Making the most of high inflation

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Abstract

The article examines the relationship between the real effects of inflation and its level in countries with frequent episodes of high inflation. The real effects are computed as asymmetric impulse responses of output to inflation separately for the regimes with different signs of the differences between the expected inflation and the predicted output-neutral inflation. It is found that, with the increase in inflation, such effects increase for the regime with the positive sign, relatively to the effects for the regime with the negative sign. It is also shown that this finding is valid for most countries with high inflation episodes, where inflation is greater than 4.8% for at least 25% of quarterly observations. This leads to a simple policy prescription that, in economies with frequent high inflation episodes, anti-inflationary monetary decisions are least damaging for output if undertaken in the periods when the difference between the expected and output-neutral inflation is negative.

I. Introduction[1]

Investigation of the nature and strength of the relationship between inflation and the real sphere is, so far, not close to being conclusive. On the theoretical side, there are two main streams of the literature on this topic: (1) following Tobin's (1965) argument that under high inflation wealth is likely to be reallocated from money to physical capital, which stimulates growth, and (2) following Sidrauski (1967), that the Tobin effect is offset by increased consumption (as holding real balances is costly), creating superneutrality of inflation. Even more pessimistic views have been developed from the early papers by Brock (1974) that endogenous labour supply stimulates a negative inflation-output relationship by reducing the cost of leisure and from Stockman's (1981) 'cash in advance' approach, in which investment transactions becomes more costly under rising inflation and therefore negatively affect output. The empirical findings are mostly on the side of the pessimists. The statement that loosely defined 'high' inflation is bad for growth seems to be widely confirmed, by the comparative survey of early results by Braumann (2000) and also by later findings (see e.g. Mallik and Chowdhury, 2001; Grier and Grier, 2006, Gillman and Harris, 2010 for the developing and transition economies). However, results by Bruno and Easterly (1996 and 1998) indicate that periods of high inflation (albeit, not hyperinflation) were often followed by growth in the long-run. Also, for some Asian countries more recent empirical findings point out at the neutrality of inflation (Kun, 2012). It is, therefore, quite natural that the empirical literature focuses on finding the threshold above which inflation might be harmful to growth. Most of the research implicitly assumes that such threshold is common for a

[1] This article was presented at the third ISCEF (Paris, April, 10-12, 2014,www.iscef.com)

relatively large group of countries and applies the cross-sectional or panel data methods in order to identify it (see e.g. Sarel, 1995; Khan and Senhadji; 2001; Rousseau and Wachtel, 2002; Vaona and Schiavo, 2007; Basco *et al*., 2009; Bick, 2010; López-Villavincencio and Mignon, 2011; Kremer *et al*., 2013; Amisano and Fagan, 2013).

The problem tackled in this article is similar; we aim at identifying a regime in which a positive inflationary shock to inflation contributes to output increase stronger than a shock induced in a different regime. Rather than to evaluate the optimal inflation threshold common for a specific group of countries, we assume, after Fischer and Modigliani (1978), that the institutional country-specific effects like taxation, financial systems, corruption levels, differences in reporting (resulting in different money illusion effects) *etc*. are important enough to create individual conditions for the development of inflationary real effects. We conjecture that this is particularly true for countries where the episodes of high inflation are relatively frequent. For these countries we aim at identification of the inflation regimes for which, depending on the magnitude and frequency of high inflation episodes, inflationary shocks might have different real effects. In order to identify such regimes we distinguish between the expected inflation, in the rational expectations sense, and the predicted output-neutral inflation. Then we define the output-active inflation (denoted by *OAI* further on) as the difference between the expected inflation and the predicted output-neutral inflation. The different forward-looking inflation regimes are identified by the signs of *OAI*'s. The article shows that the cumulative balance of such real effects in different *OAI* regimes is positively related to the magnitude of inflation in countries that experience periods of high inflation relatively often. This is, in fact, the development of the Hartmann and Roestel (2013) finding that the low inflation countries lose more than the high inflation countries from raising inflation, in terms of output. Our results lead to practical policy prescriptions. If a country experiences high inflation, knowledge of *OAI* regimes might suggest the best moment for undertaking the anti-inflationary policy which would hurt the real sphere the least. Analogously, it might also lead to the identification of the conducive moment for output-stimulating decision.

A simple vector autoregressive model (VAR) for inflation and output is applied as the initial device. Using the decomposition of the inflation and output shocks identified from this VAR (see Blanchard and Quah, 1989, Quah and Vahey, 1995), we compute two *ex-ante* inflation indicators: expected inflation and output-neutral (predicted) inflation, and, with their use, *OAI*. Next, we evaluate the cumulative asymmetric impulse responses separately for the periods of positive and negative *OAI*'s and analyse their balance (that is, the difference between the cumulative impulse responses of output to inflationary shocks for these two regimes).

From the initial set of 45 countries, 17 for which the 0.75th quantile of annual inflation is equal to at least 7.5% have been originally selected. These countries are referred to as *countries with frequent episodes of high inflation*. Later on, the group of countries with high inflation episodes has been gradually enlarged by lowering the 7.5% criterion. For the countries selected, *OAI*'s have been computed, and the asymmetric impulse responses of output to symmetric inflationary shocks evaluated. Strong positive correlation between the differences in these cumulative impulse responses and the logarithm of the 0.75^{th} quantile of inflation, measuring the magnitude of high inflation episodes, has been found. This leads to the conclusion that, for a country with a history of high inflation episodes, identification of the forward-looking inflation regimes is relevant for undertaking monetary policy decisions. More precisely, an anti-inflationary decision should be made in the periods where such regime is negative, that is when the expected inflation is below the predicted output-neutral inflation. It is also found that the higher inflation becomes, the stronger is the conclusion above, as reducing the limit of 7.5% for the 0.75^{th} quantile lowers the correlation strength. Nevertheless, the correlation becomes significant down to the limit of 4.8%.

Further structure of the article is as follows. Section 2 contains the main concepts, definitions, and derivation of *OAI* for a simple output-inflation vector autoregressive model. Section 3 briefly discusses the data and introduces our selection of countries with high inflation episodes. Further on it outlines results of the impulse response estimation and presents more detailed results for three benchmark countries: Indonesia, Malaysia and Pakistan. It also contains the main empirical results of the article, which is evidence of significant positive correlation between the magnitude of frequent high inflation and the cumulative balances of inflationary real effects from shocks in different forward-looking inflation regimes. The robustness of the results is checked by (*i*) evaluation of a possible misspecification effect due to overlaps of periods of positive *OAI* with periods of high inflation and (*ii*) gradual relaxing the definition of high inflation episodes. Section 4 provides policy conclusions.

II. Methodology

The intuition of *OAI* can be explained by a simple representation of a typical aggregate supply function, supported indirectly or directly, by a plethora of papers from the seminal works of Lucas (1972) and Bull and Frydman (1983) to thoroughly microfounded approaches by Golosov and Lucas (2007) and Midrigan (2011):

$$y_t = \theta(\pi_t - \pi_t^n) \quad , \quad \theta > 0 \quad , \tag{1}$$

where y_t is a measure of output dynamics (net of long-run effects), π_t is the headline (observed) inflation and the expected at t-1 output-neutral inflation is π_t^n. Evidently:

$$\pi_t = \pi_t^e + \upsilon_t \quad , \tag{2}$$

where π_t^e is inflation expected at t-1 and υ_t is a shock unexpected at t-1. It is usually assumed that $\pi_t^n = \pi_t^e$. However, in an economy with sticky prices, some individual relative prices cannot be fully adjusted after a shock and could have long-lasting effects on output, even if fully expected. Consequently, another decomposition of π_t is:

$$\pi_t = \pi_t^n + \omega_t \tag{3}$$

where ω_t is the non-neutral component of inflation. As π_t^n is also based on information available at time t-1, in order to avoid confusion, π_t^e is termed the expected inflation and π_t^n the output-neutral predicted inflation.[2] Referring to the seminal literature on inflation decomposition, π_t^e is similar to core inflation in the sense of Eckstein (1981), i.e. the systematic (predictable) component of the increase in production costs. In turn, π_t^n is analogous to core inflation in the sense of Quah and Vahey (1995), i.e. the component of expected inflation that does not cause a real effect in the medium and long-run.

Substituting (2) in (1) and bearing in mind that output-neutral component of inflation is evaluated on the basis of information available at time t-1, we get:

$$E_{t-1}y_t = \theta \cdot E_{t-1}(\pi_t^e + \upsilon_t - \pi_t^n) = \theta \cdot (\pi_t^e - \pi_t^n) \quad , \tag{4}$$

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[2] Strictly speaking, π_t^e is affecting output only if it is not equal to π_t^n.

where E_{t-1} denotes an expected value conditional on observations available at time *t*-1.

The relationship (4) gives rise to defining the *output-active inflation, OAI,* as:

$$OAI_t = \pi_t^e - \pi_t^n \quad , \tag{5}$$

so that, interpreting (4), the positive difference between the expected and predicted output-neutral inflations indicates that an increase in output is expected for time *t*. This observation gives rise to using this difference as a simple indicator of possible real effects of pro-inflationary and anti-inflationary shocks.

The practical way of computing π_t^n is illustrated below by the example of a simple two-equation output-inflation vector autoregressive model (VAR). Suppose that such *VAR* model can be written as

$$A(L)Z_t = K + U_t \quad , \tag{6}$$

where $Z_t^{'} = [y_t \quad \pi_t]$, $A(L)$ is the lag polynomial operator, $K' = [k_1 \quad k_2]$ is the vector of constants and $U_t^{'} = [u_{1t} \quad u_{2t}]$ are innovations with zero expectations and variance-covariance matrix Σ.

Since Z_t is stationary, its moving average representation is unique and can be recovered by inverting (6) as:

$$Z_t = M + C(L)U_t \quad , \tag{7}$$

where L is the lag operator, $C(L) = A^{-1}(L) = I + C^{(1)}L + C^{(2)}L^2 + \dots$, I being the identity matrix, and $M = [m_1 , m_2]' = EZ_t = C(1)K$. Then the expected inflation π_t^e defined by (2) can be recovered from (7) by applying the operator E_{t-1} (so that $E_{t-1}(U_t) = 0$) and taking the second component, that is:

$$\pi_t^e = [0,1] \cdot \left(M + \sum_{i=1}^{t-1} C^{(i)} L^i U_i \right) \quad . \tag{8}$$

Recovering the output-neutral predicted inflation π_t^n defined by (3) is based on the methodology suggested by Blanchard and Quah (1989) and then modified further by Gartner and Wehinger (1998) and Charemza and Makarova (2006). Under the assumption of long-run output neutrality of π_t^n, a stationary process Z_t can be decomposed into the unitary innovations given by:

$$Z_t = M + \Gamma(L)\Phi_t \quad , \tag{9}$$

where: $\Gamma(L) = \Gamma^{(0)} + \Gamma^{(1)}L + \Gamma^{(2)}L^2 + \ldots$, $\Phi_t = [\varphi_{1t}, \varphi_{2t}]'$, $E\Phi_t\Phi_t' = I$ and, additionally, with zero restrictions imposed by the long-run output-neutrality of inflation on the upper-right element of the long-run matrix $\Gamma(1) = \sum_i \Gamma^{(i)}$, that is:

$$\Gamma(1) = \Gamma^{(0)} + \Gamma^{(1)} + \Gamma^{(2)} + \ldots = \begin{bmatrix} \gamma_{11} & 0 \\ \gamma_{21} & \gamma_{22} \end{bmatrix} \quad .$$

Matrix $\Gamma(1)$ can be easily computed as the lower-triangular Cholesky factor of $C(1)\Sigma C(1)'$. The element φ_{2t} can be interpreted as output-neutral component of innovations in (9) and therefore vector $\Phi_t^n = [0, \varphi_{2t}]'$ can be interpreted as output-neutral part of unitary innovations Φ_t. The corresponding output-neutral component U_t^n of moving average innovations U_t given by (7) can then be identified by comparing (7) with (9) as:

$$U_t^n = C(1)^{-1}\Gamma(1)\Phi_t^n = C(1)^{-1}\Gamma(1)\begin{bmatrix} 0 & 0 \\ 0 & 1 \end{bmatrix}\Gamma(1)^{-1}C(1)U_t. \tag{10}$$

Then output-neutral component of inflation is recovered by combining (7) with (10) as:

$$\pi_t^n = [0,1] \cdot \left(M + \sum_{i=1}^{t-1} C^{(i)} L^i U_i^n \right) \tag{11}$$

So that, OAI_t can be derived as:

$$OAI_t = \pi_t^e - \pi_t^n = [0,1] \cdot \sum_{i=1}^{t-1} C^{(i)} L^i \left(U_i - U_i^n \right) \quad . \tag{12}$$

Consequently, evaluation of *OAI* consists of (*i*) estimation of the *VAR* model (6) and its moving average representation (7), (*ii*) computing the expected and output-neutral inflations sing (8) and (11) and (*iii*) computing *OAI* from (12).

In order to evaluate the balance of real effects in periods of positive and negative *OAI*'s, asymmetric impulse response analysis has been applied. Impulse response (*IR*) is defined as a response of one variable to an impulse in another variable (see e.g. Hamilton, 1994; Lütkepohl, 2006). Under stationarity (data used here are tested positively for stationarity; see Section 3) *IR*'s are time invariant. Let the impulse response $IR_x(z,h)$ denotes an expected change in *x* in reaction to the shock δ_z of magnitude *v* in variable *z*, after *h* periods (*h*=1,2,…,*H*), that is:

$$IR_x(z,h) = E(x_h \mid \delta_z = v) - E(x_h \mid \delta_z = 0) \quad ,$$

and the cumulative impulse response is $CIR_x(z,H) = \sum_{h=1}^{H} IR_x(z,h)$.

In order to distinguish between inflationary shocks in different forward-looking inflation regimes, the periods of positive and negative *OAI*, defined by (5), we denote $\pi_t^+ = \pi_t$ if *OAI* > 0; 0 otherwise, and $\pi_t^- = \pi_t$ if *OAI* < 0; 0 otherwise. Clearly $\pi_t = \pi_t^+ + \pi_t^-$. Following Hatemi-J (2014), the asymmetric impulse response functions are $IR_y(\pi^+,h)$ and $IR_y(\pi^-,h)$ where: $\pi^+ = \{\pi_t^+\}$; $\pi^- = \{\pi_t^-\}$. This understanding implies that a symmetric shock can produce different outcomes depending on whether it happened in a particular forward-looking inflation regime. Direct linear projection method of Jordà (2005 and 2009) of computing impulse responses has been applied here. The method consists of computing the effects of a shock in time *t* to *t*+*h* by forecasting of y_{t+h} with and without a shock. For the sake of comparison, impulse responses have also been computed from the moving average representation of the *VAR*, through orthogonalizing the errors and collecting relevant coefficients (see e.g. Lütkepohl, 2006). Further in the text the direct projection cumulative *IR*'s are denoted by $CIR_x^D(z,H)$ and orthogonal cumulative *IR*'s by $CIR_x^O(z,H)$. We have computed $CIR_y^D(\pi^+,H)$, $CIR_y^D(\pi^-,H)$, $CIR_y^O(\pi^+,H)$ and $CIR_y^O(\pi^-,H)$ from the 3-equation *VAR*'s formulated for y_t, π_t^+ and π_t^-.[3]

III. Empirical Results

The main database consists of quarterly data on annual *inflation* and annual *GDP growth* for 45 countries, comprising both advanced and developing countries. All of the data were obtained from two main sources: IFS database and the OECD database. The data end in 2011q4 and the length of the series varies between 124 observations (since 1981q1) for most countries to 60 (for Ireland, since 1997q1). For 24 countries in this group (Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, Korea, Mexico, Netherland, Norway, Pakistan, Peru, Philippines, Portugal, Spain, South Africa, Sweden, Switzerland, United Kingdom and United States) complete sets of quarterly data are available from 1981:Q1 or earlier. For the remaining 21 countries (Argentina, Brazil, Chile, Czech Republic, Columba, Hong Kong, Hungary, India, Indonesia, Ireland, Israel, Luxembourg, Malaysia, Morocco, New Zealand, Poland, Romania, Russia, Slovak Republic,

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[3] We have adopted GAUSS procedures written by Òscar Jordà for the computing direct *IR*'s and available at http://www.econ.ucdavis.edu/faculty/jorda/pubs.html and Thierry Roncalli's procedures for the orthogonal *IR*'s (see Roncalli, 1995). The GAUSS program and procedures are available on request.

Thailand and the Turkey) the earlier data has either not been available (often due to the fact that some of these countries, or their separate statistical systems, have been established at a later time), or due a change in the statistical recording system, which makes earlier comparisons impossible. Consequently, we have used data for the period from 1981:Q1 (or later) to 2011Q4 even if for some series earlier data are known. Such choice minimizes the amount of interpolation of the annual data into the quarters, and it excludes periods of the chaotic volatility resulting from oil shocks of the 1970s. Moreover, it makes the panel of data more balanced, as in this case the longest series is about twice long as the shortest series; otherwise this difference would be much bigger, which would affect the comparability. For countries outside the group selected, consistent series of comparable data of the length which merit efficient vector autoregressive analysis, are not available. Within the entire set of data selected, we have identified countries with relatively frequent episodes of high inflation. We have initially defined such countries as such where, within the data span, the 0.75^{th} quantile of inflation was at least equal to 7.5%. In other words, a country with *frequent episodes of high inflation* (*FEHI*) is where in at least 25% cases annual inflation was higher than 7.5%. There are 17 such countries in our dataset. They are listed in Appendix A.

At the initial stage, hypothesis of the stationarity of all the series has been checked. The GLS-detrended and optimal point unit root tests have been applied (see Ng and Perron, 2001 and Perron and Qu, 2007), allowing for the presence and absence of the structural breaks under the null and alternative (see Carrion-i-Silvestre *et al*., 2009).[4] As the possible structural breaks are likely to cause non-normality, the Rachev *et al*.

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[4] We have applied the GAUSS procedures made by Carrion-i-Silvestre and available at http://people.bu.edu/perron/code.html .

(1998) test which tests the $I(1)$ hypothesis against the alternative of $I(0)$ assuming infinite variance of the disturbances has been additionally applied (see Charemza *et al.*, 2005). Detailed results show the prevalence of the stationarity hypothesis for both inflation and output series; respectively 90% of cases for inflation and 86% for *GDP*.[5] There are only two countries for which the null hypothesis of the series being $I(1)$ is not rejected by all the tests for both *inflation* and *GDP*: Chile and Ireland. Neither of these countries enters the initial *FEHI* group.

For each country in the database, we have computed *OAI* using (12). Parameters of the *VAR* model (6) are estimated by the multivariate least squares method. Summary of estimation results is given in Table 1. The moving average representation has been obtained from (7) truncating after the 1,000th elements. The optimal lags of the *VARs* have been selected by the criterion of the minimum autocorrelation of the residuals. This deviates somehow from the established tradition of using information criteria (Akaike and Schwartz Bayesian criteria). The reason for this was that for estimation of *OAI* it is essential to have residuals with a minimum of autocorrelation, as this is a crucial assumption in identifying π_t^e and π_t^n from (8) and (11). The optimal lag length under this criterion is usually shorter than that given by the information criteria, which is important for the relatively short series of data we use. More precisely, as the lag selection criterion we have used the maximum *p*-value of the Hosking (1980) modification of the multivariate Ljung-Box portmanteau test, which seems to have better small sample properties than the alternatives (see Hatemi-J, 2004; for description see Lütkepohl, 2006). Table 1 contains the evaluated lag lengths of the *VARs*, absolute values of the roots of the polynomials of the *VAR* parameters matrices (as measures of *VAR* stability, see Lütkepohl, 2006), *p*-values

[5] These results are not reported here but available on request.

for univariate (separately for output and inflation) and bivariate (joint) Ljung-Box autocorrelation portmanteau statistics, and *p*-values of Jarque-Bera normality statistics. For all countries the minimal root of the polynomial is outside the unit circle, which indicates stability. For the overwhelming majority of countries there is no indication of autocorrelation in residuals of individual series, although the results of the joint test are less favourable. As it is known that distributions of inflation and output growth are rarely normal, the *p*-values of the normality statistics for the residuals often suggest non-normality, which in turn could make further testing more difficult.

INSERT TABLE 1 ABOUT HERE

After evaluating *OAI* for all 45 countries, cumulative (for 24 periods) asymmetric, that is separate for π_t^+ and π_t^-, impulse responses of the inflationary shocks on output, have been computed by two methods: direct projection and orthogonalization. Tables 2 and 3 contain the results of the evaluation of the cumulative impulse responses from direct projection (Table 2) and the orthogonal decomposition (Table 3).

INSERT TABLES 2 AND 3 ABOUT HERE

Table 2 also contains the *F*-statistics and their corresponding *p*-values for the hypothesis that $CIR_y^D(\pi^+,24) = CIR_y^D(\pi^-,24)$ and $V\{IR_y^D(\pi^+)\} = V\{IR_y^D(\pi^-)\}$; see Jordà (2009). More precisely, it gives the cumulative impulse responses obtained by direct projection, that is $CIR_y^D(\pi^+,24)$, $CIR_y^D(\pi^-,24)$, cumulative variance decomposition of particular shocks in proportion of the total cumulative variance of y_t denoted as $V\{IR_y^D(\pi^+)\}$, $V\{IR_y^D(\pi^-)\}$, and Jordà's (2009) statistics (with *p*-values) for testing the null hypotheses that (*i*) $IR_y^D(\pi^+,h) = IR_y^D(\pi^-,h)$ jointly for all *h* (the *joint test*), and (*ii*) $CIR_y^D(\pi^+,24) = CIR_y^D(\pi^-,24)$ (the *cumulative test*). Table 3 shows the orthogonal cumulative impulse responses $CIR_y^O(\pi^+,24)$, $CIR_y^O(\pi^-,24)$ and corresponding variance decompositions $V\{IR_y^O(\pi^+)\}$, $V\{IR_y^O(\pi^-)\}$. The joint significance test rejects the null at the 10% level for only 4 countries: Hong Kong, Israel, Peru and Slovak Republic. The cumulative test rejects the null more frequently: for Belgium, Finland, Hong Kong, Ireland, Morocco, Philippines, Slovak Republic, Spain, Turkey, UK and USA. Likely reason for such surprisingly low level of significant results can be the underlying assumption that the impulse responses have joint multivariate normal distribution which, in case of relatively short time series and clearly non-normal distribution of *VAR* residuals, might be somewhat stretchy.

More detailed results of *OAI* estimation for three representative Asian countries, Indonesia, Malaysia, and Pakistan, are presented below.[6] During the period investigated Indonesia and Pakistan exhibit evidence of high inflation and, using the classification introduced in Section 2, are regarded as countries with frequent episodes of high inflation (*FEHI*), while Malaysia, with markedly lower average inflation, is used as a benchmark for comparison. Below we outline briefly the development of inflation and causes for its increases in these three countries.

Indonesia

Indonesia was in a deep economic recession due to the 1997-98 Asian financial crisis. As a result, it experienced a massive depreciation in its currency causing the stock market to collapse. The economy was in unstable financial position because of Indonesian corporations' foreign currencies borrowing practices without hedging

[6] Results for other countries are available on request.

against devaluation. The rate of inflation increased sharply and reached about 80% in mid-1997. In response, Bank of Indonesia raised the interest rate to around 70%. Indonesian GDP growth rapidly declined witnessing negative economic growth of over 13% in 1998. After the crisis, Indonesia has introduced a wide range of institutional reforms and redirected monetary policy towards maintaining price and exchange rate stability. As a result, price stability has been, to an extent, reinstated. However, the annual economic growth rate in 2001 slipped to about 3.5% with the inflation rate of around 13%. In the fourth quarter of 2005 Indonesia experienced a minor crisis due to international oil shock coupled with high imports. The Indonesian government was forced by *IMF* to cut its oil subsidies to stabilize the economic situation, but the economy responded by sharp inflation rise of 17%. After that, economic growth started to increase. The Bank of Indonesia had officially launched its inflation targeting policy in July 2005. In the wake of the economic crisis, the Bank of Indonesia has been granted independence as part of conditionality of the International Monetary Fund's rescue package. It is now regarded as a country belonging to the so-called inflation control group (see Lin and Ye, 2009, but definitions and classifications vary; see e.g Brito and Bystedt, 2010).

Malaysia

Unlike Indonesia and Pakistan, Malaysian economy has not experienced episodes of substantially high inflation. Since 1991 inflation rate averaged 2.9%. In 1990, oil price shock as a result of Gulf war increased Malaysian inflation merely to 4.75% in 1991. Malaysia has been comparatively successful in balancing strong economic growth with moderate levels of inflation in the periods preceding and following the Asian financial crisis. During the Asian crisis in 1997-98 inflation was well controlled and increased only to around 5%. After facing an economic recession for about two years since 1997, Malaysian economy has begun to pick up again from the third quarter of 1999. Inflation rate started to accelerate slightly since 2005 when the world oil prices rose, but it exceeded 5% only occasionally.

Pakistan

Low and moderate inflation had been typical for the Pakistan economy until the end of 2007. Average annual inflation was above 11% for only 8 out the past 28 years. Average annual real per capita income growth was 2.8%. However, years after 2007 have been more turbulent. Inflation triggered by increasing worldwide petrol prices reached 25% in the second half of 2008. In 2009-2011, inflation was slightly reduced but was still above 10%, due to increase in agriculture prices and industrial uncertainties caused by political instability. At the same time, the GDP growth was remarkably stable, at around 7.5% with little variation.

Figure1 shows confidence intervals (± two standard deviations around the computed value of *OAI*) obtained by pairwise bootstrap applied to the residuals of the *VAR* model for 1,000 resamplings. For most periods, the confidence intervals include zero, which means that the hypothesis that the true values of *OAI* is equal to zero cannot be rejected. However, for Indonesia *OAI* is highly significant for the period 1998q3-1999q1. Inflation in this period was not markedly higher than for the remaining quarters of 1998 and 1999. For Malaysia there are some signs of significance of *OAI* for 1995q2-1997q3, and for Pakistan for 1997q2-q3 and 2007q3-2008q1. For Malaysia, as for Indonesia, inflation in the period of significant *OAI* was in line with inflation in the neighbouring quarters. For Pakistan, in 1997*, OAI* significance corresponds to a local peak in inflation and, for 2007q3-q4, it coincides with a period of gradually rising inflation, which reached its peak in the second half of 2008.

INSERT FIGURE 1 ABOUT HERE

Figure 2 presents $IR_y^D(\pi^+,h)$ and $IR_y^D(\pi^-,h)$, h = 1, 2,…, 24, together with confidence intervals around $IR_y^D(\pi^+,h)$ for the representative countries: Indonesia, Malaysia and Pakistan. If $IR_y^D(\pi^-,h)$ is outside such intervals, it is roughly interpreted as individual significance of the differences between $IR_y^D(\pi^+,h)$ and $IR_y^D(\pi^-,h)$. We present the simultaneous Scheffé bands and conditional Jordà bands (for the detailed description of both, see Jordà, 2009). The Scheffé bands are in the form of a fan-chart (respectively with 95%, 50% and 25% confidence intervals) and Jordà bands are for the 90% confidence interval. The reason for plotting different Scheffé bands is due to their construction as simultaneous bands, particular different intervals might cross, so that presenting different confidence intervals gives a clearer picture of the uncertainty related to the impulse responses.

INSERT FIGURE 2 ABOUT HERE

Despite the fact that the conditional Jordà bands are narrower than the marginal bands (not reported here) or Scheffé bands, they still include zero for most of the cases. For Pakistan, for $IR_y^D(\pi^+,h)$ increases and becomes significantly positive for the horizons of 3 to 5 quarters (according to Scheffé bands) and 21 to 23 quarters, (according to both Scheffé and Jordà bands). For the same horizons, $IR_y^D(\pi^-,h)$ decreases and becomes negative. Hence, it can be concluded that for Pakistan there is a positive difference in real effects of inflationary shocks in the periods of positive and negative *OAI*'s after 3-5 quarters, with a possible additional long-delayed effect after 21-23 quarters. Similar pattern can be observed for Indonesia for the horizons of 8-10 quarters. For Malaysia, a country without high inflation episodes, impulse responses are mainly insignificant, except for the horizons of 16-20 quarters, where $IR_y^D(\pi^-, h)$ is significant and higher than $IR_y^D(\pi^+, h)$.

As an aggregate benchmark of the balance between the real effects in periods of positive and negative *OAI*'s a simple measure of *OAI* gain (*IGAIN*) defined as $IGAIN^i = CIR_y^i(\pi^+, H) - CIR_y^i(\pi^-, H)$, i={D,O} has been computed. It is interpretable as the total real gain (in the sense of output) from inflationary shock which takes place in the period of positive *OAI* in relation to the same happening in the period of negative *OAI*.

For comparison, we have checked whether gains in periods of high and low inflation gives the real effects similar to that described by *IGAIN*. We have defined gains from high inflation (*HGAIN*) as $HGAIN^i = CIR_y^i(\pi^{M+}, H) - CIR_y^i(\pi^{M-}, H)$, i={D,O}, where $CIR_y^i(\pi^{M+}, H)$, $CIR_y^i(\pi^{M-}, H)$ denote respectively the cumulative impulse responses of output on inflationary shocks in the periods where inflation is above and below its median. The technique used for computing *HGAIN* is analogous to that explained in Section 2. For each country a 3-equation *VAR* for y_t, π_t^{M+} and π_t^{M-} has been formulated, where $\pi_t^{M+} = \pi_t$ if $\pi_t > median(\pi_t)$; and 0 otherwise and $\pi_t^{M-} = \pi_t - \pi_t^{M+}$. This model is then estimated and used for evaluation of *HGAIN*. Clearly, if, for a given country, all periods of positive and negative *OAI* correspond exactly to the periods of inflation being higher and lower than the median, then *IGAIN*=*HGAIN*. Figures 3 and 4 show the correlation of the logarithms of the 0.75th quantile of inflation with *IGAIN* and *HGAIN* respectively. On Fig. 3, the existence of positive correlation is evident. However, analogous correlation of inflation with the balance (differences) of cumulative impulse responses in periods of inflation being above and below the median, shown at Fig. 4, is negative.

INSERT FIGURES 3 AND 4 ABOUT HERE

Somewhat stylized reflection here could be that the best situation (in terms of output stimulation) in *FEHI* countries is where positive inflationary shocks occur in periods of positive *OAI*. It seems to be in line with the Bruno and Easterly (1996 and 1998) conjecture that high inflation stimulates growth. However, our interpretation is more modest. We only claim that high inflation might be a lesser obstacle to growth in an economy with frequent high inflation episodes if shocks happen in the period of positive *OAI*. Moreover, the higher inflation becomes during such episodes, the stronger is such effect. This finding is in line with the results obtained by Fidrmuc and Tichit (2013) for a smaller group of countries and with the use of different methodology and data. Our results do not contradict findings related to the existence of the optimal inflation threshold. Such threshold might indeed exist. However, if inflation is above such threshold and the regime is of a negative *OAI*, anti-inflationary effects of a contractionary monetary policy may harm output in a lesser way than in the periods of positive *OAI*.

It has also been checked to what extent the results depend on our, rather arbitrary, definition of the episodes of high inflation. Perhaps the positive relationship between *OAI* and inflation holds even if the 'high inflation' is not really that high after all? For checking this, the 7.5% high inflation limit for *FEHI* has been gradually lowered so that the *FEHI* group incorporates more countries. First, the country with the highest 0.75^{th} quantile of inflation in the non-*FEHI* group is included to the FEHI group, then the country with the second highest quantile is added, *etc.*. For this gradually enlarging group correlation coefficients as in Fig. 3, between *IGAIN* and the 0.75^{th} quantile of inflation, have been computed. The results are shown at Fig. 5.

INSERT FIGURE 5 ABOUT HERE

This figure indicates that, generally, the higher is the 0.75th quantile of inflation, the higher is the correlation of its logarithm with the inflationary real effect for periods of positive *OAI*, in comparison with periods of negative *OAI*. It also shows that the results are reasonably tolerant regarding the definition of *FEHI*. If the definition is relaxed and the *FEHI* country is redefined as such where the 0.75th quantile of inflation is greater than 4.8% rather than 7.5%, the main result of the study, that the correlation between the logarithm of this quantile and *IGAIN* remains high, holds unchanged. However, if we relax the *FEHI* definition further still, this correlation weakens markedly. This confirms indirectly the Hartmann and Roestel (2013) finding given in Section 1 that rising inflation in countries with already high inflation is not as bad for output as in countries with low inflation. Our results extend this; it has been shown above that the lower 0.75th quantile of inflation becomes, the lower is an increase in real gain in periods of high inflation.

IV. Conclusions and Simple Policy Prescriptions

Our results suggest a way of making the most of rising inflation in countries where there are already frequent episodes of high inflation by undertaking anti-inflationary monetary decisions in periods of negative forward-looking inflation regime, when the difference between the expected and predicted output-neutral inflations is negative. Analogously, the output-stimulating policy should have the best effects if undertaken in the reverse situation, when there is a positive difference between the expected inflation and output-neutral inflation. Somewhat more general reflection here, in the mood of Bruno and Easterly (1996 and 1998), is that high, and even increasing, inflation might not necessarily be bad for growth if the timing of applying the brakes is wise. More importantly, it can facilitate further institutional reforms leading to further recovery (see Drazen and Easterly, 2001; Cavallo and Cavallo, 2010).

We have identified the following limitation of the proposed approach. Firstly, our findings are valid for most countries with markedly high inflation (over 4.8% in at least every fourth quarter on average) and not for countries with intrinsically lower inflation. Secondly, it is not conclusive whether inflationary shocks in the periods when expected inflation exceeds output neutral inflation increases or decreases output volatility.

The model we use is very simple and with an obvious room for improvement. Output neutral inflation can be computed in a much more sophisticated way from disaggregated components of output and inflation or, as some measures of core inflation are constructed, by identifying price-controlled components in the consumers' price index. If a disaggregated model is used and, presumably, when the assumption of the multivariate normal distribution is relaxed, impulse response analysis and testing can be done more precisely. We are leaving this for further research.

Appendix A. Basic characteristics of the dataset

The dataset consists of data on *GDP growth* and *inflation* for 45 countries. The *GDP growth* is defined as the percentage change of the real *GDP* in a given quarter over the real *GDP* in the corresponding quarter of the previous year. Inflation is defined by the percentage change of the consumer price index (*CPI*) over the last year's level in the corresponding quarter. Real *GDP* figs have been computed by deflating the nominal *GDP* by each country's *GDP* deflator (source: from *IMF* International Financial Statistics, IFS, http://esds80.mcc.ac.uk/ wds_ifs/) except for Indonesia, where the consumers' price index, *CPI*, has been used as the deflator. For countries other than that of OECD and Brazil, India, Indonesia, Russia and South Africa, *inflation* has been computed from the original *CPI* data. Data on *inflation* for the 30 OECD countries and 5 non-OECD countries listed above are from the OECD (http://stats.oecd.org/). The *GDP* deflators for all 45 countries are from the IFS. Data for the nominal *GDP* for the non-OECD countries except for Brazil, India, Russia and South Africa have been obtained from the IFS, and for the remaining countries from OECD. For India and Pakistan some quarterly *GDP* data are converted from annual to quarterly frequencies using the polynomial quadratic interpolation. For India, annual *GDP* is interpolated for the period from 1991q1 to 1996q1, with the remaining data in this series from OECD. The annual nominal *GDP* series for India and Pakistan have been obtained from the OECD and IFS respectively, while the *GDP* deflators are from IFS.

INSERT TABLE A1 ABOUT HERE

Appendix B. Country symbols used in figs 3 And 4

Country	Symbol	Country	Symbol
Argentina	AR	Romania	RO
Hong Kong	HK	Russia	RU
Indonesia	IA	Slovak Republic	SR
Mexico	ME	Slovenia	SV
Peru	PE	Turkey	TR

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Table 1. Summary of estimation results

Country	VAR lag	Root	Ljung-Box, *p*-values			Jarque-Bera, *p*-values	
			Output	Inflation	Joint	Output	Inflation
Argentina	7	1.093	0.812	0.981	0.391	0.000	0.000
Australia	6	1.052	0.926	0.629	0.071	0.207	0.000
Austria	7	1.131	0.161	0.820	0.010	0.000	0.000
Belgium	8	1.137	0.666	0.492	0.000	0.000	0.208
Brazil	5	1.266	0.954	0.096	0.369	0.000	0.000
Canada	8	1.081	0.190	0.598	0.000	0.068	0.197
Chile	5	1.132	0.906	0.899	0.051	0.002	0.858
Czech Republic	6	1.086	0.939	0.087	0.097	0.107	0.379
Columbia	6	1.113	0.193	0.585	0.002	0.587	0.367
Denmark	8	1.102	0.007	0.369	0.000	0.487	0.508
Finland	8	1.066	0.781	0.732	0.032	0.000	0.724
France	8	1.094	0.100	0.301	0.000	0.043	0.980
Germany	8	1.096	0.812	0.466	0.002	0.000	0.000
Hong Kong	8	1.071	0.918	0.823	0.150	0.081	0.002
Hungary	3	1.108	0.375	0.398	0.047	0.003	0.735
India	8	1.063	0.960	0.901	0.173	0.000	0.095
Indonesia	8	1.078	0.196	0.931	0.006	0.000	0.000
Ireland	2	1.165	0.229	0.365	0.307	0.017	0.458
Israel	6	1.093	0.619	0.644	0.017	0.305	0.077
Italy	8	1.115	0.359	0.866	0.021	0.011	0.001
Japan	8	1.067	0.168	0.370	0.001	0.000	0.008
Korea	8	1.082	0.521	0.043	0.001	0.000	0.000
Luxembourg	6	1.113	0.674	0.989	0.337	0.739	0.937
Malaysia	8	1.118	0.543	0.980	0.091	0.000	0.000
Mexico	8	1.054	0.456	0.984	0.010	0.000	0.000
Morocco	5	1.050	0.044	0.672	0.006	0.119	0.000
Netherland	7	1.138	0.192	0.499	0.020	0.000	0.896
New Zealand	6	1.213	0.942	0.514	0.007	0.745	0.000
Norway	8	1.045	0.232	0.431	0.004	0.385	0.003
Pakistan	8	1.096	0.001	0.701	0.000	0.000	0.000

Peru	8	1.084	0.245	0.670	0.000	0.000	0.000
Philippines	8	1.140	0.475	0.823	0.250	0.000	0.000
Poland	4	1.096	0.906	0.444	0.090	0.038	0.598
Portugal	4	1.099	0.027	0.000	0.000	0.000	0.000
Romania	4	1.250	0.912	0.492	0.049	0.000	0.000
Russia	5	1.255	0.632	0.955	0.028	0.046	0.000
Slovak Republic	7	1.131	0.982	0.730	0.096	0.000	0.000
Spain	8	1.001	0.135	0.995	0.003	0.000	0.419
South Africa	7	1.050	0.181	0.486	0.006	0.010	0.107
Sweden	8	1.069	0.811	0.337	0.000	0.000	0.000
Switzerland	7	1.105	0.633	0.691	0.036	0.027	0.745
Thailand	6	1.116	0.642	0.401	0.022	0.478	0.258
Turkey	4	1.040	0.176	0.037	0.002	0.000	0.000
United Kingdom	8	1.113	0.058	0.504	0.001	0.163	0.000
United States	8	1.095	0.202	0.904	0.000	0.389	0.000

Table 2. Cumulative impulse responses from direct projections

Country	(1)	(2)	(3)	(4)	Joint test		Cumulative test	
					F-stat	*p*-Value	F-stat	*p*-value
Argentina	11.359	6.225	0.311	0.304	24.510	0.474	0.171	0.682
Australia	-2.145	0.270	0.595	0.354	16.782	0.838	1.810	0.182
Austria	-0.021	-0.165	0.659	0.349	18.564	0.757	0.008	0.929
Belgium	6.033	-7.574	0.571	0.195	16.000	0.868	4.501	0.037
Brazil	2.661	1.009	0.449	0.250	38.819	0.120	0.786	0.384
Canada	-3.023	0.663	0.432	0.492	23.464	0.504	2.747	0.102
Chile	-0.028	-0.341	0.395	0.345	31.538	0.265	0.072	0.791
Czech Republic	0.317	0.133	0.457	0.508	39.574	0.117	0.008	0.930
Columbia	1.877	-0.237	0.296	0.345	19.410	0.699	0.638	0.431
Denmark	0.366	-1.475	0.491	0.196	24.026	0.476	0.871	0.354
Finland	1.363	-8.251	0.516	0.164	20.960	0.635	4.527	0.037
France	1.553	0.154	0.518	0.233	20.357	0.666	0.679	0.413
Germany	-1.222	0.896	0.424	0.329	26.724	0.351	0.779	0.380
Hong Kong	7.226	-1.020	0.647	0.507	38.556	0.064	2.800	0.099
Hungary	5.177	3.426	0.347	0.174	30.099	0.278	0.551	0.464
India	-1.622	-0.581	0.535	0.180	32.699	0.198	0.414	0.524
Indonesia	-12.537	2.578	0.385	0.320	23.403	0.515	1.964	0.169
Ireland	-3.771	13.093	0.534	0.363	13.580	0.913	4.013	0.057
Israel	-1.204	0.260	0.512	0.592	44.445	0.072	0.405	0.531
Italy	-1.414	-0.557	0.528	0.229	28.562	0.279	0.517	0.475
Japan	-0.271	-2.641	0.561	0.681	11.762	0.974	0.348	0.557
Korea	6.688	5.493	0.493	0.336	30.544	0.213	0.035	0.852
Luxembourg	-0.663	0.143	0.453	0.272	23.203	0.533	0.094	0.762
Malaysia	0.175	1.526	0.618	0.355	28.560	0.314	0.311	0.581
Mexico	-0.422	2.156	0.576	0.316	16.384	0.853	0.263	0.610
Morocco	1.879	-5.756	0.564	0.462	18.333	0.758	3.932	0.053
Netherland	2.232	-2.339	0.605	0.311	15.697	0.879	2.545	0.115
New Zealand	-3.327	-1.852	0.337	0.434	14.890	0.899	0.294	0.590
Norway	-1.448	0.210	0.767	0.451	15.962	0.869	0.772	0.382
Pakistan	-0.272	-2.300	0.554	0.171	27.482	0.320	0.782	0.379

Peru	5.059	-11.855	0.358	0.812	56.614	0.003	2.386	0.127
Philippines	4.037	-4.809	0.519	0.190	26.053	0.380	5.568	0.021
Poland	0.844	-1.106	0.382	0.343	17.175	0.795	0.536	0.471
Portugal	1.110	0.537	0.672	0.434	16.161	0.863	0.048	0.826
Romania	-8.289	-21.994	N/A	0.174	20.312	0.658	0.174	0.680
Russia	2.023	0.921	0.559	0.091	34.767	0.182	0.064	0.802
Slovak Republic	1.289	-4.518	0.735	0.344	45.594	0.050	3.649	0.066
Spain	-0.006	-4.953	0.728	0.447	20.106	0.679	2.905	0.092
South Africa	-1.205	-2.179	0.563	0.666	18.293	0.770	0.104	0.748
Sweden	2.454	1.693	0.710	0.199	26.706	0.352	0.169	0.683
Switzerland	-0.867	-0.462	0.476	0.207	23.057	0.525	0.051	0.822
Thailand	0.098	2.798	0.488	0.311	24.754	0.462	0.508	0.481
Turkey	15.824	-6.553	0.435	0.252	30.643	0.221	3.538	0.065
United Kingdom	3.751	-5.682	0.620	0.102	32.135	0.170	11.819	0.001
United States	-5.028	-0.527	0.497	0.512	21.435	0.610	3.844	0.054

Legend: (1) $CIR_y^D(\pi^+, 24)$, (2) $CIR_y^D(\pi^-, 24)$, (3) $V\{IR_y^D(\pi^+)\}$, (4) $V\{IR_y^D(\pi^-)\}$

Table 3. Cumulative orthogonal impulse responses

Country	(1)	(2)	(3)	(4)
Argentina	4.536	23.227	0.017	0.414
Australia	-1.484	0.031	0.129	0.009
Austria	0.236	-0.665	0.034	0.093
Belgium	-1.527	-4.818	0.038	0.168
Brazil	-1.044	-0.593	0.037	0.175
Canada	-2.500	1.033	0.061	0.050
Chile	-1.179	-3.157	0.109	0.375
Czech Republic	-0.311	-4.541	0.050	0.119
Columbia	0.073	-2.181	0.139	0.186
Denmark	-0.082	-4.676	0.010	0.237
Finland	-0.585	-10.661	0.056	0.238
France	-0.915	-1.231	0.041	0.284
Germany	-0.546	-0.029	0.081	0.073
Hong Kong	0.416	-0.036	0.033	0.075
Hungary	-1.856	0.953	0.024	0.020
India	-2.273	-1.755	0.189	0.181
Indonesia	-12.957	-2.933	0.255	0.111
Ireland	1.133	-4.438	0.012	0.108
Israel	-1.203	0.783	0.065	0.098
Italy	-1.086	0.449	0.072	0.057
Japan	-0.519	-3.084	0.058	0.222
Korea	1.810	-1.193	0.018	0.038
Luxembourg	1.710	-2.929	0.030	0.082
Malaysia	-3.225	-0.256	0.145	0.263
Mexico	1.225	0.550	0.012	0.039
Morocco	0.133	-2.090	0.039	0.058
Netherland	-1.160	-0.915	0.019	0.026
New Zealand	0.322	-4.563	0.014	0.168
Norway	-2.710	0.331	0.067	0.014
Pakistan	0.505	-2.832	0.107	0.069
Peru	5.388	-16.713	0.147	0.236

Philippines	1.298	-4.521	0.042	0.051
Poland	-0.498	-2.313	0.067	0.140
Portugal	-0.101	3.260	0.046	0.117
Romania	-37.515	-41.623	0.048	0.072
Russia	0.667	-1.619	0.024	0.021
Slovak Republic	0.839	-7.061	0.063	0.224
Spain	-0.764	-3.015	0.012	0.071
South Africa	-1.886	-3.496	0.077	0.193
Sweden	2.290	-3.658	0.066	0.092
Switzerland	-2.120	-1.719	0.042	0.071
Thailand	-1.209	-1.083	0.087	0.070
Turkey	6.620	-8.318	0.136	0.023
United Kingdom	1.992	-4.350	0.071	0.162
United States	-2.134	0.432	0.122	0.063

Legend: (1) $CIR_y^o(\pi^+, 24)$, (2) $CIR_y^o(\pi^-, 24)$, (3) $V\{IR_y^o(\pi^+)\}$, (4) $V\{IR_y^o(\pi^-)\}$.

Table A1. Basic data characteristics and sources of data for *FEHI* countries

Country	N. obs.	First obs.	Last obs.	Inflation		Data source for:		
				Average	0.75 quantile	Inf.	Deflator	Nom. GDP
Argentina	76	1993q01	2011q04	6.650	9.553	B	B	B
Brazil	68	1995q01	2011q04	14.440	7.794	A	B	A
Columbia	72	1994q01	2011q04	10.060	17.770	B	B	B
Hong Kong	121	1981q04	2011q04	4.570	9.164	B	B	B
Hungary	68	1995q01	2011q04	9.680	10.600	A	B	A
India	84	1991q01	2011q04	7.760	10.130	A	B,D	A,D
Indonesia	88	1990q01	2011q04	11.080	10.210	A	B,E	B
Mexico	124	1981q01	2011q04	30.220	33.770	A	B	A
Pakistan	124	1981q01	2011q04	8.510	10.920	B	B,D	B,D
Peru	124	1981q01	2011q04	421.00	88.080	B	B	B
Philippines	124	1981q01	2011q04	8.944	10.030	B	B	B
Poland	68	1995q01	2011q04	7.287	9.886	A	B	A
Portugal	124	1981q01	2011q04	8.107	11.490	A	B	A
Romania	69	1994q04	2011q05	29.9575	42.2708	B,C	B,C	B,C
Russia	68	1995q01	2011q04	32.100	20.950	A	B	A
Slovak Republic	76	1993q01	2011q04	7.134	8.331	A	B	A
Turkey	100	1987q01	2011q04	47.600	69.800	A	B	A

Legend:

A: data source: OECD

B: data source: IFS, inflation recomputed from CPI data

C: 1995q1-2000q4: data obtained directly from the Romanian Central Statistical Office

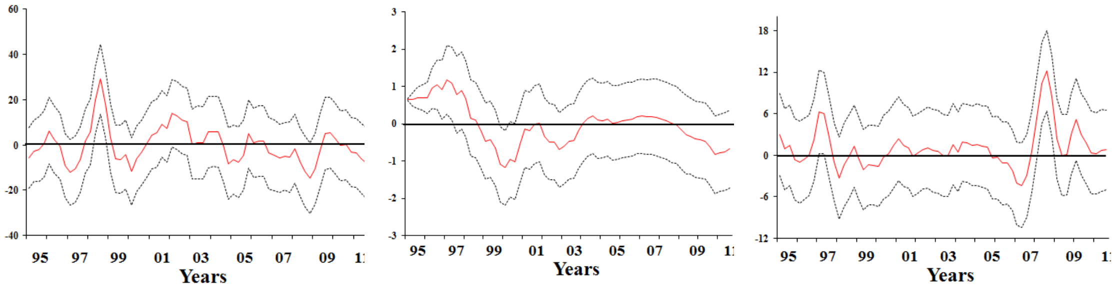
D: Interpolated from annual data

E: CPI index used as the deflator

Entries for countries included in the FEHI group are boldfaced

Fig. 1. Estimates of output-active inflation, *OAI*

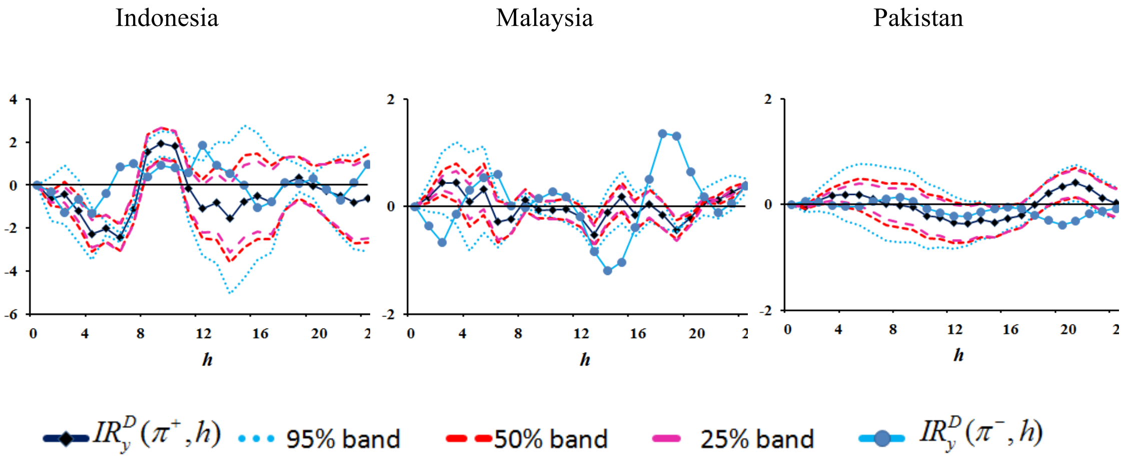
Indonesia Malaysia Pakistan



Legend: *OAI* is represented by the middle line, between 2× SD bootstrapped confidence intervals.

Fig. 2. Direct projection impulse responses of y_t to shocks in π_t^+ and π_t^- with their confidence bands

Scheffé bands



Jordà bands

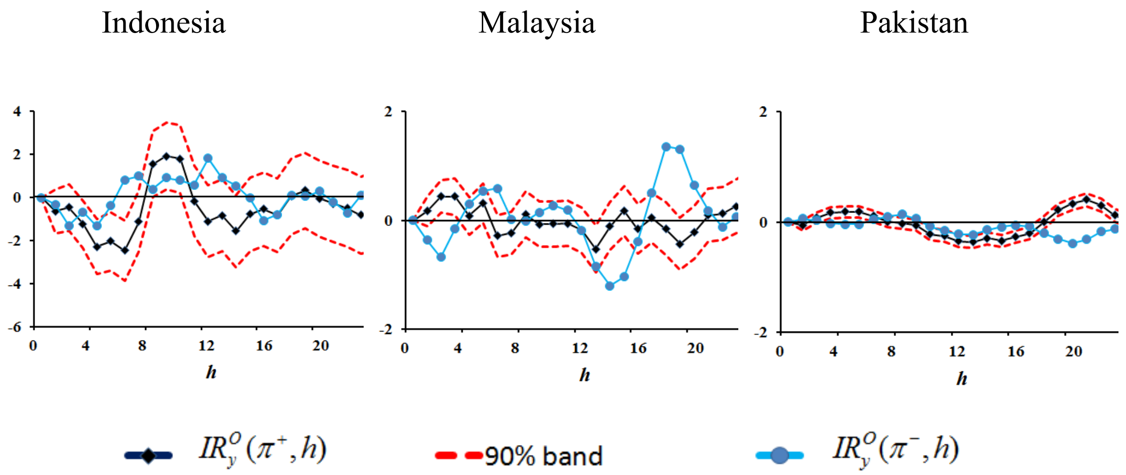
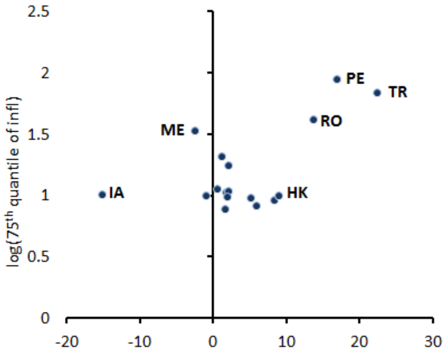
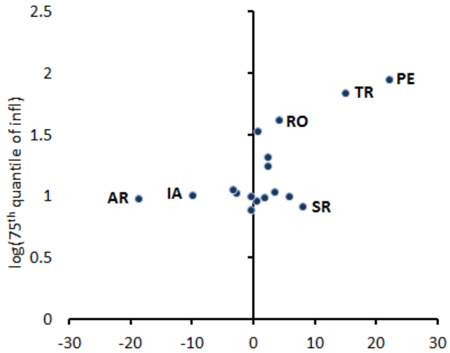


Fig. 3. Correlation between the log of 0.75th quantile of inflation and *IGAIN* for *FEHI* countries

$IGAIN^{D}$ $IGAIN^{O}$



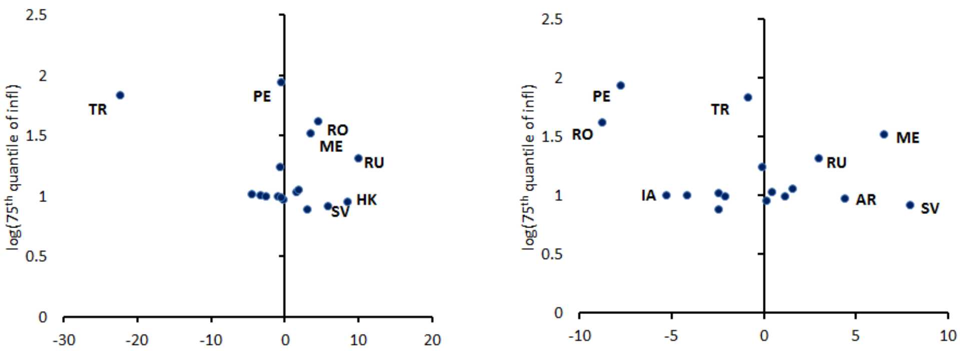


corr.coefficient = 0.61 corr.coefficient=0.66

Legend: For some correlation points, country symbols, explained in Appendix B, are printed.

Fig. 4. Correlation between the log of 0.75th quantile of inflation and *HGAIN* for *FEHI* countries

$HGAIN^D$ *$HGAIN^O$*

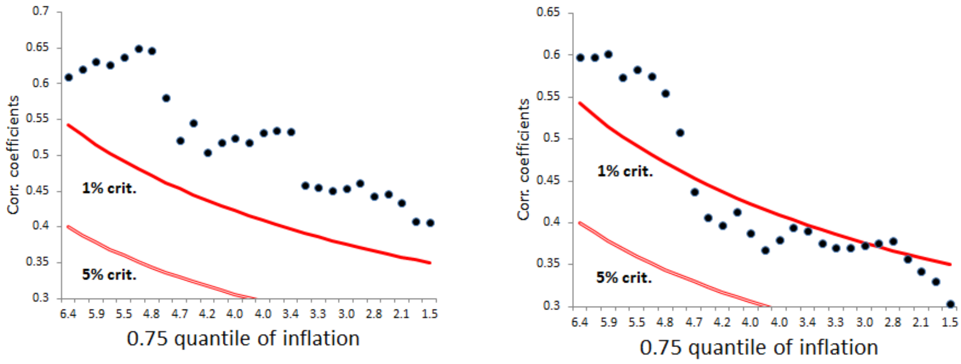


corr.coefficient = -0.30 corr.coefficient=-0.26

Legend: For some correlation points country symbols, explained in Appendix B, are printed.

Fig. 5. Correlation coefficients of *IGAIN* and log of 0.75th quantile of inflation:

$IGAIN^D$ $IGAIN^O$



Legend: *FEHI* group is increasing by gradually lowering the 0.75th percentile from 6.4% to 1.5%. Solid upper line represents the upper critical bound of the correlation coefficient around zero at 1% level of significance, and the lower line at 5% level of significance.