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Evidence on time-varying inflation synchronization

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ABSTRACT

Most studies on global inflation are conducted on homogeneous, advanced-economy, low-frequency samples and present evidence favouring the global inflation paradigm. I challenge this consensus view by quantifying price co-movements across a large, heterogeneous sample of countries, while accounting for volatility clustering in monthly inflation data. Estimation results broadly validate the global dimension of inflation but reveal that the strength of the link between global and domestic inflation is time-varying. Price co-movements have continued to be strongest for advanced economies and have increased considerably in emerging economies in recent years. However, they have remained feeble for low-income countries in the last two decades. Inflation synchronization tends to increase due to oil price shocks affecting most economies in a similar way, global economic expansions or recessions spilling over across economies and owing to more coordinated monetary policy of major central banks. Thus, marked price co-movements indicate the prevalence of common factors affecting inflation across countries.

1. Introduction

One of the most vivid discussion held in the academia with far-reaching economic policy implications concerns the domestic and global determinants of inflation. Since the paradigm of global inflation was established in the influential works by Borio and Filardo (2007) and Ciccarelli and Mojon (2010), a voluminous literature has been devoted to quantifying the extent to which consumer prices are affected by domestic and foreign – potentially global – determinants. The question on the nature of inflation remains an object of intense scrutiny for policy makers (Bernanke, 2007; Carney, 2015; Draghi, 2016; Yellen, 2017) for several reasons. Firstly, central banks adopting the inflation targeting framework are particularly interested in determining the provenance of shocks affecting consumer prices in order to form an optimal policy response in a timely manner. Secondly, effective domestic inflation forecasting requires acknowledging country’s global economic context. Finally, understanding the reasons behind consumer price fluctuations enables a more effective communication policy.

Over time an abundance of approaches have been developed to quantify the global dimension of domestic inflation. A considerable body of empirical literature employs variants of dynamic factor models to measure the contribution of the global, regional and idiosyncratic components to the overall variability in domestic inflation (e.g. Ha et al., 2019; Parker, 2018; Auer et al., 2019; Förster and Tillmann, 2014; Mumtaz and Surico, 2012; Mumtaz et al., 2011; Neely and Rapach, 2011; Ciccarelli and Mojon, 2010; Monacelli and Sala, 2009). With the exception of the study by Förster and Tillmann (2014) providing little support for the hypothesis of global inflation, the evidence put forward in most of these studies suggests that the common component contributes heavily to the overall variation in domestic inflation rates, at least in industrialized countries. Other popular frameworks used to analyse consumer price determinants employ either reduced-form or structural Phillips curve models (e.g. Jašová et al., 2020; Kabukçuoglu and Martínez-Garcia, 2018; Auer et al., 2017; Szafranek, 2017; Milani, 2010; Borio and Filardo, 2007) or various types of vector autoregressions (e.g. Bobeica and Jarociński, 2019; Szafranek and Hałka, 2019; Forbes et al., 2018; Bianchi and Civecci, 2015). These studies underline i) the importance of the global inflation and real economic slack affecting domestic consumer prices either directly or indirectly (Kabukçuoglu and Martínez-Garcia, 2018; Milani, 2010; Borio and Filardo, 2007),...
ii) the sensitivity of inflation to changes in commodity prices both in advanced and emerging economies (Jašová et al., 2020; Bobeica and Jarociński, 2019; Szafranek and Halka, 2019), iii) the increased exposure of economies to common demand and supply shocks due to expanding global value chains (Parker, 2018; Auer et al., 2017) and v) increased financial linkages (Neely and Rapach, 2011).

Most studies on global inflation apply rather homogeneous, advanced-economy, low-frequency samples, which may result in obtaining relatively high estimates of the global component’s contribution to the developments in domestic consumer prices. Recent works by Parker (2018) and Ha et al. (2019) on extensive datasets provide contrasting evidence with respect to the established consensus view on the global dimension of inflation. Drawing from the earlier frameworks provided by Neely and Rapach (2011) as well as Förster and Tillmann (2014), Parker (2018) claims that the explanatory power of global inflation becomes much less pronounced once the sample accounts for middle and low-income economies. Further evidence provided by Ha et al. (2019) corroborates this view. The authors show that although the contribution of the global inflation factor accounts for more variation in inflation rates in 2001–2017 than over 1970–2000, the impact of the global component on domestic consumer prices is still much more considerable in the median advanced economy than in the emerging or low-income ones. Nonetheless, according to Ha et al. (2019) inflation co-movements have recently strengthened and have become more broad-based and sizeable across different inflation measures. Finally, while most studies are conducted on either quarterly or annual data displaying relatively less variation, Altansukh et al. (2017) study inflation co-movements on monthly inflation data, though on a limited dataset of OECD countries. They conclude that energy inflation and food inflation are important drivers of short-run co-movements in aggregate inflation, while core inflation synchronization remains less marked.

Another limitation in the aforementioned literature is the neglect of the volatility clustering present in inflation data. In the seminal work by Engle (1982) the ARCH model is applied to account for conditional heteroskedasticity in UK inflation. Subsequent generalization of this framework by Bollerslev (1986) has led to an eruption of univariate and multivariate extensions of GARCH models, comprehensively reviewed by Bauwens et al. (2006) as well as Silvennoinen and Teräsvirta (2009). Among frameworks frequently used to model conditional heteroskedasticity in the multivariate setting, the Dynamic Conditional Correlation Generalized Autoregressive Conditional Heteroskedasticity model (henceforth DCC-GARCH) introduced by Engle (2002) offers flexibility while alleviating the curse of dimensionality of standard multivariate GARCH models, comprehensively reviewed by Bauwens et al. (2006) as well as Silvennoinen and Teräsvirta (2009). Among frameworks frequently used to model conditional heteroskedasticity in the multivariate setting, the Dynamic Conditional Correlation Generalized Autoregressive Conditional Heteroskedasticity model (henceforth DCC-GARCH) introduced by Engle (2002) offers flexibility while alleviating the curse of dimensionality of standard multivariate GARCH models. The DCC-GARCH is especially popular for modelling exchange rates co-movements, contagion and interdependencies between financial assets (Daelemans et al., 2018; Kim and Sun, 2017; Hemche et al., 2016; Tamakoshi and Hamori, 2013; Celik, 2012). However, several authors rely on this approach to describe the time-varying nexus between macroeconomic variables. These applications include modelling the dynamic relationship between inflation, output and uncertainty (Jones and Olson, 2013), time-varying exchange-rate pass-through (Ozkan and Erden, 2015), business cycle synchronization across countries (Lukenova and Tondl, 2017), systemic risk (Popescu and Turcu, 2017) or the relation between money and output (Ghosh and Parab, 2019).

In this paper, I challenge the well-established consensus view on global inflation given the two aforementioned limitations in the empirical literature and contribute to this strand of economic discussion across two margins. Firstly, I employ a comprehensive and heterogeneous dataset of consumer prices. Specifically, I make use of monthly inflation data for 51 high-income economies, 65 emerging economies and 8 low-income countries. By accounting for a large number of countries, I believe that I can approximate global inflation more precisely and deliver robust conclusions regarding the link between global and domestic inflation. Secondly, I depart from the mainstream approach to modelling global inflation and focus instead on quantifying inflation synchronization. In my approach, for each economy I quantify the unobservable, time-varying link between the global inflation proxy and inflation measured at the country level by employing the convenient DCC-GARCH framework provided by Engle (2002) which retains both flexibility and relative parsimony of specification, while accounting for the volatility clustering present in inflation data. My methodological approach is further motivated by the fact that the literature on measuring inflation synchronization directly is underdeveloped at the current juncture, especially when compared to a large body of studies devoted to global and regional business cycle synchronization and its determinants (e.g. Kollmann, 2019; Ahlborn and Wortmann, 2018; Belke et al., 2017; Ductor and Leiva-Leon, 2016; Duval et al., 2016; Kose et al., 2012; Déès and Zorell, 2012; Darvas and Szapáry, 2008).

Previewing the results, I corroborate the long-standing claim that global and domestic inflation correlate strongly in tightly interconnected, industrialized economies, especially in the euro area countries. Importantly, I unveil a markedly strengthening nexus for emerging economies which may indicate their gradual convergence towards industrialized countries. For low-income countries I show that price co-movements remain weak. An important contribution of this paper consists in establishing the time-varying nature of the relationship between inflation measured at the national level and on a global scale. The evidence presented on a comprehensive dataset validates the paradigm of the global nature of inflation established in previous studies, but shows that the strength of the link between global and domestic inflation is time-varying and can abruptly change throughout major economic periods. Moreover, I show that median inflation synchronization in all country groups (except for the low-income economies) is related to changes in the global economic activity, the level of monetary policy coordination and the variability in oil prices. Lastly, I argue that price co-movements have intensified in all country groups after the global financial crisis. These conclusions remain robust with respect to different measures of global inflation. Additional evidence suggests that core inflation across OECD economies is synchronized quite strongly as well.

From a policy perspective, the measure of inflation synchronization derived from monthly data may serve as a fairly simple and timely indicator of the dependence between domestic and global inflation. Moreover, the development of this measure may indicate the persistence of either idiosyncratic or global factors affecting consumer prices in a given country. Finally, I argue that strong and enduring synchronization of domestic inflation with its global counterpart may either signal that the central bank’s control over the development in consumer prices has weakened or that the central bank’s monetary policy target became more consistent with the targets of most countries.

The remainder of this article is structured as follows. In section 2 I briefly summarize the DCC-GARCH framework. Data used in this study are described in section 3. I report and discuss the main results in section 4, whereas their robustness is checked in section 5. The final section concludes and provides policy recommendations.

2. The model

In order to measure inflation synchronization at the country level I employ the Engle’s (2002) multivariate DCC-GARCH model. Alleviating the curse of dimensionality present in multivariate GARCH frameworks (Bauwens et al., 2006), this approach combines flexibility with relative parsimony of specification, as it allows to model conditional heteroskedasticities and correlations separately and directly. Therefore, it is well suited to quantify the unobservable and time-varying correlation coefficient between the proxy of global inflation and inflation observed at the country level in a straightforward manner, while accounting for volatility clustering in the data.

I start by denoting $y_t = [x_t^*, x_t]'$ as a bivariate vector of global and domestic inflation, respectively, over the sample $t = 1, ..., T$. In the baseline specification domestic inflation $x_t^*$ is defined as the log monthly change in the seasonally and outlier adjusted consumer price.
index, whereas the global inflation proxy $\pi^*_t$ is defined as the first static principal component calculated on standardized domestic headline inflation measures available for all the considered economies using singular value decomposition (alternative proxies for global inflation are considered in section 5). Both $\pi^*_t$ and $\pi_{lt}$ measures enter the model in a standardized form.

The specification of the DCC-GARCH(1,1) model can be represented as:

$$
\gamma_t = \epsilon_t, \quad \epsilon_t \sim IID(0, H_t)
$$

$$
\epsilon_t = H_t^{1/2} z_t
$$

$$
H_t = D_t R_t D_t
$$

$$
D_t = \text{diag}(\sqrt{h_{1t}}, \sqrt{h_{2t}})
$$

$$
Q_t = \sqrt{Q_t} = h_{jt} = a_0 + a_j \epsilon_{jt-1}^2 + b_j h_{jt-1}, \quad j \in \{1, 2\}
$$

$$
R_t = Q_t^{-1} Q_t R_t Q_t^{-1}
$$

$$
Q_t = (1 - a - b)Q + a Q_{t-1} + b Q_{t-1}
$$

In this specification $\epsilon_t$ denotes the vector of innovations characterized by the following properties $E[\epsilon_t] = 0$ and $\text{Cov}[\epsilon_t] = H_t$, i.e. the time-varying conditional variance-covariance matrix $H_t$. $D_t$ represents $2 \times 2$ matrices of conditional standard deviations of $\epsilon_t$ from univariate GARCH(1,1) models and $R_t$ denotes the symmetric dynamic correlation matrix which contains conditional correlation coefficients of the standardized disturbances $Z_t = D_t^{-1/2} \epsilon_t \sim IID(0, R_t)$. Next, $Z_t$ denotes a vector of iid errors such that $E[Z_t] = 0$ and $E[Z_t Z_t'] = I$, whereas $Q = \text{Cov}[Z_t Z_t'] = T^{-1} \sum_{t=1}^T Z_t Z_t'$ denotes the time-invariant unconditional variance-covariance matrix of the standardized errors $Z_t$ and $Q = \text{diag}(\sqrt{Q_{11}}, \sqrt{Q_{22}})$ is the diagonal matrix containing the square root of the diagonal elements of $Q_t$. The parameters $a \geq 0, b \geq 0$ are non-negative scalars governing the process for $Q_t$ and satisfying the stationarity condition $a + b < 1$. Similarly to Lukmanova and Tondl (2017) in my approach I dispense with specifying the mean equation for $\gamma_t$ since I am interested in measuring the time-varying correlation between observed inflation rates instead of idiosyncratic shocks affecting price developments across countries. Given the standardization of the time series estimating a constant is also unnecessary.

The parameters of the DCC-GARCH(1,1) model are estimated in a two-step procedure with the use of the maximum-likelihood method, introduced by Bollerslev et al. (1988) and returning consistent (Bauwens et al., 2006) and – under certain conditions – asymptotically normal estimators (Engle and Sheppard, 2001). However, they are not fully efficient since they are limited information estimators. While maximizing total likelihood provides estimates that are asymptotically efficient (Bauwens et al., 2006), due to the computational burden two-step approaches are widely employed. In such instances, Bauwens et al. (2006) show that the likelihood of the model is decomposed into the sum of the volatility part (depending on the vector $\theta^*$ of univariate model parameters) and the correlation part (depending on the vector of parameters $\theta^*_c$ governing the evolution of the variance-covariance matrix).

By replacing the correlation matrix $R_t$ with an identity matrix, one can represent the quasi log-likelihood function for the first step as a sum of log-likelihood functions of univariate models:

$$
QL1(\theta^*_t) = -\frac{1}{2} \sum_{t=1}^T \sum_{j=1}^2 \{ \log(h_{jt}) + \epsilon_{jt}^2 / h_{jt}^2 \}
$$

Given $\theta^*_t$ the parameters governing the correlation structure may be estimated by maximizing the log-likelihood function for the second step:

$$
QL2(\theta^*_t | \theta^*_t) = -\frac{1}{2} \sum_{t=1}^T \{ \log|R_t| + \epsilon_{t-1}^2 / R_t^{-1} \epsilon_{t-1} \}
$$

In the first step I conduct the estimation separately for each $\pi^*_t$ and each $\pi_{lt}$ univariate GARCH(1,1) models with skewed Student t distribution due to the observed data patterns presented in Table 1. Although the Bayesian information criterion (BIC) suggests that a more parsimonious specification of the model is preferable, pointing mostly to the normal distribution of errors, the outcomes of the adjusted chi-squared goodness of fit test (Vlaar and Palm, 1993) indicate that in the overwhelming majority of cases the empirical distribution of the standardized residuals is in line with the theoretical distribution from the chosen density. In the second step I estimate the parameters governing the correlation structure allowing for different multivariate error distributions. Specifically, I choose between the multivariate normal and Student t distribution based on the BIC. Previewing the results, this criterion only rarely indicates that a distribution potentially accounting for heavy tails should be considered.

Following the estimation, I extract the dynamic conditional correlation coefficient $\rho_{lt} = \rho(\pi^*_t, \pi_{lt})$ which I interpret as a measure of time-varying inflation synchronization between the global inflation proxy and inflation measured at the national level for country $t$. For each $\rho_{lt}$ I calculate the 95% confidence intervals using a deleted-d-jackknife with $d = \sqrt{T} \pm 15$ randomly deleted observations and $n_0 = 1000$ replications.

In the final step, to examine potential differences in the development in inflation synchronization across the globe I categorize all the coun-

| Table 1 | Summary of the descriptive statistics for headline inflation data. |
|---------|---------------------------------------------------------------|
|         | mean | std. dev. | skewness | kurtosis | JB | ADF | ARCH | $\rho$ |
| FR      | 0.001 | 0.001 | ~0.14 | 2.39 | 0.84 | ~10.19 | 18.98 | 0.014 |
| ES      | 0.002 | 0.002 | 0.19 | 2.75 | 3.36 | ~8.10 | 64.80 | 0.253 |
| DE      | 0.002 | 0.003 | 0.42 | 3.03 | 8.67 | ~6.69 | 157.15 | 0.402 |
| IN      | 0.005 | 0.005 | 0.72 | 3.54 | 23.41 | ~5.65 | 390.20 | 0.561 |
| NL      | 0.009 | 0.008 | 2.00 | 9.75 | 604.04 | ~3.96 | 818.24 | 0.683 |
| FR      | 0.584 | 1.000 | 0.920 | 0.888 |

Note: Table 1 presents the descriptive statistics across all headline inflation series prior to standardization using selected percentiles ($p_{05}$) along with the fraction of $H_0$ rejections (FR) for the normality test (Jarque-Bera, JB), unit root test (augmented Dickey-Fuller, ADF), Ljung-Box ARCH test (ARCH) and Pearson’s correlation coefficient ($\rho$) test between global headline inflation approximated by the first static principal component and headline inflation measured at the country level. Significance is checked at $a = 0.05$. For the augmented Dickey-Fuller test a specification with a constant and one lag is chosen. Alternatively, a Phillips-Perron unit root test is conducted with a constant and twelve lags included for the calculation of the Newey-West variance-covariance matrix (results available upon request). According to both tests monthly headline inflation is stationary in all the countries, at least at $a = 0.10$ (unit root tests are notorious for their low power).
tries in the sample according to the World Bank income classification. To this end, apart from investigating inflation co-movements across the whole sample (ALL), I divide the countries in the sample into three groups: highly developed economies (HDC, high-income countries), emerging economies (EME, upper middle-income and lower middle-income countries) and low-income countries (LIC). Additionally, I also examine a tightly integrated group of euro area countries (EA). For these groups I report the median of the dynamic conditional correlation coefficients, which I denote as \( \rho_{ij} \), along with the fraction of country groups in a simple linear framework.

Naturally, more sophisticated univariate and multivariate GARCH models exist. However, often a simple specification is already adequate to capture time-varying volatility and dependence between time series. Simple DCC-GARCH models with multivariate normal error distribution are frequently used in empirical approaches (e.g. Lukmanova and Tondl, 2017; Hemche et al., 2016; Ozkan and Erden, 2015). Moreover, estimating a large number of models for all countries and inflation measures requires some unifying assumptions regarding the specification of the model. Therefore, in this empirical exercise I hold the dynamic specification of the model relatively simple and fixed across countries. Further details regarding DCC-GARCH models and their estimation are discussed in Bauwens et al. (2006) and Silveranninen and Teräsvirta (2009).

### Table 2

Summary of the descriptive statistics for core inflation data.

|        | mean | std. dev. | skewness | kurtosis | JB     | ADF     | ARCH        | ρ     |
|--------|------|-----------|----------|----------|--------|---------|-------------|-------|
| p05    | 0.000| 0.001     | -0.41    | 2.35     | 1.26   | -10.07  | 66.22       | -0.091|
| p25    | 0.001| 0.001     | 0.00     | 2.53     | 3.37   | -7.24   | 141.02      | 0.217 |
| p50    | 0.001| 0.001     | 0.16     | 2.69     | 6.77   | -6.08   | 280.48      | 0.422 |
| p75    | 0.002| 0.002     | 0.59     | 3.32     | 18.11  | -4.99   | 582.50      | 0.598 |
| p95    | 0.004| 0.004     | 1.02     | 4.11     | 47.80  | -3.44   | 1188.54     | 0.713 |
| FR     | 0.564| 1.000     | 0.00     | 2.53     | 3.37   | 0.544   | 1.000       | 0.872 |

Note: Table 2 presents the descriptive statistics across all core inflation series prior to standardization using selected percentiles (\( p_i \)) along with the fraction of \( H_0 \) rejections (FR) for the normality test (Jarque-Bera, JB), unit root test (augmented Dickey-Fuller, ADF), Ljung-Box ARCH test (ARCH) and Pearson’s correlation coefficient (\( \rho \)) test between global core inflation approximated by the first static principal component and core inflation measured at the country level. Significance is checked at \( \alpha = 0.05 \). For the augmented Dickey-Fuller test a specification with a constant and one lag is chosen. Alternatively, a Phillips-Perron unit root test is conducted with a constant and twelve lags included for the calculation of the Newey-West variance-covariance matrix (results available upon request). According to both tests monthly core inflation is stationary in all the countries, at least at \( \alpha = 0.10 \) (unit root tests are notorious for their low power).

### 3. Data

I draw the headline inflation series from the IMF International Financial Statistics database, the BIS Consumer Prices database and the OECD Main Economic Indicators database. In total, data for 124 economies are available over the period from January 2000 to September 2019 (covering roughly 96% of the world GDP throughout the sample). At the expense of a shorter sample, I provide evidence for a large number of countries. This choice is in line with recent arguments by Ha et al. (2019) and Parker (2018), who claim that the use of a comprehensive database allows to approximate global inflation more precisely. The countries entering the sample are listed and categorized into respective groups in Table S1 (in the Supplementary Material, available on-line).

As one of the robustness checks – described in detail further in the text – I also inspect time-varying price co-movements across the core inflation sample. Throughout the paper core inflation is defined as headline inflation net of energy and food prices. Table S2 (in the Supplementary Material, available on-line) reports the 38 countries entering the core inflation sample and accounting for around 64% of the world GDP. The data on core inflation are drawn from the OECD Main Economic Indicators database.

### Table 1

Summary of the descriptive statistics for headline inflation data.

|        | mean | std. dev. | skewness | kurtosis | JB     | ADF     | ARCH        | ρ     |
|--------|------|-----------|----------|----------|--------|---------|-------------|-------|
| H      | 402  | 280.48    | -0.07    | 8.15     | 0.05   | 0.004   | 0.422       | 0.422 |
| F      | 10.07| 66.22     | 0.59     | 3.32     | 18.11  | -4.99   | 582.50      | 0.598 |
| L      | 6.08 | 280.48    | 1.07     | 8.15     | 0.05   | 0.004   | 0.422       | 0.422 |
| L      | 10.07| 66.22     | 0.59     | 3.32     | 18.11  | -4.99   | 582.50      | 0.598 |
| L      | 6.08 | 280.48    | 1.07     | 8.15     | 0.05   | 0.004   | 0.422       | 0.422 |

Note: Table 1 presents the descriptive statistics for all the headline inflation series prior to standardization using selected percentiles (\( p_i \)) along with the fraction of \( H_0 \) rejections (FR) for the normality test (Jarque-Bera, JB), unit root test (augmented Dickey-Fuller, ADF), Ljung-Box ARCH test (ARCH) and Pearson’s correlation coefficient (\( \rho \)) test between global and domestic inflation rates. Significance is checked at \( \alpha = 0.05 \).

Inflation behaves heterogeneously across the countries considered. In particular, it displays relatively large variation relative to its average throughout the sample. The distribution of inflation for most economies is often right-skewed and leptokurtic, deviating from normality for the majority of series (58.4%). All series in the sample are stationary and almost all of them are characterized by volatility clustering, in line with the conclusions from the seminal paper by Engle (1982) for UK inflation. They are also significantly correlated with the global inflation proxy with the median \( \rho \) amounting to 0.402. To conserve space, summary statistics for all the headline inflation series are reported in the Supplementary Material (Table S3).

In a similar fashion, Table 2 reports the descriptive statistics for all the core inflation series prior to standardization. Core inflation across the countries available in the OECD database is similarly heterogeneous as headline inflation, with skewed and leptokurtic distribution, albeit to a smaller extent. Nonetheless, for most series the distribution deviates from normality. The ADF statistic shows that all series are stationary and characterized by the prevalence of the ARCH effect. They are also significantly correlated with the global core inflation proxy, with the median \( \rho \) at 0.422, an estimate slightly higher than for the headline inflation dataset due to a more homogeneous group of countries considered. This anecdotal evidence remains in line with the stylized facts provided by Wang and Wen (2007), who show that the correlation between core inflation measures across highly developed economies remains high, even after accounting for oil shocks. Due to space constraints, I report summary statistics for all the core inflation series in the Supplementary Material (Table S4).

### 4. Results

I start this section by shortly discussing the DCC-GARCH(1,1) estimates for the baseline specification of the model. Next, I describe the development of the time-varying inflation synchronization measure across the whole sample and in specific country groups, which is the main focus of this paper. Further, I establish whether inflation synchronization strengthened after the global financial crisis (henceforth GFC) with the means of the simple Student t-test. I end this section by examining the correlates of inflation synchronization in distinguished country groups in a simple linear framework.
DCC-GARCH estimates. I start the discussion of the results by briefly inspecting the summary statistics for the estimated model parameters across all the countries. These results are reported in Table 3. Parameters $\omega$, $\alpha$, and $\beta$ quantify the constant, the ARCH and the GARCH effect in the univariate part of the DCC-GARCH(1,1) for the global (denoted in the table as GL) and domestic (denoted as CT) inflation, whereas parameters $\lambda$ and $\kappa$ describe the skewness and shape of the skewed Student $t$ distribution of errors. In turn, $a$ and $b$ govern the correlation dynamics in the DCC part of the model, whereas $\Gamma$ defines the shape of the multivariate Student $t$ distribution. ES documents the results of the Engle and Sheppard (2001) test of non-constant correlation. Significance is checked at $a = 0.05$. In the table, $\alpha$, $\beta$ denote the parameters from the standard GARCH with skewed Student $t$ error distribution with parameters $\lambda$ (skewness) and $\kappa$ (shape) for the global headline inflation (calculated as the first static principal component and denoted here as CT), whereas $a$ and $b$ denote the parameters from the DCC model. $\Gamma$ denotes the shape of the Student $t$ multivariate error distribution if chosen by the Bayesian information criterion. $T$ denotes the sample size and $N$ denotes the number of countries in the respective group.

**Table 3** Summary statistics for the estimation of the DCC-GARCH(1,1) model for all countries in the headline inflation sample and across country groups.

| Statistic | $\omega_{GL}$ | $\alpha_{GL}$ | $\beta_{GL}$ | $\lambda_{GL}$ | $k_{GL}$ | $\omega_{CT}$ | $\alpha_{CT}$ | $\beta_{CT}$ | $\lambda_{CT}$ | $k_{CT}$ | $a$ | $b$ | $\Gamma$ | ES |
|----------|----------------|---------------|---------------|----------------|--------|--------------|---------------|---------------|----------------|--------|-----|-----|-------|---|
| 5th percentile | 0.001 | 0.000 | 0.051 | 1.016 | 0.012 | 0.240 | 0.283 | 3.312 | 0.007 | 0.000 | 4.363 | 11.983 |
| 50th percentile | 0.274 | 0.705 | 0.001 | 0.981 | 60.000 | 0.073 | 0.128 | 0.815 | 1.256 | 0.189 | 0.851 | 65.451 | 63.290 |
| 95th percentile | 0.554 | 0.688 | 0.999 | 2.412 | 60.000 | 0.539 | 0.956 | 14.242 | 544.443 |
| Fraction significant | 1.000 | 1.000 | 0.000 | 1.000 | 1.000 | 0.258 | 0.677 | 0.815 | 1.000 | 0.702 | 0.669 | 1.000 | 0.823 |

| Statistic | $\omega_{GL}$ | $\alpha_{GL}$ | $\beta_{GL}$ | $\lambda_{GL}$ | $k_{GL}$ | $\omega_{CT}$ | $\alpha_{CT}$ | $\beta_{CT}$ | $\lambda_{CT}$ | $k_{CT}$ | $a$ | $b$ | $\Gamma$ | ES |
|----------|----------------|---------------|---------------|----------------|--------|--------------|---------------|---------------|----------------|--------|-----|-----|-------|---|
| 5th percentile | 0.003 | 0.000 | 0.004 | 1.016 | 0.012 | 0.240 | 0.180 | 0.981 | 1.448 | 0.000 | 0.112 | 5.669 | 11.312 |
| 50th percentile | 0.274 | 0.705 | 0.001 | 0.981 | 60.000 | 0.093 | 0.180 | 0.741 | 1.380 | 0.060 | 0.252 | 5.885 | 74.536 |
| 95th percentile | 0.342 | 0.475 | 0.999 | 7.006 | 60.000 | 0.425 | 0.944 | 11.484 | 241.892 |
| Fraction significant | 1.000 | 1.000 | 0.000 | 1.000 | 1.000 | 0.216 | 0.667 | 0.902 | 1.000 | 0.824 | 0.764 | 1.000 | 0.804 |

| Statistic | $\omega_{GL}$ | $\alpha_{GL}$ | $\beta_{GL}$ | $\lambda_{GL}$ | $k_{GL}$ | $\omega_{CT}$ | $\alpha_{CT}$ | $\beta_{CT}$ | $\lambda_{CT}$ | $k_{CT}$ | $a$ | $b$ | $\Gamma$ | ES |
|----------|----------------|---------------|---------------|----------------|--------|--------------|---------------|---------------|----------------|--------|-----|-----|-------|---|
| 5th percentile | 0.003 | 0.000 | 0.004 | 1.016 | 0.012 | 0.240 | 0.180 | 0.981 | 1.448 | 0.000 | 0.112 | 5.669 | 11.312 |
| 50th percentile | 0.274 | 0.705 | 0.001 | 0.981 | 60.000 | 0.093 | 0.180 | 0.741 | 1.380 | 0.060 | 0.252 | 5.885 | 74.536 |
| 95th percentile | 0.342 | 0.475 | 0.999 | 7.006 | 60.000 | 0.425 | 0.944 | 11.484 | 241.892 |
| Fraction significant | 1.000 | 1.000 | 0.000 | 1.000 | 1.000 | 0.216 | 0.667 | 0.902 | 1.000 | 0.824 | 0.764 | 1.000 | 0.804 |

Note: Table 3 presents the summary statistics for the estimation of the DCC-GARCH(1,1) model for all the economies and country groups using selected percentiles along with the fraction of $H_0$ rejections (FR) for coefficients tests and the Engle and Sheppard (2001) test (ES) of non-constant correlation. Significance is checked at $a = 0.05$. In the table, $\omega$, $\alpha$, $\beta$ denote the parameters from the standard GARCH with skewed Student $t$ error distribution with parameters $\lambda$ (skewness) and $\kappa$ (shape) for the global headline inflation (calculated as the first static principal component and denoted here as CT), whereas $a$ and $b$ denote the parameters from the DCC model. $\Gamma$ denotes the shape of the Student $t$ multivariate error distribution if chosen by the Bayesian information criterion. $T$ denotes the sample size and $N$ denotes the number of countries in the respective group.

The development of inflation synchronization across the globe. I continue with the discussion of the results by inspecting the developments in the time-varying inflation synchronization across the economies. Fig. 1 displays the time-varying correlation for the median development in the time-varying inflation synchronization across the countries. The upper panel depicts median price co-movements across all the countries (ALL, $\rho_{t\rightarrow m{ALL}}$), whereas further panels illustrate inflation synchronization in the following country groups: highly developed countries (HDC, $\rho_{t\rightarrow m{HDC}}$), emerging economies (EME, $\rho_{t\rightarrow m{EME}}$), low-income countries (LIC, $\rho_{t\rightarrow m{LIC}}$) and euro area countries (EA, $\rho_{t\rightarrow m{EA}}$). Further results with the distribution of $\rho_{t\rightarrow m}$ across countries are available upon request. Additionally, in Table 4...
Fig. 1. Median headline inflation synchronization across the globe and different country groups. Note: The estimates for $\rho_{jt}^{m}$, $j \in \{\text{ALL, HDC, EME, LIC, EA}\}$ across all country groups are marked with the black solid line. Confidence intervals are estimated using a deleted-d-jackknife by generating $n_d = 1000$ subsamples with randomly deleted $d = \sqrt{T} \approx 15$ observations and are illustrated by grey shaded areas. Abbreviations used: ALL - all economies, HDC - highly developed countries, EME - emerging economies, LIC - low-income countries, EA - euro area economies.

I report the mean of the inflation synchronization measures $\rho_{jt}^{m}$, $j \in \{\text{ALL, HDC, EME, LIC, EA}\}$ during major macroeconomic events. These include the recession in the early 2000s (from March 2001 to November 2001, as indicated by NBER recession dates), the final stage of the commodity price boom (from January 2007 to June 2008), the GFC (from December 2007 to June 2009, again indicated by NBER recession dates), the missing disinflation period from January 2009 to December 2011 as in Coibion and Gorodnichenko (2015) and Bobeica and Jarociński (2019), the missing inflation period from October 2011 to December 2014 as in Bobeica and Jarociński (2019) and the oil prices plunge (from June 2014 to January 2016). In choosing these time periods, I draw loosely from Ha et al. (2019), who argue that price co-
movements should rise due to real demand shocks triggering a global recession as well as major oil shocks, as they influence inflation rates similarly across all the importing countries (Choi et al., 2018).

Inflation synchronization across all the economies evolved at a moderate level before the outburst of the GFC. Initially low and stable, price co-movements \( \rho_{t,ALL} \) transitorily spiked during the short recession in 2001 and averaged around 0.31 throughout 2000–2007. In 2008 \( \rho_{t,ALL} \) increased sharply which coincided with the final stage of the commodity price boom and the GFC. This measure exceeded 0.65 in June 2008, indicating strong similarities between inflation rates across the countries. Throughout the initial phase of the recession price co-movements remained robust but the following subsequent recovery in the global economy major increases in inflation synchronization eventually reversed, falling below 0.3 in 2012. Since then, they started to intensify again and during the missing inflation period coupled with oil price plunges of 2014–2016 they approached 0.60 one more time. Overall, after the GFC \( \rho_{t,ALL} \) increased to around 0.41, with the measure reaching 0.5 at the end of the sample (in September 2019).

In HDCs inflation synchronization was already solid before the GFC. The mean estimate of \( \rho_{t,HDC} \) hovered around 0.46 throughout 2000–2007, much higher than \( \rho_{t,ALL} \). The link was even stronger for highly integrated euro area economies, with \( \rho_{t,EA} \) fluctuating around 0.61 in 2000–2007. Between the recession in the early 2000s and the commodity prices boom inflation synchronization remained moderate and relatively stable. Conversely, both in the case of HDC as well as EA countries during the surge in commodity prices, the commonality in inflation rates grew substantially, with \( \rho_{t,HDC} \) and \( \rho_{t,EA} \) peaking at 0.73 and 0.78 in June 2008, respectively. During the recession as well as the subsequent missing inflation period the strength of price co-movements fell to around 0.4 and 0.5 in November 2011 for HDC and EA countries, respectively. Similarly to the mean for all economies in the sample, throughout the low inflation period and plummeting oil prices inflation synchronization increased sharply again, approaching 0.7 and 0.8 for HDC and EA countries in January 2015, respectively. Overall, after the GFC the estimate of \( \rho_{t,HDC} \) increased to around 0.52, whereas the estimate of \( \rho_{t,EA} \) amounted to 0.64. For both these country groups inflation synchronization exceeded 0.5 at the end of the sample.

For the emerging economies price synchronization was initially low, with the \( \rho_{t,EME} \) estimate evolving around 0.23. It bolstered considerably in 2007, which, again, coincided with surging commodity prices. This increase reversed rather quickly throughout the recession in the global economy. Following the pattern for HDCs, \( \rho_{t,EME} \) began to increase steadily during the low inflation period and the oil price plunges, with the measure remaining elevated around 0.4–0.5 since 2015. After the GFC the estimate for \( \rho_{t,EME} \) rose to around 0.36 and approached 0.5 at the end of the sample, in September 2019, running at a similar level as \( \rho_{t,HDC} \). Moreover, a positive trend is visible for inflation synchronization in EMES after the GFC, which could signal their increasing integration into the world economy and gradual synchronization of business cycles (Ductor and Leiva-Leon, 2016). Recently, Jašová et al. (2020) indicate as well that emerging economies may catch up with advanced economies on inflation dynamics.

Conversely, a feeble link between domestic and global inflation persists in LICs with the estimate for \( \rho_{t,LIC} \) fluctuating around 0.15 in 2000–2007, somewhat below the measure obtained for EMES. In 2008 a short spike in the median of \( \rho_{t,LIC} \) to above 0.4 is visible, coinciding with a presence of a common shock (the GFC). After the decline following the global recession, the estimate for \( \rho_{t,LIC} \) slightly increased and averaged 0.22. In contrast to HDCs, EA and EMES, no pattern is visible throughout the missing inflation period and plunging oil prices. Moreover, at the end of the sample \( \rho_{t,LIC} \) stabilized. This suggests that LICs are still detached from the developments in the global economy and the link between domestic and global inflation remains weak. A caveat should be added here, as only 8 countries constitute this group. Therefore, I treat these results as tentative and check their robustness in the next section.

Additional statistics reported in Table 4 illustrate that throughout major global economic events headline inflation exhibited strong commonalities, especially in HDCs and EA, but to a lesser extent in EMES as well. This signals the importance of common shocks spilling across countries in shaping the co-movements in consumer prices. On the other hand, price co-movements in LICs remained low and occasionally insignificant, despite real demand and oil price shocks hitting the global economy. The evidence provided here compliments the view of Ha et al. (2019), who claim that inflation synchronization after 2000 has intensified, predominantly owing to the two exceptional periods, i.e. the global recession and the collapse in oil prices of 2014–2016.

### Table 4

Mean level of inflation synchronization throughout specific global economic events in particular country groups.

| Recession in the U.S. economy | Commodity price boom | Global financial crisis | Missing disinflation period | Missing inflation period | Oil price plunge |
|-------------------------------|----------------------|------------------------|----------------------------|-------------------------|-----------------|
| 2001:03–2001:11 | 2007:01–2008:06 | 2007:12–2009:06 | 2009:01–2011:12 | 2011:11–2014:12 | 2014:06–2016:01 |
| ALL | 0.311 | 0.456 | 0.534 | 0.362 | 0.368 | 0.503 |
| [0.173–0.415] | [0.335–0.555] | [0.425–0.659] | [0.253–0.492] | [0.244–0.472] | [0.384–0.606] |
| HDC | 0.488 | 0.586 | 0.626 | 0.490 | 0.509 | 0.636 |
| [0.379–0.604] | [0.498–0.659] | [0.562–0.712] | [0.398–0.594] | [0.416–0.600] | [0.579–0.702] |
| EME | 0.213 | 0.362 | 0.473 | 0.311 | 0.293 | 0.408 |
| [0.058–0.355] | [0.213–0.493] | [0.346–0.622] | [0.177–0.450] | [0.153–0.404] | [0.272–0.534] |
| LIC | -0.016 | 0.251 | 0.259 | 0.158 | 0.237 | 0.264 |
| [-0.145–0.116] | [0.124–0.344] | [0.175–0.364] | [0.060–0.249] | [0.146–0.311] | [0.183–0.331] |
| EA | 0.663 | 0.665 | 0.716 | 0.636 | 0.639 | 0.745 |
| [0.606–0.743] | [0.602–0.726] | [0.658–0.786] | [0.569–0.699] | [0.565–0.700] | [0.790–0.803] |
| OECD | 0.575 | 0.447 | 0.506 | 0.363 | 0.423 | 0.534 |
| [0.469–0.677] | [0.311–0.560] | [0.384–0.624] | [0.241–0.514] | [0.289–0.545] | [0.436–0.633] |

Note: Table 4 presents the mean of the time-varying inflation synchronization in particular country groups throughout several specific global economic periods. In square brackets the average of the 95% confidence levels are reported. The dates for the recession in the U.S. economy in the early 2000s and during the global financial crisis are chosen in accordance with the NBER recession dates. The period for the commodity prices boom denotes the final stage of the oil prices run-up. The missing disinflation period is defined as in Colibin and Gorodnichenko (2015), whereas the missing inflation period is defined as in Colibin and Gorodnichenko (2015) and Bobeica and Jarociński (2019). The oil price plunge denotes the period of oil prices falling steadily to their lowest level in 2016. The statistics provided in this table are reported for the baseline model, where the global headline inflation is proxies by the first static principal component. The final row of the table reports the statistics for a robustness check, where core inflation data are considered in the OECD sample, as described in section 5. Abbreviations used: ALL - all economies, HDC - highly developed countries, EME - emerging economies, LIC - low-income countries, EA - euro area economies.
To answer this question I make use of the simple Student t-test for the equality of means: 

\[ H_0 : \bar{\mu}_1 = \bar{\mu}_2 \]  

\[ H_1 : \bar{\mu}_1 < \bar{\mu}_2 \]  

where: \( \bar{\mu}_1 \) denotes the mean inflation synchronization for country group \( j \) on the sample preceding the GFC (from February 2000 to November 2007) and \( \bar{\mu}_2 \) denotes the mean inflation synchronization for country group \( j \) on the sample following the GFC (July 2009–September 2019). The null hypothesis states that the strength of inflation synchronization has not changed (the difference in means is equal to zero), whereas the alternative says that inflation synchronization has increased. The first (second) sample uses estimates of \( \rho_j \), \( j \in {\text{ALL, HDC, EME, LIC, EA}} \) for each country group on the sample February 2000–November 2007 (July 2009–September 2019). Abbreviations used: ALL - all economies, HDC - highly developed countries, EME - emerging economies, LIC - low-income countries, EA - euro area economies. Global inflation is proxied either by the first static principal component (PC), median headline inflation (MED) or weighted average inflation across all the countries in the sample (WGT) with weights resembling countries’ real GDP (in 2010 US dollars) drawn from the World Bank’s World Database Indicators. Asterisks ***, ** denote significance at \( \alpha = 0.01 \) and \( \alpha = 0.05 \), respectively.

The data for this exercise have either been collected from OECD and FRED databases or provided by Baumeister and Hamilton (2019).

Table 6 summarizes the outcomes of the fitting procedure using OLS with robust standard errors. Median inflation synchronization tends to increase during expansionary periods in the global economy for all country groups. When headline inflation is considered, rising similarities in major central banks’ monetary policy (reflected in the falling INT measure) strengthen median price co-movements calculated for all country groups, except LICs. This is in line with the claim by Clarida et al. (2002) and Neely and Rapach (2011) who argue that coordinated monetary policy can lead to increased similarities in price developments across countries. Furthermore, oil price shocks tend to increase median inflation synchronization (again in all country groups except the LIC), as they propagate similarly to domestic inflation in most countries (Choi et al., 2018). Median price co-movements calculated for all country groups except the LICs attune as well during recession in the US economy. Finally, with inflationary pressure falling median inflation synchronization tends to increase across all the economies and for EMEs and LICs (for HDCs and EA this link is also negative, but statistically insignificant at \( \alpha = 0.05 \)). This would suggest that price co-movements intensify when domestic (idiosyncratic) inflationary pressures are contained. The evidence presented in this simple exercise indicates that global economic conditions affect inflation synchronization in most country groups in line with the economic intuition. Interestingly, inflation developments remain detached from global movements in low-income countries. It should be emphasised here that a more detailed analysis of the sources of inflation synchronization could be extended to panel frameworks. This, however, remains beyond the scope of this paper and is an interesting avenue for future research.

5. Sensitivity analysis

In this section I test the robustness of the baseline results from the previous section by conducting four additional exercises. Firstly, I re-estimate all models using other proxies for global inflation. This is justified by the concern that a relatively large number of high-income and emerging economies in comparison to low-income countries may skew the results once the PCA approach is used to proxy global inflation. To this end, I approximate the global inflation by the median seasonally and outlier adjusted headline inflation across all the countries in the dataset. I also calculate the global inflation as the weighted headline inflation with weights resembling countries’ real GDP (in 2010 US dollars) drawn from the World Bank’s World Database Indicators. In this approach large countries would affect the global inflation proxy more heavily. The development in these three measures is illustrated on Figure S1 (in the Supplementary Material, available on-line). I also
check the robustness of the estimate of the first principal component as inflation data are non-Gaussian by employing the robust PCA approach proposed by Hubert et al. (2009), a method accounting for outliers and skewness present in the data.

Secondly, to double-check the feeble inflation synchronization in low-income economies I re-estimate the model on a shorter sample (spanning the period January 2002–December 2017), as there is a considerable publication lag for some least developed economies. This enabled me to extend the low-income country group to 17 economies and include 153 economies in total, accounting for up to 97% of world GDP. Thus, the fraction of least developed countries rises from around 6.5% in the baseline sample to around 11% in the shorter sample.

Thirdly, to assess the synchronization of inflation measures without the potentially considerable impact of volatile commodity prices on headline inflation, I conduct my exercise on core inflation data. This is motivated by the fact that food and energy inflation may be an important source of headline inflation co-movement (Parker, 2018; Altansukh et al., 2017), whereas core inflation should better resemble idiosyncratic factors affecting consumer prices in each country.

Finally, I perform a similar exercise for producer price inflation. This analysis has been done for 54 countries accounting for roughly 71% of world GDP. However, given space constraints and due to the fact that I am primarily interested in the co-movement between consumer prices the results on producer price inflation synchronization are available upon request. In what follows, I discuss the results from the first three robustness checks.

**Employing different proxies of global inflation.** I start the sensitivity analysis by focusing on the development of inflation synchronization $\rho_{ij}^t$ when global inflation is proxied by other measures. These results are depicted in Fig. 2. The main takeaway from this robustness check is that the results remain qualitatively similar, irrespective of the global inflation proxy. When the median inflation is considered (red solid line), the estimates of $\rho_{ij}^t$ (for $j \in \{\text{ALL, HDC, EME, LIC, EA}\}$) closely follow the estimates for the baseline specification. In turn, inflation synchronization is occasionally somewhat weaker when the weighted headline inflation approximates global inflation (blue solid line), but no apparent changes in the pattern throughout time emerge. This results from the fact that the development of weighted headline inflation is somewhat different from the other measures (Figure S1 in Supplementary Material, available on-line) due to the applied weights. Similar anecdotal evidence is presented by Ciccarelli and Mojon (2010). The results of the simple Student $t$-test for the equality of means (Table 5, rows 2–3) do not change when alternative measures of global inflation are considered. Finally, I show that the estimates of the global inflation proxy using a standard PCA approach and a framework accounting for skewness and outliers in the data (Hubert et al., 2009) are very similar (Figure S2 in the Supplementary Material, available on-line). For the sake of brevity, DCC-GARCH model estimates with alternative global inflation measures are not reported either here or in the Supplementary Material and are available upon request.

**Inflation co-movements for low-income countries on a shorter sample.** In the next step I check whether inflation synchronization is affected by the sample composition. Figure S3 (in the Supplementary Material, available on-line) compares the baseline estimates of price co-movements across chosen country groups (black solid line) with the estimates obtained on a restricted sample (red solid line). The results obtained on a shorter sample remain qualitatively the same, especially across all the countries and in the highly developed and emerging economies. Inflation synchronization remains feeble for LICs after accounting for a larger number of the least developed countries, with the estimate slightly rising only temporarily during the GFC. This outcome is similar to previous findings in the literature. Ha et al. (2019) conclude that the factor loadings on the global factor are not statistically significantly different from zero for several least developed countries. Moreover, the evidence presented by Ha et al. (2019) and Parker (2018) also shows that the explanatory power of global inflation in LICs is negligible. On a side note, extending the sample to account for the most recent data (up to June 2020) and thus effectively making the sample more homogeneous (by reducing the number of emerging and low-income countries) increases the degree of inflation synchronization throughout the sample while retaining the overall tendency of the measure (with a significant spike after the outbreak of the COVID-19 pandemic). This corroborates the

**Table 6**

Correlates of the median inflation synchronization across country groups.

|         | ALL (1) | HDC (2) | EA (3) | EME (4) | LIC (5) | OECD (6) |
|---------|---------|---------|--------|---------|---------|----------|
| BH      | 3.279*** | 2.882*** | 1.387** | 3.128*** | 2.463*** | 3.477*** |
| INT     | −0.203** | −0.207*** | −0.172*** | −0.277*** | 0.013 | −0.147 |
| OILV    | 1.882*** | 1.697*** | 1.314*** | 2.019*** | −0.081 | 1.367*** |
| USR     | 0.143*** | 0.134*** | 0.088*** | 0.146** | 0.013 | 0.176*** |
| INF     | −0.037* | −0.029* | −0.013 | −0.041** | −0.039*** | −0.044*** |
| Constant| 0.221*** | 0.363*** | 0.529*** | 0.149*** | 0.204*** | 0.295*** |

R$^2$: 0.365 0.355 0.245 0.304 0.164 0.242
F Statistic: 26.477*** 25.287*** 14.961*** 20.096*** 9.027*** 14.688***
Observations: 236 236 236 236 236 236

Note: Table 5 presents the results of a simple ordinary least square regression of the median inflation synchronization in a specific country group on several variables approximating global macroeconomic conditions. Abbreviations used: BH – Hodrick-Prescott filtered value of the novel Baumeister and Hamilton (2019) index of global economic activity, INT – a measure of central banks’ monetary policy similarity calculated as the standard deviation of the change in the policy rates across five major central banks, i.e. the FED, the ECB, the BoJ, the BoE and the BoC, OILV – conditional volatility of oil prices calculated from an exponential GARCH(1.1) model with skewed Student t distribution, USR – a dummy variable depicting recession dates in the US and INF denoting the first static principal component proxying the level of global inflation. Robust standard errors calculated using HAC estimators are reported in parentheses. Asterisks ***, ** and * denote significance at $\alpha = 0.01$, $\alpha = 0.05$ and $\alpha = 0.10$, respectively.
Fig. 2. Sensitivity of the measure of headline inflation synchronization with respect to different global inflation proxies. Note: The black solid line illustrates the estimate for $\rho_{m,j}^{PC}$, $j \in \{ALL, HDC, EME, LIC, EA\}$, when global inflation is proxied with the first static principal component (PC), the red solid line illustrates the median estimate for $\rho_{m,j}^{MED}$ when global inflation is proxied with the median inflation across all the countries (MED) and the blue solid lines depicts the median estimate for $\rho_{m,j}^{WGT}$ when global inflation is proxied with weighted average inflation across all the countries in the sample (WGT) with weights resembling the country’s real GDP (in 2010 US dollars) drawn from the World Bank’s World Database Indicators. Abbreviations used: ALL - all economies, HDC - highly developed countries, EME - emerging economies, LIC - low-income countries, EA - euro area economies. (For interpretation of the references to colour in this figure legend, the reader is referred to the Web version of this article.)
argument that in predominantly homogeneous, advanced-economy samples the contribution of the global component to the developments in domestic consumer prices is relatively high. The estimates obtained on most recent data are available upon request.

**Additional results for core inflation.** Next, I turn my attention to the synchronization of core inflation. Firstly, I inspect the estimates of the DCC-GARCH models for the core inflation measure. Table 7 presents the summary statistics for the estimation of the DCC-GARCH models across all the countries in the core inflation sample in a similar manner to Table 3.

Contrary to the headline inflation case, in both the univariate and multivariate part of the model the GARCH effects remain stronger than the ARCH effects. The sum of the $a$ and $b$ as well as $a$ and $b$ coefficients approaches 1, suggesting high variance persistence of the modelled univariate and multivariate processes. The estimate $\lambda$ indicates that the correction for skewness is necessary, but parameter $\kappa$ suggests that heavy tails only rarely have to be accounted for. Moreover, in the multivariate part the Student t distribution is never preferred as parameter $\Gamma$ is not estimated – this indicates that the tails of the multivariate distribution are lighter with respect to models for headline inflation, which follows the economic intuition regarding the variability of core inflation measures. Finally, the results of the Engle and Sheppard (2001) test imply that in 95% of countries inflation synchronization is in fact time-varying. Due to space constraints, model estimates for all the economies are reported in the Supplementary Material (Table S6).

In the next step I investigate the pattern of the time-varying synchronization of core inflation. Fig. 3 illustrates this development across all the economies in the core inflation sample (black solid line) along with the 95% deleted-d-jackknife confidence intervals (grey shaded area). For core inflation price co-movements were already strong at the beginning of the sample with several countries undergoing a disinflation process in the early 2000s. This was followed by a period of exceptionally weak and occasionally statistically insignificant price co-movements, which indicates a strong influence of idiosyncratic components on domestic core inflation measure. Throughout the surge in commodity prices, their subsequent collapse with the GFC and the period of waning global demand and plunging oil prices, core inflation synchronization strengthened again (Table 4, last row), even exceeding 0.58 in May 2015. Following further recovery in the global economy, co-movements of core inflation gradually waned. Overall, except for the early 2000s, the development in core inflation synchronization closely follows the measure of headline inflation commonality. Moreover, price co-movements intensified as well after the GFC (Table 5). Table 6 shows that they correlate positively (negatively) with the global business cycle and oil shocks (the level of global core inflation). The evidence provided here stands in slight contrast to Altansukh et al. (2017) and Parker (2018), who show that price co-movements rather stem from the global dimension of energy and, to a lesser extent, food components in the inflation basket.

**6. Conclusion**

In this study I have applied the dynamic conditional correlation GARCH framework to analyse the time-varying relationship between global and domestic inflation measured at a monthly frequency on a comprehensive dataset of 124 countries. The presented evidence corroborates the established consensus view on a robust link between global and domestic inflation for tightly integrated high-income economies. This nexus remains consistently strongest in the euro area countries. Importantly, I unveil increasing international infla-

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**Table 7**

Summary statistics for the estimation of the DCC-GARCH(1,1) models for all the countries in the core inflation sample.

| Statistic | $\alpha_{a}$ | $\alpha_{b}$ | $\beta_{a}$ | $\beta_{b}$ | $\lambda_{a}$ | $\lambda_{b}$ | $\kappa_{a}$ | $\kappa_{b}$ | $a$ | $b$ | $\Gamma$ | ES |
|-----------|--------------|--------------|-------------|-------------|--------------|--------------|-------------|-------------|-----|----|--------|-----|
| 5th percentile | 0.001 | 0.000 | 0.079 | 0.736 | 0.526 | 0.025 | 0.326 | . | 24.802 |
| 50th percentile | 0.028 | 0.258 | 0.741 | 1.872 | 60.000 | 0.602 | 0.124 | 0.819 | 1.114 | 60.000 | 0.156 | 0.777 | . | 145.265 |
| 95th percentile | 0.294 | 0.836 | 0.999 | 2.607 | 60.000 | 0.482 | 0.928 | . | 836.022 |
| Fraction significant | 0.000 | 1.000 | 1.000 | 1.000 | 1.000 | 0.237 | 0.632 | 0.947 | 1.000 | 0.789 | 0.816 | 0.947 | . | 0.947 |

Note: Table 7 presents the summary statistics for the estimation of the DCC-GARCH(1,1) model for all economies in the core inflation sample using selected percentiles along with the fraction of $H_0$ rejections (FR) for coefficient tests and the Engle and Sheppard (2001) test (ES) of non-constant correlation. Significance is checked at $\alpha = 0.05$. $\alpha$, $\beta$, $\lambda$ denote parameters from the standard GARCH with skewed Student t error distribution with parameters $\kappa$ (skewness) and $\kappa$ (shape) for global core inflation (calculated as the first static principal component and denoted here as GL) and core inflation measured at the country level (here denoted as CT), whereas $a$ and $b$ denote parameters from the DCC model. $\Gamma$ denotes the shape of the Student t multivariate error distribution if chosen by the Bayesian information criterion. $T$ denotes sample size and $N$ denotes the number of countries for which core inflation data is readily available in the OECD database.

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![Fig. 3](image-url) Median core inflation synchronization across the globe. Note: The estimates for $\rho_{M,OEC}$ is marked with the black solid line. Confidence intervals are estimated using a deleted-d-jackknife by generating $a_d = 1000$ subsamples with randomly deleted $d = \sqrt{T} \approx 15$ observations and are illustrated by grey shaded areas.
tion synchronization in emerging economies, which may reflect their gradual convergence towards developed economies. For low-income countries I show that this link remains weak. Moreover, the evidence suggests that after a period of very low comonality in core inflation rates across countries, their synchronization increased. Obtained estimates reveal that the degree of price co-movements is time-varying. In particular, inflation synchronization is more marked during the commodity price run-up, throughout the GFC and the oil price plunge, but following the subsequent normalization of economic condition it becomes less distinct. By employing a simple linear framework I show that the measure of inflation synchronization correlates positively with changes in the global economic activity and with the variability in oil prices and negatively with the measure of monetary policy synchronization for all country groups, except for the LICs, while the level of inflation affects median inflation synchronization in emerging and low-income countries. This can be interpreted that potentially global economic shocks increase similarities in inflation rates across the majority of countries. Finally, a simple testing procedure shows that inflation synchronization has become stronger after the GFC. The results presented in the paper are robust with respect to different measures of global inflation.

This study reaches three important conclusions from a policy maker’s perspective. First, the measure of inflation synchronization derived from monthly data may serve as a fairly simple and timely indicator of the dependence between domestic inflation and its global counterpart. Secondly, the development of this measure in time may indicate the persistence of either domestic or global factors affecting inflation. Furthermore, in economies with independent monetary policy strong inflation synchronization for a prolonged period of time may signal that the central bank’s control over domestic inflation has weakened. However, strong inflation commonalities may also indicate that monetary policy is less oriented towards domestic price developments or that it has become more consistent with the external environment. Finally, with inflation remaining persistently at an undesirable level and the global inflation synchronization remaining high, it may turn out that only coordinated changes in monetary policy across major economies may be able to influence consumer price developments globally.

Further research on inflation synchronization may take two directions. Firstly, the question remaining beyond the scope of this paper regards the determinants of price co-movements measured at the country level. Evidence based on simple linear framework as well as the claim presented by Ha et al. (2019) indicate that shifts in inflation synchronization stem from common shocks spreading simultaneously across economies, e.g. real demand shocks sparking global recessions or oil shocks, or structural changes in economies illustrated by the increasing trade integration, especially through global supply chains, financial integration as well as technological advances. While this paper provides only partial evidence as to the potential sources of inflation synchronization – indicating that indeed throughout global economic expansions and recessions or during excessive volatility in commodity prices inflation synchronization increases sharply – some studies have already determined the cross-sectional differences between countries affecting said co-movements (Ha et al., 2019; Parker, 2018; Neely and Rapach, 2011). That said, further research could extend these frameworks to a panel setting to provide a more comprehensive view. Secondly, literature has indicated that commonalities in nominal variables are far stronger than the synchronization of output (e.g. Ha et al., 2019; Henrilsen et al., 2013; Muntaz et al., 2011; Wang and Wen, 2007). Therefore, further research might also focus on explaining these divergences.

Declaration of competing interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.econmod.2020.09.013.

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