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Publication bias in the price effects of monetary policy: A meta-regression analysis for emerging and developing economies

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ABSTRACT

Using 43 studies conducted between 2001 and 2019, we employ a meta-regression analysis (MRA) to synthesize literature findings on the effects of monetary policy on price levels in 32 emerging and developing countries. We find strong evidence of a negative publication bias for all types of price effects (short-term, medium-term and maximum effects). Primary studies published in academic journals tend to report stronger negative effects. A cluster analysis and a mixed-effect multilevel model confirm the null hypothesis of a genuine price effect. Employing the "best practice" method, we find that the genuine effect is negative. In the other words, increasing policy interest rates appears to be effective in controlling inflation in emerging and developing countries. In comparison with the genuine price effect in advanced countries reported by Rusnak et al. (2013), our study indicates that the genuine price effects in emerging and developing countries are weaker than in advanced countries.

1. Introduction

Price stability is widely considered as one of the primary objectives of monetary policy (Friedman, 1995). A tightening of monetary policy (e.g. captured by an increase in interest rates) is largely perceived by policymakers as an effective tool to curb inflationary pressures (Mishkin, 1995; Christiano et al., 1999). Since the early 1990s, an increasing number of countries adopted an inflation-targeting framework when designing their monetary policies[1] (Bernanke & Mishkin, 1997; Roger, 2010). Consequently, there has been considerable interest in measuring the effect of a tighter monetary policy on price levels. To this end, and initiated by Sims (1980), vector auto-regressive models (VARs) have been widely used to assess such impacts (Walsh, 2017). Unfortunately, the empirical evidence to date has been very inconsistent and inconclusive. Fig. 1 demonstrates the large heterogeneity in measured price responses (of a one-percentage point increase in the interest rate) based on estimates from the 43 studies included in our meta-analysis (all of which have an explicit focus on emerging and developing economies; 32 in total). Although the majority of reported responses (clustered per year of publication for the sake of easier exposition) correspond to negative effects, the results are very diverse both with respect to sign as well as magnitude. This revealed heterogeneity has prompted a vivid academic debate on the sign of effect (Hanson,

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[1] According to (Roger, 2010), there were 26 countries until 2010 that explicitly adopted an inflation targeting framework for their monetary policies.

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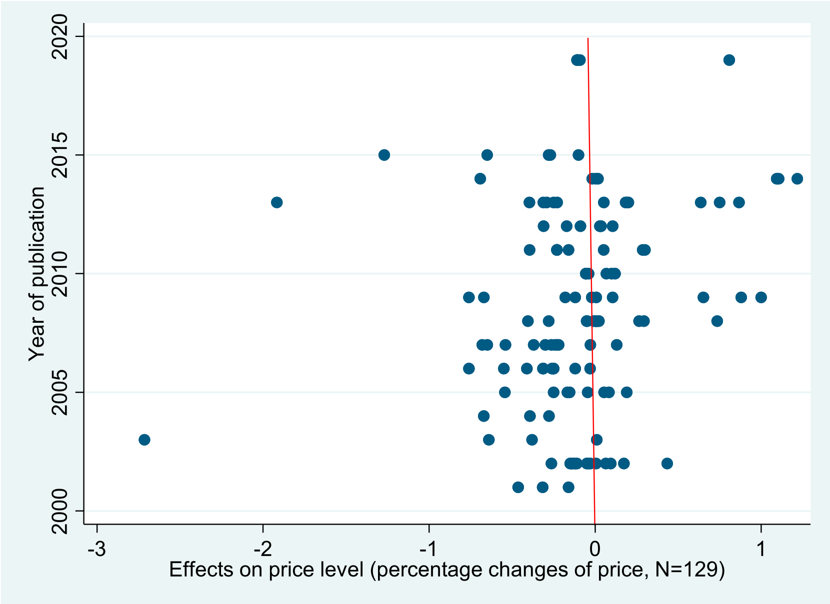

Fig. 1. Heterogeneity in price responses (after 12 months) of a one-percentage point increase in the interest rate across emerging and developing countries.

2004; Sims, 1992), the applied methodology (Bernanke et al., 2005; Boivin et al., 2010; Kim & Roubini, 2000) and the role of country/structural factors (De Haan & Kooi, 2000; Friedman & Woodford, 2010; Mishra & Montiel, 2013).

Several scholars (De Long & Lang, 1992; Doucouliagos & Stanley, 2009; Stanley, 2005, 2008) have emphasized the role of a publication selection bias in the observed heterogeneity in reported effects. A publication bias (where editors, referees and authors often prefer larger and more significant estimates) has been found to exaggerate genuine effects in several research fields (e.g. see Ioannidis et al., 2017 for the summary of publication bias in economics; Doucouliagos & Stanley, 2009 for publication bias in minimum wages; Stanley, 2005 for publication bias in trade union productivity and price elasticities). Hence, it is possible that a publication selection bias may also be present in the case of the monetary policy effects on price levels. In recent years, the development of meta-regression analysis (MRA) has allowed researchers to detect publication bias and correct for it. The key advantage of the MRA method is that it synthesizes and explains variation in a logical, transparent, and statistical way (Stanley & Jarrell, 1989; Stanley, 2001; DouCouliagos, 2016). These features are useful and important when investigating the genuine effects of monetary policy (and the corresponding heterogeneity of observed estimates).

Employing the MRA method, several scholars (e.g, De Grauwe & Costa Storti, 2004; Ridhwan et al., 2010; Havranek and Rusnak, 2012; Rusnák, Havranek, and Horvath, 2013; Papadamou et al., 2019) have synthesized the effects of monetary policy in advanced countries. In the developing countries, the effect of monetary policy is less predictable and effective (Mishra et al., 2010) necessitating hence further research in the field. Mishra et al. (2010) and Mishra and Montiel (2013) survey the impact of monetary policy in developing countries using a traditional (narrative) review of the literature. They tried to explain the relationship between the effectiveness of monetary policy and financial conditions. Unfortunately, these studies are not based on empirical assessments, but rather employ qualitative and narrative summary methods. Therefore, neither the average effect nor the heterogeneity of results are investigated. Nguyen (2019) has recently employed a meta-regression analysis on the effect of monetary policy in the context of EDCs, but this study focuses on the effects on output. Therefore, to our knowledge, there is no systematic analysis on the effect of monetary policy on price level in the context of EDCs. The contribution of this study is to provide a meta-regression analysis that synthesizes the different results.

Our meta-analysis builds on the earlier work by Rusnak et al. (2013) - they employ MRA to detect a publication selection bias in the reported effects of an interest rate increase on price levels. However, their meta-analysis limits its scope to the context of developed countries and includes exclusively published studies. According to De Long and Lang (1992), meta-analyses that include only published studies rely on a biased and limited interpretation of the relevant literature; several non-published studies have been carried out by highly qualified experts from reputable financial institutions (IMF, central banks etc).

To our knowledge, this is the first empirical attempt to systematically synthesize the effects of monetary policy on price levels in the context of emerging and developing countries (EDCs). Our paper contributes to the literature by measuring the corresponding genuine effects of monetary policy, as well as publication selection bias, for non-developed economies. We make use of 43 studies (27 published and 16 non-published) to synthesize their estimates on the effect of a tighter monetary policy on price levels in 32 emerging and developing countries. We follow the reporting guidelines on meta-regression analysis in economics by the Meta-Analysis of Economics

Research network (MAER-net)[2] and, in particular, the analytical approach by Stanley (2005, 2008) and Doucouliagos and Stanley (2009). We aim to test the null hypothesis of the negative genuine effect of an increase in interest rate on price level by providing answers to the following research questions:

- What is the summary (average) effect of an interest rate increase on price levels in EDCs based on the recent empirical literature?
- Is this average effect subject to a publication bias in primary studies?
- What is the genuine average effect after correcting for any publication bias and controlling for other potential explanatory factors (study/specification/country characteristics)?

The structure of the paper is as follows. Section 2 reviews the relevant literature and summarizes earlier findings. Section 3 describes our data collection and section 4 presents a descriptive analysis of all identified price effects. Section 5 employs several tests for the presence of publication bias and genuine effects. Section 6 consists of an extended meta-regression analysis and "best practice" analysis that estimates genuine effects after the publication bias and the misspecification is filtered out and other explanatory factors are accounted for. Section 7 concludes.

2. Literature review

Despite the increasing number of meta-analyses on economic issues in recent years, very few concentrate attention on monetary policies. Three of them focus on the output effects of monetary policy; out of these, two focus on advanced countries (De Grauwe & Costa Storti, 2004; Ridhwan et al., 2010) and one on emerging and developing countries (Nguyen et al., 2019). Similar to our analysis, Havranek and Rusnak (2013) and Rusnák et al. (2013) assess the effect of monetary policy on price level, albeit with an exclusive focus on developed economies. Klomp and De Haan (2010) examine the relationship between inflation and the independence of Central Banks, while Velickovski and Pugh (2011) discuss the flexibility of exchange rates. Papadamou et al. (2019) also employ a meta-analysis to discuss the effects of unconventional monetary policy on output and inflation in advanced countries. Serval studies employing meta-analysis regression method for emerging and developing countries, but on other topics (i.e. Iwasaki & Tokunaga, 2014 for the macroeconomic impacts of foreign direct investment; Iwasaki & Kočenda, 2017 for the role of ownership in privatized firms; Tokunaga & Iwasaki, 2017 for the determinants of foreign direct investment).

Data constraints are likely to account for the small number of meta-analyses on the effects of monetary policy. Quite often, neither the effect size nor the standard errors of estimates are directly available from primary studies. The effects are typically depicted by graphs of impulse response functions (IRFs). Therefore, in order to uncover the reported effect size, meta-analysts have to inspect IRF graphs and measure the reported outcomes. Extracting the standard errors from IRF graphs is only possible if confident interval bounds are depicted. Hence, translation of visual data (of the graphs of primary studies) into numerical data that can be used in meta-regression is a very time consuming and meticulous exercise. Several authors (e.g. De Grauwe & Costa Storti, 2004; Ridhwan et al., 2010; Havranek and Rusnak, 2012) could neither test for the publication bias nor investigate the "true" underlying effects due to the absence of reported standard errors.

The first meta-analysis on the price effects of monetary policy is found in De Grauwe and Costa Storti (2004). Their analysis is based on 43 published primary studies that cover 17 advanced economies (14 from the EU as well as the US, Japan and Australia). They analyze the impacts of a one-percentage point increase in the interest rate on output and price levels in both the short and long term. The authors conclude that there is a wide discrepancy in reported effects among countries (regarding sign, magnitude) and that their study is a preliminary attempt that can only explain parts of this variation. Regarding methodology, they find that, overall, VAR models report stronger long-term effects in comparison to SVAR estimates. They also claim that effects tend to be weaker in countries with a high inflation rate. However, a notable limitation of the study is the lack of inclusion of standard errors (of reported estimates). This naturally hinders investigating a publication bias and the "true" underlying effects (as a result of ignoring the precision of included study values, captured by their inverse standard errors, see Doucouliagos & Stanley, 2009).

A second meta-analysis by Havranek and Rusnak (2012) has a similar focus and explores the transmission lags of monetary policy effects on price levels. Their analysis is based on 67 published primary studies (on developed economies). According to their meta-analysis, the average transmission lag of monetary policy is about 29 months; they also conclude that a longer lag is found in countries characterized by higher levels of financial development, measured by the total outstanding credit to private sectors as percentage of GDP). Similar to the paper by De Grauwe and Costa Storti (2004), their study also does not collect the standard errors of reported effects, precluding hence the estimation of publication bias.

A third meta-analysis by Rusnák et al. (2013) (based on 70 primary studies) investigates the "puzzle" response of prices to a one-percentage point increase in interest rates (where, prices in many occasions, and contrary to intuition, increased following a tightening of monetary policy). In comparison to the previous two analyses, this study explicitly investigates the presence of a publication bias in the short, medium and long term. In addition, the study confirms that model misspecification could cause price "puzzle" effects in the short term, while in the long-term, price responses largely depend on macro-economic conditions. However, this meta-analysis also restricts its focus on developed countries. Furthermore, it is based exclusively on published articles, ignoring, hence, results from the so-called grey literature (working paper, mimeos etc).

[2] For more information on MAER-net, visit the network's website: https://www.hendrix.edu/maer-network/default.aspx?id=15088. The reporting guidelines for meta-regression analysis in economics can be found in Stanley et al. (2013).

Table 1
Search criteria for primary studies.

Criteria	Requirements
Country	Emerging and developing country
Model	Vector-autoregressive models
Policy shock	An increase (or decrease*) in interest rates
Proxy for economic activities	Price levels or index
Graph of impulse response functions	Reported interval confidence (to calculate standard errors). Accumulated responses are excluded.

Note: (*) In our sample, all primary studies report the responses of price level to an increase in the interest rate. No primary study reporting the response of price level to a decrease in the interest rate was found.

A third meta-analysis carried out by Papadamou et al. (2019) synthesizes 16 studies (15 published and 1 unpublished) that employed vector auto-regression specifications to measure the impacts of unconventional monetary policies (quantitative easing shocks) on output and inflation in advanced countries. The results show that FAVAR specifications predict stronger effects on output in all time horizons and on prices in the short term. In contrast, recursive identification is likely to report a weaker effect on prices. The studies on European unconventional monetary policies tend to come up with a weaker output effect. The meta-analysis by Papadamou et al. (2019) is, however, limited to advanced countries and largely relies on published studies; in addition, it does not investigate any publication bias and cannot reveal the genuine effects of unconventional monetary policy in advanced countries.

The fourth meta-analysis is carried out by Nguyen (2019) on the output effects of monetary policy in emerging and developing countries. This study is based on 45 studies conducted between 2001 and 2014 and synthesizes vector-autoregressive findings on the output effects of a tightening of monetary policy in 32 emerging and developing countries. The findings indicate a significant publication bias. However, after correcting for the publication bias, a genuine negative effect of a tightening of monetary policy on output remains. Primary studies that include commodity price variable(s) tend to report stronger negative effects. Output effects are shown to be more negative in an economy with a developed financial system, while monetary policy is less effective in an economy with high inflation volatility. Nguyen (2019) employs a meta-regression analysis and therefore assesses the publication bias and genuine effect, but only in relation to the output effects of monetary policy.

To the best of our knowledge, no other meta-analysis has been conducted to synthesize the price effects of monetary policy across emerging and developing economies. An earlier study by Mishra and Montiel (2013) makes use of qualitative techniques to survey empirical evidence on the effectiveness of monetary policy in developing countries using 39 primary studies. Their analysis provides limited support to sizeable monetary transmission effects, especially on financial development indicators. Furthermore, the study has not provided explanation on the heterogeneity of reported effects or examine the presence of a publication bias, given the qualitative approach of the analysis.

3. Data collection

We collected a series of suitable studies (published and unpublished) that report comparable effects. If a meta-analysis includes only published studies, the literature itself reflects an inherent bias (De Long & Lang, 1992, p.1452). We followed a four-step search strategy to identify as many as possible potential primary studies (see Nguyen, 2019 for the detail steps). We set search criteria (as presented in Table 1) to find primary studies that examine individual emerging or developing countries, and employ vector auto-regressive models to estimate responses of *price levels* to a shock from policy interest rates of monetary policy (primary studies that report responses of *inflation* are excluded).

Initially our search process ended in 2015 but upon the request of the referees of this journal we extended the research period to include studies up to 2019 and went through the 4 steps again. All in all, we identified a total of 43 primary studies (27 published and 26 non-published) conducted by 95 authors. 56% of the authors are academic researchers, 18% are employees of central banks, and 26% work for international financial institutions (International Monetary Fund, Bank for International Settlements). The primary studies cover 20 emerging and 12 developing countries[3].

In our meta-analysis, the unit of observation is at the level of reported impulse responses (rather than a single study). We collected a total of 133 impulse response function graphs based on our 43 primary studies. From the collected IRFs, we measure the effects of monetary policy on price levels. Based on the patterns and the horizontal axes of the IRFs, we measured and interpreted a total of 119 short-term, 119 medium-term, and 99 bottom effects. In the end, there were four less short-term effects which we could include in our meta-analysis. This is due to the short subsamples and very small magnitude of short-term effects (of four IFRs) found in the study by Fung (2002), which prevented us from calculating reliable effect estimates based on the corresponding graphs.

[3] The 43 identified primary studies are: Agha Agha, Ahmed, Mubarikm, & Shah, 2005, Aleem, 2010, Al-Mashat & Billmeier, 2008, Anzuini & Levy, 2007, H. Berument, 2007, M. H. Berument et al., 2014, Bhattacharya et al., 2011, Cheng, 2006, Chuku, 2009, Cocriş & Nucu, 2013a, Cocris & Nucu, 2013b, Cysne, 2004, Dabla-Norris & Floerkemeier, 2006, Disyatat & Vongsinsirikul, 2003, Elbourne & de Haan, 2006, Elbourne & de Haan, 2009, Fung, 2002, Gottschalk & Moore, 2001, He et al., 2013, Ibrahim, 2005, Jarociński, 2010, Kabundi & Ngwenya, 2011, Khundrakpam & Jain, 2012, Khundrakpam, 2012, Kubo, 2008, Lungu, 2007, Minella, 2003, Ngalawa & Viegi, 2011, H. T. Nguyen, 2014, C. P. Nguyen & Xuân Vinh, 2014, Oros & Romocea-Turcu, 2009, Parrado, 2001, Popescu, 2012, Samkharadze, 2008, Simic & Malesevic-Perovic, 2012, Starr, 2005, Sun et al., 2010, Tsangarides, 2010, Vonnák, 2005, Wróbel & Pawlowska, 2002, Nguyen, 2019, Pérez, 2015, Okur et al., 2019.

To investigate the effects of monetary policy on price levels over time, we examine the reported effects in the short and medium term (i.e. at 12 and 24 months after a tightening of monetary policy - denoted by y_{12} and y_{24} respectively). In addition, we examine the *bottom effect* (i.e. the maximum drop of price levels, denoted by y_{ne-max}) and the time it takes for this bottom effect to materialize (denoted by t_{ne-max}). The standard errors of reported effects are typically not directly available and need to be computed as mentioned in Nguyen (2019):

$$"\mathrm{Se}_{\mathrm{ij}} = \frac{|\mathrm{y}_{\mathrm{ij}} - \mathrm{y}_{\mathrm{ijb}}|}{"implicit\ t - value"\ \mathrm{of} \mathrm{y}_{\mathrm{ij}}} \tag{1}$$

where y_{ij} is the reported effect i of primary study j, Se_{ij} is the corresponding standard error, and y_{ijb} is the bound effect of y_{ij}". First, we measure the distances from any estimated point to its upper and lower confidence interval bounds (y_{ijb}). These distances reflect the size of the standard deviation (SD). After that, we divided the distance from the point estimate to each selected bound by the "implicit t-value" to acquire the size of the standard deviation.

We then used the *Plot Digitizer* software[4] (a Java program) to measure the magnitude of impulse responses. The software allows us to first enlarge and extract the IRF graphs, and then measure responses at 12 and 24 months, as well as the trough. The upper and lower confident intervals corresponding to each point estimate are also measured to provide the statistical significance and standard errors of the reported effects. All reported effects and standard errors were standardized so as to ensure that they correspond to the same interest rate increase (i.e. to a one-percentage point change).

4. Descriptive analysis of price effects

Table 2 summarizes the reported price effects (of a one-percentage-point interest rate increase) appearing in the 43 identified primary studies. There is considerable heterogeneity in the reported effects, although the majority of them appear to be negative and statistically insignificant. In the short term, 26% of all reported effects are statistically significant (22% negative and significant, 4% positive and significant). In the medium-term, 30% of all reported effects are statistically significant (23% negative and significant, 7% positive and significant). 45% of all bottom effects appear to be statistically significant.

Table 3 provides detailed information on average reported effects at the country level. Overall, the average reported effects are negative for most of the countries. This implies that a tighter monetary policy (measured as an increase in the policy interest rate) brings the price level down. However the magnitudes of the average reported effects differ from country to country. For example, in Brazil, after a one percentage point increase in interest rates, on average, price levels decline by 0.52% and 0.26% in the short and medium term correspondingly, but in Thailand, the price levels decline by 0.18% and 0.20% respectively. The differences in the magnitude of the effects could be due to country, study, and/or data characteristics, which will all be investigated in the later parts of this paper. The overall average price effects for the whole sample of emerging and developing countries are −0.10% and −0.12% in the short and medium term respectively. The average maximum negative effect corresponds to a price decline by 0.41% about 14 months since the initial interest rate increase[5]. The average reported effects in emerging countries are a bit stronger than in developing countries.

In comparison to developed countries (based on results by the earlier meta-analyses by Rusnak et al. (2013) and De Grauwe & Costa Storti, 2004), the average reported effects of monetary policy on price levels in emerging and developing countries (EDCs) tend to be much weaker. Fig. 2 illustrates this comparison. The magnitude of average reported effects in the short and medium term, as well as the maximum negative effects, tend to be twice as large (and persist longer) in the case of developed economies.

4.1. Weighted average effects on prices

The simple average effects calculated above treat all estimates equally regardless of their precision. This can be corrected by estimating the weighted-average effects of monetary policy on prices using the following formula:

$$\overline{y}_t = \frac{\sum w_{ij} \times y_{ij}}{\sum w_{ij}} \tag{2}$$

where $\overline{y_t}$ is the weighted average effect at time t, y_{ij} is the reported effect i of primary study j, and w_{ij} is the weight attached to effect y_{ij} (equal to the inverse of the standard error of the estimate (see Doucouliagos & Stanley, 2009). Alternatively, one can rely on sample sizes as weights in case standard errors are unavailable (see Stanley & Doucouliagos, 2012). Table 4 presents the measured weighted average effects according to these two weighting methods.

When using sample sizes as weights, results are largely in line with the average effects calculated in Table 3 (for the overall sample). This might be attributed to the rather small differences in sample size across primary studies (with of the mean number of observations per study being 105, and the standard deviation equal to 50). In contrast, weighted average effects based on inverse standard errors are considerable smaller in size. The substantial difference in weighted average effects using the two weighting schemes urges the need for

[4] *Plot digitizer* is a Java program to digitize scanned graphs of functional data. More information and to download visit http://plotdigitizer.sourceforge.net/

[5] Some additional summary statistics of reported effects are presented in Appendix 1.

Table 2
Composition of reported effects on prices.

	Statistical Significance (at 5%)		Statistical Insiginificance (at 5%)	
	Obs	%	Obs	%
Short-term effects	**33**	**26%**	**96**	**74%**
Negative	28	22%	55	43%
Positive	5	4%	41	32%
Bottom effects	**45**	**45%**	**54**	**55%**
Medium-term 35		**30%**	**84**	**71%**
Negative	27	23%	60	50%
Positive	8	7%	24	20%

Table 3
Price percentage changes after a tightening of monetary policy by country.

Country	Obs	Data ranges	**Short-term** (%)	**Medium-term** (%)	**Bottom effects** (%)	**Time lag of bottom effects** (months)
1. Brazil	7/7/3	1980–2013	−0.52 (0.40)	−0.26 (0.87)	−1.29 (1.85)	23
2. Bulgaria	2/2/1	2004–2012	0.04 (0.10)	−0.04 (0.00)	−0.09 (0.00)	4
3. China	11/7/11	1998–2013	0.13 (040)	−0.07 (0.08)	−0.25 (0.18)	16
4. Chile	3/3/2	1991–2015	−0.35 (0.10)	−0.62 (0.09)	−0.67 (0.02)	26
5. Croatia	1/1/1	2001–2011	0.04 (0.00)	0.02 (0.00)	−0.02 (0.00)	2
6. Columbia	1/1/0	1999–2013	−0.28(0.00)	−0.35 (0.00)		
7. Egypt	3/3/3	1996–2005	0.43 (0.26)	0.10 (0.39)	−0.18 (0.11)	2
8. Hungary	13/13/9	1992–2007	−0.12 (0.48)	−0.03 (0.57)	−0.43 (0.21)	17
9. India	6/6/2	1997–2012	−0.22 (0.07)	−0.16 (0.11)	−0.20 (0.16)	12
10. Indonesia	5/4/4	1986–2009	0.17 (0.49)	−0.03 (0.44)	−0.22 (0.17)	15
11. Malaysia	6/6/7	1985–2009	−0.55 (1.08)	−0.26 (0.34)	−0.75 (1.44)	7
12.Mexico	1/1/0	1999–2013	−0.10 (0.00)	−0.37 (0.00)		
13.Peru	1/1/1	2002–2013	−0.65 (0.00)	−0.89 (0.00)	−0.89 (0.00)	29
14. Philippines	2/3/3	1983–2001	0.14 (0.06)	0.21 (0.16)	−0.10 (0.11)	5
15. Poland	14/14/12	1992–2004	−0.11 (0.46)	−0.11 (0.53)	−0.33 (0.21)	18
16. Romania	6/5/6	1994–2012	−0.08 (0.11)	−0.16 (0.29)	−0.31 (0.48)	6
17. Russia	1/1/1	1995–2003	−0.15 (0.00)	−0.50 (0.00)	−0.14 (0.00)	18
18. South Africa	3/3/3	1985–2007	−0.04 (0.30)	−0.37 (0.22)	−0.70 (0.42)	30
19. Taiwan	2/1/3	1989–2001	−0.13 (0.02)	−0.34 (0.00)	−0.53 (0.69)	3
20. Thailand	9/9/9	1986–2006	−0.18 (0.66)	−0.20 (0.47)	−0.38 (0.62)	21
21. Turkey	4/2/0	1986–2017	0.54 (0.56)	0.27 (0.36)	–	–
Emerging Economies	**98/93/76**		**−0.11 (0.50)**	**−0.20 (0.55)**	**−0.42 (0.65)**	**14**
22. Armenia	2/2/2	2000–2005	−0.25 (0.18)	−0.03 (0.01)	−0.29 (0.01)	8
23. Belarus	1/1/1	1995–2003	−0.04 (0.00)	−0.003 (0.00)	−0.04 (0.00)	12
24. Georgia	2/0/2	2002–2007	0.01 (0.01)	–	−0.11 (0.00)	2
25. Kenya	4/4/4	1997–2005	−0.40 (0.11)	−0.23 (0.12)	−0.44 (0.09)	10
26. Malawi	4/4/2	1996–2006	−0.18 (0.44)	0.48 (1.17)	−1.65 (1.61)	14
27. Mauritius	4/4/3	1999–2009	−0.02 (0.06)	0.004 (0.04)	−0.07 (0.02)	3
28. Namibia	1/1/1	1990–2006	−0.65 (0.00)	−0.18 (0.00)	−0.65 (.000)	12
29. Nigeria	1/1/1	1986–2008	−0.67 (0.00)	−0.42 (0.00)	−0.79 (0.00)	9
30. Ukraine	1/1/1	1985–2003	0.08 (0.00)	−0.01 (0.00)	−0.01 (0.00)	21
31. Vietnam	10/7/5	1998–2017	0.11 (0.56)	0.53 (1.24)	−0.23 (0.29)	13
32. Zambia	1/1/1	1990–2006	−0.03 (0.00)	−0.10 (0.00)	−0.10 (0.00)	36
Developing Economies	**31/26/23**		**−0.10 (0.40)**	**0.15 (0.80)**	**−0.38 (0.57)**	**11**
Overall	**129/119/99**		**−0.10 (0.48)**	**−0.12 (0.63)**	**−0.41 (0.63)**	**14**

Note: Standard errors in parentheses; The Obs column presents the number of observations of short-term, medium-term, and bottom effects respectively. Source: author calculations.

undertaking precision effect tests (PET). This is the focus of the next section, which proposes methods to detect and test for publication bias (FAT) and genuine effects (PET and MST).

5. Publication bias and genuine effect test (MST and FAT – PET)

Several scholars have accentuated the importance of a publication selection bias in empirical research (De Long & Lang, 1992; Ioannidis et al., 2017; Stanley, 2005). Such a bias occurs when the publication of research papers depends on the nature of their results. Editors, referees and authors might, for instance, prefer larger and more significant effects that are in line with common theoretical predictions. The publication bias (i.e. the urge to publish in good journals) might drive researchers to work intensively until they can produce good/publishable results with low standard errors. However, it can be the case that a good study (with a lot of effort and a large

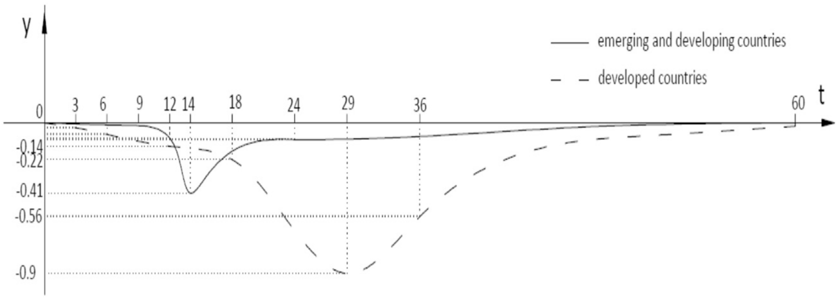

Fig. 2. Average reported effects of a one-percentage point interest rate increase on price levels in emerging and developing vs developed countries. Source: based on author calculations for the depicted effects for emerging and developing countries; depicted effects for developed countries are based on Rusnak et al. (2013) for effects at 3, 6, 9, 12, 18 and 36 months and De Grauwe and Costa Storti (2004) for effects at 60 months.

Table 4
Weighted average price effects.

Effects on prices	Sample size as weight		1/SE as weight	
	Mean	Std.Err	Mean	Std.Err
Short-term effect (%)	−0.10	0.04	−0.07	0.02
Medium-term effect (%)	−0.17	0.05	−0.08	0.02
Maximum negative (bottom) effect (%)	−0.40	0.06	−0.19	0.03
Months to reach maximum negative effect	14	1	14	1

dataset) produces estimates with lower standard errors that are less affected by a publication bias (i.e. good results end up in good journals). A third factor that influences both the precision of estimates and the publication selection could be the years of experience of a researcher (this may reflect both his/her ability to do good research as well as his/her reputation within journal circles). In general, the more precise estimates (smaller standard errors) are less affected by publication selection. Meanwhile, the results from the less precise estimates (larger standard errors) can vary a lot given the wide range of different specifications used. Therefore, in the presence of publication bias, there is correlation between effect size and its standard error. A funnel plot can detect publication bias visually (Egger et al., 1997; Stanley & Doucouliagos, 2012). In this type of graph, the horizontal axis depicts effect size against its precision (1/SE) on the vertical axis. The more precise the effect is, the closer it is to the true underlying effect. In the absence of publication bias, the funnel shape is symmetric. The more asymmetric the funnel plot, the more likely it is that publication bias occurs.

Fig. 3 depicts the funnel plots of price effects in the short (12 months) and medium term (24 months). The imbalance depicted by the larger proportion of negative effects (in both funnel plots) hints on the existence of a negative publication bias in the reported effects of primary studies.

However, a funnel plot is simply a visual aid to investigate the publication bias, and its interpretation could be subjective. For this reason, it is necessary to perform a statistical test to confirm the outcomes of our visual inspection. Following Stanley and Doucouliagos (2012), we employ the following meta-regression to test for publication bias:

$$y_{ij} = \beta_0 + \beta_1 Se_{ij} + \varepsilon_{ij} \tag{3}$$

where y_{ij} and Se_{ij} are the reported effect i of primary study j and its standard error, β_0 is the genuine average effect (after correcting for the publication bias), β_1 is the publication bias itself, and ε_{ij} is the error term. In the absence of publication bias, there is no correlation between effect and standard error (i.e. $\beta_1 = 0$). Effects are then distributed around the "true" underlying effect, or in other words, the

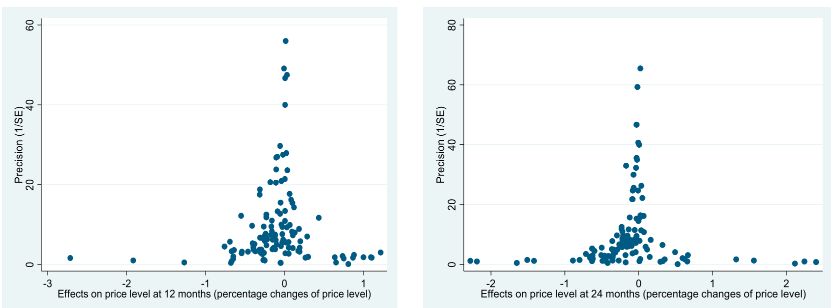

Fig. 3. Funnel plots of reported effects.

Table 5
Publication bias (FAT-PET) tests.

	Mixed-effect multilevel model			Cluster data analysis		
	Short-term	Medium-term	Bottom effect	Short-term	Medium-term	Bottom effect
Bias/FAT (β_1)	−0.440**	−0.592***	−1.741***	−0.618**	−0.706***	−1.773***
	(0.281)	(0.188)	(0.295)	(0.233)	(0.205)	(0.300)
Genuine effect/PET (β_0)	−0.014	−0.012	−0.010	−0.006	−0.008	−0.015
	(0.014)	(0.010)	(0.017)	(0.021)	(0.014)	(0.013)
N	129	119	99	129	119	99
Within-study correlation	0.62	0.58	0.85	–	–	–
# Studies	43	39	35	43	39	35

Note: Standard errors in parentheses; *p < 0.10, **p < 0.05, ***p < 0.01.

expected effect $E(y_{ij})$ is equal to β_0.

Due to the variance of the effect size and its error term, Equation (3) should not be estimated by ordinary least square (OLS) because of heteroscedasticity. Weighted least squares (WLS) are instead commonly used (Doucouliagos & Stanley, 2009). Dividing by Se_{ij}, Equation (3) becomes:

$$t_{ij} = \frac{y_{ij}}{Se_{ij}} = \beta_1 + \beta_0\left(\frac{1}{Se_{ij}}\right) + \phi_{ij} \quad (4)$$

where $\phi_{ij}(\phi_{ij}|SE_i \sim N(0, \delta^2))$. In Equation (4), the constant β_1 denotes the "true" underlying effect after correcting for any publication bias. The coefficient β_0 now denotes the publication bias. One can test for the presence of a publication bias and the "true" underlying effect by examining the statistical significance (t-statistic) of β_1 and β_0. These two tests are commonly referred to as the precision effect test (PET) and the funnel-asymmetry test (FAT) respectively (Stanley & Doucouliagos, 2012, p. 78). In addition:

- If β_1 is statistically different from zero, we fail to reject that there is no correlation between the effect size and the standard error. This suggests the presence of a publication bias in the effect size and $\widehat{\beta_0}$ provides the magnitude and the direction of this bias.

- If β_0 is statistically different from zero and the sign (direction) of $\widehat{\beta_1}$ ("true" underlying effect) is consistent with $\widehat{\beta_1}$ (publication bias), one can confirm the presence of a genuine effect on price levels (after a 1% interest rate increase).

Equation (4) is the WLS version of Equation (3) and can be directly estimated with OLS (Doucouliagos & Stanley, 2009). Stanley and Doucouliagos (2017) argue that this estimation method outperforms random or fixed effects models in the context of meta-analysis. In addition, and to account for within study independence, cluster data analysis or alternatively mixed-effect multilevel models should be applied (Rusnák et al., 2013; Stanley and Doucouliagos, 2012, 2017). Therefore, our preferred estimation method is WLS with cluster data analysis, and, in addition, we estimate a mixed-effect multilevel model as a further robustness check.

The estimations of Equation (4) and the corresponding FAT-PET tests are presented in Table 5 (both for cluster data analysis with standard errors clustered at the study level, as well as for the mixed effect multilevel models). The constants (β_1's) are consistently negative and statistically significant (at the 1% level) indicating that the reported effects suffer by a negative publication bias. In addition, the β_o's are statistically insignificant for all effects (i.e. short-term, medium-term and bottom effects) and for all estimation models. We can, hence, conclude that there is no evidence of a genuine effect when the publication bias is filtered out. However, apart from the publication selection bias, several other factors may explain the observed heterogeneity in reported effects and, should be accounted for when estimating the average genuine effect (Doucouliagos & Stanley, 2009). This will be the explicit focus of our next section.

The literature also proposes an alternative test for the existence of a genuine effect called the Meta-Significance-Test (MST) (see Card & Krueger, 1995; and Stanley, 2001, 2005):

$$E\left(\ln|t_{ij}|\right) = \alpha_0 + \alpha_1 \ln df_{ij} \quad (5)$$

where t_{ij} is the t-value of reported effect *i* of study *j*, and df_{ij} is the degrees of freedom of the y_{ij} estimation[6]. Given that precision typically increases in line with sample size, one expects a positive and statistically significant correlation ($\alpha_1 > 0$) between the natural logarithm of the t-value of effect and its corresponding degrees of freedom. Such a positive and statistically significant correlation will also indicate the presence of a genuine effect. Estimating Equation (5) with clustered (at the study level) robust standard errors yields the results presented in Table 6.

As reported in Table 6, we fail to reject the null hypothesis ($\alpha_1 \leq 0$) for the bottom effects. This suggests that there is no evidence of a genuine effect in bottom effect (in line with our earlier findings of Table 5). However, the coefficientα_1 is statistically significant at the 10% level for the short-term effect and at the 5% level for the medium-term effect, which suggests the existence of a genuine effect in the short and medium-term. Nevertheless, MST generally suffers from limitations, especially regarding a higher probability of a type I error (i.e., incorrectly rejecting the true null hypothesis) – it is for this reason that the extended FAT-PET tests (discussed in the following

[6] Degrees of freedom (df) are calculated for multi-regressions according to: $df = n - k - 1$where n is. the number of observations and k is the number of variables included in the VAR models (for estimations from FAVAR, we use the number of factors).

Table 6
Meta-significance-test.

	Short-term	Medium-term	Bottom effect
Constant (α_0)	−1.550*	−2.036*	−0.544
	(0.813)	(0.944)	(0.619)
Lndf (α_1)	0.309*	0.424**	0.197
	(0.187)	(0.198)	(0.135)
N	129	119	99
R2/Adj R2	0.022/0.014	0.054/0.046	0.016/0.005
# Studies	43	39	35

Note: Standard errors in parentheses; *p < 0.10, **p < 0.05, ***p < 0.01.

section) are generally considered more reliable when testing for the presence of genuine effects (Stanley, 2005).

(vi) Expanded meta-regression analysis

5.1. Explanatory variables

Here, we follow Stanley & Doucouliagos, 2012; Rusnák et al., 2013 and Rusnák et al., 2013 in setting up an expanded meta-regression model that also incorporates a vector of *k* additional explanatory variables (Z_k). We, hence, include in Equation (4) all these additional factors that are likely to explain the observed heterogeneity in reported effects and, should be accounted for when estimating average genuine effects. Equation (4) then becomes:

$$t_{ij} = \frac{y_{ij}}{SE_{ij}} = \beta_1 + \beta_0 \frac{1}{SE_{ij}} + \cdot \sum_{k=1}^{K} \partial_k \frac{Z_{ijk}}{SE_{ij}} + \varepsilon_{ij} \quad (6)$$

where Z_k denotes the *k* meta-explanatory variables alleged to affect reported price effects, ∂_k is a vector of meta-coefficients reflecting

Table 7
Variables used in meta-regressions.

Variable	Obs	Description	Mean	Std. Dev.
Dependent variables				
Short-term effect at 12 months (inverse of standard errors in parentheses)	129	Price response 12 months after a one percentage point increase in interest rates	−0.10 (9.14)	0.48 (10.31)
Medium-term effect at 24 months (inverse of standard errors in parentheses)	119	Price response 24 months after a one percentage point increase in interest rates	−0.12 (10.03)	0.63 (12.32)
Maximum negative effect (bottom effect) (inverse of standard errors in parentheses)	99	Maximum price response after a one percentage point increase in interest rates	−0.41 (10.39)	0.63 (10.07)
Explanatory variables				
Group 1: Study characteristics				
Number of observations (Nobs)	133	Total number of observations used in primary study	111	58
Data frequency (Freq)	133	Dummy variable = 1 if monthly data is used	0.79	0.41
Publication status (ISI-journal)	133	Dummy variable = 1 if study published in an ISI-listed journal	0.47	0.50
Publication year (Year)	133	Year of publication since 2000	8	4
Affiliation (Affil)	133	Dummy variable = 1 if author(s) work for a central bank	0.22	0.41
Group 2: Specification characteristics				
SVAR model (SVAR)	133	Dummy variable = 1 if primary study uses SVAR model	0.31	0.46
VAR model (VAR)	133	Dummy variable = 1 if primary study uses VAR model with Cholesky decomposition	0.57	0.50
Inclusion of commodity price variable (COM)	133	Dummy variable = 1 if primary study includes at least one commodity price variable	0.54	0.50
Inclusion of exchange rate variable (Exchange)	133	Dummy variable = 1 if primary study includes an exchange rate variable	0.82	0.49
Inclusion of variable control for external shocks (External)	133	Dummy variable = 1 if primary study includes at least one variable controls for external shocks.	0.60	0.49
Type of policy interest rate (INT)	133	Dummy variable = 1 if primary study looks at responses to short-term interest rates	0.86	0.35
Group 3: Structural or country characteristics				
CPI volatility (CPI volatility)	133	Standard deviation of the CPI index over the period of each study	14.12	8.5
Exchange rate regime (Float)	133	Dummy variable = 1 in case of a floating exchange rate regime	0.09	0.29
Single or multiple exchange rate regime (Single exch)	133	Dummy variable = 1 if same exchange rate regime present during entire period	0.50	0.50
Financial development (Fd)	133	Financial market development index (Svirydzenka, 2016)	0.33	0.12
Financial openness (Fo)	133	The Chinn-Ito (or KAOPEN) financial openness index	−0.10	1.17
Independence of the central bank (Ind)	133	Index of central bank independence (Arnone et al., 2009)	0.62	0.19

the effect of each meta-explanatory variable on the reported estimates, and ε_{ij} is the error term. β_0 captures the genuine effect after correcting for the publication bias and is conditional on the effects of all other meta-explanatory variables. We define the Z_k vector based on the academic literature on the relationship between monetary policy, price levels and other mediating factors. Table 7 provides summary statistics and descriptions for all variables appearing in our meta-analysis.

Explanatory variables are categorized into three groups. The first group controls for primary study characteristics; here, we follow Klomp and De Haan (2010) and Rusnak et al. (2013) and include the following five variables:

Number of observations (Nobs): reported effects might be sensitive to sample size. This explanatory variable detects the possible correlation between effect size and the number of observations used in primary studies. *Data frequency (Freq)*: this variable tests whether negative price responses might be more common in primary studies using higher frequency data. *Publication status (ISI-journal):* this variable test whether there is a tendency to report more negative price effects in primary studies published in ISI (Institute for Scientific Information) journals. *Publication year (Year):* this variable examines whether there is a systematic relationship between the year of publication and reported price effects. *Affiliation (Affil)*: we look at whether authors working for central banks tend to report stronger negative price effects (in our sample, 18% of the primary studies were conducted by authors affiliated to a central bank).

The second group (of six explanatory variables) controls for the specification characteristics of our sampled estimates. According to Walsh (2017), specification characteristics (such as the type of estimated model, included variables and adopted lag length) might influence estimation outcomes. This group of regressors consists of:

VAR and SVAR models: the VAR framework has evolved overtime from the reduced-form VAR to structural VAR (SVAR), Bayesian VAR (BVAR), and factor augmented VAR - (FAVAR) models. Kim and Roubini (2000) suggest that price "puzzle" effects (where prices increase in the aftermath of a tightened monetary policy) are more common in SVAR models. In our sample, 58%, 33%, 5%, 3% and 1% of all primary estimations employ VAR, SVAR, FAVAR, VECM and BVAR models respectively. We include dummy variables to control for the use of the more frequently adopted VAR and SVAR models. *Commodity price (COM):* Sims (1992) and Castelnuovo and Surico (2010) argue that price "puzzles" might be the result of omitted commodity prices that capture inflation expectations. On the other hand, Hanson (2004) and Giordani (2004) find no supporting empirical evidence[7]. To test for the validity of this hypothesis, we include a dummy variable to control for the presence of commodity price variables. *Exchange rate (Exch)*: Gali and Monacelli (2005) argue that there could be a trade-off between nominal exchange-rate stability and price stability in a small open economy. Nasir et al. (2020) use data from the Czech Republic to provide support of an "exchange-rate pass through" mechanism, through which the exchange rate can generate inflationary expectations and, hence, raise the prices of goods and services. In addition, Sims (1992) points out that the inclusion of exchange-rate variables in VAR models makes price "puzzle" responses to disappear. Therefore, we include a dummy variable to control for the inclusion of exchange-rate variables. *External variables (External):* Boivin and Giannoni (2008) argue that globalization could dampen the effect of monetary policy on economy. Globalization means that a national economy integrates into the world economy (Anwar & Nguyen, 2018). *Short-term interest rate (INT)*: We include a dummy variable to check whether there is a differentiated price response to changes in short-term vs. long-term interest rates.

The third group (of six explanatory variables) accounts for country-specific heterogeneity in financial dimensions, exchange rate regimes and price volatility (data sources are provided in Appendix 2). The motivation behind the inclusion of country-specific controls lies in the substantial variation in average reported effects across countries as observed in Table 3. The data on country-specific is taken from several macroeconomic data sources. This group of country-specific controls consists of:

Price volatility (CPI volatility): several scholars argue that price volatility may generally affect the effectiveness of monetary policy (Ascari & Ropele, 2007; Mishkin, 2009; Reifschneider & Williams, 2000). We use the standard deviation of the CPI index over the entire period of each study to capture price volatility. *Floating exchange-rate regime (Float)*: exchange-rate volatility can influence the price responses of monetary policy (Bleaney & Fielding, 2002; Taylor, 2001) and Reinhart and Rogoff (2004) find that floating regimes are typically associated with lower inflation rates. We include a dummy variable in order to control for the different types of exchange-rate regimes (and their mediating role on the price effects of monetary policy). This takes a value of 1 when a country adopts a floating exchange rate regime (instead of a fixed exchange rate); Reinhart and Rogoff (2004) provide annual data on exchange-rate regime classifications, and we rely on their updated series between 1946 and 2016[8]. *Single exchange-rate regime (Single exch):* this dummy variable take a value of 1 if the same exchange regime has been present during the entire period of analysis. *Financial development (Fd)*: financial development plays a significant mediating role when implementing monetary policies - policy signals are largely transmitted through financial markets before affecting the real economy. In that vein, Friedman and Woodford (2010) find that such transmission mechanisms seem to be more effective in countries with advanced financial development (Friedman & Woodford, 2010); in other words, lower levels of financial development in developing countries could result in less effective monetary policy (Gül & Taştan, 2020; Mishra et al., 2010; Mishra & Montiel, 2013). Here, we use the index of financial development by Svirydzenka (2016) that examines the development of both financial markets and financial institutions in terms of their depth, efficiency and access. *Financial openness (Fo)* (Romer, 1993; Terra, 1998): suggest that, in a more open economy, the effect of monetary policy is less effective and, hence, inflation can be higher, other things equal. Here, we use an index of financial openness, originally developed by Chinn and Ito (2006); this index measures capital account openness by country based on IMF reports on restrictions on cross-border financial transactions, exchange rate arrangements and exchange regulations. *Independence of central bank (Ind)*: there is wide consensus regarding a positive correlation

[7] On the other hand, Giordani (2004) has shown that the output gap could explain the "price puzzle". Unfortunately the primary studies in our sample do not include output gap variables (this may be due to the unavailability of data for emerging and developing countries).

[8] The dummy variable takes a value of 0 when exchange rate regimes are classified as peg, band, crawling, or managed float (coarse grid categories 1, 2, or 3).

between the independence of the central bank and price stability (Alesina & Gatti, 1995; De Haan & Kooi, 2000; Debelle & Fischer, 1994). We use an index of central bank (both political and economic) independence by Arnone et al. (2009) and test how this might influence the reported effects of monetary policy on price levels.

5.2. General-to-specific approach

We now proceed to estimate our expanded meta-regression model (Equation (6)). One needs to keep in mind, that, on the one hand, an omitted variable bias might arise in case of excluding important explanatory variables. On the other hand, due to the large number of regressors, multicollinearity issues may become present if one includes all variables simultaneously in the estimated model(s). Therefore, we follow the general-to-specific approach proposed by several meta-analysts (Doucouliagos & Stanley, 2009; Klomp & De Haan, 2010; Rusnák et al., 2013). Some other approaches could be the Baysian model average (see Havranek et al., 2015), or the use of winsorizing in meta-analysis and frequentist model averaging (see Havranek et al., 2017). We opt for the general-to-specific approach[9] because this is the standard practice in the field and prescribed by the MAER-net protocol. This approach begins with all potential explanatory variables; progressively, the least statistically significant variables are removed, one by one, until the model includes only statistically significant (at least at the 10% level) regressors (Cocris & Nucu, 2013b). Similar to Doucouliagos and Stanley (2009), we first apply the general-to-specific method in the case of cluster data analysis (with standard errors clustered at the study level); once we come up with a specific model that includes only statistically significant regressors, we re-estimate the specification with the use of a mixed effect multilevel model. Table 8 presents the estimations of these two models.

The FAT-PET test results of the extended meta-regression model in Table 8 are interpreted as follows[10]:

The constants ($\beta_1{}'s$) for all price effects (short-term, medium-term and bottom effects) are all negative and statistically significant (and similar in magnitude to the ones in Table 5). This is the case both for the cluster data analysis, as well as for the mixed effect multilevel model. These results revalidate our earlier findings on a significant and negative publication bias (Table 5).

Let us focus now on the $\partial_k{}'s$ (the coefficients of the other meta-explanatory variables) – here, we refer to the subset of coefficients that remain (jointly) significant according to the general-to-specific method. The coefficient of *ISI-journal* is consistently negative and statistically significant across all specifications – in other words, authors who publish their studies in academic (ISI-listed) journals tend to report stronger negative effects. We also find that the inclusion of commodity prices (*COM*) correlates positively with the size of reported price effects. This finding corroborates Hanson (2004) and Giordani (2004), who also find that price puzzles are not the result of omitted commodity prices. On the other hand (and contrary to earlier results by Boivin & Giannoni, 2008), we find that studies that include external indices (*External*) tend to report stronger negative price effects. Last (and in line with Romer, 1993 and Terra, 1998) we find that a tighter monetary policy tends to be less effective (in lowering price levels) in economies characterized by a floating exchange-rate regime (*Float*) and fewer regulatory restrictions on financial/exchange transactions (*Fo*).

Turning to the coefficients of the inverse standard errors ($\beta_o{}'s$) which provide the genuine underlying effect after correcting for publication bias and accounting for the role of other explanatory variables (Z_k). We still find no evidence of a negative genuine effect of monetary policy on prices. However, coefficients (β_0) is positive and statistical (especially for the cluster data analysis) indicating the existence of genuine effects at short-term, medium-term, and of the bottom effect. The sign is positive which can be due to the simultaneous presence of a relatively small negative "true" effect and a strong (larger in magnitude) negative publication bias (Doucouliagos & Stanley, 2009). More importantly, the coefficient (β_0) does not purely reflect the size and direction of the genuine effects because it is conditional on the coefficient of the publication bias and the coefficient of all other significant explanatory variables. To figure out the magnitude of the "true" underlying effects, one needs to define the preferred values of additional variables (Z_k variables in Equation (6)) by "best practice" method (Doucouliagos & Stanley, 2009).

5.3. "Best practice" analysis

We follow several previous studies (Doucouliagos & Stanley, 2009; Havranek & Irsova, 2011; Rusnák et al., 2013; Havránek, 2015) to apply the "best practice" method to discover the sign and magnitude of the genuine effects of monetary policy on prices. The "best practice" method allows to estimate the "true" underlying effects from the "ideal" parameters of all other explanatory variables (Z_k variables) to eliminate the misspecifications (Doucouliagos & Stanley, 2009). We define the "best-practice" based on the previous empirical outcomes and the implication about the "best-practice" in the literature. In terms of model specification, we prefer the inclusion of commodity and exchange rate, and foreign variables (Hanson, 2004; Sims, 1992). We opt for monthly data frequency. In terms of study characteristics, we select the peer-reviewed studies that are published in high quality journals (A & B journal or IMF papers). We prefer data covering single exchange regime rather than multi exchange rate regimes (Taylor, 2001). Other country characteristics variables are set to their sample means. Table 9 reports the estimated average genuine effects implied by "best practice" and their narrow 95% confidence intervals.

Table 9 reports the genuine average effects according to several identifications of "best practice". In all identifications, the estimated genuine effect suggested by the "best practice" are negative and significant at short-term, at medium-term and at the bottom-effects. The 95% confidence intervals of these effects also point to a negative effect. For example (when using monthly data and including

[9] See Campos et al. (2005) for a review of the literature on general-to-specific modelling in economics.

[10] The tests for multicollinearity (variance inflation factor - VIF tests) provide no evidence of a multicollinearity bias in the reported coefficients (the variance inflation factors are smaller than 10).

Table 8
Expanded meta-regression models.

	Cluster data analysis			Mixed effect multilevel model		
	Short-term	Medium-term	Bottom effect	Short-term	Medium-term	Bottom effect
Bias/FAT (β_1)	−0.568**	−0.626***	−1.761***	−0.655**	−0.597***	−2.011***
	(0.220)	(0.198)	(0.319)	(0.277)	(0.225)	(0.449)
Genuine effect/PET (β_0)	0.082**	0.059***	0.102**	0.059	0.046**	0.049
	(0.034)	(0.022)	(0.037)	(0.041)	(0.023)	(0.068)
ISI-journal	−0.164***	−0.110***	−0.110***	−0.131***	−0.094***	−0.054*
	(0.032)	(0.019)	(0.042)	(0.032)	(0.025)	(0.037)
VAR			0.065*			0.021
			(0.033)			(0.027)
COM	0.102***	0.047***	0.099*	0.085**	0.030	0.070
	(0.031)	(0.018)	(0.049)	(0.033)	(0.028)	(0.051)
External	−0.112***	−0.065***	−0.143**	−0.081**	−0.042*	−0.088*
	(0.028)	(0.017)	(0.049)	(0.031)	(0.025)	(0.048)
Float	0.188***	0.093***		0.211**	0.074*	
	(0.055)	(0.029)		(0.072)	(0.043)	
Fo	0.015*	0.022**		0.009	0.009	
Inte	(0.010)	(0.009)	−0.072** (0.039)	(0.012)	(0.011)	−0.042 (0.050)
N	129	119	99			
R2 adjusted/within study	0.35	0.26	0.19	0.49	0.51	0.84
Variance factor (mean)	3.38	3.18	5.32			
No. of studies	43	39	35	43	39	35

Note: Standard errors in parentheses; *p < 0.10, **p < 0.05, ***p < 0.01.

Table 9
Estimated price responses implied by the "best practice".

	Linear combination		
	Short-term	Bottom effect	Medium-term
1. Average country characteristics with commodity			
Estimated effect	−0.039** (0.017)	−0.040**(0.016)	−0.034***(0.16)
95% confidence interval	[-0.074; −0.004]	[-0.073; −0.008]	[-0.067; −0.001]
2. Average country characteristics with commodity, exchange rate, foreign variable			
Estimated effect	−0.09** (0.03)	−0.10**(0.03)	−0.08***(0.02)
95% confidence interval	[-0.15; −0.03]	[-0.16; −0.03]	[-0.12; −0.04]
3. Average country characteristics with commodity, exchange rate, foreign variable and SVAR specification			
Estimated effect	−0.11* (0.05)	−0.13**(0.05)	−0.08 (0.05)
95% confidence interval	[-0.23; −0.10]	[-0.23; −0.03]	[-0.19; −0.02]
4. Commodity, exchange rate, foreign variable with the best country characteristics			
Estimated effect	−0.13** (0.05)	−0.19**(0.08)	−0.14***(0.03)
95% confidence interval	[-0.24; −0.02]	[-0.36; −0.2]	[-0.21; −0.07]

Note: The values represent the percentage change of output to a one-percentage point increase in the policy interest rate. Standard errors in parentheses, *p < 0.10, **p < 0.05, ***p < 0.01.

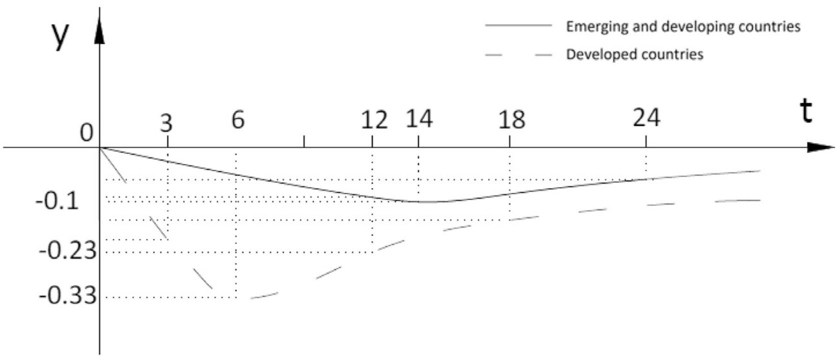

Fig. 4. Genuine price effects of monetary policy in EDCs vs developed countries.
Source: author own's depiction. Genuine price effects of emerging and developing countries based on the findings as reported in Table 9. Genuine price effects of developed countries (used the same methodology) based on the study by Rusnak et al. (2013).

commodity prices, exchange rate data and foreign variables with average country characteristics) the genuine effects are on average −0.09%, −0.10%, and −0.08% at short-term, at bottom effect, and at medium-term, respectively. The outcomes indicate that after filtering out the publication bias, the misspecification, and conditional on explanatory factor, the price effects of an increase in interest rate (suggested by the "best practice") are negative and statistically significant. The finding suggests that a tighter monetary policy (measured as a one-percentage point increase in the policy interest rate) is able to confront inflation in emerging and developing countries.

In comparison to advanced countries, the genuine price effects of an increase in interest rate (measured as a one-percentage point increase in interest rate) in emerging and developing countries weaker. Fig. 4 compares the price effects of monetary policy between advanced countries and EDCs.

The magnitude of both the simple average price effects and the genuine price effects after filtering out the publication bias and the misspecifications in EDCs countries are smaller in advanced countries (as depicted in Figs. 2 and 4, respectively). The findings of our study together with the findings of Rusnak et al. (2013) provide econometrical evidence that according to the existing studies (that use VAR models to measure the response of price level to an increase in the interest rate), the monetary policy in EDCs is less effective than in advanced countries.

5.4. Robustness check

To check for the robustness of our empirical outcomes, we run additional MRAs that cluster the standard errors at the countries level to test for the publication bias and genuine. The outcomes are presented in Appendix 3 (FAT-PET tests) and Appendix 4 (expanded MRAs), which are very similar to the outcomes of the models that cluster the standard errors at the study level (as discussed in the main analysis).

In addition, the most recent advanced meta-analysis method (WAAP – Weighted Average of only Adequated Powered) includes only adequate powered estimates in the MRA models (see Stanley, Doucouliagos and Ioannidis, 2017). Adequate powered estimates (defined according to Cohen's standard) have the probability of making type II error (β) smaller than 20% or the statistical power (1- β) greater than 0.8. Unfortunately, the VAR primary studies often report the graphs of impulse response functions instead of the tables of regression outcomes. Thus, no information about r2 is available to calculate the β and further (1- β) of VAR estimates. Nevertheless, the primary studies report the significant level of the estimates (α or the confidence intervals of the impulse response function). In our sample, α could be 0.01, 0.05, 0,1, or 0.32. However, the significant level of the estimate (α) also serves as another indicator of statistical power. The significant level of the estimate (or the precision of the estimation) reflects the probability of making type I error). Using the Cohen's standard "four-to-one" ratio between β and α, one can use the α level (equal to or smaller than 0.05) as the proxy for adequate power ((1- β) equal to or greater than 0.8). We re-estimate our MRA models for the subsample restricts to only adequate power estimates (α level could be 0.05 or 0.01). Appendices 5, 6, and 7 report the outcomes of publication bias tests (FAT-PET tests), expanded meta-regression models, and the genuine effect suggested by the "best practice" using WAAP method, which support our earlier findings.

6. conclusions

To our knowledge, this is the first meta-regression analysis that systematically reviews the price effects of a tightening of monetary policy in the context of emerging and developing countries (based on reported estimates across 40 primary studies). We synthesize all reported price effects and measure the corresponding genuine effects of monetary policy, as well as publication selection bias, for non-developed economies.

Our literature review of existing price effects (appearing in studies with a focus on developing and emerging economies) points to a substantial heterogeneity across reported estimates (see Section 4). The majority of them appear to be negative and statistically insignificant (and of relatively small size; namely, −0.07%, −0.08% and −0.19% for the weighted average of short-term, medium-term and bottom effects).

Our meta-analysis provides evidence of a strong negative publication bias for all types of price effects (short-term, medium-term and maximum negative effects; see Sections 5 and 6). In addition, we find that several other control factors can help explain the observed heterogeneity in reported price effects. On average, studies published in academic (ISI-listed) journals tend to report stronger negative effects (the same holds also for studies that include external indices in their empirical specifications). A tighter monetary policy tends to be less effective (in lowering price levels) in economies characterized by a floating exchange-rate regime (and fewer regulatory restrictions on financial/exchange transactions).

Furthermore, we confirm the null hypothesis of the genuine effect of an increase in interest rate on price level. Employing the "best practice" approach, we found that after correcting for the publication bias, misspecification, and controlling for additional meta-regressors, empirical evidence of a negative genuine effect of a tighter monetary policy on prices remains. The maximum drop of price level is about 0.10% and occurs at about 14 months after an increase in the policy interest rates. However, the genuine price effect in EDCs is weaker than in advanced countries. These analytic findings support the conventional view that an increase in interest rate can bring down inflation and the view that monetary policy is less effective in less developed countries than it is in advanced countries.

CRediT author contribution statement

Thi Mai Lan Nguyen: Conceptualization, Data curation, Formal analysis, Writing. **Elissaios Papyrakis:** Supervision, Writing - review & editing. **Peter A.G. van Bergeijk:** Methodology, Supervision, Writing - review & editing.

Appendix 1. Summary statistics of price level responses to a one percentage point increase in interest rates

Variable	Obs.	Mean	Median	Std.dev	Min.	Max.
Short-term effect (%)	**129**	**−0.10**	**−0.09**	**0.48**	**−2.71**	**1.22**
Short-term upper confidence interval (CI) effect (%)	129	0.38	0.15	0.79	−1.43	5
Short-term lower CI effect (%)	129	−0.69	−0.39	1.05	−5.69	0.56
Medium-term effect (%)	**119**	**−0.12**	**−0.12**	**0.63**	**−2.28**	**2.40**
Medium-term upper CI effect (%)	119	0.44	0.12	1.01	−0.60	6.90
Medium-term lower CI effect (%)	119	−0.79	−0.38	1.00	−6.71	0.26
Maximum negative effect (bottom effect) (%)	**99**	**−0.41**	**−0.24**	**0.62**	**−4**	**−0.01**
Maximum negative upper CI effect (%)	99	0.01	0.04	0.44	−3.43	0.92
Maximum negative lower CI effect (%)	99	−0.90	−0.53	1.26	−6.83	−0.05
Time lags of maximum negative effect (months)	**99**	**13**	**10**	**11**	**1**	**51**

Appendix 2. Country-specific controls (Group 3) and their data sources

Variable	Sources of data
CPI volatility *(CPI volatility)*	Standard deviation of CPI index by country and period of investigation. Data on CPI index (2010 = 100) provided by the World Development Indicators: https://data.worldbank.org/indicator/FP.CPI.TOTL (assessed on Oct 15th, 2017)
Financial development	Financial market development index by (Svirydzenka, 2016). Average values for the entire period of analysis. Data retrieved from https://www.imf.org/en/Publications/WP/Issues/2016/12/31/Introducing-a-New-Broad-based-Index-of-Financial-Development-43621 (assessed on Oct 15th, 2017)
Financial openness	The Chinn-Ito (or KAOPEN) financial openness index (averaged over the entire period of analysis). Data retrieved from: http://web.pdx.edu/~ito/Chinn-Ito_website.htm (assessed on Oct 15th, 2017)
Float exchange rate regime *(Float)*	Annual data on exchange-rate regime classifications between 1946 and 2016 (updated series of Reinhart and Rogoff (2004). Dataset retrieved from http://www.carmenreinhart.com/data/browse-by-topic/topics/11 (assessed on Oct 15th, 2017)
Single exchange rate regime	
Central bank independence	Central bank independence index by (Arnone et al., 2009)

Appendix 3. : Publication bias (FAT-PET) tests of reported price effects (cluster the standard errors at the country level)

	Mixed-effect multilevel model			Cluster data analysis		
	Short-term	Medium-term	Bottom effect	Short-term	Medium-term	Bottom effect
Bias/FAT (β_1)	−0.440	−0.592***	−1.741***	−0.618***	−0.706***	−1.773***
	(0.281)	(0.188)	(0.295)	(0.215)	(0.189)	(0.027)
Genuine effect/PET (β_0)	−0.014	−0.012	−0.010	0.006	−0.008	−0.015
	(0.014)	(0.010)	(0.017)	(0.002)	(0.014)	(0.014)
N	119	109	93	129	119	99
Within-study correlation	0.62	0.58	0.85	–	–	–
Countries	32	31	29	32	31	29

Note: Standard errors in parentheses; *p < 0.10, **p < 0.05, ***p < 0.01, cluster the standard errors at the country level.

Appendix 4: Expanded meta-regression models of price effects (cluster the standard errors at the country level)

	Cluster data analysis		
	Short-term	Medium-term	Bottom effect
Bias/FAT (β_1)	−0.556**	−0.626***	−1.760***
	(0.227)	(0.190)	(0.308)
Genuine effect/PET (β_0)	0.082**	0.059***	0.160**
	(0.034)	(0.021)	(0.013)
ISI-journal	−0.164***	−0.110***	−0.102**
	(0.033)	(0.020)	(0.037)
VAR			0.065*
			(0.037)
COM	0.102***	0.047***	0.090*
	(0.031)	(0.021)	(0.046)
External	−0.112***	−0.065***	−0.142**
	(0.031)	(0.019)	(0.050)
Float	0.188***	0.093***	

(*continued on next page*)

(*continued*)

	Cluster data analysis		
	Short-term	Medium-term	Bottom effect
	(0.052)	(0.033)	
Fo	0.015	0.022**	
	(0.011)	(0.010)	−0.071* (0.037)
Inte			
N	129	119	99
Adj R2	0.32	0.22	0.19
No. of countries	32	31	29

Note: Standard errors in parentheses; *p < 0.10, **p < 0.05, ***p < 0.01.

Appendix 5. WAAP FAT- PET tests of price effects

	Weighted Least Square		
	Short-term	Medium-term	Bottom effect
Precision (1/SE) (β_0)	−0.009	−0.004	−0.004
	(0.024)	(0.013)	(0.014)
Constant (bias) (β_1)	−0.461	−0.551*	−1.898***
	(0.321)	(0.299)	(0.454)
N	90	82	65
No. of studies	31	28	24

Standard errors in parentheses, *p < 0.10, **p < 0.05, ***p < 0.01; Cluster the standard errors at the study level.

Appendix 6. WAAP expanded meta-regression models of price effects

	Short-term	Medium-term	Bottom effect
Bias/FAT ($\boldsymbol{\beta}_1$)	−0.273	−0.492	−1.990***
	(0.297)	(0.297)	(0.048)
Genuine effect/PET ($\boldsymbol{\beta}_0$)	0.057	0.049**	0.074
	(0.044)	(0.023)	(0.047)
ISI-journal	−0.160***	−0.104***	−0.154**
	(0.031)	(0.021)	(0.054)
VAR			0.104*
			(0.053)
COM	0.112***	0.053***	0.116**
	(0.035)	(0.021)	(0.052)
External	−0.131***	−0.062***	−0.182***
	(0.033)	(0.020)	(0.062)
Float	0.179***	0.093***	
	(0.064)	(0.033)	
Fo	0.011	0.023***	
Inte	(0.011)	(0.010)	−0.014 (0.028)
N	90	82	65
Adj R2	0.39	0.17	0.18
No. of studies	31	28	24

Note: Estimated by OLS, standard errors in parentheses; *p < 0.10, **p < 0.05, ***p < 0.01; Cluster the standard errors at the study level.

Appendix 7. WAAP estimated responses implied by the "best practice"

	Linear combination		
	Short-term	Bottom effect	Medium-term
Estimated effect	−0.121** (0.038)	−0.204**(0.079)	−0.090*(0.048)
95% confident interval	[-0.201; −0.044]	[-0.369; −0.040]	[-0.189; 0.009]

Note: The values represent the percentage change of output to a one-percentage point increase in the policy interest rate. Standard errors in parentheses, *p < 0.10, **p < 0.05, ***p < 0.01.

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