Exchange Rate Pass-Through and the Inflation Environment in Industrialized Countries: An Empirical Investigation*

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May 2005

* For helpful comments and suggestions on earlier versions of this paper, we would like to thank participants at the Bank of Canada seminar series, the Canadian Economic Association Meetings, the European Central Bank seminar series, the Society for Computational Economics Meetings and the European Economic Association Meetings, and particularly Bob Anderton, Hafedh Bouakez, René Lalonde, James Powell, Eric Santor, and Larry Schembri. We would also like to thank Aileen Zhang for her excellent research assistance. Part of this research was conducted while Eiji Fujii was a visiting scholar at the Bank of Canada. He gratefully acknowledges the Bank's support and hospitality. Any remaining errors or omissions are ours. The views in this paper are those of the authors and should not be attributed to the Bank of Canada.
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Abstract

This paper investigates the question of whether a transition to a low-inflation environment, induced by a shift in monetary policy, results in a decline in the degree of pass-through of exchange rate movements to consumer prices. It differs from previous empirical work in its focus on the identification of changes in the inflation environment and its use of a panel-data approach. Evidence from a panel-data set of 11 industrialized countries over the period from 1977 to 2001 supports the hypothesis that exchange rate pass-through declines with a shift to a low-inflation environment brought about by a change in the monetary policy regime. More specifically, our empirical results suggest that pass-through to import, producer, and consumer price inflation declined following the inflation stabilization that occurred in many industrialized countries in the early 1990s. Interestingly, however, we find no evidence of a decline in pass-through following a similar episode of inflation reduction in the 1980s. Several potential explanations for this finding are discussed focusing on the important issue of the credibility of new monetary policy regimes.

JEL Classification: E31; E42; F31
Key Words: Inflation and prices, exchange rate pass-through, monetary policy

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1. Introduction

Starting in the early 1990s, many industrialized countries reduced their inflation rates and entered a period of relative price stability. Although several factors are thought to have contributed to this trend, it is generally agreed that a shift towards more credible monetary policy regimes played an important role.[1] In some countries, such as Australia, Canada, and the United Kingdom, the enhanced credibility was achieved through the adoption of an inflation-targeting framework for monetary policy. In others, such as the United States, monetary policy credibility was boosted through a sustained commitment to maintaining low inflation following a disinflation. Regardless of how monetary policy was improved, however, the outcome was similar across countries. It resulted in an environment in which inflation is lower and more stable.

This low-inflation period in industrialized countries has also coincided with several episodes in which countries have experienced large exchange rate depreciations, which, based on historical experience, had much smaller effects on consumer prices than anticipated.[2] This common experience has led to the belief, shared by central bankers in many industrialized countries, that the extent of exchange rate pass-through (ERPT) into consumer prices has declined.[3] Furthermore, the fact that this potential decline has coincided with a transition to a low-inflation environment has popularized the view that these two phenomena could be linked. Taylor (2000) was one of the first to formally articulate this view and put forth the hypothesis that the low-inflation environment in many industrialized countries, which was brought about by more credible monetary

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[1] Other possible factors include favorable shocks, structural change, and increased international competition.
[2] See for instance, Cunningham and Haldane (1999) for the experiences of the United Kingdom (1992), Sweden (1992), and Brazil (1999). Also, Laflèche (1996), Filion and Léonard (1997), and Kichian (2001) document the Canadian experience in the 1990s.
[3] See, for example, Bank of Canada (2000), p.9.

policies, has successfully reduced the degree of ERPT to domestic prices. He argued that ERPT is primarily a function of the persistence of exchange rate and price shocks, which tend to be reduced in an environment where inflation is low and monetary policy is more credible.

In this study, we empirically examine Taylor's hypothesis that the move to a low-inflation environment has reduced the rate of ERPT to consumer prices. Unlike the previous studies that focus on explaining cross-country variations in pass-through elasticity, we opt for a panel-data approach to address the issue of declining pass-through as a common phenomenon among industrialized.[4] Specifically, we estimate the average rate of ERPT, in both the short and long run, in 11 industrialized countries, by using a generalized method of moments (GMM) estimator for dynamic panel-data models. We then examine whether the rate of ERPT exhibits a significant decline as a shift to a low inflation environment occurs. To investigate the question thoroughly, we consider ERPT to consumer, producer, and import prices.

There are two main advantages to using a panel-data approach in the current context. First, the time-series dimension of the data enables us to investigate whether ERPT has declined as a result of a *change* in the inflation environment. Second, the cross-sectional dimension of the data allows us to draw on the experience of a group of countries that have undergone similar events at different time periods, and thus pool a larger number of observations to test our hypothesis.

In addition to our panel-data approach, our study also differs from previous empirical work in that we focus on the identification of changes in the inflation

[4] There have been a handful studies attempting to empirical evaluate Taylor's hypothesis. The majority of them are, however, cross-sectional in nature. Although informative, these studies cannot address the question of whether ERPT has declined in response to a change in the inflation environment.

environment that are the result of a change in the monetary policy regime and we allow for the possibility of multiple breaks. We do this by formally testing for structural breaks in the inflation series in our sample countries—using a test developed by Bai and Perron (1998) that allows for the identification of multiple breaks—and then ensuring that these breaks correspond with a change in the monetary policy regime. Allowing for the possibility of multiple shifts in the inflation environment seems appropriate in this context, given that many industrialized countries experienced two inflation stabilization periods in the post-Bretton Woods era, both of which were achieved by significant shifts in monetary policy.[5]

To preview our results, we do find evidence to support the hypothesis that ERPT declines with a shift to a low-inflation environment brought about by a change in the monetary policy regime, using annual data for 11 industrialized countries from 1977 to 2001. More specifically, our results suggest that pass-through to all three price indexes declined following the inflation stabilization period that occurred in many industrialized countries in the early 1990s, but not following a similar episode that occurred in the 1980s.

This paper is organized as follows. Section 2 provides an overview of the literature on ERPT and the inflation environment. Section 3 discusses our empirical methodology. Specifically, it presents the specification of our cross-country pass-through equation, describes the data, and discusses the estimation method. In section 4, we present preliminary estimation results while abstracting away from any effects of a change in the inflation environment. In section 5, we first identify shifts in the inflation

[5] Indeed, prior to this most recent episode, inflation was reduced significantly in the early 1980s in many industrialized countries, after rising to double-digit figures as a result of the oil-price shocks of the 1970s and the accommodative policy response to these shocks.

environment that occur as a result of a change in the monetary policy regime. We then investigate whether the estimated rates of pass-through vary significantly in the different inflation environments identified. Section 6 concludes.

2. Overview of the Literature

Although the degree to which exchange rate movements are reflected in prices has been of interest in international economics for a long time, the question of whether pass-through can be influenced by macroeconomic factors, such as monetary policy, is a more recent occurrence.[6] Namely, new developments in the open-economy macroeconomics literature (Obstfeld and Rogoff 1995; Betts and Devereux 1996, 2000) have brought an explicit macroeconomic perspective on ERPT by examining the extent to which it can depend on a country's inflation performance or monetary policy.

Several recent papers theoretically highlight the link between pass-through and monetary policy in dynamic general equilibrium models. Emphasizing channels such as a decline in the expected persistence of cost and price changes (Choudhri and Hakura 2001; Murchison 2004), a fall in the frequency of price changes (Devereux and Yetman 2002), or an increase in the prevalence of local currency pricing (Devereux, Engel, and Storgaard 2003), these studies show that a transition to a low-inflation environment—that

[6] Traditional open-economy macroeconomic models paid little attention to pass-through, given that typically in such models markets are characterized by perfect competition, prices are assumed to be fully flexible, and purchasing-power parity (PPP) holds at all times, implying that ERPT is complete and immediate. With such a stylized depiction of the pass-through process being a feature of many mainstream open-economy macroeconomic models, it is not surprising that, until very recently, most of the research in this area was more microeconomic in nature. The bulk of the pass-through literature approached the question from an industrial organization perspective and emphasized how pass-through could be incomplete in an environment characterized by imperfect competition and pricing to market (Kreinin 1977; Hooper and Mann 1989; Marston 1990; Krugman 1987). Goldberg and Knetter (1997) provide a comprehensive review of this literature.

comes about as a result of a more credible/stable monetary policy—can lead to a lower degree of ERPT.

However, empirical evidence on the relationship between ERPT and the inflation environment has yet to be firmly established. Among a handful preceding attempts, Choudhri and Hakura (2001) and Devereux and Yetman (2002) investigate the role of inflation variables in accounting for cross-country differences in ERPT in a large sample of countries. Their approach involves estimating a first-stage regression for each country in their sample to obtain an estimate of the average pass-through elasticity over a certain time period (usually 25 or 30 years). Then a second-stage specification is estimated where these country-specific average pass-through elasticities are regressed on various explanatory variables, such as inflation performance, exchange rate variability, and openness to trade. Thus, these second-stage regressions do not have a time-series component and focus exclusively on explaining cross-country variations in pass-through. Using this approach on a large sample of countries (both industrialized and developing) over the post-Bretton Woods period, both studies find that cross-country differences in estimated ERPT coefficients can be explained partly by differences in inflation performance.[7]

Although these studies are informative because they shed light on what might explain cross-country variations in pass-through elasticities, they cannot address the question of whether ERPT has declined in response to a change in the inflation environment. A purely cross-sectional analysis cannot tackle this question, given that it

[7] Campa and Goldberg (2002) use a similar approach in their study of pass-through in Organisation of Economic Co-operation and Development (OECD) countries, although they focus on import rather than on consumer prices. Although they also find a positive association between inflation and ERPT, they conclude that microeconomic factors related to the composition of imports are relatively more important in explaining cross-country differences in pass-through to import prices.

uses country-specific measures of pass-through that are averaged over the sample period, and are hence held constant.

Gagnon and Ihrig (2002) address this issue in their study of the link between consumer prices and monetary policy in a sample of 20 industrialized countries over the period from 1971 to 2000. In addition to using a cross-sectional approach, they test whether pass-through declined in each country in the sample following a change in the inflation regime.[8] One regime change was identified for each country using a combination of casual inspection of the data and judgment.[9] For each country, pass-through equations were then estimated on two subsamples (i.e., pre- and post-regime change). In most cases, the pass-through coefficients were smaller in the second subsample, which the authors interpret as evidence that ERPT has declined in the countries, and that this decline is attributable to the change in the inflation regime.

Our study differs from existing work in two important ways. First, our panel-data approach enables us to exploit both the time- and cross-sectional dimensions of the data. Consequently, we are able to investigate whether ERPT has declined as a result of a *change* in the inflation environment as a common experience shared by a group of industrialized countries. Second, in identifying changes in the inflation environment, we conduct the formal statistical test of multiple structural breaks, and supplement the results with a careful review of institutional facts on the monetary policy regime changes.

3. Empirical Methodology and Data Description

[8] In their cross-sectional analysis, Gagnon and Ihrig (2002) also find a systematic relationship between estimated rates of pass-through and inflation. They also examine the link between the pass-through coefficients and parameters estimated from Taylor-type monetary policy rules, but fail to find a robust relationship.

[9] The identified break dates in the sample countries cover a period starting in 1981 (for countries such as Japan, the United Kingdom, and the United States) and ending in 1993-94 (for Australia and Sweden).

3.1 Econometric specification

As a preliminary step to testing for a decline in ERPT, we first derive our benchmark specification that assumes that the rate of pass-through is stable over the sample period regardless of the inflation environment. This is a useful starting point which allows us to check that our cross-country pass-through estimates are reasonable and comparable to others in the literature. Later in section 5.2, we present a modified specification that allows changes in the inflation environment to influence the rate of ERPT.

Our approach is to use the standard reduced form specification adopted in the pass-through literature while adapting it so that it is suitable to estimate pass-through at the aggregate level for all three price indexes considered and to incorporate the influence of the inflation environment on ERPT. We begin with a simple profit maximization problem by a foreign firm that exports its product to the domestic country as commonly seen in the literature.[10] The exporting firm solves the following profit-maximization problem:

$$\max_{p} \quad \pi = s^{-1} pq - C(q) \qquad (1)$$

where π denotes profits (expressed in the foreign currency), s is the exchange rate measured in units of the domestic currency per unit of the foreign currency, p is the price of the good (denominated in the domestic currency), C(.) is the cost function (in foreign currency units) and q is the quantity demanded for the good.

Solving equation (1) yields the following first-order condition:

$$p = s C_q \mu \qquad (2)$$

[10] See, for instance, Dornbusch (1987), Knetter (1989), and Marston (1990).

where C_q is the marginal cost and μ is the markup of price over marginal cost. The markup is further defined as $\mu \equiv \eta/(\eta - 1)$ where η is the price elasticity of demand for the good.[11] The expression for the price level in equation (2) emphasizes that the local currency price of the good can vary as a result of a change in the exchange rate, a change in the firm's marginal cost, and/or a change in the firm's markup. Note that the firm's marginal cost and markup may change independently of the exchange rate. For instance, a change in the cost of a locally provided input (in the foreign country) can shift the marginal cost. Also, demand shocks in the importing country can alter the exporter's markup. It is thus important to take into account movements in these other determinants of the price when estimating pass-through to properly isolate the effects of exchange rate changes on import prices.

Consequently, a simple log-linear, reduced-form equation may be expressed as follows:

$$p_t = \alpha + \lambda s_t + \tau w_t + \eta y_t + \varepsilon_t \tag{3}$$

where w_t and y_t are measures of the exporter's marginal cost and the importing country's demand conditions, respectively. The coefficient λ thus measures ERPT. As discussed in Goldberg and Knetter (1997), variants of (3) are widely used as empirical specifications in the pass-through literature.

In adapting this specification to be suitable for estimating ERPT at the aggregate level for all three price indexes, there are several issues that need to be considered. First, the aggregate price level and the exchange rate are often found to be best described as

[11] Note that the expression in equation (2) is applicable to a variety of market structures. In a perfectly competitive industry, η is infinite so that μ is always one. On the other hand, in a monopolistically competitive environment, the exporting firm may have some leverage to raise the price above marginal cost.

I(1) series. Consequently, it is common to use a specification with these variables in the first-difference when estimating an aggregate pass-through equation—thus one ends up estimating an inflation equation.

Second, the literature on inflation dynamics has emphasized the need to account for the observed inertial behavior of inflation. However, as pointed out by Galí and Gertler (1999), it has been difficult for theoretical models to capture this persistence in inflation without appealing either to some form of ad hoc stickiness in inflation or to adaptive expectations. We follow a common practice in the literature to include lags of inflation as explanatory variables.

Third, whereas equation (3) was developed for import prices, we want to use a specification that is also suitable for consumer and producer prices. This is accomplished by using the output gap as a proxy for changes in domestic demand conditions in the inflation equations for all three price indexes. Thus, the resulting pass-through equation used for CPI inflation has all the elements of a backward-looking Phillips curve.

Given the aforementioned discussion, the following equation depicts the benchmark specification for our cross-country ERPT regression model:

$$\Delta p_{i,t} = \alpha_i + \eta_t + \sum_{j=1}^{2} \phi_j \Delta p_{i,t-j} + \lambda \Delta s_{i,t} + \tau \Delta ulc\_row_{i,t} + \delta\, gap_{i,t} + \varepsilon_{i,t}. \qquad (4)$$

where $\Delta p_{i,t}$ is the rate of change in the relevant aggregate price index for country *i* in time period *t*, α_i is a country-specific effect, η_t is a time dummy, $\Delta s_{i,t}$ is the rate of change in the nominal effective exchange rate for country *i* and time period *t*[12], $\Delta ulc\_row_{i,t}$ and $gap_{i,t}$ are control variables that capture changes in foreign producer

[12] The exchange rate is defined in terms of local currency units per unit of the (composite) foreign currency. Therefore, this variable will take on a positive (negative) value in the case of a depreciation (appreciation).

cost and domestic demand conditions for country *i* and time period *t*, respectively, and $\varepsilon_{i,t}$ is an independent and identically distributed (i.i.d.) error term.[13]

The country-specific intercepts α_i are designed to capture any unobservable or missing characteristics that vary across the sample countries (but not over time) and that influence inflation rates. For example, they could capture cross-country differences in measurement error in the construction of the price indexes or in institutional preferences for low inflation (as long as the differences between countries are constant over time).[14] The time dummy, η_t, is intended to capture the effects of global shocks on inflation rates.

3.2 Data description

Our panel dataset consists of annual observations for the following eleven industrialized countries: Australia, Belgium, Canada, Denmark, Finland, France, Italy, Netherlands, Spain, United Kingdom, and United States. The sample period is 1977-2001 in general, but shorter in some cases due to data limitations.[15] We provide estimates of ERPT for three aggregate price indexes: the consumer price index (CPI), the producer price index (PPI), and the import price index (IPI). Although our primary interest is in studying the behavior of ERPT to consumer prices, we also investigate whether changes in the domestic inflation regime that may affect consumer prices are also evident in the

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[13] We checked the stationarity of the individual series in equation (4) using the augmented Dickey-Fuller (ADF) test. According to the ADF test results, these variables are appropriately described as stationary series. In addition, we also checked whether the variables in levels were cointegrated, using the Johansen test, and found no such evidence.

[14] The country-specific effects could be either fixed (i.e., a constant that varies for each cross-sectional unit) or random (i.e., a random variable drawn from a common distribution). Although the country-specific effects are probably best modeled as fixed in the current case—given that our panel is more accurately described as a sample containing most of the countries of interest (i.e., industrialized countries) rather than a random sample from a larger group of countries—the estimation technique that we use makes it possible to estimate the coefficients of interest without having to restrict the country-specific effects to being either fixed or random.

[15] The sample countries and periods are determined by data availability. Appendix A provides more details on the data.

behavior of pass-through to import and producer prices. In the three cases, we include two lags of the dependent variable as explanatory variables to account for price inertia.[16]

To capture movements in the costs of foreign producers that export to the domestic market, we construct a foreign exporters' unit labor cost (ULC) series for each country, using the domestic ULC series and the ULC-based real effective exchange rate series. More specifically, the ULC series for foreign producers is defined as $ulc\_row_t \equiv q_t^{ULC} - s_t + ulc\_dom_t$ (where q_t^{ULC} is the ULC-based real effective exchange rate, s_t is the nominal effective exchange rate, and ulc_dom_t is the ULC of the domestic country). Since the nominal and real effective exchange rate series are trade weighted, Δulc_row_t effectively gauges the rate of change in the ULC of the exporters to the domestic country. As discussed earlier, we use the output gap as a proxy for changes in domestic demand conditions.

3.3 Estimation methodology

To estimate equation (4), we need to use a technique that is suitable for dynamic panel-data models. Complications do arise in estimating such models, stemming from the fact that the lagged dependent variable is correlated with the disturbance term. This renders estimation of both the random- and fixed-effects models using standard techniques problematic.

In the case of the random-effects formulation, it is well-known that the standard estimator will be biased if any of the explanatory variables are correlated with the unobservable individual effects. This is necessarily so with a dynamic panel-data model given that, by construction, the lagged dependent variable will be correlated with the

[16] Two lags were necessary to remove autocorrelation in the residuals.

random effects (and hence with the compound disturbance term). And in the case of a dynamic panel data model with fixed effects, it has been shown that the standard estimator generates estimates that are biased when the time dimension of the panel is small (see, for instance, Nickell (1981)).[17]

As is common in the literature, we address this issue by using a dynamic GMM panel-data estimation developed by Arellano and Bond (1991) based on work by Anderson and Hsiao (1981) and Holtz-Eakin, Newey, and Rosen (1988). This approach involves combining all available lagged values of the dependent variable with current and lagged values of the differences of the exogenous variables into an instrument matrix. The GMM estimator then makes use of the moment conditions that these instruments will be orthogonal to the disturbance term. The methodology also relies on the assumption that there is no second-order correlation in the first-differenced errors. Using this instrument matrix, Arellano and Bond (1991) derive a GMM estimator as well as two specification tests for this estimator that can be used to test the validity of the instruments: a test of second-order autocorrelation in the first-differenced residuals (the m_2 test for autocorrelation) and a Sargan test of over-identifying restrictions.

Arellano and Bond's dynamic panel-data GMM estimator is also appealing because it can accommodate a situation where one or more of the explanatory variables are assumed to be endogenous rather than exogenous.[18] This is useful in the context of estimating equation (4), given that the exchange rate term could be considered an

[17] And as Judson and Owen (1999) have shown using a Monte Carlo approach, this bias can be sizeable even when the number of observations per cross-sectional unit (T) reaches 20 and 30. Therefore, given that our panel-data set has T=25, estimating equation (4) using the standard fixed-effects model would yield biased estimates.

[18] In this case, the endogenous variables are treated similarly to the lagged dependent variable in that lagged levels of the endogenous variables are used as instruments for their respective first-differences.

endogenous variable. Indeed, if one believes that PPP holds and thus that relative price levels drive the exchange rate, then there could be a two-way causality between the rate of change of the exchange rate and inflation in our pass-through equation. Given the difficulties in modelling exchange rates, however, it is unclear whether this is the case in practice in our sample countries. Thus, we prefer to take an agnostic approach and consider both cases (i.e., we treat the exchange rate as both exogenous and endogenous).

Despite the fact that the GMM estimator does appear to be the most appropriate choice in this context, it is important to note that there may be an important drawback to using it in practice to estimate a dynamic panel-data model such as equation (4). Indeed, as with any instrumental-variable approach, the dynamic panel-data GMM estimator will suffer from large finite-sample biases if the instruments are weak. Thus, if the lagged values of the endogenous variables are only weak instruments for subsequent first differences, the GMM estimator could be poorly behaved. We acknowledge this potential drawback and address it by reporting estimation results for the pooled OLS and fixed-effects estimators, as well, as a check on the reliability of the GMM estimates. In addition, we report estimation results for two versions of the GMM estimator: "GMM1" instruments only for the lagged dependent variable (and hence assumes that the exchange rate is exogenous), whereas "GMM2" instruments for both the lagged dependent variable and the rate of change in the exchange rate.

4. Panel Pass-Through Estimates Assuming No Effects from Changes in the Inflation Environment

As a useful starting point, we begin our empirical exploration by examining the estimation results for equation (4) which abstracts away from any effects that changes in the inflation environment may have on the average rate of aggregate pass-through in our sample countries.

Table B.1 reports the pass-through estimates, in both the short and long run,[19] for each of the three price indexes that we consider: IPI, PPI, and CPI. The results were obtained by using two versions of the dynamic GMM panel-data estimator, GMM1 and GMM2, described above. As discussed, we also present estimation results using the pooled OLS and fixed-effect estimators as a check on the reliability of the GMM estimations. As shown in Table B.1, all of the pass-through estimates are statistically significant and of the expected (positive) sign, and the results are fairly robust across estimation techniques. The results are also robust to the presence of outliers. Tables D.1 through D.3 report more complete estimation results for equation (4).[20]

As expected, the size of the pass-through estimates varies substantially across price indexes. Indeed, the point estimates for import prices are much larger, reflecting the fact that the import price index is driven entirely by prices of tradable goods, whereas producer and consumer price indexes are driven by a combination of domestically produced and imported goods.[21] Therefore, the extent of pass-through to producer/consumer prices will depend on the rate of pass-through to import prices, the share of imports in the producer/consumer price indexes, and the response of domestically produced goods to movements in the exchange rate.

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[19] Our measure of long-run pass-through is defined by $\lambda/(1-[\sum_{j=1}^{2}\phi_j]))$. It is intended to capture the feedback effects resulting from the inclusion of the lagged dependent variable terms (i.e., the effects of an exchange rate change in period t will influence inflation over several periods subsequent to this as a result of these feedback effects).
[20] Tables D.1 through D.3 also report the results of the two specification tests used to check the validity of the instruments for the GMM estimations (i.e., the Sargan test and the m_2 test for autocorrelation). As shown, the results of these tests suggest that the instruments are indeed orthogonal.
[21] And in the case of the CPI, the index also includes the price of services (the majority of which are non-tradeable).

Two other reasons may also explain why the rate of pass-through to consumer prices is relatively smaller than that to import prices. First, local distribution costs—such as transportation costs, marketing, and services—can drive a wedge between import prices and consumer prices, and this wedge will fluctuate if distributors adjust their profit margins in response to movements in the exchange rate. Pass-through to consumer prices becomes smaller if distributors decide to compress their profit margins to offset (either partially or fully) the increase in the price of the good in the local currency. Second, as discussed in Bacchetta and van Wincoop (2002), differences in the optimal pricing strategies of foreign wholesalers and domestic retailers can also make the rate of consumer price pass-through less than proportional to that of import price pass-through. Specifically, this can occur if foreign exporting firms price their goods in the exporter's currency, while domestic retailers resell these goods priced in domestic currency.

Our estimation results indicate that pass-through to import prices in industrialized countries is high in the short run and complete (or near complete) in the long run. Indeed, the point estimates of 0.7493 and 0.9131 reported for short- and long-run pass-through (for GMM2), respectively, suggest that a 1 per cent increase in the annual rate of depreciation of the trade-weighted nominal exchange rate in industrialized countries leads to, on average, a 0.75 per cent increase in the annual rate of inflation of import prices in that same year, and a 0.91 per cent increase in the long run.[22] These results are in line

[22] As discussed in Bailliu and Bouakez (2004), the import price series for Canada—one of our sample countries—suffer from measurement error in that a number of Canadian import prices are constructed by multiplying the foreign-currency price by the nominal exchange rate. Given this, we checked whether our estimates of pass-through to import prices were upwardly biased by the inclusion of Canada in our sample. We conducted robustness checks by excluding Canada from the sample and found the effects on the estimation results to be negligible.

with estimates in the literature of exchange rate pass-through into import prices for industrialized countries.[23]

Our estimation results for producer and consumer prices are also consistent with the literature. For GMM2, the point estimates suggest that the short-run pass-through rates are 8 per cent for consumer prices and 20 per cent for producer prices (increasing to 16 per cent and 30 per cent, respectively, in the long run). Our estimate for long-run pass-through to consumer prices is comparable to that obtained by Gagnon and Ihrig (2002).[24]

5. Exchange Rate Pass-Through in Different Inflation Environments

5.1 Identifying changes in the inflation environment

A formal investigation of Taylor's hypothesis requires a comparison of pass-through estimates under alternative inflation environments, where the shift results from a change in monetary policy. To identify such shifts in the inflation environment, we use a two-step approach. First, we use the multiple break test developed by Bai and Perron (1998) to test for the presence of structural breaks in the inflation series in each country in the sample and, if breaks are identified, to determine the timing of these shift(s). This first step thus finds changes in the inflation environment that are significant enough to appear in the data. Second, we check whether these identified breaks line up with changes in the monetary policy regime. This second step ensures that the changes in the inflation

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[23] For instance, Campa and Goldberg (2002) find that the average rate of pass-through into import prices across their sample of 25 OECD countries over the period 1975-1999 is 0.61 in the short run and 0.77 in the long run. Anderton (2003) finds a pass-through rate of between 0.5 and 0.7 for extra-euro area imports. While we do find a higher degree of exchange rate pass-through in the short run, this can be explained by the fact that the short run refers to one year in our analysis unlike one quarter in the other studies.

[24] Gagnon and Ihrig (2002) find an average long-run pass-through rate of 23 per cent in their sample of 20 industrialized countries.

environment identified in the data are indeed the result of a change in the monetary policy regime and not the result of other factors.

The results of the Bai and Perron test, as well as a description of this methodology, are provided in Appendix C.[25] In addition, graphs depicting the CPI inflation series for each country are shown in Figure C.1 in Appendix C, along with vertical lines representing the dates at which the structural breaks were identified. As shown in these figures, we found evidence of at least one break in all countries, and for most of the countries, two breaks were identified. In the majority of cases, the first break coincides with the Volker-era disinflation in the early 1980s and the second one with the more recent inflation stabilization period that began in the early 1990s.

As outlined in Appendix E, most of these identified structural breaks in the inflation series line up with a change in the monetary policy regime in the country in question. The one exception is for Spain, where the Bai-Perron test identified three structural breaks in the 1990s. However, only one of these break dates coincides with a change in the monetary policy regime and therefore we consider only this shift in our analysis.

All of the sample countries experienced a substantial increase in inflation in the 1970s as a result of the oil shocks and the accommodative policy response to these shocks. Consequently, they all took steps to reduce inflation in the first half of the 1980s. Although they took different approaches with varying degrees of success, all brought about a disinflation by making significant changes to their monetary policy regime. In most of the European countries—with the notable exception of the United Kingdom—

[25] The structural breaks tests are conducted on the CPI series, since this is where we would expect a change in the domestic inflation environment to manifest itself. In order to have enough degrees of freedom to conduct the test, we use quarterly rather than annual data.

most of the policy changes focused on using the exchange rate as a nominal anchor, and essentially importing German monetary policy. The most extreme example of this type of regime shift is the Netherlands, where a "hard peg" was adopted versus the Deutsche Mark (DM) in the early 1980s. From that point until the launch of the euro, the Netherlands held the peg. Belgium adopted a similar approach, although the shift was more gradual. These two countries entered a low-inflation environment in the 1980s and stayed there.

The other countries in the sample entered a low-inflation environment in two steps, the first in the 1980s and the second in the 1990s. Whereas the focus of policy changes in the 1980s seemed to be on reducing inflation from the high levels it reached in the 1970s, the emphasis in the 1990s appeared to be on inflation control (i.e., achieving and maintaining low and stable rates of inflation). Many countries—such as Australia, Canada, Spain, and the United Kingdom—adopted inflation-targeting regimes to achieve this goal. Other countries, like the United States, relied on a flexible version of inflation control. And finally, most of the European countries continued to rely on exchange rate pegs as a means of importing the low inflation rate of the core country of the Exchange Rate Mechanism (ERM), Germany.

5.2 Panel pass-through estimates in different inflation environments

Based on the identified structural breaks, we modify our cross-country pass-through equation to allow the rate of pass-through to vary with the inflation environment. Specifically, we construct two dummy variables: $regime\_80_{i,t}$ and $regime\_90_{i,t}$, which capture a shift in the inflation environment in the 1980s and 1990s, respectively. In each case, the dummy variable takes on the value one starting in the period in which the country experienced a structural break (and for all subsequent years), and zero otherwise.[26] We then interact these two dummy variables with the exchange rate term, and include the two interaction terms in the specification, as follows:

$$\Delta p_{i,t} = \alpha_i + \eta_t + \sum_{j=1}^{2} \phi_j \Delta p_{i,t-j} + \lambda \Delta s_{i,t} + \lambda_{reg\_80} (\Delta s_{i,t} * regime\_80_{i,t}) \\ + \lambda_{reg\_90} (\Delta s_{i,t} * regime\_90_{i,t}) + \tau \Delta ulc\_row_{i,t} + \delta\ gap_{i,t} + \varepsilon_{i,t}. \qquad (5)$$

There are thus three coefficients of interest in equation (5): the coefficient on the rate of change in the exchange rate (i.e., λ) and the two coefficients on the interaction terms described above (i.e., λ_{reg_80} and λ_{reg_90}). The former captures the average rate of short-run exchange rate pass-through in our sample countries (for each relevant price index), whereas the latter two capture any incremental effects due to a change in the inflation environment that starts either in the 1980s or the 1990s.[27]

Estimation results for equation (5) are reported in Table B.2, focusing on the three coefficients of interest. As shown in the table, the coefficients on the interaction terms are statistically significant in the case of all three price indexes but, importantly, only for the interaction term with the dummy variable that identifies a change in the inflation environment in the 1990s. This suggests that the effects of exchange rate movements on import, producer, and consumer prices were dampened after the shift in the inflation environment that occurred in our sample countries in the 1990s—for those countries that did experience such a shift. For instance, in the case of import prices (for GMM2), the average short-run pass-through rate is equal to roughly 86 per cent prior to the shift, and is reduced to around 71 per cent following a change in the inflation environment in the

[26] If the country did not experience such a break, then the dummy variable in question would take on the value zero for all periods for this country.
[27] Given our data frequency, the "short-run" here refers to a one-year period.

1990s. As for producer prices, the average short-run pass-through rate is equal to roughly 18 per cent prior to the shift, and is reduced to around 8 per cent following a change in the inflation environment in the 1990s. Finally, in the case of consumer prices, the change in ERPT from one environment to the next is fairly dramatic. Indeed, the short-run pass-through rate falls from around 11 per cent before the shift to around 5 per cent in the low-inflation environment that started in the 1990s.

There are a few potential explanations for why pass-through might have declined in the 1990s but not in the 1980s. First, it is possible that the changes in the monetary policy regimes that were implemented in the 1990s were perceived as being more credible than those that were carried out in the 1980s. This could be due to the fact that many of the sample countries made more substantial reforms to their monetary policy regimes in the latter decade (for example, by adopting inflation targeting). Moreover, as discussed in Paulin (2000), the 1990s were characterized by a period in which central banks acquired greater operational independence to pursue their policy objectives and became more open institutions. This trend most likely contributed to increasing both the effectiveness and credibility of policy actions.

Another potential explanation is that credibility is not built overnight, but takes time to acquire. Thus, it is possible that although inflation fell as a result of monetary policy changes implemented in the 1980s, credibility was not enhanced until the 1990s, as a result of the cumulative impact of a range of initiatives and/or time needed for agents in the economy to be convinced of the credibility of the new regime.

It is also possible that firms only revise their pricing strategy significantly when inflation falls below a certain threshold (i.e., a level at which inflation is considered to be "low enough" to be consistent with relative price stability). As a result, variations in ERPT in response to a change in the inflation environment would occur in a discrete, rather than a continuous, fashion. For instance, consider the case CPI inflation in our sample countries. According to our identified inflation breaks, the average annual inflation rate in our sample countries was around 5 percent in the inflation regime that followed the first inflation stabilization period (i.e., the one that occurred in the 1980s). Although this was a substantial improvement over the 1970s—the corresponding figure for the pre-break period being roughly 10 per cent—it is not considered low by current (or historical) standards nor would it be considered to be consistent with relative price stability. On the other hand, in the low-inflation regime that started in the 1990s, the average inflation rate was further reduced to around 2 percent—a level of inflation consistent with price stability and more likely to be at (or below) this threshold level that induces firms to change their pricing behavior.[28]

6. Conclusion

The rate at which exchange rate fluctuations translate into price movements is an important policy issue, particularly for open economies. While the rate of pass-through is essentially determined by the pricing behaviour of firms, the macroeconomic environment in which they make pricing decisions can shift significantly over time. In particular, the inflation environment in many industrialized countries has improved substantially over the few past decades with the establishment of more effective and credible monetary policy regimes. Such changes in the inflation environment may well

[28] Inflation variability also declined substantially in the low-inflation regime that started in the 1990s.

influence the pricing behavior of firms and hence, the rate of ERPT. More specifically, it has been conjectured that an environment of low and stable inflation established by shifts in the monetary policy regime has significantly reduced the rate of ERPT in many industrialized countries.

In this study, we investigated whether or not the rate of ERPT has indeed declined as a result of a shift in monetary policy using a panel-data set of 11 industrialized countries over the period from 1977 to 2001. A few key findings are worth highlighting. First, we find that almost all the countries we examined have experienced more than one break in the inflation environment during the sample period. More specifically, most countries experienced one break in the 1980s and another in the 1990s as they stabilized inflation in two stages. Furthermore, these break points generally correspond with significant policy changes that suggest a shift in the respective country's monetary policy regime.

Second and more importantly, when the rate of ERPT is allowed to vary with the inflation environment, our panel estimates provide us with significant evidence to support the hypothesis that ERPT declines with a shift to a low-inflation environment induced by a change in the monetary policy regime. There is, however, an important caveat to the pattern of decline in ERPT. Our results indicate that the rate of ERPT declined following the inflation stabilization that occurred in many industrialized countries in the early 1990s, but not following a similar episode in the 1980s. This pattern of decline in pass-through is found consistently across three different prices—consumer, producer, and import prices.

This finding related to the pattern of exchange rate pass-through in the past two decades raises an important question. Why is it the case that the inflation stabilization in the 1990s resulted in a significant decline in the rate of ERPT while that in the 1980s had no similar effect? We conjecture that this divergence is most likely due to perceived differences in the credibility of the monetary policy regimes in the two decades. This important issue, however, warrants further investigation. We leave this to future research.

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Appendix A: Data Appendix

Sample countries

Australia, Belgium, Canada, Denmark, Finland, France, Italy, Netherlands, Spain, United Kingdom, United States

Sources and Definitions of Variables

Dependant variable

1. Rate of change in the relevant annual aggregate price index[1]
   (Source: Organisation of Economic Co-operation and Development's (OECD) *Main Economic Indicators* and *Monthly Statistics of International Trade)*
   - calculated as the log difference in the level of the average annual price index

Explanatory variables

2. Rate of change in the annual nominal effective exchange rate
   (Source: Bank for International Settlements (BIS))
   - calculated as the log difference in the level of the average annual exchange rate

3. Rate of change in the foreign exporters' unit labor cost (ULC)
   (Source: BIS)
   - calculated as the log difference in the level of the average annual foreign ULC using the domestic ULC and the ULC-based real effective exchange rate
4. Output gap
   (Source: OECD's *Main Economic Outlook*)
   - measured as the deviation of actual output from potential as a percentage of potential output
   - potential output is calculated using the "production function method" (see OECD's *Economic Outlook Sources and Methods* for more details)

[1] We use the following three aggregate price indexes: (1) consumer price index; (2) producer price index; and the (3) import price index.

Table A1: Sample Period by Country

Country	Estimations where dependent variable is rate of change in[1]:		
	Consumer Price Index	Producer Price Index	Import Price Index
Australia	1977-2001	1977-2001	1977-2001
Belgium	1981-2001	1981-2001	n.a.[2]
Canada	1981-2001	1981-2001	1981-2001
Denmark	1977-2001	1977-2001	1977-2001
Finland	1977-2001	1977-2001	1977-2001
France	1977-1998	1977-1998	1977-1998
Italy	1977-2001	1981-2001	1977-2001
Netherlands	1984-2001	1984-2001	1984-1996[3]
Spain	1980-2001	1980-2001	1980-2001
U.K.	1977-2001	1977-2001	1977-2001
U.S.A.	1977-2001	1977-2001	1982-2001

Notes:
1. All data are annual.
2. The import price index series for Belgium is only available starting in 1993.
3. The import price index series for the Netherlands was obtained from the BIS.

Appendix B: Estimates of Exchange Rate Pass-through

Table B.1: Panel Pass-through Estimates for 11-country sample over 1977-2001

Estimation technique	GMM1	GMM2	Pooled OLS	Fixed Effects
A. IPI Inflation				
Short-run	0.7598**	0.7493**	0.7482**	0.7476**
	(0.0624)	(0.0524)	(0.0383)	(0.0431)
Long-run	0.8503**	0.9131**	0.9093**	0.8963**
F test / Chi-square test	128.60	149.91	242.31	195.69
B. PPI Inflation				
Short-run	0.2112**	0.2023**	0.2137**	0.2052**
	(0.0287)	(0.0289)	(0.0246)	(0.0259)
Long-run	0.2764**	0.3012**	0.3700**	0.3318**
F test / Chi-square test	32.51	31.62	50.30	44.19
C. CPI Inflation				
Short-run	0.0650**	0.0804**	0.0806**	0.0809**
	(0.0238)	(0.0202)	(0.0156)	(0.0166)
Long-run	0.1259*	0.1600**	0.1705**	0.1634**
F test / Chi-square test	5.28	11.53	26.01	21.70

Notes:

1. This table reports estimates of exchange rate pass-through obtained from estimating equation (7). The short-run estimate is $\hat{\lambda}$ whereas the long-run estimate is $\hat{\lambda}/(1-(\hat{\phi}_1+\hat{\phi}_2))$. See Table D.1 in Appendix D for more complete estimation results for equation (7).
2. The figures in parentheses are robust standard errors.
3. The F and Chi-square test statistics reported for the long-run pass-through estimates correspond to the values of these respective test statistics for the hypothesis that $\hat{\lambda}/(1-(\hat{\phi}_1+\hat{\phi}_2))=0$.

4. Both GMM1 and GMM2 refer to estimations carried out using the Arellano-Bond one-step dynamic panel-data first-difference robust estimator. GMM1 instruments only for the lagged dependent variable whereas GMM2 instruments also for the rate of depreciation. . In both cases, a restricted version of the estimator is used in that the maximum number of available lagged values of the endogenous variables used as instruments is set to 3. See Section 3.2 for more details on these estimators.
5. "**", "*", and "#" indicate statistical significance at the 1, 5, and 10% levels, respectively.

Table B.2: Short-run Panel Pass-through Estimates for 11-country sample over 1977-2001
(using interaction terms to account for inflation regime shifts)

Estimation technique	GMM1	GMM2	Pooled OLS	Fixed Effects
A. IPI Inflation				
Δs_t	0.8628**	0.8615**	0.8402**	0.8585**
	(0.1617)	(0.1371)	(0.1370)	(0.1456)
$\Delta s_{t\,*}$regime_80	-0.0334	-0.0619	-0.0418	-0.0615
	(0.1328)	(0.1065)	(0.1407)	(0.1482)
$\Delta s_{t\,*}$regime_90	-0.1845**	-0.1481**	-0.1437*	-0.1465#
	(0.0573)	(0.0530)	(0.0705)	(0.0769)
B. PPI Inflation				
Δs_t	0.1486#	0.1844**	0.1819**	0.1969**
	(0.0821)	(0.0603)	(0.0563)	(0.0637)
$\Delta s_{t\,*}$regime_80	0.1246	0.0600	0.0775	0.0508
	(0.0955)	(0.0743)	(0.0667)	(0.0745)
$\Delta s_{t\,*}$regime_90	-0.1365*	-0.1037#	-0.1091*	-0.1084*
	(0.0618)	(0.0556)	(0.0473)	(0.0493)
C. CPI Inflation				
Δs_t	0.0907*	0.1085**	0.1154**	0.1114**
	(0.0443)	(0.0244)	(0.0405)	(0.0391)
$\Delta s_{t\,*}$regime_80	-0.0122	-0.0116	-0.0182	-0.0145
	(0.0574)	(0.0329)	(0.0443)	(0.0442)
$\Delta s_{t\,*}$regime_90	-0.0427	-0.0544#	-0.0565#	-0.0543#
	(0.0302)	(0.0295)	(0.0305)	(0.0317)

Notes:
1. This table reports the coefficient estimates of the three listed variables obtained from the estimation of equation (4).
2. See notes (2), (4), and (5) from Table B.1.

Appendix C: Changes in the Inflation Environment in the Sample Countries

Testing for Multiple Structural Breaks: Methodology

To test for the presence of breaks in the inflation series of our sample countries, we use the endogenously-determined multiple break test developed by Bai and Perron (1998). This methodology tests for the presence of breaks in a series when neither the number nor the timing of breaks is known *a priori*. More specifically, this approach allows us to test for the presence of m breaks in the mean inflation rate of each country at unknown times using the following model:

$$\Delta p_t = \mu_j + \eta_t \qquad t = T_{j-1}+1,\ldots,T_j \text{ and } j = 1,\ldots,m+1$$

where $\mu_j (j = 1,\ldots,m+1)$ is the regime-specific mean inflation rate, η_t is an error term, and $T_0 = 0$ and $T_{m+1} = T$. In essence, the testing procedure searches for m optimal breaks that achieve a global minimal of the total of the sum of squared residuals in each regime.

We test for breaks in the CPI series given that this is where we would expect a change in the domestic inflation environment to manifest itself. In conducting the test, we adopt two kinds of test statistics. The first statistic, denoted by $Sup_F(m)$, evaluates the null hypothesis of no structural break against the alternative of m structural breaks. The second statistic, $Sup_F(m+1|m)$, tests the null hypothesis of m breaks against the alternative of (m+1) breaks.

Table C.1: Results of Structural Break Tests on the CPI Inflation Series, by Country

	Australia	Belgium	Canada	Denmark	Spain	Finland	France	U.K.	Italy	Nether.	U.S.
A. $Sup_F(m)$											
(1)	80.09**	95.54**	56.04**	78.62**	80.25**	48.94**	86.43**	48.23**	91.73**	57.05**	16.59**
(2)	55.22**	51.63**	59.18**	72.54**	61.90**	53.78**	72.82**	39.26**	107.15**	31.22**	13.99**
(3)	30.15**	34.71**	40.50**	50.54**	78.94**	34.55**	175.99**	26.96**	77.23**	24.75**	14.71**
(4)	22.96**	31.75**	31.66**	37.13**	76.62**	24.61**	135.43**	20.09**	74.33**	18.70**	12.69**
B. $Sup_F(m+1\|m)$											
(2|1)	12.74*	2.84	39.68**	44.52**	31.35**	25.63**	34.01**	22.01**	46.01**	8.04	9.83*
(3|2)	1.60	3.15	0.82	3.30	36.50**	0.90	5.86	3.18	6.04	0.93	4.14
(4|3)	0.00	0.61	0.57	0.94	14.86**	0.09	8.025	0.63	6.04	0.81	1.48
(5|4)	0.00	0.00	0.00	0.00	0.57	0.00	5.86	0.00	5.80	5.53	0.00
C. Break dates											
	1982Q3	1985Q1	1982Q3	1982Q3	1983Q4	1984Q1	1985Q1	1982Q1	1983Q4	1982Q2	1981Q2
	1990Q3		1990Q4	1989Q3	1991Q4	1991Q1	1992Q1	1992Q1	1996Q1		1990Q4
					1995Q1						
					1996Q1						

Notes: 1. Panel A reports the Sup_F test statistics for the null hypothesis of no structural break against *m* (*m*=1,...,4) breaks.

2. Panel B presents the Sup_F statistics for the null hypothesis of *m* (*m*=1,...,4) structural breaks against (*m*+1) structural breaks.

3. Panel C provided suggested break dates based on the results of these two tests.

4. "**" and "*" indicate statistical significance at the 1 and 5 % levels, respectively.

Figure C.1: Identified Structural Breaks in the CPI Inflation Series, by Country

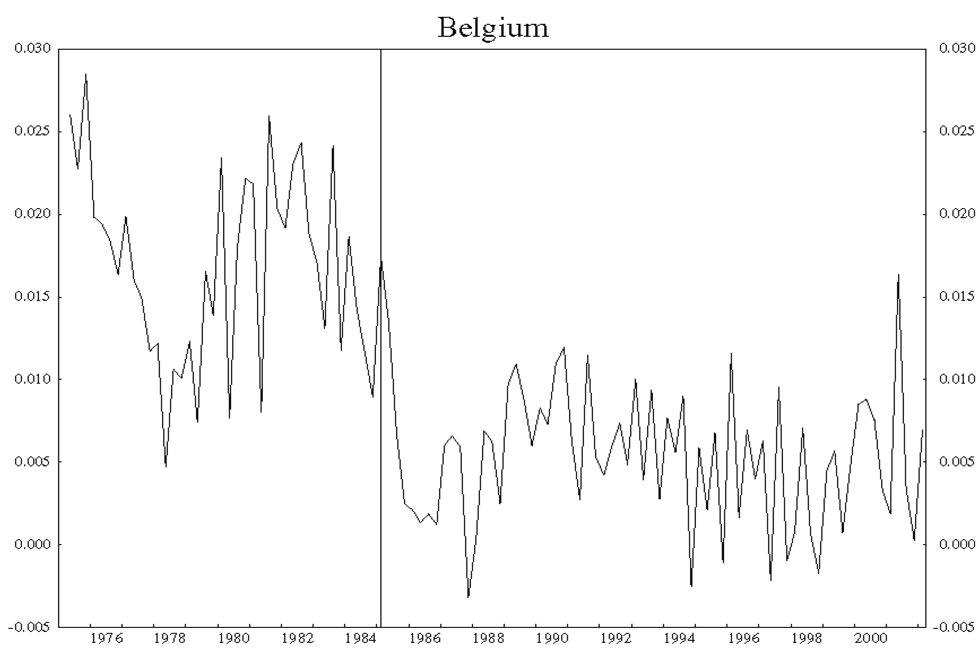
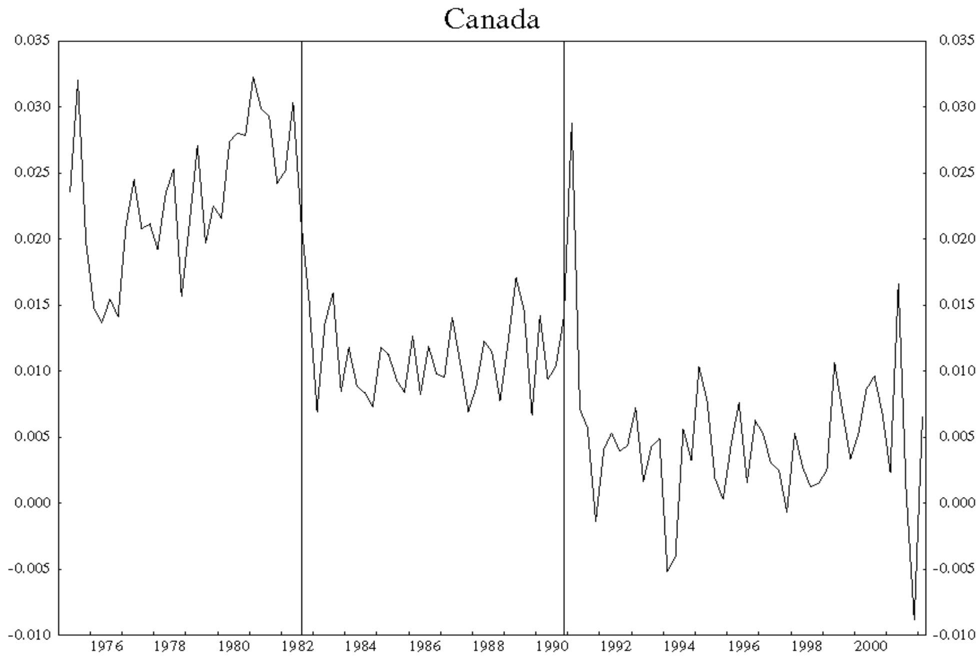


Figure C.1 (Cont'd)
Identified Structural Breaks in the CPI Inflation Series, by Country



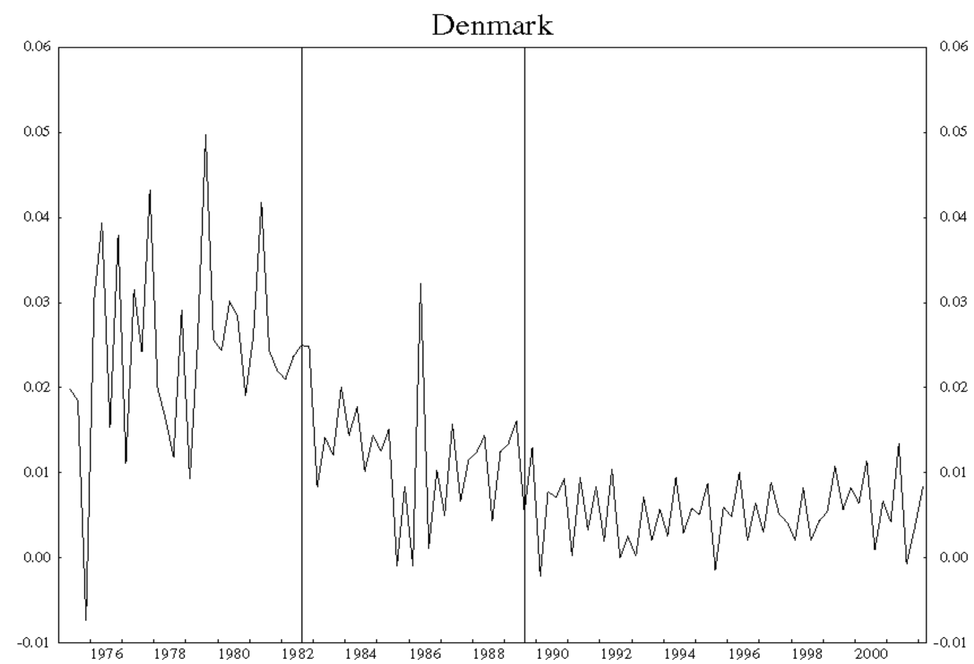


Figure C.1 (Cont'd)
Identified Structural Breaks in the CPI Inflation Series, by Country

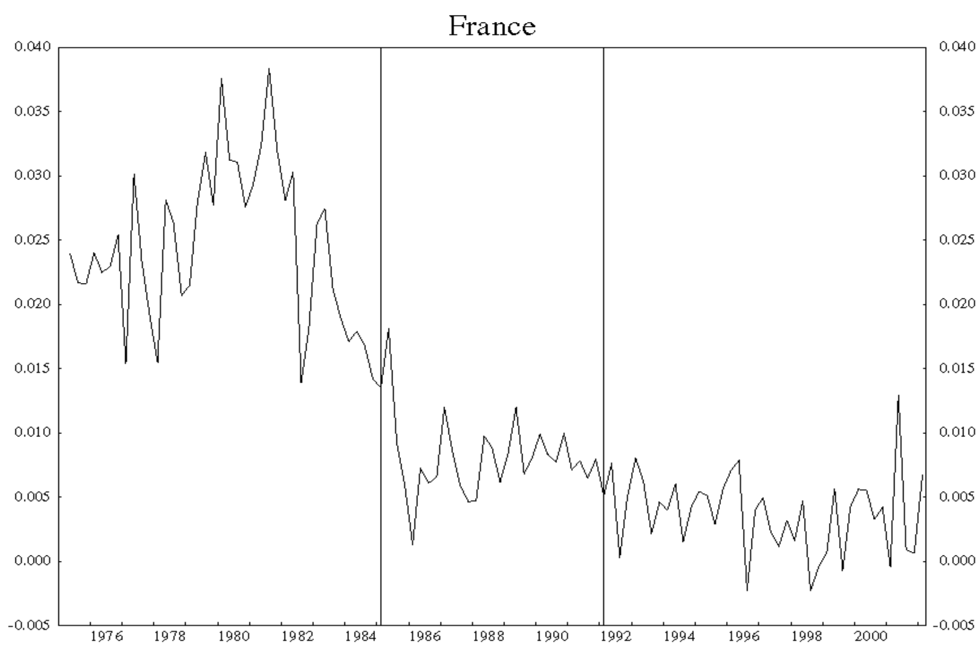
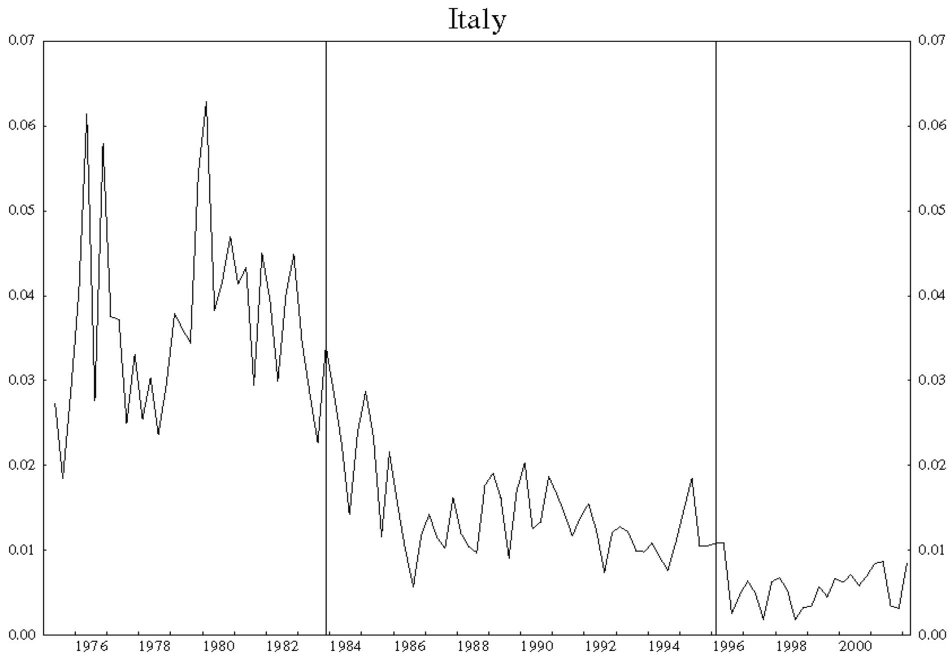


Figure C.1 (Cont'd)
Identified Structural Breaks in the CPI Inflation Series, by Country



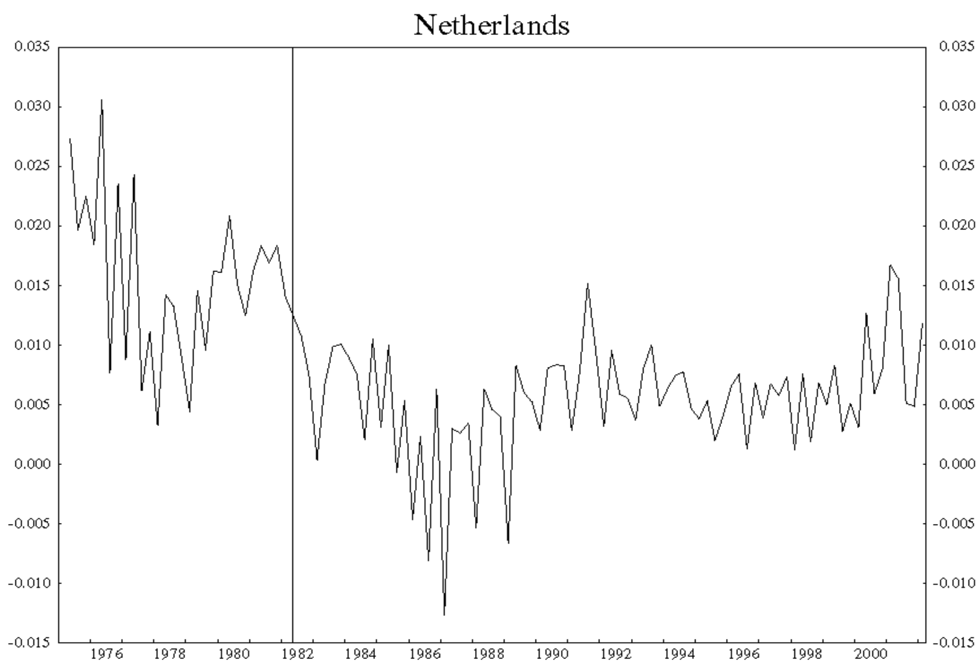


Figure C.1 (Cont'd)
Identified Structural Breaks in the CPI Inflation Series, by Country

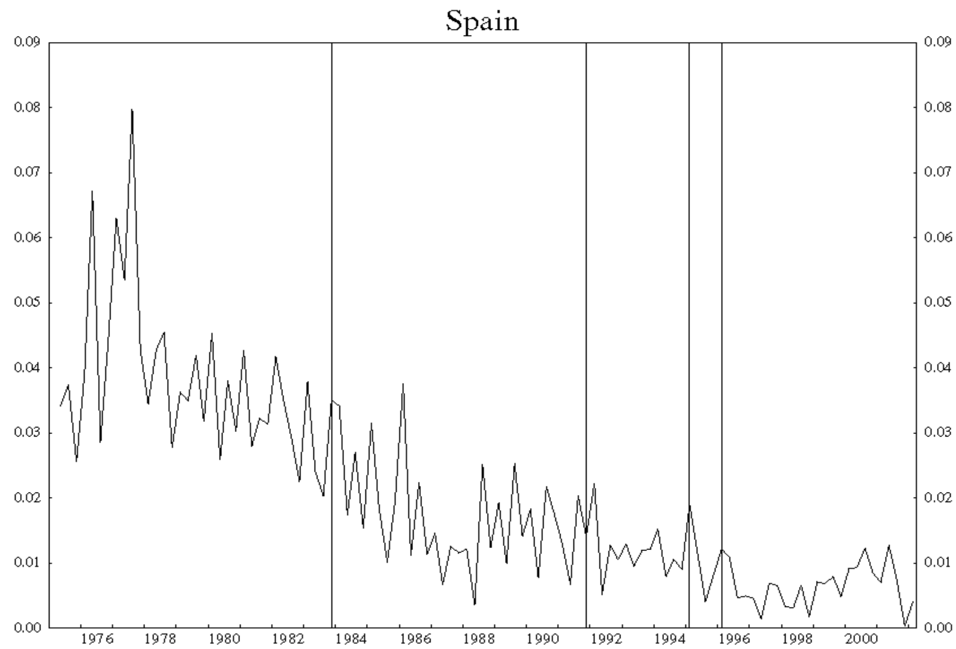
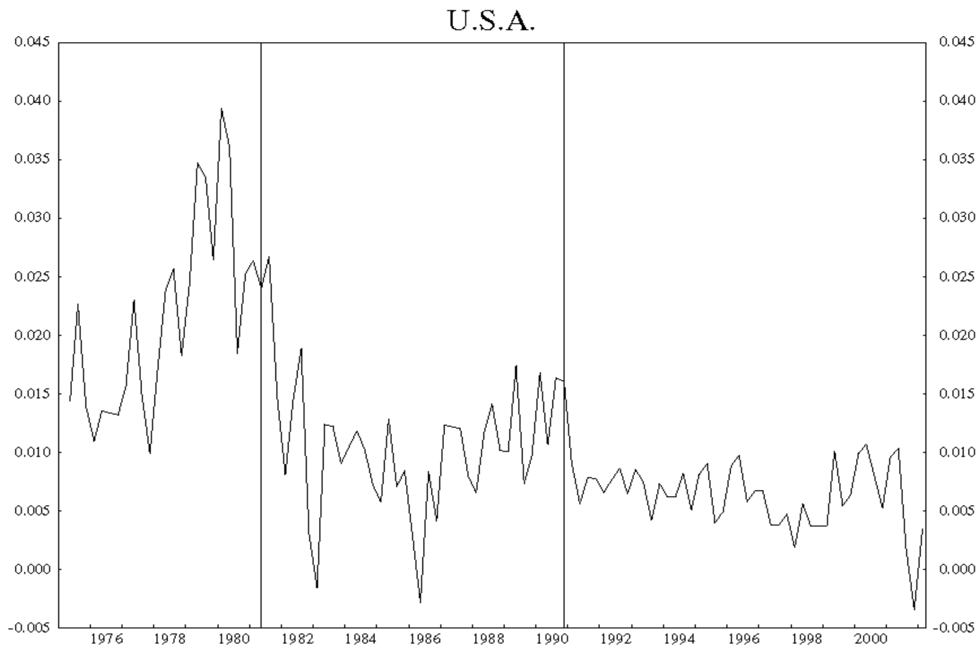


Figure C.1 (Cont'd)
Identified Structural Breaks in the CPI Inflation Series, by Country



Appendix D: Estimation Results for Equation (4)

Table D.1: Dependent Variable: Rate of Change in Annual Import Price Index
Panel Estimates for 11-country sample over 1977-2001

Estimation technique	GMM1	GMM2	Pooled OLS	Fixed Effects
$\Delta p_{i,t-1}$	0.1754**	0.2179**	0.2349**	0.2237**
	(0.0478)	(0.0559)	(0.0371)	(0.0378)
$\Delta p_{i,t-2}$	-0.0689**	-0.0385	-0.0577**	-0.0578**
	(0.0204)	(0.0253)	(0.0113)	(0.0109)
$\Delta s_{i,t}$	0.7598**	0.7493**	0.7482**	0.7476**
	(0.0624)	(0.0524)	(0.0383)	(0.0431)
$\Delta ulc\_row_{i,t}$	-0.0429	-0.0203	-0.0257	-0.0146
	(0.0448)	(0.0339)	(0.0392)	(0.0538)
$gap_{i,t}$	0.2509*	0.1547[#]	0.1432	0.1512
	(0.1241)	(0.0923)	(0.1183)	(0.1225)
No. of observations	207	207	218	218
Sargan test	90.37*	185.61		
m_2 test for autocorrelation	-0.54	-0.67		
Wald test for joint sign. of expl. variables	471.56**	370.76**		
Adjusted R^2			0.9195	0.9225
F test for joint sign. of fixed effects				1.01
Test for joint sign. of time effects				
F test			28.49**	25.49**
Wald test	546.22**	249.81**		

Notes:
1. This table reports estimation results for equation (7).
2. The figures in parentheses are robust standard errors.
3. Both GMM1 and GMM2 refer to estimations carried out using the Arellano-Bond one-step dynamic panel-data first-difference

robust estimator. GMM1 instruments only for the lagged dependent variable whereas GMM2 instruments also for the rate of depreciation. In both cases, a restricted version of the estimator is used in that the maximum number of available lagged values of the endogenous variables used as instruments is set to 3. See Section 3.2 for more details on these estimators.
4. "**", "*", and "#" indicate statistical significance at the 1, 5, and 10% levels, respectively.

Table D.2: Dependent Variable: Rate of Change in Annual Producer Price Index
Panel Estimates for 11-country sample over 1977-2001

Estimation technique	GMM1	GMM2	Pooled OLS	Fixed Effects
$\Delta p_{i,t-1}$	0.3434**	0.4206**	0.4538**	0.4138**
	(0.0713)	(0.0621)	(0.0486)	(0.0502)
$\Delta p_{i,t-2}$	-0.1075**	-0.0922*	-0.0312*	-0.0322*
	(0.0312)	(0.0402)	(0.0134)	(0.0128)
$\Delta s_{i,t}$	0.2112**	0.2023**	0.2137**	0.2052**
	(0.0287)	(0.0289)	(0.0246)	(0.0259)
$\Delta ulc\_row_{i,t}$	0.0066	0.0505	0.0759**	0.0413
	(0.0538)	(0.0471)	(0.0278)	(0.0360)
$gap_{i,t}$	0.1251	0.0374	0.0063	0.0513
	(0.0793)	(0.0705)	(0.0514)	(0.0606)
No. of observations	229	229	240	240
Sargan test	117.25**	191.83		
m_2 test for autocorrelation	-1.56	-1.46		
Wald test for joint sign. of expl. variables	204.78**	190.97**		
Adjusted R^2			0.8644	0.8701
F test for joint sign. of fixed effects				0.75
Test for joint sign. of time effects				
F test			19.12**	19.17**
Wald test	1189.62**	936.37**		

Notes: See notes for Table D.1.

Table D.3: Dependent Variable: Rate of Change in Annual Consumer Price Index
Panel Estimates for 11-country sample over 1978-2001

Estimation technique	GMM1	GMM2	Pooled OLS	Fixed Effects
$\Delta p_{i,t-1}$	0.5061**	0.5228**	0.5426**	0.5201**
	(0.0638)	(0.0468)	(0.0463)	(0.0506)
$\Delta p_{i,t-2}$	-0.0222	-0.0255*	-0.0154**	-0.0153**
	(0.0148)	(0.0102)	(0.0042)	(0.0044)
$\Delta s_{i,t}$	0.0650**	0.0804**	0.0806**	0.0809**
	(0.0238)	(0.0202)	(0.0156)	(0.0166)
$\Delta ulc\_row_{i,t}$	0.1155**	0.1337**	0.1300**	0.1346**
	(0.0317)	(0.0293)	(0.0191)	(0.0227)
$gap_{i,t}$	0.1370*	0.0887#	0.0737#	0.0890*
	(0.0628)	(0.0475)	(0.0375)	(0.0415)
No. of observations	237	237	248	248
Sargan test	78.96	192.13		
m_2 test for autocorrelation	-0.46	-0.25		
Wald test for joint sign. of expl. variables	229.15**	698.87**		
Adjusted R^2			0.9417	0.9445
F test for joint sign. of fixed effects				1.14
Test for joint sign. of time effects				
F test			8.97**	8.49**
Wald test	851.69**	228.62**		

Notes: See notes for Table D.1.

Appendix E: Identified Structural Breaks in Inflation and Corresponding Policy Change

Country	Date	Policy Change
Australia	1982Q3	Following a high-inflation period in the 1970s, there was a gradual shift in Australian monetary policy towards an approach that "…generally articulated a goal of disinflation" (Gruen and Stevens, 2000: 43).
Australia	1990Q3	Although the formal announcement of an inflation target occurred in September 1994 in Australia, the shift in monetary policy towards a strategy focused on inflation control started a few years earlier (see Bernanke et al. (1999) for more details).
Belgium	1985Q1	After Belgium joined the European Monetary System (EMS) in 1979, there were frequent downward realignments of the Belgian franc until the mid-1980s, when Belgium started pursuing "…a progressively tighter exchange rate policy" (Halikias, 1993:1).
Canada	1982Q3	As was the case in the United States, the Bank of Canada raised interest rates substantially in the early 1980s in order to reduce inflation. See Freedman (1982) for more details. This period also coincided with the end of the practice of targeting monetary aggregates at the Bank of Canada (officially cancelled in November 1982).
Canada	1990Q4	The adoption of inflation targeting in Canada was announced in February 1991.
Denmark	1982Q3	In 1982, a new government took office in Denmark and announced radical measures to tackle the economic crisis the country was experiencing, including a strong commitment to a fixed exchange rate policy. See Dahl and Hansen (2002) for more details.
Denmark	1989Q3	The exchange rate policy in effect since 1982 was further strengthened in the late 1980s when parities against the strongest currencies in the ERM were fixed. See Christensen and Topp (1997) for more details.
Spain	1983Q4	Starting in 1978, the Spanish central bank began to take an active role in monetary policy by publicly announcing monetary growth target rates as a means of reducing inflation. Around 1983-84, Spanish monetary policy started de-emphasizing the targeting of monetary aggregates in favor of a focus on the exchange rate (see Ayuso, Kaminsky and Lopez-Salido (2003) for more details).
Spain	1995Q1	The adoption of inflation targeting in Spain was announced in November 1994 (inflation targets were introduced in January 1995).
Finland	1984Q1	Starting in the early 1980s, the exchange rate was used as a nominal anchor in Finland in an attempt to eliminate the inflation-devaluation cycle that had afflicted the country for most of the post-war period (see Honkapohja and Koskela (1999) for more details).
Finland	1991Q1	Finland abandoned its "hard-currency" exchange rate policy in 1991 following the economic crisis brought on by the collapse of the Soviet Union (see Honkapohja and Koskela (1999) for more details).

France	1985Q1	Several important changes were made to French monetary policy over the period 1983 to 1987 including the adoption of a policy of competitive disinflation (which began in 1983 but took several years to complete) and a major reform of French financial markets which profoundly changed the operating procedures of French monetary policy. See Mojon (1999) for more details.
France	1992Q1	France successfully defended its peg during the ERM crisis in 1992, demonstrating its commitment to the ERM and the impending European Monetary Union.
United Kingdom	1982Q1	Starting in mid-1979, monetary policy shifted significantly in the U.K. towards a much more restrictive policy aimed at bringing about a disinflation through higher interest rates. The disinflation occurred gradually over the period from 1980 to 1983. See Nelson and Nikolov (2002) for more details.
United Kingdom	1992Q1	The U.K. exited the ERM in the wake of the crisis in September 1992 and announced the adoption of inflation targeting in October 1992.
Italy	1983Q4	Following its decision to join the EMS in 1979, Italy began implementing an inflation stabilization program based on commitment to an exchange rate target. Although the exchange rate anchor was rather weak initially owing to frequent realignments, the system became more stable around the middle of the 1980s (see Detragiache and Hamann (1997) for more details). Spinelli and Tirelli (1993) identify 1984 as a key date in the transition to a new regime for monetary policy, since this is when the Bank of Italy began announcing targets for M2.
Italy	1996Q1	After being forced off during the crisis in 1992, Italy re-enters the ERM in 1996.
Netherlands	1982Q2	The exchange rate has been a key element in Dutch monetary policy throughout the post-war period. Following the breakdown of the Bretton-Woods system, the Dutch authorities decided to stabilize the guilder in terms of the Deutsche mark. Although there were several devaluations in the 1970s, the peg stabilized in the early 1980s (the last devaluation was in 1983). See Hilbers (1998) for more details.
United States	1981Q2	Paul Volker was appointed Chairman of the Federal Reserve in 1979 and orchestrated a disinflation by raising interest rates (they peaked in early 1981).
United States	1990Q4	A shift in U.S. monetary policy in the 1990s has been identified and characterized as one in which the Federal Reserve responded more aggressively to rising inflation than in previous decades (see Mankiw (2001) for more details).