Exchange Rate Pass-Through and Monetary Policy: A Cross-Commodity Analysis

Jui-Chuan Chang and Ching-Chuan Tsong

ABSTRACT: This paper investigates how a change in monetary policy affects the degree and the speed of exchange rate pass-through to import prices in the emerging market economy, using a newly constructed data set from Taiwan's trading commodities. First, the analytical framework is set up following Goldberg and Knetter (1997) and Campa and Goldberg (2005). Next, the period-by-period and the multiple-period cumulative effects of monetary policy on the degree of exchange rate pass-through can be traced out. The dynamic panel data model is then estimated by Bun and Carree's (2005) bias-corrected approach, which enjoys easy calculation and robust testing performances, leading to more reliable empirical results. Our cross-commodity evidence strongly supports the partial pass-through in the short run and the complete pass-through in the long run. Moreover, following a change in monetary policy, this pass-through effect increases during several initial periods and declines to zero over time.

KEY WORDS: dynamic panel, emerging market, exchange rate pass-through, monetary policy.

The degree to which changes in the exchange rate pass through to import prices has long been of great interest to policymakers. A small exchange rate pass-through indicates that the movements in the nominal exchange rate may dampen the expenditure-switching effects of the domestic monetary policy and consequently reverse the cross-country correlation of output. In contrast with traditional literature focusing on the role of market power, Taylor (2000) pioneered the link between the exchange rate pass-through and macroeconomic factors and put forth the hypothesis that the low-inflation environment, which was brought about by more credible monetary policies, successfully reduced the degree of exchange rate pass-through to domestic prices.

Since Taylor (2000), voluminous theoretical literature has concentrated on this issue. However, the relationship between the exchange rate pass-through and macroeconomic factors has been examined recently in only a handful of empirical studies, for example, those by Bailliu and Fujii (2004), Campa and Goldberg (2005), Choudhri and Hakura (2006), and Gagnon and Ihrig (2004). Methodologically, most of these studies employed a cross-sectional framework and focused on explaining cross-country variations in pass-through elasticities. Although informative, they cannot address the question of how a change in monetary policy influences the sectoral exchange rate pass-through. Moreover,

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Jui-Chuan Chang (jccdella@mail.ncyu.edu.tw) is a professor in the Department of Applied Economics, National Chia Yi University. Ching-Chuan Tsong (tcc126@ncnu.edu.tw) is a professor in the Department of Economics at National Chi Nan University. The authors are indebted to Jürgen von Hagen, Eric M. Leeper, Ilir Miteza, Tai-Kuang Ho, and participants in the Western Economic Association International 82nd Annual Conference at Seattle and the Mini-Workshop at National Chung Cheng University for their insightful comments and suggestions. They also thank Chung-Hwa Shih for her assistance in compiling the data set. Financial support from the National Science Council is gratefully acknowledged (NSC95-2415-H-260-007). The usual disclaimer applies.

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the research into pass-through effects has focused mainly on larger or industrialized economies. As seen in Menon (1995), more than 50 percent of the estimates of the pass-through effects relate to the United States, Germany, and Japan. In contrast, smaller and more trade-dependent emerging economies draw less attention.

This paper attempts to fill these gaps and provide cross-commodity evidence on how a change in monetary policy affects the degree and the speed of the exchange rate pass-through to import prices in emerging market economies such as Taiwan. Known as one of the Asian Four Dragons, Taiwan is highly open and dependent on international trade. Therefore, it offers an interesting case study, complementary to the existing literature on this subject.

The contribution of this paper is threefold. First, with the newly constructed panel data set, this is the first paper to address the issue for Taiwan in terms of a cross-commodity analysis in a dynamic panel data model. As mentioned, previous studies mainly focused on developed countries, and even those that included developing countries neglected the case of Taiwan (e.g., Choudhri and Hakura 2006). Whether the results for developed countries can be extended to Taiwan is still unknown.

Second, we derive our analytical framework from a variant of the firm pricing behavior model developed by Campa and Goldberg (2005) and Goldberg and Knetter (1997). By incorporating dynamic adjustment into the model, we construct the panel data model for this empirical study. Next, the period-by-period and the multiple-period cumulative effects of monetary policy on the degree of exchange rate pass-through can be easily traced out, in contrast to the current literature, almost all of which focuses only on the short-run and the long-run effects. Therefore, this method can effectively reveal the transitional adjustment of the exchange rate pass-through period by period after a monetary shock.

Third, taking sixteen trading commodities in Taiwan over the period from 1996:M10 to 2004:M12, we estimate the fixed effects dynamic panel data model using the bias-corrected (BC) approach proposed by Bun and Carree (2005). In contrast to the conventional least-squares dummy variable (LSDV) and generalized method of moments (GMM) approaches commonly used in the literature, the BC estimation has the advantage of easy calculation, and the corresponding statistics offer a robust performance of type I error that leads to more reliable empirical results. Therefore, this contribution to the existing literature on the exchange rate pass-through is in terms of the econometric methodology.

Our cross-commodity evidence strongly supports the partial exchange rate pass-through in the short run and the complete pass-through in the long run. Moreover, following a change in monetary policy, this pass-through effect increases during several initial periods and declines to zero over time. The results of this paper not only provide an understanding of the link between monetary policy and cross-commodity exchange rate pass-through to import prices in Taiwan but also may serve as references for conducting and designing monetary policy, for economic stabilization, for international transmissions, and so on.

Analytical Model

The microfoundations of a firm's optimal pricing behavior by exporters are a useful starting point for understanding the dynamics of the exchange rate pass-through to import prices. The import prices for any country, P^I, are a transformation of the export prices of that country's trading partners, P^X, using the exchange rate S (defined as the foreign currencies per unit of domestic currency): $P^I = P^X/S$.

Based on the pricing-to-market model of Goldberg and Knetter (1997) and Campa and Goldberg (2005), our analytical framework considers a representative profit-maximizing exporting firm that produces goods for sale in foreign markets. The profit function of a monopolist is assumed to be the difference between its revenue and its cost, as follows:

$$\pi_t = S_t P_t^I h\left(P_t^I, Y_t\right) - C\left\{h\left(P_t^I, Y_t\right), Q_t^W\right\}, \tag{1}$$

where π denotes the firm's profit (expressed in the foreign currency) and $h(\cdot)$ is the quantity of demand for goods, determined by some factors such as domestic income Y and its selling price P^I. $C(\cdot)$ is the cost function (denominated in the foreign currency), taking into account the demand $h(\cdot)$ and the input price of production Q^W (such as labor wages).

Taking the first-order derivative of Equation (1) with respect to P^I, the import prices can be rearranged as

$$P_t^I = \frac{1}{S_t} MC\left\{h\left(P_t^I, Y_t\right), Q_t^W\right\}\left(\frac{\eta}{\eta - 1}\right), \tag{2}$$

where $\eta = -(\partial h_t(\cdot)/\partial P_t^I)(P_t^I/h_t(\cdot))$ denote the markup of the foreign price over the importer's marginal cost, $MC\{\cdot\}$. The expression in Equation (2) emphasizes that the local currency price of goods can vary as a result of a change in the exchange rate, a change in the firm's marginal cost, or a change in the firm's markup. To estimate the pass-through, therefore, it is important to take into account the movements in these other determinants of import prices.

Using lower-case letters to reflect logarithms, we rewrite Equation (2) as follows:

$$pi_t = \gamma_0 + \gamma_1 s_t + \gamma_2 y_t + \gamma_3 qw_t, \tag{3}$$

where the coefficient γ_1 measures the degree of the exchange rate pass-through. As proposed by Goldberg and Knetter (1997) and others, variants of Equation (3) are widely used as empirical specifications in the discussion of producer versus local currency pricing. If $\gamma_1 = -1$, producer currency pricing takes place; if $\gamma_1 = 0$, local currency pricing does, and exporters fully absorb the fluctuations in exchange rates in their own markups.

This effect has been examined often in studies on Organisation for Economic Co-operation and Development (OECD) countries; however, there is little evidence that emerging economies exhibit this link at the sectoral level, which could partly be due to lack of sectoral data. The fact that Taiwan has such disaggregate data allows this paper to consider the sectoral pass-through effects on Taiwan's import prices. The advantages of panel structure, namely, the possibility of the sectoral disaggregation of the data and the incorporation of time and cross-sectional effects, offer the rationale for the solution of our investigation. The cross-commodity effects could be either fixed (that is, a constant that varies for each cross-sectional unit) or random (that is, a random variable drawn from a common distribution). We choose the fixed effects model rather than its random effects counterpart because our empirical study covers all the available trading commodities in Taiwan and does not take a random sample from a large group of commodities.

Moreover, the literature on inflation dynamics has emphasized the importance of accounting for the observed inertial behavior of inflation. In our case, the effects of an exchange rate change in period t will be memorized, and these feedback effects will subsequently influence import price inflation over several periods. By extending Calvo's (1983) price setting to allow for a backward-looking rule of thumb, inertia in inflation is accomplished by including a lagged inflation as one of the explanatory variables in our empirical work. Such a dynamic model is able to capture the feedback effects and properties of time series—serial correlation.

Finally, as suggested by the new open economy macroeconomics literature, we wish to account for the fact that pass-through may be a function of macroeconomic variables. We do this by including interaction terms in our pass-through regression model between the rates of changes in the exchange rate and monetary variable that capture the changes in the inflation environment.

Accordingly, the specification of fixed effects dynamic panel data appears credible, so Equation (3) can be modified as follows:

$$pi_{k,t} = a_k + a_1 pi_{k,t-1} + a_2 s_t + a_3 y_{k,t} + a_4 qw_{k,t} + a_5 \left(s_t * m_t\right) + \nu_{k,t}, \tag{4}$$

where k denotes the type of trading commodity, the intercept a_k captures the commodity-specific effect, $(s * m)$ is the interaction term between the exchange rate and monetary variable, and ν is the error term.

Empirical Analysis

Data Description

To conduct our empirical analysis, we used a newly constructed panel data set that comprises sixteen trading commodities in Taiwan,[1] with the accessible series commencing from 1996:M10 and extending up to 2004:M12. All the raw data we needed were obtained from the Taipei Foreign Exchange Market Development Foundation, the *Taiwan Economic Journal,* the Central Bank of China, the Bureau of Foreign Trade, and the Datastream and AREMOS databases. Table 1 provides additional details on data description and computation.

The variables were selected on the basis of the following considerations. First, as a price variable, the commodity-specific import price index (pi_k) was used. Previous studies such as those by McCarthy (2000) and Ito et al. (2005) attempted to incorporate other types of price variables (that is, producer price index and consumer price index) in their models in order to analyze how external shocks are transmitted from one price stage to the next. Owing to data accessibility for each commodity (k), we incorporate only one price variable in the regression.

Second, although many studies have used the bilateral exchange rate vis-à-vis the U.S. dollar, a nominal effective exchange rate (s) was used in our model. The effective exchange rate was more appropriate in our study because the total effect of the exchange rate changes was used as a measure in the case of a country with diversified trading partners. In this case, an increase in the nominal effective exchange rate indicates an appreciation, in conformity with the definition provided by the Taipei Foreign Exchange Market Development Foundation.

Third, money supply (m) was used to allow for the effects of monetary policy on inflation. The Federal funds rate or call rate is typically used in the literature, such as in Clark (1999) and Hahn (2003), but the rate reflects a substantial fluctuation in crisis-hit countries, especially for Asian countries. Thus, we included the growth rate of the narrowed monetary aggregate M1B, as opposed to the Federal funds or call rate, as a proxy variable.

Fourth, we attempted to allow for supply and demand shocks in our estimation. qw was a primary "control" variable representing exporter production costs, and y was a vector of

Table 1. Data description

Notation	Variable (frequency)	Source
pi_k	Commodity-specific import price index (M)	AREMOS
s	Nominal effective exchange rate index (M)	Taipei Foreign Exchange Market Development Foundation
m	Narrowed money supply–M1B (M)	Central Bank of China
y_k	Commodity-specific gross domestic product (A)	AREMOS
IIP_k	Commodity-specific gross industrial product index (M)	AREMOS
PPI_c	Country-specific producer price index (M)	Taiwan Economic Journal, Datastream, AREMOS
TV_c	Country-specific exports and imports trading volumes (A)	Bureau of Foreign Trade
WPI_k	Commodity-specific wholesale price index (M)	AREMOS
TV_k	Commodity-specific exports and imports trading volumes (A)	AREMOS

Notes: The sample covers the period 1996:M10–2004:M12. These trading commodities (k) are animal products, vegetable products, prepared foodstuffs and beverages, minerals products, chemical or allied products, plastics and articles thereof, rubber and articles thereof, raw hides and skins—leather, fur skins—and articles thereof, wood and articles of wood and allied products, pulp, paper and printing products, textile products, basic metals and articles thereof, machinery equipment, electrical equipment, transport equipment, and precision instruments. Countries (c) selected were based on the proportion of their trade with Taiwan accounting for over 1 percent of the total trading volumes in Taiwan. There are eighteen countries, including the United States, Japan, Hong Kong, China, South Korea, Germany, Singapore, Malaysia, the Netherlands, the Philippines, the United Kingdom, Thailand, Indonesia, France, Canada, Italy, Saudi Arabia, and Australia. At the end, the Australian data were excluded due to a short sample period, and the data for Japan, Singapore, and the Philippines were replaced by the wholesale price index instead.

other controls, such as the real gross domestic product (GDP) of the destination market, representing the domestic demand conditions. Biased estimation of the pass-through coefficient could arise if foreign labor wages or GDP were correlated with exchange rates but omitted from the regression. Moreover, following McCarthy (2000) and Ito et al. (2005), the output gap ($ygap_k$) for each commodity (k) was calculated as the residuals from the regression of the log of the industrial production index (IIP_k) on constant plus linear and quadratic time trends. In our case, the industrial production index was employed due to the availability of monthly series.

To be consistent with the characteristics of multilateral trade in Taiwan, the commodity-specific costs of foreign production (pf_k) were calculated using the formula proposed by Pollard and Coughlin (2004): $pf_k = (qw/dw) * WPI_k$, where

$$qw = \sum_{c=1}^{c=17}\left\{\left[\left(TV_c\right)\Big/\sum_{c=1}^{c=17}\left(TV_c\right)\right]* PPI_c\right\}$$

represents the foreign labor costs, TV_c is the country-specific trading volumes,[2] and PPI_c is the country-specific producer price index. In addition,

$$dw = \sum_{k=1}^{k=16}\left\{\left[\left(TV_k\right)\Big/\sum_{k=1}^{k=16}\left(TV_k\right)\right]* WPI_k\right\}$$

denotes the domestic labor cost, TV_k is the commodity-specific trading volume, and WPI_k is the commodity-specific wholesale price index.

Model Specification and Econometric Method

Based on Equation (4) and with the inclusion of some lags for dynamic adjustment, our empirical model can be specified as follows:

$$\begin{aligned} dlpi_{k,t} &= \alpha_k + \tau dlpi_{k,t-1} + \beta_0 dls_t + \sum_{i=1}^{i=17} \beta_i dls_{t-i} + \gamma dygap_{k,t} + \lambda dlpf_{k,t} \\ &+ \mu_0 \left(dls_t * dlm_t\right) + \sum_{i=1}^{i=17} \mu_i \left(dls_{t-i} * dlm_{t-i}\right) + \varepsilon_{k,t}, \end{aligned} \tag{5}$$

where the notations "1" and "*d*" indicate variables in logarithmic and first-difference forms,[3] respectively, $k = 1, 2, ..., 16$ denotes the type of trading commodity, ε is the error term, and the optimal lag length is 17, which is selected using the top-down method with a given maximum lag of 18.

Our empirical exploration gets started by examining the properties of parameters in Equation (5), because it allows us to check whether or not the estimation results are reasonable and to compare our pass-through estimates with others in the literature.[4] In a stationary dynamic model, the absolute value of the first-order autoregressive coefficient (τ) should be less than one. This specification shrinks to a static version once τ equals zero. Besides, *dygap* refers to the changes in the domestic demand conditions. Other things being equal, the smaller the demand for imports, the lower the import prices—that is, γ is negative. While the costs of foreign production (*dlpf*) increase, ceteris paribus, import prices rise; therefore, λ is positive.

Moreover, without considering the monetary policy (that is, $dlm = 0$), the degree of the exchange rate pass-through is captured by the relevant coefficients on the change in the exchange rate, that is, β_i for $i = 0, 1, 2, ..., 17$. The lags of *dls* are included in order to capture the adjustment effects. The expenditure-switching effect, whereby the exchange rate appreciation reduces the import prices, implies that β_i is supposedly negative. Further, the complexity of dynamic adjustment and the heterogeneity of commodities do not necessarily imply that each β_i is negative, but indicate that at least one of them will be negative. The larger the absolute value of β_i, the greater will be the degree of exchange rate pass-through at that point of time. To clarify the period-by-period effect of the exchange rate pass-through, let $ERPT_{t+h}$ denote the degree of a change in exchange rate at period t passing through into import prices at period $t + h$, which can be expressed as

$$ERPT_{t+h} = \frac{\partial dlpi_{k,t+h}}{\partial dls_t} = \begin{cases} \sum_{i=0}^{i=h} \tau^{h-i} \beta_i & 0 \le h \le 17 \\ \sum_{i=0}^{i=17} \tau^{h-i} \beta_i & h > 17. \end{cases} \tag{6}$$

From Equation (6), we can easily observe that the dynamic adjustment of the exchange rate pass-through depends on τ as well. Of particular significance is that the size of τ determines the diminishing speed at which the effect in the earlier period contributes to $ERPT_{t+h}$. The smaller the τ, the faster is the declining speed, and vice versa. Further, we can investigate the h-period cumulative effect of the exchange rate pass-through, denoted as $CERPT_{t+h}$, by summing up each period of $ERPT_{t+h}$ up to the h periods. That is,

$$CERPT_{t+h} = \sum_{j=0}^{j=h} ERPT_{t+j},$$

where $ERPT_{t+j}$ is defined in Equation (6). Finally, we can compute the long-run exchange rate pass-through, denoted as *LR_ERPT,* according to the following:

$$LR\_ERPT = \sum_{h=0}^{h=\infty} CERPT_{t+h} + \frac{\sum_{i=1}^{i=17} \beta_i}{1-\tau}. \tag{7}$$

Once again, the *LR_ERPT* is associated with τ as well as β_i. When τ increases, indicating that the period-by-period effect becomes greater, or

$$\left|\sum_{i=0}^{i=17} \beta_i\right|$$

is larger, the degree of *LR_ERPT* escalates. In summary, without considering the monetary policy, the short-run and the long-run exchange rate pass-through are measured by β_0 and

$$\left(\sum_{i=0}^{i=17} \beta_i\right) \Big/ (1-\tau),$$

respectively.

On the other hand, by including the interaction terms associated with the exchange rate and monetary variable ($dls_{t-i} * dlm_{t-i}$ for $i = 0, 1, 2, ..., 17$), we explore the macroeconomic determinants of the exchange rate pass-through in the specification of Equation (5). According to the results of the two-stage least squares estimation in Choudhri and Hakura (2006) and Devereux and Yetman (2002), the inflation environment and the exchange rate pass-through are positively correlated. Given the comovement between the monetary policy and the inflation environment, the sign of the interaction terms μ_i is supposed to be the same as that of β_i, which is less than zero. The larger the absolute value of μ_i, the greater is the degree to which the monetary policy affects the exchange rate pass-through. Similarly, the period-by-period effect and the *h*-period cumulative effect of monetary policy on the degree of exchange rate pass-through are represented by $ERPTM_{t+h}$ and $CERPTM_{t+h}$, respectively, as follows:

$$ERPTM_{t+h} = \frac{\partial}{\partial dlm_t}\left(\frac{\partial dlpi_{k,t+h}}{\partial dls_t}\right) = \begin{cases} \sum_{i=0}^{i=h} \tau^{h-i}\mu_i & 0 \le h \le 17 \\ \sum_{i=0}^{i=17} \tau^{h-i}\mu_i & h > 17 \end{cases} \tag{8}$$

$$CERPTM_{t+h} = \sum_{j=0}^{j=h} ERPTM_{t+j}. \tag{9}$$

This indicates that the short-run effect of the monetary policy depends on μ_0, while the long-run effect (*LR_ERPTM*) is captured by

$$\left(\sum_{i=0}^{i=17} \mu_i\right) \Big/ (1-\tau).$$

In this paper, we will conduct Bun and Carree's (2005) BC method for the estimation and hypothesis testing for the dynamic panel data model in Equation (5). As pointed out in Bun and Carree (2005), the BC estimator, based on the bias correction of the LSDV estimator, is easy to compute and without the choice of appropriate instruments, so that it yields more robust results than the GMM estimation. Moreover, the BC estimator has small finite-sample bias, leading to more robust testing statistics than the counterparts from the GMM procedure. These imply that more reliable empirical results can be obtained from the BC procedure.[5]

Empirical Results

Results of Estimation and Testing

Table 2 presents the empirical results. We mainly focus on the BC estimation and hypothesis testing procedures, but the LSDV and GMM counterparts are also included for comparison.[6] One reason for this is that the BC estimation originates from the LSDV estimation, and the testing results are more reliable. More important, the Wald test does not reject the null hypothesis of homoskedasticity, due to its statistic and *p*-value equal to 0.0346 and 0.8525, respectively. Moreover, the result of the LM test for serial correlation, its statistic and *p*-value being 0.2576 and 0.4972, respectively, suggests that the error terms are serially uncorrelated.[7] Accordingly, the error terms satisfy the requirements for homoskedasticity and no serial correlation, implying that the fixed effects dynamic panel data model is well specified under the BC estimation.

As we can see in Table 2, the first-order autoregressive coefficients (τ) are 0.2708, 0.2577, and 0.2495, corresponding to the BC, LSDV, and GMM estimations, respectively. They are all statistically significant even at the 1 percent level and are less than one. This indicates that, apart from the misspecification of the static counterpart, our dynamic panel data model is properly specified in the stationary version. Besides, the BC estimate of the first-order autoregressive coefficient is greater than the LSDV estimate,[8] which confirms the findings in the dynamic panel data literature. This reveals that the LSDV method usually underestimates the feedback effects of changes in the exchange rate, incorrectly indicating that the economy adjusts to the long-run equilibrium at a rapid speed.

In addition, for clarification, the relevant coefficients on the changes in the exchange rate (β_i and μ_i) are illustrated in Figure 1. Panels (a), (c), and (e) on the left of Figure 1 depict the estimated coefficients of the current and lagged changes in the exchange rate (β_i), whereas panels (b), (d), and (f) on the right plot the estimated coefficients of current and lagged interaction terms between the changes in the exchange rate and the money supply (μ_i). At a glance, the results of the GMM and BC estimations are similar throughout the graphs.

As shown in Table 2, the BC estimates of the parameters γ and λ, corresponding to the variables *dygap* and *dlpf,* are significant and consistent with our expectation. Furthermore, since the relevant coefficients on the changes in the exchange rate (β_i and μ_i) play crucial roles in our investigation, it is worth looking at them in some detail. First, the BC estimate of the short-run exchange rate pass-through (β_0) is equal to –0.4988, implying that a 1 percent increase in the depreciation rate of the nominal exchange rate leads to a 0.5 percent increase in the inflation rate of import prices at that point in time. Besides, this estimate is significantly different from zero at the 1 percent level and satisfies the expected (negative) sign, indicating that the short-run pass-through is partial. Second, taking a closer look at Figure 1(a), from the BC estimation we will observe that 9 out of 17 estimates of β_i are significantly different from zero in addition to β_0. It is noteworthy that there are six parameters significant during the last 8 periods from the 10-lag period to the 17-lag period, indicating that the impacts of the exchange rate movements on cross-commodity import prices are prolonged. Third, the long-run pass-through estimate of –1.0163 meets the expected (negative) sign and suggests that a 1 percent increase in

Table 2. Estimation results

Variable	BC	LSDV	GMM
$dlpi_{k,t-1}$	0.2708***	0.2577***	0.2495***
	(0.0248)	(0.0246)	(0.0787)
dls_t	–0.4988***	–0.4979***	–0.5155***
	(0.0935)	(0.0935)	(0.0907)
$dygap_{k,t}$	Figure 1(a)	Figure 1(c)	Figure 1(e)
$dygap_{k,t}$	–0.0094**	–0.0092**	–0.0070
	(0.0047)	(0.0047)	(0.0070)
$dlpf_{k,t}$	0.3110***	0.3119***	0.4829**
	(0.0217)	(0.0217)	(0.2384)
$(dls_t * dlm_t)$	–0.0320	–0.0326	–0.0345
	(0.0285)	(0.0285)	(0.0197)
$(dls_{t-i} * dlm_{t-i})$	Figure 1(b)	Figure 1(d)	Figure 1(f)
Adj. R^2	0.3359	0.3244	0.3035
Wald test for heteroskedasticity (p-value)	0.0346		
	(0.8525)		
LM test (p-value)	0.2576		
	(0.4972)		
J-statistic of Sargan test			607.98

Null hypothesis	Coefficient	Wald test
$H_0 : \left(\sum_{i=0}^{i=17} \beta_i\right) \Big/ (1-\tau) = -1$	–1.0163	0.0013
$H_0 : \left(\sum_{i=0}^{i=17} \mu_i\right) \Big/ (1-\tau) = 0$	0.1632	0.6240

Notes: The model

$$dlpi_{k,t} = \alpha_k + \tau dlpi_{k,t-1} + \beta_0 dls_t + \sum_{i=1}^{i=17} \beta_i dls_{t-i} + \gamma dygap_{k,t} + \lambda dlpf_{k,t} + \mu_0 \left(dls_t * dlm_t\right) + \sum_{i=1}^{i=17} \mu_i \left(dls_{t-i} * dlm_{t-i}\right) + \varepsilon_{k,t}$$

is estimated by the method of BC, LSDV, and GMM, respectively. The numbers in parentheses are standard errors. For GMM estimation, the valid instrumental variables include $dlpi_{k,t-2}$, ..., $dlpi_{k,t-5}$, $dygap_{k,t-1}$, ..., $dygap_{k,t-5}$, $dlpf_{k,t-1}$, ..., $dlpf_{k,t-5}$, dls_{t-18}, ..., and dls_{t-24}. ***, **, and * indicate that the null hypothesis is rejected at the 1 percent, 5 percent, and 10 percent significance levels, respectively.

the exchange rate depreciation raises the inflation rate of import prices to 1.02 percent in the long run. Furthermore, the Wald test cannot reject the null hypothesis of its absolute value equal to one, supporting that the pass-through is complete in the long run.

Using the quarterly data from 1975 to 2003, Campa and Goldberg (2005) find that the average rate of pass-through into import prices across twenty-three OECD countries is approximately 0.46 in the short run and 0.64 in the long run. Moreover, in Anderton's (2003) study of the exchange rate pass-through in eurozone countries, he shows a pass-through rate of between 0.5 and 0.7 for extra-eurozone imports. They all propose that a partial pass-through is the best description for import price response in the short run, whereas a complete pass-through is generally supported as a longer-run characterization. In addition, Choudhri and Hakura (2006) explore seventy-one countries, both industrial and developing, with emphasis on different inflation regimes over the period

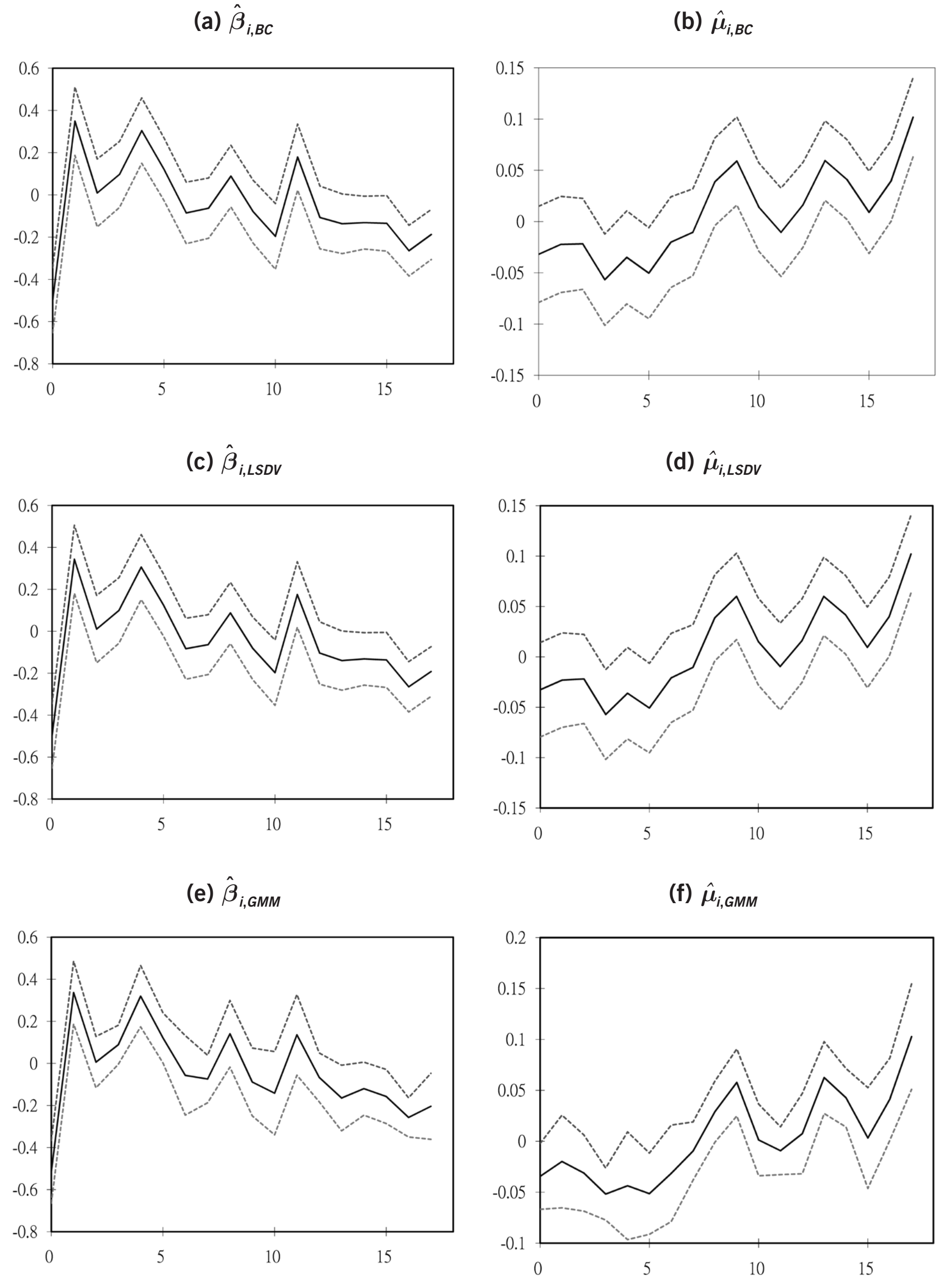

Figure 1. The relevant estimates of changes in exchange rate

Notes: The bold lines in panels (a), (c), and (e) depict the coefficients of the current and lagged changes in the exchange rate (β_i), and panels (b), (d), and (f) plot the coefficients of current and lagged interaction terms between the changes in the exchange rate and the money supply (μ_i), where $i = 0, 1, ..., 17$. The dotted lines refer to the 90 percent confidence bands.

1979–2000. They note that all industrial countries classified in the low-inflation group provide relatively low degree of pass-through, whereas developing countries appeared in all different groups, giving rise to a relatively high rate. Besides, Choudhri and Hakura (2006) state that the average pass-through for each inflation group increases in a longer term, without rigorously testing for the short-run and long-run effects. Accordingly, our results are consistent with the literature.

For the monetary policy, on the other hand, the point estimate of –0.0320 reported in Table 2 for the short-run effect suggests that a 1 percent increase in the growth rate of money supply leads to a 0.03 percent increase in the degree of exchange rate pass-through in the short run. Despite the estimate of μ_0 being insignificant at the 10 percent level, 7 out of 17 estimates of μ_i are significantly different from zero. This can be easily observed in Figure 1(b). Analogous to the estimates of β_i, there are four parameters significant during the last 5 periods from the 13-lag period to the 17-lag period, implying that the monetary policy conducted usually generates a deferred effect on the cross-commodity exchange rate pass-through. Conversely, the estimate of the long-run effect is equal to 0.1632, indicating that a 1 percent increase in the growth rate of money supply causes a 0.16 percent decline in the degree of exchange rate pass-through in the long run. However, the value of the Wald test for the null hypothesis of no long-run effect is 0.6240, revealing the ineffectiveness of the monetary policy in the long run.

Corresponding to the findings surveyed by McCarthy (2000), evidence supporting local currency pricing would indicate that the monetary policy should be used to deal with shocks. On the other hand, evidence in support of the producer currency pricing mechanism would suggest that monetary policy is ineffective in dealing with shocks. Recall that our results corroborate the partial exchange rate pass-through in the short run and the complete exchange rate pass-through in the long run. As such, our empirical results associated with monetary policy are likely to be effective in the short run and ineffective in the long run. Consequently, our results are in conformity with the literature.

Results of Calibration

To conduct further analysis over a period of up to 25 months, we take relevant coefficients on exchange rate changes based on the BC estimation to calibrate the period-by-period and cumulative adjustment effects. In Figure 2, panels (a) and (c) illustrate the period-by-period effects, $ERPT_{t+h}$ and $ERPTM_{t+h}$, respectively, and panels (b) and (d) reflect the cumulative effects, $CERPT_{t+h}$ and $CERPTM_{t+h}$, respectively. First, Figure 2(a) shows that the impact of a change in the nominal exchange rate on cross-commodity import prices is approximately –0.5. Next, such impacts turn positive between the 1st and 5th periods after the shift in the exchange rate. Then, they become negative except in the 8th and 11th periods and gradually decay to zero after the 16th period. By aggregating these impacts in panel (a), panel (b) first shows that the cumulative effects of the exchange rate pass-through to cross-commodity import prices are negative in the first four periods. After that, they switch and become positive between the 5th and 13th periods, then decline and become negative after the 14th period. Finally, they moderately move toward the long-run degree of the exchange rate pass-through.

The reasons for this are the following. At the start of the depreciation in the exchange rate, foreign exporters are likely to maintain the same export prices, driving up the import prices. The increased import prices relative to domestic goods dampens the demand for imports; in order to rescue their market share, foreign exporters may begin

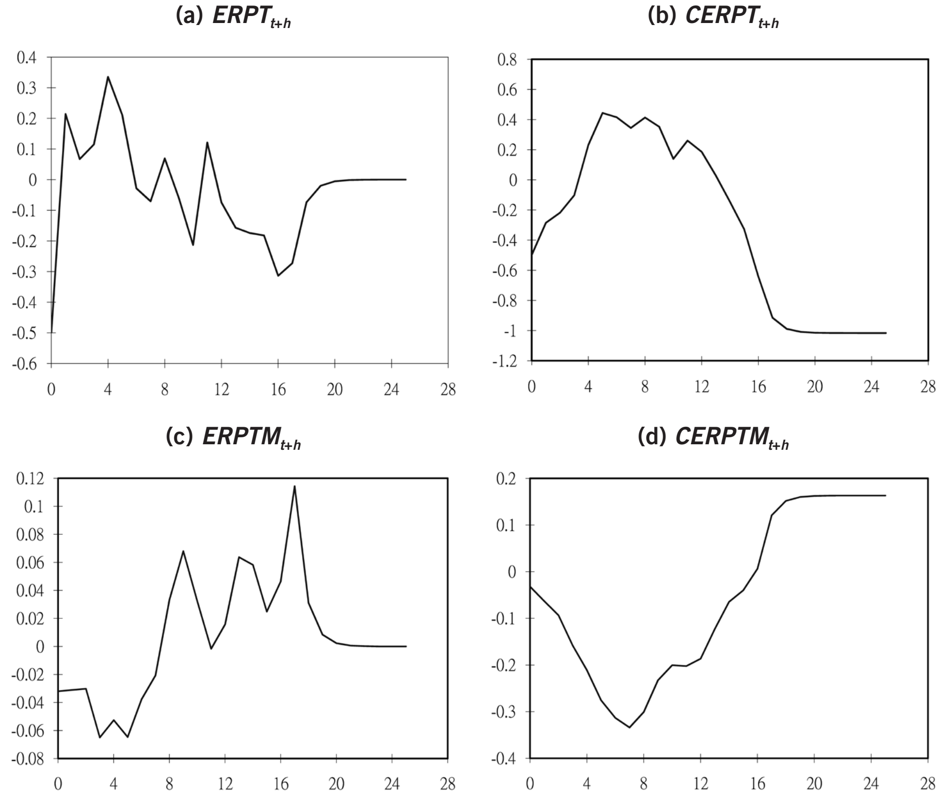

Figure 2. Calibration results

Notes: Given $h = 25$, $ERPT_{t+h}$, $CERPT_{t+h}$, $ERPTM_{t+h}$, and $CERPTM_{t+h}$ denote the calibrations for the period-by-period effect of exchange rate pass-through, the cumulative effect of exchange rate pass-through, the period-by-period effect of monetary policy, and the cumulative effect of monetary policy, respectively.

reducing their export prices and, consequently, their import prices in domestic currencies. Subsequently, if foreign exporters cannot balance certain profit margins by cutting their prices to maintain their market share, or if some firms cannot bear with transitory losses that result from pricing competition and decide to exit the market, the existing firms will resort to raising their prices, resulting in the completeness of the exchange rate pass-through for that period.

On the other hand, Figure 2(c) displays how the degree of the exchange rate pass-through to cross-commodity import prices fluctuates in response to a policy shock for each period. Following a monetary loosening, such impacts become negative for seven periods, subsequently turn positive (except for a sharp decline to a negative value in the 11th period), and finally fluctuate until they reach zero after the 17 periods. Figure 2(d) sums up the period-by-period effects and demonstrates that during the first several periods, the expansionary policy affects the degree of the exchange rate pass-through to an increasing extent (–0.032 in the current period, extending to –0.334 in the 7th period); subsequently, such impacts begin to move in the other direction, and after 21 months, they finally approach the long-run level of 0.163, which, as mentioned earlier, is insignificantly different from zero.

As documented by Faust and Rogers (2003), following a monetary expansion, the exchange rate patterns are likely to depreciate first and then appreciate, returning to their original equilibrium. Accordingly, import prices will first rise and then fall, offsetting the extent of increases in import prices accumulated initially until they reach zero. As a result, the monetary policy does not influence the degree and speed of exchange rate pass-through in the long run.

Conclusions

The main purpose of this paper is to investigate whether or not monetary policy affects the degree of the exchange rate pass-through into import prices based on Taiwan's trading commodities over the period from 1996:M10 to 2004:M12. We estimated a fixed effects dynamic panel data model using the LSDV, BC, and GMM approaches and conducted hypothesis testing and calibration through the BC procedure. Our cross-commodity evidence strongly supports the partial exchange rate pass-through to import prices in the short run and complete exchange rate pass-through in the long run. Moreover, we find that following a monetary policy change, the degree of exchange rate pass-through increases during the first several periods and thereafter declines to zero over time. Our results correspond with the evidence on the new open economy macroeconomics literature.

Thus, there are two policy implications. First, if government authorities intend to influence trade balances and macroeconomic activities through monetary policies, then the effects could be minor due to the incomplete exchange rate pass-through effects in the short run and gradually decay to zero, declaring the monetary policy is inoperative in the long run. Furthermore, if the central bank targets mild inflation pressure, then stabilizing import prices would be an effective method by stabilizing the exchange rate fluctuation.

Finally, one of the contributions of this paper is to newly construct a large data set with prices of sixteen commodities. We have checked the degree of the exchange rate pass-through into disaggregated import prices separately; however, we concentrate on the cross-commodity analysis here and leave detailed discussion for each specific commodity in our future work. We realize that such a disaggregated exploration would be very interesting and would open a wide new empirical agenda.

Notes

1. These trading commodities (k) are animal products, vegetable products, prepared foodstuffs and beverages, minerals products, chemical or allied products, plastics and articles thereof, rubber and articles thereof, raw hides and skins—leather, fur skins—and articles thereof, wood and articles of wood and allied products, pulp, paper and printing products, textile products, basic metals and articles thereof, machinery equipment, electrical equipment, transport equipment, and precision instruments.

2. Countries (c) selected were based on the proportion of their trade with Taiwan accounting for over 1 percent of the total trading volumes in Taiwan. There are eighteen countries, including the United States, Japan, Hong Kong, China, South Korea, Germany, Singapore, Malaysia, the Netherlands, the Philippines, the United Kingdom, Thailand, Indonesia, France, Canada, Italy, Saudi Arabia, and Australia. At the end, the Australian data were excluded due to a short sample period, and the data for Japan, Singapore, and the Philippines were replaced by the wholesale price index instead.

3. These variables are integrated of order one and without any cointegration relations investigated with augmented Dickey–Fuller and Johansen cointegration tests. These results are available from the authors upon request.

4. For comparison, we also reestimate Equation (5) to separately explore the exchange rate pass-through for each commodity. The information can be found in the previous version of this paper.

5. For brevity, the BC procedure and the detailed comparisons among BC, LSDV, and GMM estimators are omitted. They can be found in the previous version of this paper and are available upon request.

6. Based on the Sargan tests, these instrumental variables, including $dlpi_{k,t-2}$, ..., $dlpi_{k,t-5}$, $dygap_{k,t-1}$, ..., $dygap_{k,t-5}$, $dlpf_{k,t-1}$, ..., $dlpf_{k,t-5}$, dls_{t-18}, ..., and dls_{t-24}, are valid under the GMM estimation.

7. To the authors' knowledge, the Hausman test used for discriminating the fixed effects model from the random effects model is valid only under the static panel setting. In a dynamic model, its asymptotic distribution may differ from its conventional counterpart. Thus, this paper ignores such a testing procedure and takes a plausible fixed effects model based on the above-mentioned premise.

8. In this aspect, however, the literature addresses nothing pertaining to the GMM estimate.

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