Monetary Policy "Contagion" in the Pacific Basin: A Historical Inquiry

Sebastian Edwards

This paper analyzes the way in which Federal Reserve policy actions have affected monetary policy in a group of countries in the Pacific Basin—three in Asia and three in Latin America—during the period 2000–08. The results indicate that historically there has been policy "contagion" and that during the period under study these countries tended to "import" Fed policies. The paper also finds that the pass-through has been higher in the Latin American countries than in the Asian countries and that the extent of "contagion" for the Latin American nations has been independent of their degree of capital mobility. In contrast, in East Asia there is some evidence suggesting that greater capital mobility has been associated with greater policy rate pass-through.

1. Introduction

In Samuel Beckett's play *Waiting for Godot*, two friends—Vladimir and Estragon—wait, in vain, for someone called Godot. They wait and wait, but he never comes. In many ways this play is about the "anxiety of waiting." During most of the year 2015, central bankers from around the world—and especially those from the emerging markets—felt as if they were living inside that play. They waited for the Federal Reserve to make a move and raise interest rates, and as time passed without the Fed taking action, they became increasingly anxious. The first sign of apprehension came in June 2013 during the so-called taper tantrum,[1] when an increasing number of influential central bankers from emerging market economies (EMEs) called for the Fed to begin normalizing monetary policy by raising rates. They wanted the "waiting game" to be over. For example, the governor of the Reserve Bank of India, Ragu Rajan, told the *Wall Street Journal* on August 30, 2015, "from the perspective of emerging markets . . . it's preferable to have a move early on and an advertised, slow move up rather

Author's note: *I have benefited from conversations with a number of former central bankers and policymakers, including John Taylor, Vittorio Corbo, José De Gregorio, and Guillermo Ortiz. I thank Alvaro Garcia for research assistance. I thank my discussants Linda Tesar and Woon Gyu Choi as well as conference participants for very helpful comments.*

than the Fed be forced to tighten more significantly down the line." The wait finally ended on December 17, 2015, when the Fed raised the federal funds policy target from near zero to a range of 0.25 to 0.50 percent. In the weeks that followed, many Latin American countries—Chile, Mexico, and Peru, for example—followed suit, and their respective central banks raised interest rates.[2] In contrast, during that same period most East Asian central banks remained "on hold." How emerging markets react—or will react—when the Fed begins raising interest rates raises important policy questions. What will be the spillover effects? How fast will higher global interest rates be transmitted into local financial markets and, more importantly, how will central banks react to the new reality of tighter global financial markets?[3]

According to the workhorse model of international macroeconomics—the Mundell-Fleming model—countries are able to undertake independent monetary policies under flexible exchange rates; that is, in principle central bank actions in emerging markets need not follow or even take into account the policy stance of advanced nations, such as the United States, as long as they are willing to operate with flexible exchange rates.[4] More recently, however, some authors, such as Taylor (2009, 2013, 2015) and Edwards (2015), have argued that even under flexible exchange rates there is policy interconnectedness across countries. That is, in a highly globalized economy, even when there are no obvious domestic reasons for raising interest rates, some central banks will follow the Fed. This policy "contagion" could be the result of a number of factors and considerations, including the desire to "protect" exchange rates from "excessive" depreciation, or a "fear of floating."[5] The late Ron McKinnon was a strong exponent of the existence and importance of policy "contagion." In May 2014, he stated at a conference held at the Hoover Institution that "there's only one country that's truly independent and can set its monetary policy. That's the United States."[6]

At the end, however, the extent of monetary policy independence is an empirical matter. If, for whatever reason, a particular central bank feels that it needs to mimic (or follow) advanced-country policy actions, then there will be policy "contagion" and the actual—as opposed to theoretical—degree of monetary policy autonomy will be greatly reduced.

There are many possible ways to analyze the extent of policy "contagion." One approach is to build a dynamic stochastic general equilibrium (DSGE) model of the world economy, with (some) large and (some) small countries, international trade, imperfect asset substitutability, and capital mobility. This setting could be used to address a number of important issues, including the transmission of the business cycle, "contagion," the way in which smaller countries accommodate and deal with real and monetary shocks, the propagation of crises to the real sector, saving and investment decisions, portfolio diversification, and the role of global banks in the transmission of disturbances, among others.[7] This type of model also could be used to gain insights into how central banks in small emerging countries are likely to react to a tightening of monetary policy in a center country, such as the United States.[8] The answer depends on the structure of the smaller economy, the degree of asset substitutability and capital mobility, pass-through coefficients, and the preferences of policymakers. An alternative approach is to use historical data to investigate how central banks in EMEs have in fact reacted in the past to changes in monetary policy in the United States.

This paper uses weekly data from six Pacific Basin countries—Chile, Colombia, and Mexico in Latin America; and Korea, Malaysia, and the Philippines in East Asia—to analyze the issue of policy "contagion" from a historical perspective. The sample period extends from January 2000 through early June 2008, before the onset of the recent global financial crisis. Error correction models are estimated both for individual nations and for pooled regional panels. Thus the sample period excludes the turmoil that followed the collapse of Lehman Brothers and led the Federal Reserve to pursue policies based on zero interest rates and quantitative easing (QE).[9]

The Latin American countries in the sample have several important characteristics in common: all followed inflation targeting and allowed some degree of exchange rate flexibility during the period under study (2000–08), all had a relatively high degree of capital mobility, and all had independent central banks. In that sense, they constitute a somewhat homogenous group. The three East Asian nations constitute a slightly more varied group. Korea and the Philippines had inflation-targeting regimes and some degree of currency flexibility, while during most of the period under study Malaysia had a fixed exchange rate (relative to the U.S. dollar). All three Asian central banks were de facto (but not necessarily de jure) quite independent from political pressure.[10]

The approach taken in this paper differs from the existing literature in several ways. First, other papers typically rely on either pooled (panel) data for a group of countries—often pooling countries as diverse as Argentina and India—or have based their simulations on a "representative EME." This paper focuses on individual countries with contrasting experiences. Second, other papers use monthly or quarterly data, while this paper uses short-term (weekly) time-series data, permitting analysis of the transmission from U.S. interest rates to interest rates in EMEs at high frequency. However, the approach taken in this paper does have some limitations. In particular, dealing with individual countries means that it is not possible to exploit cross-country variability in some variables, such as the extent of capital controls. In addition, the use of weekly data means that suitable proxies for real economic activity must be constructed. This paper uses data on commodity prices as an indicator commonly and strongly associated with real expansions and/or contractions in emerging markets.

The rest of the paper is organized as follows: Section 2 provides some background discussion and discusses the data. Section 3 deals with the theoretical underpinnings for the analysis. Section 4 reports the regression results for Latin America and East Asia using least-squares and instrumental-variables estimation, with a variety of controls. It also presents robustness and extension exercises. Section 5 provides preliminary results on the possible role of capital mobility in the pass-through process. Section 6 analyzes the extent to which Federal Reserve policies have affected market deposit rates in the three Asian and three Latin American countries. Section 7 offers concluding comments and suggests areas for future research. Section 8 lists the data sources used in the paper.

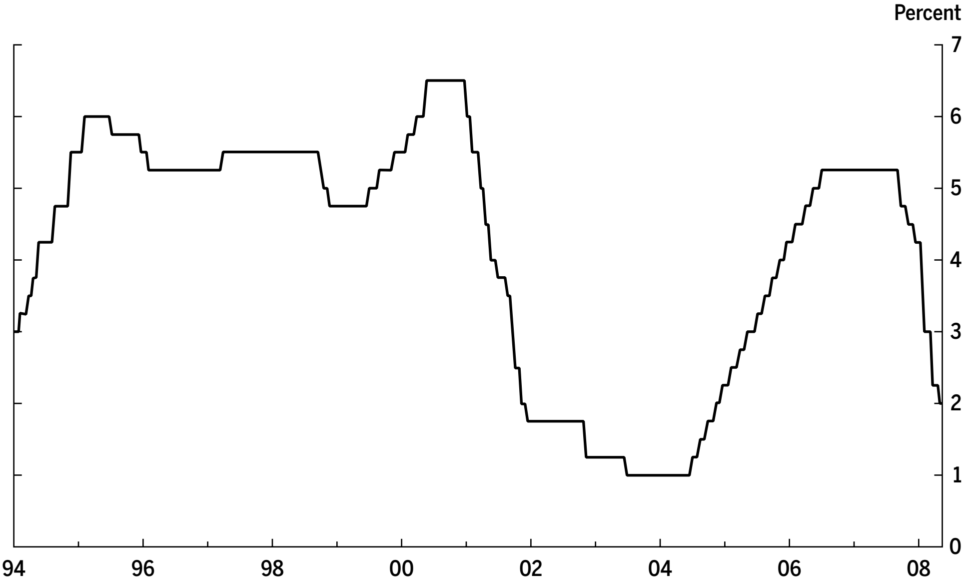
2. Background

Figure 1 plots the federal funds target rate with weekly data over the period 1994 through early June 2008, just before it was reduced to (near) zero and QE was subsequently implemented. Figure 2 shows weekly data on the policy rate for the six countries in the analysis: Chile, Colombia, Mexico, Korea, Malaysia, and the Philippines. As noted, the key question in this paper is the extent to which the central banks in these EMEs took into account the Fed's policy stance when determining their own monetary policy. In other words, with other things given, did (some of) these countries take into account Fed actions when deciding on their own policies, or did they act with complete independence?

Between January 2000 and September 2008 the federal funds policy target rate was changed 40 times. Twenty actions involved federal fund rate hikes, nineteen by 25 basis points (bps) and one by 50 bps (during the week of May 19, 2000). The other 20 policy actions involved cuts in the federal funds rate. In 7 cases it was cut by 25 bps, in 11 cases by 50 bps, and on 2 occasions by 75 bps (both in early 2008: the weeks of January 25 and March 21).

Standard tests indicate that it is not possible to reject the null hypothesis that the policy interest rate has a unit root. For this reason an error correction specification is used in the analysis that follows. This is standard in the literature on interest rate dynamics.[11] In addition, and not surprisingly, it is not possible to reject the hypothesis that the federal fund's rate "Granger causes" the

FIGURE 1
U.S. Federal Funds Target Rate, $r^{us}_{ff,t}$, 1994–2008



EMEs' policy rates; on the other hand, the null that these rates "cause" Fed policy actions may be rejected at conventional levels. The details of these tests are not reported due to space considerations, but are available on request.

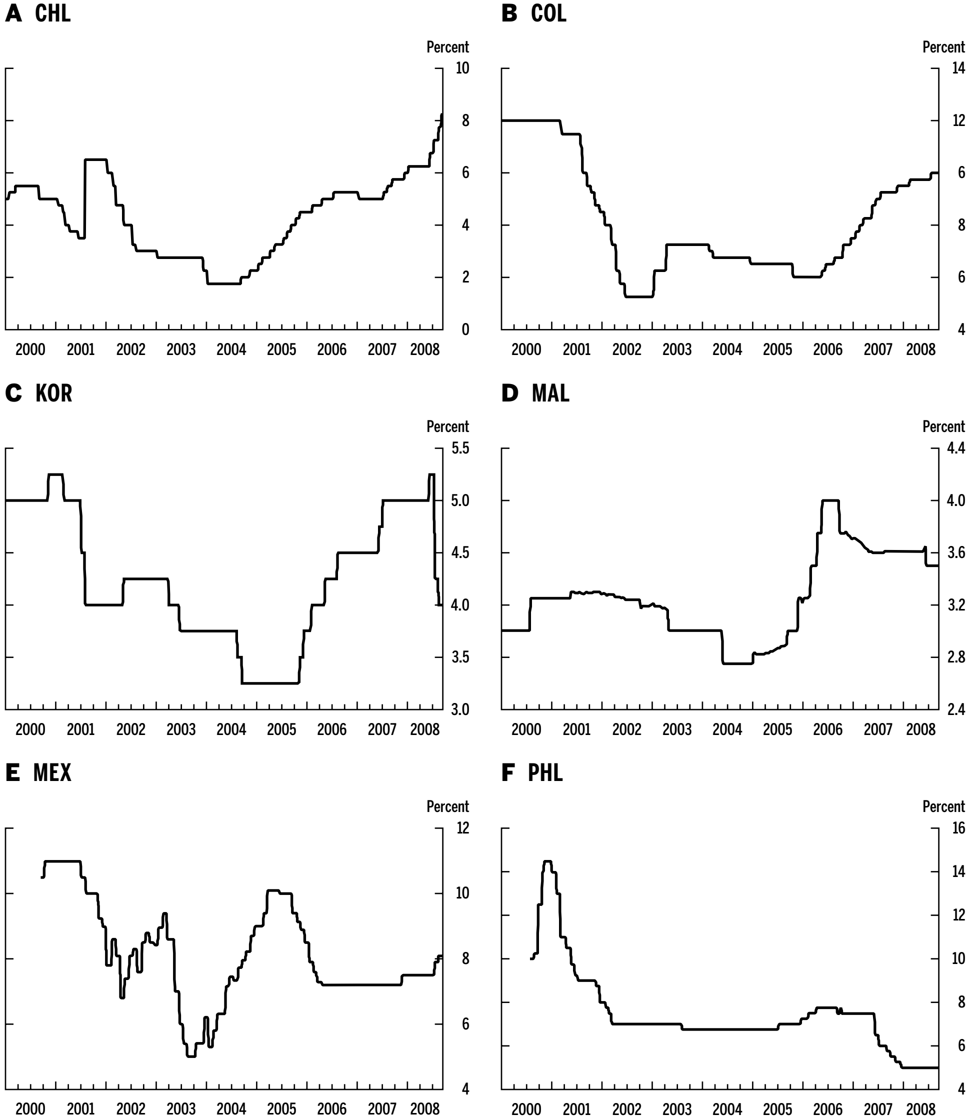
3. Policy "Contagion"

Consider a small open economy with risk-neutral investors. Assume further in order to simplify the exposition that there are controls on capital outflows in the form of a tax rate of τ.[12] Then, the following condition will hold in equilibrium:

$$\frac{r_t - r_t^*}{(1 + r_t^*)} = E_t\{\Delta e_{t+1}\} - (1 + E_t\{\Delta e_{t+1}\})\tau, \tag{1}$$

where r_t and r_t^* denote domestic and foreign interest rates for securities of the same maturity and equivalent credit risk, respectively, and $E_t\{\Delta e_{t+1}\}$ is the expected rate of depreciation of the domestic currency. This assumes perfect substitutability of local and foreign securities. If domestic and foreign assets are not perfect substitutes, r_t^* could be multiplied by some parameter $\theta \neq 1$. In a country with a credible fixed exchange rate, $E_t\{\Delta e_{t+1}\} = 0$, and full capital mobility, i.e., $\tau = 0$, then $r_t \approx r_t^*$. That is, the local interest rate (in domestic currency) will not deviate from the foreign interest rate. Under these circumstances, changes in the foreign interest rate will be transmitted one-to-one to the local

FIGURE 2
Monetary Policy Rates in Latin America and East Asia, 2000–08



economy rate. It is in this sense that there cannot be an independent monetary policy with (credible) pegged exchange rates; i.e., the local central bank cannot set the domestic rate of interest in the long run. With limited capital mobility, i.e., $\tau \geq 0$, then there will be an equilibrium wedge between domestic and international interest rates.

Under flexible exchange rates, however, $E_t\{\Delta e_{t+1}\} \neq 0$, and local rates may deviate from the international interest rate. Assume that there is a tightening of monetary policy in the foreign country—i.e., the Fed raises the federal funds target rate—that results in a higher r_t^*. Under pegged exchange rates this would be translated into a one-to-one increase in r_t; this is so even if $\tau \geq 0$. However, if the exchange rate is flexible, it is possible that r_t will remain at its initial level and that all of the adjustment takes place through an expected appreciation of the domestic currency, $E_t\{\Delta e_{t+1}\} < 0$. As Dornbusch (1976) argued in his famous "overshooting" paper, for this to happen it is necessary for the local currency to depreciate on impact by more than in the long run. Under flexible rates then, the exchange rate will act as a "shock absorber" and will fluctuate volatilely in response to foreign shocks.[13]

In order to avoid "excessive" exchange rate volatility, the local central bank may take into account the foreign central bank's actions when determining its own policy rate. That is, its policy rule (e.g., the Taylor rule) will include a term related to the foreign central bank's policy rate.[14] In a world with two countries, this situation is captured by the following two policy reaction equations, where r_p is the policy rate in the domestic country, r_p^* is the policy rate in the foreign country, and x and x^* are vectors with other determinants of policy rates, such as the deviation from an inflation target and the deviation of the unemployment rate from its "natural" rate, and the coefficients β, β^* reflect the sensitivity of each country's policy rate to the other country's rate:

$$r_p = \alpha + \beta r_p^* + \gamma x \tag{2}$$

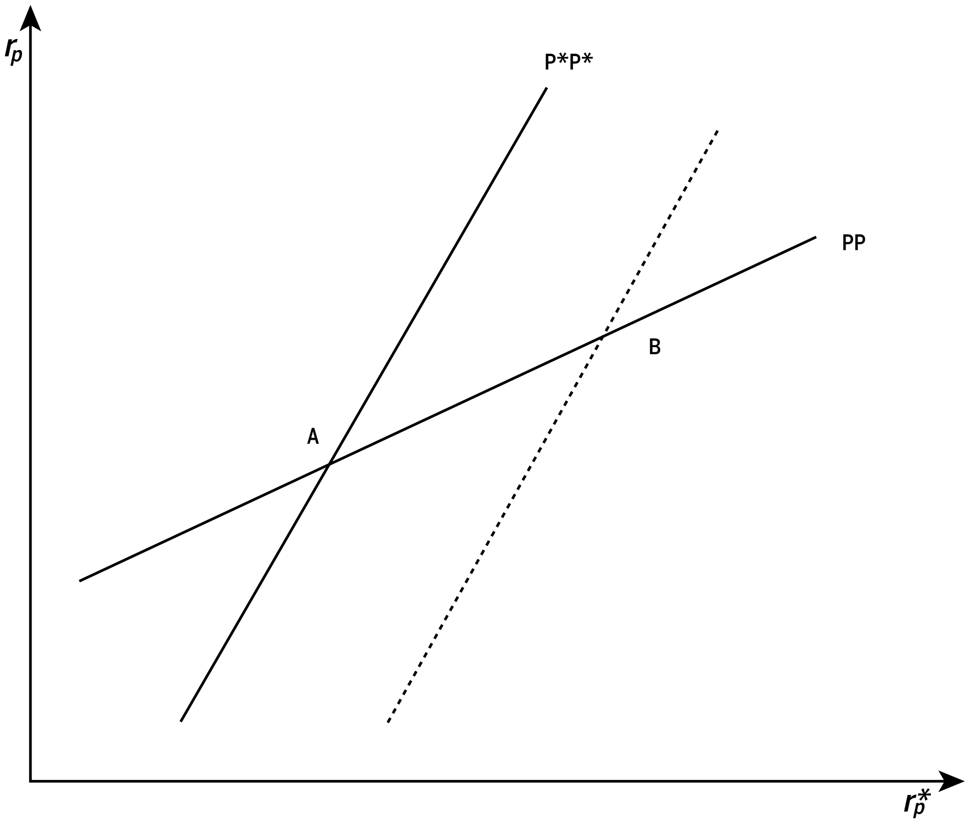
$$r_p^* = \alpha^* + \beta^* r_p + \gamma^* x^*. \tag{3}$$

Solving equations (2) and (3) gives the equilibrium monetary policy rate in each country. For the domestic country the equilibrium policy rate is (there is an equivalent expression for the foreign country)[15]

$$r_p = \frac{\alpha + \beta\alpha^*}{1-\beta\beta^*} + \left(\frac{\gamma}{1-\beta\beta^*}\right)x + \left(\frac{\beta\gamma^*}{1-\beta\beta^*}\right)x^*. \tag{4}$$

Equation (4) shows how changes in the drivers of the foreign country's policy interest rate, such as α^* or x^*, will affect the equilibrium domestic policy rate. This interdependence is illustrated in Figure 3, which includes both reaction functions (2) and (3), where PP depicts the policy function for the domestic country and P*P* for the foreign nation. The initial equilibrium is at point A. A higher level of x^* (because, say, there is a gap between the actual and target inflation rate in the foreign country) will result in a shift to the right of P*P*,

FIGURE 3
Policy Rate Equilibrium with Policy "Contagion" and Large Countries



resulting in a higher equilibrium policy rate in both countries; the new equilibrium is given by point B.[16] Notice that in this case the increase in the equilibrium foreign policy rate is *amplified*; it is larger than what was originally intended by the foreign central bank, as given by the horizontal shift of the P*P* curve.

Figure 3 assumes both countries take into consideration the other nation's actions, i.e., $\beta > 0$ and $\beta^* > 0$. But this need not be the case. Indeed, if one country is large (e.g., the United States) and the other one is small (e.g., Colombia), we would expect policy "contagion" to be a one-way phenomenon. In this case, if the foreign country is the large one, β^* in equation (2) will be zero, and the P*P* curve will be vertical. A hike in the foreign country's policy rate then will impact the domestic country rate, but there will be no feedback to the large nation and thus no amplifying effect.[17] The magnitude of the "policy spillover" depends on the slope of the PP curve, as given by the β parameter. The steeper is this curve, the larger is the extent of policy "contagion." If, on the contrary, the PP curve is very flat, policy "contagion" is minimal. In the limit, when there is complete policy independence in both countries, the PP curve is horizontal and the P*P* curve is vertical.

In traditional analyses central banks take into account the direct determinants of inflation (and unemployment, if that is part of their mandate) and there is no role for the foreign policy rate when determining the domestic policy stance, i.e., $\beta = \beta^* = 0$. The simple theoretical framework above suggests how, if $\beta > 0$ and/or $\beta^* > 0$, policy "contagion" may exist.[18] Ultimately, the extent to which specific countries are affected by policy "contagion" is an empirical matter. The rest of this paper deals with this empirical question.

4. An Empirical "Contagion" Model of Monetary Policy

This section reports results from the estimation of monetary policy rate equations for the six countries in the sample. It is assumed that each central bank has a policy function of the form of equation (2), and that it does not necessarily adjust its policy rate instantaneously to new information, including to changes in policy rates in the advanced nations. More specifically, the following error correction model allows central banks to make adjustments at a gradual pace:

$$\Delta r_{p,t} = \alpha + \beta r_{ff,t}^{us} + \theta \Delta r_{p,t-1} + \delta r_{p,t-1} + \Sigma y_j x_{j,t} + \varepsilon_t, \tag{5}$$

where $r_{p,t}$ is the policy rate in each of the six countries in period t, $r_{ff,t}^{us}$ is the federal funds policy target rate, and $x_{j,t}$ contains other variables that affect central bank actions and would typically be included in a Taylor-type policy rule, such as domestic inflation pressure and the unemployment gap. If there is policy "contagion" from abroad, the estimated β would be significantly positive. The extent of long-term policy spillover is given by $-(\frac{\beta}{\delta}) > 0$, as long as $\delta < 0$, i.e., changes in the policy rate are mean-reverting. If $-(\frac{\beta}{\delta}) = 1$, then there is full importation of Fed policies into the domestic policy rate. In this case, monetary autonomy would be greatly reduced. The parameter θ measures the extent to which the equilibrium policy rate is cyclical, with current changes depending on past changes. In equation (5) the timing of the explanatory variables is contemporaneous with the dependent variable. However, in the estimations alternative lag structures are considered for different explanatory variables.

As noted above, the vector x in equation (5) may include the direct effect of domestic inflation and unemployment gaps on the domestic policy rate. It may also include variables that indirectly depend on the U.S. policy rate. For example, domestic inflation (or its deviations from target) may depend on import prices and the rate of depreciation of the domestic currency, a variable that, in turn, depends on the interest rate differential between the domestic country and the United States. Another possible channel may exist when EME monetary authorities believe that the Federal Reserve has superior knowledge and/or information about world economic conditions, including global monetary pressures and/or the evolution of commodity prices. In this case the EME central bank may follow the Fed's action in a way similar to how firms follow a "barometric price leader" in the industrial organization literature. The analysis below seeks to disentangle these effects and assess whether the federal funds rate has an independent effect on EME policy rates, even when other variables (e.g., domestic inflationary pressures, U.S. expected inflation, and expected depreciation of the domestic currency) are held constant.

The following section reports results for the six countries in the sample for a basic specification where the only covariate, in addition to lagged values of the policy rate (in levels and first differences), is the federal funds target rate. These bivariate estimates provide a preliminary look at the correlation between the policy rates in these countries and the U.S. rates. This is followed by multivariate regressions with additional covariates added for the Latin American countries (Section 4.2) and Asian countries (Section 4.3).

4.1. Basic Results

Table 1 reports estimation results for a basic bivariate error correction specification of equation (5) for all six countries, using least squares. The federal funds variable is entered with a one-week lag.

The main insights from Table 1 may be summarized as follows:

- In five of the six countries the estimated coefficient β for the impact effect of a change in the federal funds rate, $r^{us}_{ff,t}$, is positive and significant, the

TABLE 1

Monetary Policy Rates in Latin America and East Asia, 2000–08

	Chile	Colombia	Mexico	Korea	Malaysia	Philippines
$\Delta r^{us}_{ff,t}$	0.016	0.016	0.004	0.007	0.002	0.018
	[2.384]**	[3.373]***	[0.590]	[3.215]***	[2.363]**	[2.146]**
C	0.044	0.055	0.090	0.062	0.022	0.088
	[1.505]	[2.055]**	[1.589]	[2.609]***	[1.494]	[1.646]
$r_{p,t-1}$	–0.024	–0.015	–0.013	–0.020	–0.008	–0.020
	[–2.610]***	[–3.588]***	[–1.854]*	[–3.072]***	[–1.695]*	[–2.377]**
$\Delta r_{p,t-1}$	0.005	–0.027	0.004	–0.010	0.159	–0.000
	[0.100]	[–0.525]	[0.073]	[–0.202]	[3.171]***	[–0.004]
Memo: $-\beta/\delta$	0.667	1.067	0.308	0.350	0.250	0.900
Observations	390	387	403	387	387	357
R-squared	0.019	0.038	0.009	0.030	0.043	0.018

Note: *, **, and *** refer to significance at 10 percent, 5 percent, and 1 percent, respectively.

exception being Mexico. This provides some preliminary evidence suggesting that during the period under study there may have been some policy "contagion" from the United States to these EMEs. Though generally significant, the magnitude of the impact effect during the first week of a Fed change on foreign policy rates is small. This is not surprising, as the timing of central bank meetings does not necessarily coincide across countries.

- Only Malaysia has a significant coefficient for $\Delta r_{p,t-1}$, suggesting non-monotonic adjustment of the equilibrium policy rate. This finding may be related to Malaysia's regime of maintaining a relatively rigid exchange rate during this period, in contrast to the other countries in the sample.
- The estimated long-run "contagion" effect, $-(\frac{\beta}{\delta})$, ranges from 0.25 to 1.0. The magnitude of this long-run effect is 0.67 for Chile, 1.07 for Colombia, 0.35 for Korea, 0.25 for Malaysia, insignificantly different from zero for Mexico, and 0.90 for the Philippines (throughout the paper I use rounding when discussing the results). The result that U.S. policy did not affect Mexico's central bank stance during this period is somewhat surprising, given the proximity of the two countries and the traditional dependence of Mexico's economy on U.S. economic developments. This issue is investigated in greater detail in the next subsection.
- Finally, it should be noted that the R-squared in all cases is quite low, as is usually the case for interest rate regressions in first differences.

4.2. Latin America

This subsection reports results from multivariate estimates of equation (5) for the three Latin American countries in the sample that includes, in addition to the federal funds target rate and dynamic terms, the following covariates $x_{j,t}$: (1) INFL_t, the year-over-year inflation rate, lagged between four and six weeks. Its coefficient is expected to be positive, as central banks tighten policy when domestic inflation increases; (2) $\mathrm{EXP\_DEPREC}_t$, expected currency depreciation, measured as the annualized difference between the three-month forward exchange rate and the spot exchange rate relative to the U.S. dollar, lagged one to three periods. To the extent that central banks are concerned about the value of the currency, its coefficient is expected to be positive; (3) $\mathrm{US\_EXP\_INFL}_t$, a measure of expected global inflationary pressures, defined as the breakeven spread between the U.S. five-year Treasury securities and five-year U.S. Treasury Inflation-Protected Securities (TIPS). This is entered with a one-period lag, and its coefficient is expected to be positive.[19] In addition, some of the regressions include an indicator of regional risk EMBI_LATAM, defined as the

Emerging Market Bond Index (EMBI) spread for Latin America over U.S. Treasury securities, lagged one period. Its expected sign is not determined a priori and depends on how a central bank reacts to changes in perceived regional risk.

4.2.1. OLS Estimates

The Latin American country results using least squares are reported in Table 2 (analogous results for the Asian countries are reported in Section 4.3). As may be seen, most coefficients are significant at conventional levels and have the expected signs. The most salient findings in Table 2 may be summarized as follows:

- The coefficient on the federal funds rate (β) is significantly positive for all three Latin American countries, indicating that during the period under study there was pass-through of U.S. policy rates to central banks' policy interest rates. These coefficients are positive and significant, even when expected devaluation, country risk, and global covariates are included in the regressions. Interestingly, once these other covariates are included, the coefficient on the federal funds rate for Mexico becomes significantly positive (remember that it was insignificant in the reduced-form estimates in Table 1). This suggests that changes in Fed policy rates may

TABLE 2

Monetary Policy Rates in Latin America, 2000–08

	Chile	Colombia	Mexico	Chile	Colombia	Mexico
$r^{us}_{ff,t-1}$	0.014	0.035	0.017	0.026	0.030	0.020
	[1.983]**	[3.572]***	[1.807]*	[2.679]***	[2.939]***	[1.631]*
C	–0.113	–0.375	–0.282	–0.342	–0.222	–0.339
	[–1.306]	[–3.001]***	[–1.536]	[–2.220]**	[–1.536]	[–1.469]
$r_{p,t-1}$	–0.031	–0.047	–0.053	–0.035	–0.050	–0.054
	[–3.007]***	[–3.941]***	[–4.191]***	[–3.302]***	[–4.188]***	[–4.198]***
US_EXP_INFL$_{t-k}$	0.059	0.115	0.169	0.124	0.070	0.187
	[1.844]*	[3.156]***	[2.943]***	[2.569]**	[1.644]*	[2.592]***
EXP_DEPREC$_{t-k}$	0.008	–0.002	0.026	0.005	–0.001	0.026
	[1.698]*	[–0.468]	[2.984]***	[1.174]	[–0.175]	[2.879]**
$\Delta r_{p,t-1}$	0.004	–0.043	–0.018	–0.002	–0.050	–0.019
	[0.078]	[–0.843]	[–0.351]	[–0.047]	[–0.994]	[–0.357]
INFL$_{t-k}$	0.018	0.059	0.027	0.018	0.065	0.026
	[1.647]*	[4.368]***	[1.735]*	[1.674]*	[4.712]***	[1.343]
EMBI_LATAM$_{t-1}$	—	—	—	0.012	–0.010	0.004
				[1.974]**	[–2.079]**	[0.409]
Memo: $-\beta/\delta$	0.452	0.745	0.321	0.743	0.600	0.370
Observations	389	387	351	389	387	351
R-squared	0.035	0.086	0.068	0.043	0.096	0.069
F-statistic	2.324	5.924	4.197	2.463	5.740	3.613
Durbin-Watson	1.996	1.994	2.006	1.994	1.993	2.011

Note: *, **, and *** refer to significance at 10 percent, 5 percent, and 1 percent, respectively.

be transmitted to domestic policy rates through channels other than the effect of $r^{us}_{ff,t}$ on exchange rates and/or domestic inflation.[20]

- The extent of long-term policy "contagion," measured by $-(\frac{\beta}{\delta})$, is large. The point estimates for the long-run effect range from (.017/.053=) 0.32 for Mexico, to 0.45 for Chile, to 0.74 for Colombia. However, in none of the three cases is the pass-through one-to-one. The null hypothesis that $-(\frac{\beta}{\delta}) = 1$ is rejected at conventional levels.
- The coefficients on the interest rate terms in equation (5)—β, δ, γ—can be used to calculate the medium-term effects on the domestic policy rate over time. Consider a 50 basis point increase in the federal funds rate. According to the estimates in the first three columns in Table 2, the pass-through after 24 weeks is 13 bps in Chile, 27 bps in Colombia, and 12 bps in Mexico. After 52 weeks, the policy rate is almost 25 bps in Chile, 38 bps higher in Colombia, and 16 bps higher in Mexico.[21]

The coefficients of the other variables (γ_i) generally have the anticipated signs and are significant at conventional levels. These results indicate that, as expected, with other things given, inflationary pressures—both domestic and global—result in higher policy rates. The same is true for expected depreciation in Chile and Mexico. Higher regional risk affects the policy rate significantly and positively in Chile and negatively in Colombia.

A possible limitation of the results in Table 2 is that the regressions do not include a measure of domestic activity and hence do not take into account the possibility that the central bank reacts to the evolution of the real economy. This is because there are no high-frequency (weekly) indicators for real output. In order to allow for this possibility, Table 3 reports variant regressions that include the change of each country's main commodity export price (P_COMMOD) as a proxy for real activity.[22,23] The variable is defined as the cumulative change in the commodity price over the prior six months, lagged one period.

The results in Table 3 indicate that the finding of policy "contagion" reported in Table 2 is maintained, as are the results for the other regressors. In addition, the coefficient of the change in export prices is positive as expected in all regressions, though significant only for Colombia and Mexico. This suggests that as the global market for commodity exports strengthens and prices increase and boost the local economy, the central bank will tend to reduce liquidity through a hike in its own policy rate.

4.2.2. Robustness and Extensions

This subsection reports various extensions and robustness checks.

TABLE 3
Monetary Policy Rates in Latin America, 2000–08: Role of Commodity Prices

	Chile	Colombia	Mexico	Panel
$r^{us}_{ff,t-1}$	0.026	0.029	0.025	0.014
	[2.675]***	[2.933]***	[2.048]**	[3.107]***
C	–0.342	–0.209	–0.213	–0.187
	[–2.219]**	[–1.456]	[–0.933]	[–2.011]***
$r_{p,t-1}$	–0.035	–0.050	–0.055	–0.020
	[–3.300]***	[–4.186]***	[–4.246]***	[–4.662]***
US_EXP_INFL$_{t-1}$	0.124	0.066	0.157	0.076
	[2.568]**	[1.662]*	[2.154]**	[2.580]***
EMBI_LATAM$_{t-1}$	0.012	–0.010	0.004	0.002
	[1.792]*	[–2.090]**	[0.449]	[0.494]
EXP_DEPREC$_{t-k}$	0.006	–0.001	0.031	0.006
	[1.180]	[–0.135]	[3.595]***	[2.254]***
$\Delta r_{p,t-1}$	–0.002	–0.033	–0.008	–0.003
	[–0.038]	[–0.649]	[–0.159]	[–0.086]
INFL$_{t-k}$	0.018	0.064	0.005	0.015
	[1.572]*	[4.657]***	[0.271]	[2.964]***
Δln(P_COMMOD)$_{t-1}$	0.003	0.007	0.081*	0.030
	[0.609]	[2.773]***	[1.595]	[1.297]
Memo: $-\beta/\delta$	0.743	0.580	0.455	0.700
Observations	389	387	351	1127
R-squared	0.044	0.114	0.071	0.027
F-statistic	2.198	6.072	3.270	3.854

Notes: Commodity prices measured by copper price for Chile and WTI oil price for Colombia and Mexico. *, **, and *** refer to significance at 10 percent, 5 percent, and 1 percent, respectively.

Instrumental Variables. Given the structure of lags and the nature of the covariates included in the analysis, it is unlikely that the results are affected by endogeneity issues. For countries such as Chile, Colombia, and Mexico, the federal funds rate, the TIPS rate, and global commodity prices are all clearly exogenous to their monetary policy decisions. Someone could argue, however, that in a specification where some covariates enter contemporaneously, the expected currency depreciation variable may be endogenous. In order to address this concern, Table 4 reports instrumental variable versions of the equations in Table 2 and 3 for individual countries and a pooled panel with fixed effects. The results confirm the results reported previously, in that during the period under consideration all three countries were subject to considerable policy "contagion."[24]

Lag Structure. In the regressions in Tables 1, 2, and 3, the federal funds rate is entered with a one-period (week) lag. However, the results are virtually identical if the contemporaneous federal funds rate is used instead.[25] The results are also basically unaffected if the estimation period is altered somewhat, or if the *effective* federal funds rate is used instead of the *target* policy rate.[26]

Additional Global Financial Variables. An interesting question is whether other policies related to global economic conditions enter the policy rules of

TABLE 4
Monetary Policy Rates in Latin America, 2000-08: Controlling for Endogeneity

	Chile	Colombia	Mexico	Panel
$r^{us}_{ff,t}$	0.025	0.021	0.034	0.015
	[2.532]**	[1.705]*	[2.182]**	[3.093]***
C	–0.353	–0.197	–0.301	–0.202
	[–2.249]**	[–1.315]	[–1.185]	[–2.187]**
$r_{p,t-1}$	–0.028	–0.040	–0.066	–0.022
	[–2.401]**	[–2.467]**	[–3.277]***	[–3.989]***
US_EXP_INFL$_{t-1}$	0.113	0.062	0.201	0.081
	[2.345]**	[1.432]	[2.812]***	[2.768]***
EMBI_LATAM$_t$	0.015	–0.010	0.006	0.002
	[2.006]**	[–2.066]**	[0.679]	[0.580]
EXP_DEPREC$_t$	–0.009	–0.004	0.042	0.007
	[–0.702]	[–0.505]	[1.785]*	[1.969]**
$\Delta r_{p,t-1}$	–0.006	–0.051	–0.026	–0.007
	[–0.119]	[–1.012]	[–0.481]	[–0.228]
INFL$_{t-4}$(–4)	0.020	0.058	0.001	0.015
	[1.679]*	[4.095]***	[0.018]	[3.071]***
Memo: $-\beta/\delta$	0.83	0.525	0.515	0.682
Observations	378	384	351	1119
R-squared	0.040	0.091	0.063	0.020
F-statistic	2.165	5.333	2.719	3.950
Durbin-Watson	1.992	1.997	2.005	2.013

Notes: Equations estimated with instrumental variables. Pooled panel equation includes country fixed effects. *, **, and *** refer to significance at 10 percent, 5 percent, and 1 percent, respectively.

the Latin American countries. This issue is addressed by considering two additional covariates: the yield on the U.S. 10-year Treasury bond, $r^{us}_{tb10,t}$, and the (log of the) euro–U.S. dollar exchange rate. Adding the 10-year Treasury yield allows investigation of whether Latin American central banks react to changes in the longer-maturity portion of the global yield curve. The results (available upon request) suggest that this is not the case. However, for Colombia and Mexico the coefficient of the (one-period lagged) euro–dollar exchange rate is significantly positive. The inclusion of these variables does not affect the main findings regarding policy "contagion" earlier.

Negative and Positive Policy "Contagion." Another important question is whether the extent of policy "contagion" depends symmetrically on federal funds rate increases and decreases. Investigating this issue by separating out positive and negative funds rate changes in the regressions does not support the existence of asymmetrical responses.

Short-Term Deposit Rates. Also investigated is the extent to which Fed policies were translated into (short-run) financial market rates. The results of this exercise—also available on request—show that there is a significant and fairly rapid pass-through from Federal Reserve policies to three-month CD rates in the three Latin American countries. This is the case even after controlling for expected depreciation, country risk, and global financial conditions such as the dollar–euro exchange rate and commodity prices. For a detailed analysis of this issue see, for example, Edwards (2012) and the literature cited there.

4.3. East Asia

This subsection reports results for the three East Asian countries in the sample—Korea, Malaysia, and the Philippines—and compares them with the results for Latin America.

4.3.1. OLS Estimates

The basic multivariate results for the East Asian countries are reported in Table 5. Several aspects are worth noticing.

- First, even after controlling for other variables, the coefficient of the federal funds rate is significantly positive, indicating that during the period under study there was "contagion" from the United States to these three Asian nations.
- Second, the long-run effects measured by $-(\frac{\beta}{\delta})$ are somewhat smaller than in the case of the Latin American countries, and range from 0.66

TABLE 5
Monetary Policy Rates in East Asia, 2000–08

	Korea	Malaysia	Philippines	Panel
$r^{us}_{ff,t-1}$	0.015	0.002	0.054	0.003
	[1.662]	[2.489]*	[3.384]**	[1.303]
C	0.144	0.064	0.937	0.067
	[2.302]*	[1.981]*	[2.982]**	[1.067]
$r_{p,t-1}$	–0.032	–0.017	–0.083	–0.005
	[–2.173]*	[–2.579]*	[–4.013]**	[–1.989]*
US_EXP_INFL$_{t-1}$	–0.010	–0.001	–0.222	–0.004
	[–0.682]	[–0.275]	[–2.330]*	[–0.237]
EMBI_ASIA$_{t-1}$	–0.018	–0.005	–0.035	–0.015
	[–2.542]*	[–1.342]	[–0.627]	[–1.238]
EXP_DEPREC$_{t-1}$	0.033	–0.003	0.080	0.023
	[1.022]	[–0.452]	[2.873]**	[1.953]
$\Delta r_{p,t-1}$	–0.012	0.142	–0.157	–0.034
	[–0.215]	[2.766]**	[–1.812]	[–0.962]
INFL$_{t-3}$	–0.009	0.007	0.251	–0.022
	[–0.359]	[0.545]	[1.921]	[–0.699]
Memo: $-\beta/\delta$	0.469	0.118	0.651	0.600
Observations	378	378	357	1269
R-squared	0.036	0.061	–0.886	–0.206
F-statistic	2.701	3.442	8.817	2.024

Notes: Pooled panel equation includes country fixed effects. *, **, and *** refer to significance at 10 percent, 5 percent, and 1 percent, respectively.

for the Philippines to 0.13 for Malaysia, the only country in the sample with a relatively rigid exchange rate for most of the sample period. A possible explanation is that Malaysia had restricted capital mobility during this period (see the discussion in Section 5). The first difference of the lagged policy rate is only significant for Malaysia. Additional lagged terms were insignificant in all three countries. The estimates for the additional covariates are less precise than for the Latin American countries. The coefficient of U.S. expected inflation is significant for the Philippines but is surprisingly negative. The coefficient of domestic inflation is significantly positive for Malaysia and in one of the Korean regressions. Expected depreciation is significant also only in the case of the Philippines.

An important difference between the Latin American and East Asian nations is that the former are commodity exporters, while the latter export manufacturing goods. Table 6, the East Asia analogue of Table 3, includes a measure of changes in commodity prices, P_COMMOD_t, proxied by an international index for energy prices. As before, the impact coefficients for the federal

TABLE 6
Monetary Policy Rates in East Asia, 2000–08: Role of Commodity Prices

	Korea	Malaysia	Philippines	Panel
$r^{us}_{ff,t-1}$	0.009	0.002	0.025	0.002
	[2.521]**	[2.873]***	[3.019]***	[3.504]***
C	0.150	0.051	0.322	0.047
	[2.452]**	[2.149]**	[2.044]**	[2.213]**
$r_{p,t-1}$	–0.026	–0.015	–0.037	–0.010
	[–2.792]**	[–2.836]**	[–3.807]**	[–3.399]**
$US\_EXP\_INFL_{t-1}$	–0.015	–0.001	–0.106	0.000
	[–1.144]	[–0.142]	[–2.278]**	[0.109]
$EMBI\_ASIA_{t-1}$	–0.016	–0.004	0.040	–0.004
	[–2.623]**	[–1.557]	[1.473]*	[–1.691]*
EXP_DEPREC_{t-1}	0.002	–0.000	0.011	–0.000
	[0.939]	[–0.332]	[3.549]**	[–0.483]
$\Delta r_{p,t-1}$	–0.025	0.141	–0.020	0.041
	[–0.493]	[2.853]***	[–0.406]	[1.448]
INFL_(–3)	–0.013	0.009	0.009	0.004
	[–1.780]*	[2.227]**	[0.319]	[1.249]
$\Delta \ln(P\_COMMOD)_{t-6}$	0.139	–0.012	–0.062	0.007
	[1.732]*	[–0.388]	[–0.221]	[0.258]
Memo: $-\beta/\delta$	0.346	0.133	0.676	0.200
Observations	387	439	409	1287
R-squared	0.061	0.059	0.066	0.020
F-statistic	3.092	3.393	3.554	2.647

Notes: Pooled panel equation includes country fixed effects. *, **, and *** refer to significance at 10 percent, 5 percent, and 1 percent, respectively.

funds rates are significantly positive but small. The long-run point estimates for the pass-through from U.S. monetary policy to domestic monetary policy (with other things given) are 0.35 for Korea, merely 0.13 for Malaysia, and 0.68 for the Philippines.

One of the most salient results in this analysis comes from comparing the panel results for Asia with the panel estimates for Latin America (the last columns in Tables 6 and 3, respectively). As may be seen, in both cases the coefficient of the federal funds rate is significantly positive. However, it is much smaller in Asia than in Latin America (0.002 versus 0.015), indicating that the impact effect of a federal funds rate on policy has historically been significantly stronger in the Latin American countries. This is also the case for the long-run effects of policy "contagion." According to these results, in the long run the three Latin American countries "imported" more than two-thirds of the federal funds changes. Pass-through for the Asian countries is significantly lower, at 0.20. That is, with everything else constant, a federal funds increase of 350 bps would have been translated, on average, into a policy rate increase of 240 bps in Latin America. In Asia the pass-through would have amounted, on average, to a mere 70 bps.

4.3.2. Robustness and Extensions

The data for the East Asian countries were subject to the same battery of extensions and robustness tests as those for Latin America. Instrumental variable estimates are presented in Table 7. As may be seen, the main message from the previous results is maintained. In particular, even after controlling for other relevant variables, there is evidence of "partial contagion" from the Fed to these three Asian nations. In all three countries the long-run pass-through coefficient is significantly lower than one, though the magnitude varies across countries. In the Philippines the extent of pass-through was around two-thirds. In Malaysia it is barely higher than 10 percent (0.13).

The results on the extensions and robustness tests discussed above maintained the basic finding reported in Tables 5, 6, and 7, in the sense that during the period under consideration there indeed has been some—although far from full—policy "contagion" for this group of East Asian nations.

5. "Contagion" and Capital Mobility

The specification of equation (1) assumed that there was a tax of rate τ on capital outflows leaving the country. Alternatively, it is possible to think that there is a tax on capital inflows, of the type popularized by Chile during the 1990s.[27] In this case equation (1) becomes[28]

TABLE 7
Monetary Policy Rates in East Asia, 2000–08: Controlling for Endogeneity

	Korea	Malaysia	Philippines	Panel
$r^{us}_{ff,t}$	0.015	0.002	0.054	0.003
	[1.662]*	[2.489]**	[3.384]***	[1.303]
C	0.144	0.064	0.937	0.067
	[2.302]**	[1.981]**	[2.982]***	[1.067]
$r_{p,t-1}$	–0.032	–0.017	–0.083	–0.005
	[–2.173]**	[–2.579]**	[–4.013]***	[–1.989]**
US_EXP_INFL$_{t-1}$	–0.010	–0.001	–0.222	–0.004
	[–0.682]	[–0.275]	[–2.330]**	[–0.237]
EMBI_ASIA$_t$	–0.018	–0.005	–0.035	–0.015
	[–2.542]**	[–1.342]	[–0.627]	[–1.238]
EXP_DEPREC$_t$	0.033	–0.003	0.080	0.023
	[1.022]	[–0.452]	[2.873]***	[1.973]**
$\Delta r_{p,t-1}$	–0.012	0.142	–0.157	–0.034
	[–0.215]	[2.766]**	[–1.812]	[–0.962]
INFL$_{t-1}$	–0.009	0.007	0.251	–0.022
	[–0.359]	[0.545]	[1.921]*	[–0.699]
Memo: $-\beta/\delta$	0.469	0.118	0.651	0.600
Observations	378	378	357	1269
R-squared	0.036	0.061	–0.886	–0.206
F-statistic	2.701	3.442	8.817	2.024

Notes: Equations estimated by instrumental variables. Pooled equations include country fixed effects. *, **, and *** refer to significance at 10 percent, 5 percent, and 1 percent, respectively.

$$r_t - r^*_t(1-\iota) + \iota = E_t\{\Delta e_{t+1}\}, \qquad (1')$$

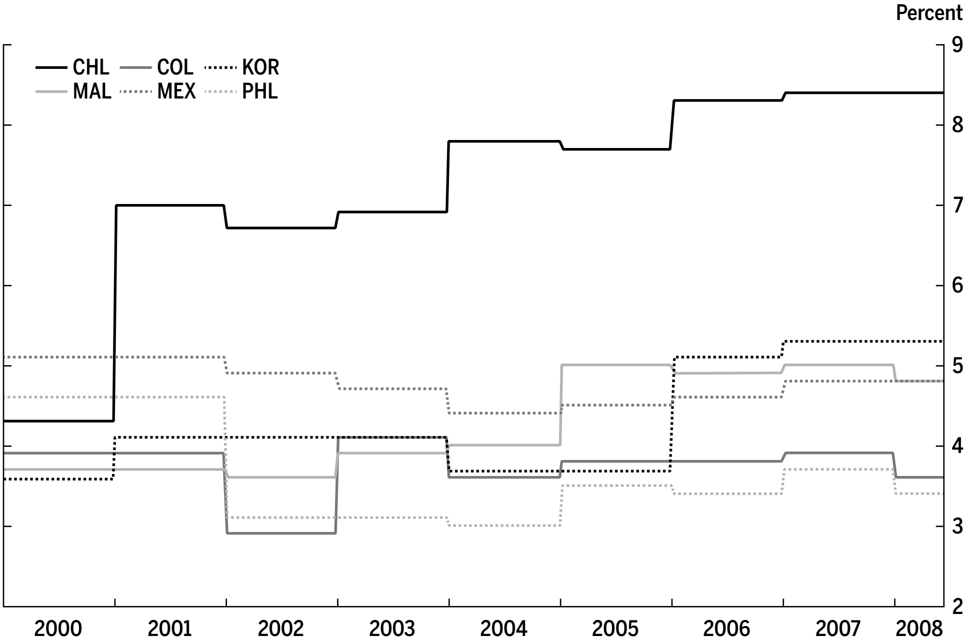
where ι is the tax rate on capital inflows.

As pointed out above, the six countries in this study had varying degrees of capital mobility during the period 2000–08, with Chile being the most open and the Philippines being the least open to capital movement. In addition, during the (almost) 500 weeks covered by the analysis, there were some adjustments to the extent of mobility in all six nations. This was especially the case of Chile, which, after concluding a free trade agreement with the United States in early 2001, opened its capital account further.

Figure 4 shows the evolution of a comprehensive measure of capital mobility for each of the countries in the analysis, constructed using data from the Fraser Institute. A higher number denotes a higher degree of capital mobility.

An important question is whether the degree of capital mobility affects the extent of pass-through from federal funds rates to policy interest rates in emerging countries. In order to address this issue, a number of reduced-form equations are estimated similar to those reported in Table 1, with two additional regressors: an index of capital mobility (CAP_MOBILITY) and a variable that interacts this index with the federal funds rate. The capital mobility index varies for the Asian countries in the sample from a minimum of 3 for the

FIGURE 4
Capital Mobility Index for Selected Latin American and Asian Countries, 2000–08



Philippines to a maximum of 5.3 for Korea, with an average of 4.0 and median of 3.9. The rather limited variation in the mobility index within each country means that it is difficult to estimate country-specific regressions. For this reason, results are reported for two pooled panels: one for the three Latin America nations and one for the three East Asian countries. The results reported in Table 8 should be considered preliminary and subject to further research for a number of reasons, including, in particular, the fact that the index of capital mobility is an aggregate summary that includes different types of capital controls. To understand better the role of mobility for interest rate pass-through, it is necessary to construct more detailed and granular indexes. Second, in order to investigate this issue fully, a broader sample that includes countries with greater restrictions is necessary.

The results in Table 8 confirm the findings reported in previous tables of pass-through from the federal funds rate to domestic policy rates. However, it is important to interpret the results with caution. As may be seen, the capital mobility measure is positive and significant when entered on its own in the Latin America sample; in this case the federal funds coefficient continues to be

TABLE 8
Monetary Policy Rates in Latin America and East Asia, 2000–08: The Role of Capital Mobility

	Latin America	Latin America	Latin America	East Asia	East Asia	East Asia
$r^{us}_{ff,t-1}$	0.014	0.008	0.023	0.002	–0.003	0.017
	[3.993]***	[1.064]	[2.324]**	[3.178]***	[–1.337]	[0.886]
C	–0.028	0.050	–0.062	0.029	0.054	0.015
	[–0.694]	[2.645]***	[–1.163]	[1.705]*	[3.332]***	[0.183]
$r_{p,t-1}$	–0.016	–0.015	–0.016	–0.012	–0.013	–0.019
	[–4.610]**	[–4.420]**	[–4.679]**	[–3.489]**	[–3.790]**	[–3.566]**
$\Delta r_{p,t-1}$	–0.009	–0.006	–0.010	0.044	0.043	–0.004
	[–0.313]	[–0.214]	[–0.332]	[1.486]	[1.461]	[–0.111]
CAP_MOBILITY$_t$	0.016	—	0.023	0.005	—	0.019
	[2.199]**		[2.246]**	[2.409]**		[0.817]
CAP_MOBILITY$_{t^*}$	—	0.001	–0.002	—	0.001	–0.003
$r^{us}_{ff,t-1}$		[0.872]	[–0.981]		[2.698]**	[–0.562]
Memo: $-\beta/\delta$	0.875	0.533	1.438	0.167	–0.231	0.895
Observations	1127	1127	1127	1131	1131	744
R-squared	0.024	0.020	0.024	0.023	0.024	0.023
F-statistic	4.501	3.794	4.001	4.400	4.636	2.882

Notes: *, **, and *** refer to significance at 10 percent, 5 percent, and 1 percent, respectively. Pooled equations include country fixed effects.

significant and positive as well. The interactive variable, however, is not significant in any of the Latin American regressions.

The most interesting result in Table 8 is for the Asian countries. The coefficient of the federal funds rate on its own is not significantly different from zero, with a point estimate of –0.0028. However, the coefficient of the federal funds rate interacted with the capital mobility term is positive, with a point estimate of 0.0013. This suggests that countries with higher mobility had higher pass-through. The capital mobility index varies for the Asian countries in the sample from a minimum of 3 for the Philippines to a maximum of 5.3 for Korea. This suggests that the extent of pass-through in countries with capital mobility within this span would range from 0.30 to 0.53. Once again these estimates are significantly lower than those for the Latin American nations, suggesting that the extent of "contagion" for Asia has historically been lower than for Latin America.

6. The Fed and Short-Term Market Interest Rates in Latin America and Asia

This section expands the analysis by investigating the extent to which Federal Reserve actions affected short-term (90-day) deposit rates in Latin America and Asia. More specifically, it addresses four questions: First, what has been the effect of Fed actions on domestic market interest rates? Second, what are the channels through which these effects work? Do they work solely through the effects on domestic policy rates that were unearthed in the previous section, or is there an additional channel?[29] Third, how (if at all) have changes in the slope of the U.S. yield curve affected domestic interest rates? Fourth, has the transmission of Fed actions to domestic market interest rates differed in magnitude and speed in Latin America and Asia? These questions are addressed by adding the yield on the 10-year U.S. Treasury note as an additional covariate in the estimations.

6.1. The Data

Tables 9 and 10 present data on the impact and lagged average changes in the short-term (three-month) deposit rate in Latin America, Asia, and the United States during the weeks following changes in the federal funds policy rate. To take into account possible asymmetric responses, Table 9 gives results for increases in the federal funds rate, while Table 10 gives results for cuts in the

TABLE 9

Average Cumulative Changes in Short-Term Deposit Interest Rates in Latin America, East Asia, and United States in Weeks following Federal Funds Policy Rate Increases, 2000–08

(in basis points; standard deviation in parentheses)

	Impact	1 week	2 weeks	3 weeks	6 weeks
Latin America	14	14	14	12	15
	(111)	(99)	(100)	(115)	(150)
East Asia	2	5	7	5	9
	(29)	(33)	(39)	(54)	(78)
United States	3	6	9	12	25
	(2)	(3)	(4)	(6)	(12)

Note: Average Federal Reserve funds rate increase during sample period is 26 bps.

TABLE 10

Average Cumulative Changes in Short-Term Deposit Interest Rates in Latin America, East Asia, and United States in Weeks following Federal Funds Policy Rate Cuts, 2000–08

(in basis points; standard deviation in parentheses)

	Impact	1 week	2 weeks	3 weeks	6 weeks
Latin America	17	5	15	9	–12
	(82)	(89)	(115)	(145)	(161)
East Asia	–1	8	4	–15	–6
	(84)	(121)	(93)	(72)	(113)
United States	–20	–21	–24	–30	–54
	(22)	(26)	(28)	(29)	(29)

Notes: Latin America is defined as Chile, Colombia, and Mexico; East Asia is defined as Korea, Malaysia, and the Philippines. Average Federal Reserve fund rate cut over sample period is 44 bps.

federal funds rate. The average increase in the Fed policy rate during the sample period was 26 bps, while the average cut was 44 bps. The key results are as follows:

- Not surprisingly, the response of deposit rates differs between the countries in Latin America and Asia. Changes in short-term rates in Latin America are greater (in absolute terms) than in Asia, on average, particularly in the case of federal funds rate hikes. However, none of these effects—either after federal funds hikes or cuts—are statistically significant.
- After six weeks there appears to be a one-to-one transmission of the Fed's action to U.S. 90-day deposit rates. This is the case for both federal funds increases and federal funds cuts: The average rate hike of 26 bps is followed by a deposit rate increase of 26 bps, while the average rate cut of 44 bps is followed by a deposit rate decrease of 54 bps.
- Following a federal funds target rate reduction, short-term interest rates in both Latin America and Asia are somewhat volatile, exhibiting both increases and decreases. After six weeks, however, there is a net decline in short-term rates in both regions. These decreases average 12 bps in Latin America and only 6 bps in Asia.

To summarize, Tables 9 and 10 indicate that after six weeks federal funds policy rates changes are transmitted fully into changes in short-term deposit rates in the United States. In contrast, the changes in short-term deposit rates abroad are very small; indeed, they are not significantly different from zero in either region. The next section analyzes whether these results are maintained in regressions that control for the role of other covariates.

6.2. Regression Results

This section reports the results from error correction models for short-term market deposit rates in Latin America and Asia. The specifications reported here are similar (but not identical) to those for policy rates in Section 4. The dependent variable is the first difference of weekly interest rates, and (most) of the covariates are the same as those used in the preceding analysis with policy interest rates. The regressions in Table 11 do not include the domestic policy rate in order to enable estimation of the total pass-through from the Fed policies to market rates, after the local central bank has reacted fully to the Fed action. With regard to the question of interest, the main findings may be summarized as follows: In every country, except Korea, there is evidence during the sample period of pass-through from Fed policy to domestic short-term deposit interest rates. In most cases the impact effect is very small, not surprising in

TABLE 11
Short-Term Deposit Rates in Latin America and East Asia, 2000–08

	Chile	Colombia	Mexico	Korea	Malaysia	Philippines
$r^{us}_{ff,t}$	0.050	0.047	0.116	0.001	0.002	0.210
	[4.119]***	[5.296]***	[5.108]***	[0.111]	[1.721]*	[5.499]***
C	0.250	0.336	0.271	0.085	0.045	–0.009
	[2.684]***	[4.773]***	[1.743]*	[1.350]	[2.106]**	[–0.023]
EXP_DEPREC	0.015	0.017	0.074	–0.015	–0.001	0.108151
	[1.629]*	[3.040]***	[5.080]***	[–4.668]**	[–0.602]	[6.929]***
$r_{90d,t-1}$	–0.058	–0.037	–0.099	–0.015	–0.014	–0.255
	[–3.949]***	[–4.252]***	[–6.363]***	[–2.387]**	[–2.314]**	[–8.8015]**
$EMBI_t$	–0.029	0.022	0.025	0.001	–0.0032	0.140
	[–1.33]	[3.756]***	[0.646]	[0.042]	[–1.110]	[2.718]***
$\Delta r_{90d,t-1}$	0.0621	0.029	0.011	0.444	0.206	0.032
	[1.096]	[0.619]	[0.244]	[7.134]***	[4.304]***	[0.643]
$r^{us}_{tb10,t}$	–0.027	–0.078	–0.062	–0.005	–0.000	0.069
	[–1.186]	[–3.751]***	[–1.400]	[–0.448]	[–0.049]	[0.864]
Memo: $-\beta/\delta$	0.859	1.283	1.178	0.041	0.153	0.824
Observations	397	452	452	187	422	345
R-squared	0.123	0.095	0.089	0.404	0.063	0.208
F-statistic	6.813	7.809	7.273	20.341	4.679	14.814

Notes: The dependent variable is the change in the 90-day deposit rate, $\Delta r_{90d,t}$, with the exception of Malaysia, where the one-year deposit rate is used. *, **, and *** refer to significance at 10 percent, 5 percent, and 1 percent, respectively.

view of the weekly data frequency. However, the effect builds over time, and in the long run it appears to be quite sizable in all countries, with the exception of Malaysia. The long-run point estimates exceed one, but are not significantly different from one, in Colombia and Mexico, and are slightly lower in Chile, 0.86, and the Philippines, 0.82. For Malaysia, the point estimate of the long-run pass-through is a low 0.15.

Table 12 presents results when the domestic policy rate is included as a covariate. This allows analysis of whether there was interest rate pass-through from the Fed to local market rates, even when the domestic policy rate is held constant. Alternatively, these results provide information as to the extent to which the domestic central bank can neutralize the effect of Fed policy changes on local interest rates. The coefficient for the domestic policy rate is significant and positive in four of the six countries; surprisingly, it is marginally negative in the case of Mexico and the Philippines, though it is insignificant in the latter case. A comparison of Tables 11 and 12 shows that in all countries the point estimate for the federal funds rate effect is smaller when the policy rate is included in the regression than when it is excluded (in the Philippines the coefficients are not significantly different). The simultaneous inclusion of the domestic and Fed policy rates allows decomposition of the transmission effect of federal funds rate changes into direct and indirect effects. The results suggest that

TABLE 12
Short-Term Deposit Rates and Domestic Monetary Policy Rates in Latin America and East Asia, 2000–08

	Chile	Colombia	Mexico	Korea	Malaysia	Philippines
$r^{us}_{ff,t}$	0.027808 [2.1204]***	0.006751 [0.6288]	0.100406 [4.6195]***	0.000057 [0.0106]	0.003572 [2.9946]***	0.213631 [5.0225]***
C	0.519667 [4.5813]***	0.237618 [3.4168]***	0.593594 [2.7694]***	0.124370 [1.9452]*	0.014436 [0.6966]	–0.008582 [–0.0214]
EXP_DEPREC$_t$	0.005755 [0.5484]	0.006032 [1.0850]	0.077020 [5.1340]***	–0.014230 [–4.5506]***	–0.000378 [–0.3671]	0.108628 [6.8655]***
$r_{90d,t-1}$	–0.140211 [–5.5776]***	–0.090198 [–7.5380]***	–0.070061 [–4.2336]***	–0.051005 [–3.3588]***	–0.109426 [–6.5196]***	–0.254177 [–8.6837]***
EMBI	–0.093599 [–3.5060]***	0.036215 [5.9054]***	–0.014894 [–0.4174]	–0.015447 [–0.4707]	0.001639 [0.6630]	0.145301 [2.5072]**
$\Delta r_{90d,t-1}$	0.096566 [1.7255]*	0.010856 [0.2429]	0.017768 [0.3648]	0.478838 [7.6434]***	0.213766 [4.7022]***	0.030802 [0.6175]
$r_{p,t}$	0.121452 [3.9753]***	0.073390 [6.1999]***	–0.042073 [–1.7891]*	0.041804 [2.5999]**	0.096787 [6.0635]***	–0.011630 [–0.1945]
$r_{tb10,t-1}$	–0.075028 [–2.9870]***	–0.055419 [–2.7190]***	–0.085927 [–1.8729]*	–0.014209 [–1.1368]	0.006096 [2.0367]**	0.078558 [0.8329]
Memo: $-\beta/\delta$	0.198	0.075	1.433	0.011	0.033	0.840
Observations	397	452	416	187	430	345
R-squared	0.1690	0.1673	0.0856	0.4257	0.1381	0.2083
F-statistic	8.3953	12.7476	5.4566	18.9584	9.6599	12.6673

Notes: The dependent variable is the change in the 90-day deposit rate, $\Delta r_{90d,t}$, with the exception of Malaysia, where the one-year deposit rate is used. *, **, and *** refer to significance at 10 percent, 5 percent, and 1 percent, respectively.

the transmission mechanism varies across countries. In Chile and Malaysia the effect of changes in the federal funds rate on the local deposit rate takes place both indirectly, through changes in the domestic policy rate (see the discussion in the preceding sections), and directly, as shown in Table 12. For Korea and Colombia there is no direct effect, only an indirect channel through the domestic policy rate, since once the domestic policy rate is included in the regression, the coefficient for the federal funds rate is no longer significant. For Mexico, in contrast, the effect appears to be only direct. In unreported regressions, the inclusion of the capital mobility index does not help explain these differences. A possible explanation could be related to the role of the banking sector and the extent to which global banks play a role in each country.

To illustrate the nature of these results more fully, consider the case of Chile reported in Table 12. As may be seen, the coefficients of both the federal funds and domestic policy rates are significantly positive, though the magnitudes of the point estimates are very different—0.028 and 0.121, respectively. The long-run pass-through effect of a change in the federal funds rate is (.028/.14=) 0.2, implying that an increase in the federal funds rate of 100 bps would be translated in the long run into an increase in the short-term deposit rate of 20 bps, as long as the domestic policy rate and other covariates remain given. However, the analysis in the preceding sections suggests that in the case of Chile there is a long-run pass-through into the domestic policy rate of approximately 0.67 (see Table 1). If one adds both effects, the total pass-through from the federal funds rates into short-term market rates would be about 76 bps in the long run. This estimate is slightly smaller than the one obtained as a total effect from Table 11. Another way of thinking about the results is that the Central Bank of Chile could offset the effect of a 100 bps hike in the federal funds rate by reducing its own policy rate by 22 bps. Of course, this discussion regarding Chile does not necessarily apply to the other countries in the sample, as given the differences in results, each needs to be analyzed individually.

Similarly to the case of policy rates, the results for short-term deposit rates were subjected to a battery of robustness tests, including panel and instrumental variables estimation. These results are not reported here due to space considerations, but they confirm prior findings.

7. Concluding Remarks

In November 2015 it was expected that the Federal Reserve would raise interest rates in December. That was to be the first federal funds rate hike since 2006. An important question was—and continues to be—how the tightening process will affect the emerging markets. This paper attempts to provide an answer to this question by investigating the extent to which Fed policy actions have in the past been passed into monetary policy interest rates in three East Asian and three Latin American nations—Korea, Malaysia, the Philippines, Chile, Colombia, and Mexico.

The basic estimates suggest that Federal Reserve interest rate changes are imported, on average, into all six countries in the analysis: by 45 percent in Korea, only 12 percent in Malaysia (the only country in the sample with a rigid exchange rate during most of the period), 65 percent in the Philippines, 74 percent in Colombia, more than 45 percent in Chile, and 32 percent in Mexico. Thus, if the Federal Reserve were to hike rates by a cumulative total of 325 bps—bringing the federal funds rate to 3.5 percent—we could expect that (with other things given) Colombia would hike policy rates by 250 bps, Chile by approximately 150 bps, and Mexico by more than 100 bps; in Asia the estimates for the average policy rate adjustments (with everything else constant) would be 146 bps in Korea, merely 40 bps in Malaysia, and 200 bps in the Philippines. There is no evidence that the extent of policy "contagion" depends on the degree of capital mobility in Latin America. In contrast, there is some evidence that the more open countries in East Asia have had a higher degree of pass-through.

The finding of a non-zero pass-through from the Fed to monetary policy in the five countries in the sample with exchange rate flexibility is important for the debate on optimal exchange rate regimes. Indeed, according to traditional models one of the key advantages of flexibility is that it permits a country to conduct its own monetary policy. The results in this paper question that principle by finding that in almost all countries in the sample there is a fairly high degree of policy “contagion.” A possible explanation for the results reported in this paper is a “fear of floating” that is not captured fully by the covariates included in the analysis.[30] According to models in the Mundell-Fleming tradition, if there is less than perfect capital mobility, a hike in the global interest rate—generated by, say, Federal Reserve action—will result in an incipient external deficit and a depreciation of the domestic currency. Indeed, currency adjustment is what reestablishes equilibrium. If, however, there is a “fear of floating,” the local authorities will be tempted to tighten their own monetary stance by hiking the policy rate as a way of preventing the weakening of the domestic currency. Further investigation along these lines could shed additional light on the question of the “true” degree of monetary independence in small countries with flexible exchange rates. A particularly important point that follows from this analysis is that to the extent that smaller countries import policies from advanced-country central banks—such as the Federal Reserve—that are destabilizing, this may create a more volatile macroeconomic environment in EMEs.[31]

8. Data Sources

The following are the data sources used in the paper:

Interest rates: Policy rates were obtained from various issues of the national central bank of each country. Data on U.S. Treasury securities and the federal funds rate were also obtained from Datastream. All figures correspond to the Friday of a given week.

Exchange rates: Defined as domestic currency per U.S. dollar. Expected devaluation is constructed as the 90-day forward discount, also relative to the dollar. The euro–dollar rate is defined as euros per dollar. Source: Datastream.

Commodity price indexes: Source: JP Morgan.

Regional risk: Defined as the Emerging Market Bond Index (EMBI) premium for Latin America or Asia over U.S. Treasury securities, measured in percentage points. Source: Datastream.

REFERENCES

Aizenman, Joshua, Mahir Binici, and Michael M. Hutchison. 2014. "The Transmission of Federal Reserve Tapering News to Emerging Financial Markets." National Bureau of Economic Research Working Paper 19980.

Calvo, Guillermo A., and Carmen M. Reinhart. 2000. "Fear of Floating." National Bureau of Economic Research Working Paper 7993.

Cetorelli, Nicola, and Linda S. Goldberg. 2011. "Global Banks and International Shock Transmission: Evidence from the Crisis." *IMF Economic Review* 59(1), pp. 41–76.

Claessens, Stijn, Hui Tong, and Shang-Jin Wei. 2012. "From the Financial Crisis to the Real Economy: Using Firm-Level Data to Identify Transmission Channels." *Journal of International Economics* 88(2), pp. 375–387.

Claro, Sebastián, and Luis Opazo. 2014. "Monetary Policy Independence in Chile." Bank for International Settlements Paper 78.

De Gregorio, José, Sebastian Edwards, and Rodrigo O. Valdés. 2000. "Controls on Capital Inflows: Do They Work?" *Journal of Development Economics* 63(1), pp. 59–83.

Dincer, N. Nergiz, and Barry Eichengreen. 2013. "Central Bank Transparency and Independence: Updates and New Measures." Working Paper.

Dornbusch, Rudiger. 1976. "Expectations and Exchange Rate Dynamics." *Journal of Political Economy* 84(6), pp. 1161–1176.

Edwards, Sebastian. 2006. "The Relationship between Exchange Rates and Inflation Targeting Revisited." National Bureau of Economic Research Working Paper 12163.

Edwards, Sebastian. 2012. "The Federal Reserve, the Emerging Markets, and Capital Controls: A High-Frequency Empirical Investigation." *Journal of Money, Credit and Banking* 44(s2), pp. 151–184.

Edwards, Sebastian. 2015. "Monetary Policy Independence under Flexible Exchange Rates: An Illusion?" National Bureau of Economic Research Working Paper 20893.

Edwards, Sebastian, and Eduardo Levy Yeyati. 2005. "Flexible Exchange Rates as Shock Absorbers." *European Economic Review* 49(8), pp. 2079–2105.

Edwards, Sebastian, and Roberto Rigobón. 2009. "Capital Controls on Inflows, Exchange Rate Volatility and External Vulnerability." *Journal of International Economics* 78(2), pp. 256–267.

Eichengreen, Barry J., and Poonam Gupta. 2014. "Tapering Talk: The Impact of Expectations of Reduced Federal Reserve Security Purchases on Emerging Markets." World Bank Policy Research Working Paper 6754.

Frankel, Jeffrey, Sergio L. Schmukler, and Luis Servén. 2004. "Global Transmission of Interest Rates: Monetary Independence and Currency Regime." *Journal of International Money and Finance* 23(5), pp. 701–733.

Glick, Reuven, Ramon Moreno, and Mark M. Spiegel. 2001. *Financial Crises in Emerging Markets.* Cambridge University Press.

Goldberg, Linda S. 2009. "Understanding Banking Sector Globalization." *IMF Staff Papers* 56(1), pp. 171–197.

Obstfeld, Maurice, Jay C. Shambaugh, and Alan M. Taylor. 2005. "The Trilemma in History: Tradeoffs among Exchange Rates, Monetary Policies, and Capital Mobility." *Review of Economics and Statistics* 87(3), pp. 423–438.

Prasad, Eswar S., Kenneth Rogoff, Shang-Jin Wei, and M. Ayan Kose. 2003. "Effects of Financial Globalization on Developing Countries: Some Empirical Evidence." International Monetary Fund Occasional Paper 220, March 17.

Rey, Hélène. 2013. "Dilemma not Trilemma: The Global Financial Cycle and Monetary Policy Independence." Paper presented at the Federal Reserve Bank of Kansas City Economic Symposium "Global Dimensions of Unconventional Monetary Policy," August 21–23, Jackson Hole, WY.

Spiegel, Mark M. 1995. "Sterilization of Capital Inflows through the Banking Sector: Evidence from Asia." Federal Reserve Bank of San Francisco *Economic Review* (3), pp. 17–34.

Taylor, John B. 2009. "Globalization and Monetary Policy: Missions Impossible." In *International Dimensions of Monetary Policy*, eds. Jordi Galí and Mark Gertler. University of Chicago Press and NBER, pp. 609–624.

Taylor, John B. 2013. "International Monetary Coordination and the Great Deviation." *Journal of Policy Modeling* 35(3), pp. 463–472.

Taylor, John B. 2015. "Rethinking the International Monetary System." Remarks at the CATO Institute Monetary Conference on Rethinking Monetary Policy, Washington, DC, November 12.

Tesar, Linda L. 1991. "Savings, Investment and International Capital Flows." *Journal of International Economics* 31(1), pp. 55–78.

NOTES

1 See, for example, Aizenman, Binici, and Hutchison (2014) and Eichengreen and Gupta (2014) for discussion of the effects of the tapering on EMEs.

2 For most Latin American countries the Fed action was seen as a contributing factor to the depreciation of their currencies. In mid-January 2016, Agustín Carstens, the governor of the Bank of Mexico, suggested that Latin American central banks should form a common front to deal with further Fed action.

3 The International Monetary Fund's 2015 *World Economic Outlook* contains a long discussion of this issue.

4 On the trilemma, see Obstfeld, Shambaugh, and Taylor (2005) and Rey (2013).

5 On "fear of floating" see Calvo and Reinhart (2000).

6 I thank John Taylor for making the transcript of Ron McKinnon's remarks available to me.

7 See Glick, Moreno, and Spiegel (2001) and Prasad et al. (2003) on crises in emerging markets. On the role of the banking sector in a context of capital mobility, see Spiegel (1995) and Goldberg (2009). On capital flows in a context that emphasizes savings and investment decisions, see Tesar (1991). On the propagation of financial distress to the real economy, see Claessens, Tong, and Wei (2012).

8 Indeed, both discussants of this paper—Linda Tesar and Woon Gyu Choi—have worked on large-scale models to analyze related issues. Their approaches, however, differ from the one taken in this paper in various respects. In particular, and importantly, this paper deals with the effects of monetary policy transmission at the individual country level. Most work on the subject either pools data for many nations or considers a "representative" emerging country.

9 Aizenman, Binici, and Hutchison (2014) analyzed how the announcement of Federal Reserve tapering in 2013 affected financial conditions in emerging markets.

10 For indexes of central bank transparency and independence, see Dincer and Eichengreen (2013).

11 See, for example, Frankel, Schmukler, and Servén (2004) and Edwards (2012) for analyses of the transmission of interest rate shocks. Those studies differ from the current paper in a number of respects, including the fact that they concentrate on market rates and do not explore the issue of policy "contagion." Other differences are the periodicity of the data—this paper uses weekly data—and the fact that individual countries are analyzed. Rey (2013) deals with policy interdependence, as does Edwards for the case of a single country, Chile.

12 Parts of this section draw on Edwards (2015).

13 The shock absorber role of the exchange rate goes beyond cushioning against monetary disturbances. For example, Edwards and Levy Yeyati (2005) show that countries with more flexible rates are better able to accommodate terms-of-trade shocks.

14 Edwards (2006) argues that many countries include the exchange rate as part of their policy (or Taylor) rule. Taylor (2009, 2013) has argued that many central banks include other central banks' policy rates in their rules. The analysis that follows in the rest of this section owes much to Taylor's work.

15 The stability condition is $\beta\beta^* < 1$. This means that in Figure 3 the P*P* schedule is steeper than the PP schedule.

16 The new equilibrium will be achieved through successive approximations, as in any model with reaction functions of this type, where the stability condition is met.

17 Of course, if neither country considers the other foreign central bank's actions, PP will be horizontal and P*P* will be vertical.

18 Notice that "contagion" is in quotation marks, both here and in the paper's title. This is because many central banks strongly resist the notion that their decisions are affected by what other central banks are doing. From a theoretical point of view, it is easy to derive models where the optimal policy would include reacting to "the world" interest rate.

19 However, it is possible to argue that once the federal funds rate is included, the coefficient of the spread between Treasury securities and TIPS should be zero, since the federal funds rate already incorporates market expectations of U.S. inflation.

20 In a recent paper Claro and Opazo (2014) argue that the Central Bank of Chile has been fully independent and has not directly responded to Fed policy moves.

21 Most (but not all) central banks conduct policy by adjusting their policy rates in multiples of 25 bps. The estimates discussed here refer to *averages*. Thus, they need not be multiplied by 25 bps.

22 There are no weekly data on real activity. However, there is significant evidence that the evolution of prices of a major commodity export is a good leading indicator of economic performance in emerging market economies.

23 The weekly price of copper is used for Chile, a combination of coffee and oil prices for Colombia, and the West Texas Intermediate (WTI) oil index for Mexico. All data are from JP Morgan.

24 The instrument set includes the log of lagged commodity prices (copper, coffee, metals, energy, WTI oil), the lagged U.S. dollar–euro rate, lagged (six weeks) effective depreciation, lagged expected depreciation, and lagged rates for U.S. government assets of varying maturities.

25 The data refer to the end of the week (Friday). Since the Federal Open Market Committee never meets on a Friday, Fed actions precede in time the end-of-week recording of the local interest rate.

26 In order to check whether the last few months in the sample—the months leading to the Lehman Brothers crisis—affected the results, I reestimated the regressions for a shorter period. No significant changes were found.

27 On the Chilean tax on capital inflows, see De Gregorio, Edwards, and Valdés (2000) and Edwards and Rigobón (2009).

28 See, for example, Edwards (2012).

29 An interesting question that is beyond the scope of this paper is whether global banks play a role in the magnitude or speed of transmission of interest rate shocks. On this issue see, for example, Goldberg (2009) and Cetorelli and Goldberg (2011).

30 Calvo and Reinhart (2000) is the classic reference on this subject.

31 For a discussion along these lines see, for example, Edwards (2012), Rey (2013), and Taylor (2013).