heteroskedasticity test Null: no heteroskedasticity	Chi-square: 375.42	0.09

Source: Author's calculation.

4.3. Forecast evaluation

Here, we consider a one-step-ahead forecast based on rolling regressions, using OLS and IIS separately. The model is first estimated for a period of 20 years (i.e., 80 quarters) and then a one-step-ahead forecast is recorded (for the 81st quarter). The sample is then rolled over one step while keeping the window of 80 quarters unchanged. The model is again estimated and the forecast recorded. This gives the second observation. By proceeding in this manner, a forecast series of 52 observations—beginning in the first quarter of 1997 and ending in the fourth quarter of 2009—is generated. In the IIS-based forecast series, Autometrics was allowed to pick impulse indicators as the sample was rolled over. It was observed that the IIS technique picked up the indicators as shown in Table 2 as the rolled over sample was re-estimated each time. With the two forecast series in hand, we computed the root mean squared error (RMSE) and the mean absolute error (MAE) for each of them. The RMSE based on forecast using the OLS estimates is 0.14, while the same using the IIS estimates is 0.12. This corresponds to about 14 percent or about one-third of one standard deviation improvement over the OLS-based forecast. Similarly, the MAE using the OLS-based forecast is 0.094, while it is 0.086 using the IIS-based forecast. This improvement is about 8.5 percent or roughly one-tenth of one standard deviation over the OLS-based forecast.

Finally, we apply a forecast-accuracy test to examine if the improvement in the IIS-based forecast is statistically significant. Diebold and Mariano (1995) proposed a test to compare the accuracy of two forecast series based on a loss function, which uses loss differential (squared error loss) from the two series. The null hypothesis under the test is that of no improvement in forecast accuracy or the null of equal predictive accuracy. The test result (DM stat = 4.25, p-value <0.01) shows that this null is rejected. Thus, the difference observed in terms of forecast performance is statistically significant, which offers further evidence that using IIS added strength to the empirical analysis.

5. Conclusion

Many countries, especially in a non-floating exchange-rate regime, face pressure on their currency. Only a few of them suffer a currency collapse. It is possible that by establishing a proper role for the monetary policy, the authorities can adopt preemptive corrective actions to reduce pressure on a currency and avert any serious crisis. The EMP framework based on a monetary model allows investigating the impact of policy changes on the exchange rate.

In this study, Nepal's exchange-rate with India was investigated using the EMP framework based on a monetary model. Despite weak fundamentals, Nepal continues to maintain a pegged exchange-rate regime with support from exogenous sources, like remittances, which have underpinned the balance of payments for a long time. By using domestic credit as a policy variable, a positive impact of expansionary monetary policy on EMP was found. The coefficients on domestic credit and the money multiplier had positive signs, but were well below one. Output growth was found to lower EMP as expected. Foreign inflation, however, was insignificant. Despite finding a lesser magnitude effect than predicted under the monetary model, these results suggest that the monetary policy is relevant in Nepal, and a contractionary policy can reduce pressure on the home currency.

One important dimension of the study was the application of the IIS technique. By allowing impulse dummies for each observation, the technique enabled us to identify outliers in the data in a systematic way. The IIS technique was used to select impulse indicators and the lags of all variables using the GETS approach in Autometrics. In most of the regressions, the selection of impulse indicators significantly improved the model's performance by confirming to various specification tests and producing theory-consistent effects. There were five impulse indicators in the chosen model, with the first one corresponding to the first quarter of 1993, which is important in view of several economically significant events occurring at the time. The rest of the impulse indicators merely picked up big fluctuations in the data. While the structural break test did not reveal any statistically significant breaks in the data, the IIS estimates were found to have better predictability using one-quarter ahead forecast.

Appendix A

A.1. VAR identification

As an alternative to the single-equation estimation, a system-based vector autoregression (VAR) model is considered. A VAR model is generally used to accommodate feedback effects in macroeconomic variables. Using a VAR model allows us not only to observe the effects of conventional monetary policy but also to analyze how the authorities respond to pressure on the domestic currency. The possibility of reverse causality (endogeneity) can be addressed in a VAR framework. Formally, a reduced form VAR of order p is given as follows:

$$z_t = A_1 z_{t-1} + \cdots + A_p z_{t-p} + \varepsilon_t, \quad (13)$$

where z_t is a vector of all domestic variables (emp_t, dc_t, mm_t, y_t) and ε_t is a vector of i.i.d. shocks. One important step in the VAR estimation is identification, more commonly known as the causal ordering of variables. All domestic variables enter the VAR specification as endogenous variables. Some ad-hoc assumptions are needed to complete identification. In line with several previous studies, the exogeneity restriction or the Choleski ordering of the variables is as shown below:

$$\varepsilon_{y_t} = e_{y_t}, \quad (14)$$

$$\varepsilon_{dc_t} = e_{dc_t} + e_{y_t,dc_t}, \quad (15)$$

$$\varepsilon_{mm_t} = e_{mm_t} + e_{y_t,mm_t} + e_{dc_t,mm_t}, \quad (16)$$

$$\varepsilon_{emp_t} = e_{emp_t} + e_{y_t,emp_t} + e_{dc_t,emp_t} + e_{mm_t,emp_t}. \quad (17)$$

In Eqs. (14)–(17), income growth is treated as the least endogenous variable, while EMP is the most endogenous variable. Income (y_t) is affected in the current period only by its own shocks, while other shocks have a lagged effect. On the other hand, innovations in EMP are correlated with all other shocks and are therefore affected by contemporaneous as well as lagged feedback from all other shocks. Similarly, domestic credit is affected by the current and past shocks of its own and income growth, while EMP has a lagged effect, which is reasonable given that the authorities are likely to respond to a rise in EMP with some time lag rather than acting concurrently.

The economic relationship underpinning the causal ordering in Eqs. (14)–(17) cannot be firmly established if the alternative ordering produces significantly different implications. Therefore, an order-invariant approach as in generalized impulse response by Pesaran and Shin (1998) is also adopted.

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