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Have the driving forces of inflation changed in advanced and emerging market economies?*

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Abstract

We construct a balanced panel dataset for 47 advanced and emerging market economies over a sample period from 1996 to 2018 to empirically investigate possible changes in the driving forces of inflation. Using an open economy hybrid Phillips curve model of inflation and formally testing for structural breaks, we find relatively little significant change in the underlying driving forces or their quantitative effects for most economies, even after the Great Financial Crisis. However, one notable change has been an increase in the average weight on expected future inflation, measured using professional forecasts, for both advanced and emerging market economies. We find very heterogeneous but significant effects of inflation expectations, domestic and foreign output gaps, exchange rate passthrough, and oil prices, with generally higher sensitivities to external driving forces for emerging market economies. Consistent with the model, the behavior of the various inflation drivers, especially what appear to be better anchored inflation expectations, can explain patterns of changes in the level and volatility of inflation across different economies.

Keywords: open economy Phillips curve; structural breaks; inflation expectations; exchange rate passthrough; inflation volatility

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1 Introduction

Inflation has remained remarkably stable in many economies since the Great Financial Crisis (GFC). This stability of inflation has been attributed to a flattening of the Phillips curve and/or an anchoring of inflation expectations (e.g., IMF 2013). Another hypothesized source of structural change that could explain the relative stability of inflation is increased foreign competition and trade integration (e.g., Forbes 2018). With the recent COVID-19 crisis, interest in what drives inflation has only increased as many central banks reconsider how their monetary policy frameworks relate to inflation.

In this context, we empirically investigate possible changes in the driving forces of inflation before, during, and after the GFC for 47 advanced and emerging market economies. To do so, we construct a balanced panel dataset with key variables that include expected future inflation measured using professional forecasts and a foreign output gap measured using economy-specific trade weights. These variables allow us to estimate an open economy hybrid Phillips curve model of inflation and we formally test for structural breaks in the model parameters. We also track what the behavior of variables in the model would predict for the level and volatility of inflation over time.

Perhaps surprisingly, we find relatively little significant change in the driving forces of inflation or their quantitative effects for most of the 47 economies over the sample period from 1996 to 2018. However, a notable change has been an increase in the average weight on expected future inflation for both advanced and emerging market economies. We find very heterogeneous but significant effects of inflation expectations, domestic and foreign output gaps, exchange rate passthrough, and oil prices, with generally higher sensitivities to external driving forces for emerging market economies. Consistent with the open economy hybrid Phillips curve model of inflation, the behavior of the various inflation drivers, especially what appear to be better anchored inflation expectations, can explain patterns of changes in the level and volatility of inflation across different economies.

Our paper relates to a number of strands of literature on what drives inflation, including on differences for advanced and emerging market economies (e.g., Blanchard, Cerutti and Summers 2015, IMF 2016; 2018, Miles et al. 2017, Jorda and Nechio 2018, Ha, Kose and Ohnsorge 2019, Kamber and Wong 2020), forward-looking inflation expectations (e.g., Fuhrer 2012, Coibion and Gorodnichenko 2015, Cecchetti et al. 2017, Ball and Mazumder 2019), global factors (e.g., Borio and Filardo 2007, Monacelli and Sala 2007, Ciccarelli and Mojon 2010, Guerrieri, Gust and Lopez-Salido 2010, Ihrig et al. 2010, Milani 2010, Mumtaz and Surico 2012, Bianchi and Civelli 2015, Auer, Borio and Filardo 2017), and exchange rate passthrough (e.g., Choudhri and Hakura 2006, Mihaljek and Klau 2008, Jasova, Moessner and Takats 2016).

The rest of this paper is organized as follows: Section 2 provides background to our analysis by surveying the extensive empirical literature on the behavior of inflation in advanced and emerging market economies. Section 3 describes the data used in our analysis. Section 4 presents our empirical model and discusses the econometric methods employed in our analysis. Section 5 reports our empirical results. Section 6 concludes.

2 Background

The empirical literature on the behavior of inflation in advanced and emerging market economies has proliferated over the past decade (see, among many others, Ha, Kose and Ohnsorge 2019, Miles et al. 2017, IMF 2013; 2016; 2018). In this section, we provide an overview of some stylized facts about inflation based on recent findings in the literature.

The first stylized fact relates to the relationship between inflation and the real economy, or the slope of the Phillips curve. After being on a downward path for much of the 1990s and the

2000s, inflation in many economies has remained relatively stable since the GFC, even though this period would seem to have led to large negative output gaps worldwide. The fact that inflation did not fall as much as might have been expected given a high degree of economic slack during the crisis or rise much in the recovery afterwards has raised questions about what has contributed to the recent inflation stability and whether the slope of the Phillips curve has flattened.

A second stylized fact about inflation that has emerged in recent discussions is the importance of common factors in its evolution. Principal components analysis of inflation across economies suggests a single common factor can explain a high share of variance of inflation (on the order of 45-50%). The evidence presented by IMF (2016) and Ha, Kose and Ohnsorge (2019) suggests that this common component of inflation is highly correlated with global economic forces.

A third stylized fact about inflation that has played a prominent role in recent discussions and empirical models of inflation relates to the information content of inflation for its own future evolution. This issue was highlighted by Stock and Watson (2007) when discussing the predictive properties of inflation in a univariate framework. Decomposing inflation into a trend and a transitory component in a time varying trend-cycle model, Stock and Watson (2007) show that, while the volatility of shocks to the transitory component of US inflation has remained unchanged, the volatility of shocks to trend inflation has fallen since the mid-1980s. One implication of this finding is that, with fluctuations to trend inflation fading away, inflation has become less persistent as well as less predictable over the past three decades. Cecchetti et al. (2007) extend Stock and Watson's analysis to other G7 economies and note that the decline in volatility of trend inflation is not just a US development but a common phenomenon across advanced economies since the mid-1980s, appearing to coincide with changing public policy preferences towards low and stable inflation, particularly with the shift of monetary policy to inflation-targeting regimes.

Two dominant hypotheses have emerged over the past decade to explain these stylized facts. The first one, suggested by Bernanke (2007) and Mishkin (2007), relies on the premise that more stable and better anchored inflation expectations have helped to keep inflation itself more stable in the face of large negative demand shocks, or what is referred to as the 'anchored inflation expectations' hypothesis. This view has stressed the role of monetary factors in driving recent changes in the behavior of inflation. A second hypothesis, emphasized by Forbes (2018), has pointed to the importance of structural changes in the global economy, including the growing integration of labour and product markets across different economies, driving changes in the relationship between inflation and the real economy.

A standard theoretical construct used to examine these hypotheses is an open economy hybrid New Keynesian Phillips curve, which takes the following general form:

$$\pi_t = (1-\gamma)\pi_{t-1} + \gamma E_t[\pi_{t+1}] + \kappa \tilde{y}_t + \lambda z_t \tag{1}$$

where π_t is current inflation, π_{t-1} is lagged inflation and expectations for backward-looking firms, $E_t[\pi_{t+1}]$ is the rational expectation of future inflation and expectations for forward-looking firms, $\tilde{y}_t$ is the domestic output gap serving as a proxy for marginal costs for domestic factors of production, and z_t represents a set of global demand and supply variables serving as a proxy for marginal costs for foreign factors of production.[1] The views supporting structural changes in the global economy have generally pointed to a smaller value of κ and a larger value of λ in the determination of inflation. Meanwhile, those supporting a stronger role for monetary policy in determining the inflation process have pointed to a larger value of γ going hand in hand with a smaller value of κ.

[1]See Roberts (1995) and Gali and Gertler (1999) for theoretical foundations behind the expectations-augmented New Keynesian Phillips curve.

One implication of the anchored inflation expectations hypothesis is that, with current inflation being well-guided by agents' expectations about future price developments in the economy, inflation could become less responsive to demand and supply shocks. According to Bernanke (2007), "If people set prices and wages with reference to the rate of inflation they expect in the long run and if expectations respond less than previously to variations in economic activity, then inflation will become relatively more insensitive to the level of economic activity – that is the conventional Phillips curve will become flatter."

Several recent studies have reported estimates that are consistent with the anchored inflation expectations hypothesis, although the exact channel through which it may have influenced the behavior of inflation remains unclear. For instance, Ball and Mazumder's (2019) estimate suggest that γ has become closer to one for the United States in a model that includes a short-term unemployment gap, lagged inflation, and the 10- year ahead survey of professional forecasters' inflation expectations. The structural break tests reported in the study show that the break to the parameter occurred in 1998. Accounting for this break, the model produces stable Phillips curve estimates and US inflation forecasts that are close to actual inflation. Blanchard, Cerutti and Summers (2015) report Phillips curve estimates for 23 advanced economies while accounting for the hysteresis effects of financial crisis by allowing the natural rate and the coefficients to vary over time. They show that the median estimates of γ has increased since the mid-1980s, with the Phillips curve becoming flatter over time. However, they report that there have been no major changes to κ since the early 1990s. Jorda and Nechio (2018) compare Phillips curve estimates before and after the GFC for a large number of crisis-hit and crisis-missed economies to find the role of common factors in global inflation development. They suggest that, in most economies, the Phillips curve flattened after the GFC, with a gradual increase in the role of forward-looking inflation expectations in the determination of inflation.[2]

In two recent studies, the IMF (2016; 2018) extend the analysis of Blanchard, Cerutti and Summers (2015) to 44 advanced and emerging market economies, covering the period 1990-2016. The results reported in these studies are again suggestive of the view that the parameters of the Phillips curve are relatively stable over time, particularly since the mid-1990s. However, the model residuals have increased, pointing to possible underestimation of economic slack in the model. The estimates also suggest that the coefficient on forward-looking inflation expectations increased steadily in advanced economies in the run up to the GFC, but fell subsequently to their early 1990s level, possibly reflecting the zero lower bound constraints on monetary policy in many economies.

In apparent contrast to the experience of advanced economies, there has been a significant degree of heterogeneity of inflation behavior documented for emerging market economies. According to the IMF (2016), while the median estimates for the slope of the Phillips curve in emerging market economies has been stable over time, the coefficient on forward-looking inflation expectations has fallen significantly since the GFC. The bulk of the recent inflation in these economies has been attributed to factors such as labour market conditions, import prices, and, to a lesser extent, changes in the exchange rate.[3] Consistent with this, the IMF (2018) reports that while anchoring of inflation expectations improved substantially in emerging market economies in the early 1990s, it remains weaker than that in advanced economies. Ha, Kose and Ohnsorge (2019) report similar evidence, while noting that inflation expectations in emerging market economies appear more sensitive to domestic and global shocks than advanced economies, possibly reflecting weaker monetary policy credibility in these economies.

[2] Other studies reporting evidence on the role of forward-looking inflation expectations are Williams (2009), Fuhrer (2012), IMF (2013), Coibion and Gorodnichenko (2015), Cecchetti et al. (2017).

[3] A number of studies have suggested that exchange rate passthrough has fallen over the past two decades in many emerging market economies – see, for instance, Choudhri and Hakura (2006), Mihaljek and Klau (2008), Jasova, Moessner and Takats (2016).

That said, several recent studies have also pointed to a possible misspecification bias in the Phillips curve estimates that either do not properly account for shifts in supply factors or fail to adequately reflect crisis-induced structural changes in the economy, including possible nonlinear relationship between inflation and economic activity. For example, Gordon (2013) demonstrates that the conventional backward-looking Phillips curve does a better job of predicting US inflation than a forward-looking Philips Curve when estimated with a short-term unemployment gap rather than standard unemployment gap and accounting for the shifts in relative price of food and energy as well as the trend growth of productivity. Stock and Watson (2010) suggest a new 'unemployment recession gap' measure (difference between current unemployment rate over the minimum unemployment rate over past 12 months) for improving predictability of the short term inflation, while Gilchrist et al. (2017) highlight the need to account for financial frictions, particularly the behavior of liquidity-constrained firms, in explaining the counter-cyclical behavior of prices after the GFC.

Turning to the second hypothesis, the role of global factors in the driving inflation behavior has been a subject of intense debate in recent years (see Borio and Filardo 2007, Guerrieri, Gust and Lopez-Salido 2010, Ihrig et al. 2010, Milani 2010, Forbes 2018, Auer, Borio and Filardo 2017). Studies have generally been inconclusive about the extent to which global factors matter and the channels through which they might affect inflation. One line of investigation looks at the source of inflation variance and whether it can be better explained by a model augmented with global factors than a standard closed-economy Phillips curve.[4] For instance, Ciccarelli and Mojon (2010) show that including a measure of global inflation (the average inflation of OECD economies) in the individual OECD economy-level inflation model not only helps to explain the variation in inflation better, but it also helps to sharply improve the forecasting power of the conventional backward-looking Phillips curve. This is consistent with common shocks leading to a high degree of co-movement in inflation across economies.

Mikolajun and Lodge (2016) explore the role of global factors in OECD economies in the context of an open economy Phillips curve, allowing for changes in oil prices and the exchange rate. Their results suggest that, while global inflation clearly plays a role, it is more important when inflation is high than when it is low (i.e., 1970s and 1980s compared to the 1990s and 2000s). The statistical significance of global inflation declines and the model's predictive power weakens once it is augmented with long-term inflation expectations. Kamber and Wong (2020) examine the same hypothesis by analyzing the impact of global factors on the trend and the cyclical component of inflation in the context of advanced and emerging market economies. They argue that monetary policy regime is likely to play a role and global factors tend to have more pronounced effect on trend inflation in economies without a credible monetary regime than in economies with a well-established inflation target.

Another line of enquiry explores the role of global factors having an independent effect on the slope of the Phillip's curve beyond their direct impact through channels such as common shocks. Forbes (2018) notes two key channels through which globalization can affect the slope of the Phillips curve. The first channel is greater foreign competition and outsourcing of production to foreign locations that have the potential effect of reducing the pricing power of firms, squeezing margins, and putting pressure on the mark up. At the same time, with increased trade integration, domestic costs are likely to move closely with foreign costs. Auer, Levchenko and Saure (2016) show how these effects have been strengthened by increased use of global value chains, leading to greater synchronization of producers prices. Guerrieri, Gust and Lopez-Salido (2010) provide a formal model to examine the response of inflation not only to real marginal costs but also to the ratio of import prices to domestic costs. The second channel is the increased integration of emerging market economies into the global economy, leading to

[4] See, for example, Mumtaz and Surico (2008), Monacelli and Sala (2007), Ciccarelli and Mojon (2010), Kamber and Wong (2020).

intensified competition in domestic labour markets. This has been hypothesized to reduce the bargaining power of workers, dampening the response of wages and prices to domestic economic slack (Auer and Fischer 2010).

In empirical models, a key variable often used to capture these effects is a measure global economic slack. Recent studies have reached very different conclusions about the impact of a global output gap on inflation and whether it has changed the slope of the Phillips curve. Among the studies reporting a positive impact of a global output gap on inflation are Borio and Filardo (2007), Bianchi and Civelli (2015), Auer, Borio and Filardo (2017). The most recent estimates are by Forbes (2018), who examines the impact global factors in the context of a Phillips curve regression using data for 43 advanced and emerging market economies covering the period 1990 to 2017. In pooled regressions with random effects, she shows that forward-looking inflation expectations have a significant effect on inflation (with a coefficient estimate of 0.67), while domestic and global output gaps have weak though positive effects on inflation (with coefficient estimates of 0.09 and 0.07, respectively). However, allowing the coefficients to vary over time, she reports that there has been a substantial increase in the coefficient of the global output gap, particularly after the GFC. In contrast, she finds that the impact of domestic output gap on inflation, which was positive before the GFC, is zero or negative in most years since 2008. Meanwhile, several recent papers (Ihrig et al. 2010, Milani 2010, Mikolajun and Lodge 2016) provide evidence to the contrary, noting that a global output gap has either very little effect or no effect at all on the slope of the Phillips curve, particularly when using data for the most recent decades.

3 Data

For our empirical analysis, we construct a balanced panel dataset for 47 advanced and emerging market economies that includes quarterly observations for inflation, inflation expectations, real GDP, exchange rates, and oil prices. As our benchmark case, we consider inflation based on headline CPI, expected future inflation based on professional forecasts, a domestic output gap based on the HP filter ($\lambda = 1,600$) applied to log real GDP for a given economy, a foreign output gap also based on the HP filter applied to log real GDP of other economies and trade weights, the nominal effective exchange rate index, and the WTI oil price index.[5] However, we also consider robustness of our results to an alternative approach of trend-cycle decomposition based on the BN filter with dynamic demeaning that Kamber, Morley and Wong (2018) show provides more reliable estimates of the output gap than the HP filter in real time.

Unlike some of the studies discussed in the previous section which focus on a global output gap, we define the foreign output gap for economy i as a weighted average of its trading partners' output gaps

$$\tilde{y}^*_{it} = \sum_{j=1}^{10} \omega^i_{jt}\tilde{y}_{jt} \tag{2}$$

where $\tilde{y}_{jt}$ is the output gap in economy j and ω_{jt} is the trade weight. For each economy pair i and j, we calculate the weight of economy j in economy i's foreign output gap as

$$\omega^i_{jt} = \frac{imports_{ij} + exports_{ij}}{imports_i + exports_i} \tag{3}$$

That is, trade weights are defined as the ratio of trade openness between economies i and j divided by the total trade openness of economy i. We calculate the foreign output gaps using

[5]Headline CPI and real GDP series were obtained from various national sources. Professional forecasts of inflation are from Consensus Economics and the BIS. The nominal effective exchange rates are from Bruegel Datasets. The WTI oil price index is from Datastream.

output gaps from 10 largest trading partners and update the weights annually, as in Borio and Filardo (2007).[6]

Expected future inflation is measured as one year ahead inflation expectations based on monthly surveys of professional forecasters.[7] Each month, Consensus Economics collects year-end inflation forecasts for the current and the following years. As these are fixed-event forecasts, the forecast horizon varies with the month in which the forecasts have been made. For example, while forecasts made in January are 12 and 24 month ahead forecasts, those in February are 11 and 23 month ahead forecasts, etc... In order to convert these fixed-event forecasts to fixed-horizon 12 month ahead forecasts, we follow the literature (see Yetman 2018) and average the forecasts for the current and next calendar years, weighted by their shares in the forecast period.

We have a balanced panel for all 47 economies from 1996Q1 to 2018Q3.[8] Based on BIS classification, we are considering 20 advanced economies (AEs) and 27 emerging market economies (EMEs). However, the 'emerging market' categorization is broad, as evidence by the fact that, based on IMF classification, the economies correspond to 30 AEs and 17 EMEs. Basically, there are three groups using standard two-letter codes:[9]

1. AT, BE, CH, DE, DK, ES, FI, FR, GB, IE, IT, NL, NO, PT, SE, AU, CA, JP, NZ, US
2. CZ, EE, GR, LT, LV, SK, SI, KR, HK, SG
3. HU, PL, RU, TR, ZA, CN, ID, IN, MY, PH, TH, AR, BR, CL, CO, MX, PE

The BIS classifies the economies in first group as AEs and the economies in second and third groups as EMEs, while the IMF classifies the economies in first and second groups as AEs and the economies in third group as EMEs. We consider robustness of any results that delineate AEs and EMEs to these two classifications.

4 Model specification and estimation

For each economy i, we consider an open economy hybrid Phillips curve regression model specification for inflation at time t:

$$\pi_{it} = \beta_{0i} + \beta_{1i}\pi_{it-1} + \beta_{2i}E_t^s\left[\pi_{it+4}\right] + \beta_{3i}\tilde{y}_{it} + \beta_{4i}\tilde{y}_{it}^{*} + \beta_{5i}\Delta_4 e_{it} + \beta_{6i}\Delta_4 p_{t-1}^{oil} + \varepsilon_{it} \tag{4}$$

where π_{it} is a quarterly measure of year-on-year inflation (the four-quarter change in 100 times the log consumer price index at time t), $E_t^s\left[\pi_{it+4}\right]$ is a measure of expected future inflation

[6]Trade weights were obtained from UN Comtrade and the BIS. We consider the largest trading partners that are also in our dataset.

[7]Our use of survey data allows us to avoid having to specify a full model of the economy that would be necessary to determine the rational expectation of future inflation in the New Keynesian Phillips curve. Although the survey forecasts may have some backward-looking element to their formation, the coefficient on expected future inflation should capture the impact of the forward-looking element of the survey forecasts given that we control for lagged inflation in our regression analysis. Meanwhile, consistent with the standard representation of the New Keynesian Phillips curve in equation (1), we focus on relatively short-term inflation expectations in determining current inflation. We note, however, that the survey measures we use are highly correlated with longer-term measures of expected future inflation from surveys when they are available for a given economy, but they allow us to consider a broader set of economies than would be possible for longer-term inflation expectations given data availability issues.

[8]It should be noted that inflation forecast surveys only became available for Estonia, Latvia, and Lithuania after 1998Q1. However, given the availability of survey data for all other economies in our sample from 1996Q1 and to keep a balanced panel, we use a weighted-average of current and lagged inflation to proxy for expected future inflation over the short sample period from 1996Q1 to 1998Q1 for these three economies.

[9]The listed order for each group is alphabetical within each region of Europe (includes South Africa), Asia-Pacific (includes Canada and the United States), and Latin America, respectively.

based on one year ahead survey forecasts (i.e., the expected four-quarter change in 100 times the log consumer price index at time $t+4$), $\tilde{y}_{it}$ is the domestic output gap (100 times detrended log real GDP), $\tilde{y}^*_{it}$ is the foreign output gap (100 times trade-weighted detrended log foreign real GDP), $\Delta_4 e_{it}$ is a quarterly measure of the year-on-year percent change in the exchange rate (the four-quarter change in 100 times the log of the nominal effective exchange rate index, where an increase corresponds to an appreciation of the domestic currency), $\Delta_4 p^{oil}_{t-1}$ is the lagged quarterly measure of the year-on-year percent change in oil prices (the four-quarter change in 100 times the log of the oil price index), and ε_{it} is a residual inflation shock that is possibly heteroskedastic with assumed scale-equivariant long-run variance $\bar{\sigma}^2_i$.[10]

In our econometric analysis, we follow much of the empirical literature and assume that the right-hand-side variables in equation (4) are either exogenous or at least predetermined in the sense that the inflation shock ε_{it} only affects them with a lag.[11] However, this assumption is debatable for some variables. In particular, it is at least plausible that inflation shocks could affect expected future inflation, the output gaps (although presumably domestic more so than foreign), or the exchange rate on impact. Such endogeneity would lead to biased and inconsistent results, at least under the assumption that the variables in equation (4) are stationary. Reassuringly, though, we have found that our results are generally robust, albeit not as statistically significant when considering lagged expected future inflation, output gaps, and exchange rate, which are predetermined by construction. Meanwhile, we include lagged oil prices in our main specification because we want to control for the possibility that inflation shocks for some large economies could feed into contemporaneous changes in oil prices. However, we have also found that our results are robust to considering contemporaneous oil prices.

To reduce the number of independent parameters that we need to estimate, which is particularly helpful when allowing for multiple structural breaks at unknown breakdates, we also consider a restricted version of equation (4) that imposes $\beta_{1i} + \beta_{2i} = 1$. This restriction gives the parameters β_{1i} and β_{2i} direct interpretation as respective weights on backward-looking and forward-looking expectations like with the hybrid New Keynesian Phillips curve in equation (1). Estimates for the restricted specification are obtained using the following regression equation:

$$\pi_{it} - E^s_t\left[\pi_{it+4}\right] = \beta_{0i} + \beta_{1i}\left(\pi_{it-1} - E^s_t\left[\pi_{it+4}\right]\right) + \beta_{3i}\tilde{y}_{it} + \beta_{4i}\tilde{y}^*_{it} + \beta_{5i}\Delta_4 e_{it} + \beta_{6i}\Delta_4 p^{oil}_{t-1} + \varepsilon_{it} \quad (5)$$

where inferences about β_{2i} can be made indirectly using the restriction $\beta_{2i} = 1 - \beta_{1i}$. Notably, the variables in this regression equation are all plausibly stationary (other than possible infrequent structural breaks in long-run mean and/or variance that we allow for in some of our estimation), even if inflation and expected future inflation are I(1) but cointegrated with vector $[1 - 1]'$. Stationarity is important for the validity of most structural break analysis with unknown breakdates (e.g., Bai and Perron 1998; 2003, Qu and Perron 2007).

We begin our analysis by considering panel estimation of equations (4) and (5) assuming slope coefficients are the same across all economies in the panel (i.e., $\beta_{ji} = \beta_j$ for $j = 1, ..., 6$), but allowing for heterogeneity in the intercepts (i.e., β_{0i} is unrestricted).[12] As part of our analysis,

[10] Technically, any serial correlation in the residuals would lead to biased and inconsistent estimates given the presence of a lagged dependent variable in equation (4). However, allowing for serial correlation in estimation means that inferences about pseudo-true values of parameters (i.e., estimation values in the limit) will be more accurate if any serial correlation is actually present. In practice, given the inclusion of a lagged dependent variable, there is little remaining serial correlation in the residuals, but for robustness we always allow for both heteroskedasticity and serial correlation when making econometric inferences about parameters.

[11] In the case of strict exogeneity, equation (4) can be thought of as an autoregressive distributed lag model, where the long-run effects of shocks to the other variables would be equal to the short-run effects multiplied by $1/(1 - \beta_{1i})$.

[12] Slope heterogeneity can be addressed by using fixed or random effects. We follow Forbes (2018) in using random effects, which corresponds to the assumption that economy-specific variation in β_{0i} is uncorrelated with the explanatory variables.

we allow the slope coefficients to vary over time (i.e., β_{jt} for $j = 1, ..., 6$) and, motivated by Forbes (2018), we consider how they have changed with the onset of the GFC by interacting the explanatory variables with a dummy variable $D_t = 1$ for $t \geq 2007Q1$ and 0 otherwise.

In addition to the panel estimation, we also consider economy-by-economy least squares estimation of equation (5) to allow for heterogeneous slope coefficients, as well as heterogeneous intercepts, across economies. Despite a potential lack of precision in economy-by-economy estimation, heterogeneity is checked by formal tests of coefficient values to determine if they are significantly different than panel estimates (i.e., t tests of $H_0 : \beta_{ji} = \hat{\beta}_j$). Meanwhile, a benefit of the economy-by-economy approach is that it easily allows consideration of potential multiple structural breaks in regression coefficients and/or error variance at unknown breakdates according to Qu and Perron (2007) procedures.[13] That is, we allow the model parameters to vary over time (i.e., β_{jit} and $\bar{\sigma}_{it}$).

As is standard when estimating structural breaks at unknown breakdates, we use trimming that restricts the minimum length between breaks to be at least 15% of the total sample period in order to determine critical values for tests with reasonable power. The adjusted estimation sample period is 1996Q2-2018Q3, meaning that the first possible breakdate is 1999Q3 and the last possible breakdate is 2015Q1. Note that we follow the convention in the structural break literature that the parameters change to new values in the period after the breakdate. Tests and estimation of structural breaks allow for heteroskedasticity and serial correlation in the residuals.

Breakdates are estimated sequentially and breaks are allowed for in regression coefficients only if the change across regimes is significant at the 5% level based on a likelihood ratio test, although changes in error variance are always allowed with each significant break. In principle, we allow up to 5 breaks across the sample period. However, given the short subsample periods with sequential testing for breaks, we are often restricted to fewer breaks in practice. For our benchmark specification, we allow for only 2.7 breaks on average, while finding 1.7 breaks on average, with the estimated total number of breaks equal to the maximum number for 51% of the economies in the sample. However, we test whether breaks apply to specific parameters and generally find fewer breaks for slope coefficients than the total, so the maximum allowable breaks is almost never binding for the estimated number of breaks in slope coefficients.

We record the estimated parameters at each point of time for each economy based on estimated breakdates and report the mean, median, and quartiles of the distribution of estimates across economies. Because these estimates treat the breakdates as known, we also record weighted averages of estimated changes in parameters at each point of time based on 95% confidence sets constructed via inverted likelihood ratio tests (see Eo and Morley 2015), with weights proportional to the precision of the confidence sets in order to address uncertainty about the timing of breaks. Specifically, the weighted-average changes consider estimated changes (including zero) in each period, dividing an estimated change that could have occurred in a given period by the the number of periods in a confidence set. To address confidence sets that are contiguous with the beginning or end of the trimmed sample period, the estimated effect in these cases is divided by the number of trimmed periods at the beginning or end (i.e., 15% of the length of the total sample period).

[13] The methods in Qu and Perron (2007) are valid in a general multi-equation regression model setting and so could potentially be applied in panel estimation. However, there are practical challenges with estimating a large panel model with heterogeneous slopes and potentially different structural breaks in parameters for different economies. Thus, we consider economy-by-economy estimation for simplicity.

Table 1: Panel estimates

	(a) HP filter gaps				(b) BN filter gaps			
Variable	Unrestricted	Restricted	Unrestricted	Restricted	Unrestricted	Restricted	Unrestricted	Restricted
π_{it-1}	0.584***	0.583***	0.588***	0.580***	0.587***	0.586***	0.587***	0.578***
	(0.045)	(0.044)	(0.051)	(0.049)	(0.048)	(0.046)	(0.055)	(0.053)
$E_t^s[\pi_{it+4}]$	0.414***	0.417***	0.391***	0.420***	0.413***	0.414***	0.395***	0.422***
	(0.056)	(0.044)	(0.058)	(0.049)	(0.059)	(0.046)	(0.064)	(0.053)
$\tilde{y}_{it}$	0.074**	0.075**	0.138**	0.136**	0.054	0.055	0.187*	0.183
	(0.035)	(0.037)	(0.068)	(0.067)	(0.054)	(0.057)	(0.111)	(0.113)
$\tilde{y}_{it}^*$	0.099**	0.097**	-0.118**	-0.082	0.141**	0.140**	0.034	-0.020
	(0.041)	(0.045)	(0.058)	(0.061)	(0.058)	(0.060)	(0.116)	(0.125)
$\Delta_4 e_{it}$	-0.085***	-0.085***	-0.109***	-0.098***	-0.084***	-0.084***	-0.110***	-0.098***
	(0.024)	(0.023)	(0.030)	(0.027)	(0.025)	(0.024)	(0.032)	(0.029)
$\Delta_4 p_{t-1}^{oil}$	0.005***	0.005***	0.005***	0.003	0.005***	0.005***	0.003	0.001
	(0.001)	(0.001)	(0.002)	(0.002)	(0.001)	(0.001)	(0.002)	(0.003)
$D_t \cdot \pi_{it-1}$			-0.001	0.022			0.033	0.048
			(0.040)	(0.045)			(0.043)	(0.055)
$D_t \cdot E_t^s[\pi_{it+4}]$			0.102***	-0.022			0.090**	-0.048
			(0.036)	(0.045)			(0.036)	(0.055)
$D_t \cdot \tilde{y}_{it}$			-0.118	-0.113			-0.213*	-0.215*
			(0.074)	(0.078)			(0.113)	(0.117)
$D_t \cdot \tilde{y}_{it}^*$			0.257***	0.260***			0.243**	0.275**
			(0.069)	(0.072)			(0.119)	(0.133)
$D_t \cdot \Delta_4 e_{it}$			0.071**	0.047			0.074**	0.048
			(0.029)	(0.029)			(0.031)	(0.032)
$D_t \cdot \Delta_4 p_{t-1}^{oil}$			0.001	0.003			0.003	0.006**
			(0.002)	(0.002)			(0.003)	(0.003)
Constant	-0.126*	-0.136***	-0.211**	-0.124***	-0.130*	-0.133***	-0.231**	-0.097***
	(0.070)	(0.020)	(0.089)	(0.014)	(0.071)	(0.021)	(0.104)	(0.010)
Random Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	4230	4230	4230	4230	4230	4230	4230	4230
Adjusted R^2	0.949	0.709	0.952	0.718	0.948	0.704	0.952	0.715

Notes: 'Unrestricted' refers to the model in equation (4) and 'Restricted' refers to the model in equation (5). The adjusted R^2s are not comparable across the restricted and unrestricted models given different dependent variables. For panel estimation, we assume $\beta_{ji} = \beta_j \ \forall j$, but we allow for economy-specific random effects. For the restricted model, estimates for both β_1 and β_2 are reported using $\beta_2 = 1 - \beta_1$. The dummy variable $D_t = 1$ for $t \geq 2007Q1$ and 0 otherwise allows for a structural break in model parameters β_{jt} for $j = 1, ..., 6$ with the onset of the GFC and changes in parameters are reported. Robust standard errors with clustering by economy reported in parentheses and significance is based on two-tailed t tests (***significant at 1%, **significant at 5%, *significant at 10%).

5 Empirical results

5.1 Panel estimates

The panel estimates are reported in Table 1. Despite some differences in the data and model specification, the estimates are generally in line with those from a similar panel analysis in Forbes (2018).[14] All of the variables in the model appear to be significant driving forces of inflation. There is a higher estimated weight on lagged inflation than expected future inflation, although this result was somewhat ambiguous in Forbes (2018) depending on the measure of inflation considered. In our unrestricted estimation, the estimated weight on expected future

[14] Forbes (2018) considers a broadly similar, but not identical, set of economies and sample period. In her baseline specification, she also considers quarterly inflation, both headline and core, rather than year-on-year inflation for the dependent variable, although she considers lagged year-on-year inflation, as we do, to measure backward-looking expectations. She uses a five-year-ahead survey forecast to measure forward-looking inflation expectations, different measures of domestic and foreign output gaps, the percent change in the real effective exchange rate over a two-year horizon, and lagged percent change in world oil prices over a one-quarter horizon. She also includes commodity prices and a measure of world producer price dispersion in her baseline model.

inflation increased after the onset of the GFC. In our restricted estimation, the estimated slope of domestic Phillips curve flattened, although this is not significant, which was also the case for headline CPI in Forbes (2018). In both cases, the estimated slope of the foreign Phillips curve increased, while the estimated degree of exchange rate passthrough decreased, significantly so in the unrestricted case, while the decrease was not significant in Forbes (2018).

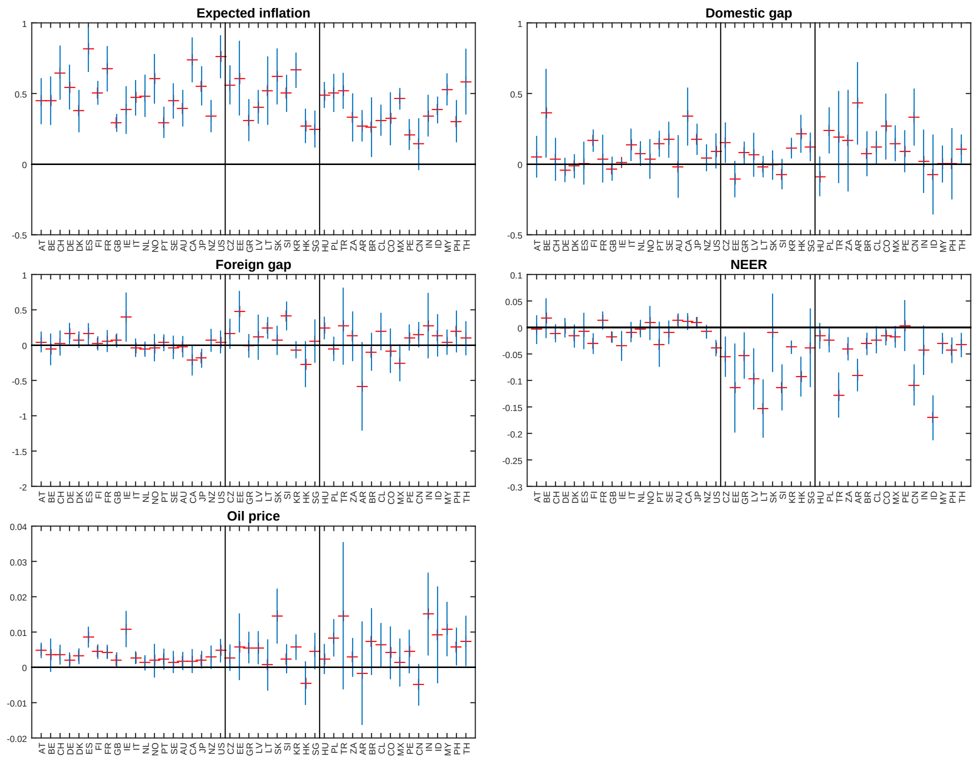
Our unrestricted and restricted estimates are generally quite similar to each other, supporting the imposition of the theoretically-motivated restriction that the coefficients on lagged inflation and expected future inflation correspond to weights that sum to one. Not surprisingly, imposing the restriction leads to more precise estimates of the effects of lagged/expected future inflation, although it is more mixed about the precision of the estimates for the other variables. The major difference is that, when allowing for a structural break, the restricted estimates do not suggest a significant change in the coefficient on expected future inflation. The implication is that the data contain more precise information about the effect of lagged inflation than expected future inflation, so the restricted estimates of the possible change in weights on lagged and expected future inflation are dominated by this information. We note that the adjusted R^2s are not comparable across the restricted and unrestricted models given different dependent variables in equations (4) and (5). However, we argue that the general similarity of the estimates supports imposing the restriction, which we take as our benchmark specification in subsequent analysis.

Another result from the panel estimation that we highlight is that the quantitative estimates are generally robust to different approaches for measuring output gaps, especially in the no-break case. Lagged inflation has relatively more weight than expected future inflation, with estimates close to a 60/40% split for the no-break case and the pre-GFC sample period when allowing for a structural break. The estimated slope of the foreign Phillips curve is steeper than the domestic Phillips curve for the no-break case and the foreign Phillips curve steepens while the domestic Phillips curve flattens (though not significantly so using the HP filter gaps) when allowing for a structural break. Estimates for exchange rate passthrough and the impact of oil prices generally suggest economically large and quantitatively similar effects for different measures of output gaps, with the degree of exchange rate passthrough decreasing significantly according to the unrestricted estimates when allowing for a structural break.[15] In this case, the adjusted R^2s are comparable across the different approaches to measuring output gaps. Although the model fits are quite similar and some of the estimates seem more reasonable for the BN filter (e.g., the pre-GFC estimated slope of the foreign Phillips curve), we use the HP filter for our benchmark specification in subsequent analysis given slightly higher R^2s and better comparability to other studies that have also used the HP filter. However, we check the sensitivity of results in our subsequent analysis to consideration of the more reliable BN filter and find that the main results are largely robust, suggesting that real-time issues are less important for examining historical driving forces of inflation than would likely be the case for understanding current inflation pressures.[16]

[15] Given the normalization of the nominal effective exchange rate index such that an increase in it corresponds to an appreciation of the domestic currency, we note that the initial negative estimate for β_5 and the estimated increase in the parameter to a less negative value with the structural break corresponds to a decline in the degree of passthrough. We also note that our consideration of nominal effective exchange rates implicitly assumes passthrough is dominated by movements in nominal exchange rates with major trading partners rather than foreign price levels or necessarily the bilateral US dollar exchange rate in particular, although see Gopinath et al. (2020) on the importance of the latter in driving global trade prices.

[16] We note that this robustness is also consistent with the idea that economic agents set prices based on many signals that capture economic slack rather than the real-time HP filter estimate of the output gap per se.

Figure 1: Full-sample economy-by-economy estimates and 95% confidence intervals



Notes: Results are reported for expected future inflation, the domestic output gap, the foreign output gap, the nominal effective exchange rate, and oil prices, with two-letter economy codes listed on the x-axis (RU excluded given much larger scale for confidence intervals). Point estimates are red crosses and confidence intervals are blue lines. Confidence intervals are based on inverted t tests using HAC standard errors calculated according to Andrews and Monahan (1992).

5.2 Economy-by-economy estimates

A key issue with the panel estimates in Table 1 is that they do not allow for heterogeneity in the slope coefficients or in the existence and timing of structural changes in the slope coefficients. We find that allowing for these forms of heterogeneity is critically important for inferences about the effects of different potential driving forces of inflation over time.

5.2.1 Coefficient heterogeneity

Figure 1 reports the full-sample economy-by-economy point estimates and 95% confidence intervals for the slope coefficients in equation (5).[17] For each coefficient, the estimates are sorted by the three different groups of economies, with the the left/right of the first line corresponding to AEs/EMEs according to the BIS classification and the left/right of the second line corresponding to AEs/EMEs according to the IMF classification. Given these classifications, it is easy to see that the estimates generally imply larger (albeit more heterogeneous and less precisely estimated) effects of the variables for EMEs than AEs, except in the case of expected future inflation for which the weights appear somewhat higher on average for the AEs. Interestingly, the economies in the middle group sometimes look more like those always classified as AEs in

[17]For the economy-by-economy analysis, we always consider the restricted case that imposes that the coefficients on lagged inflation and expected future inflation are weights that sum to one. Thus, for simplicity, we do not report results for the coefficient on lagged inflation, which are directly implied by the results for the coefficient on expected future inflation.

Table 2: Rejection rates for parameter tests at 5% level

	Hypothesis		
Variable	$H_0 : \beta_{ji} = 0$	$H_0 : \beta_{jit} = 0, \forall t$	$H_0 : \beta_{ji} = \hat{\beta}_j$
$E_t^s [\pi_{it+4}]$	98%	98%	26%
$\tilde{y}_{it}$	40%	66%	21%
$\tilde{y}_{it}^*$	17%	66%	15%
$\Delta_4 e_{it}$	49%	74%	74%
$\Delta_4 p_{t-1}^{oil}$	43%	72%	21%

Notes: For $H_0 : \beta_{ji} = 0$ and $H_0 : \beta_{ji} = \hat{\beta}_j$, we consider two-tailed t tests using HAC standard errors calculated according to Andrews and Monahan (1992). For $H_0 : \beta_{jit} = 0$, $\forall t$, we reject if we can reject either the t test for $H_0 : \beta_{ji} = 0$ or a likelihood ratio test for a structural break in a given parameter (i.e., $H_0 : \beta_{jit} = \beta_{ji}$) based on Qu and Perron (2007) procedures.

terms of high weights on expected future inflation and relatively flat domestic Phillips curves and other times look more like those always classified as EMEs in terms of a relatively steep foreign Phillips curves, more exchange rate passthrough, and larger impact of oil prices.

Although there is a lot of apparent heterogeneity in the estimates in Figure 1, it is notable that the estimated weight on expected future inflation (and, therefore, lagged inflation) always lies between 0 and 1, the estimated slope of the domestic Phillips curve is never significantly negative, the estimated slope of the foreign Phillips curve is only significantly negative in 1 out of 47 cases, the estimated coefficient on the exchange rate is only significantly positive in 1 out of 47 cases, while the estimated impact of oil prices is never significantly negative. Thus, despite being less precise than the panel estimates, there is a sense in which the economy-by-economy estimates are consistent with what would be expected for an open economy hybrid Phillips curve model, at least when taking sampling error into account. Furthermore, Table 2 suggests that all of the variables are useful to include when capturing inflation for a significant fraction of the economies in the sample. In particular, we can reject a t test for $H_0 : \beta_{ji} = 0$ at a higher rate than the 5% level of the test for all of the parameters, ranging from 17% for the foreign output gap to 98% for expected future inflation. Notably, we can reject no role for the exchange rate for about half of the economies in the sample. As discussed in more detail below, the support for considering all of these variables to explain inflation is even stronger when allowing for structural breaks.

A reasonable question to ask is whether it is worthwhile giving up on the precision of the panel estimates to allow for heterogeneity in slope coefficients. The general answer from our analysis is "Yes". Table 3 reports the average effects implied by the panel estimates versus economy-by-economy estimates. Focusing on the 'no break' case for now, we can see that average effects implied by panel and economy-by-economy estimation are quite similar for oil prices, although less so for expected future inflation, the output gaps, or exchange rates. The larger differences presumably result from unequal weighting of economies in the panel estimation given random effects. However, this result clearly illustrates a potential sensitivity of estimates to which economies are included in a panel given the presence of underlying heterogeneous effects.

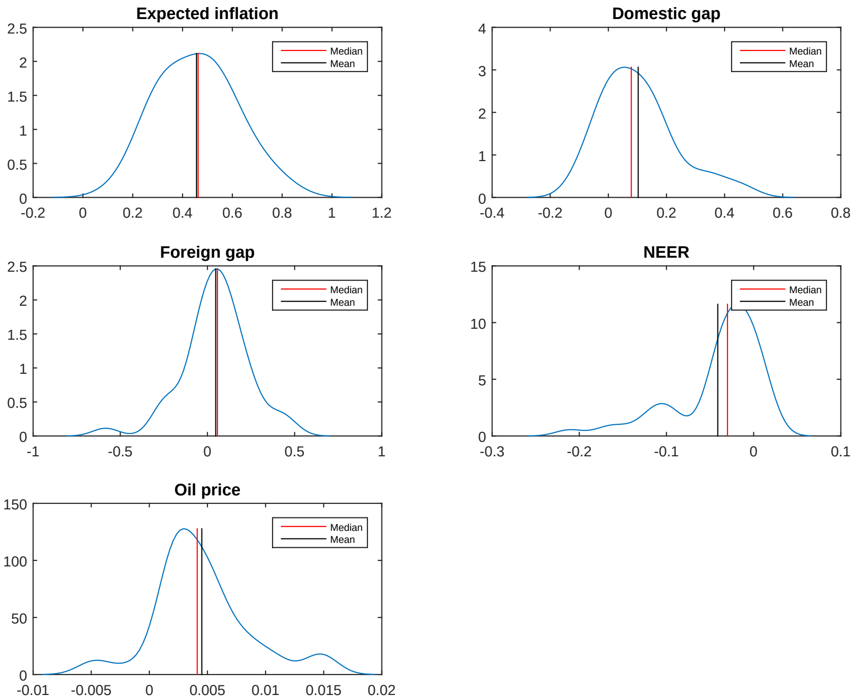
Figure 2 reports the distribution of the full-sample economy-by-economy point estimates, with the mean and median estimates noted. The mean, median, and modal estimates are close for expected future inflation and the foreign output gap, but there are economically meaningful differences for the domestic output gap, exchange rate, and oil prices. For the domestic output gap, the average estimated effect of 0.103 is clearly influenced by some large positive outlier estimates, with the median estimate lower and the modal estimate only about half the mean

Table 3: Average estimated effects

Variable	Panel		Economy-by-economy	
	No Break	Break	No Break	Breaks
$E_t^s\left[\pi_{it+4}\right]$	0.417	0.408	0.456	0.528
$\tilde{y}_{it}$	0.075	0.077	0.103	0.068
$\tilde{y}_{it}^*$	0.097	0.054	0.047	0.076
$\Delta_4 e_{it}$	-0.085	-0.073	-0.041	-0.036
$\Delta_4 p_{t-1}^{oil}$	0.005	0.004	0.005	0.004

Notes: The average estimated effects are $\frac{1}{T}\sum_t \hat{\beta}_{jt}$ for the panel estimation and $\frac{1}{NT}\sum_i \sum_t \hat{\beta}_{jit}$ for the economy-by-economy estimation.

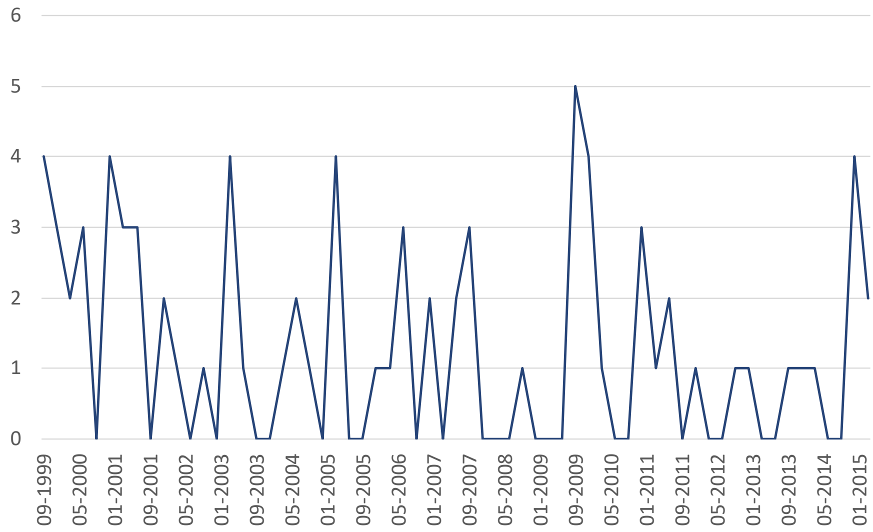
Figure 2: Distributions for full-sample estimates across different economies



Notes: Smoothed distributions are calculated using the default nonparametric smoothing kernel in Matlab. The black line corresponds to the mean estimate and red line corresponds to the median estimate.

estimate. This suggests that removal of a few economies with very high estimates evident in Figure 1 from the panel could have a big impact on inferences about the importance of the domestic output gap and could possibly explain the differences in the panel estimate and the mean economy-by-economy estimate if these economies receive less weight in estimation due to random effects. For the exchange rate, the negative mean estimate of -0.041 that corresponds to a reasonably substantial degree of passthrough is also clearly influenced by some large negative outlier estimates, also visible in Figure 1. In this case, the panel estimate implies an even higher degree of passthrough, so these economies are presumably getting more weight in estimation due to random effects. However, Figure 2 makes it clear that median

Figure 3: Estimated number of breaks across different economies in each quarter



Notes: Structural breaks are tested for and breakdates estimated based on Qu and Perron (2007) procedures. 15% trimming means that breaks are only allowed to be estimated between 1999Q3 and 2015Q1.

estimated degree of passthrough is smaller and the modal estimated degree of passthrough is close to zero. For oil prices, the mean and median effects are reasonably close, but the modal effect is only half as large. Yet, it is clear from Figure 2 that there is a lot of heterogeneity in the impact of oil prices on inflation and, indeed, in the effects of all of the possible inflation drivers.

Confirming this heterogeneity, we find that we can reject a t test for $H_0 : \beta_{ji} = \hat{\beta}_j$ (i.e., that a given estimated effect for each economy is the same as the panel estimate) far more than the 5% level of the test. In particular, returning to Table 2, we can reject 26% of the time for expected future inflation, 21% of the time for the domestic output gap, 15% of the time for the foreign output gap, 74% of the time for exchange rate passthrough, and 21% of the time for oil prices. Notably, the rejection rate for homogeneous exchange rate passthrough is higher than the rate for rejecting that there is no passthrough, strongly suggesting that the panel estimate provides a misleadingly large estimate of the degree of passthrough for most of the economies under consideration.

5.2.2 Heterogeneity in the existence and timing of structural breaks

Given these results supporting heterogeneous slope coefficients, we continue our economy-by-economy analysis to consider possible heterogeneous structural changes in model parameters over the full sample period.

Figure 3 reports the estimated number of structural breaks across different economies in each quarter. Although the largest estimated number of breaks occurs in 2009, there are many breaks that occur at other times before and after the GFC too, especially near the beginning of the trimmed sample period. Meanwhile, Table 4 makes it clear that the estimated number of breaks is heterogeneous across economies and also varies in terms of which coefficients undergo structural change. In particular, the median estimated number of breaks for a given economy is 2, but there are many economies for which there is only 1 break. Furthermore, the breaks are often in terms of the volatility of the inflation shock, not changes in slope coefficients. The median estimated number of breaks for each slope coefficient is 0 in every case but the foreign

Table 4: Distributions for estimated number of breaks across different economies

Variable	Mean	25^{th} Percentile	Median	75^{th} Percentile
$E_t^s[\pi_{it+4}]$	0.47	0	0	1
$\tilde{y}_{it}$	0.62	0	0	1
$\tilde{y}_{it}^*$	0.66	0	1	1
$\Delta_4 e_{it}$	0.64	0	0	1
$\Delta_4 p_{t-1}^{oil}$	0.64	0	0	1
Total no. of breaks (incl. in $\bar{\sigma}_i$)	1.70	1	2	2
Maximum allowed	2.74	2	3	4

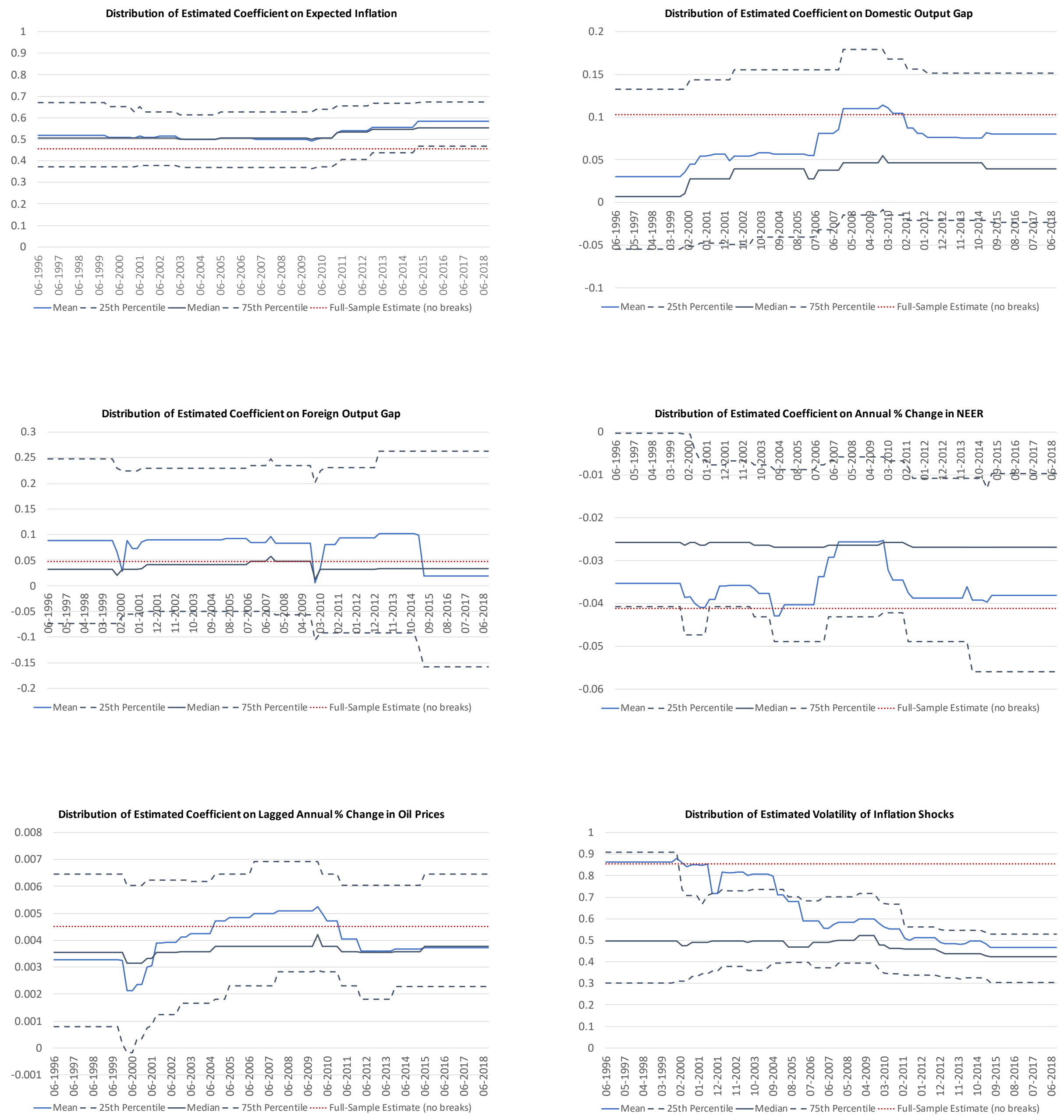
Note: The presence of structural breaks and whether they apply to a given slope coefficient is tested for based on Qu and Perron (2007) procedures.

output gap, for which it is 1, although there are also many economies for which there is at least 1 break in a given slope coefficient. So the notable takeaway is that, while there are breaks for some economies, there is also a remarkable stability in the effects of the various possible inflation drivers across the full sample period for many of the economies.

Despite the apparent stability of many of the slope coefficients for many economies, it is clear that it is important to account for breaks in making inferences across the full set of economies. Returning to Table 3, we can see that allowing for breaks affects inferences about average effects over the full sample period. In particular, it leads to more weight on average for expected future inflation than lagged inflation. Notably, it also leads to a smaller average estimated effect of the domestic output gap than the foreign output gap. The estimated degree of passthrough and impact of oil prices are slightly smaller when allowing for breaks. So one implication is that expected future inflation and the foreign output gap are more important than would be suggested by ignoring any structural change or, looking back at the panel estimates in Table 1, from assuming only a simple one-time structural change in all parameters during the GFC. Meanwhile, even though some of the average estimated effects are smaller when allowing for breaks (e.g., the domestic output gap, the exchange rate, and oil prices), allowing for breaks reveals that the various possible inflation drivers are important for a much higher proportion of economies than when not allowing for breaks. Returning to Table 2, we are able to reject the hypothesis $H_0 : \beta_{jit} = 0, \forall t$ much more often than $H_0 : \beta_{ji} = 0$ (under the assumption of no breaks) for the output gaps, the exchange rate, and oil prices. In particular, the existence of structural breaks means that the effects of the possible inflation drivers are not always equal to zero across the full sample period, so a composite test of $H_0 : \beta_{ji} = 0$ and whether there are structural breaks in a slope coefficient is more powerful than the basic test of variable relevance when not allowing for breaks.

Allowing for heterogenous slope coefficients and heterogeneous structural change suggests that an open economy hybrid Phillips curve model of inflation is quite useful for a large number of economies, with stable slope coefficients in many cases. Of course, allowing for breaks also provides inferences about how the parameters which have changed have actually done so. Figure 4 reports how estimated slope coefficients have evolved over time. In particular, the panels plot at each point of time the mean, median, and quartiles of the distribution of estimates across economies for a given parameter. Starting with the weight on expected future inflation, we can see that the estimates are reasonably dispersed, but centered around almost exactly 0.5 at the beginning of the sample period and through until 2010, when the whole distribution shifts up over the next 4-5 years, with the mean reaching close to 0.6. The slope of the domestic Phillips curve also has a fair amount of dispersion, with the mean and especially the median starting close to zero and gradually increasing over the 2000s. The mean increases

Figure 4: Distributions for parameter estimates over time



Notes: Estimated parameters are conditioned on estimated breakdates and whether a likelihood ratio test suggests that a structural break applies to a given slope coefficient based on Qu and Perron (2007) procedures. Volatility is measured in standard deviation terms.

a lot in 2006-2007, but falls back part way in 2010-2011. The difference between mean and median suggests the presence of some outlier economies that saw large changes in the slope of the domestic Phillips curve during and after the GFC. The slope of the foreign Phillips curve shows a high degree of dispersion across economies and displays no clear pattern of change, although the mean is notably almost always higher than the median suggesting some outlier economies that have particularly high sensitivity to the foreign output gap. The coefficient on the exchange rate shows little change in terms of the median, but there is huge dispersion and the mean effect corresponds to more passthrough for most of the sample period, although

Table 5: Average estimated effects using BN filter output gaps

	Panel		Economy-by-economy	
Variable	No Break	Break	No Break	Breaks
$E_t^s[\pi_{it+4}]$	0.414	0.396	0.427	0.539
$\tilde{y}_{it}$	0.055	0.071	0.059	0.058
$\tilde{y}_{it}^*$	0.140	0.123	0.121	0.107
$\Delta_4 e_{it}$	-0.084	-0.073	-0.037	-0.038
$\Delta_4 p_{t-1}^{oil}$	0.005	0.004	0.004	0.004

Notes: The average estimated effects are $\frac{1}{T}\sum_t \hat{\beta}_{jt}$ for the panel estimation and $\frac{1}{NT}\sum_i \sum_t \hat{\beta}_{jit}$ for the economy-by-economy estimation.

it briefly converges to the median level during the GFC. The coefficient on oil prices shows some pattern of increase in the 2000s, especially in the mean and the quartiles, but less so the median, with the mean reverting back to the lower median after 2010.

Figure 4 also reports on the distribution of the estimates of the volatility of inflation shocks, measured in standard deviation terms, in the bottom-right panel. It is clear that the typical size of inflation shocks has fallen across the board from a mean level close to the full-sample estimate at the beginning of the sample period to about half that by the end of the sample period. Median volatility is lower and more stable, suggesting the decline in the size of shocks is a phenomenon for economies that started with a high degree of volatility. This could reflect the adoption of inflation targeting by many EMEs during this particular sample period. It should also be noted econometrically that the more flexible models which allow for structural breaks will inherently imply smaller shock volatility on average. However, it is worth recalling that the structural breaks are statistically significant, suggesting that the implied lower average estimate of volatility is not just a statistical artefact, but reflects an actual economic phenomenon.

The inferences in Figure 4 take the estimated breakdates as given. However, in some cases there is considerable uncertainty about the exact timing of a structural break for a given parameter. To address this, Figure 5 reports the weighted averages of estimated changes in parameters at each point of time based on 95% confidence sets constructed via inverted likelihood ratio tests (see Eo and Morley 2015), with weights proportional to the precision of the confidence sets. This approach can reveal if structural changes have actually been more gradual than the discrete estimated changes implied by Figure 4. The first thing to notice in Figure 5 is that the weighted-average estimated changes are generally quite small relative to the average estimated effects. This finding confirms the idea that the open economy hybrid Phillips curve model provides a reasonably stable structure for understanding inflation determination from one quarter to the next. The second thing to notice is that there are distinct periods in which parameters are estimated to be effectively unchanged and periods in which they change for at least some economies. Comparing back to Figure 4, there are a few noticeable differences from taking uncertainty about timing into account. First, there is little certainty about the exact timing of the apparent average flattening of the foreign Phillips curve, which may have occurred in the mid-2000s or from 2012 on or both. Second, there is also relatively little certainty about the timing or speed of the apparent decrease in the average degree of exchange rate passthrough and its subsequent reversal. Third, there is actually relatively more certainty that much of the apparent decrease in the average volatility of the inflation shock occurred prior to the GFC.

Figure 5: Weighted-average changes in parameter estimates over time

Notes: Estimated changes are averaged relative to the precision of the confidence set for the timing of a breakdate based on Eo and Morley (2015). If a confidence set is contiguous with the beginning or end of the trimmed sample period, the weighted-average effect is calculated by dividing by the number of trimmed periods at the beginning or end (i.e., 15% of the length of the total sample period). Volatility is measured in standard deviation terms.

5.3 Robustness and results for different groups of economies

The results in the previous subsection are generally robust to consideration of BN filter estimates of output gaps instead of HP filter estimates. Table 5 reports average effects for panel and economy-by-economy estimates when using BN filter output gaps. As with Table 3 based on HP filter output gaps, we can again see that accounting for coefficient and break heterogeneity reveals that, on average over the whole sample period, expected future inflation has relatively more weight than lagged inflation, the foreign output gap is more important than the domestic output gap, and exchange rate passthrough is less important than implied by panel estimates. Meanwhile, inferences about the impact of oil prices are quite robust across all specifications and the different measures of output gaps.

Some of the patterns and disparate findings in the previous subsection in terms of how parameters have changed over time could be due to our consideration of a very heterogeneous set of economies. Just as this heterogeneity argues against considering panel analysis with homogenous slope coefficients to estimate average effects across the whole set of economies, it also suggests that it makes sense to look at patterns of changes for different groups of economies rather than for the whole set of economies under consideration.

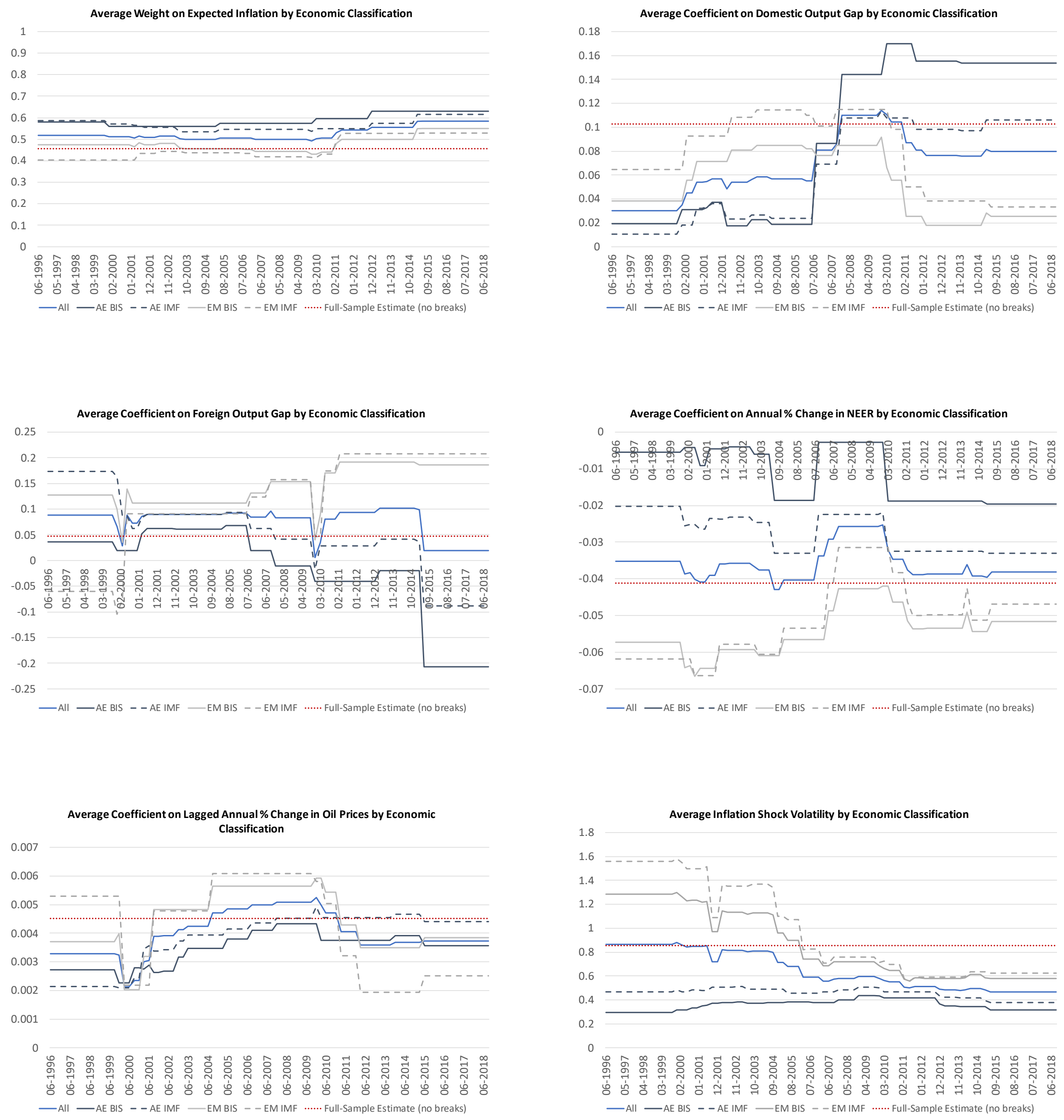
Figure 6 reports average effects over time by economy groups based on BIS and IMF classifications using the HP filter output gaps (again, we found similar results when using BN filter output gaps, which are not reported for brevity). The increased weight on expected future inflation from 2010 on is apparent for both AEs and EMEs. AEs have an increase in the slope of domestic Phillips curve during the GFC, especially for the BIS classification given in Section 3, while it sharply flattens for EMEs after the GFC. AEs have a decrease in the slope of the foreign Phillips curve from the GFC on, while it steepens for EMEs after the GFC. The degree of exchange rate passthrough is larger but decreases during the GFC for EMEs, while it is small and changes little for AEs. The impact of oil prices increases for both AEs and EMEs throughout the 2000s, but reverses somewhat for both groups afterwards. The volatility of the inflation shock falls dramatically for EMEs throughout the 2000s and levels off from 2011 on, while it is low and stable for AEs, however defined.

Another useful delineation of economies is by region. Figure 7 reports average effects over time for Asian EMEs and Latin America using the HP filter output gaps (again, results are similar for BN filter output gaps and not reported for brevity). For Asian EMEs, the shift in weight to expected future inflation is particularly notable, going from just below 0.40 to 0.65 by the end of the sample period. By contrast, there appears to have been no such shift for Latin American economies. Asian EMEs show an increase in the slope of the domestic Phillips curve, while Latin American economies show a decrease from high levels in the early 2000s. With the foreign Phillips curve, the main change is an increase in the slope for Latin American economies from a negative slope in the late 1990s that may have reflected the financial market contagion with the Peso and Asian crises delinking the fortune of Latin American economies from their major trading partners. The degree of exchange rate passthrough decreased for Asian EMEs, especially for the IMF classification that excludes Korea, Hong Kong, and Singapore with already lower exchange rate passthrough than other Asian economies, while it increased for Latin America.[18] The impact of oil prices is similar for Asian EMEs as for the whole set of economies, especially for the BIS classification, while it declines dramatically from a very high initial level for Latin America. The volatility of the inflation shock declines for both regions from the beginning to the end of the sample period, but with a big jump up in the early 2000s for Latin America, possibly related to the economic crisis in Argentina at the time, and with some temporary increase during the GFC for Asia.

Because the patterns of structural change are so heterogeneous and reasonably complicated, it is also useful to consider average effects calculated pre-GFC and afterwards in order to compare to the results in Forbes (2018). Table 6 reports average effects by different groups for the subsample periods of 1996Q2-2006Q4 and 2007Q1-2018Q3. This confirms the main findings from Figures 6 and 7, with increased weights on expected future inflation for all but

[18] If we take the exchange rate as exogenous, the implied long-run passthrough becomes higher on average for Latin America relative to Asian EMEs given comparatively larger average coefficients on lagged inflation at the end of the sample period.

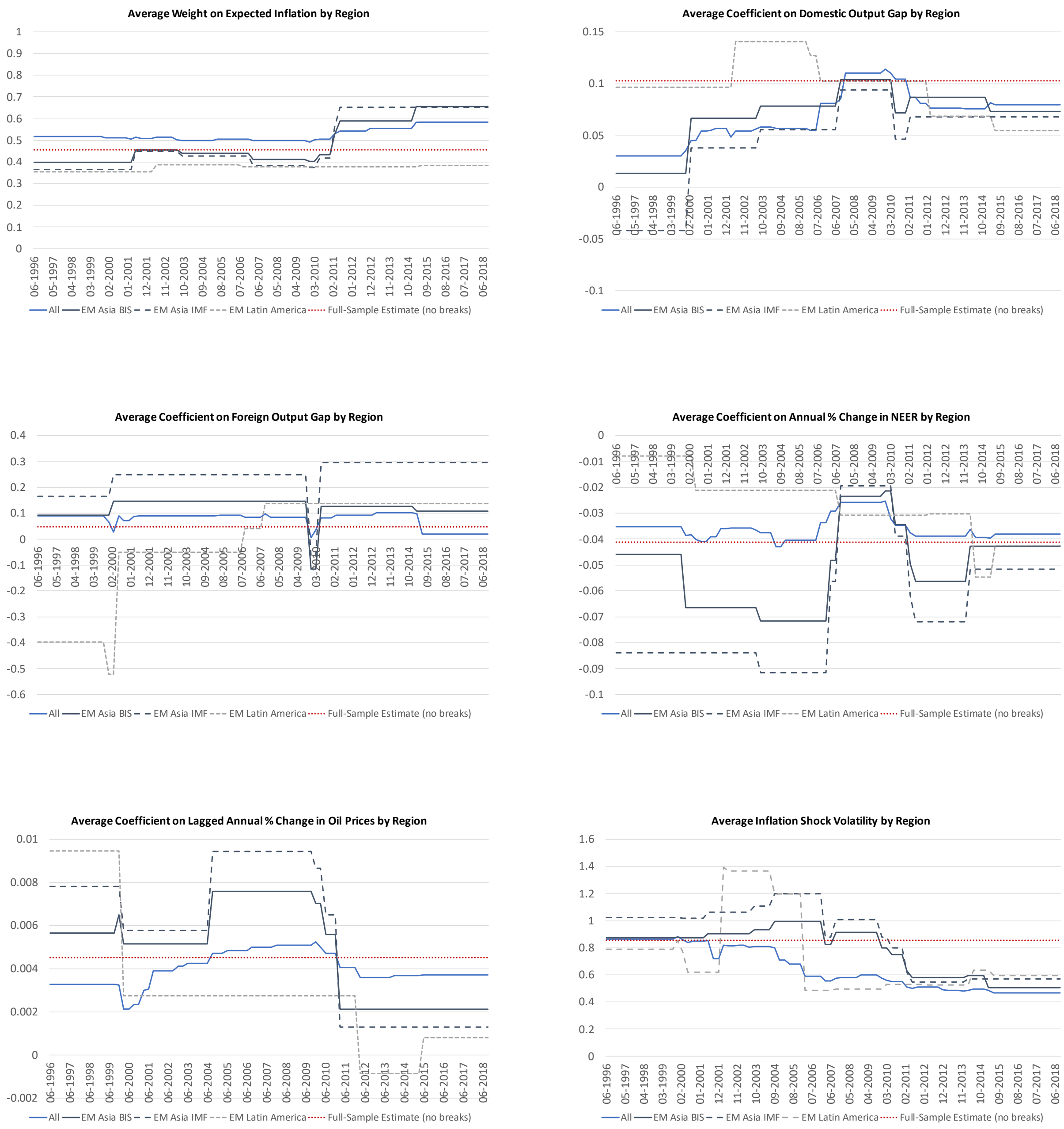
Figure 6: Average parameter estimates over time by different classifications of economies



Notes: 'AEs' stands for advanced economies and 'EMEs' stands for emerging market economies. Estimated parameters are conditioned on estimated breakdates and whether a likelihood ratio test suggests that a structural break applies to a given slope coefficient based on Qu and Perron (2007) procedures. Volatility is measured in standard deviation terms.

Latin America, increased slopes for domestic Phillips curves for all but Latin America, increased slopes of foreign Phillips curves for EMEs but slopes close to zero for AEs (BIS classification), decreased degree of exchange rate passthrough for emerging markets except Latin America and very little exchange rate passthrough for AEs (BIS classification), and increased impacts of oil prices for AEs but decreased impacts for Asian EMEs and Latin America. For completeness, the table also reports results for the BN filter output gaps and shows that they are highly robust.

Figure 7: Average parameter estimates over time by different regions



Notes: 'AEs' stands for advanced economies and 'EMEs' stands for emerging market economies. Estimated parameters are conditioned on estimated breakdates and whether a likelihood ratio test suggests that a structural break applies to a given slope coefficient based on Qu and Perron (2007) procedures. Volatility is measured in standard deviation terms.

5.4 What explains changes in the level and volatility of inflation?

Despite some notable structural changes in the effects of the various possible inflation drivers, the most striking result from our analysis so far is how stable most of the parameters of the open economy hybrid Phillips curve model seem to be over the sample period. In particular, the results in Table 4 suggest that most slope coefficients in the model remained constant for most economies, while the results in Figure 5 suggest that the economic magnitude of any structural

Table 6: Average estimated effects by different groups of economies before and after the GFC

Variable	Group	(a) HP filter gaps				(b) BN filter gaps			
		1996Q2-2006Q4		2007Q1-2018Q3		1996Q2-2006Q4		2007Q1-2018Q3	
		Classifications							
		BIS	IMF	BIS	IMF	BIS	IMF	BIS	IMF
$E_t^s[\pi_{it+4}]$	All	0.511	0.511	0.543	0.543	0.509	0.509	0.567	0.567
	AEs	0.569	0.561	0.608	0.574	0.578	0.549	0.645	0.598
	EMEs	0.468	0.422	0.495	0.490	0.458	0.439	0.509	0.512
	Asian EMEs	0.424	0.404	0.550	0.561	0.436	0.451	0.532	0.553
	Latin America	0.369	0.369	0.380	0.380	0.404	0.404	0.397	0.397
$\tilde{y}_{it}$	All	0.048	0.048	0.088	0.088	0.020	0.020	0.092	0.092
	AEs	0.025	0.023	0.150	0.102	-0.018	-0.016	0.123	0.112
	EMEs	0.064	0.091	0.041	0.062	0.048	0.083	0.069	0.057
	Asian EMEs	0.052	0.016	0.085	0.072	0.060	-0.009	0.202	0.204
	Latin America	0.114	0.114	0.079	0.079	0.166	0.166	0.095	0.095
$\tilde{y}_{it}^*$	All	0.087	0.087	0.067	0.067	0.090	0.090	0.122	0.122
	AEs	0.047	0.116	-0.077	-0.001	-0.015	0.062	-0.013	0.093
	EMEs	0.116	0.035	0.173	0.186	0.168	0.139	0.222	0.173
	Asian EMEs	0.127	0.219	0.115	0.269	0.158	0.291	0.308	0.524
	Latin America	-0.182	-0.182	0.133	0.133	-0.313	-0.313	-0.192	-0.192
$\Delta_4 e_{it}$	All	-0.037	-0.037	-0.035	-0.035	-0.042	-0.042	-0.035	-0.035
	AE	-0.008	-0.025	-0.015	-0.030	-0.006	-0.032	-0.009	-0.030
	EM	-0.059	-0.060	-0.050	-0.043	-0.069	-0.060	-0.054	-0.043
	Asian EMEs	-0.061	-0.086	-0.041	-0.048	-0.061	-0.085	-0.050	-0.055
	Latin America	-0.016	-0.016	-0.036	-0.036	-0.021	-0.021	-0.035	-0.035
$\Delta_4 p_{t-1}^{oil}$	All	0.004	0.004	0.004	0.004	0.003	0.003	0.004	0.004
	AE s	0.003	0.003	0.004	0.005	0.003	0.002	0.004	0.004
	EMEs	0.004	0.005	0.004	0.004	0.003	0.006	0.004	0.004
	Asian EMEs	0.006	0.007	0.004	0.004	0.004	0.004	0.002	0.002
	Latin America	0.005	0.005	0.001	0.001	0.008	0.008	0.003	0.003

Notes: 'AEs' stands for advanced economies and 'EMEs' stands for emerging market economies, with corresponding economies listed in the data description in Section 3. The average estimated effects are $\frac{1}{NT}\sum_i \sum_t \hat{\beta}_{jit}$ for economy-by-economy estimation.

changes are small compared to the average estimated effects.

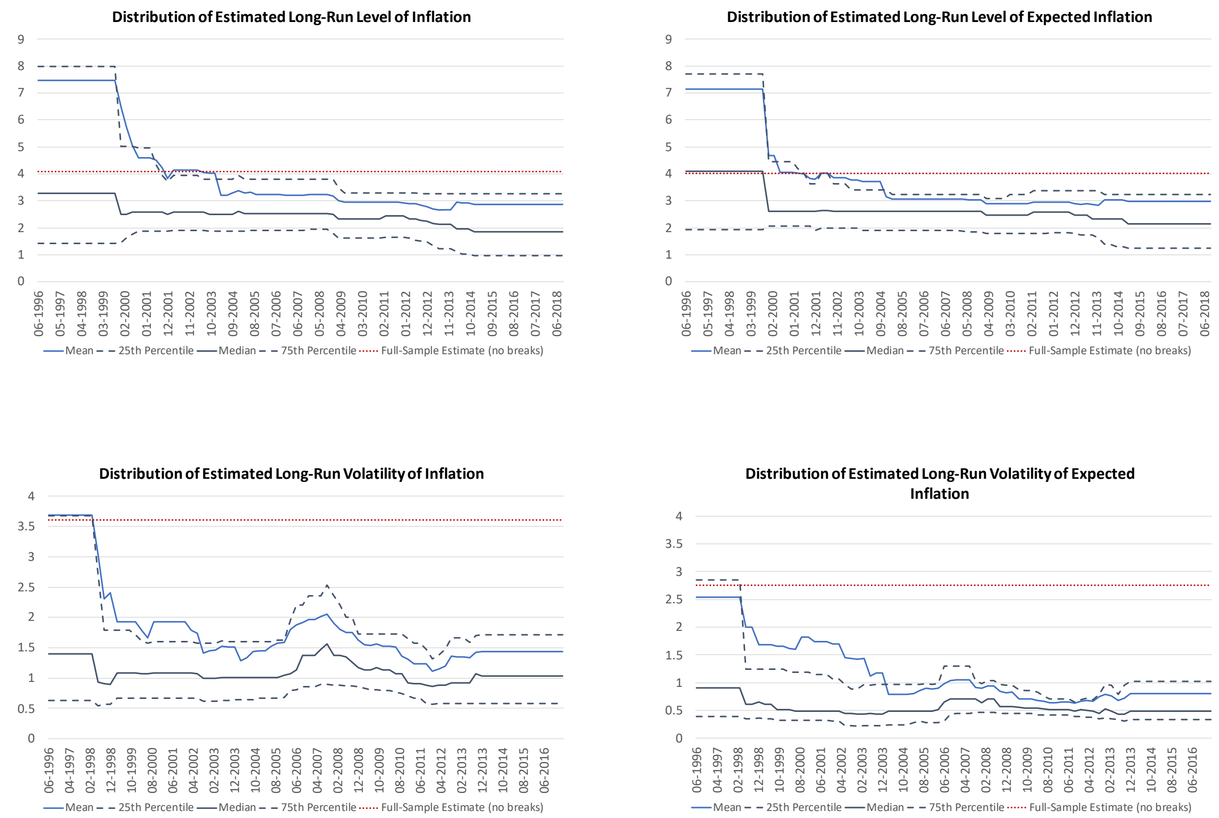
Given this relative stability of the open economy hybrid Phillips curve model, we examine the extent to which the behavior of the variables in the model can explain patterns of changes in the level and volatility of inflation over time. To investigate this, we consider the presence of structural breaks in the mean and variance of each possible driving variable and see if they are related to changes in the mean and variance of inflation. Using these results, we are then able to conduct a variance decomposition to examine the roles of the possible inflation drivers in explaining the behavior of inflation.

We first conduct univariate analysis of structural breaks for each variable in our open economy hybrid Phillips curve model using a simple regression on a constant:

$$x_{it} = \mu_i + u_{it} \tag{6}$$

where $x_i = \{\pi_i, \pi_i^e, \tilde{y}_i, \tilde{y}_i^*, \Delta_4 e_i, \Delta_4 p^{oil}\}$, μ_i is the mean, and the residual u_{it} is allowed to be heteroskedastic and serially correlated with long-run variance $\bar{\sigma}_{xi}^2$. The allowance for general serial

Figure 8: Distributions for estimated long-run level and volatility of inflation and expected inflation over time



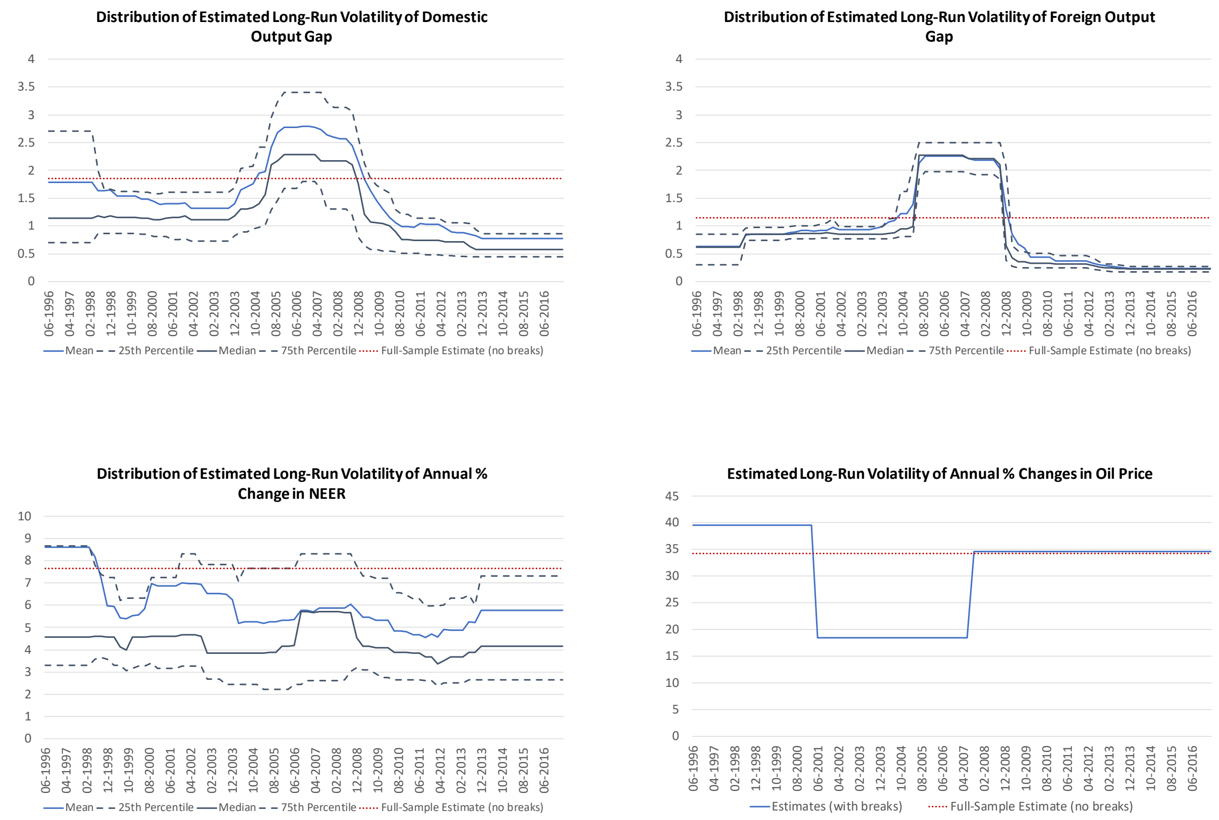
Notes: Estimated long-run levels and volatilities are conditioned on estimated breakdates for univariate analysis of structural breaks in mean and/or variance based on Qu and Perron (2007) procedures. Volatility is measured in standard deviation terms.

correlation in the residual makes this basic regression model consistent with the open economy hybrid Phillips curve model in equations (4) and (5) under the assumption of stationarity other than the possible infrequent structural breaks. Again, we use the methods in Qu and Perron (2007) to test for and estimate structural breaks in each variable.

Figure 8 reports how the long-run level and volatility of inflation and expected future inflation across different economies have evolved over time conditioning on estimated breakdates by plotting the mean, median, and quartiles of the distributions for estimates for the univariate model across economies. The mean estimate of the long-run level of inflation fell dramatically in the early 2000s from 7.5% to 3% by the end of the sample period. The median estimate is more stable, but also fell from just over 3% to just under 2%. Consistent with an anchoring of inflation expectations and suggestive of a proximate cause for the reduction in the long-run level of inflation, there are similar estimated reductions in the long-run level of expected future inflation.[19] Likewise, the mean and median estimates of the long-run volatility of inflation and inflation expectations fell from the beginning to the end of the sample period, suggesting again an anchoring of inflation expectations and that expected future inflation is a proximate cause

[19]We also found a reduction in the average level exchange rate depreciation for our panel of economies that is consistent with the convergence of inflation rates and long-run purchasing power parity, although we note the usual caveats about possible endogeneity in attributing causality from exchange rate passthrough to explaining the corresponding reduction in inflation.

Figure 9: Distributions for estimated long-run volatility of output gaps, exchange rates, and oil prices over time



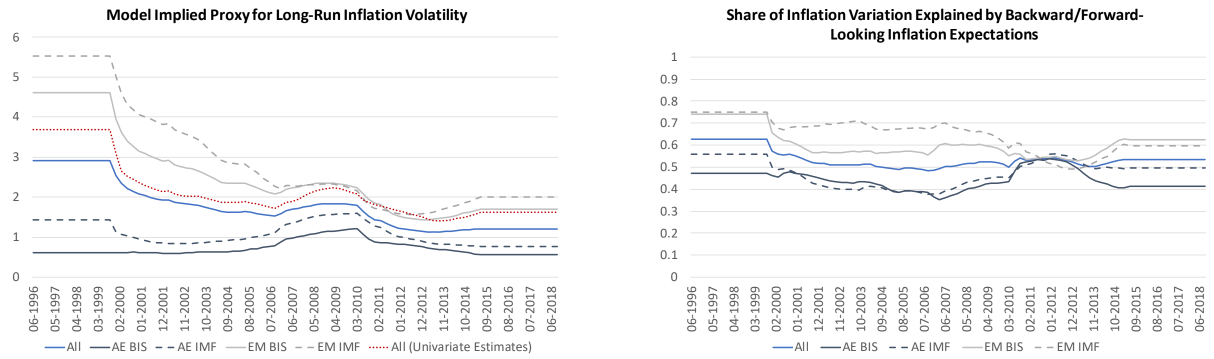
Notes: Estimated unconditional mean and volatility are conditioned on estimated breakdates for univariate analysis of structural breaks in mean and/or variance based on Qu and Perron (2007) procedures. Output gaps are measured using the HP filter. Volatility is measured in standard deviation terms.

for the reduction in inflation volatility. However, it is notable that inflation volatility increased more during the GFC than the volatility of inflation expectations, which is generally lower throughout the sample period. Looking back at Figure 5, the gradual decrease in the mean and little change in the median of the volatility of the inflation shock suggests that other drivers must have played some role in determining inflation volatility.

Figure 9 reports how the long-run volatility estimates for the other variables have evolved (we report results for HP filter output gaps only, although we found that the results for BN filter output gaps were very similar). The large increases in the long-run volatility of the domestic and foreign output gaps during the GFC can clearly explain the increase in inflation volatility at that time. This is consistent with the idea that the Phillips curve reasserted its presence during the GFC given the large and persistent negative output gaps at the time (see Stock and Watson 2010). The changes in the long-run volatility estimates for exchange rates and oil prices are consistent with the pattern of changes in long-run inflation volatility, but causality is unclear for exchange rates and, based on timing, oil prices are clearly not the main driver.

Although these results are consistent with a relatively stable open economy hybrid Phillips curve, where large estimated weights on expected future inflation appear to explain most changes in the long-run level of inflation and the long-run volatilities of expected future inflation and the output gaps appearing to explain most changes in long-run inflation volatility,

Figure 10: Average long-run volatility estimates and variance share of inflation expectations by different classifications of economies



Notes: Estimates are based on underlying weighted-average estimates of parameters for each economy using HP filter output gaps. Volatility is measured in standard deviation terms, while shares are based on decomposition of model-implied proxy variance of inflation.

a formal variance decomposition for inflation based on the open economy hybrid Phillips curve model would be useful to quantify the exact roles of the different drivers of inflation. However, a strict variance decomposition is not possible given possible correlation between the driving variables. To address this, we construct and decompose a model-implied "proxy variance" of inflation for economy i, $\tilde{\sigma}^2_{\pi i} \approx \sigma^2_{\pi i}$, that ignores any such correlation:

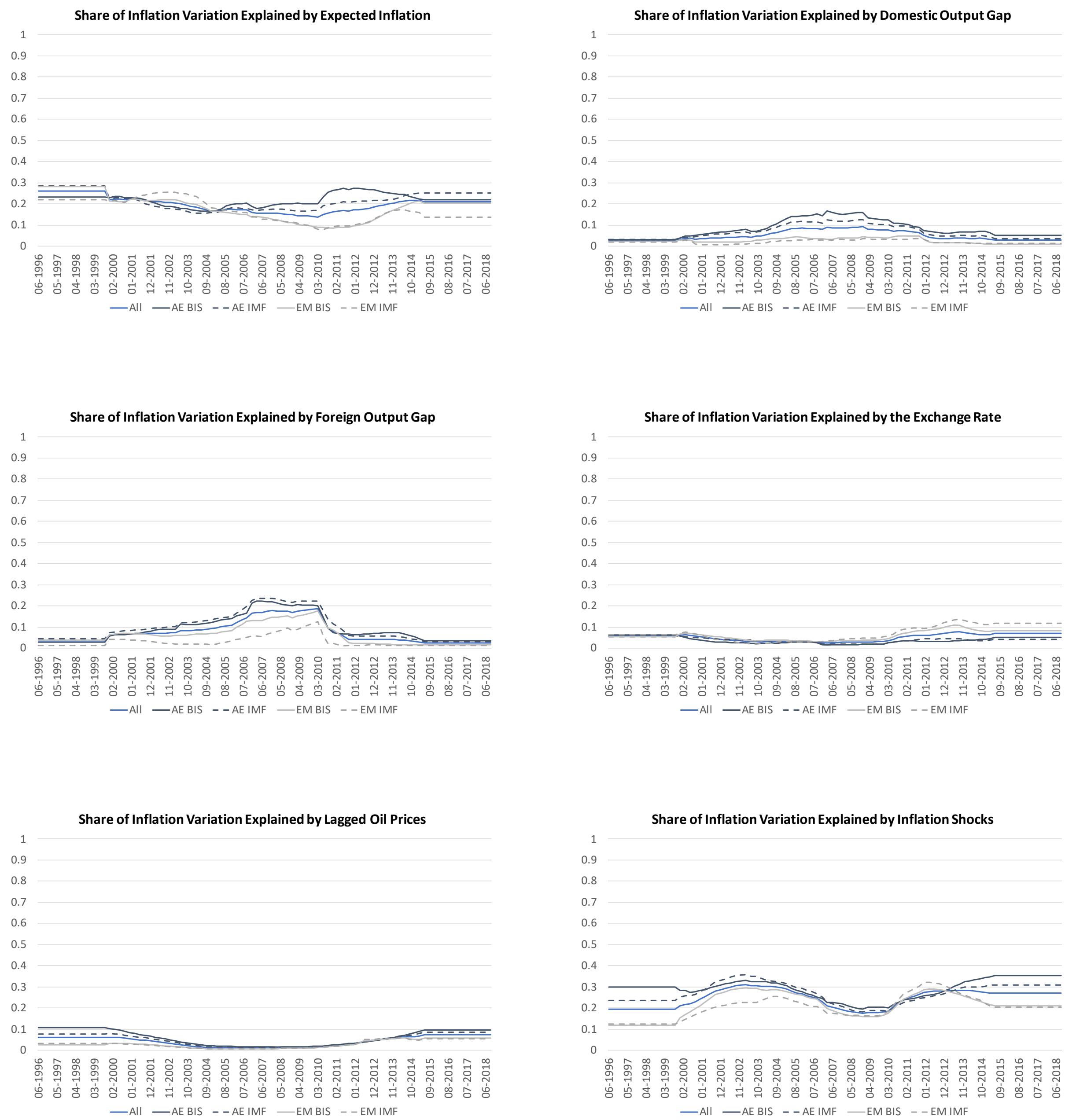
$$\tilde{\sigma}^2_{\pi i} = \sum_j \beta^2_{ji}\bar{\sigma}^2_{ji} + \bar{\sigma}^2_{\varepsilon i} \tag{7}$$

where β_{ji} and $\bar{\sigma}^2_{\varepsilon i}$ are from the model in equation (5) and $\bar{\sigma}^2_{ji}$ is from the univariate model in equation (6). We use estimates of these parameters when allowing for structural breaks to calculate variance shares for each determinant of inflation in the model, including inflation shocks, over time as $\hat{\beta}^2_{jit}\hat{\bar{\sigma}}^2_{jit}/\hat{\tilde{\sigma}}^2_{\pi it}$ and $\hat{\bar{\sigma}}^2_{\varepsilon it}/\hat{\tilde{\sigma}}^2_{\pi it}$, where weighted-averaged estimates are used to account for uncertainty about the timing of structural changes, as in Figure 5 (again, we only report the results for HP filter output gaps, although we found that the results for BN filter output gaps were very similar).

Figure 10 reports the average long-run volatility and variance shares of overall inflation expectations by different classifications of economies. The first thing to notice is that the approximation $\hat{\tilde{\sigma}}_{\pi it} \approx \hat{\sigma}_{\pi it}$ holds reasonably well in practice, albeit with fairly stable downward bias due to ignored (implied net positive) correlation between the possible inflation drivers. Importantly, the patterns of changes in volatility over time are the same for the proxy measure. Meanwhile, overall inflation expectations (both forward- and backward-looking measures) explain a substantial portion of the overall variation in inflation, with generally a greater role for EMEs that has declined somewhat from the beginning to the end of the sample period. However, it is clear that overall inflation expectations are the main driving force for inflation in both advanced and emerging market economies according to our open-economy hybrid Phillips curve model.

Figure 11 reports the variance shares for various possible inflation drivers in the model, including inflation shocks. The specific role of expected future inflation (i.e., forward-looking expectations) is generally but not always higher for AEs than EMEs, declining in early 2000s, increasing in early 2010s, and eventually settling around 20% for both AEs and EMEs based on

Figure 11: Variance shares of inflation driving forces by different classifications of economies



Notes: Estimates are based on underlying weighted-average estimates of parameters for each economy using HP filter output gaps. Volatility is measured in standard deviation terms, while shares are based on decomposition of model-implied proxy variance of inflation.

BIS classifications. The importance of the domestic and foreign output gaps is relatively low at the beginning and end of the sample period, but increases during the GFC, especially for AEs, when the combined shares peak at about 35%. Exchange rate passthrough explains less than 5% of inflation variation during most of the 2000s, but increases from 2010 on to peak about 10% for emerging markets. Despite the fact that we consider headline inflation, oil prices only ever explain a small share of the variance of inflation (less than 10% on average) and as little as 2% during the GFC even though oil prices moved dramatically at the time. Meanwhile, inflation shocks explain a nontrivial portion of the overall variation in inflation, with a rise in the early 2000s, a fall during the GFC, and a rise again after the GFC to as high as 35% for AEs and 20% for EMEs based on BIS classifications. However, it is notable that the high volatility of inflation in the GFC is not attributed to inflation shocks and there is no pattern of an ongoing increasing share for inflation shocks that one might expect if we had somehow failed to capture all of the underlying structural changes in the open economy hybrid Phillips curve model.

6 Conclusions

For most of the 47 economies that we consider, we find relatively little significant change in the driving forces of inflation or their quantitative effects over the sample period from 1996 to 2018. Forward- and backward-looking inflation expectations and domestic and foreign output gaps are all important drivers of inflation, albeit with very heterogeneous effects across different economies. Our finding of a relatively stable open economy hybrid Phillips curve model for most economies might seem surprising when one considers the dramatic change in policy environment that occurred with the onset of the GFC along with the Lucas (1976) critique, although our results do suggest there has been an increase in the average weight on expected future inflation for both advanced and emerging market economies. However, as would be predicted by a relatively stable open-economy hybrid Phillips curve model, structural changes in expected future inflation, domestic and foreign output gaps, and, to a lesser extent, exchange rates and oil prices can explain explain patterns of changes in the level and volatility of inflation across different economies.

The broad policy implication is that central banks should focus on inflation expectations as the key driver of inflation, but also allow for a substantial influence of the degree of economic slack both domestically and for major trading partners. Put simply, the Phillips curve is not dead, but it needs to account for forward-looking expectations and a foreign output gap. Doing so provides a surprisingly stable and useful structure for understanding the behavior of inflation in both advanced and emerging market economies.

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