JSPS Grants-in-Aid for Scientific Research (S)

Understanding Persistent Deflation in Japan

Working Paper Series

No. 056

December 2014

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

Atsushi Sekine

Takayuki Tsuruga

UTokyo Price Project
702 Faculty of Economics, The University of Tokyo,
7-3-1 Hongo, Bunkyo-ku, Tokyo 113-0033, Japan
Tel: +81-3-5841-5595
E-mail: watlab@e.u-tokyo.ac.jp
http://www.price.e.u-tokyo.ac.jp/english/

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

Atsushi Sekine[†] and Takayuki Tsuruga[‡]

December 18, 2014

Abstract

Since 2000s, large fluctuations in non-energy commodity prices have become a concern among policymakers about price stability. Using local projections, this paper investigates the effects of commodity price shocks on inflation. We estimate impulse responses of the consumer price indexes (CPIs) to commodity price shocks from a monthly panel consisting of 120 countries. Our analyses show that the effects of commodity price shocks on inflation are transitory. While the effect on the level of consumer prices varies across countries, the transitory effects on inflation are fairly robust, suggesting that policymakers may not need to pay special attention to the recent fluctuation in non-energy commodity prices. Employing the smooth transition autoregessive models that use the past inflation rate as the transition variable, we also explore the possibility that the effect of commodity price shocks is influenced by the inflation regimes. In this specification, commodity prices may not have transitory effects when a country is less developed and its currency is pegged to the U.S. dollar. However, the effect remains transitory in developed countries with exchange rate flexibility.

JEL Classification: E31; E37; Q43

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

*We are grateful to Takatoshi Ito, Yoichi Matsubayashi, and Yoshiyuki Nakazono for helpful comments and discussion. We also thank participants at the Asian Meeting of the Econometric Society and Summer Workshop on Economic Theory for comments. Takayuki Tsuruga acknowledges the financial support for the Grants-in-Aid for Scientific Research and JCER.

[†]Graduate School of Economics, Kyoto University

[‡]Graduate School of Economics, Kyoto University

1 Introduction

Fluctuations in the non-energy commodity prices since the early 2000s have renewed policymakers' attention to their effects on inflation.[1] One of the issues for policymakers is how monetary policy should respond to the commodity price shocks. Among others, Yellen (2011) argues that commodity price shocks have only modest and transitory effects on U.S. inflation and that a recent surge of commodity prices does 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 price stability due to rising inflation expectations triggered by commodity price shocks.

The objective of this paper is to investigate the effects of non-energy commodity price shocks on the Consumer Price Indexes (CPIs) by estimating impulse responses (IRs). We use a monthly panel of CPIs consisting of 120 countries during the 2000s to address the following questions: How do commodity price shocks affect CPIs and inflation across the globe? Do commodity price shocks, as Yellen (2011) argued, have only transitory effects on inflation? What factors matter for the effects of commodity price shocks? To answer these questions, we estimate the IRs of the CPI to a commodity price shock using local projections developed by Jordà (2005) and see how the IRs vary across the economic factors characterizing countries. In addition to the benchmark linear projections, we estimate IRs based on a smooth transition autoregressive (STAR) model to explore the possibility of regime-dependent IRs. While many previous studies on commodity prices focus on the pass-through of commodity prices to inflation with standard linear regressions, we primarily focus on the IRs of consumer prices to commodity price shocks, using both linear and nonlinear local projections.[2]

As emphasized by Jordà (2005), the local projections have several advantages in estimating IRs. First, this approach is relatively robust to misspecification 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 misspecification

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

[2] These influential previous studies include Cecchetti and Moessner (2008), International Monetary Fund (2008, 2011), Rigobon (2010), and Gelos and Ustyugova (2012). 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 the pass-through of international commodity prices to food price indexes in the euro area.

would be advantageous particularly in our context. Second, in comparison to the panel vector autoregressions (VAR), which are one of alternative estimation strategies, we can economize on the number of estimated parameters. Third, the local projections can be made using the least squares. Therefore, we can easily extend the benchmark linear estimations to those with interaction terms and/or nonlinear estimations.

Using the local linear projections, we show that the effects of commodity price shocks on inflation are transitory, as Yellen (2011) argued regarding U.S. inflation. Under the benchmark estimation, the CPIs increase by 1.4 to 1.9 percent in 12 months in response to a 10 percent commodity price shock, if we measure the price responses with the 95 percent confidence intervals. But in the subsequent year, changes in consumer prices become substantially small, implying that commodity price shocks have only a transitory effect on inflation.

For robustness, we also quantify the effect of economic factors that may affect IRs. We measure cross-sectional differences in price responses, based on the exchange rate flexibility vis-à-vis the U.S. dollars, a country's adoption of inflation targeting (IT), and the degree of economic development. Consistent with previous studies, the magnitudes of price responses differ substantially across country groups. We find that exchange rate flexibility and the adoption of IT dampen price responses during the first year after a commodity price shock, contributing to price stability in terms of commodity price shocks. The higher economic development also matters in reducing price responses to commodity price increases, and the effect is observed only in the medium or long run. Nevertheless, after we control for these economic factors affecting the IRs, the effects of commodity price shocks on inflation remain transitory under the linear projections.

We further explore the possibility that responses of consumer prices to non-energy commodity price shocks depend on the lagged inflation rate in each economy. We use the STAR model, which takes time-varying regression coefficients into account. A variety of empirical studies provide evidence on the possibility of time-varying regression coefficients.[3] We introduce time-varying regression coefficients into the estimations and focus on the IRs of the high- and low-inflation

[3] In the literature on oil price shocks, Chen (2009), Clark and Terry (2010), Shioji and Uchino (2011), and Baumeister and Peersman (2013) estimated the effects of oil price shocks with time-varying regression coefficients. Auerbach and Gorodnichenko (2012a, 2012b) estimated IRs of various macroeconomic variables 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 the exchange rate pass-through changes depending on the past inflation rate in the U.S. economy.

regimes.

The estimation results suggest some interesting responses of the CPIs that differ between high- and low-inflation regimes. When a country is less developed and its currency is pegged to the U.S. dollar, the IRs differ between inflation regimes. In particular, when the country is experiencing high inflation relative to its average inflation, commodity price shocks tend to increase inflation substantially, and turn to decrease afterwords. Under the low-inflation regime, the effect of commodity price shocks on inflation persists, although inflation *per se* is not substantially high. However, we emphasize that such differences are not observed when countries are developed and have exchange rate flexibility. In particular, IRs in the developed countries with exchange rate flexibility exhibit very similar patterns, regardless of inflation regimes.

This paper is organized as follows. Section 2 describes the methodology and the data. Section 3 shows the main results. In Section 4, we extend the estimation equation to the STAR model to explore the nonlinearity in regressions. Section 5 concludes.

2 Estimating IRs from Local Projections

2.1 Benchmark regressions

To introduce local projection methods proposed by Jordà (2005), consider a panel AR(q) process of the form

$$p_{j,t} - p_{j,t-1} = \alpha_j + \sum_{i=1}^{q} \beta_i (p_{j,t-i} - p_{j,t-i-1}) + \gamma u_{c,t} + u_{j,t}, \tag{1}$$

where $p_{j,t}$ represents the logarithm of the CPI for country j in period t. In this equation, we decompose shocks to inflation into international commodity price shocks $u_{c,t}$ and a linear combination of other shocks to inflation $u_{j,t}$, both of which are unexpected in period t, serially uncorrelated, and orthogonal to each other. For now, we leave the description of how unobservable $u_{c,t}$ is estimated to the subsequent section and assume that it is observable. Note that the left-hand-side variable is country j's inflation: $\pi_{j,t} = p_{j,t} - p_{j,t-1}$. Also, α_j includes the country fixed effects and β_i captures persistence of inflation.

The purpose of our analysis is to investigate the effects of commodity price shocks on the CPIs and inflation via local projections. We specify the estimation equation for local projections as follows:

$$p_{j,t+k} - p_{j,t-1} = \alpha_{j,k} + \sum_{i=1}^{q} \beta_{i,k}(p_{j,t-i} - p_{j,t-i-1}) + \gamma_k u_{c,t} + u^k_{j,t+k}, \tag{2}$$

where $k = 0, 1, 2, ..., K$ denotes the forecasting horizons.[4] In the equation, the estimated parameters are $\alpha_{j,k}$, $\beta_{i,k}$, and γ_k. Here, shocks to inflation are again decomposed into $u_{c,t}$ and $u^k_{j,t+k}$, but the latter typically includes $u_{j,t+i}$ for $i = 0, 1, ..., k$ and $u_{c,t+i}$ for $i = 1, 2, ..., k$, implying that $u^k_{j,t+k}$ follows the k-th order moving average process. In the estimation, we set K at 24 months and the maximum number of lags q is determined by the Bayesian Information Criterion (BIC).

The key parameter in (2) is γ_k, which represents the response of k-period-ahead consumer prices to a current commodity price shock. Another interpretation is the cumulative impulse response of inflation over k periods because $p_{j,t+k} - p_{j,t-1} = \pi_{j,t+k} + \pi_{j,t+k-1} + ... + \pi_{j,t}$. In our local projections, we directly estimate the IRs of the CPI to a commodity price shock. That is, the IR of consumer prices for k-th period after a one percent increase in commodity price shocks can be written as

$$IR(k) = \gamma_k, \tag{3}$$

for $k = 0, 1, ..., K$. Note that all coefficients in (2) are separately estimated for each horizon k. While the standard method of computing $IR(k)$ calls for the estimates of β_i and γ in (1), our local projections directly estimate γ_k without relying on estimates of $\beta_{i,k}$. This direct estimation tends to be more robust to misspecification of the stochastic process of inflation. (See Jordà, 2005 and Teulings and Zubanov, 2014.) The lagged inflation on the right-hand side of the equation is introduced only to control for the inflation persistence rather than to estimate the dynamic effects of commodity price shocks. We estimate (2) using least squares dummy variable (LSDV) estimator with Newey-West heteroskedasticity and autocorrelation consistent covariance matrix, because $u^k_{j,t+k}$ for $k > 0$ follows a k-th order moving average process.

It is well known that the presence of 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). Although the serial correlation of the dependent variables may be high, the sample period used for the estimation is relatively long ($T = 95$)

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

compared to the cross-sectional dimension ($N = 120$). Considering the relatively long time series, we proceed with the LSDV estimator.

2.2 Economic factors affecting IRs

Equation (3) assumes that the IRs are the same across all countries in our benchmark estimation. To relax this restrictive assumption, we introduce economic factors that may affect IRs of the CPIs to a commodity price shock. Among others, we consider (a) whether the currency is pegged to the U.S. dollar, (b) the adoption of IT, and (c) the degree of economic development.

The role of exchange rate variations Rigobon (2010) finds a stabilizing role of nominal exchange rate variation on consumer prices in response to commodity price shocks.[5] Because commodity prices are often denominated in U.S. dollars, the pass-through of commodity prices to the CPIs would be affected by exchange rate variations. If appreciations of a country's currency take place with increases in commodity prices, *ceteris paribus*, the CPI's response in this country would be smaller in response to a commodity price shock than in a country with a currency pegged to the U.S. dollar. In this case, the exchange rate variations stabilize the country's CPI.

Inflation targeting It has been argued that commodity price shocks destabilize inflation expectations, which in turn affect actual inflation.[6] If inflation expectations are well anchored by IT, inflation expectations respond less to a commodity price shock in IT countries than in non-IT countries. We would expect that IT countries have smaller IRs of the CPI than non-IT countries.

The degree of economic development International Monetary Fund (2008, 2011) also notes that the pass-through of commodity price shocks to headline CPIs are negatively related to the degree of economic development. (See also Gelos and Ustyugova, 2012.) Because the expenditure shares on commodities are much higher in less developed countries (LDCs) than in developed countries (DCs), the effect of this structural difference on the CPI in LDCs may be larger than the DCs, at least in the long run. If differences in the expenditure share on commodities are well

[5]See also, De Gregorio, Landerretche, and Neilson (2007).

[6]See International Monetary Fund (2008, 2011) and Cecchetti and Moessner (2008). Also, Levin, Natalucci, and Piger (2004) show evidence that IT stabilizes long-run inflation expectations.

captured by the degree of economic development, the magnitude of price responses to a commodity price shock could be larger in LDCs than in DCs.

Estimation To assess the impact of these factors on the IRs to commodity price shocks, we augment the estimation equation with interaction terms: the product of $u_{c,t}$ and a dummy variable. Our estimation equation is given by

$$p_{j,t+k} - p_{j,t-1} = \alpha_{j,k} + \sum_{i=1}^{q} \beta_{i,k}(p_{j,t-i} - p_{j,t-i-1}) + \gamma_{j,k} u_{c,t} + u^{k}_{j,t+k}, \qquad \text{for } k = 0, 1, 2, ..., K, \quad (4)$$

where $\gamma_{j,k}$ is equal to the IR for country j and is specified as

$$IR(k, j) = \gamma_{j,k} = \gamma_k + \gamma_{USD,k} D_j^{USD} + \gamma_{IT,k} D_j^{IT} + \gamma_{LDC,k} D_j^{LDC}. \quad (5)$$

Here D_j^{USD} is the dummy variable that takes one if country j pegs the currency to the U.S. dollar and zero otherwise. Similarly, we define D_j^{IT} as the dummy variable taking one for IT countries and D_j^{LDC} as the dummy variable taking one if a country is grouped with the LDCs. The intercept $\alpha_{j,k}$ also includes these dummies as well as the fixed effect for each country.[7]

The coefficients on the dummy variables measure the difference across country groups. If $\gamma_{USD,k}$ takes a positive (negative) value, we interpret this to mean that the U.S. dollar exchange rate is stabilizing (destabilizing) consumer prices. Also, if $\gamma_{IT,k}$ is estimated to be negative, this would be consistent with the hypothesis that IT contributes to anchoring inflation expectations. Finally, we expect that $\gamma_{LDC,k}$ should be positive at least in the long run.

2.3 Commodity price shocks

We can estimate the IRs of consumer prices only when commodity price shocks $u_{c,t}$ are observable. In this paper, we take the commodity prices as given and rely on the forecasting equation used by

[7]While we express dummies as being independent of t for notational simplicity, these new dummy variables are in practice time-dependent. For example, some countries move from an exchange rate regime to another regime during the sample period; and some LDCs experiencing high economic growth become categorized into DCs during the sample period. Our estimation allows for this time dependence, since it is not only more precise than time-independent dummies, but also allows us to avoid colinearity between country-specific fixed effects and the newly introduced dummy variables.

Chen, Rogoff, and Rossi (2010) to obtain a proxy of $u_{c,t}$:

$$\pi_{c,t} = a + b\pi_{c,t-1} + c_{AUS}\Delta s_{t-1}^{AUS} + c_{CAN}\Delta s_{t-1}^{CAN} + c_{NZ}\Delta s_{t-1}^{NZ} + \varepsilon_{c,t}, \tag{6}$$

where $\pi_{c,t}$ denotes commodity price inflation and Δs_t^j is the nominal exchange rate growth in country j vis-à-vis the U.S. for $j = AUS, CAN, NZ$. Here, a, b, and c_j are the parameters estimated and $\varepsilon_{c,t}$ is the error term. Chen, Rogoff, and Rossi (2010) show that the nominal exchange rate growth of resource-rich countries such as Australia (AUS), Canada (CAN), and New Zealand (NZ) has strong forecasting power for commodity price inflation.[8] In (6), the lag length of explanatory variables is determined by the BIC, but the main results in this paper are robust to the length of the lags.

We assume that commodity price shocks $u_{c,t}$ can be represented by forecast errors $\varepsilon_{c,t}$. Of course, we may suffer from bias in estimating $u_{c,t}$ because we might be excluding some important variables that are helpful in forecasting $\pi_{c,t}$ (e.g., forward commodity prices and macroeconomic variables.). However, the literature suggests that finding such variables is not an easy task. For example, Chen, Rogoff, and Rossi (2010) argue that nominal exchange rate growth is much more useful indicator for spot commodity price movements than the forward premium. Groen and Pesenti (2011) also find that the factor augmented regressions which replace Δs_t^j with the principal components from the factor model often perform poorly compared to the Chen, Rogoff, and Rossi's (2010) forecasting models.

Our estimate of the commodity price shock is $\hat{\varepsilon}_{c,t}(= \pi_{c,t} - \hat{\pi}_{c,t})$. However, the use of $\hat{\varepsilon}_{c,t}$ leads to a generated regressor's problem in our estimation. In this case, the usual standard error estimation for the coefficients in (2) and (4) 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. (See Hardin, 2002).

[8]Chen, Rogoff, and Rossi (2010) interpret their finding in terms of the difference in the informational content embodied in commodity prices and exchange rates. Commodity prices are sensitive to the global demand and supply and thus are more likely to be inaccurate in terms of containing the market conditions in the future. In contrast, the exchange rate is very sensitive to future market conditions, including expectations on commodity markets. Because of the information difference, exchange rates can help forecast commodity prices.

2.4 Data

We construct a balanced panel for the headline CPIs taken from the International Financial Statistics (IFS) of the International Monetary Fund from January 2000 to December 2010.[9] Because the raw data of the CPIs are seasonally unadjusted, we use the X-12-ARIMA procedure to make a seasonal adjustment. The IFS has reported 146 countries' CPIs since January 2000. Of these 146 countries, some have missing values during the above sample period. We exclude them from the dataset, reducing the number of countries to 141. The CPIs in some countries are only available on a quarterly basis.[10] We interpolate each quarterly CPI with linear interpolation to obtain the monthly CPI.

We also identify the exchange rate regimes, IT countries, and LDCs. For the exchange rate regime, we rely on the *de facto* classification constructed by Ilzetzki, Reinhart and Rogoff (2010).[11] For the choice of IT countries, we follow the definition of Roger (2009). Finally, for the degree of economic development, we define the low- and middle-income economies in the World Bank's classification as LDCs. In the process of constructing the panel data with these country-specific factors, the number of countries is further reduced to 120. These countries and the country-specific factors are summarized in Table 1.

To estimate commodity price shocks, we use the non-energy Commodity Price Index published by the World Bank. It is a monthly index denominated in the nominal U.S. 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 Datastream.

As a preparatory analysis, Figure 1 plots the commodity price inflation and our forecasts from (6) in the upper panel and the resulting forecast error $\hat{\varepsilon}_{c,t}$ in the lower panel. All series in the figure are expressed at an annual rate, and commodity price shocks are estimated based on the

[9]Some central banks pay close attention to "core" measure of inflation that excludes 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 monitor the headline measure due to the large expenditure shares on food and energy. Hence, our analysis throughout the paper relies on the headline CPI for all countries. See also De Gregorio (2012) for the reason why most central banks focus on headline inflation.

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

[11]Ilzetzki, Reinhart and Rogoff (2010) report 15 detailed classifications of exchange rate regimes (e.g., hard peg, crawling peg, and managed floating, etc.), together with the reference currency for the peg. In constructing the dummy variable for countries with D_j^{USD}, we define D_j^{USD} as unity, if country j adopts hard pegs to the U.S. dollar, crawling pegs to the U.S. dollar, or *de facto*, pre-announced or crawling bands around the U.S. dollar with $\pm$ 2 percent bands.

sample period from February 2001 to December 2008.[12] Both commodity price inflation and its forecast errors are very volatile and show large declines particularly during the global financial crisis (September and October 2008). We confirm that the nominal exchange rate growth in (6) Granger-causes commodity price inflation, consistent with Chen, Rogoff, and Rossi (2010), although the frequency of our data differs from theirs. In addition, the standard Ljung-Box statistics for $\hat{\varepsilon}_{c,t}$ for lags 1 to 12 are all insignificant at conventional significance levels, suggesting that serial correlation would not exist in the estimated commodity price shocks.

3 Results

3.1 IRs from the benchmark regressions

Figure 2 plots the estimated IRs of the CPIs to a 10 percent increase in a commodity price shock in period $t = 0$ based on (2). 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 IRs indicate that the CPIs increase by 1.66 percent (at an annual rate) until period 9 and the estimated responses after this period range from 1.51 to 1.79 percent. The 95 percent confidence intervals are narrow for all forecast horizons, suggesting that the IRs are estimated quite precisely.

Recall that the estimated IRs are defined not for inflation in a single month but for the cumulative sum of inflation. This means that inflation responses are measured by the slope of the IR function in Figure 2. According to Figure 2, the slope is substantially steep until period 9 but becomes nearly flat after this period. In fact, while the average response of inflation between periods 0 and 9 is 0.166 percent, that for the remaining periods (periods 10 to 24) is close to zero: 0.01 percent. Therefore, the effects of commodity price shocks on inflation are transitory as discussed by Yellen (2011), because the effect on inflation effectively disappears in 10 months.

While the effects on inflation are transitory, the impact on the price level is not negligible. As shown in Figure 1, our estimates of the commodity price shock $\hat{\varepsilon}_{c,t}$ are substantially volatile. In fact, at the annual rate, the standard deviation of $\hat{\varepsilon}_{c,t}$ is estimated at 33.8 percent from February

[12]This sample period is determined by the facts that we allow for 24 leads on the left-hand side of (2) and (4) and for the maximum lags of order 12 on the right-hand side. The resulting length of the sample period used for estimation is from February 2001 to December 2008. We also use in-sample forecasts, because the standard error correction for generated regressors requires in-sample forecasts rather than out-of-sample forecasts.

2001 to December 2008. Hence, if a one-standard-deviation shock to commodity prices hits an economy, the CPI increases by 6.05 percent in period 24. This magnitude of the price responses is not negligible when we allow for the volatility of commodity price shocks.

A few remarks should be made 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 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 a pass-through of 17.9 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, the result is approximately 18.2 percent, consistent with our estimated pass-through.[13] Cecchetti and Moessner (2008) also estimate the 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 the pass-through is consistent with the previous studies.

Second, our dataset includes countries where the inflation dynamics are unstable. Because these countries could substantially affect the estimated IRs as outliers, we select only countries where the standard deviation of inflation is below 10 percent.[14] Panel (b) of Figure 2 plots the estimated IRs of the CPI for the selected countries. The estimated IR function suggests that countries with large standard deviations of inflation do not substantially affect our benchmark results in panel (a).

3.2 Roles of economic factors

We next turn to the estimation results from the estimation equation with dummies of D_j^{USD}, D_j^{IT}, and D_j^{LDC} to see whether commodity price shocks continue to have a transitory effect on inflation, even after controlling for country-specific economic factors. The dummy variables specified in the previous section lead to eight combinations of IR functions, since we have three dummy variables that take zero or one. To simplify the discussion, we take four IR functions: (a) inflation-targeting developed countries (IT DCs) with flexible exchange rates; (b) non-inflation targeting developed

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

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

countries (non-IT DCs) with flexible exchange rates; (c) non-IT DCs with an exchange rate pegged to the U.S. dollar; and (d) non-IT LDCs with an exchange rate pegged to the U.S. dollar.[15] The reason why we focus on these four country groups is that we can easily isolate the impacts of the above three economic factors only by comparing the IRs. For example, differences in the IRs between country groups (a) and (b) arise from $\hat{\gamma}_{IT,j}$, those between (b) and (c) arise from $\hat{\gamma}_{USD,j}$, and so on. Although our choice of country groups does not fully cover the 120 countries, the selected country groups still account for nearly 70 percent of them (83 countries). Using Table 1, we see that group (a) includes 12 countries (e.g., Australia, Canada, and the U.K.), (b) includes 19 countries (e.g., the euro area and Japan), (c) consists of 8 countries, including the U.S., and (d) comprises 44 countries, including many Latin American, African and Asian countries and some European countries.

Figure 3 plots the IRs of the CPIs for the above country groups to a 10 percent increase in commodity price shocks. All panels of the figure suggest a transitory effect of commodity price shocks, because the slopes of the IR functions have positive steep slopes during periods 0 to 9 but become nearly flat afterward.

Comparisons across the four panels reveal that the IR functions widely vary across the country groups. In panel (a), price increases in IT DCs with flexible exchange rates are the smallest among the four panels. Panel (b) highlights the effect of IT on consumer prices. The slope of the IR function in panel (b) is steeper than that in panel (a), implying that the impact of commodity price shocks is amplified if a country does not adopt IT in its monetary policy. This result is consistent with the hypothesis that inflation expectations are well anchored by IT. Moving to panel (c), we see the effect of exchange rate variations on the IRs. If the exchange rate is pegged to the U.S. dollar, the slope of the IR function for the first 10 months becomes steeper, suggesting that the exchange rate variations dampen the inflation responses. This finding is similar to Rigobon (2010), who finds a stabilizing role of the exchange rate variations in the price response to an oil price shock. In our case, when the non-energy commodity price index excluding oil prices is used, the exchange rate has a stabilizing role on consumer prices. Finally, panel (d) shows the IR functions in LDCs, while other factors remain the same as the country group in panel (c). We observe only a small difference in the slope of the IR functions for the first 10 months between panels (c) and

[15] The remaining four IR functions are provided in the Appendix, which is available upon request.

(d). But the price responses increase, particularly in the subsequent months.

The differences in the IRs can also be investigated more comprehensively in Table 2. The table reports $\hat{\gamma}_{USD,k}$, $\hat{\gamma}_{IT,k}$, and $\hat{\gamma}_{LDC,k}$ for each forecast horizon k, along with the Newey-West standard errors below the estimates. As shown in panel (a) of the table, we confirm the stabilizing role of the exchange rate in all forecast horizons. The signs of $\hat{\gamma}_{USD,k}$ are all positive and their magnitudes are economically significant, compared to $\hat{\gamma}_k$.[16] More specifically, $\hat{\gamma}_{USD,k}$ is statistically significant for $k = 2, 3, ..., 13$, 23, and 24 at the 5 percent significance level. Hence, a stabilizing effect of the exchange rate on consumer prices can economically and statistically be significant particularly in short run, the first year after a commodity price shock.

Next, the estimates of $\gamma_{IT,k}$ are shown in the middle panel of Table 2 and are negative for all forecast horizons except for $k = 0$. In terms of the point estimate, $\hat{\gamma}_{IT,k}$ peaks in period 13 and is statistically significant for $k = 6, 7, ..., 17$. Thus, the stabilizing role of IT may not be effective in reducing temporary price increases. However, IT plays a significant role in stabilizing prices in the medium run.

Finally, the estimates of $\gamma_{LDC,k}$ are positive and become stronger as k increases. They are statistically significant in the relatively long run (i.e., $k = 13, 14, ..., 24$). In contrast to the case of $\hat{\gamma}_{USD,k}$ and $\hat{\gamma}_{IT,k}$, the effect of economic development may not strongly influence the IRs of the CPI in the first year after a commodity price shock, but the effect is larger in the longer run. This is also consistent with the interpretation that the effects of larger expenditure share on the CPI should appear in the medium and long run.

4 Evidence from the STAR model

So far, we have considered only differences across country groups but not cyclical factors in accounting for the IRs of the CPIs and inflation. Recent empirical and theoretical studies on inflation dynamics, however, demonstrate that prices may depend on the state of the economy. For example, in the literature on the exchange rate pass-through, Taylor (2000) argues that the pass-through or pricing power of firms can be influenced by the level of inflation, since high inflation increases the

[16] The magnitude of $\hat{\gamma}_k$ can be computed by dividing the IRs in panel (b) of Figure 3 by 10, where $D_j^{USD} = D_j^{IT} = D_j^{LDC} = 0$.

frequency that firms re-optimize their prices.[17] Empirically, Shintani, Terada-Hagiwara, and Yabu (2013) use the STAR model and find that the exchange rate pass-through varies depending on the past inflation rate in the U.S. economy. These studies suggest that prices and inflation may depend on the state of the economy such as the level of inflation in the economy.

In this section, we explore the possibility that IRs vary across the states of the economy. In particular, we use the logistic STAR model with the transition variable of each country's lagged inflation to see whether IRs vary between high- and low-inflation regimes. The questions we ask are: (a) Does the effect of commodity price shocks remain transitory? (b) How different are the IRs between the high- and low-inflation regimes? (c) Does the stabilizing effect of exchange rates, IT, and the degree of economic development remain present?

4.1 Estimation

To answer these questions, we extend the linear model (4) to the STAR model discussed in Teräsvirta (1994). Let $\gamma_{j,k}^{L}$ and $\gamma_{j,k}^{H}$ be the coefficient on $u_{c,t}$ for low-inflation (L) and high-inflation (H) regimes, respectively. Our estimation equation is

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

for each forecast horizon $k = 0, 1, 2, ..., K$. 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. We also allow for a difference in inflation persistence by $\beta_{i,k}^{L}$ and $\beta_{i,k}^{H}$. 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 superscript 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 superscript L are dominating. Likewise, if $\delta \to 0$, then $F(z_{j,t-d})$ converges to 1/2, which is effectively equivalent to the linear regressions (4), because parameters

[17] See, for example, Sheshinski and Weiss (1977), and Golosov and Lucas (2007), among others.

with superscript H and L can no longer be identified.

Taking $F(z_{j,t-d})$ as given, the IRs remain easy to compute 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 $IR(k,j,t) = F(z_{j,t-d})\gamma^L_{j,k} + [1 - F(z_{j,t-d})]\gamma^H_{j,k}$. Since the IR function for each period and country is a weighted sum of $IR^L(k,j)$ and $IR^H(k,j)$, we focus on $IR^L(k,j)$ and $IR^H(k,j)$ in what follows. For the transition variable $z_{j,t-d}$, we specify it as the standardized past inflation rate in country j:

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

where $\bar{\pi}_j$ and $\hat{\sigma}_j$ are the time-series average and the time-series standard deviations of the inflation rate in country j, respectively. In this specification of the transition function, we assume that the past inflation rate affects the IRs to commodity price shocks, because the inflation rate may affect firms' price setting as in the model of state-dependent pricing.

Our estimation strategy basically follows Auerbach and Gorodnichenko (2012a) who investigate whether the fiscal multipliers depend on output growth, based on the multi-country panel data. Following them, we do not include the location parameter in the function $F(\cdot)$. Instead, we standardize the transition variable so that $z_{j,t-d}$ has zero mean and a unit variance, following Auerbach and Gorodnichenko (2012a). This standardization is somewhat restrictive, but allows us to estimate (7) as a linear function if δ is fixed in (8). While Auerbach and Gorodnichenko (2012a) parameterize δ at a single value in their analysis, we allow for more flexible parameterizations of δ by a 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 best δ that minimizes the sum of squared residuals. We repeat this procedure for delay parameter d to choose a pair of (δ, d) that minimizes the sum of squared residuals.

4.2 Results from the STAR model

Table 3 reports the results of the grid search of δ and d as a preparatory analysis. Our grid search is based on the intervals $\delta \in (0, 10]$ and $d = 1, 2, ..., 6$.[18] The grid search for δ suggests that δ which

[18]We set the upper bound of δ at 10 because, when δ exceeds 10, even a small change in the standardized inflation rate generates $F(z_{t-d})$ that takes a value close to either zero or one. As a result, even if we consider a larger upper bound for δ, the estimation results are essentially unchanged.

minimizes the sum of squared residuals varies between 0.33 and its upper bound of 10.00. The grid search for d results in $d = 1$ for $k < 12$ and $d = 2$ for $k \geq 12$.

As in the previous section, we report the IRs for (a) IT DCs with flexible exchange rates, (b) non-IT DCs with flexible exchange rates, (c) non-IT DCs with an exchange rate pegged to the U.S. dollar, and (d) non-IT LDCs with an exchange rate pegged to the U.S. dollar. We report IRs under the high- and low-inflation regimes (as extreme cases), which can be measured by the estimates of $\gamma_{j,k}^{H}$ and $\gamma_{j,k}^{L}$ in (7), respectively.

Figure 4 plots the IRs of the CPI under the high- and low-inflation regimes. Panel (a) of the figure shows the IRs for IT DCs with flexible exchange rates. In this country group, we do not observe significant difference in the IRs between the two regimes. The magnitude of the IRs remains the smallest, compared to the other panels in the same figure. The results are similar even if we remove the stabilizing effect of IT in the monetary policy. In panel (b), the two IR functions are only modestly steep for the first 12 or 13 months and become nearly flat for the remaining months. Therefore, in these country groups, the effect of commodity price shocks is only transitory, even if we allow for the possibility that IRs depend on the level of the past inflation rate.

By contrast, however, when we remove the effect of exchange rate variations on the price responses, we see substantially different IRs between the high- and low-inflation regimes. Panel (c) of Figure 4 shows that, in response to a commodity price shock, the CPI under the high-inflation regime continues to increase sharply until period 8 and declines afterward. The overshooting IR function implies that the effect of commodity price shocks on inflation may not be transitory under the high-inflation regime, since the IR function does not become flat over the periods investigated. Also, the slope of the IR function under the low-inflation regime implies that inflation is low in response to the same shock, but the effect of a commodity price shock persists. The difference between the two IR functions is more evident in the case of LDCs shown in panel (d).

We can test whether the IRs are statistically the same between the high- and low-inflation regimes. More specifically, if we take an example of panel (a) (i.e., IT DCs with flexible exchange rates), the null hypothesis is that $\gamma_k^H + \gamma_{IT,k}^H = \gamma_k^L + \gamma_{IT,k}^L$. The Wald test statistics are repeatedly computed for each k. Similarly, if we test whether the IRs in panel (d) are the same across the two regimes, the null hypothesis is defined as H_0: $\gamma_k^H + \gamma_{USD,k}^H + \gamma_{LDC,k}^H = \gamma_k^L + \gamma_{USD,k}^L + \gamma_{LDC,k}^L$.

Not surprisingly, we cannot reject the null hypothesis of the same IRs in both panels (a) and

(b) at the conventional significance level in all periods, suggesting that the regime-dependent IRs of the CPIs cannot be identified in the country groups of DCs with flexible exchange rates. For panel (c), the test statistics reject the null hypothesis for only four periods $k = 6, 7, 8, 9$ at the 5 percent significance level. Therefore, for the country group of non-IT DCs with an exchange rate pegged to the U.S. dollar, the evidence for regime-dependent IRs is not necessarily strong. In contrast, we have much stronger evidence for regime-dependent IRs in panel (d). The test statistics reject the null hypothesis for $k = 1, 2, ..., 12$. Hence, at least during the first year after the shock, the IRs of the CPIs in the country group of LDCs with an exchange rate pegged to the U.S. dollar are statistically different in the high- and low-inflation regimes.

We also investigate how the roles of economic factors are affected by the inflation regimes. Tables 4 and 5 report the estimates of the coefficients on the interaction terms for the high- and low-inflation regimes, respectively. The stabilizing role of exchange rate variations is stronger and more effective in the short run (e.g., $k = 6$) under the high inflation regime than the corresponding estimates under the low-inflation regime. (See panel (a) of Table 5.) In particular, while the statistically significant stabilizing role of exchange rate variations begins to appear at $k = 18$ under the low-inflation regime, the effect under the high-inflation regime becomes significant for $k = 2, 3, ..., 12$. This may be because the exchange rate pass-through to inflation is larger under the high-inflation regime, perhaps due to the higher frequency of price changes in an inflationary environment than in a low-inflation environment.

For the role of IT, the stabilizing role is substantially weaker under the high-inflation regime than the low-inflation regime. Interestingly, the effect of IT is statistically significant only under the low-inflation regime (for $k = 7, 8, ..., 12$ and $k = 16$). To interpret this, recall that the transition variable is defined by $(\pi_{j,t-d} - \bar{\pi}_j)/\hat{\sigma}_j$. If the average inflation $\bar{\pi}_j$ is higher than the target inflation rate, a positive standardized inflation rate implies a large deviation of the inflation rate from the target, while a negative standardized inflation rate means a small deviation from the target inflation rate. Our results on the asymmetry may imply that IT does not work well if the actual inflation rate deviates substantially from the target, perhaps due to the difficulty of anchoring inflation expectations.

Finally, we also find asymmetry in the effect of the degree of economic development. The effect is large in the point estimates and statistically significant only under the high-inflation regime.

This result is reasonable if the frequency of price changes under the low-inflation regime is low and the price adjustment is still in process. As a result, the difference in the expenditure shares on commodity prices across country groups is unlikely to appear even after 24 months.

5 Conclusion

Using local projections, we estimated the IRs of the CPIs to a commodity price shock and explored the implications for inflation. We found that, in most countries, the effect of commodity price shocks on inflation is transitory. The finding is robust even when we control for exchange rate variations, adoption of IT, and the degree of economic development, which suggests that policymakers may not need to pay special attention to the recent fluctuations in commodity prices. The above factors affect the response of the level of consumer prices but the impacts on inflation are short-lived in response to commodity price shocks. Based on the STAR model that uses the past inflation rate as the transition variable, we found that commodity price shocks may have non-transitory effects on inflation, particularly in LDCs with an exchange rate pegged to the U.S. dollar. Nevertheless, commodity price shocks continue to have transitory effects in the DCs with exchange rate flexibility.

Our study focused on the responses of the CPI and inflation. Toward deeper understanding of the impacts of commodity price shocks on the economies as a whole, investigating policy responses by central banks and the effects on output would be helpful. It would call for identifying demand and supply shocks in commodity markets. Exploring these would be an interesting avenue for future research.

References

[1] Auerbach, A. and Y. Gorodnichenko, 2012a, "Fiscal Multipliers in Recession and Expansion," in A. Alesina and F. Giavazzi eds., *Fiscal Policy After the Financial Crisis*, 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, S.-S., 2009, "Oil Price Pass-Through into Inflation," *Energy Economics*, 31 (1), pp. 126-33.

[7] 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.

[8] 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.

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

[10] De Gregorio, J., Landerretche, O., and Neilson, C., 2007, "Another Pass-Through Bites the Dust? Oil Prices and Inflation," Central Bank of Chile Working Papers, No. 417.

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

[12] 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.

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

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

[15] Golosov M. and R. E. Lucas, 2007, "Menu Costs and Phillips Curve," *Journal of Political Economy*, 115 (2), pp. 171-199.

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

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

[18] 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.

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

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

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

[22] Levin, A. T., F. M. Natalucci, and J. M. Piger, 2004, "The Macroeconomic Effects of Inflation Targeting," *Federal Reserve Bank of St. Louis Review*, 86 (4), pp. 51-80.

[23] 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.

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

[25] Rigobon, R., 2010, "Commodity Prices Pass-Through," Central Bank of Chile Working Papers, No. 572.

[26] Roger S., 2009, "Inflation Targeting at 20: Achievement and Challenges," IMF Working Paper, WP/09/236.

[27] Sheshinski, E. and Y. Weiss, 1977, "Inflation and Cost of Price Adjustment," *Review of Economic Studies*, 44 (2), pp. 287-303.

[28] 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. 512-527.

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

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

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

[32] Teulings, C. N. and N. Zubanov, 2014, "Is Economic Recovery a Myth? Robust Estimation of Impulse Responses," *Journal of Applied Econometrics*, 29 (3), pp. 497-514.

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

Table 1: List of Countries

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

Notes: List of countries used in the analysis. In total, the dataset has 120 countries. Countries with an asterisk (*) are those with the exchange rate pegged to the U.S. dollar as of December 2010, countries with a dagger (†) are those that adopt inflation targeting, and countries with a section sign (§) are less-developed countries.

Table 2: The estimates of $\gamma_{USD,k}$, $\gamma_{IT,k}$, and $\gamma_{LDC,k}$ based on the local linear projection with dummy variables

(a) Estimates of $\gamma_{USD,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
0.008	0.012	0.025	0.038	0.055	0.054	0.061	0.070	0.070	0.071	0.068	0.061	0.062
(0.005)	(0.008)	(0.010)	(0.013)	(0.016)	(0.018)	(0.021)	(0.023)	(0.024)	(0.026)	(0.028)	(0.029)	(0.029)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.061	0.058	0.056	0.051	0.052	0.053	0.053	0.057	0.062	0.059	0.068	0.073	
(0.030)	(0.031)	(0.031)	(0.031)	(0.032)	(0.032)	(0.033)	(0.034)	(0.034)	(0.034)	(0.035)	(0.035)	

(b) Estimates of $\gamma_{IT,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
0.003	-0.007	-0.011	-0.012	-0.018	-0.032	-0.040	-0.048	-0.056	-0.061	-0.062	-0.061	-0.064
(0.004)	(0.006)	(0.009)	(0.012)	(0.015)	(0.017)	(0.019)	(0.021)	(0.023)	(0.025)	(0.026)	(0.027)	(0.027)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
-0.064	-0.060	-0.062	-0.060	-0.056	-0.054	-0.050	-0.042	-0.039	-0.038	-0.040	-0.034	
(0.027)	(0.028)	(0.027)	(0.028)	(0.028)	(0.028)	(0.028)	(0.028)	(0.028)	(0.028)	(0.028)	(0.029)	

(c) Estimates of $\gamma_{LDC,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
-0.001	0.004	0.008	0.009	0.015	0.026	0.029	0.027	0.030	0.035	0.042	0.041	0.043
(0.004)	(0.006)	(0.008)	(0.010)	(0.012)	(0.014)	(0.016)	(0.018)	(0.019)	(0.02)	(0.021)	(0.022)	(0.023)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.047	0.048	0.048	0.052	0.051	0.060	0.062	0.061	0.058	0.062	0.065	0.068	
(0.023)	(0.024)	(0.024)	(0.024)	(0.024)	(0.025)	(0.025)	(0.026)	(0.026)	(0.026)	(0.026)	(0.026)	

Notes: Estimated from (4). The numbers in parentheses are heteroskedasticity- and autocorrelation-consistent standard errors. Each of the dummy variables is defined as a variable that takes one (a) if a country's currency is pegged to the U.S. dollar and zero otherwise; (b) if a country adopts inflation targeting in its monetary policy and zero otherwise; and (c) if a country is less developed and zero otherwise.

Table 3: The results of a grid search for the STAR models

	$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
q	12	12	12	12	11	11	11	10	9	7	7	7	3
δ	0.320	2.192	4.816	6.848	6.800	6.752	7.344	6.000	7.696	8.384	7.312	7.856	10
d	1	1	1	1	1	1	1	1	1	1	1	1	1
	$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
q	1	1	1	1	1	1	1	1	1	1	1	1	
δ	10	10	10	10	10	10	10	10	10	10	10	10	
d	2	2	2	2	2	2	2	2	2	2	2	2	

Notes: The parameter δ in the transition function (8) is chosen from grid search over $(0, 10]$ and the delay parameter d is chosen from $d = \{1, 2, ..., 6\}$. The values of δ and d reported in the table minimize the sum of squared residuals in (7) for each grid. The lag lengths of q in (8) are selected by the BIC.

Table 4: The estimates of $\gamma^H_{USD,k}$, $\gamma^H_{IT,k}$, and $\gamma^H_{LDC,k}$ under the high-inflation regime

(a) Estimates of $\gamma^H_{USD,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
-0.02	0.013	0.037	0.058	0.073	0.080	0.091	0.111	0.117	0.110	0.109	0.111	0.098
(0.043)	(0.016)	(0.018)	(0.021)	(0.025)	(0.028)	(0.031)	(0.035)	(0.037)	(0.039)	(0.041)	(0.042)	(0.040)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.072	0.066	0.070	0.064	0.058	0.047	0.044	0.052	0.049	0.043	0.054	0.056	
(0.040)	(0.041)	(0.041)	(0.042)	(0.042)	(0.043)	(0.044)	(0.045)	(0.045)	(0.046)	(0.047)	(0.047)	

(b) Estimates of $\gamma^H_{IT,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
-0.016	-0.016	-0.010	-0.010	-0.011	-0.020	-0.030	-0.032	-0.043	-0.043	-0.042	-0.038	-0.055
(0.033)	(0.014)	(0.018)	(0.021)	(0.024)	(0.027)	(0.030)	(0.033)	(0.034)	(0.036)	(0.038)	(0.038)	(0.037)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
-0.061	-0.059	-0.054	-0.046	-0.044	-0.047	-0.043	-0.029	-0.021	-0.021	-0.018	-0.013	
(0.034)	(0.034)	(0.035)	(0.035)	(0.036)	(0.036)	(0.036)	(0.037)	(0.037)	(0.038)	(0.038)	(0.039)	

(c) Estimates of $\gamma^H_{LDC,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
0.052	0.023	0.025	0.025	0.037	0.052	0.059	0.056	0.052	0.057	0.066	0.064	0.062
(0.035)	(0.013)	(0.015)	(0.017)	(0.020)	(0.022)	(0.024)	(0.027)	(0.029)	(0.030)	(0.031)	(0.032)	(0.031)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.041	0.051	0.055	0.055	0.058	0.075	0.086	0.089	0.091	0.100	0.104	0.110	
(0.031)	(0.032)	(0.032)	(0.032)	(0.033)	(0.033)	(0.033)	(0.034)	(0.035)	(0.035)	(0.036)	(0.036)	

Notes: Estimated from (7). The model is estimated with the STAR model. See the footnote of Table 2 for the other detail.

Table 5: The estimates of $\gamma^L_{USD,k}$, $\gamma^L_{IT,k}$, and $\gamma^L_{LDC,k}$ under the low-inflation regime

(a) Estimates of $\gamma^L_{USD,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
0.04	0.014	0.014	0.017	0.035	0.026	0.027	0.023	0.018	0.023	0.020	0.004	0.017
(0.043)	(0.016)	(0.015)	(0.017)	(0.022)	(0.024)	(0.028)	(0.031)	(0.032)	(0.034)	(0.037)	(0.039)	(0.037)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.051	0.052	0.042	0.039	0.049	0.066	0.071	0.070	0.085	0.088	0.093	0.104	
(0.040)	(0.041)	(0.042)	(0.043)	(0.044)	(0.045)	(0.045)	(0.046)	(0.047)	(0.048)	(0.049)	(0.051)	

(b) Estimates of $\gamma^L_{IT,k}$

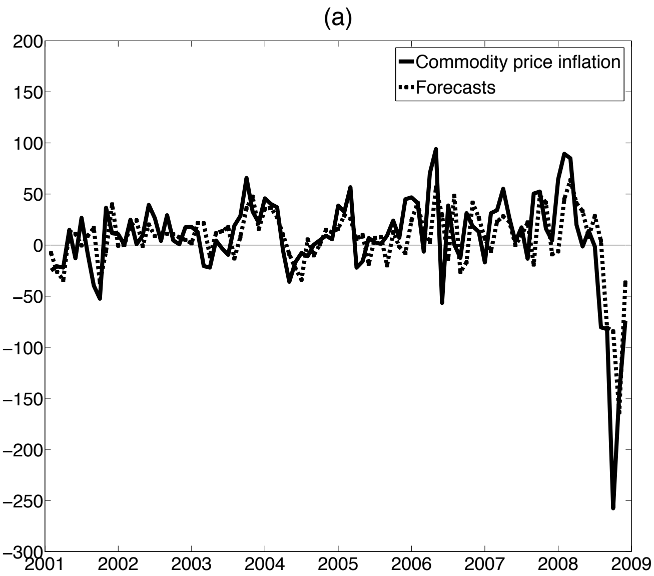
$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
0.024	0.006	-0.009	-0.009	-0.017	-0.032	-0.038	-0.053	-0.057	-0.065	-0.070	-0.075	-0.067
(0.031)	(0.011)	(0.011)	(0.014)	(0.018)	(0.020)	(0.024)	(0.026)	(0.028)	(0.030)	(0.032)	(0.033)	(0.033)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
-0.069	-0.060	-0.069	-0.077	-0.068	-0.059	-0.052	-0.050	-0.052	-0.051	-0.059	-0.051	
(0.037)	(0.039)	(0.038)	(0.038)	(0.038)	(0.038)	(0.039)	(0.039)	(0.039)	(0.040)	(0.040)	(0.041)	

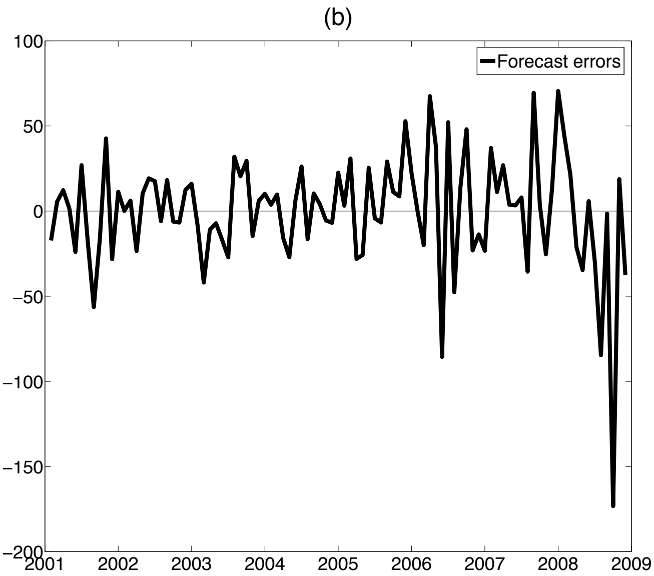
(c) Estimates of $\gamma^L_{LDC,k}$

$k=0$	$k=1$	$k=2$	$k=3$	$k=4$	$k=5$	$k=6$	$k=7$	$k=8$	$k=9$	$k=10$	$k=11$	$k=12$
-0.056	-0.016	-0.009	-0.006	-0.003	0.005	0.005	0.003	0.013	0.015	0.019	0.021	0.030
(0.034)	(0.013)	(0.013)	(0.014)	(0.016)	(0.018)	(0.021)	(0.024)	(0.024)	(0.026)	(0.028)	(0.030)	(0.029)
$k=13$	$k=14$	$k=15$	$k=16$	$k=17$	$k=18$	$k=19$	$k=20$	$k=21$	$k=22$	$k=23$	$k=24$	
0.058	0.047	0.043	0.054	0.046	0.043	0.035	0.029	0.021	0.023	0.026	0.023	
(0.030)	(0.031)	(0.031)	(0.032)	(0.033)	(0.033)	(0.034)	(0.035)	(0.035)	(0.036)	(0.037)	(0.038)	

Notes: Estimated from (7). The model is estimated with the STAR model. See the footnote of Table 2 for the other detail.

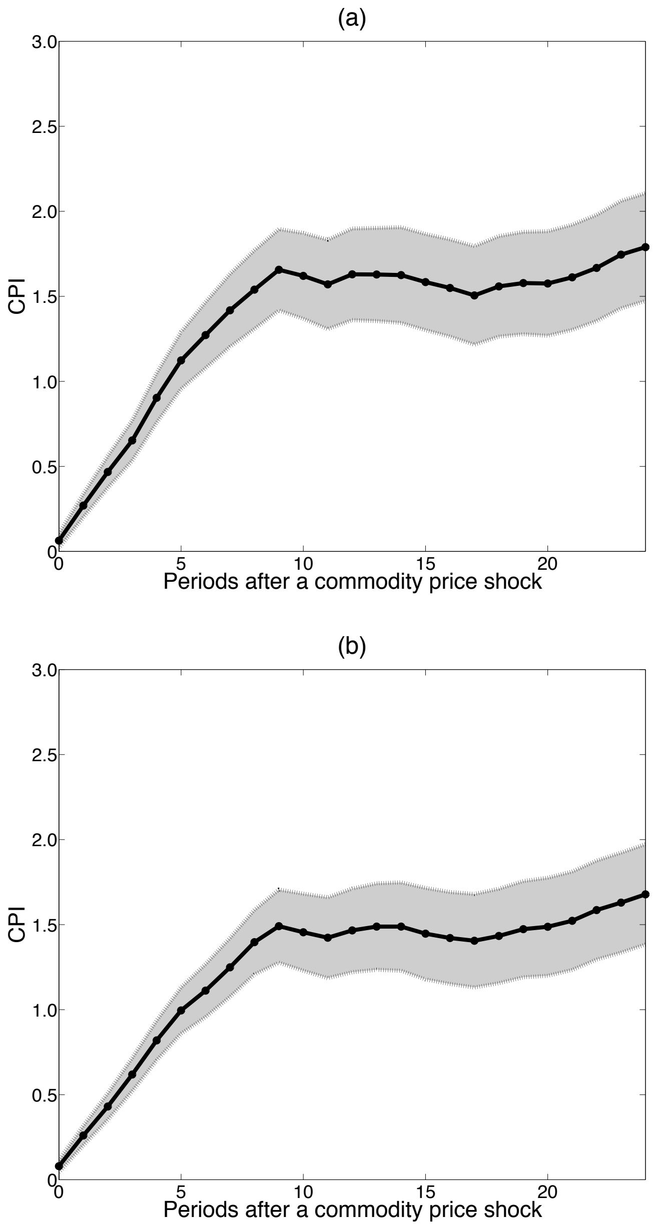
Figure 1: Commodity price inflation and the estimated commodity price shocks





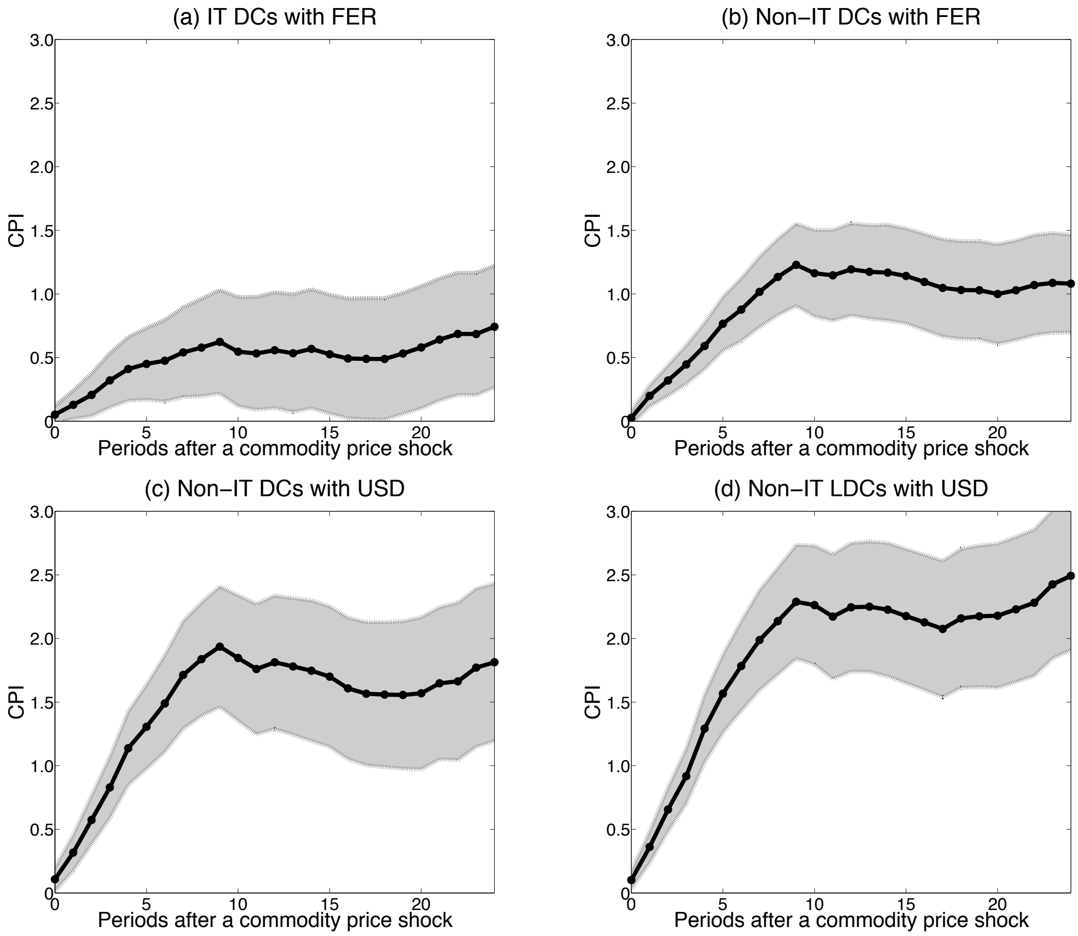
Notes: Panel (a) shows commodity price inflation and its forecasts based on (6). Panel (b) plots the residuals in (6).

Figure 2: Impulse responses of the CPI: Benchmark regressions



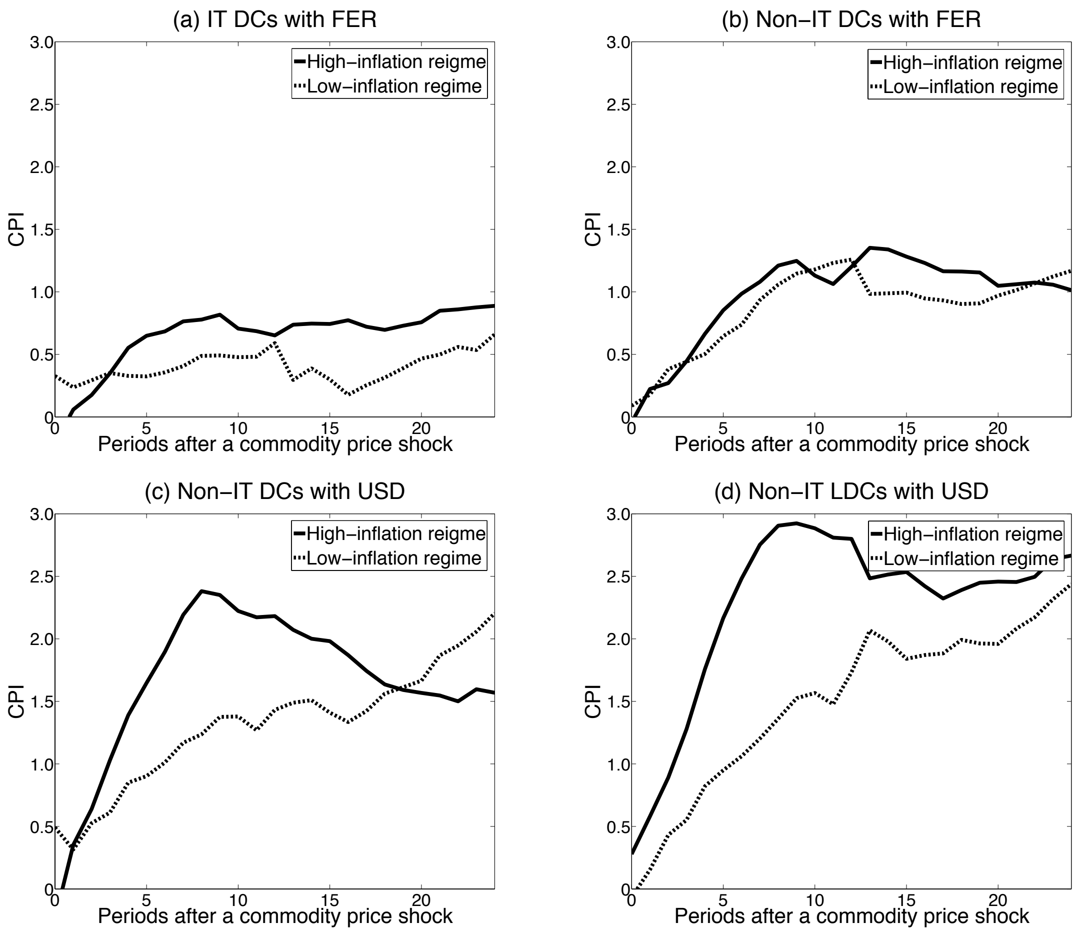
Notes: The panels plot impulse responses of the CPI to a 10 percent increase in commodity price shocks. The impulse responses are estimated from (2). The unit of the responses is percent. Panel (a) shows the CPI responses estimated from the 120-country panel. Panel (b) shows those estimated from the 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 interaction terms



Notes: The panels plot impulse responses of the CPI to a 10 percent increase in commodity price shocks. The impulse responses are estimated from (4). The unit of the responses is percent. Panels (a) and (b) show the responses of the CPI in inflation-targeting developed countries with flexible exchange rates (IT DCs with FER) and non-inflation-targeting developed countries with flexible exchange rates (Non-IT DCs with FER), respectively. Panels (c) and (d) correspond to non-inflation targeting developed countries with an exchange rate pegged to the U.S. dollar (Non-IT DCs with USD) and non-inflation targeting less developed countries with an exchange rate pegged to the U.S. dollar (Non-IT LDCs with USD), respectively. The shaded areas represent the 95 percent confidence intervals.

Figure 4: Impulse responses of the CPI with interaction terms: High- and low-inflation regimes



Notes: The panels plot impulse responses of the CPI to a 10 percent increase in commodity price shocks under the high- and low-inflation regimes. The impulse responses are estimated from (7). The unit of the responses is percent. Panels (a) and (b) show the responses of the CPI in inflation-targeting developed countries with flexible exchange rates (IT DCs with FER) and non-inflation-targeting developed countries with flexible exchange rates (Non-IT DCs with FER), respectively. Panels (c) and (d) correspond to non-inflation targeting developed countries with an exchange rate pegged to the U.S. dollar (Non-IT DCs with USD) and non-inflation targeting less developed countries with an exchange rate pegged to the U.S. dollar (Non-IT LDCs with USD), respectively.