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# Not All Exchange Rate Movements Are Alike: Exchange Rate Persistence and Pass-Through to Consumer Prices

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**Abstract**

This study develops a framework to identify persistent and transitory shocks in exchange-rate movements and to estimate the shock-specific exchange-rate pass-through to domestic prices. The framework combines a dataset of a long time series of exchange-rate forecasts since the 1980s with a range restriction that is a natural generalization of the standard sign restriction. The empirical results show that exchange rate pass-through is higher when a persistent shock dominates exchange-rate movements. The composition of persistent and transitory shocks varies over time. This study asserts that time variations of exchange rate pass-through are at least partly attributable to differences in shock-specific pass-through rates and variations in the composition of shocks over time. Applying our identification procedure to disaggregated prices of the CPI, we also find that a correlation between pass-through coefficients and frequencies of price adjustments is shock dependent. Specifically, the positive correlation, which is reported in Gopinath and Itskhoki [2010], disappears, when exchange-rate movements are transitory.

*JEL classification*: F41, C32
*Keywords*: exchange-rate pass-through; exchange-rate persistence; range restriction; survey expectation;

# 1 Introduction

Since policy rates have reached an effective lower bound in many advanced economies, the exchange-rate channel of monetary policy transmission mechanisms has become important. For an exchange-rate's expenditure switching mechanism, the key parameter is the responsiveness of domestic prices to exchange-rate changes, i.e., "the exchange rate pass-through."

When estimating this pass-through coefficient[1], conventional studies have simply regressed domestic prices on exchange rates with relevant covariates. This empirical strategy, which Forbes et al. [2015] call the "rules-of-thumb" pass-through measurement, is

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[1]Numerous studies on exchange rate pass-through have been conducted, including the early theoretical works of Dornbusch [1987], Krugman [1987], Giovannini [1988], and Froot and Klemperer [1989], the cross-country empirical comparisons of Campa and Goldberg [2005], Goldberg and Campa [2010], and Ito

effective only when the exchange rate is driven by a single exogenous factor or when domestic-price responses are indifferent, irrespective of the drivers behind exchange-rate movements. However, not all exchange-rate movements are alike. Similar movements in exchange rates are regarded as persistent at certain times and as noisy fluctuations at other times. Forward-looking firms may change the pass-through behavior depending on the shock characteristics behind exchange-rate movements. For example, in case of a persistent depreciation, a firm with nominal frictions, such as menu costs of price adjustments, is likely to raise its price, anticipating a persistent increase in imported-goods prices for intermediate inputs and hence also in marginal costs. By contrast, in a transitory depreciation, the firm may hesitate to raise its price because it is not worth the cost of a price adjustment. This example suggests that an exchange-rate pass-through could be shock-specific and calls for an identification of the drivers behind exchange-rate movements when estimating a pass-through coefficient.

This study's contribution is to present a new framework for estimating a shock-specific pass-through. Our framework identifies transitory and persistent shocks in exchange-rate movements separately, combining a dataset of a long time series of yen-dollar exchange-rate forecasts since 1985, with a range restriction that is a natural extension of Uhlig [2005]'s sign restriction.[2] Intuitively, we regard a perceived shock as persistent if the shock moves the spot exchange rate and the exchange-rate forecast in the same direction and in a similar magnitude. Similarly, we regard a perceived shock as transitory if the shock moves the spot exchange rate but has little impact on the exchange-rate forecast. Since this framework requires data to include a plenty of episodes of persistent and transitory shocks, we use long-term historical data beginning in the 1980s.

The main message of this study is that the pass-through coefficient can differ, depending on shock characteristics that influence exchange-rate movements. An empirical assessment of Japan's data shows that the exchange rate pass-through is higher when persistent shocks dominate exchange-rate movements. The composition of persistent and transitory shocks varies over short periods of time. This study finds that time variations of exchange-rate pass-through are at least partly caused by the differences in shock-specific pass-through rates and the variations in shock composition. The findings suggest that time variations in exchange-rate pass-through is caused not only by slow-moving structural changes, such as composition of imports (Campa and Goldberg [2005]), monetary policy framework (Gagnon and Ihrig [2004]), trade integration (Gust, Leduc, and Vigfusson [2010]), and production structure (Shioji [2015]), but also by fast-moving changes in the composition of persistent and transitory shocks. This finding is valuable, especially for policymakers who make decision over different phases of the business cycles. In addition, by applying our identification procedure to disaggregated prices of the CPI, we find

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and Sato [2008], and the micro-data studies on the source of incomplete pass-through such as Goldberg and Hellerstein [2008], Gopinath and Itskhoki [2010], Gopinath, Itskhoki, and Rigobon [2010], and Nakamura and Zerom [2010]. More recent studies on exchange rate pass-through have begun focusing on time variations in the short-run (Shambaugh [2008], Forbes, Hjortsoe, and Nenova [2015], and Berger and Vavra [2013]). Burstein and Gopinath [2014] provide a comprehensive survey of prior studies on the exchange rate pass-through.

[2] This unique dataset has been collected by the Japan Center for International Finance (JCIF) since May 1985. See later sections for details about the JCIF survey.

that a correlation between pass-through coefficients and frequencies of price adjustments is shock dependent. The positive correlation, which is reported in Gopinath and Itskhoki [2010], disappears, when exchange-rate movements are transitory.

Since the earliest stage of the literature on this topic, theoretical studies have focused on the persistence of exchange-rate movements. Giovannini [1988] examines a firm's dynamic pricing behavior and finds that persistent exchange-rate variations lead to higher pass-through. Burstein and Gopinath [2014] exposit it using a New Keynesian sticky price model. Froot and Klemperer [1989] and Ravn, Schmitt-Grohe, and Uribe [2010] reach a similar conclusion, using a dynamic customer market model in partial and general equilibrium settings. Other studies such as Krugman [1987], Baldwin [1988], and Dixit [1989] posit that exchange-rate pass-through could be nearly zero when exchange-rate movements are small and transitory, when sunk costs exist for market entry and exit.

In contrast to theoretical studies, a limited number of empirical studies have tried to identify underlying drivers behind exchange-rate movements when estimating pass-through to domestic prices.[3] Shambaugh [2008] and Forbes et al. [2015] are few exceptions.[4] Specifically, Shambaugh [2008] and Forbes et al. [2015] identify structural shocks, such as monetary policy shocks, technology shocks, or demand shocks, employing a structural vector autoregression (SVAR) with long-run and other identification restrictions.[5] They find that pass-through can be shock-specific and that changes in the composition of shocks are the source of short-run time variations in exchange-rate pass-through.

Although our study is in line with these preceding studies that estimate shock-specific pass-through, it has a different focus.[6] Specifically, our study clarifies that characteristics of shocks are important, as well as structural sources of shocks, as noted in earlier theoretical studies. For empirical reasons, it is important to focus on shock characteristics because the principal driver of nominal exchange rates may not be structural shocks. For example, Lubik and Schorfheide [2006] report that more than 90 percent of nominal exchange-rate movements are attributable to a non-structural PPP shock that is designed to capture deviations of the model from the data, based on an estimated open-economy DSGE model. The identifying restrictions derived from a structural model, which preceding studies rely on, may be misleading.

The identification procedure developed in this study has several technical advantages over existing analyses of shock-specific pass-through. First, it is more robust regarding

[3]Klein [1990] points out the potential need for shock identification in estimating exchange rate pass-through. Some studies such as Ito and Sato [2008], An and Wang [2012], Shioji [2012], and Shioji [2015] estimate exchange rate pass-through, identifying exogenous variations in exchange rates with an (S)VAR model but do not distinguish the characteristics of shocks behind exchange-rate fluctuations. The analysis in this study intends to go one step further to these previous studies by making a distinction between persistence and transitory shocks.

[4]Corsetti, Dedola, and Leduc [2008], using hypothetical data generated by an open-economy DSGE model, show that shock identification is necessary to estimate the true pass-through coefficient.

[5]Shambaugh [2008] relies on long-run restrictions and Forbes et al. [2015] use a combination of short-run, long-run, and sign restrictions. In both studies, identification assumptions are derived from an open economy macroeconomic model.

[6]Amstad and Fischer [2010] identify the pass-through coefficient using an event study approach, which implicitly assumes that the coefficient can vary over time.

model misspecification. Previous studies specified a full model structure to obtain ideas about identifying restrictions. In contrast, our procedure is model free, and therefore, does not necessarily rely on a specific model structure. Thus, our approach is minimalistic and only imposes range restrictions on instantaneous responses of a few variables. Second, our procedure does not rely on long-run restrictions that demand strict stationarity of the data. Therefore, information loss from over-differentiation, suggested in Sims, Stock, and Watson [1990], can be avoided. Considering the above properties, this procedure has the potential to be a useful toolkit for policy practitioners who demand a robust and handy analytical framework. Persistent and transitory shocks are easily identified, using the spot rate and forecast data that are available in high frequencies.

The remainder of this article is organized as follows: Section 2 lays out a sticky price model to illustrate firms' pricing behavior in the presence of persistent exchange-rate shocks. Section 3 discusses the identification strategy of persistent and transitory shocks behind exchange-rate movements and shock-specific pass-through. Section 4 explains Japan Center for International Finance (JCIF) data and other data used for the empirical exercise. Section 5 presents the major results of shock-specific pass-through, as well as time variations in aggregated pass-through. Section 6 provides a robustness analysis and Section 7 is the conclusion.

# 2 Simple Model of Firms' Dynamic Pricing

In this section, a dynamic model of infrequent price adjustments illustrates how persistent exchange-rate movements affect pass-through behaviors. Although a dynamic pricing firm can be modeled using several different setups, such as time-dependent pricing with infrequent arrivals of price-reset probabilities (Calvo [1983]) and state-dependent pricing with an existence of price-adjustment cost (Dotsey, King, and Wolman [1999] and Golosov and Lucas [2007]), we adopt a partial-equilibrium Calvo pricing model just for analytical simplicity.

The model is characterized by unit mass monopolistic producers of differentiated goods, $C_t(k)$, which are compiled in a CES aggregator and sold as a final good: $C_t = [\int_0^1 C_t(k)^{(\theta-1)/\theta} dk]^{\theta/(\theta-1)}$. A firm $k$ resets its price when an exogenous price-reset signal arrives as in Calvo [1983]. The price-adjusting firm's profit maximization is denoted as follows.

$$\max_{p_t(k)} \sum_{j=0}^{\infty} (\alpha\beta)^j \, \mathbb{E}_t \Pi_{t,t+j}(k), \tag{1}$$

where $\mathbb{E}_t$ is an expectation operator, $\Pi_{t,t+j}(k)$ is a firm $k$'s profit at $t + j$ keeping the reset price at time $t$, $p_t(k)$, unchanged, $0 < \alpha < 1$ is an exogenous price non-adjustment probability, and $0 < \beta \leq 1$ is a discount factor. The log-linearized aggregate price level is derived as follows:

$$p_t = \underbrace{(1-\alpha)(1-\alpha\beta)\sum_{j=0}^{\infty}(\alpha\beta)^j\,\mathbb{E}_t mc_{t+j}}_{\text{the optimal reset price}} + \alpha \cdot \underbrace{p_{t-1}}_{\text{non-reset prices}}, \tag{2}$$

where $mc_t$ is a log deviation of marginal cost from its steady state value. A firm combines labor inputs and foreign-intermediate inputs with Cobb-Douglas technology to produce a differentiated good. With a common technology and common factor markets, marginal costs are identical for all firms. The marginal cost in a logarithm can be approximated around a steady state as

$$mc_t = (1-\gamma)\,w_t + \gamma \cdot m_t, \tag{3}$$

where $w_t$ and $m_t$ denote log deviations of the price of domestic labor inputs and foreign intermediate inputs from their respective steady states. Parameter $0 \leq \gamma < 1$ represents the share of foreign inputs in the cost of production. Following Gopinath and Itskhoki [2010], we postulate that wages are not sensitive to exchange-rate changes and that the main driver of $m_t$ is the logarithm of exchange rate $s_t$: $m_t = s_t$. Finally, we also presume that the process of $s_t$ is described as an autoregressive of order one, AR(1), process: $s_{t+1} = \rho s_t + \epsilon_{t+1}$ in which $0 < \rho < 1$ and $\epsilon_t \sim N(0, \sigma^2)$.

Then, if $w_t$ is constant for all $t$, an exchange rate pass-through at $t+1$ is:

$$\frac{\mathbb{E}_t\left[p_{t+1} \mid \epsilon_{t+1} = 1\right] - \mathbb{E}_t\left[p_{t+1}\right]}{\mathbb{E}_t\left[s_{t+1} \mid \epsilon_{t+1} = 1\right] - \mathbb{E}_t\left[s_{t+1}\right]} = \gamma\frac{(1-\alpha)(1-\alpha\beta)}{1-\alpha\beta\rho}. \tag{4}$$

(4) suggests that the persistence of exchange rates, which is denoted as $\rho$, is one of the key parameters for the responsiveness of domestic prices to exchange rate movements. This feature comes from the dynamic aspect of the firms' pricing behavior. (4) implies that once nominal rigidities disappear ($\alpha = 0$), the firms' pricing problem reduces to a static one and the persistence parameter does not affect the impulse response.

# 3 Identification Strategy

## 3.1 Persistence in exchange-rate movements

This section describes how to identify persistence in exchange-rate movements using an SVAR model with a range restriction. Here, we identify an ex ante perception of persistence in exchange rates, not ex post persistence. Standard procedures of ex post trend-cycle decomposition, such as in Beveridge and Nelson [1981], are not suitable for our purpose. Thus, we develop an alternative methodology by utilizing spot exchange rates and exchange-rate forecasts.

First, apart from a simple AR(1) specification in the previous section, we explicitly model persistent and transitory shocks as drivers of exchange-rate movements. Specifically, we presume that the (log) exchange rate is a linear combination of a persistent component $s_t^P$ and a transitory component $s_t^T$ with both components following the autoregressive processes of (5) and (6),[7]

$$\begin{aligned} s_t &= s_t^P + s_t^T, \\ s_{t+1}^P &= \rho_P s_t^P + \epsilon_{t+1}^P, \qquad (5) \\ s_{t+1}^T &= \rho_T s_t^T + \epsilon_{t+1}^T, \qquad (6) \end{aligned}$$

where $0 \leq \rho_T < \rho_P \leq 1$; $\epsilon_t^P$ and $\epsilon_t^T$ are mean zero serially uncorrelated fundamental shocks. Given a realization of each shock, conditional forecasts of n-period ahead exchange rates are denoted as follows:

$$\mathbb{E}_t[s_{t+n} \mid \epsilon_{t+1}^P] - \mathbb{E}_t[s_{t+n}] = (\rho_P)^n \, \epsilon_{t+1}^P, \tag{7}$$
$$\mathbb{E}_t[s_{t+n} \mid \epsilon_{t+1}^T] - \mathbb{E}_t[s_{t+n}] = (\rho_T)^n \, \epsilon_{t+1}^T. \tag{8}$$

Since an exchange rate is generally described as an I(1) process with noisy fluctuations (Baillie and Bollerslev [1989]), $\rho_P$ is approximately one and $\rho_T$ is approximately zero. Then, (7) and (8) suggest that a persistent shock affects the exchange-rate forecast and the spot exchange rate to a similar amount, as long as $n$ is finite and not so large, whereas a transitory shock has a negligible impact on the exchange-rate forecast.

## 3.2 Identification procedure

To recover these shocks from observable data and to estimate shock-specific pass-through to domestic prices, we formulate a system of three variables, $\boldsymbol{Y_t} = [s_{t,t+n}^e, s_t, p_t]'$, which consists of an n-period ahead forecast of exchange rate $s_{t,t+n}^e$, a spot exchange rate $s_t$, and a domestic price index $p_t$:

$$\boldsymbol{B_0 Y_t} = \sum_{l=1}^{L} \boldsymbol{B_{1,l} Y_{t-l}} + \boldsymbol{\epsilon}_t, \tag{9}$$

where $\boldsymbol{\epsilon_t} = [\epsilon_t^P, \epsilon_t^T, \epsilon_t^O]'$, $\boldsymbol{B_{1,l}}$ is the parameter matrix, and $\boldsymbol{B_0}$ is the nonsingular matrix of contemporaneous interactions between the model's variables. The covariance matrix of the fundamental shocks is normalized, such that $\mathbb{E}(\boldsymbol{\epsilon_t \epsilon_t}') \equiv \boldsymbol{\Sigma_\epsilon} = \boldsymbol{I}_{[3\times3]}$. We write the relationship between residuals of a reduced-form VAR model and fundamental shocks as $\boldsymbol{u_t} = \boldsymbol{B_0^{-1} \epsilon_t}$ where $\boldsymbol{u_t} = [u_t^{s^e}, u_t^s, u_t^p]'$.

(7) and (8) motivate us to assign a range restriction on each element of $\boldsymbol{B_0^{-1}}$ as in (10):

$$\boldsymbol{B_0^{-1}} \equiv \begin{bmatrix} b_{11} & b_{12} & * \\ b_{21} & b_{22} & * \\ b_{31} & * & * \end{bmatrix}, \tag{10}$$

where asterisks denote unrestricted elements and the the other elements suffice the following: $b_{11} \geq \phi_1 \cdot b_{21}$, $b_{12} \leq \phi_2 \cdot b_{22}$, $b_{21} > 0$, $b_{22} > 0$, and $b_{31} > 0$. We also presume that $\phi_{i\in[1,2]}$ are arbitrary constants that suffice $0 \leq \phi_2 \leq \phi_1$. The range restrictions on $b_{11}$

[7] AR(1) specifications of (5) and (6) are just for simplification.

and $b_{12}$ are generalizations of Uhlig [2005], because (10) is reduced to a standard sign restriction when $\phi_1$ and $\phi_2$ are set to zero.

If appropriately calibrated, (10) classifies a shock as (a) persistent when the shock affects the spot exchange rate and the exchange-rate forecast at the same time and for a similar amount, and (b) transitory when the shock affects the spot exchange rate but does not have as much of an impact on the exchange-rate forecast. To facilitate the identification, we also assume that a persistent depreciation (appreciation) shock affects domestic prices positively (negatively), while a domestic price response to a transitory shock remains unrestricted. In the benchmark estimation, we set $\phi_1 = \phi_2 \equiv \phi = 0.9^n$, which implies that a monthly auto-regressive coefficient of a persistent shock is 0.9 or higher. In our later empirical analysis, we use six-month forecast data ($n = 6$).

By using private agents' forecast, our empirical procedure only postulates a simple, short-run restriction, even though it intends to estimate whether or not current exchange-rate movements are perceived persisting in the future. One advantage is that the procedure can avoid relying on long-run restrictions first developed by Blanchard and Quah [1989], which require strict stationarity of the data (even in a VAR framework[8]) and has been criticized by Chari, Kehoe, and McGrattan [2007] for its strong assumption that imposes an infinite, zero impulse response in the long run.

To estimate the model, we employ a Bayesian approach. Following Kadiyala and Karlsson [1997] and Sims and Zha [1998], a prior distribution is a Normal Inverted Wishart:

$$p\left(vec\left(\bar{\boldsymbol{A}}\right) \mid \boldsymbol{\Sigma_u}\right) \sim N\left(vec\left(\boldsymbol{A^0}\right), \boldsymbol{\Sigma_u} \otimes \boldsymbol{\Psi^0}\right), \quad p\left(\boldsymbol{\Sigma_u}\right) \sim IW\left(\boldsymbol{\Sigma_u^0}, \boldsymbol{\tau^0}\right), \tag{11}$$

where $\bar{\boldsymbol{A}}$ is a coefficient matrix of a reduced-form VAR of $\boldsymbol{Y_t} = \sum_l^L \boldsymbol{A_l Y_{t-l}} + \boldsymbol{u_t}$, the prior parameters $\boldsymbol{A^0}$, $\boldsymbol{\Psi^0}$, $\boldsymbol{\Sigma_u^0}$, and $\boldsymbol{\tau^0}$ are chosen so that the prior expectations and variances of $\bar{\boldsymbol{A}}$ and $\boldsymbol{\Sigma_u}$ coincide with the generalized Minnesota prior discussed in Mumtaz and Zanetti [2012].[9] In the implementation, we impose the prior by adding dummy observations to the data, as is common in the literature (e.g. Banbura, Giannone, and Reichlin [2010]).[10]

By implementing a range restriction, we basically follow Algorithm 2 in Rubio-Ramirez, Waggoner, and Zha [2010], which is a standard algorithm for a sign restriction. Specifically, denoting an eigenvalue decomposition of the covariance matrix of a VAR as $\boldsymbol{\Sigma_u = PDP'}$, we define $\boldsymbol{\tilde{B}_0^{-1} = PD^{1/2}}$. Then, we draw a $3 \times 3$ matrix from an independent, normal distribution and compute its QR decomposition to create an orthogonal impulse matrix, $\boldsymbol{Q}$. Following Uhlig [2005], we obtain a structural impact matrix as $\boldsymbol{B_0^{-1} = \tilde{B}_0^{-1} Q'}$. Finally, we retain $\boldsymbol{B_0^{-1}}$ only when it satisfies (10). We repeat the above sequence until we acquire a certain number of $\boldsymbol{B_0^{-1}}$ for each Gibbs sampling. Our analyses of impulse responses and historical decompositions are based on the $\boldsymbol{B_0^{-1}}$ matrix that is in a closest

[8] As stressed in Sims et al. [1990], the over differentiation may miss important information in the data. A level VAR can avoid this over-differentiation problem. As long as our interest remains on impulse responses, a level VAR can generate a consistent estimate.

[9] Different from Litterman [1986]'s Minnesota prior, (11) does not presume a diagonal, fixed, and known covariance matrix.

[10] We set a loose prior on VAR coefficients by assigning a large hyperparameter value of 10, which divides the prior mean for each VAR coefficient in the dummy observations.

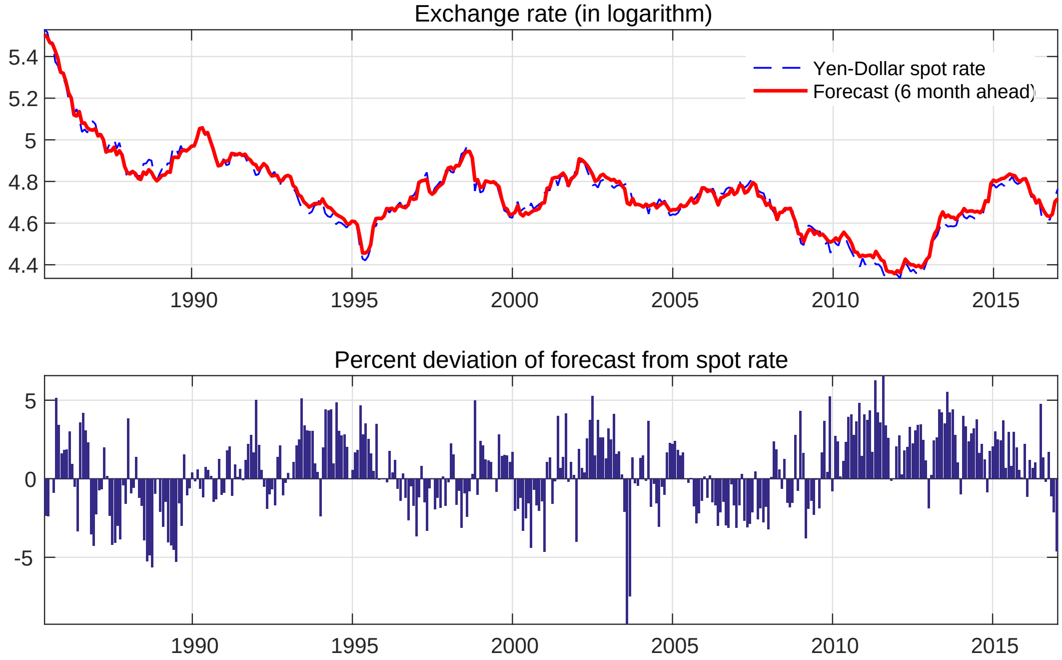
distance from the median of the estimated distribution of $\boldsymbol{B}_0^{-1}$ for each draw from the posterior.

Our procedure is different from the standard sign restriction in that the specific value of a restriction could be different at each draw. This is because $b_{1i}$, which is the threshold value of the range restriction on a six-month forecast, could vary with a draw of $(2, i)$ element of $\boldsymbol{B}_0^{-1}$ for $i = 1, 2$. A simple Monte Carlo experiment in Appendix A shows that our identification procedure is successful in identifying persistent and transitory shocks in artificial time-series data.

# 4 Data

Our empirical strategy requires a long time series of exchange-rate forecasts. The JCIF provides a unique dataset of market experts' exchange-rate forecasts. It has conducted telephone surveys twice a month (in the middle and at the end of the month) since May, 1985.[11] Respondents, consisting of exchange-rate-market participants at approximately 50 companies, (including banks, brokers, securities companies, life insurance companies, trading companies, exporters and importers), provide point forecasts of yen-dollar exchange rate for the one, three, and six-month horizons. The price index, which is our primary interest, is released once a month, and we use the end-of-the-month survey as a forecast of the month. Our estimation uses a mean of individual respondents' 6-month forecasts as an aggregate forecast.

Figure 1: JCIF market experts' forecasts on exchange rate




[11]Ito [1990] studies the expectation formation in foreign exchange-rate markets using JCIF surveys.

The upper panel of Figure 1 depicts the JCIF forecast and the spot yen-dollar exchange rate. It shows that these two series move closely, suggesting that the spot rate strongly affects professional forecasts. However, it is rare that professional forecasts match the spot rate perfectly. The percentage deviation of forecasts from the spot rate in the lower panel of Figure 1 illustrates that there are significant differences between these two series for almost the entire period. We will use these deviations to identify persistent and transitory shocks.

The JCIF survey provides market-participants' forecasts only for yen-dollar rates in a long historical time series. An exchange rate in this study is, therefore, a spot yen-dollar rate consistent with the JCIF survey's expectations. It may be more appropriate to use the effective exchange rate that is a trade-weighted average of multiple exchange rates. However, the bias that stems from using the yen-dollar rate instead of the effective exchange rate is relatively small, because recent empirical works using granular data on prices[12] find that prices are rigid for significant durations in the invoiced currency and the majority of trade is invoiced in very few currencies. As for Japan, approximately 70 percent of imports are invoiced in U.S. dollars, and 25 percent of imports are invoiced in Japanese yen.[13]

Regarding other data, a domestic price in the Japan's consumer price index excludes fresh foods. We also use a real crude oil price index, denominated by the U.S. PCE deflator, as in Kilian [2009], and a monthly index of industrial production as covariates.[14]

The sample period is January, 1986 to December, 2016.[15] The model is constructed by 12 lags of logged endogenous variables. The number of draws in Gibbs sampling is 20,000 and the first 15,000 of the draws are discarded as burn-ins. The number of second-stage draws of $\boldsymbol{B}_{\mathbf{0}}^{-\mathbf{1}}$ is 1,000.

# 5 Results

## 5.1 Shock-specific exchange-rate pass-through

The identification procedure developed in the above section allows us to estimate impulse responses to persistent and transitory shocks, distinctively. The resulting responses in Figure 2 show that the degrees of exchange rate pass-through to the consumer price index are different, depending on the characteristics of underlying shocks.[16]

[12]e.g. Goldberg and Tille [2008], Gopinath [2015], Gopinath and Rigobon [2008], and Fitzgerald and Haller [2014]

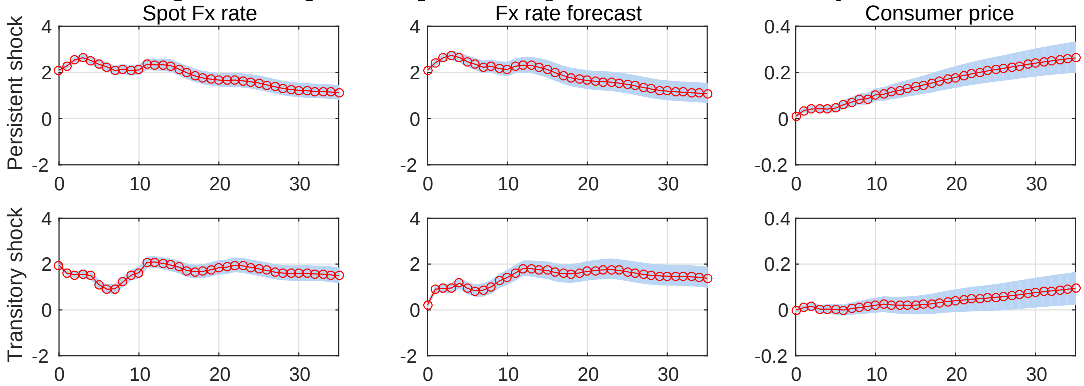
[13]The Ministry of Finance started to publish invoiced currencies for imports and exports in the second half of the fiscal 2000. Its figures show that 70.7 percent and 23.5 percent of imports in 2000 were invoiced in U.S. dollar and Japanese yen, respectively. Import shares in U.S. dollars and Japanese yen are similar in the latest data: 66.9 percent and 26.1 percent of imports were invoiced in their respective currencies in the first half of fiscal 2016.

[14]See Appendix B for details of data sources.

[15]In the first several rounds, the JCIF made some modification in survey methodology. Thus, we set the sample beginning at January, 1986.

[16]Initial responses of spot exchange rates and exchange-rate forecasts in Figure 2 differ for persistent shocks versus transitory shocks. This reflects our identification assumption. The upper panels in Figure 2 show that a persistent shock has a similar impact on the spot exchange rate and the six-month exchange-

Figure 2: Impulse responses to persistent and transitory shocks




Note: The red lines and the shaded areas depict the medians and interquartile ranges of distributions, respectively.

Although initial responses of the spot exchange rate are similar for both types of shocks, consumer-price responses are different. Consumer prices increase up to 0.25 percent after three years, to a one standard deviation persistent shock. In contrast, consumer prices increase 0.1 percent or less, to a one standard deviation transitory shock. These results suggest that the impulse response of consumer prices is shock-specific and larger for a more persistent underlying shock. However, the comparison of impulse responses in Figure 2 does not show how rigorously exchange-rate movements pass through to consumer prices, because the size of the shock and the dynamic responses of exchange rates after the initial period are not the same in the upper and lower panels of Figure 2. Thus, it is necessary to examine the impulse response of exchange-rate pass-through, which normalizes the consumer-prices response by the exchange-rate response. Specifically, similar to Ito and Sato [2008][17], we define the dynamic exchange rate pass-through as follows:

$$PT_l = \hat{p}_l^i/\hat{s}_l^i, \tag{12}$$

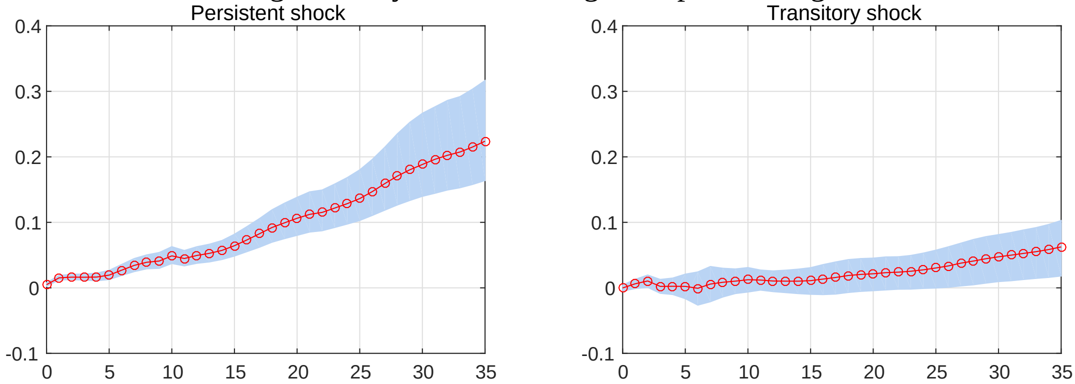
where $\hat{p}_l^i$ represents an impulse response of consumer prices to an $i$ shock after $l$ periods, and $\hat{s}_l^i$ represents the corresponding impulse response of exchange rates.

Figure 3 is the central result of this study. It illustrates how much exchange-rate movements pass through to consumer prices. The pass-through rate varies depending on the characteristics of underlying shocks. As suggested in the model's prediction in Section 2 and in other previous literature, such as Giovannini [1988] and Froot and Klemperer [1989], a persistent shock exerts a greater impact on consumer prices than a transitory

rate forecast in the initial period, implying that market participants' forecast of current depreciation would persist for a considerable period of time. Meanwhile, the lower panels in Figure 2 show that a transitory shock raises the spot exchange rate in the initial period, but has a minimal effect on the six-month exchange-rate forecast in the same period, suggesting that exchange-rate depreciation is perceived to be transitory. Our range restriction, with threshold values of $\phi_1 = \phi_2 = 0.9^6$, is successful in identifying persistent and transitory shocks, separately. In a later section, we will examine the sensitivity of results to alternative threshold values.

[17]The definition in Ito and Sato [2008] is slightly different from ours. They use cumulative impulse responses of price changes and exchange-rate changes because they use the first-difference VAR model.

Figure 3: Dynamic exchange rate pass-through

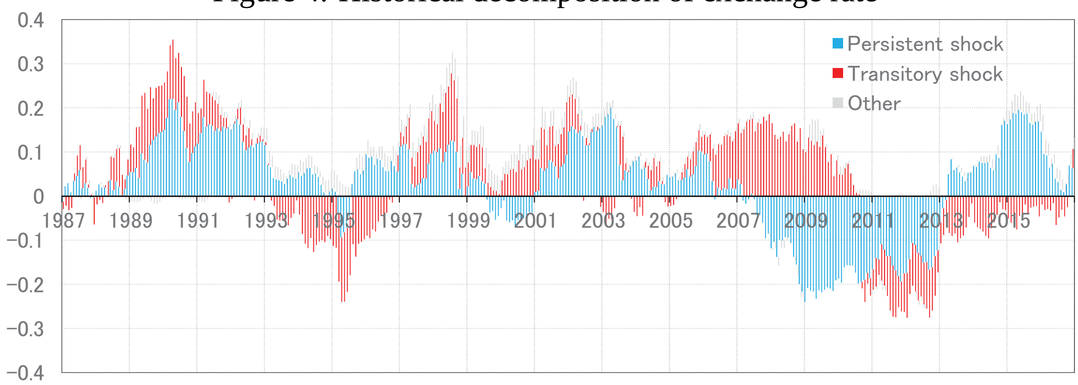



Note: The figures depict dynamic behaviors of exchange-rate pass-through after persistent and transitory shocks. The red lines and shaded areas represent medians and interquartile ranges of distributions, respectively.

shock. Specifically, as shown in the left-hand-side panel, the dynamic pass-through starts to rise gradually after a persistent shock hits the economy and reaches 0.22 after three years. In contrast, for a transitory shock in the right-hand-side panel, the dynamic pass-through stays around zero and approaches 0.06 after three years.

Our results are quantitatively similar to or slightly higher than Shioji [2014]'s (long-run) pass-through coefficients, which, using a two percent exchange-rate shock in a standard VAR model, show coefficient values that range between 0.0069 and 0.2827. Making a distinction between persistent and transitory shocks may contribute to higher pass-through for persistent shocks in our results. Another characteristics in Figure 3 is the delayed pass-through. As suggested in Nakamura and Zerom [2010], a gradual pass-through response is consistent with a menu-cost model in the retail sector. In summary, Figure 3 clearly indicates that different shocks, causing the same amount of depreciation, have different effects on consumer prices.

Figure 4: Historical decomposition of exchange rate




Note: The spot exchange rate in the logarithm is decomposed. Contributions of the constant term are omitted.

To understand the relative importance of transitory shocks and persistent shocks over time, Figure 4 depicts the historical decomposition of exchange rates. It indicates that the dominant sources of exchange-rate movements vary over time. For example, Figure 4 suggests that a transitory shock was the major driver of exchange-rate developments in 1995, when yen appreciation was considered ”beyond the levels justified by underlying economic conditions in major countries”[18]. In contrast, our model suggests that a permanent shock drove up the yen-dollar exchange rate after the first half of 2013, when the yen depreciation was caused by the regime change in Japan's monetary policy operation.[19] Accordingly, the regime shift in monetary policy was considered to persist and the consensus forecast of long-term inflation rose from 0.7 to 1.4 percent.[20] Historical decomposition of exchange rate shows that the estimated shocks are consistent with these episodes.

## 5.2 Implications for time-varying exchange-rate pass-through

To examine whether changes in shock compositions cause time variations in pass-through,[21] we will define the following implicit pass-through, using an estimated time series of shocks and pass-through ratios,

$$impPT_t \quad = \quad \sum_i \bar{PT}^i \frac{\Delta s_t^i}{\Delta s_t}, \tag{13}$$

where $\Delta s_t^i$ is an exchange-rate change induced by a shock $i$ in Figure 4 and $\bar{PT}^i$ is $i$ shock-specific pass-through coefficients in Figure 3. (13) represents how much of pass-through variations can be attributable to changes in shock composition.

Figure 5 illustrates that implicit pass-through ratios move up and down in short periods of time. Conventional research has attempted to explain time-varying pass-through as a result of slow-moving structural changes. In contrast, the result presented in Figure 5 offers alternative possibilities of fast-moving pass through, by focusing on changes in shock compositions.

Table 1 suggests that the part of pass-through variations over time is attributable to changes in shock composition. The table compares our results with those of Shioji [2014] and shows pass-through ratios were low in 1995, but increased in 2012. These results are also consistent with findings in Shioji [2015]. However, time-variation in our results are smaller than that in Shioji [2014]. Thus, our explanation complements Shioji [2015]'s slow-moving structural-change hypotheses, which suggest that changes in input-output

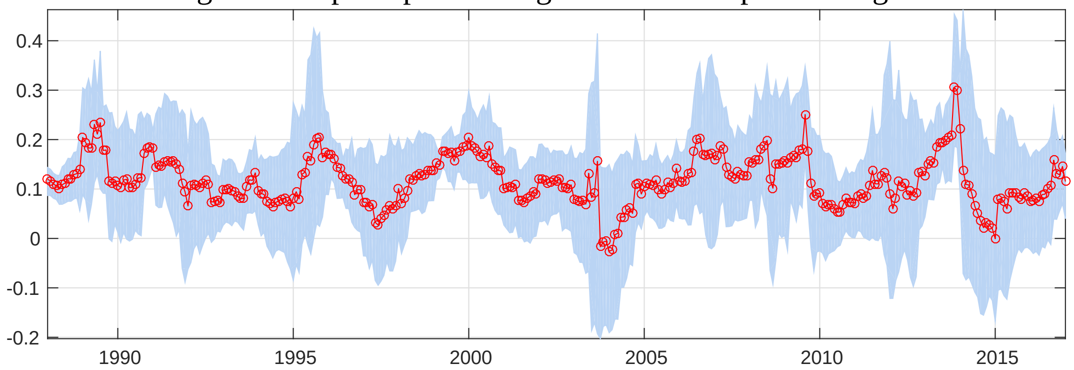
[18]The statement at the G8 Finance Minister meeting, April 25, 1995

[19]The Bank of Japan adopted a new policy regime called quantitative and qualitative monetary easing, in which Governor Kuroda declared to ”do whatever is necessary to overcome deflation” and ”dramatically change(ing) the expectations of market participants as well as firms and households” (”Quantitative and qualitative monetary easing,” speech at a meeting held by the Yomiuri International Economic Society in Tokyo, April 12, 2013).

[20]The figures are two-to-six-year ahead expected inflation of the ESP forecast in 2013/Q1 and 2014/Q4.

[21]Campa and Goldberg [2005], Otani, Shiratsuka, and Shirota [2003], Otani, Shiratsuka, and Shirota [2006], Shioji [2014], and Shioji [2015] found that pass-through coefficients for domestic and import prices are time-varying. These studies mainly consider slow-moving pass-through changes over decades.

Figure 5: Implied pass-through to consumer price changes




Note: Relative contributions of shocks to exchange-rate changes are calculated based on the historical decompositions in Figure 4. Changes in exchange rates are year-on-year changes. The pass-through coefficient assigned to each shock is the one with a 3-year horizon.

structure in production is the source of pass-through variations over time.

Table 1: Time variation in implied pass-through

| | (1) Jan. 1995 | (2) Oct. 2012 | (2)-(1) |
|---|---|---|---|
| Implicit ERPT | 0.0822 | 0.1325 | 0.0503 |
| Shioji [2014] | -0.0132 | 0.1632 | 0.1764 |

Note: Shioji [2014]'s pass-through coefficients are cumulative impulse responses of CPI goods to an exchange-rate shock, based on a time-varying-parameter VAR model.

A closer look at Figure 5 provides an interesting implication about the initial success of the monetary policy shift in April, 2013.[22] Figure 5 shows that the implied pass-through jumped up from 0.086 in April, 2011 to 0.185 in April, 2013. Using Figure 4, one can see that currency depreciation, which was associated with massive monetary easing, was perceived to persist for some time, and firms decided to incorporate these exchange-rate movements into their prices more aggressively.

## 5.3 Application to disaggregated prices

We apply our identification procedure to disaggregated prices for two purposes. First, Campa and Goldberg [2005] claim that time variations in exchange-rate pass-through are at least partly attributable to the aggregation bias. So, we need to check whether our results hold for not only the aggregate CPI but also disaggregated prices. The second purpose is to explore a relation between the exchange-rate pass-through and the frequency of price adjustments. Gopinath and Itskhoki [2010] first find this systematic relationship

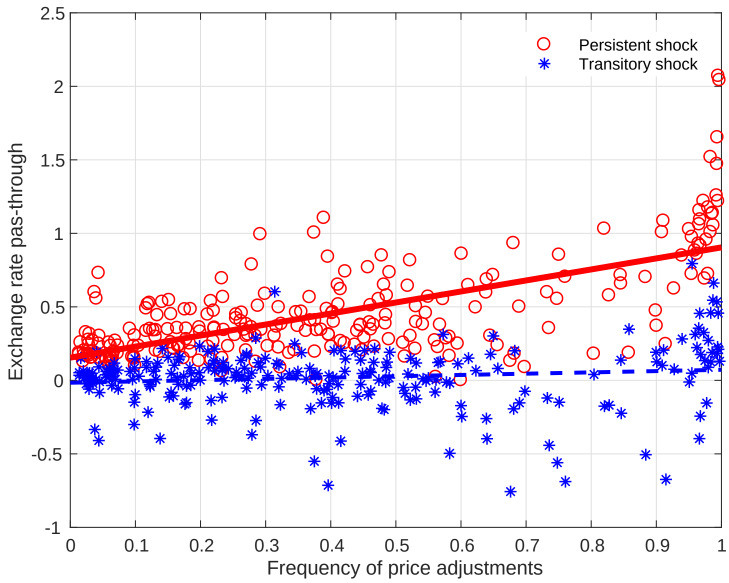
[22] After the shift of monetary policy stance, the CPI inflation rate increased from zero percent to more than one percent.

and study a firm's pricing behavior. They, however, do not consider underlying shocks behind exchange-rate movements.

In this exercise, we adopt 248 disaggregated categories of the CPI index, which are consecutively available in the sample period of January, 1986 to December, 2016.[23] We estimate a dynamic pass-through coefficient of (12), using the VAR model of (9) although $p_t$ in a VAR is replaced with each disaggregated price.[24]

Figure 6 shows that our baseline conclusion holds for most disaggregated prices. Disaggregated domestic prices are more responsive to exchange-rate movements when underlying shocks are persistent.[25] According to previous studies, various attributes of individual prices such as invoiced currencies, domestic distribution wedges, production structures, etc can affect the degree of pass-through. The results in the figure suggest that the pass-through is still shock-specific even if these attributes are controlled.

Figure 6: Shock-specific pass-through for disaggregated prices




Note: The exchange-rate pass-through is the dynamic pass-through of (12) with a three-year horizon.

Figure 6 also presents an interesting relation between shock-specific pass-through and frequencies of price adjustments. The frequency of price adjustments has a positive correlation with pass-through in case of a persistent shock, as is pointed out in Gopinath and Itskhoki [2010]. However, it does not so in case of a transitory shock.[26]

Based on the above results, we can deduce firms' pricing behavior as follows. First, firms' pricing decision is dynamic because high-frequency adjusters have a pass-through

[23]See Appendix C for details of the disaggregated dataset.

[24]Identified shocks are basically indistinguishable from those of the baseline model.

[25]A difference of pass-through between persistent and transitory shocks is negative only for six series.

[26]Although Figure 6 depicts the dynamic pass-through in a three-year horizon, similar patterns can be observed in one-year and two-year horizons.

that is higher than that of low-frequency adjusters, at least in the case of persistent shocks. Gopinath and Itskhoki [2010] argue that a dynamic pricing with a combination of nominal and real rigidities is a source of higher pass-through of high-frequency price adjusters. Our results are in line with the findings in previous studies. Second, it can be pointed out that the marginal cost of final-goods firms may hardly change for temporary shocks. As inferred in Burstein, Neves, and Rebelo [2003], Goldberg and Campa [2010], Ito and Sato [2008], and Shioji and Uchino [2011], multiple stages of production hamper the pass-through of exchange-rate movements to downstream sectors. An effect of a temporary shock, which changes a firms' desired price to a small extent, may disappear in the process of production chain.[27]

# 6 Robustness Analysis

To examine the sensitivity of empirical results to parameter settings and model specifications, we have estimated several variations of the baseline model. First, we set two alternative threshold values for the range restriction. One is the higher value, equal to 0.99. The other is the lower value, equal to 0.5. The latter figure is almost equivalent to Uhlig [2005]'s original sign restriction, because the resulting threshold value of the six-month forecast is approximately zero ($0.5^6 = 0.0156$).

Second, we expand the set of variables in the VAR model, adding crude oil prices and industrial production. Both variables are expected to control variations in marginal costs.

Third, we do not restrict responses of domestic prices. Different from the identification restrictions of the baseline case in (10), the $(3, 1)$ element of this alternative case in (14) is unrestricted. This exercise can be called "agnostic" in a sense of Uhlig [2005].

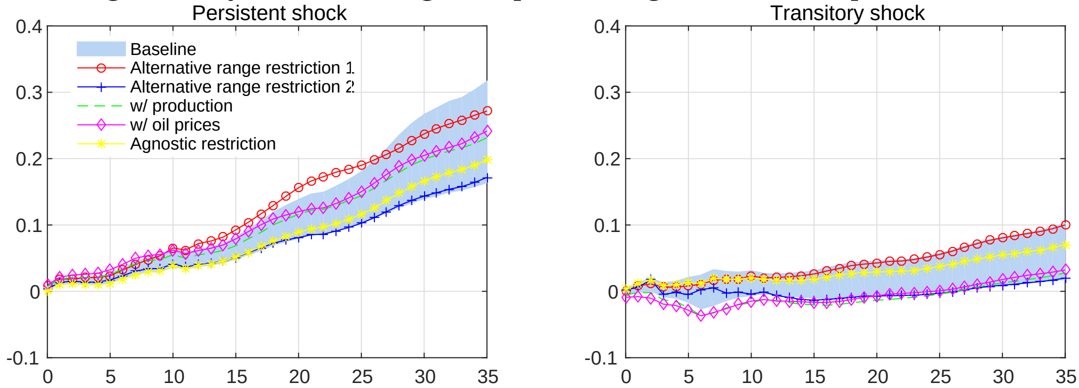
$$\boldsymbol{B}_{\mathbf{0}}^{-\mathbf{1}} \equiv \begin{bmatrix} b_{11} & b_{12} & * \\ b_{21} & b_{22} & * \\ * & * & * \end{bmatrix}, \tag{14}$$

Figure 7 presents point estimates of dynamic pass-through for all cases. Consistent with the theoretical prediction of the dynamic pricing model in Section 2, the higher threshold value of range restriction leads to a higher exchange-rate pass-through. However, in general, the pass-through in Figure 7 are basically similar and stay within an interquartile range of the baseline estimation. For the "agnostic" restriction, the persistent shock pass-through is approximately three times larger than the transitory shock pass-through, even if no restriction is imposed on domestic prices' response. We conclude that our baseline results are robust to these changes.[28]

[27] In order to analyze this point, it is necessary to measure the influence on the marginal cost of each item that is difficult to observe. I would like to make it a challenge of the future.

[28] The results of disaggregated prices are also robust to the series of the robustness checks in this section.

Figure 7: Dynamic exchange rate pass-through: alternative specifications




Note: The shaded areas are interquartile ranges of the baseline estimation.

# 7 Conclusion

This study develops a new framework to identify persistent and transitory shocks in exchange-rate movements and estimates a shock-specific pass-through to domestic prices. The framework combines a dataset of exchange-rate forecasts from the 1980s with a range restriction that is a generalization of Uhlig [2005]'s sign restriction.

The empirical results show that exchange-rate pass-through becomes higher when shocks behind exchange-rate movements are more persistent. The composition of persistent and transitory shocks varies over time. Past and recent literature has analyzed time variations of exchange rate pass-through. This research clarifies that time variations of exchange rate pass-through are at least partly due to the difference in shock-specific pass-through rates and changes in shock composition over time.

Apart from conventions in the literature, this study sheds light on fast-moving pass-through, rather than slow-moving one which is induced by structural changes. This is a part of recent strand of literature, by authors such as Shambaugh [2008] and Forbes et al. [2015]. The important difference in this study is its focus on the persistence of exchange-rate movements, which is pointed out in the classic theoretical studies. Applying our identification procedure to disaggregated prices of the CPI, we also find that a correlation between pass-through coefficients and frequencies of price adjustments is shock dependent. Specifically, the positive correlation, which is reported in Gopinath and Itskhoki [2010], disappears, when exchange-rate movements are transitory. The identification strategy in this study has an advantage, because the risk of model misspecification is relatively low, since it only relies on a small number of short-run restrictions.

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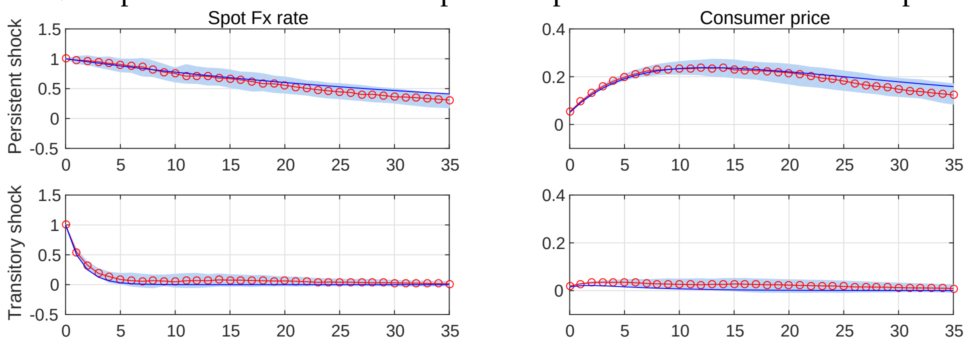
# A Monte Carlo Experiment

We conduct a simple Monte Carlo experiment to test the performance of our identification procedure. Two thousand samples of artificial time-series data are generated from the model described as (2), (5), (6), and the exogenous wage, which is the AR(1) process as follows: $w_{t+1} = \rho_w w_t + \epsilon^w_{t+1}$ where $\epsilon^w_{t+1} \sim N(0, 1)$.

At each iteration, we simulate 350 observations for the model's variables. The first 50 observations are discarded to avoid the effects of initial values. Using artificial data, we apply a range restriction and estimate impulse responses to a persistent shock and a transitory shock. For this exercise, we set the following parameters: $\rho_P = 0.975$, $\rho_T = 0.5$, and $\rho_w = 0.5$. The innovation of each shock follows a normal distribution of $N(0, 1)$.[29] The number of lags for the SVAR is six.

Figure 8: Empirical and theoretical impulse responses in a Monte Carlo experiment




Note: The red line with circles and shaded areas represent the median and interquartile ranges of estimated impulse responses. The blue solid line represents theoretical impulse responses.

The results in Figure 8 suggest that our identification procedure is capable of recovering responses to a persistent shock and a transitory shock. The figure shows that theoretical responses and the distributions of estimated responses. The estimated responses of exchange rate and domestic price to a persistent shock and a transitory shock match the theoretical responses closely in terms of magnitude and persistence. Specifically, the persistence of exchange rate are distinctively different between these cases: a persistent shock remains positive even after 3 years have passed whereas a transitory shock approaches to zero after several months. The theoretical responses of prices are also different, reflecting the calibrated pass-through parameters are not the same among two cases. The estimated responses are successful in following these theoretical responses.

[29]In this Monte Carlo experiment, the number of Gibbs sampling is 10,000, and the number of second stage draws of $\boldsymbol{B}_0^{-1}$ is 100. The first 8,000 Gibbs iterations are discarded as burn-ins.

# B Data Source

| | Description | Source |
|---|---|---|
| $S_t$ | Yen-Dollar exchange rate (End of month) | Bloomberg |
| $S_{t,t+6}$ | 6-month forecasts of Yen-Dollar exchange rate | JCIF Survey |
| $P_t$ | Consumer price index excluding fresh foods | Ministry of Internal Affairs and Telecommunication |
| $OIL$ | Crude oil price index | IFS |
| $USPCE$ | U.S. PCE deflator | IFS |
| $IIP$ | Index of Industrial Production | Ministry of Economy, Trade, and Industry |

# C Data of Disaggregated Prices

Disaggregated prices in the main text are taken from the CPI. 298 series of disaggregated prices are available for the entire sample period of January, 1986 to December, 2016.

Following, Higo and Saita [2007], we calculate frequencies of price adjustments using CPI micro data in the Retail Price Survey from 2000 to 2016. Frequencies of price adjustments in Figure 6 are averages of annual frequencies. As we exclude items that are missing in some cities, the number of items in Figure 6 is 248. The list of specific items is available upon request.