Effects of Commodity Price Shocks on Inflation: A Cross Country Analysis*

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Abstract

Using local projections, this paper investigates effects of commodity price shocks on inflation. We estimate impulse responses of the consumer price indexes (CPIs) to a commodity price shock, based on a monthly panel consisting of 120 countries. Our results from the local projections suggest that the CPIs are almost fully adjusted within a year in response to a commodity price shock and thus effects of commodity price shocks are transitory. We then explore the possibility that the responses of the CPIs may be dependent on the inflation regimes. Based on the smooth transition autoregressive models that use the past inflation rate as a transition variable, we find that commodity price shocks have more persistent effects on inflation under the low inflation regime than under the high inflation regime. Our analysis also shows that, under the high inflation regime, there are (i) stabilizing roles of the exchange rate on consumer prices; and (ii) large differences in price responses between developed and less developed countries. However, these effects are not detected under the low inflation regime. Our findings suggest that business cycle factors may play an important role in understanding effects of commodity price shocks on the CPIs.

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Keywords: Commodity prices, inflation, pass-through, local projections, smooth transition autoregressive models

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1 Introduction

The fluctuations in the non-energy commodity prices since the early 2000s have renewed policymakers' attention on their effects on inflation.[1] One of the issues for policymakers is how the monetary policy reacts to the commodity price shocks. Among others, Yellen (2011) argues that commodity price shocks have only modest and transitory effects on the US CPI inflation and that a recent surge of commodity prices do not "warrant any substantial shift in the stance of monetary policy." On the other hand, European Central Bank (2008) and International Monetary Fund (2008) express some concerns about the upside risks to the price stability due to increased commodity prices.

The objective of this paper is to investigate effects of non-energy commodity price shocks on the consumer price indexes (CPIs) by estimating impulse responses (IRs). We use the monthly panel of the CPI consisting of 120 countries during the 2000s to address the following questions: How do commodity price shocks affect the CPI and inflation across the globe? Do commodity price shocks have persistent effects on inflation? What factors matter for effects of commodity price shocks? To answer these questions, we estimate IRs of the CPI to a commodity price shock by local projections developed by Jordà (2005). In addition to the benchmark linear projections, we estimate time-varying IRs based on the smooth transition autoregressive (STAR) models. Many previous studies on the relationship between inflation and non-energy commodity prices focus on pass-through of commodity prices to inflation rather than IRs and employ the linear estimations with constant parameters. These influential works include Cecchetti and Moessner (2008), International Monetary Fund (2008, 2011), Rigobon (2010), and Gelos and Ustyugova (2012).[2]

As emphasized by Jordà (2005), the local projections have several advantages in estimating IRs. First, local projections are relatively robust to misspecifications of the data generating process. Because we use panel data consisting of many countries, it would be difficult to specify the estimation equations from the viewpoint of economic theory. Hence, the robustness to misspecifications would be advantageous particularly in our context. Second, in comparison to the panel VAR, which would be one of alternative estimation strategies, we can economize the number of estimated parameters. Because we are interested only in responses of the CPIs to a commodity price shock, the direct

[1] See Bernanke (2008), European Central Bank (2008) and International Monetary Fund (2008).

[2] One of the few exceptions is Ferrucci, Jiménez-Rodríguez and Onorante (2012) who analyze the pass-through of commodity prices with nonlinear specifications. However, their study focuses on pass-through of international commodity prices to food price indexes in Euro area.

estimation of IRs by local projections is particularly useful. Third, the local projections are made by the least squares. Hence, we can easily extend the benchmark linear estimations to non-linear estimations.

Using the local linear projections, we show that the CPIs are, on average, almost fully adjusted to the long-run level within a year in response to a commodity price shock. Hence, the effects of commodity prices on inflation are transitory, as Yellen (2011) argued regarding the US inflation. The CPIs increase by 1.3 - 1.8 percent in 12 months in response to a 10 percent increase in commodity price shocks, if we measure the price responses by the 95 percent confidence intervals. But, in the subsequent year, increases in consumer prices are substantially small. In fact, the price responses in the first year account for more than 80 percent of total price increases.

The fast speed of price adjustment found in this paper is fairly robust even if we consider economic factors that may affect the shape of the IR functions. We allow for differences in price responses across country groups, based on the regimes for the exchange rate vis-à-vis the US dollars and the degrees of economic development. Consistent with the previous studies, the magnitudes of price responses substantially differ across country groups. However, the speed of price adjustment measured by the shape of the IR functions is basically the same across the country groups.

We also address the possibility that responses of consumer prices to non-energy commodity price shocks depend on the lagged inflation rate in each economy, based on the STAR model. A variety of empirical studies have highlighted the possibility of time-varying parameters.[3] We consider the possibility that IRs may vary over time, and compare the IRs between the high and low inflation regimes.

The estimation results suggest some interesting responses of the CPIs that are different in regimes. We find that, if a country is experiencing high inflation relative to the average inflation, commodity price shocks remain to have transitory effect on inflation, as in the benchmark linear estimations. In terms of the effects on the level of the CPI, the price responses under the high inflation regime are much stronger than those under the low inflation regime and differences in

[3] In the literature of oil price shocks, Chen (2009), Clark and Terry (2010), Shioji and Uchino (2011), and Baumeister and Peersman (2013) estimated effects of oil price shocks with time-varying parameters. Auerbach and Gorodnichenko (2012a, 2012b) estimated IRs to government spending shocks, based on the smooth transition vector autoregressive models. In their estimation, the government multipliers vary between recessions and expansions. In the literature on the exchange rate pass-through, Shintani, Terada-Hagiwara, and Yabu (2013) also found that exchange rate pass-through change depending on the past inflation rate in the US economy.

the price responses arising from the exchange rate and the degree of economic development can be detected. By contrast, if an economy is faced with low inflation, the effects of a commodity price shock on inflation are more persistent. The changes in the CPIs under the low inflation regime are approximately a half of those under the high inflation regime and the differences in the price responses across the country groups cannot be detected under the low inflation regimes.

This paper is organized as follows. Section 2 describes the benchmark linear projections and their results. Section 3 extends the estimation equation to the STAR model and reports the main results of this paper. Section 4 concludes.

2 Estimating IRs from Local Projections

2.1 Methodology: Benchmark regressions

To investigate effects of commodity price shocks on the CPIs, we employ local projection methods proposed by Jordà (2005). More specifically, we estimate the following equation for each forecast horizons k:

$$p_{j,t+k} - p_{j,t-1} = \alpha_{j,k} + \sum_{i=1}^{q} \beta_{i,k}(p_{j,t-i} - p_{j,t-i-1}) + \gamma_k FE_t + u^k_{j,t+k}, \qquad \text{for } k = 0, 1, 2, ..., K, \quad (1)$$

where $p_{j,t}$ represents the logarithm of the CPI for country j in period t, FE_t is an exogenous commodity price shock, which is unexpected in period t and serially uncorrelated. The error term is $u^k_{j,t+k}$. We will provide the detailed descriptions on FE_t in the next subsection. The estimated parameters are $\alpha_{j,k}$, $\beta_{i,k}$, and γ_k. Here, $\alpha_{j,k}$ includes the country fixed effects and $\beta_{i,k}$ captures persistence of inflation. In this equation, K denotes the maximum length of forecast horizons which is set to two years (i.e., 24 months in our monthly dataset) and the maximum number of lags q is chosen based on the data.[4] When $k = 0$, (1) turns out to be the standard specification for measuring responses to exogenous shocks.[5] When $k > 0$, the dependent variables mean changes in the CPI relative to the CPI before a commodity price shock.[6] Note that the dependent variables can also be expressed as the cumulative sum of inflation: $p_{j,t+k} - p_{j,t-1} = \pi_{j,t+k} + \pi_{j,t+k-1} + ... + \pi_{j,t}$,

[4]We use Bayesian Information Criterion (BIC) to choose the lag length q.

[5]See Romer and Romer (2004), for example.

[6]A similar specification was employed by Furceri and Zdzienicka (2012) who estimated effects of debt crises on output with local projections.

where $\pi_{j,t}$ is the inflation rate in country j in period t. This formulation ensures the stationarity of the dependent variables as long as the inflation rate is stationary.

The key parameter in (1) is γ_k, which represents the response of k-period ahead consumer prices to a current commodity price shock. Owing to the local projections, we directly estimate IRs of the CPI to a commodity price shock. That is, the IR for kth period after a commodity price shock can be written as

$$IR(k) = \gamma_k, \tag{2}$$

for $k = 0, 1, ..., K$. Note that all coefficients in (1) are separately estimated for each horizon k. As discussed in the literature, we need not to use the estimates of $\beta_{i,k}$ for computing IRs and their standard errors. The lagged inflation on the right-hand side is introduced only to control for the inflation persistence rather than to estimate dynamic effects of commodity price shocks.

We estimate (1) by Least Squares Dummy Variable (LSDV) estimator with Newey-West heteroskedasticity and autocorrelation consistent covariance matrix. When $k > 0$, $u^k_{j,t+k}$ follows a k-th order moving average process. Thus, we set the truncation parameter for covariance matrix estimation to k in each equation.

It is well-known that the presence of the lagged dependent variables in panel estimations may lead to a severe bias when the serial correlation of the dependent variables is high and the time series dimension of the data is short (Nickell, 1981). Though the serial correlation of the dependent variables may be high, the sample period used for estimation is relatively long ($T = 95$) compared to the cross-sectional dimension ($N = 120$). Considering relatively long time series, we continue to proceed with the LSDV estimator.

2.2 Commodity price shocks

We can estimate $IR(k)$ only when the data of commodity price shocks FE_t are available. In this paper, we take the commodity prices as given and rely on the empirical analysis by Chen, Rogoff, and Rossi (2010) in obtaining the proxy of FE_t. They show that the nominal exchange rate growths of the resource-rich countries such as Australia (AUS), Canada (CAN) and New Zealand (NZ) have strong forecasting power of commodity price inflation. Following Chen, Rogoff, and Rossi (2010), we make forecasts of commodity price inflation and assume that commodity price shocks can be represented by forecast errors in the forecasting model.

Let $\pi_{c,t}$ be commodity price inflation. We assume that the forecasting model is given by

$$\pi_{c,t} = a + b\pi_{c,t-1} + c_{AUS}\Delta s_{t-1}^{AUS} + c_{CAN}\Delta s_{t-1}^{CAN} + c_{NZ}\Delta s_{t-1}^{NZ} + \varepsilon_{c,t}, \tag{3}$$

where Δs_t^j denotes the nominal exchange rate growth in country j vis-à-vis the US for $j = AUS, CAN, NZ$. Here, a, b and c_j are parameters estimated and the lag length of explanatory variables is determined by BIC. We denote the resulting residuals by $\widehat{FE}_t$.

One may be concerned that the estimated forecast errors may not represent commodity price shocks, because (3) does not include potentially important macroeconomic variables such as industrial production indexes, monetary policy variables, and forward commodity prices. If they are critical in forecasting commodity prices, the commodity price shocks proxied by the residuals in (3) would be a linear combination of structural shocks of these variables. However, Groen and Pesenti (2011) find that the factor augmented regressions that replace Δs_t^j by the principal components from the factor model often perform poorly, compared to the Chen, Rogoff, and Rossi's (2010) regressions. Chen, Rogoff, and Rossi (2010) also argue that the nominal exchange rate growths are much more useful indicators for spot commodity price movements than the forward premium.

In (1), $\widehat{FE}_t(= \pi_{c,t} - \hat{\pi}_{c,t})$ results in a generated regressor that is measured with the sampling error. It is well-known that, due to the sampling error, the usual standard error estimation for the coefficients in (1) is downward biased (Murphy and Topel, 1985). To address this generated regressor's problem, we use the heteroskedasticity-robust version of Murphy and Topel's (1985) standard error correction.[7]

2.3 Data

We construct the balanced panel for the headline CPIs taken from the International Financial Statistics (IFS) of the International Monetary Fund over January 2000 – December 2010.[8] Because the raw data of the CPI is seasonally unadjusted, we use X-12-ARIMA procedure to make a

[7] See Hardin (2002).

[8] Some central banks pay close attention to a "core" measure of inflation that excludes or down-weights food and energy prices in their policymaking and thus, using the core measure might be ideal for some countries. However, the central banks in most countries watch the headline measure due to the large expenditure shares on foods and energy. Hence, our analysis throughout the paper relies on the headline CPI for all countries. See also Gregorio (2012) for the reason why most central banks focus on headline inflation.

seasonal adjustment. The IFS has reported 146 countries' CPI since January 2000. Out of these 146 countries, there are some countries that have missing values during the above sample periods. We exclude these countries from the dataset and the resulting number of the countries is reduced to 141. The CPIs in some countries are only available with quarterly basis.[9] We interpolate each quarterly CPI with the linear interpolation to obtain the monthly CPI. In our extended regression analysis, we also use the exchange rate regime classifications constructed by Ilzetzki, Reinhart and Rogoff (2010) and the data of the income classifications (i.e., high-, middle-, and low-income countries) published by the World Bank. In constructing the panel data with the exchange rate regimes and income classifications, the number of countries is further reduced to 120 countries. The list of these countries is reported in Table 1.

To estimate $\widehat{FE}_t$, we use non-energy Commodity Price Index published by the World Bank. Non-energy Commodity Price Index is a monthly index based on nominal US dollars and comprises Metals (31.6 percent), Agriculture (64.9 percent) and Fertilizer (3.6 percent). The nominal exchange rates s_t^{AUS}, s_t^{CAN} and s_t^{NZ} are taken from the Datastream.

As a preparatory analysis, Figure 1 reports the commodity price inflation and our forecasts based on (3) in the upper panel and the resulting forecast error $\widehat{FE}_t$ in the lower panel. All series in the figure are expressed at an annual rate and the forecast errors are estimated based on the sample period over February 2001 – December 2008.[10] Both commodity price inflation and its forecast errors are very volatile and show large declines particularly in the periods of the global financial crisis (September and October in 2008). We confirm that the nominal exchange rate growths in (3) Granger-cause commodity price inflation, consistent with Chen, Rogoff, and Rossi (2010), though the frequency of our data differs from theirs. In addition, the standard Ljung-Box statistics for $\widehat{FE}_t$ for lags 1 – 12 are all insignificant at conventional significance levels, suggesting that serial correlation would not exist in $\widehat{FE}_t$.

[9]These countries are Australia, New Zealand, and Papua New Guinea.

[10]In forecasting FE_t, it might be more reasonable to compute out-of-sample forecasts which use only the information available at period in which forecasts are made. However, we rely on in-sample forecasts because we allow for the standard error correction due to the sampling errors of the estimated FE_t. In this correction, the sample period must be the same as that for estimating (1). Because we allow for 24 leads in the left-hand side of the equation and the maximum lags of order 12 in the right-hand side, the resulting length of the sample period used for estimation is over February 2001- December 2008.

2.4 Benchmark results

Figure 2 plots the IRs of the CPIs to a 10 percent increase in a commodity price shock in period $t = 0$ based on (1). It also reports the 95 percent confidence interval bands represented by the shaded area. Panel (a) of the figure refers to the benchmark case. The panel indicates that the CPIs increase by 1.66 percent in period 9 and the estimated responses after this period range over 1.50 – 1.79 percent. The estimated confidence intervals of the responses show that the IRs are estimated quite precisely. In fact, the widths of the 95 percent confidence intervals are less than 1 percentage point for all forecast horizons.

The estimated IRs are defined not for inflation but for prices. In terms of inflation, the slope of the IR function in Figure 2 corresponds to inflation responses. According to Figure 2, the slope is substantially steep until period 9 but it turns out to be almost flat after the period. This implies that effects of commodity price shocks on inflation are transitory as discussed by Yellen (2011). In fact, the average response of inflation between periods 0 and 9 is 0.17 percent in response to a 10 percent increase in commodity price shocks. By contrast, the average response of inflation for the remaining periods is only 0.01 percent. This means that the CPI almost converges to the long-run level within a year after a commodity price shock.

Considering the fact that the estimated price responses after period 9 range between 1.57 and 1.79 percent to a 10 percent commodity price shock, one may be tempted to think that effects of commodity price shocks might be negligible or weak at least. This, however, may not necessarily be the case. As shown in Figure 1, our estimates of the commodity price shock $\widehat{FE}_t$ are substantially volatile. In fact, at the annual rate, the standard deviation of the estimated commodity price shocks is estimated to be 33.8 percent over February 2001 - December 2008. Hence, if a one-standard deviation shock of the commodity prices hits the economy, the CPIs, on average, increase by 6.09 percent in period 24. This magnitude of the price responses would not be negligible when we allow for the volatility of commodity price shocks.

There are a few remarks on the estimated IR function in panel (a) of Figure 2. First, our results are consistent with the estimated pass-through of commodity prices to the CPI that the previous studies have discovered. If we interpret the price responses in period 24 as the long-run price response, the price response of 1.79 percent at $k = 24$ to a 10 percent commodity price shock means the pass-through of 18.1 percent. Looking at the previous studies, Rigobon (2010) estimates the pass-through of wheat prices to the CPIs in 50 countries and reports the maximum pass-through for each country. If the maximum pass-through estimates in Rigobon (2010) are averaged over 50 countries, it leads to about 18 percent, consistent with our estimated pass-through.[11] Cecchetti and Moessner (2008) also estimate pass-through, using 27 countries and find that, for most countries, the pass-through of commodity prices to the one-year-ahead CPI ranges from 0 to 25 percent. Again, our estimate of 17.9 percent in terms of pass-through is consistent with the previous studies.

Second, our dataset includes countries with somewhat unstable inflation dynamics. Because these countries could affect the estimated IRs substantially, we eliminate these possible outliers from our sample. Defining these possible outliers as countries in which the standard deviation of inflation is more than 10 percent, we re-estimated (1) only for countries with stable inflation.[12] Panel (b) of the figure plots the estimated IRs of the CPI for the selected countries with stable inflation. The estimated IR function suggests that countries with large standard deviations of inflation do not substantially affect our benchmark results in panel (a).

Finally, as (2) indicates, our benchmark estimation assumes that the IR function is the same across all countries. This assumption would be too restrictive because country-specific factors may substantially influence the shape of each country's IR function. To consider this possibility, we extend (1) to allow for unknown country-specific effects on IRs. In particular, we run the following regression with coefficient dummies:

$$p_{j,t+k} - p_{j,t-1} = \alpha_{j,k} + \sum_{i=1}^{q} \beta_{i,k}(p_{j,t-i} - p_{j,t-i-1}) + \left(\gamma_k + \sum_{\ell=1}^{120} \gamma_{\ell,k} D_{\ell,j} \right) FE_t + u_{j,t+k}^k, \tag{4}$$

where $D_{\ell,j}$ is a country-specific dummy variable equal to unity when $\ell = j$. The parameters for the coefficient dummies $\gamma_{\ell,k}$ are estimated with the restriction that $\sum_{\ell=1}^{120} \gamma_{\ell,k} = 0$. This parameter could capture country-specific difference in responses arising from a variety of factors. In fact, the estimation of (4) reveals a large variability in $\hat{\gamma}_{\ell,k}$.[13] However, regardless of heterogeneity in the IRs, our estimation also reveals that most coefficients are not significantly different from zero. Thus, our analysis in what follows excludes country-specific coefficient dummies for FE_t and focuses on

[11]See Table A6 in Rigobon (2010).

[12]With this selection of countries, the number of countries is reduced to 91 countries.

[13]The estimation results are provided in Appendix, which is available upon request.

more specific economic factors.

3 What economic factors affect IRs of the CPIs?

We now investigate what economic factors influence the IRs of the CPIs to a commodity price shock. Among others, we consider (i) whether the currency is pegged to the US dollar; (ii) the degree of economic development; and (iii) the inflation rate in each country.

3.1 Motivation

3.1.1 The US-dollar-pegged exchange rate

Rigobon (2010) argues that there may be stabilizing roles of the nominal exchange rate variations on the CPIs. Because commodity prices are often denominated in the US dollar, the pass-through of commodity prices to the CPIs would be affected by exchange rate movements. If appreciations of a country's currency take place with increases in commodity prices, ceteris paribus, the country with the flexible exchange rate would have smaller increases of the CPI in response to a commodity price shock than a country with the currency pegged to the US dollar. In this case, the exchange rate movement is stabilizing the country's consumer prices.

In our analysis, we separate our sample countries into countries based on whether their currency is pegged to the US dollar or not. The *de facto* classification is based on Ilzetzki, Reinhart and Rogoff (2010). Regarding more detailed classification, we follow Ilzetzki, Mendoza, and Végh (2013).[14] In our definition, the country group with the pegged exchange rate includes some Latin American and Caribbean economies as well as the US.

3.1.2 The degree of economic development

The previous studies argue that pass-through of commodity price shocks to the headline CPIs are negatively related to the degree of economic development. (International Monetary Fund, 2008, 2011, and Gelos and Ustyugova, 2012). Because the expenditure shares on commodities are much higher in less developed countries than the developed countries, the medium- and long-run

[14] Ilzetzki, Mendoza, and Végh (2013) categorized countries in the fixed exchange rate regime as countries with no legal tender, hard pegs, crawling pegs, and *de facto* or pre-announced or crawling bands of $\pm$ 2 percent.

effect on the headline CPI in less developed countries may be larger than the developed countries. International Monetary Fund (2011) also argues that inflation expectations are generally less well-anchored in less developed countries than in developed countries. If these factors could be well captured by the degree of economic development, the magnitude of price responses to a commodity price shock could be smaller in developed countries than in less developed countries. In terms of the data, our analysis follows the World Bank's classification which categorizes countries by the countries' income. In particular, we define the low- and middle-income economies as less developed countries and the high-income countries as the developed countries.

3.1.3 Inflation

We further investigate whether responses of consumer prices to non-energy commodity prices depend on each country's past inflation rate. In the literature of the exchange rate pass-through, Taylor (2000) argues that the pass-through or pricing power of firms would be influenced by the inflation regimes. Shintani, Terada-Hagiwara, and Yabu (2013) use the exponential STAR model and found that exchange rate pass-through varies depending on the past inflation rate in the US economy.[15] As we will explain in the next subsection, our analysis relies on the logistic STAR model with the transition variable of each country's lagged inflation to explore the possibility that IRs may vary between the high and low inflation regimes.

3.2 Methodology

To assess the effects of the first two factors on price responses to a commodity price shock, we augment the estimation equation with dummy variables for (i) countries with the US-dollar-pegged exchange rates; and (ii) less developed countries. Our estimation equation can be extended to the following form:

$$p_{j,t+k} - p_{j,t-1} = \alpha_{j,k} + \sum_{i=1}^{q} \beta_{i,k}(p_{j,t-i} - p_{j,t-i-1}) + \gamma_{j,k} FE_t + u^k_{j,t+k}, \qquad \text{for } k = 0, 1, 2, ..., K, \quad (5)$$

[15] In contrast to our analysis, they use the US producer price index rather than the CPI.

where $\gamma_{j,k}$ is the IR for country j defined as

$$IR(k,j) = \gamma_{j,k} = \gamma_k + \gamma_{USD,k} D_j^{USD} + \gamma_{LDC,k} D_j^{LDC}, \tag{6}$$

and $\alpha_{j,k}$ includes the two fixed effects for countries whose currency is pegged to the US dollar and for less developed countries, as well as the fixed effect for each country. Here, D_j^{USD} (D_j^{LDC}) is the dummy variable which takes one if a country j is pegging the currency to the US dollar (if a country is grouped into less developed countries) and zero otherwise.[16] The coefficients for the dummy variables are $\gamma_{USD,k}$ and $\gamma_{LDC,k}$ which measures a difference across country groups. In particular, γ_k represents the IRs of the CPIs to a commodity price shock for the developed countries under the flexible exchange rate regime, while $\gamma_k + \gamma_{USD,k}$ refers to the IRs for the developed countries with the pegged exchange rates. Hence, if $\gamma_{USD,k}$ takes a positive (negative) value, we interpret that the US dollar exchange rate is stabilizing (destabilizing) consumer prices. On the other hand, $\gamma_{LDC,k}$ will be positive if less developed countries have a larger price response to a commodity price shock.

To allow for effects the inflation rate on IRs, we extend the linear model (5) to the STAR model discussed in Teräsvirta (1994). Define $\gamma_{j,k}^L$ and $\gamma_{j,k}^H$, similar to (6), as the IRs of consumer prices for low (L) and high (H) inflation regimes, respectively. Our estimation equation is

$$\begin{aligned} p_{j,t+k} - p_{j,t-1} = \alpha_{j,k} &+ F(z_{j,t-d}) \left[\sum_{i=1}^{q} \beta_{i,k}^L (p_{j,t-i} - p_{j,t-i-1}) + \gamma_{j,k}^L FE_t \right] \\ &+ [1 - F(z_{j,t-d})] \left[\sum_{i=1}^{q} \beta_{i,k}^H (p_{j,t-i} - p_{j,t-i-1}) + \gamma_{j,k}^H FE_t \right] + u_{j,t+k}^k, \end{aligned} \tag{7}$$

for each forecast horizons $k = 0, 1, 2, ..., K$. Here, $F(z_{j,t-d})$ is the transition function, and $z_{j,t-d}$ is the transition variable where d denotes the delay parameter. In this specification, the transition function is given by

$$F(z_{j,t-d}) = \frac{\exp(-\delta z_{j,t-d})}{1 + \exp(-\delta z_{j,t-d})}, \tag{8}$$

which implies that, as $z_{j,t-d} \to +\infty$, $F(z_{j,t-d}) \to 0$, meaning that the coefficients with subscript

[16] In practice, the dummy variable D_j^{USD} is not time-independent, because some countries move from an exchange rate regime to the other during the sample period. Likewise, D_j^{LDC} is not time-independent, either, due to the economic growth during the sample period. While we express dummies as being independent of t for notational simplicity, our estimation allowed for the time-dependence of the dummy variables.

H dominate the dynamics of the dependent variable. By contrast, as $z_{j,t-d} \to -\infty$, $F(z_{j,t-d}) \to 1$, implying that the coefficients with subscript L are dominating. Likewise, if $\delta \to 0$, $F(z_{j,t-d})$ converges to 1/2, which is effectively equivalent to (5), because parameters with subscript H and L cannot be identified.

Our specification follows Auerbach and Gorodnichenko (2012b) who investigate whether the fiscal multipliers depend on business cycle factors, based on the multi-country panel data. As in Auerbach and Gorodnichenko (2012b), we do not include the location parameter into the function $F(\cdot)$, in contrast to the applications of the single-equation-based STAR models. Instead, we standardize the transition variable so that $z_{j,t-d}$ has zero mean and a unit variance. This standardization is somewhat too restrictive, but allows us to estimate (7) as a linear function if one fixes the parameter δ in (8). Though Auerbach and Gorodnichenko (2012b) parameterize δ at a single value in their analysis, we allow for more flexible parameterizations of δ by grid search. In other words, we parameterize δ over a range of $\delta \in (0, \Delta]$ and run linear regressions for each grid constructed from an interval of $(0, \Delta]$. We then search for the optimal δ that minimizes the sum of squared residuals. We repeat this procedure for integer $d \leq q$ and choose a pair of (δ, d) that minimizes the sum of squared residuals.

Recall that, owing to the fact that we are estimating IRs by local projections, the IRs remain extremely easy to obtain in this nonlinear specification: $IR^L(k, j) = \gamma^L_{j,k}$ and $IR^H(k, j) = \gamma^H_{j,k}$, for $k = 1, 2, ..., K$. In general, the IR in country j can be represented by $IR(k, j, t) = F(z_{j,t-d})\gamma^L_{j,k} + [1 - F(z_{j,t-d})]\gamma^H_{j,k}$. Note that, while (6) considers differences in the IRs between country groups, the STAR model further allows us to investigate the time-varying effect of $z_{j,t-d}$ on IRs. Assuming that the inflation rate affects pass-through of shocks and the firms' pricing power as discussed in Taylor (2000), we specify the transition variable $z_{j,t-d}$ as the standardized past inflation rate:

$$z_{j,t-d} = \frac{\pi_{j,t-d} - \bar{\pi}_j}{\hat{\sigma}_j}, \tag{9}$$

where $\pi_{j,t-d} = p_{j,t-d} - p_{j,t-d-1}$, and $\bar{\pi}_j$ and $\hat{\sigma}_j$ are the time-series average and the time-series standard deviations of the inflation rate in country j, respectively.

3.3 Results

3.3.1 IRs from the linear model

We first interpret the results without allowing for the effect of the inflation regimes on IRs to confirm whether the results are consistent with the previous studies. We then crystalize the roles of the inflation regime. Figure 3 plots the IRs of the CPIs to a 10 percent increase in commodity price shocks based on (5). The figure plots four IR functions, depending on the inclusion of dummy variables $\gamma_{USD,k}$ and $\gamma_{LDC,k}$. More specifically, panel (a) plots the IRs for the developed countries with the flexible exchange rates while panel (b) corresponds to the IRs for the developed countries with the pegged exchange rates. Panels (c) and (d) are the IRs for less developed countries. As in the upper panels, these panels are separated based on the exchange rate regimes.

Overall, the figure shows that price increases in the developed countries with the flexible exchange ares the smallest among the four price response paths. For example, the price increase in period 9 in panel (a) is 1.01 percent to a 10 percent increase in commodity price shocks. However, turning to panel (b), we can confirm that the CPI in the developed countries with the US-dollar-pegged exchange rates increases by 1.83 percent in the same period to the same shock. On the other hand, less developed countries have larger price responses than developed countries as shown in panels (c) and (d). Among others, less developed countries in which their currencies are pegged to the US dollar are faced with the largest price increases to the same commodity price shock. As shown in panel (d), for example, the headline CPI increases by 2.24 percent in period 9.

The differences in the IRs in Figure 3 can also be more closely investigated by Table 2. The table reports the estimates of $\gamma_{USD,k}$ in the upper panel and $\gamma_{LDC,k}$ in the lower panel, along with Newey-West standard errors shown below the estimates. As shown in the upper panel, $\hat{\gamma}_{USD,k}$ are all positive, suggesting that exchange rate would be playing a role for stabilizing consumer prices, similar to Rigobon (2010). Rigobon (2010) finds a stabilizing role of the exchange rate in the CPI to oil price shock. The positive estimates of $\gamma_{USD,k}$ suggest that, even when non-energy commodity price index excluding oil prices is used, there would be some stabilizing role of the exchange rate on consumer prices.[17]

[17] Rigobon (2010) noted that, in contrast to oil price shocks, price responses to a wheat price shock did not benefit from the exchange rate movement. This slight difference from our results may arise because our data of the non-energy commodity price index include more broad categories of commodities such as food and metals.

However, we should emphasize that the stabilizing role of the exchange rate cannot be found to be significantly different from zero in all periods after a commodity price shock. In the upper panel, $\hat{\gamma}_{USD,k}$ are statistically significant only between periods 2 and 16 while they are statistically insignificant for the other periods except for periods 23 and 24. Thus, based on the linear model (5), a stabilizing role of the exchange rate could exist only in the short-run, but may not be in the longer-run.

Turning to $\hat{\gamma}_{LDC,k}$ shown in the lower panel of Table 2, we find $\hat{\gamma}_{LDC,k}$ to be positive in all periods after a commodity price shock. Though the signs of $\hat{\gamma}_{LDC,k}$ are the same as the previous studies such as International Monetary Fund (2008, 2011) expected, $\hat{\gamma}_{LDC,k}$ are not necessarily significantly from zero in all periods. In particular, they are significantly different from zero in periods 5, 6, 10, and $12-24$. In our interpretation, this suggests that, in contrast to the case of $\hat{\gamma}_{USD,k}$, the effects of larger expenditure share on the CPI begin to appear only in the medium- and long-run.

We next investigate whether commodity price shocks have only transitory effects on *inflation*, similar to the benchmark regression. To clarify the effects on inflation rather than the price level, we transform the IR functions to the normalized IR functions in which all price responses in the last period in Figure 3 are normalized to unity. If we assume that the long-run IRs of the CPI are well approximated by the price responses in period 24, this normalized measure helps us understand how much the CPIs have increased before reaching to the long-run responses of the CPI.

Given the above assumption, Figure 4 reports the IRs normalized to the long-run IRs for the four country groups. The figure indicates that all country groups have almost identical speed of price adjustment. In all cases, the price responses exceed more than 80 percent of the long-run responses within a year and the rest of the price increases is observed in the remaining periods. Therefore, based on (5), commodity price shocks have only temporary effect on inflation in all country groups.

3.3.2 IRs from the STAR model

So far, we have considered only country groups but not considered business cycle factors in the analysis. In this section, we ask the data on how the responses of the CPIs to a commodity price shock are influenced by the past inflation rate and analyze whether the roles of the economic factors discussed in the previous section remain effective.

Table 3 reports the results of the grid search of δ and d as a preparatory analysis. A set of our grid search is based on the intervals $\delta \in (0, 80]$ and $d = 1, 2, ..., 6$. The grid search for δ suggests that δ that minimizes the sum of squared residuals takes a variety of values between 0.33 and 49.52. In contrast, the grid search for d results in $d = 1$ for all k in the estimation equation (7).

Figures 5 and 6 show the IRs of the CPI under the high inflation regime and the low inflation regime, respectively. Note that, by IRs in the high (low) inflation regime, we mean the estimates of $\gamma^H_{j,k}$ $(\gamma^L_{j,k})$ in (7) and it is multiplied by 10 due to the size of the shock. As in the previous figures, the shaded areas correspond to the 95 percent confidence intervals.

Comparisons of the two figures show some interesting features, depending on the degree of economic development as well as the inflation regimes. Not surprisingly, the IRs under the high inflation regime are larger than those under the low inflation regime in most forecast horizons k, but the differences between the inflation regimes are larger in the less developed countries. For example, the price response in period 9 shown in panel (c) of Figure 5 (the high inflation regime) is 1.84 percent while that in panel (c) of Figure 6 (the low inflation regime) points to only 1.08 percent.[18] In fact, even the 95 percent confidence intervals for IRs do not overlap in periods 5 and 6 and thus the price responses are, at least during some periods, statistically different between the high and low inflation regimes. The statistically different price responses are more significant in the less developed countries with the pegged exchange rates. In the point estimate in period 9, the CPI in these countries increases by 1.47 percent under the low inflation regime but the corresponding increase under the high inflation regime is 2.93 percent. In terms of the confidence intervals, the IRs do not overlap from period 3 to 11. By contrast, we do not find any statistically different IRs in the developed countries between the two inflation regimes.

We next turn to each inflation regime. Under the high inflation regime, our estimation results are qualitatively similar to the results based on the linear model (5). The coefficients for the dummy variables controlling for the effects of the pegged exchange rates and the degree of the economic development are shown in Table 4. Except for the magnitude of point estimates, the results are similar to the estimate in Table 2. In particular, the statistical significance of the stabilizing role of the exchange rate can be confirmed in the short-run ($\hat{\gamma}^H_{USD,k}$ for $k = 3, 4, ..., 13, 15$). However,

[18] Appendix provides the estimates and the 95 confidence intervals of the IRs.

the stabilizing role of the exchange rate cannot be confirmed for the remaining periods, at least statistically. The effects of larger expenditure share on the CPI, represented by $\gamma^H_{LDC,k}$ become significant in period $4-12$ and $18-24$. This suggests that the effects of larger expenditure share tend to be significant for longer periods.

Turning to the low inflation regime, we fail to find both of the stabilizing roles of the exchange rate and the effects of larger expenditure share, in contrast to the high inflation regime (Table 5). For all periods after a commodity price shock, both $\hat{\gamma}^L_{USD,k}$ and $\hat{\gamma}^L_{LDC,k}$ are statistically insignificant. As a consequence, under the low inflation regime, we do not observe large difference in the IRs across country groups.

Finally, we conduct the same exercises to assess whether commodity price shocks have temporary effects on inflation through the normalized IRs, as we did in the previous subsection. In the upper panel of Figure 7, we observe that the slope of the price responses in all country groups is steep for the first several periods but much flatter for the remaining periods. Therefore, once again, commodity price shocks have only transitory effects on inflation. By contrast, however, the slope of the IRs of the CPIs under the low inflation regime is moderately steep over all forecast horizons. Also, the slopes are similar across country groups. This implies that commodity price shocks, while having modest impacts on the price responses, have persistent effects on inflation if the economy is under the low inflation environment.

4 Conclusion

Using local projections, we estimated the IRs of the CPI to a commodity price shock and explored the implications for the inflation dynamics. Toward this end, we combined the local projections with the nonlinear STAR model to analyze the price responses. We show that there are some stabilizing roles of the exchange rate on consumer prices as in the previous literature, but found that the roles are limited only in the short-run and the stabilizing role is not present, or at least weak, under the low inflation regime. We also confirmed that the degree of economic development influences the magnitude of price responses as discovered by the previous studies. Our STAR model allows us to find that the effect is present only under the high inflation regime. In terms of the effect of commodity price shocks on inflation, we found a similarity of the effects across country groups and large heterogeneity across the inflation regimes. In particular, our estimation results suggest that the effect of rising commodity prices on inflation would be transitory in inflationary economies. On the other hand, however, when the economy is surrounded by low inflation environment, the effect of commodity prices on inflation is modest but persistent. Our findings suggest that business cycle factors may play an important role in understanding effects of commodity prices.

References

[1] Auerbach, A. and Y. Gorodnichenko, 2012a, "Fiscal Multipliers in Recession and Expansion" in *Fiscal Policy After the Financial Crisis* eds. A. Alesina and F. Giavazzi, University of Chicago Press.

[2] Auerbach, A. and Y. Gorodnichenko, 2012b, "Measuring the Output Responses to Fiscal Policy," *American Economic Journal: Economic Policy*, 4 (2), pp. 1-27.

[3] Baumeister, C., and G. Peersman, 2013, "Time-Varying Effects of Oil Supply Shocks on the US Economy," *American Economic Journal: Macroeconomics*, 5 (4), pp. 1-28.

[4] Bernanke, B. S., 2008, "Outstanding Issues in the Analysis of Inflation," Federal Reserve Bank of Boston's 53rd Annual Economic Conference, Chatham, Massachusetts, June 9, 2008.

[5] Cecchetti, S. and R. Moessner, 2008, "Commodity Prices and Inflation Dynamics," *BIS Quarterly Review*, December 2008.

[6] Chen, Y.-C., K. S. Rogoff, and B. Rossi, 2010, "Can Exchange Rates Forecast Commodity Prices?" *Quarterly Journal of Economics*, 125 (3), pp. 1145-1194.

[7] Clark, T. E., and S. J. Terry, 2010, "Time Variation in the Inflation Passthrough of Energy Prices," *Journal of Money, Credit and Banking*, 42 (7), pp. 1419-1433.

[8] Chen, S.-S., 2009, "Oil Price Pass-Through into Inflation" *Energy Economics*, 31 (1), pp. 126-33.

[9] European Central Bank, 2008, *Monthly Bulletin*, September 2008.

[10] De Gregorio, J., 2012, "Commodity Prices, Monetary Policy, and Inflation," *IMF Economic Review*, 60 (4), pp. 600 - 633.

[11] Ferrucci, G., R. Jiménez-Rodríguez, and L. Onorante, 2012, "Food Price Pass-Through in the Euro Area: Non-Linearity and the Role of the Common Agricultural Policy," *International Journal of Central Banking*, 8 (1), pp. 179-217.

[12] Furceri, D., and A. Zdzienicka, 2012, "How Costly are Debt Crisis?" *Journal of International Money and Finance*, 31 (4), pp. 726-742.

[13] Gelos, G. and Y. Ustyugova, 2012, "Inflation Reponses to Commodity Price Shocks – How and Why Do Country Differ?" IMF Working Paper, WP/12/225.

[14] Groen, J. J. and P. A. Pesenti, 2011, "Commodity Prices Commodity Currencies, and Global Economic Developments," in *Commodity Prices and Markets, East Asia Seminar on Economics*, eds. T. Itoh and A. Rose, University of Chicago Press.

[15] Hardin, J. W., 2002, "The Robust Variance Estimator for Two-stage Models," *Stata Journal*, 2 (3), pp. 253-266.

[16] Ilzetzki, E., E. G. Mendoza, C. A. Végh, 2013, "How Big (Small) Are Fiscal Multipliers?" *Journal of Monetary Economics*, 60 (2), pp. 239-254.

[17] Ilzetzki, E., C. M. Reinhart, and K. S. Rogoff, 2010, "The Country Chronologies and Background Material to Exchange Rate Arrangements into the 21st Century: Will the Anchor Currency Hold?" mimeo.

[18] International Monetary Fund, 2008, *World Economic Outlook*, September 2008.

[19] International Monetary Fund, 2011, *World Economic Outlook*, September 2011.

[20] Jordà, Ò., 2005, "Estimation and Inference of Impulse Responses by Local Projections," *American Economic Review*, 91 (1), pp. 161-182.

[21] Murphy, K. M. and R. H. Topel, 1985, "Estimation and Inference in Two-step Econometric Models," *Journal of Business and Economic Statistics*, 3 (4), pp. 370-379.

[22] Nickell, S., 1981, “Biases in Dynamic Models with Fixed Effects,” *Econometrica*, 49 (6), pp. 1417-1426.

[23] Rigobon, R., 2010, “Commodity Prices Pass-through,” Central Bank of Chile Working Papers, No. 572.

[24] Romer, C. D. and D. H. Romer, 2004, “A New Measure of Monetary Shocks: Derivation and Implications,” *American Economic Review*, 94 (4), pp. 1055-1084.

[25] Shintani, M., A. Terada-Hagiwara, and T. Yabu, 2013, “Exchange Rate Pass-through and Inflation: A Nonlinear Time Series Analysis,” *Journal of International Money and Finance*, 32, pp. 512-227.

[26] Shioji, E., and T. Uchino, 2011, “Pass-Through of Oil Prices to Japanese Domestic Prices,” in *Commodity Prices and Markets, East Asia Seminar on Economics*, eds. T. Itoh and A. Rose, University of Chicago Press.

[27] Taylor, J. B., 2000, “Low Inflation, Pass-through and the Pricing Power of Firms,” *European Economic Review*, 44 (7), pp. 1389-1408.

[28] Teräsvirta, T., 1994, “Specification, Estimation, and Evaluation of Smooth Transition Autoregressive Models,” *Journal of the American Statistical Association*, 89 (425), pp. 208-218.

[29] Yellen, J., 2011, “Commodity Prices, the Economic Outlook, and Monetary Policy,” speech at the Economic Club of New York, New York, 11 April 2011.

Table 1: List of Countries

Countries			
Albania*	Ecuador	Kuwait*	Romania
Argentina	Egypt*	Latvia*	Russian Federation*
Armenia	El Salvador*	Lithuania*	Saudi Arabia*
Australia*	Estonia*	Macedonia, FYR*	Senegal*
Austria*	Finland*	Madagascar	Singapore*
Bahamas*	France*	Malawi	Slovak Republic*
Bangladesh*	Gabon*	Malaysia*	Slovenia*
Barbados*	Gambia*	Mali*	South Africa*
Belgium*	Georgia	Malta*	Spain*
Benin	Germany*	Mauritania*	Sri Lanka
Bolivia*	Ghana	Mauritius*	St. Kitts and Nevis*
Botswana*	Greece*	Mexico*	St. Lucia
Brazil*	Grenada*	Moldova*	St. Vincent & Grens.*
Bulgaria*	Guatemala*	Mongolia	Suriname
Burkina Faso	Guinea-Bissau	Morocco*	Swaziland
Burundi	Guyana	Mozambique	Sweden*
Cameroon*	Honduras*	Nepal*	Switzerland*
Canada*	Hungary*	Netherlands*	Thailand*
Cape Verde*	Iceland*	New Zealand*	Togo
Central African Rep.	India*	Nicaragua*	Trinidad and Tobago*
China,P.R.:Hong Kong*	Indonesia*	Niger	Tunisia*
Colombia*	Iran*	Nigeria	Turkey
Congo, Republic of	Ireland*	Norway*	Uganda*
Costa Rica*	Israel*	Pakistan*	United Kingdom*
Côte d'Ivoire*	Italy*	Panama*	United States*
Croatia*	Jamaica*	Papua New Guinea*	Uruguay*
Cyprus*	Japan*	Paraguay	Zambia*
Czech Republic*	Jordan	Peru*	
Denmark*	Kazakhstan*	Philippines*	
Dominica*	Kenya	Poland*	
Dominican Republic	Korea*	Portugal*	

Note: List of countries used in the analysis. In total, the dataset has 120 countries. The names of countries with an asterisk are countries where the standard deviation of the headline inflation is less than 10 percent over the sample period of January 2000 – December 2010.

Table 2: The estimates of dummy variables based on linear projections

(a) Coefficients of the dummy variable for the country group with the US-dollar-pegged exchange rates: $\gamma_{USD,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
0.005	0.009	0.021	0.031	0.046	0.047	0.055	0.068	0.074	0.076	0.074	0.071	0.072
(0.005)	(0.008)	(0.011)	(0.013)	(0.016)	(0.018)	(0.021)	(0.023)	(0.025)	(0.027)	(0.029)	(0.030)	(0.031)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.072	0.070	0.071	0.066	0.064	0.065	0.065	0.066	0.067	0.066	0.075	0.078	
(0.032)	(0.033)	(0.033)	(0.033)	(0.034)	(0.034)	(0.035)	(0.036)	(0.036)	(0.036)	(0.037)	(0.037)	

(b) Coefficients of the dummy variable for the less developed country group: $\gamma_{LDC,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
0.000	0.007	0.011	0.013	0.021	0.033	0.037	0.034	0.036	0.041	0.047	0.045	0.047
(0.005)	(0.006)	(0.009)	(0.011)	(0.013)	(0.015)	(0.017)	(0.019)	(0.020)	(0.021)	(0.023)	(0.024)	(0.024)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.050	0.051	0.049	0.053	0.052	0.061	0.062	0.061	0.059	0.061	0.065	0.067	
(0.024)	(0.025)	(0.025)	(0.026)	(0.026)	(0.026)	(0.027)	(0.027)	(0.027)	(0.028)	(0.028)	(0.028)	

Note: The numbers in parentheses are standard errors. In panel (a), the dummy variable is defined as the variable that takes one for countries whose currency is pegged to the US dollar and zero otherwise. In panel (b), the dummy variable takes one for the less developed countries and zero for developed countries.

Table 3: Parameter used for the STAR models

	$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
δ	0.328	2.276	4.670	6.596	6.514	6.416	6.992	5.715	7.452	8.106	7.062	7.536	29.994
d	1	1	1	1	1	1	1	1	1	1	1	1	1
q	12	12	12	12	11	11	11	10	8	7	7	7	3
	$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
δ	36.624	31.898	35.154	29.839	24.788	22.166	27.595	36.263	38.287	43.022	44.570	49.516	
d	2	2	2	2	2	2	2	2	2	2	2	2	
q	1	1	1	1	1	1	1	1	1	1	1	1	

Note: The parameters of δ in the transition function (8) are obtained from grid search over $(0, 80]$ and the delay parameter d is chosen from possible values of $d = 1, 2, ..., 6$. Parameter δ and d are chosen so as to minimize the sum of squared residuals after running the regression of (7) for each grid. The lag lengths of q in (7) are selected by BIC.

Table 4: The estimates of $\gamma^H_{USD,k}$ and $\gamma^H_{LDC,k}$ based on the STAR models: High inflation regime

(a) Coefficients of the dummy variable for the country group with the US-dollar-pegged exchange rates: $\gamma^H_{USD,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
-0.020	0.012	0.034	0.046	0.055	0.063	0.075	0.099	0.113	0.108	0.114	0.121	0.109
(0.041)	(0.016)	(0.018)	(0.021)	(0.025)	(0.028)	(0.031)	(0.036)	(0.038)	(0.041)	(0.043)	(0.045)	(0.044)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.086	0.081	0.088	0.083	0.076	0.068	0.062	0.067	0.062	0.056	0.066	0.068	
(0.042)	(0.043)	(0.043)	(0.043)	(0.044)	(0.045)	(0.046)	(0.047)	(0.047)	(0.048)	(0.049)	(0.049)	

(b) Coefficients of the dummy variable for the less developed country group: $\gamma^H_{LDC,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
0.056	0.027	0.028	0.031	0.045	0.061	0.070	0.067	0.062	0.066	0.073	0.068	0.071
(0.037)	(0.014)	(0.015)	(0.018)	(0.021)	(0.024)	(0.026)	(0.029)	(0.031)	(0.032)	(0.034)	(0.034)	(0.033)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.043	0.052	0.054	0.051	0.053	0.070	0.080	0.081	0.081	0.089	0.093	0.097	
(0.033)	(0.034)	(0.034)	(0.035)	(0.035)	(0.036)	(0.036)	(0.036)	(0.037)	(0.037)	(0.038)	(0.038)	

Note: The model is estimated with the STAR model. See the footnote of Table 2 for the other detail.

Table 5: The estimates of $\gamma^L_{USD,k}$ and $\gamma^L_{LDC,k}$ based on the STAR models: Low inflation regime

(a) Coefficients of the dummy variable for the country group with the US-dollar-pegged exchange rates: $\gamma^L_{USD,k}$

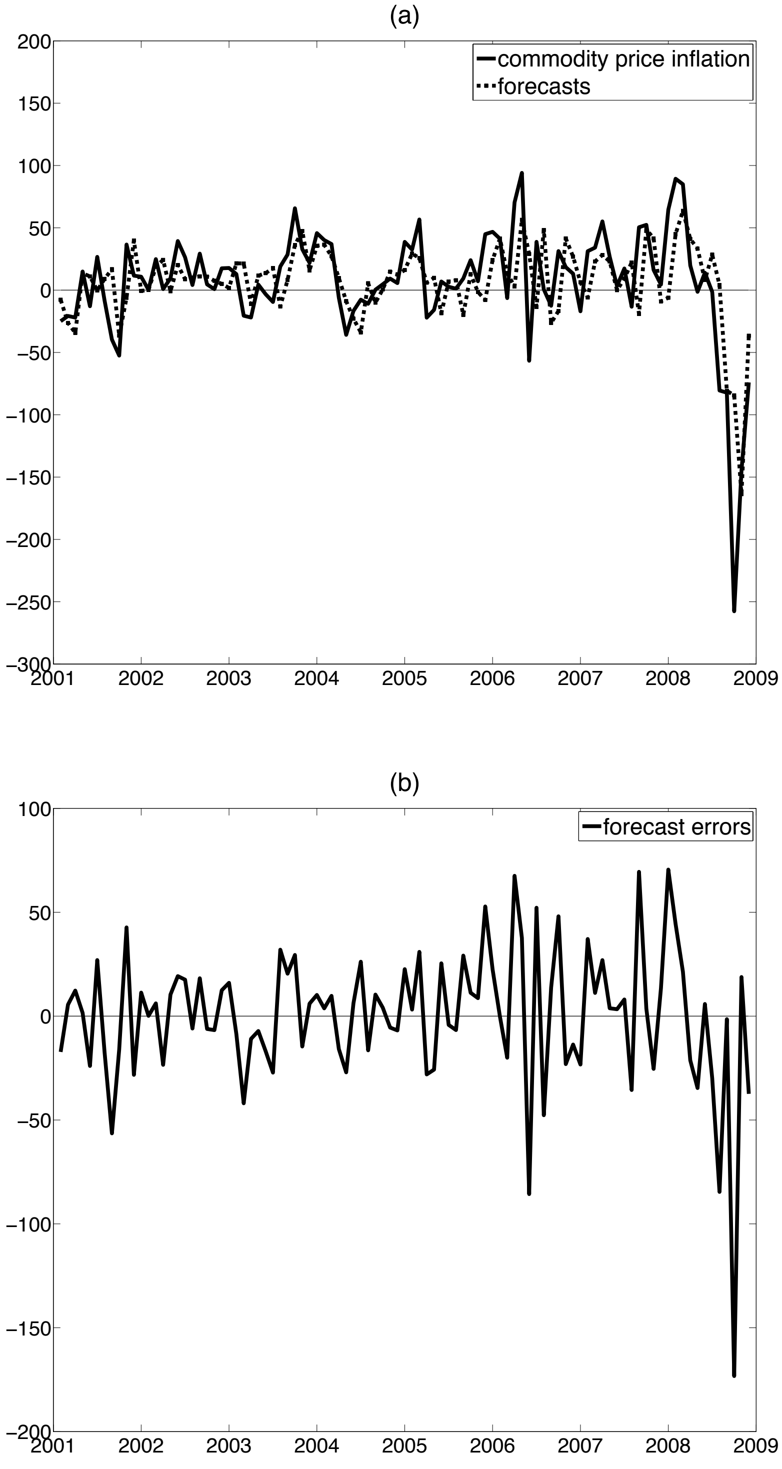
$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
0.036	0.008	0.009	0.015	0.037	0.030	0.034	0.034	0.031	0.039	0.029	0.016	0.029
(0.041)	(0.016)	(0.016)	(0.018)	(0.023)	(0.026)	(0.029)	(0.033)	(0.033)	(0.036)	(0.039)	(0.042)	(0.040)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.055	0.056	0.048	0.046	0.051	0.064	0.070	0.067	0.077	0.085	0.094	0.098	
(0.042)	(0.043)	(0.043)	(0.045)	(0.046)	(0.047)	(0.047)	(0.048)	(0.049)	(0.051)	(0.052)	(0.054)	

(b) Coefficients of the dummy variable for the less developed country group: $\gamma^L_{LDC,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
-0.059	-0.016	-0.007	-0.005	-0.003	0.005	0.005	0.004	0.013	0.014	0.021	0.022	0.028
(0.036)	(0.013)	(0.013)	(0.014)	(0.017)	(0.020)	(0.022)	(0.025)	(0.025)	(0.028)	(0.030)	(0.032)	(0.031)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.064	0.053	0.050	0.061	0.054	0.051	0.042	0.038	0.032	0.031	0.034	0.033	
(0.032)	(0.034)	(0.034)	(0.034)	(0.035)	(0.036)	(0.036)	(0.037)	(0.037)	(0.038)	(0.039)	(0.041)	

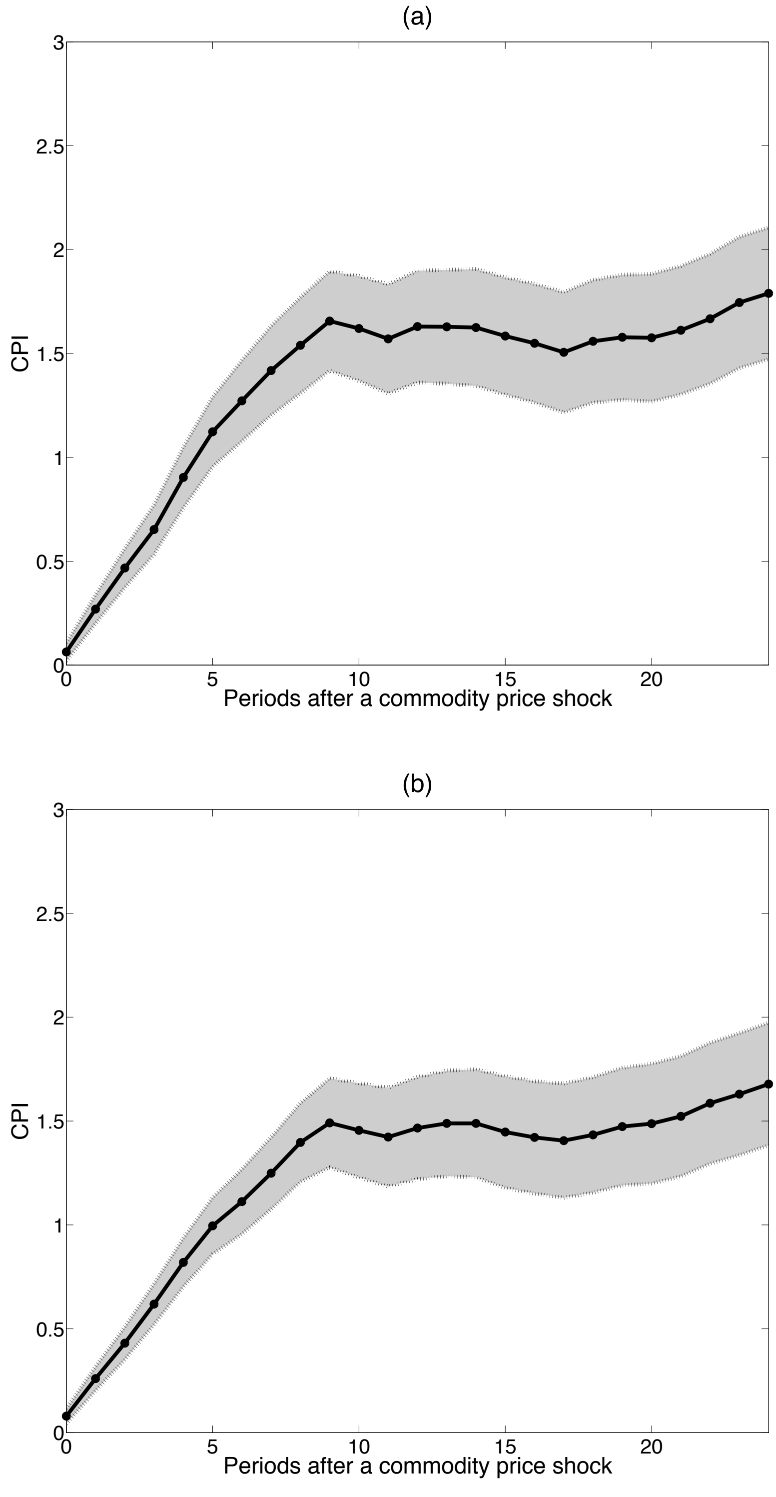
Note: The model is estimated with the STAR model. See the footnote of Table 2 for the other detail.

Figure 1: Forecast errors in commodity price inflation



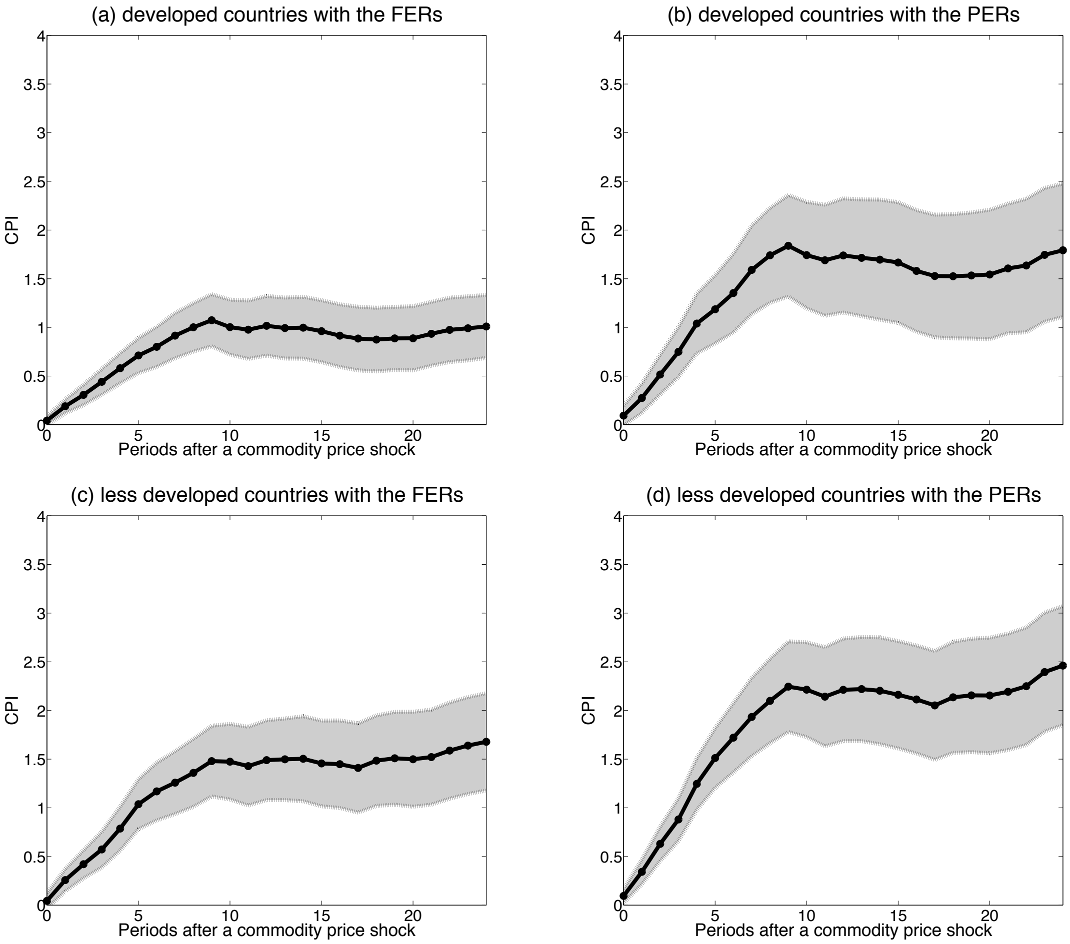
Note: Panel (a) shows commodity price inflation and its forecasts based on (3). Panel (b) plots the forecast errors.

Figure 2: Impulse responses of the CPI: benchmark regressions



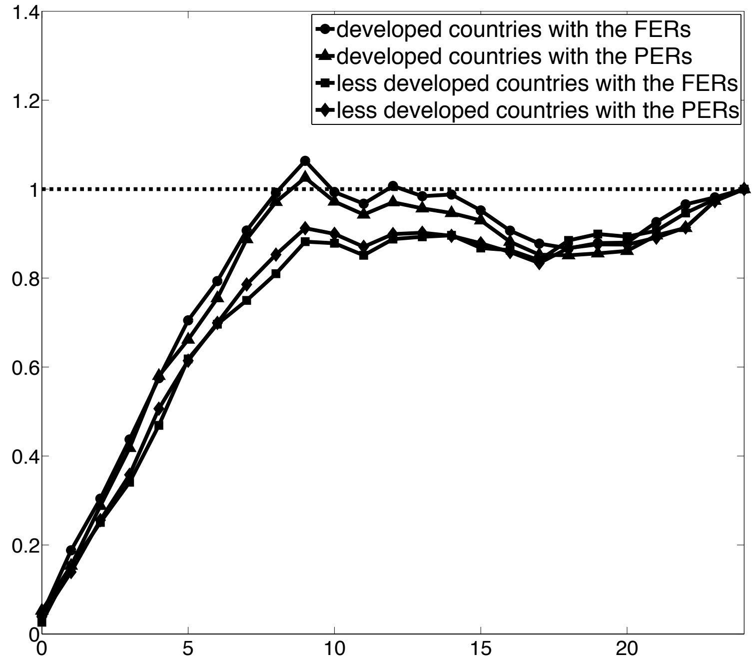
Note: The panels plot impulse responses of the CPI to a 10 percent increase in commodity prices. The unit of the responses is percent. Panel (a) shows the CPI responses estimated from 120-country panel. Panel (b) shows those estimated from the smaller panel dataset that excludes countries whose inflation is highly volatile. The shaded areas represent the 95 percent confidence intervals.

Figure 3: Impulse responses of the CPI with dummies



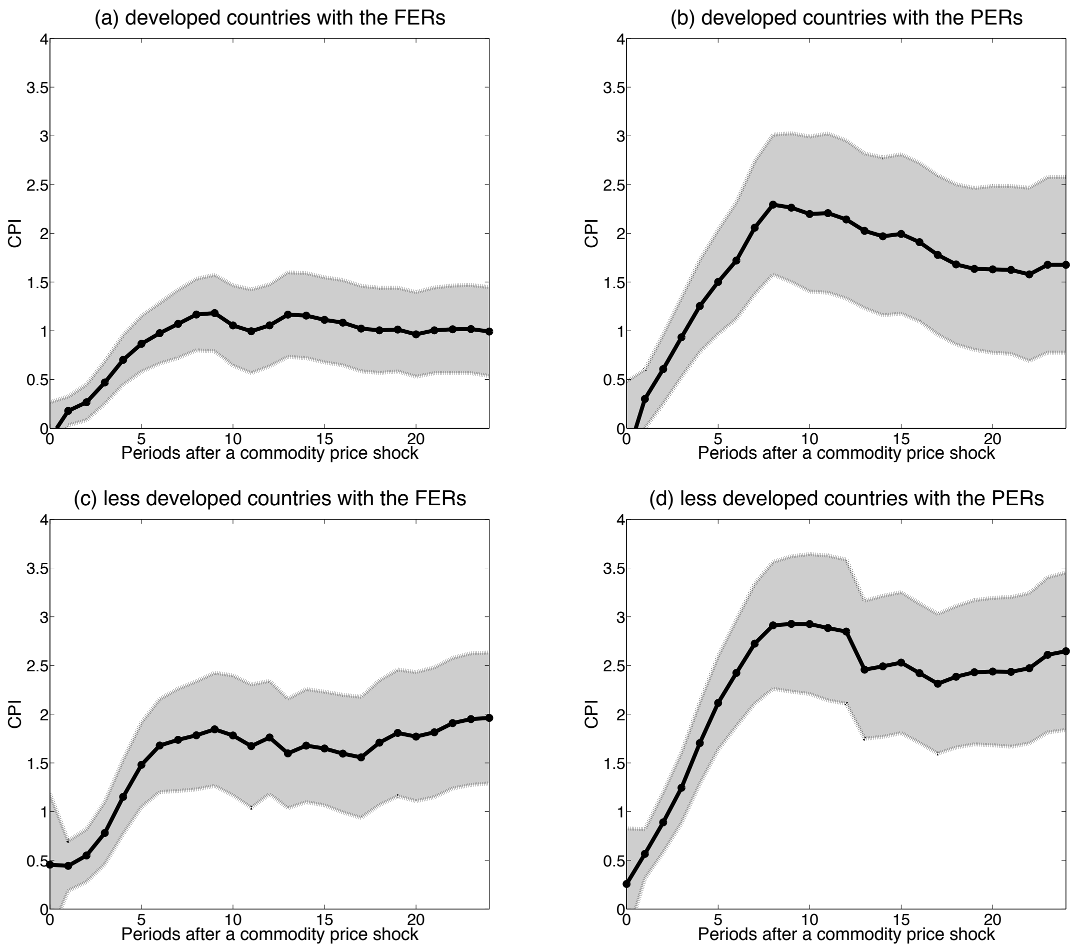
Note: The panels show impulse responses of the CPI to a 10 percent increase in commodity prices. The unit of the responses is percent. Panels (a) and (b) show the responses of the CPI in developed countries with flexible exchange rate and exchange rate pegged to the US dollar, respectively. Panels (c) and (d) correspond to those in less developed countries. The shaded areas represent the 95 percent confidence intervals. FER and PER stand for the flexible exchange rate and the pegged exchange rate, respectively.

Figure 4: Speed of price adjustment: benchmark regressions with dummies



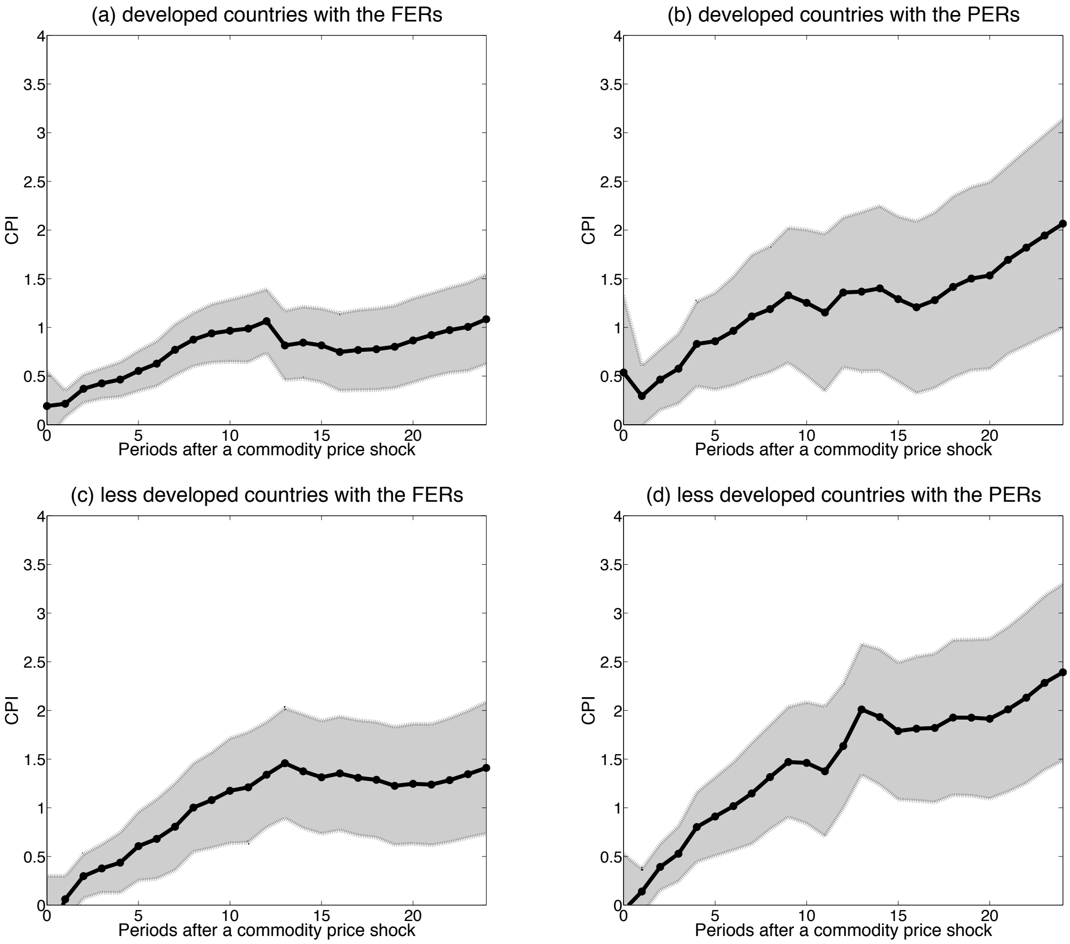
Note: The panel plots the impulse response functions normalized to the CPI responses at $k = 24$. In the figure, FER and PER stand for the flexible exchange rate and the pegged exchange rate, respectively.

Figure 5: Impulse responses of the CPI with dummies: high inflation regime in the STAR models



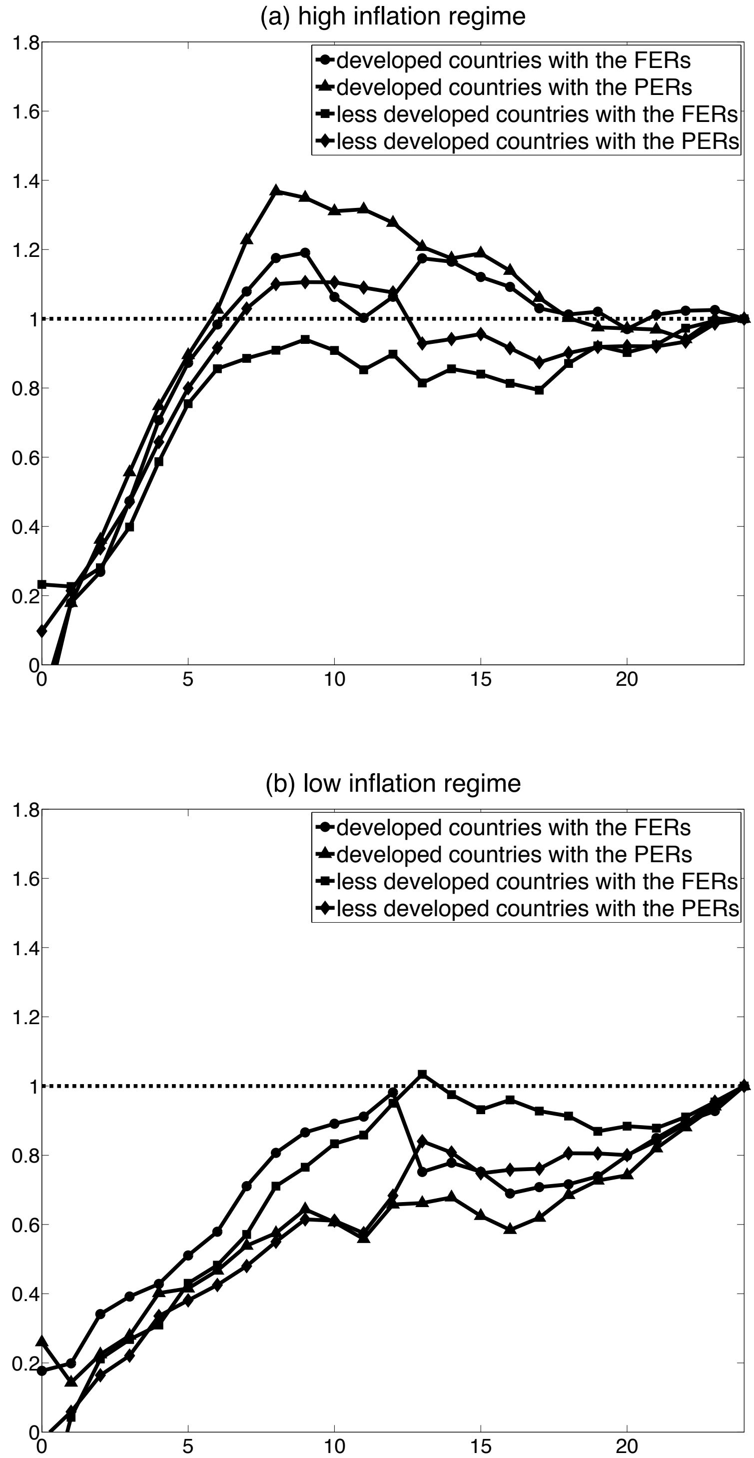
Note: The panels show impulse responses of the CPI to a 10 percent increase in commodity prices under the high inflation regime. The unit of the responses is percent. Panels (a) and (b) show the responses of the CPI in developed countries with flexible exchange rate and exchange rate pegged to the US dollar, respectively. Panels (c) and (d) correspond to those in less developed countries. The shaded areas represent the 95 percent confidence intervals. FER and PER stand for the flexible exchange rate and the pegged exchange rate, respectively.

Figure 6: Impulse responses of the CPI with dummies: low inflation regime in the STAR models



Note: The panels show impulse responses of the CPI to a 10 percent increase in commodity prices under the low inflation regime. The unit of the responses is percent. Panels (a) and (b) show the responses of the CPI in developed countries with flexible exchange rate and exchange rate pegged to the US dollar, respectively. Panels (c) and (d) correspond to those in less developed countries. The shaded areas represent the 95 percent confidence intervals. FER and PER stand for the flexible exchange rate and the pegged exchange rate, respectively.

Figure 7: Speed of price adjustment: STAR models



Note: Panels (a) and (b) plot the impulse response functions normalized to the CPI responses at $k = 24$. Panel (a) corresponds to the responses under the high inflation regime, while panel (b) shows the responses under the low inflation regime. In the figure, FER and PER stand for the flexible exchange rate and the pegged exchange rate, respectively.