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Abstract: Recent research suggests that the pass-through of exchange rate changes into domestic inflation has declined in many countries since the 1980s. We develop a theoretical model that attributes the change in pass-through (defined as the correlation of inflation with exchange rate changes) to increased emphasis on inflation stabilization by many central banks. This hypothesis is tested on eleven industrial countries between 1971 and 2000. We find widespread evidence of both a decline in pass-through and a decline in the variability of inflation in the 1990s. We also find a statistically significant link between measured pass-through and inflation variability. However, our efforts to correlate the decline in pass-through with estimated changes in monetary policy behavior are inconclusive due to poor estimates of policy behavior.

Keywords: inflation targeting, Taylor rule

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I. Introduction

In the 1990s several countries experienced episodes of large real exchange rate depreciations that did not lead to significant increases in domestic inflation. The experiences of Sweden and the United Kingdom in 1992 are two widely cited examples. A potential explanation of this phenomenon is that central banks in these countries have articulated more or less formally an enhanced commitment to keeping inflation low since the beginning of the 1990s. In such an environment, firms are less keen to pass through fluctuations in their input prices to output prices both because the central bank applies countervailing pressure to aggregate demand contemporaneously and because firms believe that the central bank will be successful in stabilizing inflation in the future.

This paper proposes that the anti-inflationary actions and credibility of the monetary authority are important factors behind the reduced pass-through of exchange rate changes into domestic inflation. We develop a simple theoretical model that explains how monetary policy influences expectations and exchange rate pass-through. In this model, when the monetary authority focuses strongly on stabilizing inflation, there is less pass-through of exchange rate movements into domestic prices. We then examine the monetary and inflation experiences of a sample of industrial countries since the early 1970s, and we find some empirical support for this hypothesis, although the data are inconclusive for the most direct test we implement.

Several previous studies have identified a reduction in exchange rate pass-through across various countries. For example, Cunningham and Haldane (1999) document the low pass-through of sterling depreciation in 1992-93 as well as the low pass-through of sterling appreciation in 1996-97. Taylor (2000) discusses the cases of Sweden and the United Kingdom in 1992-93 and Brazil in 1999. Goldfajn and Werlang (2000) examine episodes of large depreciations in seven emerging markets and five industrial countries in the 1990s. In all cases, Goldfajn and Werlang find that pass-through was less than would have been predicted by their empirical model using data for the 1980s and 1990s. Laflèche (1996-97) discusses the surprisingly low pass-through of the Canadian depreciation of 1992-94 compared with previous pass-through episodes.[1] One of the main contributions of our study is to explicitly compare the evolution of pass-through performance in a number of countries.

Both Taylor (2000) and the Bank of Canada have conjectured that changes in pass-through behavior may be due to changes in the orientation of monetary policy. According to the Bank of Canada's November 2000 *Monetary Policy Report* (p. 9) "the low-inflation environment itself is changing price-setting behavior. When inflation is low, and the central bank's commitment to keeping it low is highly credible, firms are less inclined to quickly pass higher costs on to consumers in the form of higher prices." However, to our knowledge, the theoretical mechanism behind this linkage has not been rigorously derived. Here we present a simple theoretical model that incorporates the link between policy behavior and pass-through.

We test our model's hypothesis on data from eleven countries between 1971 and 2000. We consider five countries with a history of moderately high inflation that adopted explicit and relatively low inflation targets as objectives for monetary policy in the early 1990s. In order of formal adoption of the new regimes, these countries are New Zealand, Canada, the United

[1]In a study of pass-through at different stages of the distribution process, McCarthy (2000) finds that eliminating the years prior to 1983 tends to reduce estimated pass-through in a sample of industrial countries.

Kingdom, Sweden, and Australia.[2] Because of their striking changes in policy regimes, these five countries form a natural experiment with which to test for the impact of monetary policy on exchange rate pass-through. Indeed, two of the five (the United Kingdom and Sweden) adopted inflation targets in response to the ERM crisis of 1992, which sparked a real effective depreciation of the pound and the krona by nearly 20 percent. In addition to these five inflation targeting (IT) countries, we analyze six other industrial countries (Germany, Greece, Japan, Norway, Switzerland, and the United States). Although these other countries did not adopt explicit inflation targets, some of their monetary authorities may have attached greater importance to stabilizing inflation around the same time as the IT countries.

We estimate the pass-through of exchange rate changes to inflation in each of these countries and we find that estimated pass-through declined in the 1990s for all of the IT countries and for four of the six non-IT countries. Next, we show that estimated pass-through coefficients are very significantly correlated with the standard deviation of the inflation rate, even after allowing for a break in behavior in the 1990s. Finally, as a more direct test of our theory, we estimate monetary policy rules for these countries and attempt to correlate components of these policy rules with estimated pass-through behavior. Unfortunately, the relevant components of our estimated policy rules have no significant correlation with our pass-through estimates. This result appears to be due to the poor precision with which we are able to estimate our policy rules.

We develop our theoretical model highlighting the link between monetary policy and

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[2]Freeman and Willis (1995) provide background on the early experiences of the first four of these five. While there is some evidence from long-term interest rates that the new policy regimes may not have been immediately and fully credible, inflation rates did come down faster than almost anyone expected and policy credibility grew with the observed success in fighting inflation.

pass-through in Section II. We test the model's implications in Section III. Concluding comments are in Section IV.

II. A Simple Model

This section explores the relationship between exchange rate pass-through and monetary policy in the context of a theoretical macro model with rational expectations. The model presented here incorporates a simple expectations-augmented Phillips curve. Similar results are also presented using models in which the current inflation rate depends on either lagged inflation or expected future inflation. Our objective is to show how the implied correlation between exchange rate changes and inflation depends on the policy regime. The three equations of our model are:

(1) $$\Delta p_t = {}_{t-1}\Delta p_t + \alpha(e_t + p^*_t - p_t) - \beta(i_t - {}_t\Delta p_{t+1}) + u_t$$

(2) $$e_t = {}_t e_{t+1} - i_t + i^*_t + v_t$$

(3) $$i_t = \pi + \mu(\Delta p_t - \pi) + w_t$$

where $\alpha, \beta, \mu > 0$

Equation (1) states that the inflation rate in period *t* equals the value that was expected in period *t-1* plus the net effect of all shocks in period *t*. In addition to the direct effect of the price shock, *u*, the inflation rate is affected by the action of shocks through the real exchange rate and the real interest rate. A higher real exchange rate (foreign price level compared to domestic price level) tends to increase the domestic price level both indirectly through its positive impact on aggregate demand and directly through the imported goods component of the price level. A higher real interest rate tends to lower the price level indirectly through its implicit effect on aggregate demand. (Note that aggregate demand has been solved out of this model for simplicity.)

Equation (2) is a standard uncovered interest rate parity relation. Expected exchange rate appreciation equates any difference between domestic and foreign interest rates, except for a temporary risk premium, v.

Equation (3) is a simple monetary policy rule, where π is the target inflation rate, μ represents the strength of the monetary authority's response to deviations of inflation from its target, and w is a policy shock. Note that w may also be interpreted as a temporary shock to the inflation target. We will interpret a regime shift toward "inflation targeting" as some combination of an increase in μ and a decrease in the variance of w.[3,4] For simplicity, equation (3) has no lagged adjustment term; this is not a bad approximation if one interprets the time period of the model as annual or longer.

To solve the model, we first assume that p^* and i^* are exogenous and set to zero for simplicity. We also assume that the shocks, u, v, and w are i.i.d., mean zero, and uncorrelated with each other. By substituting equation (3) into (1) and (2) and employing the method of undetermined coefficients we obtain the following solutions for domestic inflation and exchange

[3]Fair (2001) finds that the estimated value of μ nearly doubled in the United States after 1982, when policy succeeded in achieving relatively low and stable inflation.

[4]Inflation targeting has also been associated with a reduction in the mean inflation rate, π, but a permanent shift in the average level of inflation does not by itself affect the correlation of inflation and exchange rate changes in this model.

rate changes.

(4) $$\Delta p_t = \pi + \frac{u_t + \alpha v_t - (\alpha + \beta) w_t}{(1 + \alpha\mu + \beta\mu)}$$

(5) $$\Delta e_t = \pi + \frac{(1-\mu)u_t + \mu u_{t-1} + (1 + \alpha + \beta\mu)v_t - (1 + \beta\mu)v_{t-1} - (1 + \alpha + \beta)w_t + w_{t-1}}{(1 + \alpha\mu + \beta\mu)}$$

For $\mu > 1$ the above solution is stable and unique. For $\mu \leq 1$ there are a multiplicity of nonexplosive solutions. We shall restrict our analysis to the cases in which $\mu > 1$. Equation (6) displays the correlation of inflation and exchange rate changes. Our objective is to find the effect of changes to μ and σ_w^2 on this correlation.

(6) $$\mathit{Correlation}(\Delta p_t, \Delta e_t) = \frac{\mathit{Covariance}(\Delta p_t, \Delta e_t)}{\sqrt{\mathit{Variance}(\Delta p_t)}\sqrt{\mathit{Variance}(\Delta e_t)}}$$

where

$$\mathit{Covariance}(\Delta p_t, \Delta e_t) = \frac{(1-\mu)\sigma_u^2 + \alpha(1 + \alpha + \beta\mu)\sigma_v^2 + (\alpha + \beta)(1 + \alpha + \beta)\sigma_w^2}{(1 + \alpha\mu + \beta\mu)^2}$$

$$\mathit{Variance}(\Delta p_t) = \frac{\sigma_u^2 + \alpha^2\sigma_v^2 + (\alpha + \beta)^2\sigma_w^2}{(1 + \alpha\mu + \beta\mu)^2}$$

$$\mathcal{V}ariance(\Delta e_t) = \frac{[(1-\mu)^2 + \mu^2]\sigma_u^2 + [(1+\alpha+\beta\mu)^2 + (1+\beta\mu)^2]\sigma_v^2 + [(1+\alpha+\beta)^2 + 1]\sigma_w^2}{(1+\alpha\mu+\beta\mu)^2}$$

While we cannot obtain an analytic expression to determine the signs of the derivatives of equation (6) with respect to μ and σ_w^2, we can evaluate equation (6) for particular numerical values of the parameters and determine the nature of their effect over a wide range of the parameter space.

We begin by setting α, β, and μ to values consistent with the literature. For α, we consider both the direct and indirect effects of exchange rates on prices. The direct effect derives from pass-through of exchange-rate-adjusted foreign prices into domestic prices of imported goods. Goldberg and Knetter (1997) provide a discussion of the microeconomic pass-through literature. Pass-through at the micro level typically is 50 to 100 percent of an exchange rate change. The share of imports in final demand ranges from around 10 percent in Japan and the United States to 40 percent in Canada. Assuming pass-through of 75 percent and an import share in final demand of 20 percent implies a direct effect of the real exchange rate on prices of 0.19. With an import share of 10 percent, the direct effect would be 0.08. The indirect effect is a function of two components: 1) the effect of the real exchange rate on aggregate demand, and 2) the effect of aggregate demand on inflation holding expected inflation constant. At an annual frequency, the impact of a 1 percent increase in aggregate demand on prices typically is estimated to be around 0.1 percent.[5] With trade price elasticities at or below unity, and trade shares of aggregate demand well below unity, the indirect effect of the real exchange rate on prices is well below 0.1. Thus, reasonable values of α are around 0.1 for the United States and Japan, and 0.2 to 0.4 for smaller, more open, economies.[6] In our numerical solutions we use $\alpha=0.1$ and $\alpha=0.2$.

Research on monetary conditions indicators suggests that the ratio of β to α is likely to be between 2 and 10 for most industrial countries.[7] This ratio reflects the relative strengths of the interest rate and exchange rate channels of monetary transmission.

Finally, our estimate of μ comes from the empirical literature. Taylor (1993) found that a policy rule reacting to deviations of both the inflation rate and potential output with coefficients of 1.5 and 0.5, respectively, tracked the U.S. federal funds rate quite well in the 1980s. Others, including Clarida, Gali, and Gertler (1998) and Fair (2001), estimate variants of this equation for the United States and other countries. Their results yield estimates of μ between 1.1 and 2.0. Since our theoretical model solves out aggregate demand, the appropriate value of μ may be somewhat higher than would be implied by model estimates that include an aggregate demand or output gap term. We believe a reasonable estimate of μ for our model is between 1 and 3, which is consistent with the stable solution of the model.

Using these ranges of values for the parameters, Table 1 displays the predicted

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[5]See, for example, Brayton, Roberts, and Williams (1999).

[6]Another way to calibrate this equation is to note that when $\alpha=1$, real exchange rate swings are associated with contemporaneous swings of inflation of equal magnitude, whereas the evidence for most industrial countries is that real exchange rate movements are much greater than movements in inflation. Thus, values of α well below 1 are more plausible.

[7]See, for example, Freedman (1994) and IMF (1996).

correlation between exchange rate changes and inflation. The absolute magnitudes of the shock variances are irrelevant--only the relative magnitudes matter. The entries of Table 1 are arranged in groups of three rows. In each group we vary $\mu \in \{1.01, 1.5, 3.0\}$ but all other parameters are identical; thus the movement in the inflation-exchange rate correlation within each triplet captures the derivative of the correlation with respect to μ for each combination of the remaining parameters. Across the triplets, we test the sensitivity of our results to alternative values for the remaining parameters of the model one at a time.

The correlations between inflation and exchange rate changes under the basic model are displayed in the third column from the right. In every triplet, an increased emphasis on inflation in the monetary policy rule, raising the value of μ, reduces the correlation--in some cases to a negative value. By comparing the last triplet with the first triplet, it can be seen that reducing the variance of the shock to monetary policy, σ_w^2, also reduces the correlation of inflation and exchange rate changes. Hence, monetary policies that increase the emphasis on inflation stabilization reduce pass-through in our model. Looking across the various triplet experiments we see that these results are not sensitive to any of the remaining parameter values of the model.

The final two columns of Table 1 present the correlations of inflation and exchange rate changes under two variants of equation (1),[8] one of which is entirely forward-looking and the other of which is entirely backward-looking. As can be seen from Table 1, the implications of

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[8]These forward-looking and backward-looking equations are as follows:

$$\Delta p_t = {}_t\Delta p_{t+1} + \alpha(e_t + p^*_t - p_t) - \beta(i_t - {}_t\Delta p_{t+1}) + u_t$$
$$\Delta p_t = \Delta p_{t-1} + \alpha(e_t + p^*_t - p_t) - \beta(i_t - {}_t\Delta p_{t+1}) + u_t$$

changes in the policy parameters, μ and σ_w^2, in these models are qualitatively similar to those under the basic model. While the correlations under the forward-looking model are quantitatively close to those under the basic model, the correlations under the backward-looking model display an even greater sensitivity to increases in the monetary feedback from inflation, μ.

The results for the forward- and backward-looking models are especially interesting because the popular class of staggered contracts models of inflation combines elements of both backward-looking and forward-looking price adjustment. Note that with forward-looking price adjustment, the credibility of future central bank policy plays a role in reducing pass-through.

III. Evidence

Our search for evidence on the relationship between monetary policy and exchange rate pass-through proceeds as follows. First, we look at individual countries over time to identify cases in which exchange rate pass-through decreased. Then, we examine whether this change in pass-through may be related to changes in the level and/or variability of inflation. Finally, we attempt to determine whether the changes in pass-through and inflation variability can be attributed to changes in monetary policy rules as described in the preceding theoretical model.

We focus on data from five countries that implemented inflation targeting regimes in the 1990s (Australia, Canada, Sweden, New Zealand and the United Kingdom) as well as six additional developed countries (Germany, Greece, Japan, Norway, Switzerland and the United States). These countries are all the industrial countries for which we could obtain data and that did not follow a fixed exchange rate regime for a significant fraction of the past three decades.

III.A. Pass-Through in Industrial Countries

We estimate the following pass-through equation for each country over the sample 1971:Q3 through 2000:Q4:[9]

(7) $$\Delta p_t = \gamma_0 + \gamma_1 \Delta p_{t-1} + \gamma_2 \Delta(e_t + p^*_t) + \left[\gamma_3 + \gamma_4 \Delta p_{t-1} + \gamma_5 \Delta(e_t + p^*_t)\right] D_t$$

The variables p, e, and p^* are the quarterly consumer price index, trade-weighted exchange rate, and trade-weighted foreign consumer price index, respectively. All variables are seasonally adjusted. The dummy variable, D, equals zero prior to the announcement of the inflation target regime for IT countries and one afterwards. For the other countries, we set the dummy breakpoint at 1990:Q1. We also include quarterly dummy variables in some countries to control for changes in indirect taxes that affect consumer prices.[10] The coefficient γ_2 represents the immediate impact of an exchange rate change or foreign price level change on the domestic price level. The equation incorporates lagged adjustment of inflation to shocks, so that $\gamma_2/(1-\gamma_1)$ measures the long-run pass-through of exchange rate movements to overall inflation prior to the 1990s. $(\gamma_2+\gamma_5)/(1-\gamma_1-\gamma_4)$ is the long-run pass-through during the 1990s.

Table 2 reports the coefficient values and standard errors of OLS regressions for our eleven countries.[11] Prior to the 1990s, the autoregressive coefficients, γ_1, range from 0.3 to 0.8,

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[9]Due to limited data, the United Kingdom sample begins in 1975:Q3.

[10]The dates of the inflation targeting regimes and more details of the data are found in Appendix A.

[11]Because we wish to allow for the possibility of a shift in the error variance, we report heteroskedasticity-consistent standard errors. On the other hand, since we have dummy variables in these regressions, we also report the non-robust standard errors.

implying relatively quick pass-through of exchange rate changes and foreign inflation. In most countries, inflation adjustment is even faster in the 1990s, as γ_4 is negative, and in no country is adjustment significantly slower in the 1990s. F-tests indicate that the structural break in the 1990s is significant at the 5 percent level or lower for all countries except Germany and Greece. Q and LM tests, with lags from one to four quarters, do not reject the null of no autocorrelation for most countries.

Table 3 reports the long-run pass-through coefficients, which are a close analogue to the inflation-exchange rate correlations in our theoretical model once we allow for lagged adjustment. The pass-through coefficients prior to 1990 range from values insignificantly different from zero to significant positive values above 0.2 for Canada, Greece, Japan, and Switzerland. A pass-through coefficient of 0.2 implies that a 10 percent exchange rate depreciation would raise domestic prices by 2 percent. Estimated pass-through coefficients decline in the 1990s in all countries except Greece and the United States. The changes are statistically significant only for Canada and Japan, but the magnitudes of the changes are often economically significant. Long-run pass-through fell from an average value of 0.12 pre-1990 to 0.05 post-1989. For IT countries, average pass-through dropped sharply after 1989 from 0.12 to 0.01. The average pass-through of the non-IT countries declined modestly from 0.12 to 0.09.

III.B. Inflation Variability and Pass-Through in Industrial Countries

Both the mean and the standard deviation of inflation have fallen in each of our 11 countries since 1990 (or since the advent of inflation targeting for IT countries). The average rate of inflation in the 1970s and 1980s was 8.2 percent compared to 2.9 percent in the 1990s. For IT countries, the average rate fell even more, from 8.9 percent to 2.1 percent. For non-IT countries it fell from 7.6 to 3.5 percent. A similar story holds true for the standard deviation of quarterly inflation. For all 11 countries on average, the standard deviation of inflation fell from 4.7 percent in the 1970s and 1980s to 2.3 percent in the 1990s. For IT countries, the decline was greater, from 4.9 to 1.9 percent. For non-IT countries, the standard deviation fell from 4.5 to 2.5 percent.

Table 4 provides information on the relationship between estimated exchange rate pass-through coefficients and inflation behavior in these countries. The 22 long-run pass-through coefficients from Table 3 are regressed on the mean and/or standard deviation of inflation for the respective countries and sample periods. To control for any omitted common factor in the 1990s, a dummy variable is also included that equals zero prior to 1990 and one afterwards. The first column presents the preferred specification, where the standard deviation of inflation has a strongly significant (1 percent level) effect on pass-through. The coefficient implies that a 1 percentage point increase in the standard deviation of inflation would increase the pass-through coefficient by 0.034, or more than half of the average of the estimated pass-through coefficients in the 1990s.

The second column shows that this result is not sensitive to inclusion of a dummy variable for the 1990s, which is insignificant. The third column shows that inclusion of the mean rate of inflation has only a modest effect on the value of the coefficient on the standard deviation of inflation, but, due to the collinearity of the mean and standard deviation of inflation, the standard errors on both coefficients are quite high. While the mean of inflation is significant in the absence of the standard deviation (column four), the adjusted R^2 of this specification is lower than that of any specification that includes the standard deviation of inflation. In other regressions (not shown) inclusion of an additional dummy variable that equals one for IT countries in the 1990s and zero for all other countries and sample periods has no material effect on the results.

Given the above empirical results, and our theoretical model, we are inclined to the conclusion that lower variability of inflation is an important factor behind lower pass-through, and that the lower average rates of inflation are not important, *per se.* However, a reasonable conjecture for a country implementing an inflation target, where the target is lower than the pre-existing rate of inflation, is that credibility may be linked to movement in the inflation rate towards the target. We believe that, over time, market attention most likely would switch to the success of the central bank in maintaining the inflation rate within a narrow range--in other words, reducing the variability of inflation.

III.C. Monetary Policy in Industrial Countries

In this section we test for a connection between monetary policy and the pass-through results discussed above. We start by estimating a policy rule similar to equation (3) of our theoretical model and allowing for a change in the parameters in the 1990s.[12,13] For each country, we estimate:

(8) $$i_t = \theta_0 + \theta_1 i_{t-1} + \theta_2 i_{t-2} + \theta_3 \Delta p_t + \left(\theta_4 + \theta_5 i_{t-1} + \theta_6 i_{t-2} + \theta_7 \Delta p_t\right) D_t$$

where *i* is the end-of-quarter nominal interest rate on the three-month Treasury bill, *p* is the

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[12]Nelson (2000) follows a similar strategy. He splits the U.K. sample into five distinct policy regimes and finds a stronger long-run response to inflation in the post-1992 period than in earlier periods.

[13]Clarida, Gali, and Gertler (1998) use future inflation in their policy rules, but Fair (2001) reports that the estimates for the United States are not sensitive to using current inflation. Use of current inflation allows for a simpler estimation technique.

quarterly consumer price index, and D is a dummy variable for the 1990s as above.[14] All first differences are quarterly percentage changes at annual rates to be consistent with the annualized interest rates. We estimate equation (8) on quarterly data from 1972:Q2 to 2000:Q4 for each country using instrumental variables.[15] The use of instrumental variables reflects the fact that the central bank must estimate current inflation using data that is reported with a lag. The first-stage regressions share the same breakpoints as the second stage.

The inclusion of lagged interest rates in equation (8) is standard in the literature.[16] This specification reflects the slow adjustment of policy interest rates to economic news at a quarterly frequency. It is also standard practice to include a measure of the output gap or a deviation of the unemployment rate from the natural rate. We do not include such a term for two reasons: 1) our theoretical model solved out the aggregate demand effect for tractability and we want our empirical results to correspond to that model, and 2) any estimate of the output gap or natural rate of unemployment introduces a new level of complexity and measurement error to the analysis, since these concepts must be estimated under assumptions that are not universally accepted.

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[14]We include quarterly dummy variables to control for changes in indirect taxes that are included in consumer prices but are generally not targeted by central banks.

[15]The instruments for current inflation are four lags of inflation, two lags of an estimated output gap, and two lags of exchange-rate adjusted foreign inflation. Due to limited data on interest rates for New Zealand, Norway, Switzerland, and the United Kingdom, our sample for these countries begins in 1974:Q4, 1972:Q3, 1973:Q1, and 1976:Q2, respectively.

[16]The second lag of the interest rate was significant for Germany, Japan, New Zealand, and Norway. Clarida, Gali, and Gertler (1998) and Fair (2001) find two quarterly lags to be important for U.S. interest rates. Using only one lagged interest rate does not significantly change any coefficient reported in Tables 6 and 7, nor does it have any noticeable impact on the pattern of results reported in Table 7.

The coefficient θ_3 represents the immediate response of the monetary authority to inflation in the pre-1990 sample. The expression $\theta_3/(1-\theta_1-\theta_2)$ represents the long-run response to inflation over the same period in the presence of slow adjustment ($\theta_1+\theta_2>0$). In the 1990s, the long-run response is given by $(\theta_3+\theta_7)/(1-\theta_1-\theta_2-\theta_5-\theta_6)$. Table 5 reports the coefficient estimates of the policy rules for the eleven countries in our sample. Only for Greece, Japan, and Norway does the F-test indicate that the 1990s break is significant at the 5 percent level. Q and LM tests indicate that we cannot reject the null of no autocorrelation for most countries.

Our theoretical model abstracts from lagged adjustment for simplicity, and it has been calibrated roughly to correspond to an annual or lower frequency. We believe that the estimated long-run responses of monetary policy to inflation are the relevant empirical analogues to the parameter μ in the theoretical model. With typical estimated adjustment lags of around 0.8 quarterly, 60 percent of the long-run effect is transmitted within one year. For a country that shifts to an IT regime, or puts more emphasis on low inflation, we expect to find an increase in the estimated long-run inflation response. Additionally, if interest rates are responding solely to movements away from an inflation target, then the policy rule specified here should fit the data better (with a smaller residual standard deviation) after the adoption of the new policy.

Table 6 reports the long-run inflation coefficient estimates and the standard deviations of the residuals in the pre-1990 and post-1989 samples for our countries. We find that nine of the 11 countries have larger long-run inflation coefficients in the 1990s than in the pre-1990 period, including all five IT countries. The average long-run inflation coefficient for IT countries rises markedly from 0.44 to 1.08. The average for non-IT countries edges up only slightly from 0.68 to 0.72. For the IT countries on average, central banks do appear to have reacted more strongly to inflation in the 1990s.

Perhaps the most important point to make about the inflation coefficients is that they have very high standard errors in the 1990s period. This, in part, reflects the fact that the post-1989 sample period has few observations, and that movements of the regressors (as measured by their standard deviations) fell dramatically from pre-1990 to post-1989.

Another point to note is that most of the individual inflation coefficients in Table 6 are below unity, which is the lower bound for a unique and stable solution to our theoretical model. We believe that our coefficients are biased downward partly because we use the volatile "headline" consumer price index instead of core domestic prices.[17] We use the broad consumer price index because a consistent measure of core inflation is not available for many countries. Another source of downward bias may be a change in the monetary authority's inflation target within either of our subsamples. For example, a monetary authority that pursued a strategy of "opportunistic disinflation" would appear to have a weak reaction to changes in inflation when inflation is falling, as it was in the United States during the 1990s.[18] Finally, the low monetary responses to inflation prior to the 1990s in some of these countries may be part of the reason that they experienced great macroeconomic instability during the 1970s and 1980s. In other words, choosing a policy parameter that is associated with multiple or explosive solutions in the theoretical model may lead to instability in the real economy.

Turning to the policy rule residuals, the average standard deviation fell from a pre-1990

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[17]In the United States, core inflation is typically defined using the CPI excluding food and energy.

[18]See Orphanides and Wilcox (1996).

value of 1.49 to a post-1989 value of 0.62. This corresponds to a substantial decrease in σ^2_w in our model. The standard deviation declined more in the IT countries, where it fell from an average of 1.76 to 0.60. In the non-IT countries the residual standard deviation fell from an average value of 1.26 to 0.63.

These results suggest that for IT countries, at least, monetary policy behavior has shifted emphasis toward stabilizing inflation. But is there a relationship between the emphasis on inflation in monetary policy and the degree of exchange rate pass-through? Our theoretical model suggests that an increased emphasis on inflation by the monetary authority translates into less pass-through. To test this hypothesis, we regress the estimated pass-through coefficients on the inflation coefficients and the standard deviations of the estimated policy rules. We also included two dummy variables: a 1990s dummy equal to zero in the early sample period and one in the 1990s, and an IT dummy equal to one for IT countries in the 1990s and zero for other countries and other sample periods. We start with all four independent variables and successively delete the least significant. The results are presented in Table 7.

The final column shows that with all the variables in the regression, no variable is significant at any level. Sequentially deleting the least significant variable and rerunning the regression revealed that the IT dummy is significant in combination with the inflation coefficients when the other variables are excluded. However, the relationship between the inflation coefficients and the pass-through coefficients has the wrong sign--increased monetary responses to inflation tend to increase pass-through--and this effect is not statistically significant. On its own, the IT dummy is nearly significant at the 10 percent level (p-value of .13) indicating that IT regimes may have a negative effect on pass-through. Our interpretation of these results is that the estimated policy rule coefficients and standard deviations are very imprecise measures of monetary policy. This conclusion is based largely on the large standard errors of the inflation coefficients and the wide disparity in inflation coefficients across countries, as well as their theoretically implausible values in many cases.

To check on the robustness of our results, and to see if we could obtain more precise estimates of monetary policy and pass-through behavior, we tried augmenting the policy rule and pass-through regressions with the Federal Reserve Board staff's estimated output gaps, and the overall results are similar to the results presented here. We also estimated the policy rules and pass-through equations after adding oil prices and obtained similar results.

IV. Conclusion

This paper documents a decline in measured exchange rate pass-through at the macroeconomic level for many industrial countries since 1990. We develop a theoretical model to explain how such a development could be the consequence of a shift in the monetary authority's responsiveness to inflation. When agents expect the monetary authority to act strongly to stabilize the domestic inflation rate, they are less inclined to change prices in response to a given exchange rate shock. We present evidence for a sample of 11 industrial countries that supports this hypothesis indirectly by establishing a connection between pass-through behavior and inflation stability. We also show that inflation-targeting (IT) countries exhibited a marked change in monetary behavior in the 1990s and that pass-through declined more sharply in IT countries than elsewhere. However, we are unable to show a direct correlation between measures of monetary behavior and pass-through.

A natural extension of this paper is to include developing countries in the empirical analysis. In a paper that focuses primarily on the level of the real exchange rate as a contributor to excess demand, Kamin (1998) finds that the pass-through of exchange rate depreciations into domestic inflation is much greater in Latin American countries than in developing Asian or industrial countries. Kamin presents evidence that at least part of this difference in Latin American pass-through is due to an association of exchange rate depreciations with increased inflationary expectations in Latin American countries. Such a finding is consistent with the conclusions of this paper and it points out a promising avenue for future research.

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Table 1 - Theoretical Model Correlations

α	β	μ	σ_u^2	σ_v^2	σ_w^2		$Correlation(\Delta p, \Delta e)$		
							Basic	Forward	Backward
0.1	0.5	1.01	0.1	0.1	0.1		0.31	0.34	0.63
0.1	0.5	1.5	0.1	0.1	0.1		0.16	0.19	0.16
0.1	0.5	3	0.1	0.1	0.1		-0.12	-0.05	-0.15
0.2	0.5	1.01	0.1	0.1	0.1		0.39	0.40	0.51
0.2	0.5	1.5	0.1	0.1	0.1		0.24	0.26	0.17
0.2	0.5	3	0.1	0.1	0.1		-0.04	0.04	-0.10
0.1	1	1.01	0.1	0.1	0.1		0.44	0.54	0.54
0.1	1	1.5	0.1	0.1	0.1		0.30	0.44	0.19
0.1	1	3	0.1	0.1	0.1		0.07	0.28	-0.03
0.1	0.5	1.01	0.2	0.1	0.1		0.22	0.24	0.68
0.1	0.5	1.5	0.2	0.1	0.1		0.02	0.04	0.15
0.1	0.5	3	0.2	0.1	0.1		-0.28	-0.23	-0.21
0.1	0.5	1.01	0.1	0.2	0.1		0.29	0.31	0.57
0.1	0.5	1.5	0.1	0.2	0.1		0.16	0.19	0.15
0.1	0.5	3	0.1	0.2	0.1		-0.07	-0.04	-0.11
0.1	0.5	1.01	0.1	0.1	0.2		0.44	0.47	0.63
0.1	0.5	1.5	0.1	0.1	0.2		0.30	0.36	0.21
0.1	0.5	3	0.1	0.1	0.2		0.02	0.17	-0.09

Table 2 - Pass-Through Regressions, Allowing for Regime Shift

$$\Delta p_t = \gamma_0 + \gamma_1 \Delta p_{t-1} + \gamma_2 \Delta(e_t + p^*_t) + \left[\gamma_3 + \gamma_4 \Delta p_{t-1} + \gamma_5 \Delta(e_t + p^*_t)\right] D_t$$

	γ_0	γ_1	γ_2	γ_3	γ_4	γ_5	R^2
Australia	3.74*** (0.82) [1.04]	0.51*** (0.08) [0.11]	0.04* (0.02) [0.02]	-2.15** (1.19) [1.14]	-0.34 (0.25) [0.21]	-0.04 (0.05) [0.03]	0.53
Canada	1.10*** (0.45) [0.49]	0.78*** (0.06) [0.06]	0.06** (0.02) [0.02]	0.62 (0.62) [0.57]	-0.64*** (0.15) [0.09]	-0.08 (0.04) [0.03]	0.84
Germany	0.97*** (0.37) [0.29]	0.71** (0.08) [0.07]	0.02* (0.01) [0.01]	0.09 (0.57) [0.49]	-0.21 (0.15) [0.20]	0.01 (0.02) [0.02]	0.53
Greece	4.56*** (1.52) [1.76]	0.63*** (0.07) [0.12]	0.10** (0.03) [0.05]	-3.83* (2.11) [1.90]	0.21 (0.14) [0.14]	-0.03 (0.10) [0.08]	0.64
Japan	1.60** (0.57) [0.87]	0.66*** (0.07) [0.19]	0.08** (0.02) [0.04]	-1.02 (0.85) [0.91]	-0.41* (0.30) [0.24]	-0.09** (0.04) [0.04]	0.62
New Zealand	4.71*** (1.14) [1.93]	0.54*** (0.08) [0.18]	0.05** (0.02) [0.02]	-4.06** (1.50) [1.95]	0.10 (0.35) [0.21]	-0.03 (0.06) [0.03]	0.65
Norway	3.61*** (0.81) [0.81]	0.51*** (0.08) [0.09]	0.06 (0.04) [0.05]	-1.68 (1.15) [0.92]	-0.24 (0.30) [0.17]	-0.09 (0.07) [0.06]	0.65
Sweden	5.30*** (0.77) [0.83]	0.31*** (0.09) [0.11]	0.04** (0.02) [0.03]	-4.68*** (1.05) [0.88]	-0.06 (0.26) [0.18]	-0.02 (0.04) [0.28]	0.59
Switzerland	1.31*** (0.39) [0.36]	0.62*** (0.07) [0.10]	0.08*** (0.02) [0.03]	-1.02* (0.60) [0.42]	0.20 (0.16) [0.12]	-0.07** (0.03) [0.03]	0.57
United Kingdom	2.81*** (0.50) [0.59]	0.58 (0.05) [0.08]	0.02 (0.02) [0.02]	-0.03 (1.85) [0.74]	-0.66 (0.66) [0.16]	-0.01 (0.03) [0.02]	0.81
United States	1.35*** (0.45) [0.51]	0.82 (0.06) [0.08]	-0.03 (0.02) [0.02]	0.16 (0.83) [0.63]	-0.37 (0.22) [0.15]	0.08** (0.04) [0.03]	0.73

Note: D_t = 1 in Inflation Targeting regime or, for other countries, D_t=1 for 1990:Q1-2000:Q4. Non-robust standard errors in parenthesis, ( ). Heteroskedasticity-consistent standard errors in brackets, [ ]. *, **, and *** indicate significance at the10, 5, and 1 percent levels respectively.

Table 3 - Long-Run Pass-Through Coefficients

	Pre-1990	Post-1989
Australia	0.080*	0.003
	(0.043)	(0.062)
Canada	0.278***	-0.021
	(0.102)	(0.039)
Germany	0.070	0.051
	(0.042)	(0.041)
Greece	0.261***	0.404
	(0.095)	(0.554)
Japan	0.224***	-0.013
	(0.070)	(0.037)
New Zealand	0.117**	0.054
	(0.057)	(0.162)
Norway	0.114	-0.043
	(0.088)	(0.080)
Sweden	0.057**	0.023
	(0.026)	(0.048)
Switzerland	0.222***	0.056
	(0.064)	(0.156)
United Kingdom	0.036	0.006
	(0.035)	(0.020)
United States	-0.149	0.103
	(0.140)	(0.061)
Average for All countries	0.119	0.057
Average for IT countries	0.114	0.013
Average for Other countries	0.123	0.093

Note: *, **, and *** indicate significance at the 10, 5, and 1 percent levels, respectively.

Table 4 - Long-run Pass-through Coefficients and Inflation

Intercept	-0.030 (.045)	-0.062 (.074)	-0.073 (.077)	-0.039 (.072)
1990s dummy		0.032 (.057)	0.046 (.062)	0.041 (.062)
Standard Deviation of Inflation	0.034*** (.011)	0.039** (.014)	0.027 (.023)	
Mean of Inflation			0.008 (.012)	0.019** (.008)
Adj. R^2	.284	.259	.235	.222

Note: We regressed long-run pass-through coefficients (Table 3) on the listed variables. There are 22 observations based on 11 countries and 2 sample periods (pre-1990 and post-1989). The 1990s dummy equals zero in the first sample period and one in the second.

Table 5 - Empirical Policy Rules, Allowing for Regime Shift

$$i_t = \theta_0 + \theta_1 i_{t-1} + \theta_2 i_{t-2} + \theta_3 \Delta p_t + \left(\theta_4 + \theta_5 i_{t-1} + \theta_6 i_{t-2} + \theta_7 \Delta p_t\right) D_t$$

	θ_0	$\theta_1+\theta_2$	θ_3	θ_4	$\theta_5+\theta_6$	θ_7	R^2
Australia	0.61 (0.57) [0.56]	0.92*** (0.04) [0.05]	0.03 (0.04) [0.05]	0.36 (1.40) [0.64]	-0.14 (0.25) [0.08]	0.13 (0.31) [0.11]	0.89
Canada	0.99 (0.56) [0.84]	0.85*** (0.05) [0.07]	0.08 (0.06) [0.07]	-0.37 (1.16) [1.00]	-0.07 (0.14) [0.10]	0.13 (0.38) [0.19]	0.88
Germany	0.76** (0.29) [0.23]	0.80*** (0.06) [0.05]	0.16** (0.08) [0.09]	-0.65 (0.46) [0.26]	0.15 (0.10) [0.06]	-0.12 (0.15) [0.10]	0.89
Greece	-0.17 (0.48) [0.45]	0.97*** (0.02) [0.03]	0.05** (0.02) [0.04]	0.63 (0.66) [0.68]	-0.14*** (0.05) [0.06]	0.15*** (0.05) [0.06]	0.97
Japan	1.63*** (0.29) [0.55]	0.64*** (0.05) [0.10]	0.16*** (0.03) [0.05]	-1.63*** (0.34) [0.55]	0.33*** (0.09) [0.10]	-0.16 (0.16) [0.06]	0.95
New Zealand	0.62 (1.22) [1.68]	0.82*** (0.06) [0.09]	0.16** (0.06) [0.08]	0.43 (1.60) [1.72]	-0.04 (0.16) [0.11]	0.09 (0.29) [0.12]	0.84
Norway	1.91** (0.92) [0.92]	0.85*** (0.06) [0.05]	-0.01 (0.08) [0.10]	-0.98 (1.38) [1.20]	-0.02 (0.10) [0.14]	0.11 (0.40) [0.27]	0.84
Sweden	1.47 (0.78) [0.96]	0.84*** (0.05) [0.08]	0.02 (0.07) [0.12]	-1.11 (1.12) [1.00]	0.01 (0.15) [0.10]	0.27 (0.34) [0.16]	0.82
Switzerland	0.65** (0.30) [0.39]	0.72*** (0.08) [0.10]	0.11* (0.06) [0.09]	-0.48 0.43) [0.41]	0.08 (0.15) [0.12]	0.16 (0.20) [0.13]	0.79
United Kingdom	2.24*** (0.74) [0.86]	0.77*** (0.07) [0.08]	0.06 (0.05) [0.08]	-0.41 (2.13) [1.09]	-0.15 (0.32) [0.12]	0.15 (0.82) [0.28]	0.87
United States	1.34** (0.52) [0.94]	0.78*** (0.06) [0.11]	0.06 (0.05) [0.07]	-0.90 (0.98) [0.97]	0.09 (0.21) [0.12]	-0.02 (0.27) [0.11]	0.78

Note: D_t = 1 in Inflation Targeting regime or, for other countries, D_t=1 for 1990:Q1-2000:Q4. Non-robust standard errors in parenthesis, ( ). Heteroskedasticity-consistent standard errors in brackets, [ ]. *, **, and *** indicate significance at the10, 5, and 1 percent levels respectively.

Table 6 - Long-run Inflation Coefficients and Standard Deviations of Residuals

	Long-Run Inflation Coefficient		Standard Deviation of Residuals	
	Pre-1990	Post-1989	Pre-1990	Post-1989
Australia	0.39 (0.56)	0.72 (1.25)	1.50	0.46
Canada	0.52 (0.39)	0.96 (1.81)	1.46	0.80
Germany	0.92*** (0.33)	0.98 (2.00)	1.17	0.35
Greece	1.41 (1.09)	1.11*** (0.14)	0.89	0.84
Japan	0.45*** (0.05)	-0.08 (5.48)	1.04	0.30
New Zealand	0.91* (0.48)	1.16 (1.14)	2.53	0.85
Norway	-0.06 (0.51)	0.61 (2.27)	1.51	1.40
Sweden	0.13 (0.44)	2.02 (2.43)	1.79	0.56
Switzerland	0.39* (0.19)	1.31*** (0.45)	1.35	0.56
United Kingdom	0.25 (0.19)	0.56 (1.96)	1.53	0.35
United States	0.30 (0.23)	0.39 (1.89)	1.61	0.34
Average for All Countries	0.57	0.89	1.49	0.62
Average for IT Countries	0.44	1.08	1.76	0.60
Average for Other Countries	0.68	0.72	1.26	0.63

Note: *, **, and *** indicate significance at the 10, 5, and 1 percent levels, respectively.

Table 7 - Pass-through and Monetary Policy

Intercept	0.110*** (.029)	0.062 (.046)	0.059 (.042)	0.162 (.111)
IT dummy	-0.097 (.061)		-0.142** (.065)	-0.118 (.074)
Inflation Coefficient		0.037 (.054)	0.088 (.055)	0.097 (.057)
Policy Rule Standard Div.				-0.062 (.068)
1990s dummy				-0.100 (.085)
Adj. R^2	.067	-.023	.135	.107

Note: We regressed long-run pass-through coefficients (Table 3) on the listed variables, including the inflation coefficients and residual standard errors from Table 6. There are 22 observations based on 11 countries and 2 sample periods (pre-1990 and post-1989). The 1990s dummy equals zero in the first sample period and one in the second.

Appendix A

Start of the Inflation Targeting Regime for selected Countries

country	Start of IT regime
Australia	1993:2
Canada	1991:1
New Zealand	1990:2
Sweden	1993:1
United Kingdom	1992:4

R= nominal 3 month interest rate, annualized

country	series	source
Australia	13-week Treasury note yield	Haver
Canada	3-month Treasury bill rate	IFS
Germany	3-month interbank rate (3-month interbank rate)	INTL/FRB (Haver)
Greece	3-month Treasury bill rate (Commercial bank deposit rate)	IFS (IFS)
Japan	3-month Gensaki rate	Haver
New Zealand	3-month Treasury bill rate (90-day bank bill rate)	Haver (OECD)

country	series	source
Norway	3-month interbank rate (Call money rate)	BIS (IFS)
Sweden	3-month Treasury bill rate (3-month Treasury discount note rate)	INTL/FRB (IFS)
Switzer-land	Treasury bill rate (Call money rate)	IFS (OECD)
U.K.	3-month interbank rate (91-day Treasury bill tender rate)	INTL/FRB (IFS)
U.S.	3-month Treasury bill rate	IFS

Series and sources in parentheses used to estimate missing periods in primary data source.

π = quarterly domestic inflation

The series that the central bank currently targets

country	series	source
Australia	CPI*	INTL/RBA
Canada	CPI*	INTL/BOC
Germany	CPI, SA	INTL/Bun-desbank
Greece	CPI, SA	Haver
Japan	CPI, SA	Haver
New Zealand	CPI*	Haver

country	series	source
Norway	CPI*	Haver
Sweden	CPI, SA	Haver
Switzer-and	CPI, SA	Haver
U.K.	RPIX*	INTL/CSO
U.S.	CPI, SA	US/BLS

***SA by authors**

π*=Exchange-rate adjusted foreign consumer prices, quarterly rate

Constructed by authors as $\pi/\Delta RER$, where the real exchange rate, RER (SA), is measured as foreign/domestic currency.

country	π series	π source	RER series	RER source
Australia	CPI*	INTL/RBA	Real exchange rate (trade-weighted, time-varying)	REX/FRB
Canada	CPI*	INTL/BOC	Real exchange rate (trade-weighted, time-varying)	REX/FRB
Germany	CPI, SA	INTL/Bun-desbank	Real exchange rate (trade-weighted, time-varying)	REX/FRB
Greece	CPI, SA	Haver	Real effective exchange rate	OECD
Japan	CPI, SA	Haver	Real exchange rate (trade-weighted, time-varying)	REX/FRB
New Zealand	CPI*	Haver	Real effective exchange rate	OECD
Norway	CPI*	Haver	Real effective exchange rate	OECD
Sweden	CPI, SA	Haver	Real exchange rate (trade-weighted, time-varying)	REX/FRB
Switzer-land	CPI, SA	Haver	Real exchange rate (trade-weighted, time-varying)	REX/FRB
U.K.	RPIX*	INTL/CSO	Real exchange rate (trade-weighted, time-varying)	REX/FRB
U.S.	CPI, SA	US/BLS	Real exchange rate (trade-weighted, time-varying)	REX/FRB

***SA by authors**

Individual Country Tax Dummies*
(Dummies for changes in tax policies)

Country	Tax Policy Change
Australia	2000:3
Canada*	1991:1, 1994:1, 1994:2
Greece	1994:2, 1996:1
Japan	1989:2, 1997:2
Sweden	1991:1, 1992:1, 1993:1
U.K.	1979:3

* All dummies set equal to one in the appropriate quarter, except Canada's 1994 VAT change that was phased in over two quarters so we set the dummy as 1994:1 = 2/3, 1994:2 = 1/3.

D = Dummy Variable during Post Inflation Targeting Regime or 1990s

country	D=1
Australia	1993:2-2000:4
Canada	1991:1-2000:4
Germany	1990:1-2000:4
Greece	1990:1-2000:4
Japan	1990:1-2000:4
New Zealand	1990:2-2000:4
Norway	1990:1-2000:4
Sweden	1993:1-2000:4
Switzerland	1990:1-2000:4
U.K.	1992:4-2000:4
U.S.	1990:1-2000:4