# Exchange Rate Pass-Through to Domestic Prices: Does the Inflationary Environment Matter?

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## Abstract

The paper tests a hypothesis suggested by John Taylor (2000) that a low inflationary environment leads to a low exchange rate pass-through to domestic prices. To test this hypothesis, the paper derives a pass-through relation based on new open economy macroeconomic models. A large database that includes 1979–2000 data for 71 countries is used to estimate this relation. There is strong evidence of a positive and significant association between the pass-through and the average inflation rate across countries and periods. The inflation rate, moreover, dominates other macroeconomic variables in explaining cross-regime differences in the pass-through.

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# I. Introduction

The degree to which changes in the exchange rate pass through to prices is an important issue in debates about appropriate monetary and exchange rate policies. A low exchange rate pass-through is thought to provide greater freedom for pursuing an independent monetary policy and to make it easier to implement inflation targeting. There is no consensus, however, on the conditions that lead to a low pass-through. Traditional literature has been concerned with the exchange rate pass-through to import prices and has stressed the role of market power and price discrimination in international markets (pricing to market).[2] This literature suggests that the import price pass-through (at the sectoral or aggregate level) is essentially determined by microeconomic factors (e.g., demand elasticities, market structure) that are exogenous to monetary policy.[3]

An alternative view has been put forth by John Taylor (2000), who argues that the recently-observed declines in the pass-through to aggregate prices are the result of a low inflation environment. Taylor explains the link between inflation and pass-through in terms of a model of firm behavior based on staggered price setting and monopolistic competition. As firms set prices for several periods in advance, their prices respond more to cost increases (due to exchange rate depreciation or other sources) if cost changes are perceived to be more persistent. Regimes with

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[2] See, for example, Goldberg and Knetter (1997) for a survey of this literature.

[3] A related strand in the literature emphasizes the role of local versus producer currency pricing in the determination of the exchange rate pass-through. The pricing method chosen by foreign firms is generally regarded as exogenous to domestic monetary policy. See, however, Devereux and Engel (2001) who endogenize this choice and relate it to monetary stability.

higher inflation tend to have more persistent costs.[4] A high inflation environment would thus tend to increase the exchange rate pass-through. In this view, the pass-through depends on the policy regime: a credible low inflation regime will automatically achieve a low pass-through.

Recently, there has been a growing interest in examining the relationship of the exchange rate pass-through with monetary policy behavior and the inflationary environment. A number of recent studies find some empirical support for the relationship but the evidence is not conclusive. Campa and Goldberg (2001) provide evidence on the pass-through to import prices based on data for OECD countries. They find that although higher inflation and exchange rate volatility are positively associated with higher import price pass-through, microeconomic factors related to the composition of imports play a much more important role in determining the pass-through. The import price pass-through reflects the price behavior of foreign firms and this behavior may not be strongly related to the home inflationary environment. Thus evidence on the pass-through to domestic prices (e.g., CPI) would provide a more appropriate test of the Taylor view. Gagnon and Ihrig (2001) explore the relationship between CPI pass-through and inflation stabilization for eleven industrial countries.[5] They find that the pass-through generally declined in the 1990's and the change in the pass-through is significantly related to the variability of inflation. They also explore but do not find a systematic relation between the pass-through and the monetary policy behavior.

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[4] Taylor (2000) provides evidence showing that the disinflation in the US has been associated with a reduction in the persistence of inflation.

[5] McCarthy (2000) uses a VAR model to estimate the pass-through to CPI for a number of countries. He shows that the pass-through tends to be larger for countries with higher exchange rate persistence, but he does not relate the pass-through to the inflationary experience of the countries.

This paper extends the literature relating pass-through to inflation in two directions. First, the paper's empirical analysis uses a large database that comprises of 1979–2000 data for 71 countries and includes developing as well as industrial countries.[6] As the inflationary experience varies much more among developing countries than among industrial countries, our database is especially suitable for exploring how inflation affects the pass-though. The theoretical link between inflation and the pass-through has not been adequately discussed in terms of a fully-specified macroeconomic model. A second contribution of the paper is to derive such a link using a general equilibrium framework based on new open-economy macroeconomic models. The theoretical framework also suggests a general form of the pass-through relation that can be implemented empirically.

Our theoretical framework incorporates Taylor-type staggered pricing, but simplifies the dynamics of the exchange and wage rates in response to monetary shocks. Representing monetary policy by a simple feedback rule, our analysis derives a negative association between the pass-through to CPI and the degree to which a monetary regime offsets short-run price deviations from the long run path. This association arises essentially because the pass-through reflects the expected effect of monetary shocks on current and future costs. A regime that reacts more aggressively to price deviations lowers the pass-through by weakening the expected future effect of the shocks. In our empirical analysis we use the long-run inflation rate as an indicator of the aggressiveness of monetary policy response to short-run price fluctuations. The use of this proxy is motivated by the plausible assumption that regimes which make a stronger effort in stabilizing the short-run inflation rate are also able to maintain low inflation rates in the long run.

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[6] Our database is similar to that used in the Goldfajn and Werlang (2000) study of the exchange rate pass-through. Their study, however, is concerned with other determinants of the pass-

(continued…)

This assumption allows us to link the CPI pass-through in an economy to its long-run inflation performance.[7]

The paper estimates the pass-through to CPI (over short as well as long periods) using an empirical model suggested by the theory. Most of the countries in our database appear to experience a single inflation regime during the sample period. We find strong evidence of a positive and significant association between the pass-through and the average inflation rate across these countries. Further evidence in support of a robust link between inflation and the pass-through is provided by a small number of countries that experienced a dramatic shift in the inflation environment. This shift was accompanied by a marked change in the pass-through for all of these countries. We also explore the influence of other macroeconomic variables on the pass-through and find that the average inflation rate dominates these variables in explaining cross-regime variation in the pass-through.

Section II discusses the theoretical framework and derives the relation between the CPI pass-through and the inflation regime. Section III discusses the methodology and presents the results of the empirical analysis. Conclusions are offered in section IV.

## II. Theoretical Framework

This section discusses the theoretical framework underlying the pass-through relation examined in this paper. The framework is based on the new open-economy macroeconomic models and incorporates imperfect competition and price inertia. We consider a variation of

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through and does not focus on the influence of the inflationary environment.

[7] Our analysis also shows that lower variability of monetary shocks reduces the information content of the exchange rate in predicting monetary shocks, and hence decreases the pass-through. If low inflation economies are also subject to less variable monetary shocks, then this effect suggests another reason for the pass-through to be smaller under a low inflation regime.

these models, which treats final consumer goods as non-traded goods and imports as intermediate inputs into the production of consumer goods.[8] We also assume a production structure suggested by Basu (1995), in which each variety can be used both as a final good and an intermediate input. Price inertia is assumed to arise from staggered price adjustment based on the Taylor model. We assume a Stiglitz-Dixit aggregator function for differentiated consumer goods and intermediate inputs. Furthermore, we assume that prices are set in producer's currency. These assumptions abstract from the role of pricing to market and local-currency pricing but allow us to simplify the analysis of the effect of the inflationary environment on the exchange rate pass-through.[9]

### A. Price Setting and the Exchange Rate Pass-Through

We consider a model comprising of a small home and a large foreign country. Each country produces a good consisting of a continuum of differentiated varieties in the unit interval. Households in both countries consume varieties produced domestically. The home sub-utility function for the differentiated consumption good is

$$U_t = \left[\int_0^1 U_t(i)^{(\sigma-1)/\sigma} di\right]^{\sigma/(\sigma-1)}. \tag{1}$$

The price index for the consumer good is given by the unit expenditure function for $U_t$ as

$$P_t = \left[\int_0^1 P_t(i)^{1-\sigma} di\right]^{1/(1-\sigma)}. \tag{2}$$

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[8] McCallum and Nelson (1999) use such a variant, for example. This variation underscores the dependence of the CPI on the price behavior of home firms. The pass-through relation that we estimate, however, would not be affected much if the CPI also includes imported goods.

[9] See Bergin and Feenstra (2001) for a model with staggered pricing that incorporates pricing to market (based on a translog aggregator function) and local currency pricing. They use this model to explain persistence in real exchange rates.

Each variety is produced with the following Cobb-Douglas production function:

$$Y_t = AZ_t^{\alpha} L_t^{1-\alpha}, \tag{3}$$

where $Y_t$ is output, $Z_t$ is a composite intermediate input and $L_t$ is the amount of labor. The composite intermediate is obtained by aggregating bundles of domestic and imported varieties via the following aggregator function:

$$Z_t = [(1-\nu)^{1/\sigma} Z_{Ht}^{(\sigma-1)/\sigma} + \nu^{1/\sigma} Z_{Mt}^{(\sigma-1)/\sigma}]^{\sigma/(\sigma-1)}, \tag{4}$$

where $Z_{Ht} = \left[\int_0^1 Z_{Ht}(i)^{(\sigma-1)/\sigma} di\right]^{\sigma/(\sigma-1)}$ and $Z_{Mt} = \left[\int_0^1 Z_{Mt}(i)^{(\sigma-1)/\sigma} di\right]^{\sigma/(\sigma-1)}$. The elasticity between domestic and imported bundles is assumed, for simplicity, to be the same as that between varieties within each bundle. Letting $\widetilde{P}_t$ denote the unit cost of $Z_t$ and $W_t$ the wage rate, and using (3), we can express each variety's marginal cost of production (for a suitable choice of *A*) as

$$C_t = \widetilde{P}_t^{\alpha} W_t^{1-\alpha}. \tag{5}$$

Using an asterisk to denote foreign variables, we assume that relations analogous to (1)–(5) hold for the foreign economy. The home variety is supplied to final and intermediate demand at home as well as intermediate demand abroad. Let $\widetilde{P}_t^*$ represent the unit cost of $Z_t^*$ in foreign currency and $S_t$ denote the exchange rate (expressed as the price of foreign currency). The demand for a home variety can then be expressed as

$$D_t(i) = \delta_t P_t(i)^{-\sigma}, \tag{6}$$

where $\delta_t = U_t P_t^{\sigma} + (1-\nu) Z_t \widetilde{P}_t^{\sigma} + \nu^* Z_t^* (S_t \widetilde{P}_t^*)^{\sigma}$.

Assume, for simplicity, that the price of each variety is fixed for 2 periods. Let $X_t$ denote the fixed price set in *t*. (Thus, for a variety *i* whose price is set in *t*, $P_{t+\tau}(i) = X_t$, for $\tau = 0, 1$).

The value of $X_t$ is chosen to maximize the expected discounted value of profits over 2 periods and can be derived using (6) as

$$X_t = E_t \frac{\sigma \sum_{\tau=0}^{1} \beta^{\tau} \delta_{t+\tau} C_{t+\tau}}{(\sigma - 1) \sum_{\tau=0}^{1} \beta^{\tau} \delta_{t+\tau}}, \tag{7}$$

where $\beta$ is the discount factor and $E_t$ denotes the expectation operator conditional on information available at time $t$.[10] We assume that firms observe the contemporaneous value of the exchange rate (the information set at $t$ includes $S_t$). All other variables are observed with a one-period lag. We also assume a uniform overlap between prices so that prices of $1/2$ home varieties (and $1/2$ foreign varieties) are set in each period. Then using (2), we can restate the price of the home consumer good as

$$P_t = \left[ (1/2) \sum_{\tau=0}^{1} X_{t-\tau}^{1-\sigma} \right]^{1/(1-\sigma)}. \tag{8}$$

Approximate versions of the key price-setting relations can be obtained by linearizing these relations around initial steady-state values. Lower-case letters are used to denote proportional deviations of variables from their initial steady-state values (e.g., $x_t = (X_t - \overline{X})/\overline{X}$, where $\overline{X}$ represents the initial steady-state value of $X_t$). Relations (7) and (8) are linearized as

$$x_t = \frac{1}{1+\beta} E_t c_t + \frac{\beta}{1+\beta} E_t c_{t+1}, \tag{9}$$

$$p_t = \frac{1}{2} x_t + \frac{1}{2} x_{t-1}. \tag{10}$$

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[10] Firms solve the problem: *max* $E_t \sum_{\tau=0}^{1} \beta^{\tau} [X_t - C_{t+\tau}] \delta_{t+\tau} X_t^{-\sigma}$. The first order condition for this problem is used to obtain (7).

Note that if $\beta \approx 1$, the price setting in (9) gives roughly equal weight to current and next-period costs, as in the standard Taylor model. In view of (4), $\widetilde{P}_t = \left[(1-\nu)P_t^{1-\sigma} + \nu(S_t P_t^*)^{1-\sigma}\right]^{1/(1-\sigma)}$, where $P_t^*$ is the aggregate price of foreign varieties in foreign currency ($S_t P_t^*$ is thus the aggregate home-currency price of imported intermediates). Using this condition and (3), (5) can be linearized as

$$c_t = \alpha\theta p_t + \alpha(1-\theta)(s_t + p_t^*) + (1-\alpha)w_t, \tag{11}$$

where $\theta = (1-\nu)\overline{P}^{1-\sigma} / [(1-\nu)\overline{P}^{1-\sigma} + \nu(\overline{SP}^*)^{1-\sigma}]$ represents the steady-state share of home varieties in intermediates inputs.

Relations (9)–(11) represent the basic channels that transmit the effect of exchange rate changes to consumer prices. A change in the exchange rate has a direct impact on costs via (11). Given staggered price adjustment process (10), however, the exchange rate change will initially pass-through to only a fraction of goods whose prices are being reset. In resetting prices according to (9), firms will take into account the influence of the exchange rate on the expected values of future costs. Expectations thus play an important role in determining the exchange rate pass-through.

To solve for the expected costs ($E_t c_{t+\tau}$), we need to add relations that determine $w_t$ and $s_t$. We consider below a simple model of household preferences and asset markets, which simplifies the dynamics of these variables. This model can be solved analytically to obtain a pass-though relation that depends on the feedback rule of the monetary authority.

### B. The Influence of the Feedback Rule

We first briefly outline household behavior. Households allocate a fixed amount of time between work and leisure and their utility depends on the differentiated good (via the sub-utility function), real money balances and labor supply. The representative household in the home country is assumed to maximize the following simple objective function:

$$E_t \sum_{t=0}^{\infty} \beta^t \left[\log U_t + \log(M_t / P_t) - L_t\right].$$

Foreign households maximize an analogous function. We also assume, that households in both countries have access to a complete set of state contingent assets (we later consider departures from this assumption). As Bergin and Feenstra (2001) show, household optimization under these assumptions implies the following linear approximations:

$$w_t = u_t + p_t = (1-\beta) m_t + \beta E_t (u_{t+1} + p_{t+1}), \tag{12}$$

$$s_t + p_t^* - p_t = u_t - u_t^* \quad . \tag{13}$$

The first equality in (12) represents the linear version of the labor-supply condition while the second equality is based on the linearized forms of the Euler and money-demand conditions. Equation (13) gives the linearized form of the risk-sharing condition whereby the ratio of home and foreign utilities is proportional to the real exchange rate as implied by the assumption of complete asset markets.

The small home economy exerts a negligible effect on foreign prices. To isolate the effects of the home monetary regime, assume that the foreign monetary regime introduces no monetary shocks so that $m_t^* = p_t^* = 0$. Under these assumptions, (12), its foreign counterpart and (13) imply that

$$w_t = s_t = m_t + \beta E_t (s_{t+1} - s_t)/(1-\beta). \tag{14}$$

The monetary authority could choose either the exchange rate or the money stock as an instrument (The second equality in (14) would then determine the value of the variable that is not chosen as the instrument).[11] To relate the pass-through to monetary policy regime in a simple way, we assume that the exchange rate is the instrument and the feedback rule takes the following simple form:

$$s_t = -\rho x_{t-1} + v_t, \tag{15}$$

where $\rho\ (>0)$ is a parameter that measures the extent to which policy reacts to price deviations, and $v_t$ is a monetary policy shock, which is assumed, for simplicity, to be white noise.[12] As discussed below, the parameter $\rho$ plays a critical role in determining the degree of exchange rate pass-through.

As shown in the Appendix, the model comprising of (9)–(11) and (14)–(15) can be solved to obtain a solution for $x_t$ of the form: $x_t = \phi_1 x_{t-1} + \phi_2 v_t$, where $\phi_1$ and $\phi_2$ depend on $\rho$, $\alpha$ and $\theta$. Then using (15) and (10), this solution can be expressed as the following pass-through relation between the price level and the exchange rate:

$$p_t = \mu_1 p_{t-1} + \mu_2 (s_t + s_{t-1})/2, \tag{16}$$

where $\mu_1 = \phi_1 + \rho\phi_2$ and $\mu_2 = \phi_2$. The short-run exchange rate pass-through (i.e., the impact of a one unit increase in $s_t$ on $p_t$) equals $\mu_2/2$. The pass-through over a longer period would also depend on $\mu_1$, which represents the degree of persistence in *p*. The short-run pass-through

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[11] Another possible instrument is the interest rate, which (in linearized form) would equal $E_t(s_{t+1} - s_t)$.

[12] The choice of the exchange rate as the instrument simplifies the algebra, but the basic results could also be obtained by using a similar feedback rule with $m_t$ as the instrument. Also, if $s_t$ reacts to $p_{t-1}$ (instead of $x_{t-1}$), then additional values of $x_t$ would be included in (15) via (10). The lagged value of $s_t$ could also be included in the rule to represent policy inertia. These

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coefficient, $\mu_2$, reflects the effect of monetary shocks on prices set in the current period. This effect depends on the anticipated future costs and hence on the expected future exchange rate. A larger value of $\rho$ would reduce the effect of the shocks on the expected exchange rate (via the feedback rule) and thus lower $\mu_2$. A bigger $\rho$ would also be expected to decrease the persistence coefficient, $\mu_1$.[13] The Appendix shows that both $\mu_1$ and $\mu_2$ decrease in $\rho$.[14] Thus a larger value of $\rho$ would decrease the exchange rate pass-through both in the short and the long run.[15]

The above example relies heavily on the special assumption of complete asset markets. Departures from this assumption arise not only because asset markets are incomplete (only limited state-contingent assets are available) but also because uncovered interest rate parity does

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possibilities are ignored here, for simplicity, but would be accommodated in our empirical model discussed below.

[13] A stronger feedback reaction would be expected to weaken the association between $x_t$ and $x_{t-1}$, and thus decrease $\phi_1$ (the persistence in $x_t$). Note, however, that the effect on $\mu_1$ is less clear since $\rho$ exerts a positive direct effect on $\mu_1$ in addition to the negative indirect effect (via the $\phi$' s).

[14] The Appendix also shows that $\mu_1$ increases and $\mu_2$ decreases in $\alpha\theta$. The effect on $\mu_1$ arises because a larger share of intermediate inputs in costs (i.e., an increase in $\alpha$) increases price-level persistence. The effect on $\mu_2$, on the other hand, reflects the influence of a decrease in the share of imported varieties in intermediate inputs (i.e., a decrease in $1-\theta$) on the pass-through in the short run.

[15] Note that the pass-through to import prices equals 1 (the aggregate import price can be linearized as $s_t + p_t^*$) and does not depend on the home inflationary environment. A limitation of our model is that it abstracts from pricing to market and local currency pricing, and thus does not provide an explanation of an incomplete import price pass-through.

not hold for several reasons.[16] To briefly explore the implications of relaxing the complete-asset-markets assumption, we modify (13) as

$$s_t + p_t^* - p_t = u_t - u_t^* + \eta_t , \tag{17}$$

where departures from complete markets are treated simply as a stochastic error and represented by the shock $\eta_t$. To simplify the analysis, assume that $\eta_t$ is also white noise and is uncorrelated with $v_t$. Using (17), its foreign counterpart and (12), and recalling that $m_t^* = p_t^* = 0$, we modify the wage-exchange-rate relation as $w_t = s_t - \eta_t$. Thus $\eta_t$ shocks loosen the link between the wage and the exchange rate.

We assume, moreover, that the feedback rule also accommodates, at least partially, non-policy shocks.[17] We thus modify rule (15) as

$$s_t = -\rho x_{t-1} + v_t + \varphi \eta_t , \tag{18}$$

where $0 < \varphi \leq 1$. The pass-through would now reflect the effect of both $v_t$ and $\eta_t$ shocks. The solution of the modified model [consisting of (9)–(11), (18) and the relation that $w_t = s_t - \eta_t$] is also discussed in the Appendix. Letting $\sigma_v^2$ and $\sigma_\eta^2$ denote the variances of the two shocks, the Appendix shows that the model implies a relation of the same form as (16), but the values of $\mu_1$ and $\mu_2$ now also depend on $\lambda \equiv \varphi^2 \sigma_\eta^2 /(\varphi^2 \sigma_\eta^2 + \sigma_v^2)$. Both $\mu_1$ and $\mu_2$ decrease in $\rho$, as before. These coefficients also decrease in $\lambda$ and hence increase in $\sigma_v^2$. The intuition for this

[16] In addition to a (possibly small) risk premium, a liquidity premium arising from "noise traders" (Jeanne and Rose, 2000) could account for a stochastic deviation from uncovered interest parity.

result is simply that the lower variability of monetary shocks reduces the information content of the exchange rate in predicting monetary shocks and thus decreases the exchange rate pass-through.

### C. Further Considerations

The above theoretical analysis relates the pass-through coefficient to the feedback rule. To test this relation, we use the long-run inflation rate as a proxy for the feedback-rule parameter, $\rho$, in our empirical analysis. The use of this proxy is based on the plausible assumption that a regime which makes a stronger effort to offset short-term price fluctuations will also tend to achieve a lower inflation rate in the long run. This assumption implies an inverse relation between $\rho$ and the long-run inflation rate, and provides a straightforward basis for establishing a positive link between the long-run inflation rate and the exchange rate pass-through.[18]

Our empirical analysis focuses on the long-run inflation rate as the key indicator of the monetary policy regime. Other macroeconomic indicators, such as measures of the exchange rate or inflation variability, have also been suggested as determinants of the exchange rate pass-though in the literature. Our model, however, does not imply a clear relation between the pass-through and the variance of $p$ or $s$ if asset markets are not complete. As shown in the Appendix, the variance of $p$ decreases in $\rho$ and increases in $\sigma_v^2$, but the sign of the partial effect of $\sigma_\eta^2$ on

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[17] For example, the policy rule could characterize a managed float regime, under which the exchange rate would be influenced by policy actions but would also respond to shocks to financial markets.

[18] This assumption is also consistent with the evidence of positive association between the inflation rate and inflation persistence since, as discussed above, our model implies that an increase in $\rho$ leads to a decrease in the persistence of $p$.

this variance is ambiguous. In the case of the variance of *s*, even the sign of the partial effect of $\rho$ is ambiguous. Thus it is not obvious that a regime with a high variance of *p* or *s* would necessarily have higher values of $\mu_1$ and $\mu_2$. The variance measures may still be empirically useful indicators of monetary policy regimes (instead of or in addition to the inflation rate) and this issue is explored in the empirical analysis below.

The above analysis focuses on the effect of temporary exchange rate changes (induced by white-noise shocks), but it can be easily extended to examine the influence of permanent changes in the exchange rate (brought about by random-walk stochastic processes). Under complete markets (without $\eta_t$ shocks), the model implies that a permanent one percent increase in the exchange rate (arising from permanent monetary policy shocks) would increase costs and prices by one percent in the long run. This result does not hold if $\eta_t$ shocks are introduced. In this case, a permanent one percent increase in the exchange rate does not imply a permanent one percent increase in costs since the exchange rate change could reflect the effect of permanent shocks to $\eta_t$. Thus the long-run pass through could be less than one in the absence of complete markets.

Although the simple pass-through relation (16) is useful in illustrating how the inflationary environment affects the exchange rate pass-through, it is derived from a special version of the model, which assumes that the feedback rule is of a simple form, prices are fixed for only two periods, shocks follow simple stochastic processes and foreign prices equal their steady state paths ( $p_t^* = 0$ ). For our empirical analysis, however, we consider a general version of the model that relaxes these assumptions. The general version suggests a pass-through relation of the following form:

$$p_t = \pi_1(L)\, p_{t-1} + \pi_2(L)s_t + \pi_3(L)p_{t-1}^* + \varepsilon_t\,, \tag{19}$$

where $\pi_1(L)$, $\pi_2(L)$ and $\pi_3(L)$ are lag polynomials and $\varepsilon_t$ is an error term representing the residual effect of shocks (i.e., the effect not captured by the first three terms). The order of the lag polynomials would depend not only on the number of periods for which prices are fixed, but also on the stochastic processes for the exogenous variable $p_t^*$ and the shocks, $v_t$ and $\eta_t$. Our empirical analysis will allow the data for an inflation regime to determine the length of each lag polynomial. In the general version, additional variables such as the lagged values of the money stock could be used to determine expected value of shocks and thus enter (19). We, therefore, also explore an augmented version of (19), which includes lagged values of $m_t$. Incomplete markets could introduce a non-stationary component in the error term, because temporary shocks would lead to permanent wealth effects in this case.[19] As discussed further below, a non-stationary error term in (19) would require estimation of this relation in first differences.

Finally, note that we have used a very simple feedback rule to distinguish inflation regimes. In reality, monetary policy regimes differ in much more complex ways. Our simple rule is intended to capture key attributes of the inflationary environment even if the policy regime does not explicitly use the exchange rate as an instrument.[20] Our analysis assumes, however, that the policy regime is stable (i.e., follows a particular rule). Thus it may not be appropriate (as discussed below) for hyperinflation situations.

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[19] Lane and Milesi-Ferretti (2000) provide evidence on the effect of net foreign assets on the real exchange rate. Although our model abstracts from productivity shocks, the presence of such shocks could also cause the real exchange rate to be non-stationary for Balassa-Samuelson reasons.

[20] Even if the regime uses other policy instruments, rule (15) or (18) could still provide a useful as if description of the extent to which the regime tolerates deviations from the long-run path of prices.

## III. Empirical Analysis

This section describes the data and the methodology used to estimate the exchange rate pass-through and discusses the basic results on the association between inflation and the pass-through. We use a large sample that consists of time series data from 1979 to 2000 for 71 countries and includes economies from all major regions.[21]

### A. Inflation Experience

Table 1 summarizes the behavior of quarterly CPI inflation (expressed as annual rates) for our sample countries. These countries exhibit a wide range of inflation experience. The mean rate of inflation varies from a low of 1.5 for Bahrain to a high of 124.4 for Brazil. There is also considerable cross-country variation in the range and the standard deviation of the inflation rate.

Countries with high inflation are more likely to experience major changes in their inflation environment. One possibility is that inflation could get out of control and lead to hyperinflation characterized by very high and accelerating inflation rates. Another possibility is that an inflation stabilization program is able to bring down inflation rates to low or moderate levels. To explore possible shifts in inflation regimes, we examined the inflation behavior of countries with mean inflation rate above 30%. For a number of these countries, high inflation represented a chronic situation over the sample period.[22] For eight countries, however, there is a clear indication of a switch in inflation regimes. These shifts are identified in Figure 1 (and Table 3). Five countries in this group – Argentina, Bolivia, Brazil, Nicaragua and Peru – experienced

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[21] The data were obtained from the Information Notice System Database of the IMF, except where it is indicated otherwise. See Zanello and Desruelle (1997) for a complete description of the database.

hyperinflation for a certain period but were successful in stabilizing inflation and maintaining moderate rates afterwards. The remaining three countries did not undergo hyperinflation but did experience a significant change in their inflationary environment. Israel and Mexico implemented successful stabilization programs to reduce inflation from high to moderate levels. Romania, on the other hand, shifted from low to high inflation at the time of its transition to a market economy.

For the three countries that did not have hyperinflation episodes (Israel, Mexico and Romania), the time series data is simply divided into two periods and each period is treated as a distinct inflation regime. For Bolivia and Nicaragua, pre-stabilization periods are dominated by hyperinflation and are unlikely to constitute inflation regimes with stable expectations (as assumed by our pass-through model). Our empirical analysis thus focuses only on the regimes represented by post-stabilization periods for these two countries. The remaining three countries (Argentina, Brazil and Peru) experienced shorter bouts of hyperinflation (these periods are enclosed by the dotted and solid lines in Figure 1). For these countries, the entire pre-stabilization periods, or parts of these periods without the hyperinflation episodes, are viewed as separate regimes in our tests below.

### B. Estimation of the Pass-Through

For each inflation regime, we consider the following log-linear relation based on (19):

$$\log P_t = \gamma_0 + \gamma_1 t + \pi_1(L)\log P_{t-1} + \pi_2(L)\log S_t + \pi_3(L)\log P^*_{t-1} + \varepsilon_t\,, \qquad (20)$$

[22] These countries include Ecuador, Sierra Leone, Turkey, Uruguay and Zambia. Although the inflation rate for these countries fluctuates considerably over the sample period, there is no clear indication of a switch in inflation regime.

where $P_t$ represents the home CPI, $S_t$ and $P_t^*$ are the effective nominal exchange rate and the foreign CPI, and $\varepsilon_t$ is the error term in quarter *t*.[23] The exchange rate pass-through to CPI over *N* periods can be defined as the accumulated partial effect of a one-unit increase in the log of nominal exchange rate in period *t* on the log of CPI in period $t + N$ (i.e., the sum of dynamic multipliers: $\sum_{\tau=0}^{T-1} \partial \log P_{t+\tau} / \partial \log S_t$ ).[24] The *N* - period pass-through can be calculated from coefficient estimates of lag polynomials, $\pi_1(L)$ and $\pi_2(L)$.[25]

The price series used for each country is the seasonally adjusted CPI. The exchange rate series for each country represents the nominal trade-weighted effective exchange rates of domestic currency over foreign currency. A series for foreign CPI is extracted from the relative price index (RPI) reported by the INS.[26] The INS monthly series are converted to quarterly series by taking simple averages of the monthly data over each quarter.

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[23] Log deviations are used to approximate proportional deviations around steady-state values (i.e., $\log P_t - \log \overline{P}_t \approx (P_t - \overline{P}_t)/\overline{P}_t$). Initial steady-state values (in logs) are included in the constant term and a trend term is introduced to allow for possible trends in these values. As discussed in section 2, lagged values of $\log M_t$ could also enter (20). This variation is explored below.

[24] Note that this cumulative effect is the same as the effect of a permanent one unit increase in the log of the nominal exchange rate in *t* on the log of the CPI in *t* + *N*. We focus on the effect of a permanent change in the nominal exchange rate because this variable tends to be non-stationary (and thus subject to permanent shocks)

[25] The dynamic multiplier $\partial \log P_{t+\tau} / \log S_t$ equals the coefficient of the $\tau$ 'th term in the lag polynomial $\Pi(L) = [1 - \pi_1(L)]^{-1} \pi_2(L)$.

[26] The RPI is defined as the ratio of domestic CPI for a country to trade weighted averages of the CPI's of partner countries. The series for the foreign CPI is obtained from this series by subtracting the log of the domestic CPI from the log relative price index.

As the exchange rate and price levels are non-stationary variables, the appropriate form for estimating the pass-through relation depends on whether it is a cointegrating relations or not. According to the purchasing power parity theory, the real exchange rate is stationary and thus the variables in the pass-through relation are cointegrated [i.e., $\log S_t$, $\log P_t^*$ and $\log P_t$ are cointegrated with a cointgrating vector equal to (1, 1, -1)]. The issue of whether the real exchange rate is stationary or contains a unit root has not been settled despite a vast literature on testing the purchasing power parity relation. We also explore this issue for our sample. First, using time series data for individual countries, we find that for most countries Augmented Dickey-Fuller tests do not reject the null hypothesis that the log of real exchange rate ($\log S_t + \log P_t^* - \log P_t$) contains a unit root.[27] Time series in our sample cover a period of about two decades and tests based on such a short span of time have limited power. Panel data has been used to increase the power of the unit-root tests. Although this approach also raises problematic issues, we applied unit-root tests to panel data formed by combining the time series data for our sample countries.[28] A test due to Im, Pesaran and Shin (1997) did not reject the unit root null for the whole panel.[29] In view of this evidence, we first assume that variables in (20) are not cointegrated and estimate this relation in the following first difference form:

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[27] The unit root null is not rejected for 64 out of 71 countries at the 5% level. The test includes a constant and a time trend (if significant). The criterion for choosing the number of lags (up to a maximum of 6 lags) was based on a general-to-specific procedure.

[28] The basic problem with panel methods is that they test the null hypothesis that all individual series have a unit root. The null hypothesis could be modified to allow some of the series to be stationary. Tests based on such a modified hypothesis have been applied to samples with a small number of countries (e.g., Alan Taylor, 2000) but are computationally cumbersome for our sample, which contains a large number of countries.

[29] The probability level for the Im, Pesaran and Shin (1997) test (including a time trend) is 0.45. This procedure tests the unit root null against the alternative of heterogeneous roots across units.

(continued…)

$$\Delta \log P_t = \gamma_1 + \pi_1(L)\Delta \log P_{t-1} + \pi_2(L)\Delta \log S_t + \pi_3(L)\Delta \log P^*_{t-1} + \varepsilon'_t . \tag{21}$$

Later, we explore the sensitivity of our estimates to this specification and examine estimates based on the level form of the pass-through relation. The pass-though relation is estimated separately for each inflation regime.

### C. Basic Evidence

To examine the exchange rate pass-through over the short, medium and long run, Tables 2 and 3 show estimates of the *N*-period pass-through for *N = 0, 1, 4, 20,* where *N* represents the number of quarters after the change in the exchange rate *(N = 0* represents the effect of the exchange rate change in the same quarter*)*. These estimates are based on (21) with the order of each lag polynomial determined by the Schwartz criterion. Table 2 provides the pass-through estimates for countries with a single inflation regime while Table 3 gives the estimates for those with two regimes. Table 2, classifies countries into three groups based on the average inflation rate (defined as the mean of $\Delta \log P_t$, expressed as an annual rate). Low, moderate and high inflation groups are defined as consisting of countries with average inflation rates less than 10 percent, between 10 and 30 percent and more than 30 percent, respectively. Most of the countries in this table (i.e., countries with single regimes) fall into the low and moderate inflation groups. Also note that while developing countries appear in all three inflation groups, all industrial countries are included in the low inflation group. The average exchange rate pass-through for each group rises as *N* increases, as would be expected. Comparing the average pass-through

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The unit root hypothesis, however, is rejected for our panel by the Levin and Lin (1993) test, which is based on a more restrictive alternative of homogeneous roots across units. For both tests, one country and the first four observations were dropped to form a balanced panel. These tests were implemented using GAUSS programs NPT 1.2 by Chiang and Kao (2001), and COINT 2.0 by Ouliaris and Phillips.

across groups, the table indicates a positive relation between inflation and the pass-through. For each *N*, the average pass-through is the lowest for the low inflation group and the highest for the high inflation group.

Table 3 provides further evidence on the influence of inflation on the pass-through based on within-country comparisons of high and moderate inflation regimes for six countries. For Argentina, Brazil and Peru (which experienced hyperinflation episodes), two alternative definitions of high inflation regimes are used: one including and the other excluding the hyperinflation period. In all cases, the pass-through for the high inflation regime is greater than that for the moderate regime for each value of *N*. The inter-regime difference in the pass-through is pronounced in most cases. As our model of the pass-through would predict, F-tests reject the null hypothesis that coefficients in the lag polynomials of the pass-through relation are the same between the two regimes. For Bolivia and Nicaragua (where the high inflation period is largely characterized by hyperinflation), the pass-through is estimated for only the moderate inflation period. These estimates are used in the cross-regime tests discussed below.

The cross-regime association between the average inflation rate and the exchange rate pass-through is illustrated in Figure 2 for the short ($N = 0$) as well as the medium ($N = 4$) run. Figures 2a and 2b show the association across countries with single regimes. These countries tend to have low and moderate inflation rates and are bunched at the lower end of the inflation scale. Nevertheless, the figures indicate a clear positive relation between inflation and the pass-through. Figures 2c and 2d show the relation between inflation and the pass-through across regimes for countries that experienced a regime shift.[30] This sample represents a wider variation in inflation rates and suggests an even stronger relation than the single-regime sample.

Pearson and Spearman rank correlations between the average inflation rate and the *N*-period pass-through across regimes are shown in Table 4 for *N = 0, 1, 4, 20*. The correlation coefficients are computed for the total sample as well as for samples restricted to countries with either one or two inflation regimes. In all cases, the correlation coefficients are significant at the 5 percent level, using a one-tail test of positive correlation.

Calvo and Reinhart (2000) present evidence that the exchange rate pass-through tends to be larger in emerging economies as compared to developed countries. As emerging economies also tend to have higher inflation rates, this finding could reflect the effect of inflation on the pass-through. It is possible, on the other hand, that some other factors account for the pass-through difference between developed and developing countries and the correlations between inflation and pass-through are picking up the effect of these factors. This issue can be explored by examining the relation between inflation and the pass-through separately for each country group. As noted above, all industrial countries experienced low inflation during our sample period and there is very little dispersion of inflation rates among these countries. Inflation rates vary considerably, however, among developing countries (see Table 1), and Table 4 shows the correlations between inflation and pass-through for these countries. The association between the two variables is positive and significant for the total as well as the single-regime sample of developing countries. Also note that the group experiencing shifts in inflation regimes includes

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[30] The high-inflation periods for Argentina, Brazil and Peru in these figures refer to the longer periods that include hyperinflation. The relations based on the shorter high-inflation periods (excluding hyperinflation) would not be much different.

only developing countries, and as already discussed above, there is a strong association between inflation and the pass-through for this group.

Table 5 examines the sensitivity of the results to certain variations in the specification of the pass-through relation. First, as the evidence on cointegration between variables in the pass-through relation is not conclusive, we also estimate this relation in the level form. The table shows that correlation coefficients between inflation and the pass-through based on the level regressions are significant in all cases except one (the Pearson coefficient for the $N = 20$ pass-through is not significant in the case of countries with regime shifts). Next, we explore the possibility that the first-difference form of the pass-through relation includes the lagged values of money growth. Correlations between inflation and the pass-through based on these regressions are also shown in Table 5.[31] These correlations are significant in all cases. In contrast to the results for the level regressions, the first-difference regressions with money growth show that inflation is strongly correlated with the long-run pass-through for countries with two regimes. The pass-through relation could potentially include other variables such as the gap between actual and potential (or target) output.[32] Data limitations, however, make it difficult to explore such a variation for our large sample of countries.

### D. Regression Analysis

We next undertake regression analysis to further examine the effect of the average inflation rate on the pass-through and to explore the influence of other macroeconomic variables.

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[31] The source of the data for money growth is the IMF's International Financial Statistics, line 34. The reported money series is seasonally adjusted. For countries where this data are not reported, the money series is estimated as the sum of currency in circulation and demand deposits, which is then seasonally adjusted.

We consider three additional variables, which represent measures of inflation variability, exchange rate volatility and the degree of openness (to imported goods). We estimate a regression of the following form:

$$PT_R(N) = b_0 + b_1 INFA_R + b_2 INFV_R + b_3 EXRV_R + b_4 OPEN_R + e_R, \quad (22)$$

where subscript $R$ indexes the inflation regime, $PT_R(N)$ is the estimated pass-through for $N$ periods, $INFA_R$ is the average inflation rate, $INFV_R$ is the variance of the inflation rate, $EXRV_R$ is the variance of the proportional exchange rate change (i.e., $\Delta \log S_t$, expressed as an annual rate), $OPEN_R$ is the import to GDP ratio, and $e_R$ is the error term.[33] We initially estimate (22) by OLS, but later address the problem that the variance of $e_R$ may be different across regimes.

OLS estimates of (22) for the total sample are shown in Table 6 for $N = 0$ and $N = 20$.[34] In this table, we also show estimates of simple regressions of $PT_R(N)$ on each explanatory variable separately. The simple regressions show that while the pass-through (in both the short and the long run) is not significantly affected by the import to GDP ratio, it is positively and significantly related to the average inflation rate and the inflation and exchange rate variances. As would be expected, the two variances are strongly correlated with the average inflation rate and with each other.[35] The multicollinearity between these variables could potentially make it

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[32] In terms of our theoretical model, this variable could be motivated by the assumption that the monetary policy rule also includes the output gap.

[33] The data on the ratio of imports of goods and services to GDP are obtained from the World Bank's World Development Indicators.

[34] The regression equation was also estimated for $N = 1$ and $N = 4$, but the results were similar and are not reported in the table.

[35] The correlations between the four explanatory variables are as follows:

(continued…)

difficult to isolate the effect of these variables in regression (22). Estimates of (22) show, however, that the average inflation rate clearly dominates the other variables and is the only variable that exerts a significant and positive effect. Our estimates of (22) indicate that a 10% increase in the average inflation rate increases the short-run pass-through by about 0.05 and the long-run one by about 0.06. It is interesting to note that recent studies based on the OECD data do not find the effect of the inflation rate on the pass-through to be as strong as in our analysis.[36] This difference in results may be due to the fact that there is much more variation in the inflation rate across regimes in our sample than in the OECD sample.

To examine the sensitivity of our results to the data for different country groups, Table 7 estimates (22) for samples that consist of the one-regime, two-regime or developing countries. The effect of the average inflation rate differs across country groups, especially in the case of the long-run pass-through. In all cases, however, the inflation rate continues to exert a positive and significant effect on the pass-through. The effects of the inflation and exchange rate variances remain insignificant. We also explored the sensitivity of our results to alternative estimates of the

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| | INFA | INFV | EXRV | OPEN |
|---|---|---|---|---|
| INFA | 1 | 0.85 | 0.82 | -0.25 |
| INFV | | 1 | 0.96 | -0.27 |
| EXRV | | | 1 | -0.28 |
| OPEN | | | | 1 |

[36] Gagnon and Ihrig (2001), for example, find that the standard deviation of the inflation rate performs better the average inflation rate in explaining cross-regime variation in the pass-through. Campa and Goldberg (2001) show, moreover, that a measure of exchange rate volatility (similar to ours) is a significant determinant of the pass-through (to import prices) in regressions that include the average inflation rate.

pass-through based on the level regression (20) and on the first-difference regression (21) augmented by lagged money growth terms. These variations did not much effect our result that the average inflation rate is the key determinant of the pass-through.

The pass-through measure is subject to a regime-specific estimation error and the variance of this error could differ across regimes. One way to deal with this problem is to estimate (22) by weighted least squares (WLS), using the inverse of the (normalized) standard error of the pass-through estimate as weights. This weighting scheme is appealing as it gives less weight to more noisy estimates of the pass-through. One problem with this procedure, however, is that a few observations with exceptionally low standard errors of the estimate could distort the results by giving an unduly high weight to outliers. This problem can be avoided by excluding such outliers. A more general problem is that the procedure represents an over-adjustment as the error in estimating the pass-through is only a part of the overall error ($e_R$).[37] We estimate (22) by WLS, nevertheless, to examine whether WLS estimates (with an over-adjustment) differ much from the OLS estimates (with no adjustment).

Table (8) presents WLS estimates of (22) for the total sample as well as the one-regime, two-regime and developing countries samples. The regression is estimated for *N = 0* and *N = 20*,

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[37] Let $e_R = e_R^P + e_R^O$, where $e_R^P$ is the error for the pass-through estimate and $e_R^O$ is the error from other sources. Assume that $e_R^P$ and $e_R^O$ are uncorrelated and the variance of $e_R^O$ is the same for all *R*. Let $\sigma_{PR}^2$ denote the regime-specific variance of $e_R^P$ and $\sigma_O^2$ the variance of $e_R^O$. Our WLS regressions use the weights, $\xi_R = \sigma_P / \sigma_{PR}$, where $\sigma_P$ is the average of $\sigma_{PR}$ across *R*. The appropriate weight for each observation, however, is the inverse of the standard deviation of $e_R$ normalized by its average. Denoting this weight by $\chi_R$, it can be shown that $\chi_R = \xi_R \; \{[1 + (\sigma_O^2 / \sigma_P^2)]/[1 + \xi_R^2(\sigma_O^2 / \sigma_P^2)]\}^{1/2}$. For a nonzero $\sigma_O^2$, $\chi_R < \xi_R \; (> \xi_R)$ if $\xi_R > 1 \; (< 1)$. Thus the WLS regressions give too much (little) weight to observations with below-average (above-average) $\sigma_{PR}$.

using the inverse of the standard error of the estimate for $PT_R(N)$ as the weight series in each case.[38] The regressions for the short-run pass-through exclude one outlier (Burkina Faso), where the standard error of the $N = 0$ pass-through is very low and much below that of other regimes. As in the case of OLS estimates, the effect of the average inflation rate on the pass-through is positive and significant in all regressions. Other variables, however, do not have a robust effect across different regressions. The effect of the inflation rate variance, for example, is significant and negative on the long-run pass-through for the total and developing country samples but insignificant for the one- and two-regime samples. The effect of this variable on the short-run pass-through, moreover, is insignificant for all samples.

## IV. Conclusions

The paper explores a model of the exchange rate pass-through, which emphasizes the role of price inertia and expectations. As prices are set for several periods in the model, the pass-through includes the expected effect of changes in the exchange rate on future costs and prices. The expected effect depends on the inflation regime. For high inflation regimes, the effect of monetary shocks tends to be more persistent and is likely to be reflected in exchange rate changes to a larger degree. For these reasons, the model implies the hypothesis that the exchange rate pass-through is larger in high inflation regimes.

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[38]The standard error of the coefficient of $\Delta \log S_t$ from regression (21) is used to estimate the standard error of $PT_R(0)$. The standard error of $PT_R(20)$ is approximated by the standard error of the long-run pass-through, $\pi_2(1)/[1-\pi_1(1)]$. This standard error is derived from the standard errors and the covariance of $\pi_1(1)$ and $\pi_2(1)$, which are obtained from re-estimating (21) in the following transformed form:

$$\Delta \log P_t = \gamma_1 + \pi_1(1)\Delta \log P_{t-1} + \pi_2(1)\Delta \log S_t + \pi_1'(L)\Delta^2 \log P_{t-1} + \pi_2'(L)\Delta^2 \log S_t + \pi_3(L)\Delta \log P^*_{t-1} + \varepsilon_t' .$$

This paper undertakes an extensive empirical analysis based on a large data set to test this hypothesis. Our data set is especially suitable for testing the relationship between inflation and the pass-through because it includes countries with pronounced differences in inflationary experience. We find strong evidence that the relation between the pass-through and the average inflation rate is positive and significant across regimes. This relation remains robust even when we control for other macroeconomic variables. These results are based on comparisons of regimes across countries as well as across periods.

An important policy implication of these findings is that the dependence of the exchange rate pass-through on the inflation regime should be taken into account in designing monetary policy rules. This dependency would make it easier for a country to implement a policy targeting for a low inflation rate. It should be emphasized, however, that the credibility of a low inflation regime would need to be established before the benefits of  a low pass-through can be realized.

To focus on the influence of the inflationary environment on the exchange rate pass-through to CPI, this paper's estimates of the pass-through have been based on a model that abstracts from the role of microeconomic factors emphasized in the traditional literature. Exploring the relative importance of these factors in determining the pass-through to CPI (or to certain components of CPI) would be an interesting topic for future research. Our model assumes producer currency pricing and thus implies that the pass-through to import prices does not depend on the inflationary environment. One interesting extension of the paper's model would be to allow for the possibility that the choice between producer and local currency pricing depends on the inflationary environment.

## Table 1. Selected Characteristics of Inflation: Sample Countries

| Country | Mean Annual Inflation | Maximum | Minimum | Std. Dev. | Country | Mean Annual Inflation | Maximum | Minimum | Std. Dev. |
|---|---|---|---|---|---|---|---|---|---|
| | | *(in percent)* | | | | | *(in percent)* | | |
| *Industrial Countries* | | | | | | | | | |
| Austria | 3.0 | 9.7 | -0.9 | 2.0 | Ghana | 29.7 | 162.5 | -62.7 | 28.1 |
| Australia | 5.3 | 13.3 | -1.2 | 3.7 | Greece | 13.5 | 28.8 | 1.3 | 6.7 |
| Belgium | 3.4 | 10.7 | -0.8 | 2.6 | Guatemala | 11.9 | 53.6 | -10.7 | 12.2 |
| Canada | 4.2 | 12.9 | -2.7 | 3.3 | Haiti | 13.0 | 61.5 | -55.4 | 13.8 |
| Denmark | 4.4 | 19.4 | 0.2 | 3.6 | Honduras | 12.3 | 41.2 | 1.5 | 8.8 |
| Finland | 4.5 | 14.3 | -0.9 | 3.5 | Hungary | 13.9 | 38.5 | -1.8 | 8.5 |
| France | 4.4 | 15.5 | -0.3 | 4.0 | Indonesia | 10.9 | 75.0 | -10.7 | 12.6 |
| Germany | 2.6 | 8.6 | -6.3 | 2.3 | India | 7.9 | 23.7 | -1.9 | 5.3 |
| Ireland | 5.6 | 23.0 | -0.3 | 5.5 | Israel | 39.2 | 199.6 | -5.7 | 48.1 |
| Italy | 7.2 | 22.7 | 0.8 | 5.3 | Jamaica | 18.1 | 77.2 | 1.5 | 15.2 |
| Netherlands | 2.7 | 7.8 | -3.7 | 2.1 | Jordan | 5.5 | 38.0 | -8.6 | 8.3 |
| Norway | 5.1 | 13.8 | -1.7 | 3.6 | Kenya | 12.6 | 57.3 | -6.2 | 10.8 |
| New Zealand | 6.6 | 28.7 | -2.0 | 6.1 | Korea | 6.7 | 35.1 | -2.9 | 6.9 |
| Singapore | 2.3 | 14.7 | -3.6 | 3.1 | Madagascar | 15.6 | 62.1 | -3.0 | 13.5 |
| Spain | 6.9 | 17.0 | 0.5 | 4.2 | Malaysia | 3.6 | 11.4 | -1.7 | 2.6 |
| Sweden | 5.2 | 16.4 | -2.6 | 4.4 | Mexico | 32.3 | 117.0 | 2.3 | 25.7 |
| Switzerland | 2.7 | 10.7 | -1.7 | 2.5 | Morocco | 5.6 | 20.6 | -2.2 | 4.7 |
| UK | 5.4 | 28.8 | -0.9 | 4.5 | Nepal | 9.1 | 28.8 | -4.0 | 6.5 |
| USA | 4.2 | 15.0 | -1.6 | 3.1 | Nicaragua | 113.1 | 759.8 | -11.3 | 182.9 |
| *Developing Countries* | | | | | Pakistan | 7.9 | 20.3 | -0.1 | 4.3 |
| Argentina | 88.6 | 813.2 | -3.9 | 127.4 | Papua New Guinea | 8.1 | 28.1 | -2.0 | 6.5 |
| Bahrain | 1.5 | 29.9 | -9.7 | 5.1 | Paraguay | 16.4 | 46.8 | -1.5 | 10.9 |
| Bangladesh | 8.1 | 31.7 | -7.5 | 7.0 | Peru | 85.3 | 773.2 | -1.5 | 126.0 |
| Bolivia | 60.1 | 679.6 | -0.5 | 123.5 | Philippines | 10.7 | 239.3 | -242.3 | 38.2 |
| Brazil | 124.4 | 599.8 | -1.4 | 123.8 | Portugal | 10.7 | 33.9 | 1.2 | 7.9 |
| Burundi | 10.2 | 42.3 | -20.2 | 12.2 | Romania | 35.1 | 166.6 | -14.5 | 46.7 |
| Burkina Faso | 4.1 | 46.3 | -17.2 | 11.1 | South Africa | 11.2 | 23.2 | -0.6 | 4.4 |
| Cameroon | 6.4 | 68.1 | -15.8 | 12.0 | Sierra Leone | 36.8 | 158.1 | -70.5 | 44.5 |
| China | 5.9 | 35.1 | -6.8 | 7.8 | Thailand | 5.1 | 25.1 | -3.2 | 4.9 |
| Chile | 14.7 | 46.1 | -1.0 | 9.8 | Trinidad | 8.4 | 24.4 | -1.1 | 5.4 |
| Colombia | 19.8 | 31.3 | 3.8 | 5.8 | Tunisia | 6.3 | 16.7 | 1.0 | 3.4 |
| Costa Rica | 18.4 | 80.2 | -3.0 | 14.2 | Turkey | 48.5 | 132.5 | 18.7 | 19.0 |
| Ecuador | 32.1 | 98.4 | 7.3 | 17.3 | Uruguay | 39.1 | 89.6 | 2.9 | 22.1 |
| Egypt | 11.8 | 42.9 | -7.2 | 9.4 | Venezuela | 27.6 | 121.4 | 2.7 | 20.6 |
| El Salvador | 12.7 | 31.8 | -4.8 | 8.5 | Zambia | 39.7 | 165.6 | -0.3 | 32.9 |
| Ethiopia | 5.2 | 63.9 | -26.8 | 13.1 | Zimbabwe | 20.0 | 59.1 | -1.1 | 13.6 |

Table 2. Estimates of Exchange Rate Pass-Through: Countries With A Single Inflation Regime

| Country | Exchange Rate Pass-Through For: | | | | Country | Exchange Rate Pass-Through For: | | | |
|---|---|---|---|---|---|---|---|---|---|
| | N=0 | N=1 | N=4 | N=20 | | N=0 | N=1 | N=4 | N=20 |
| *Low Inflation Countries* | | | | | USA | 0.00 | -0.01 | 0.02 | 0.06 |
| Austria | 0.04 | 0.05 | 0.05 | 0.05 | *Average* | 0.04 | 0.08 | 0.14 | 0.16 |
| Australia | 0.03 | 0.06 | 0.10 | 0.14 | | | | | |
| Bahrain | -0.08 | -0.09 | -0.09 | -0.09 | *Moderate Inflation Countries* | | | | |
| Bangladesh | 0.00 | 0.08 | 0.08 | 0.08 | Burundi | 0.12 | 0.10 | 0.09 | 0.09 |
| Belgium | 0.08 | 0.10 | 0.19 | 0.26 | Chile | 0.17 | 0.18 | 0.35 | 0.38 |
| Burkina Faso | 0.16 | 0.29 | 0.38 | 0.34 | Colombia | -0.01 | -0.01 | -0.01 | -0.01 |
| Cameroon | 0.22 | 0.32 | 0.38 | 0.39 | Costa Rica | 0.03 | 0.14 | 0.57 | 0.57 |
| Canada | 0.00 | 0.04 | 0.11 | 0.19 | Egypt | 0.01 | -0.01 | 0.21 | 0.24 |
| China | 0.04 | 0.14 | 0.30 | 0.41 | El Salvador | 0.06 | 0.12 | 0.18 | 0.19 |
| Denmark | 0.06 | 0.14 | 0.22 | 0.24 | Ghana | 0.14 | 0.12 | 0.20 | 0.09 |
| Ethiopia | -0.01 | -0.06 | -0.09 | -0.10 | Greece | 0.12 | 0.25 | 0.42 | 0.48 |
| Finland | -0.01 | 0.03 | -0.02 | -0.04 | Guatemala | 0.15 | 0.33 | 0.45 | 0.46 |
| France | -0.01 | 0.03 | 0.11 | 0.18 | Haiti | 0.21 | 0.28 | 0.31 | 0.31 |
| Germany | 0.05 | 0.11 | 0.13 | 0.13 | Honduras | 0.07 | 0.12 | 0.31 | 0.38 |
| India | 0.06 | 0.09 | 0.10 | 0.10 | Hungary | 0.18 | 0.31 | 0.48 | 0.54 |
| Ireland | 0.06 | 0.12 | 0.38 | 0.54 | Indonesia | 0.10 | 0.21 | 0.41 | 0.43 |
| Italy | 0.04 | 0.06 | 0.11 | 0.18 | Jamaica | 0.17 | 0.33 | 0.49 | 0.53 |
| Jordan | 0.13 | 0.27 | 0.44 | 0.40 | Kenya | 0.09 | 0.22 | 0.35 | 0.38 |
| Korea | 0.06 | 0.08 | 0.10 | 0.12 | Madagascar | 0.01 | 0.21 | 0.31 | 0.32 |
| Malaysia | -0.01 | 0.03 | 0.05 | 0.05 | Paraguay | 0.07 | 0.19 | 0.34 | 0.34 |
| Morocco | 0.10 | 0.26 | 0.29 | 0.29 | Philippines | 0.00 | 0.33 | 0.35 | 0.35 |
| Nepal | 0.12 | 0.25 | 0.31 | 0.31 | Portugal | -0.01 | 0.08 | 0.44 | 0.54 |
| Netherlands | 0.05 | 0.10 | 0.19 | 0.23 | South Africa | 0.02 | 0.07 | 0.13 | 0.14 |
| Norway | 0.01 | 0.07 | 0.08 | 0.13 | Venezuela | 0.29 | 0.36 | 0.50 | 0.61 |
| New Zealand | 0.05 | 0.13 | 0.27 | 0.42 | Zimbabwe | 0.06 | 0.17 | 0.29 | 0.32 |
| Pakistan | -0.05 | -0.06 | -0.07 | -0.07 | *Average* | 0.09 | 0.19 | 0.33 | 0.35 |
| Papua New Guinea | 0.11 | 0.23 | 0.36 | 0.38 | | | | | |
| Singapore | -0.08 | -0.13 | -0.13 | -0.13 | *High Inflation Countries* | | | | |
| Spain | 0.00 | 0.03 | 0.11 | 0.15 | Ecuador | 0.13 | 0.34 | 0.60 | 0.68 |
| Sweden | 0.02 | 0.03 | 0.03 | 0.03 | Sierra Leone | 0.27 | 0.29 | 0.52 | 0.52 |
| Switzerland | 0.08 | 0.07 | 0.07 | 0.07 | Turkey | 0.36 | 0.46 | 0.51 | 0.52 |
| Thailand | 0.02 | 0.07 | 0.14 | 0.19 | Uruguay | 0.19 | 0.27 | 0.45 | 0.62 |
| Trinidad | 0.04 | 0.08 | 0.12 | 0.13 | Zambia | 0.15 | 0.27 | 0.41 | 0.46 |
| Tunisia | -0.03 | 0.05 | 0.09 | 0.10 | *Average* | 0.22 | 0.32 | 0.50 | 0.56 |
| UK | -0.01 | -0.01 | 0.02 | 0.03 | | | | | |

Notes: The pass-through represents the effect in quarter t+N of a unit increase in period t.

Table 3. Estimates of Exchange Rate Pass-Through: Countries With Two Inflation Regimes

| Country | Exchange Rate Pass-Through For: | | | | F-test | |
|---|---|---|---|---|---|---|
| | N=0 | N=1 | N=4 | N=20 | | |
| Argentina | | | | | | |
| Moderate Inflation (1991:3-2000:4) | 0.01 | 0.06 | -0.09 | -0.09 | | |
| High Inflation (1980:1-1991:2) | 0.67 | 1.02 | 1.01 | 1.09 | 17.02 | * |
| High Inflation (1980:2-1989:1) | 0.62 | 0.68 | 0.76 | 0.79 | 25.95 | * |
| Brazil | | | | | | |
| Moderate Inflation (1994:4-2000:4) | 0.08 | 0.17 | 0.39 | 0.44 | | |
| High Inflation (1981:1-1994:3) | 0.75 | 0.33 | 0.63 | 0.63 | 95.75 | * |
| High Inflation (1980:2-1989:2) | 0.64 | 0.76 | 0.89 | 0.92 | 26.35 | * |
| Israel | | | | | | |
| Moderate Inflation (1985:4-2000:4) | 0.25 | 0.28 | 0.28 | 0.28 | | |
| High Inflation (1979:3-1985:3) | 0.64 | 0.83 | 0.83 | 0.83 | 47.87 | * |
| Mexico | | | | | | |
| Moderate Inflation (1988:3-2000:4) | 0.09 | 0.27 | 0.27 | 0.27 | | |
| High Inflation (1980:2-1988:2) | 0.25 | 0.28 | 0.28 | 0.28 | 17.87 | * |
| Peru | | | | | | |
| Moderate Inflation (1993:3-2000:4) | -0.11 | -0.07 | -0.20 | -0.20 | | |
| High Inflation (1980:2-1993:2) | 0.49 | 0.64 | 0.64 | 0.64 | 16.83 | * |
| High Inflation (1980:3-1988:2) | 0.06 | 0.29 | 0.52 | 0.40 | 4.32 | * |
| Romania | | | | | | |
| Moderate Inflation (1981:3-1990:3) | 0.06 | -0.01 | 0.03 | 0.06 | | |
| High Inflation (1990:4-2000:4) | 0.30 | 0.47 | 0.47 | 0.47 | 5.45 | * |
| Bolivia | | | | | | |
| Moderate Inflation (1986:2-2000:4) | 0.02 | 0.08 | 0.09 | 0.09 | NA | |
| Nicaragua | | | | | | |
| Moderate Inflation (1991:3-2000:4) | 0.46 | 0.28 | 0.28 | 0.28 | NA | |

Notes: The pass-through represents the effect in quarter t+N of a unit increase in period t. The F test tests the null hypothesis that the coefficients on the nominal exchange rates, the lagged consumer prices, and the foreign prices are equal in the indicated high and moderate inflation regimes. * indicates significance at the 5 percent level.

Table 4. Correlation of Exchange Rate Pass-Through With the Average Inflation Rate: Regressions in First Differences

| Sample | No. of observations | Pearson Correlations | | | | Spearman Correlations | | | |
|---|---|---|---|---|---|---|---|---|---|
| | | N=0 | N=1 | N=4 | N=20 | N=0 | N=1 | N=4 | N=20 |
| All country sample | | | | | | | | | |
| All observations | 77 | 0.86 | 0.74 | 0.66 | 0.62 | 0.60 | 0.67 | 0.69 | 0.65 |
| Observations including only developing countries | 58 | 0.85 | 0.72 | 0.65 | 0.62 | 0.60 | 0.61 | 0.65 | 0.63 |
| Country sample with one inflation regime | | | | | | | | | |
| All observations | 63 | 0.66 | 0.58 | 0.60 | 0.56 | 0.50 | 0.58 | 0.64 | 0.59 |
| Observations including only developing countries | 44 | 0.62 | 0.50 | 0.52 | 0.52 | 0.47 | 0.46 | 0.57 | 0.54 |
| Country sample with two inflation regimes | 14 | 0.87 | 0.77 | 0.84 | 0.83 | 0.87 | 0.88 | 0.84 | 0.84 |

Notes: The pass-through represents the effect in quarter t+N of a unit increase in period t. All correlations are significant at the 5%
for the Spearman correlations are based on quantiles of Spearman's rho in Conover (1999).

Table 5. Correlation of Exchange Rate Pass-Through With the Average Inflation Rate: Additional Regressions

| | No. of | Pearson Correlations | | | | Spearman Correlations | | | |
|---|---|---|---|---|---|---|---|---|---|
| | observations | N=0 | N=1 | N=4 | N=20 | N=0 | N=1 | N=4 | N=20 |
| Regressions in Levels | | | | | | | | | |
| All country sample | 77 | 0.81 | 0.77 | 0.58 | 0.31 | 0.61 | 0.71 | 0.73 | 0.64 |
| Country sample with one inflation regime | 63 | 0.61 | 0.63 | 0.60 | 0.41 | 0.54 | 0.64 | 0.68 | 0.57 |
| Country sample with two inflation regimes | 14 | 0.78 | 0.71 | 0.47 | 0.24 [a] | 0.82 | 0.89 | 0.87 | 0.72 |
| Regressions in First Differences With Money Growth | | | | | | | | | |
| All country sample | 75 | 0.85 | 0.84 | 0.72 | 0.70 | 0.58 | 0.66 | 0.64 | 0.65 |
| Country sample with one inflation regime | 61 | 0.56 | 0.48 | 0.45 | 0.45 | 0.48 | 0.55 | 0.53 | 0.53 |
| Country sample with two inflation regimes | 14 | 0.88 | 0.96 | 0.89 | 0.91 | 0.86 | 0.94 | 0.93 | 0.89 |

Notes: The pass-through represents the effect in quarter t+N of a unit increase in period t. All correlations are significant at the 5% level using a one-tailed test of positive correlation unless otherwise indicated. Tests of significance for the Spearman correlations are based on quantiles of Spearman's rho in Conover (1999). The regressions with money growth do not include Belgium and the United Kingdom, for which comparable money growth data are not available.

a. indicates not significant at the 5% level.

Table 6. Regression Estimates: Total Sample

| | *Dependent Variable* | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Exchange Rate Pass-Through at N=0 | | | | | Exchange Rate Pass-Through at N=20 | | | | |
| Constant | 0.025 * | 0.082 * | 0.071 * | 0.145 * | 0.014 | 0.188 * | 0.249 * | 0.236 * | 0.325 * | 0.177 * |
| | (0.01) | (0.01) | (0.01) | (0.03) | (0.02) | (0.02) | (0.02) | (0.03) | (0.05) | (0.04) |
| Inflation | 0.439 * | | | | 0.528 * | 0.470 * | | | | 0.584 * |
| | (0.03) | | | | (0.06) | (0.07) | | | | (0.13) |
| Inflation Variance | | 0.314 * | | | -0.094 | | 0.325 * | | | -0.279 |
| | | (0.04) | | | (0.10) | | (0.07) | | | (0.23) |
| Exchange Rate Variance | | | 0.243 * | | -0.001 | | | 0.263 * | | 0.122 |
| | | | (0.03) | | (0.08) | | | (0.05) | | (0.17) |
| Import to GDP Ratio | | | | -0.001 | 0.0001 | | | | -0.001 | 0.000 |
| | | | | (0.001) | (0.0004) | | | | (0.001) | (0.001) |
| Adjusted R-sqaured | 0.73 | 0.54 | 0.43 | 0.01 | 0.73 | 0.38 | 0.21 | 0.23 | 0.01 | 0.37 |

Notes: The pass-through represents the effect in quarter t+N of a unit increase in period t. * indicates significance at the 5 percent level.

## Table 7. Regression Estimates: Selected Samples

| | *Dependent Variable* | | | | | |
|---|---|---|---|---|---|---|
| | Exchange Rate Pass-Through at N=0 | | | Exchange Rate Pass-Through at N=20 | | |
| | Countries With One Inflation Regime | Countries With Two Inflation Regime | Developing Countries | Countries With One Inflation Regime | Countries With Two Inflation Regime | Developing Countries |
| Constant | 0.024 | -0.103 | -0.005 | 0.143 * | -0.039 | 0.150 * |
| | (0.02) | (0.05) | (0.03) | (0.05) | (0.12) | (0.06) |
| Inflation | 0.513 * | 0.389 * | 0.514 * | 1.253 * | 0.451 * | 0.536 * |
| | (0.12) | (0.08) | (0.07) | (0.29) | (0.19) | (0.14) |
| Inflation Variance | -0.134 | 0.225 | -0.067 | 0.974 | -0.026 | -0.202 |
| | (0.44) | (0.16) | (0.12) | (1.11) | (0.38) | (0.25) |
| Exchange Rate Variance | 0.066 | -0.148 | -0.009 | -0.429 | 0.075 | 0.095 |
| | (0.12) | (0.11) | (0.09) | (0.31) | (0.27) | (0.18) |
| Import to GDP Ratio | -0.0004 | 0.007 * | 0.001 | -0.0004 | 0.005 | 0.001 |
| | (0.0004) | (0.002) | (0.001) | 0.0009 | (0.004) | (0.002) |
| Adjusted R-sqaured | 0.41 | 0.90 | 0.72 | 0.30 | 0.63 | 0.36 |
| No. of Observations | 63 | 14 | 58 | 63 | 14 | 58 |

Notes: The pass-through represents the effect in quarter t+N of a unit increase in period t. * indicates significance at the 5 percent level.

## Table 8. Regression Estimates: Weighted Least Sqaures

| | Dependent Variable | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Exchange Rate Pass-Through at N=0 | | | | Exchange Rate Pass-Through at N=20 | | | |
| | Total Sample | Countries With One Inflation Regime | Countries With Two Inflation Regime | Developing Countries | Total Sample | Countries With One Inflation Regime | Countries With Two Inflation Regime | Developing Countries |
| Constant | 0.01 | 0.02 | -0.03 | 0.03 | 0.20 * | 0.19 * | 0.15 | 0.21 * |
| | (0.01) | (0.02) | (0.04) | (0.03) | (0.03) | (0.04) | (0.11) | (0.05) |
| Inflation | 0.50 * | 0.44 * | 0.44 * | 0.49 * | 0.65 * | 0.76 * | 0.44 * | 0.58 * |
| | (0.09) | (0.16) | (0.10) | (0.10) | (0.12) | (0.35) | (0.17) | (0.13) |
| Inflation Variance | -0.08 | 0.87 | 0.07 | -0.05 | -1.00 * | 0.79 | -0.48 | -0.79 * |
| | (0.11) | (1.04) | (0.16) | (0.12) | (0.30) | (1.80) | (0.43) | (0.34) |
| Exchange Rate Variance | 0.02 | -0.08 | -0.07 | 0.001 | 0.58 * | 0.30 | 0.31 | 0.45 * |
| | (0.07) | (0.16) | (0.14) | (0.08) | (0.22) | (0.44) | (0.27) | (0.24) |
| Import to GDP Ratio | -0.0002 | -0.0003 | 0.003 | -0.001 | -0.002 * | -0.002 * | 0.002 | -0.001 * |
| | (0.0003) | (0.0003) | (0.002) | (0.001) | (0.001) | (0.001) | (0.003) | (0.001) |
| Adjusted R-sqaured | 0.45 | 0.33 | 0.77 | 0.45 | 0.53 | 0.40 | 0.84 | 0.54 |
| No. of Observations | 76 | 62 | 14 | 57 | 77 | 63 | 14 | 58 |

Notes: The pass-through represents the effect in quarter t+N of a unit increase in period t. The regression results for the N=0 pass-through exclude Burkina Faso (see text for explanation). * indicates significance at the 5 percent level.

Figure 1. Countries With Shifts in Inflation Regimes

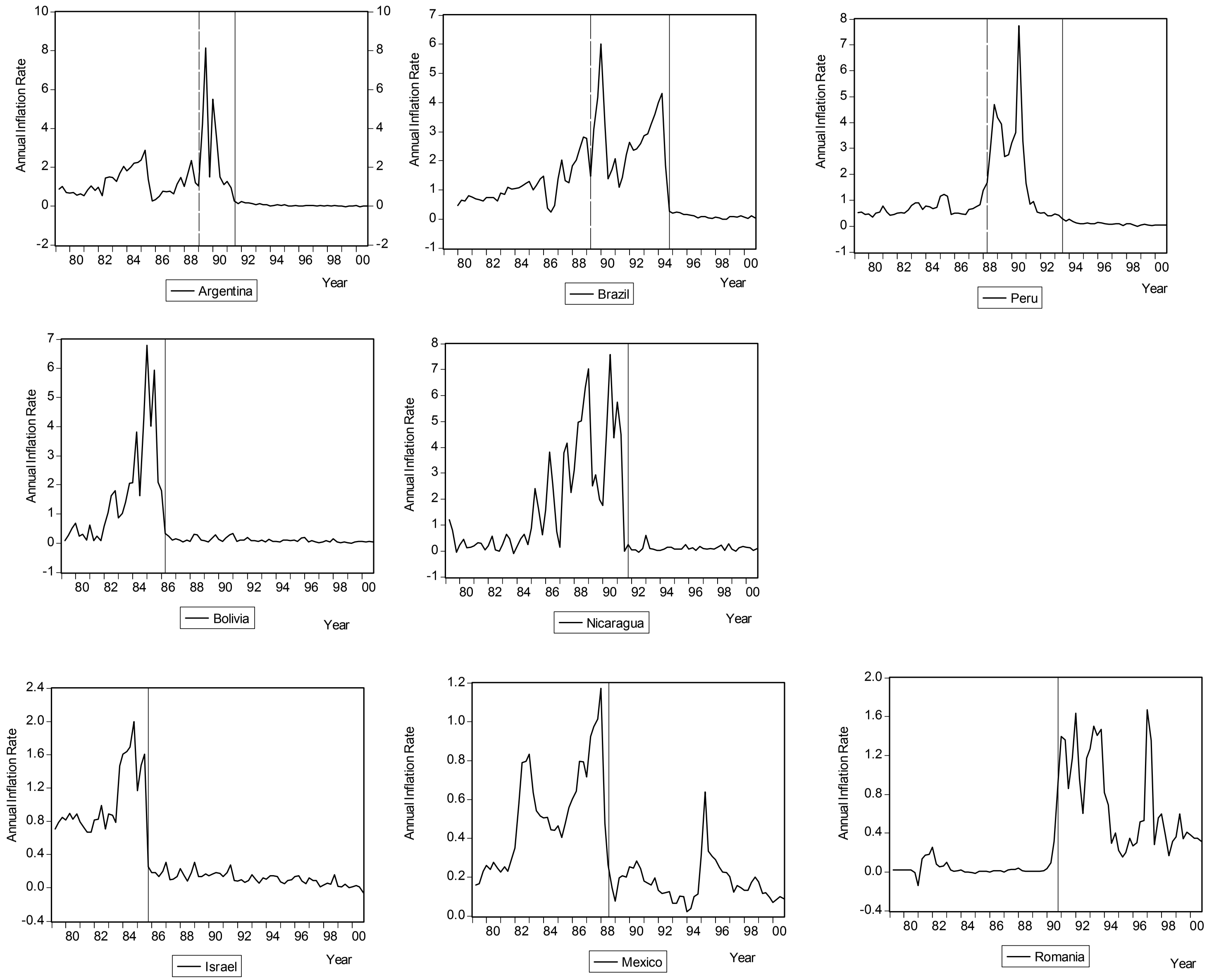
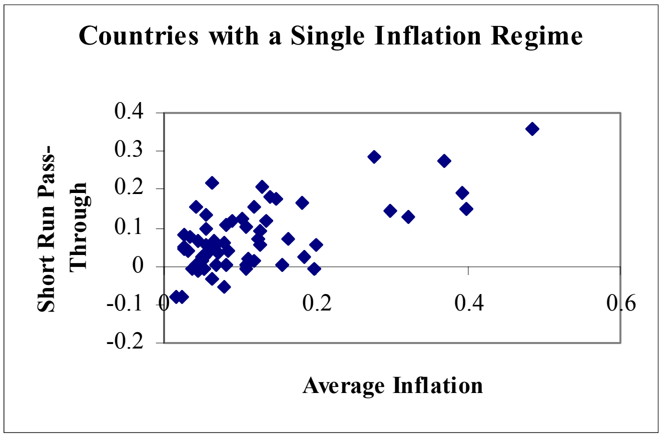
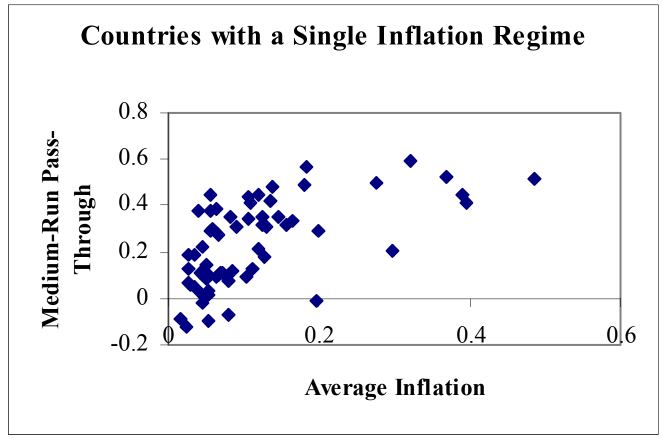


Figure 2. Association Between Average Inflation and Exchange Rate Pass-Through

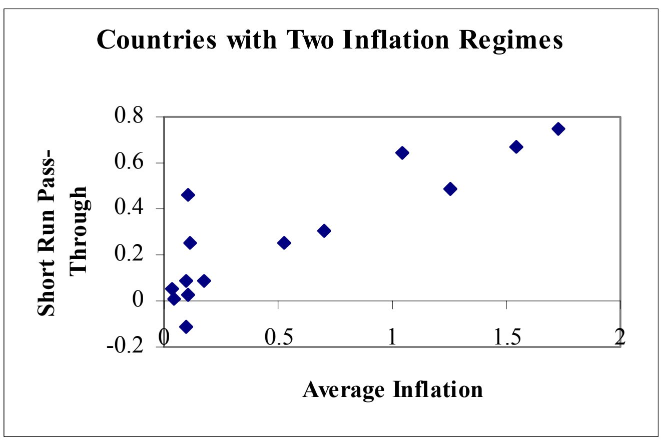
a.



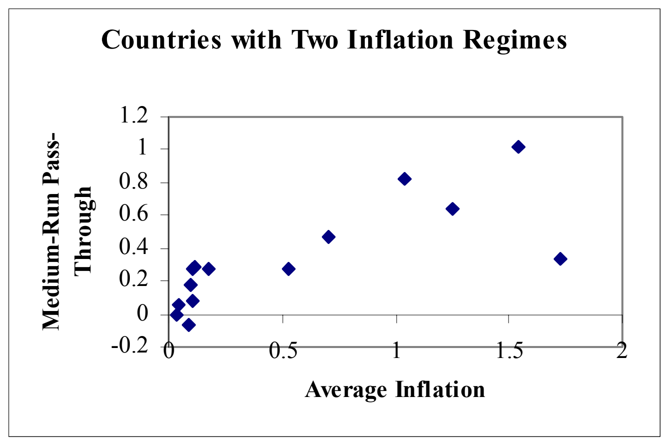

b.




c.




d.



## Appendix

### A. Complete Asset Markets

Use (9)-(11), the first equality in (14), the assumption that $p_t^* = 0$, and let $\beta = 1$, for simplicity, to obtain

$$x_t = \kappa(x_{t-1} + E_t x_{t+1}) + \omega(s_t + E_t s_{t+1}), \qquad \text{(A1)}$$

where $\kappa = \alpha\theta/(2(2-\alpha\theta))$ and $\omega = (1-\alpha\theta)/(2-\alpha\theta)$. Next use (15) and the assumption that $v_t$ is white noise to express (A1) as

$$x_t = [(\kappa - \rho\omega)x_{t-1} + \kappa E_t x_{t+1} + \omega v_t]/(1 + \rho\omega). \qquad \text{(A2)}$$

Assume a solution of the form

$$x_t = \phi_1 x_{t-1} + \phi_2 v_t. \qquad \text{(A3)}$$

Note that the value of $v_t$ can be inferred from the feedback rule (15) by observing $s_t$ and $x_{t-1}$. After solving the model to determine $\phi_1$ and $\phi_2$, use (15) to express (A3) as $x_t = (\phi_1 + \rho\phi_2)x_{t-1} + \phi_2 s_t$, and then use (10) to derive (16) with $\mu_1 \equiv \phi_1 + \rho\phi_2$ and $\mu_2 \equiv \phi_2$. Using (A3) to determine $E_t x_{t+1}$, write (A2) as

$$x_t = [(\kappa - \rho\omega + \kappa\phi_1^2)x_{t-1} + (\kappa\phi_1\phi_2 + \omega)v_t]/(1 + \rho\omega). \qquad \text{(A4)}$$

Matching coefficients between (A3) and (A4), the coefficients in (A3) are given by

$$\kappa\phi_1^2 - (1+\rho\omega)\phi_1 + (\kappa - \rho\omega) = 0, \qquad \text{(A5)}$$

$$\phi_2 = \omega/(1 + \rho\omega - \kappa\phi_1). \qquad \text{(A6)}$$

Solving (A5) and (A6) numerically (and using the stable root of the quadratic equation (A5), i.e., $|\phi_1| < 1$), it can be shown that both $\mu_1$ and $\mu_2$ (as well as $\phi_1$) decrease in $\rho$, given that $\alpha\theta$ is in the *(0, 1)* interval (note that both $\kappa$ and $\omega$ depend on $\alpha\theta$). This correspondence is illustrated in the Appendix Table for selected values of $\rho$ and $\alpha\theta$.

**B. Incomplete Asset Markets**

Now use (9)-(11), the condition that $w_t = s_t - \eta_t$, and the assumptions that $p_t^* = 0$ to obtain

$$x_t = \kappa(x_{t-1} + E_t x_{t+1}) + \omega(s_t + E_t s_{t+1}) - \psi E_t(\eta_t + \eta_{t+1}), \tag{A7}$$

where $\psi = (1-\alpha)/(2-\alpha\theta)$. Then use (18) and the assumption that $v_t$ and $\eta_t$ are white noise to express (A7) as

$$x_t = [(\kappa - \rho\omega)x_{t-1} + \kappa E_t x_{t+1} + \omega(v_t + \varphi\eta_t) - \psi E_t \eta_t]/(1 + \rho\omega). \tag{A8}$$

Assume a solution of the form

$$x_t = \tilde{\phi}_1 x_{t-1} + \tilde{\phi}_2(v_t + \varphi\eta_t), \tag{A9}$$

since only the value of the composite shock, $v_t + \varphi\eta_t$, can now be inferred from the feedback rule (18). Now (18), (A9) and (10) can be used to derive (16) with $\mu_1 \equiv \tilde{\phi}_1 + \rho\tilde{\phi}_2$ and $\mu_2 \equiv \tilde{\phi}_2$. Note that although agents do not know $\eta_t$, they can derive its expected value as $E_t\eta_t = \lambda(v_t + \varphi\eta_t)$ with $\lambda = \varphi^2\sigma_\eta^2/(\varphi^2\sigma_\eta^2 + \sigma_v^2)$. Using this expression and (A9) to determine $E_t x_{t+1}$, express (A8) as

$$x_t = [(\kappa - \rho\omega + \kappa\tilde{\phi}_1^2)x_{t-1} + (\kappa\tilde{\phi}_1\tilde{\phi}_2 + \omega - \psi\lambda)(v_t + \varphi\eta_t]/(1 + \rho\omega). \tag{A10}$$

Using (A9) and (A10) to determine $\tilde{\phi}_1$ and $\tilde{\phi}_2$, it can be shown that the coefficient of $x_{t-1}$ in (A9) is the same as before, i.e., $\tilde{\phi}_1 = \phi_1$. The coefficient of $v_t + \varphi\eta_t$, however, equals

$\widetilde{\phi}_2 = (\omega - \psi\lambda)/(1 + \rho\omega - \kappa\widetilde{\phi}_1)$. Thus $\mu_2(\equiv \widetilde{\phi}_2)$ decreases in $\lambda$ and can be shown to decrease in $\rho$ as well.

For *T = 2*, use (10) and (A9) to obtain the variance of *p* as

$$\mathrm{var}(p) = \widetilde{\phi}_2^2(\varphi^2\sigma_\eta^2 + \sigma_v^2)/2(1 - \widetilde{\phi}_1^2). \tag{A11}$$

As (A11) indicates, $\mathrm{var}(p)$ decreases in $\rho$ since both $\widetilde{\phi}_1$ and $\widetilde{\phi}_2$ decrease in $\rho$. Note that $\widetilde{\phi}_2$ decreases in $\sigma_\eta^2$, and increases in $\sigma_v^2$. Thus $\mathrm{var}(p)$ increases in $\sigma_v^2$, but the sign of the partial effect of $\sigma_\eta^2$ on $\mathrm{var}(p)$ is ambiguous. Next, use (18) and (A9) to obtain the variance of *s* as

$$\mathrm{var}(s) = (\varphi^2\sigma_\eta^2 + \sigma_v^2)[1 + (\rho^2\widetilde{\phi}_2^2)/(1 - \widetilde{\phi}_1^2)]. \tag{A12}$$

According to (A12), $\mathrm{var}(s)$ increases in $\sigma_v^2$, but the signs of the partial effects of both $\rho$ and $\sigma_\eta^2$ on $\mathrm{var}(s)$ are ambiguous.

Appendix Table

| Coefficients | $\rho$ | $\mu_1$ | $\mu_2$ |
|---|---|---|---|
| $\alpha\theta$ = 0.1 | 0.1 | 0.025 | 0.451 |
| | 0.3 | 0.023 | 0.413 |
| | 0.5 | 0.021 | 0.382 |
| | 0.7 | 0.020 | 0.354 |
| | 0.9 | 0.018 | 0.330 |
| $\alpha\theta$ = 0.5 | 0.1 | 0.165 | 0.330 |
| | 0.3 | 0.153 | 0.306 |
| | 0.5 | 0.143 | 0.286 |
| | 0.7 | 0.134 | 0.268 |
| | 0.9 | 0.127 | 0.253 |
| $\alpha\theta$ = 0.9 | 0.1 | 0.508 | 0.113 |
| | 0.3 | 0.486 | 0.108 |
| | 0.5 | 0.467 | 0.104 |
| | 0.7 | 0.451 | 0.100 |
| | 0.9 | 0.436 | 0.097 |

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