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TÜSİAD-KOÇ UNIVERSITY ECONOMIC RESEARCH FORUM
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**EXCHANGE-RATE PASS-THROUGH TO IMPORT
PRICES: NONLINEARITIES AND EXCHANGE RATE
AND INFLATIONARY REGIMES**

Rehim Kılıç

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Exchange-rate pass-through to import prices: nonlinearities and exchange rate and inflationary regimes

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Abstract

This paper investigates the relationship between exchange rate pass-through and exchange rate appreciations/depreciations and inflation by estimating nonlinear time series models. Motivated by theoretical and empirical results in the literature, the paper proposes new econometric models that can characterize nonlinear and asymmetric dynamics between import prices and exchange rate changes in a parsimonious fashion. Findings show the presence of complete and incomplete pass-through regimes depending upon the magnitude of appreciations of a currency and inflation rates both in the short-run and in the long-run. Results also reveal threshold effects and asymmetry in the pass-through relationship over appreciations/depreciations as well as inflationary and disinflationary periods. Findings have important macroeconomic policy implications.

JEL Classification: C22, F31, F41.

Keywords: Smooth Transition, Nonlinearity, Asymmetry, Exchange Rate Pass-through, Import Prices.

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1 Introduction

One of the central questions in international finance is the degree of exchange rate shocks that are passed-through to import and consumer prices. This question has generated a large body of research but a feature shared by most of the extant empirical international finance literature on this is a reliance on linear pass-through regressions. Findings of the extant literature suggest rather limited sensitivity of prices to exchange rate movements and considerable variation in the estimated elasticities over various sampling periods and countries (see Goldberg and Knetter 1997, Bailliu and Bouakez 2004, Bailliu and Fuji 2004, Gagnon and Ihrig 2001 and Campa and Goldberg 2005 among others). Recent studies also provide evidence that suggests a decline in exchange rate pass-through (ERPT) to import prices during 1990s and early 2000s in some industrialized countries (Campa and Goldberg 2005, Gust et al. 2006, Bouakez and Rebei 2008 and Goldberg and Campa 2010).

Despite its' simplicity and advantages, linear specifications may prevent empirical literature to address some important questions: Does the degree of ERPT to import prices depend upon regimes where regimes are characterized by the movements of some economic factors, including the past appreciations or depreciations of the domestic currency and shifting local cost factors as measured by the past inflation or disinflation? Do the ERPT estimates vary over appreciations and depreciations and over inflationary and disinflationary (or low inflationary) episodes? How does the degree of ERPT depend upon the size and sign of the exchange rate shock itself? Are there threshold effects in terms of macroeconomic factors which derive the regime-dependence and hence temporal variation in ERPT measures? Can we identify periods of "complete" and "incomplete" pass-through? These questions are important not only to understand implications of some economic models which suggest presence of nonlinearity in the ERPT relationship

(e.g., Krugman 1987, Baldwin 1988, Baldwin and Krugman 1989, Dixit 1989, Froot and Klemperer 1989, Knetter 1991, Kogut and Kulatilaka 1994, Taylor 2000 and Devereux and Yetman 2008 and Shintani et al. 2009 and Berman et al. 2009) but also for the conduct of the monetary policy (e.g. Smets and Wouters 2002, Corsetti and Pesenti 2005, Adolfson 2002, Sutherland 2005, and Monacelli 2005), the choice of exchange rate regime (Engel 2000, Devereux and Engel 2002 and 2003) and the international transmission of shocks (Betts and Devereux 2001).

This paper offers an econometric approach to addressing these questions. In a time series framework, we specify a novel ERPT regression which we refer to Smooth Transition Pass-through (STP) regression that allows for the the import price elasticity with respect to exchange rate shocks to depend upon smooth functions of macroeconomic factors including the past exchange rate appreciations and inflation rates. Two types of transition functions, namely the exponential and logistic forms that allow for symmetric and asymmetric dynamics respectively in the ERPT relationship, are estimated and diagnostically evaluated by using several tools. A Logistic STP (LSTP) regression specification adequately captures the nonlinearity and asymmetric dynamics in the ERPT relationship. This paper also extends the linearity tests typically used in the context of Smooth Transition Autoregressive (STAR) models to the STP framework. We show that the pass-through of exchange rates to import prices may have nonlinearities that can be driven by exchange rate appreciations/depreciations and inflationary environments such that the degree of pass-through is regime specific and varies over time.

To the best of our knowledge, ours is the first study that formally examines the relationship between ERPT and the past exchange rate appreciations and depreciations within the context of a smooth transition threshold regression framework. Except for Sintani et al. (2009), the previous empirical studies on the relationship between ERPT and inflationary environment have focused on cross-country evidence, including the anal-

yses by Calvo and Reinhart (2002), Choudhri and Hakura (2006) and Devereux and Yetman (2008). In this paper, instead of examining the relationship between the ERPT and average inflation rates across countries, we are interested in examining the role of deviations of inflation and exchange rate appreciations from estimated threshold levels in the time-varying dynamics of exchange rate pass-through using a time series modeling framework.

The major contribution of the current paper is to show that the ERPT can be characterized by the presence of smooth transition type threshold effects with two extreme regimes; one with “incomplete” and another with “complete” pass-through, depending on the degree of past exchange rate appreciations and inflation rates, for the import prices analyzed. Both short run and long run pass-through estimates show time-variation over appreciations/depreciations and over inflationary and disinflationary periods. Our empirical analysis uncovers important asymmetries in the degree of pass-through to import prices over past appreciations and depreciations as well as over low inflationary (or disinflationary) and high inflationary periods. Findings also reveal considerable variation in the degree of pass-through and speed of adjustment dynamics between incomplete and complete pass-through regimes. Our estimates of STP models and the analysis reported in the paper may reconcile the mixed empirical evidence reported in the literature for various countries over different sampling periods. For example, the decline in the pass-through during the late 1980s and most of the 1990s to import prices might be explained by the the degree of deviations of past currency appreciations and inflation rates from the estimated threshold levels -especially for Canada, Japan, Germany, the US and the UK.

The two papers most closely related to ours are Herzberg et al. (1993) and Shintani et al. (2009). There are key differences between those studies and our approach. First, both of these papers utilize STAR models to study nonlinear dynamics. Herzberg et

al. (2003) study the ERPT to UK import prices by estimating STAR models and report no evidence of nonlinearity. Shintani et al. (2009) on the other hand, investigate nonlinearity in ERPT to monthly US producer price inflation rates by estimating various forms of STAR models. The STAR models employed by Shintani et al. (2009) force price changes depend on both its own past values and current and past values of exchange rate changes in a nonlinear fashion. Therefore, in their framework nonlinear dynamics may arise because prices themselves and/or the ERPT have nonlinearity. In contrast, the STP model introduces nonlinearity into the ERPT relationship itself and thereby allows us to investigate nonlinear dynamics directly and follows very closely the specification used by Campa and Goldberg (2005) in measuring the short run and long run ERPT. Second, we consider two macroeconomic factors in characterizing the regime-dependent dynamics in ERPT and show that both the past behavior of exchange rates and shifting-cost factors (e.g. inflation rates) can identify regimes and thereby provide a link between the macroeconomic environment and the regime-dependence (and hence time-variation) in ERPT. In contrast, Shintani et al. (2009) only consider inflation rates as the driving factor in the nonlinearity. Third, our framework also controls for market size as well as import partner's cost as failing to control these factors may bias the pass-through estimates. Despite differences in terms of methodology and data set, our results complement and extend Shintani et al. (2009)'s findings by showing the presence of regime-dependence in ERPT to import prices in several countries at the quarterly frequency.

The remainder of the paper is organized as follows. Section 2 discusses briefly theoretical models which suggest presence of nonlinear dynamics in the relationship between import prices and exchange rates. Section 3 describes the empirical model. Data and estimation results are discussed in Sections 4 and 5 respectively. Some concluding remarks are made in Section 6.

2 Nonlinearity in exchange rate pass-through

In this section we briefly discuss theoretical work that implies the presence of nonlinear dynamics in the relationship between import/export prices and exchange rate changes. Theoretical work has suggested a number of potentially important factors in characterizing the nonlinear features in ERPT. First, in oligopolistic markets, the response of import prices to changes in exchange rates or other cost items depends both on the curvature of demand and the market structure (Dornbusch 1987, Knetter 1989, Bergin and Feenstra 2001, Atkeson and Burstein 2008) and the size and sign of the shocks to costs including appreciations and depreciations of a currency. Second, local costs (as measured by the inflationary environment at the aggregate level) may play an important role in determining pass-through (Sanyal and Jones 1982, Burstein et al. 2003, Corsetti and Dedola 2004). Local costs drive a wedge between prices and imported costs that is unresponsive to exchange rate fluctuations. As a consequence, if local costs are large, even a substantial increase in the price of an imported factor of production could have little impact on marginal costs thereby causing pass-through to vary over inflationary environments. Third, price rigidity and other dynamic factors have the potential to contribute to time-variation and incomplete pass-through (Giovannini 1988, Kasa, 1992, Devereux and Engel 2002, Bacchetta and van Wincoop 2003).

In this paper, we primarily focus on two key macroeconomic factors, namely the the degree of appreciations and depreciations of domestic currencies, and inflation rates which may characterize the smooth transition type nonlinear dynamics in ERPT to import prices (Taylor 2000, and Devereux et al. 2004). Within this context, one strand of literature points out the presence of menu costs, transport costs or other fixed costs of entering the export markets in generating nonlinear dynamics in exchange rates and the pass-through relationship. For example, studies by Krugman (1987), Baldwin (1988),

Baldwin and Krugman (1989), Dixit (1989), Froot and Klemperer (1989), Knetter (1993) and Kogut and Kulatilaka (1994) provide models that show how firm behavior differs when a currency is depreciating or appreciating. These studies suggest that nonlinearity in exchange rate pass-through can arise when the correlation between import prices and exchange-rate fluctuations differs over appreciations and depreciations. Moreover, these models also imply that the presence of costs (either sunk or fixed costs) associated with trade can create hysteresis type behavior in export and import decisions and hence may lead to nonlinearity in the import-price elasticities with respect to exchange rate movements as the difference between marginal cost of exporting or importing and the marginal benefit may change over appreciations and depreciations. In a recent paper, Berman et al. (2009) develop a model with heterogenous firms in terms of productivity and show that high and low productive firms react differently to exchange rate movements as they face different demand elasticities. Although they do not explore nonlinearity in the exchange rate pass-through, their model and results imply that since the size of the fixed and distribution costs faced by heterogenous firms may depend, among other factors, on the exchange rates itself, pass-through to prices should have nonlinearities as markups will vary over firms and over different degrees of exchange rate changes.

A second line of literature suggests the importance of variation in local cost factors, namely the inflationary environment in characterizing nonlinearity. Beginning with Sanyal and Jones (1982), Taylor (2000), Burstein et al. (2003) Corsetti and Dedola (2004), Devereux and Yetman (2008) and Shintani et al. (2009) several papers suggest that ERPT to import or consumer prices depends upon the inflationary environment. For example Devereux and Yetman (2008) developed a model of importing firms where ERPT is predicted to depend on the steady-state average inflation rate. On the other hand Shintani et al. (2009) extend Devereux and Yetman (2008)'s model by removing price-stickiness and infinite-horizon profit maximizing importers to show that, under a more realistic

setup, pass-through may depend upon past levels of inflation.

The research also suggests asymmetric behavior in the pass-through relationship as persistent appreciations and depreciations of domestic currency and/or persistent inflationary or disinflationary periods may induce different import/export pricing strategies and, hence, may lead to variation in the pass-through to import prices over periods of appreciations and depreciations. Moreover, firms may react to the magnitude of the appreciations and depreciations differently and develop risk management and pricing strategies that are more proactive for large changes in the value of a currency. This suggests that pass-through dynamics differ for various levels of exchange rate uncertainty. For example, Caballero and Engel (1993) constructed a microeconomic model with asymmetric price adjustment at firm level and examined the resulting aggregate time series. Their analysis showed that the aggregate price level responds less to negative shocks (say appreciation of the importing country's currency) than to positive shocks, that the size of this asymmetry increases with the size of shock, and that the number of firms changing their prices and therefore the flexibility of the price level to aggregate shocks varies endogenously over time in response to changes in economic conditions such as exchange rates or inflation.

Given the implications of the theoretical work discussed above, the conventional methods used in the extant literature fail to accurately account for the nonlinear dynamics. In the following sections, we aim to investigate more formally the nonlinearity and asymmetry in the pass-through to import prices by using a class of nonlinear time series models which allow us to model both smooth transition type threshold dynamics in pass-through over appreciations and depreciations as well as over various (low/high) inflationary regimes.¹

¹We recognize that the specific form of nonlinear dynamics in ERPT relationship implied by various studies may not be unique and hence may differ from the proposed STP models. In these instances, the STP regressions may approximate more general forms of nonlinearity in a parsimonious and intuitive

3 Smooth Transition Pass-through Regressions

To quantify the pass through of exchange rates to import prices we follow Goldberg and Campa (2005) and use the following linear benchmark regression model

$$\Delta p_t^j = \alpha + \sum_{i=0}^p \beta_i \Delta s_{t-i}^j + \sum_{i=0}^p \omega_i \Delta w_{t-i}^j + \psi \Delta gdp_t^j + u_t^j \quad (1)$$

where p_t is the log local currency import price for country j , s_t is the exchange rate, w_t is the foreign production costs (it is a primary control variable representing exporter costs), and gdp is the real GDP of the destination market. In this model, the short run relationship between the import prices and the exchange rate is given by β_0 . In other words, β_0 measures the short-run import price elasticity of contemporaneous exchange rate changes (i.e. short-term pass-through). The long-run elasticity is given by the sum of the coefficients on the contemporaneous exchange rate and p lags of exchange rate terms $\sum_{i=0}^p \beta_i$. In Campa and Golderg (2005) p is set to 4 with quarterly data. As discussed in Campa and Goldberg (2005) this specification can be obtained from the microfoundations of pricing behavior by exporters under markup pricing. Some papers do not control for the export country costs and market size (as measured by the real GDP). As discussed in Campa and Golderg (2005) ignoring these controls may induce bias in the estimates of pass-through coefficients.

In order to to capture smooth transition type threshold dynamics in ERPT relations, we extend the linear pass-through regression by the following nonlinear specification

way which allows for a very rich regime-dependent dynamics in ERPT. Moreover, we conduct evaluate proposed models by extensive diagnostic and specification tests as will be discussed in the next section.

which we call the Smooth Transition Pass-through (STP) regression:

$$\Delta p_t^j = \alpha + \left[\sum_{i=0}^p \beta_i \Delta s_{t-i}^j \right] (1 - F(\gamma, c, z_{t-d})) + \left[\sum_{i=0}^p \beta_i^* \Delta s_{t-i}^j \right] F(\gamma, c, z_{t-d}) \quad (2)$$

$$+ \sum_{i=0}^p \omega_i \Delta w_{t-i}^j + \psi \Delta g d p_t^j + u_t^j$$

where u_t is a zero mean, stationary disturbance term, and $F(\cdot)$ is the transition function which controls the nonlinear dynamics and is chosen to be either the exponential function,

$$F(\gamma, \mu, z_{t-d}) = 1 - \exp(-\gamma(z_{t-d} - c)^2), \quad (3)$$

or the logistic function,

$$F(\gamma, c, z_{t-d}) = 1 / [1 + \exp(-\gamma(z_{t-d} - c))]. \quad (4)$$

In Eqns. (3) and (4), z_t is the transition variable while d is called the delay parameter, γ is a slope parameter (or the transition parameter) and c is a location parameter (or the threshold parameter). The parameter restriction $\gamma > 0$ is an identifying restriction. Under any of these functions the d -lagged period values of transition variable z characterize the transition and pass-through dynamics.

The STP regression is related to the STAR models introduced by Granger and Teräsvirta (1993) and by Teräsvirta (1994). In these models, the time series process smoothly evolves or adjusts to an equilibrium relationship in a nonlinear fashion where the precise nature of nonlinear dynamics is governed by the past values of a predetermined (endogenous) variable which is called the transition variable. Indeed, the adjustment process in the STP model occurs in every period and the speed of adjustment is governed by the transition variable and the transition parameter which measures the speed of transi-

tion across various regimes. While there are many possible choices of transition functions, the exponential and logistic function specifications are attractive in the present context as they allow for both symmetric and asymmetric response of the import prices to changes in the exchange rate.

Both the logistic function and the exponential functions are bounded between 0 and 1 and depend on the transition variable z_t . Despite both functions are bounded between 0 and 1, the implied nonlinear dynamics under logistic and exponential functions are drastically different. The exponential transition function takes on the value of unity for very large positive and negative values of the transition variable, i.e., as $z_t \rightarrow \pm\infty$, $F(.) \rightarrow 1$, and whenever the transition variable is in the neighborhood of threshold parameter c , it approaches to 0. The logistic transition function approaches zero for very large negative values of transition variables, i.e. as $z_t \rightarrow -\infty$, $F(.) \rightarrow 0$ and as $z_t \rightarrow +\infty$ $F(.) \rightarrow 1$, and whenever the transition variable is in the neighborhood of threshold parameter c , it takes on the value of 0.5. Depending on the choice of the transition function, the STP model will be referred to as Logistic STP (LSTP) or the Exponential STP (ESTP) model.

When $\gamma \rightarrow \infty$, the logistic transition function $F(.)$ approaches a step function, as such the LSTP model effectively becomes a threshold model. Therefore, the LSTP model nests a two-regime threshold model. On the other hand, exponential function becomes flat (with an abrupt swing around c) and hence does not nest a threshold model. In both models, the parameter c can be interpreted as the threshold between the two regimes corresponding to $F(z_t; \gamma, c) = 0$ and $F(z_t; \gamma, c) = 1$ in the sense that the logistic function changes monotonically from 0 to 1 as z_t increases and that the exponential function changes from 0 to 1 as z_t increases in absolute value (both in positive and negative directions). This suggests that, despite both models imply presence of two extreme regimes for $F(z_t; \gamma, c) = 0$ and $F(z_t; \gamma, c) = 1$, the dynamics across these regimes are

substantially different under each model. The ESTP suggests that for values of z_{t-d} in the neighborhood of threshold value c , the pass through relationship will be characterized by the regime where $F(z_{t-d}; \gamma, c) = 0$. For very large positive and negative values of z (i.e. values that exceed the threshold parameter) pass through is characterized by the regime $F(z_t; \gamma, c) = 1$. On the other hand, under LSTP specification, the pass through is characterized by the regime where $F(z_t; \gamma, c) = 0$ when z_t decreases and by the regime $F(z_t; \gamma, c) = 1$ when z_t increases. Therefore, in both models regimes associated with $F(z_t; \gamma, c) = 0$ and $F(z_t; \gamma, c) = 1$ are called the lower and the upper regimes respectively. Note also that in the LSTP model there is an intermediate regime which occurs when $F(z_t = c; \gamma, c) = 0.5$.

We should also note that both the short run and long run pass through coefficients change smoothly between the lower and upper regimes and in the extreme regimes, taking on values of $(\beta_0$ and $\sum_{i=0}^p \beta_i$ respectively), and $(\beta_0^*$ and $\sum_{i=0}^p \beta_i^*$ respectively), respectively. Therefore, depending upon the estimated coefficients, STP regressions can capture time-varying pass through of exchange rate changes to import prices where the temporal dynamics are governed by predetermined economically relevant variables such as past appreciations and depreciations of a currency or the past inflationary changes. Depending upon the sign and magnitude of transition variables, the proposed models can characterize and identify complete and incomplete pass-through regimes. Thus, findings of the estimated models can be very useful in providing insights into our understanding not only of the dynamics of pass through relationship but can also reconcile findings of the empirical pass through literature. As discussed above, ESTP model would imply that the pass through to import prices is symmetric with respect to negative and positive deviations of z_t from the threshold level c while LSTP implies different pass through profiles for negative and positive deviations and can thereby be useful when there is asymmetry as well as nonlinearity in the pass through relationship. Among the important

advantages of the proposed STP specification is the rich set of dynamics which it allows to capture despite its relative simplicity, estimability via a nonlinear least squares based approach, and observability of the variable triggering regime switches which may help attach a cause to the underlying regime-dependent dynamics in the ERPT relationship.

4 Data

Following Goldberg and Campa (2005) we use import prices to investigate the pass through. Import price data is obtained from the OECD Statistical Compendium. The quarterly import price index data in local currency is available for the period 1975Q1 to 2009Q1 from International Financial Statistics (IFS) for the set of countries we investigate in this paper. In constructing inflation series, we use consumer price indices which are obtained from IFS. Nominal exchange rates are from IFS (series neu), defined in our specifications as domestic currency per unit of foreign currencies ($1/\text{neu}$), so that home-currency depreciations appear as increases in the nominal exchange rate series. Real exchange rates also are from the International Financial Statistics (series reu). The real GDP of the importing country is used as a proxy for the total demand in the importing country. The real GDP series are obtained from IFS.

Following Campa and Goldberg (2005) we construct a consolidated exporter partners' cost proxy. The cost variable is measured as $W_t = \frac{\text{neu}_t \times \text{ULC}_t}{\text{reu}_t}$ where ULC_t is the unit labor cost (obtained from OECD Statistical Compendium). This provides us with a measure of country j 's trading partners' costs where each partner is weighted by its importance in country j 's trade. Since unit labor cost measures are available after 1980Q1 for Japan, and real exchange rate series (reu) are available after 1984Q1 for Australia, we have 117 and 104 quarterly observations for Japan and Australia respectively. For all others our sample has 137 observations.

5 Empirical Results

5.1 Linear pass-through estimates and tests for nonlinearity

Before proceeding with the estimation of the nonlinear models, first we report results from the linear pass through regression models in Table 1. We find that interesting cross-country differences in pass-through into import prices. The United States and the United Kingdom have relatively low pass-through, 26% and 32%, respectively, within one quarter and 42% and 46%, respectively, over the longer run. Pass-through estimates for Canada, Germany, Australia, and Japan on the other hand are between 56% and 70%. We fail to reject the complete long-run pass-through for Germany and Japan in this sample which is consistent with the findings reported by Campa and Goldberg (2005). Contrary to the findings of Campa and Goldberg (2005), on the other hand, we reject the complete pass-through for Canada in the long-run. Indeed, the long-run pass-through estimate for Canada is smaller than the short-run pass through (about 59%). Our empirical findings are, broadly speaking, in line with the results reported in Campa and Goldberg (2005).²

The last four columns of Table 1 report p-values for testing linearity against the LSTP and ESTP type nonlinearity in the residuals of linear pass-through regressions by using lagged and average exchange rate appreciations/depreciations as well as inflation series over $d = 1, 2, 3, 4$ quarters. We use linearity tests due to Granger and Teräsvirta (1993) and Teräsvirta (2004). These tests approximate the logistic and exponential type nonlinear component by taking an appropriate Taylor series expansion around the null of linearity. We computed several versions of these tests as discussed in the literature. The reported results are based on a third order Taylor series expansion of logistic transition function and a second order Taylor series expansion of the exponential transition function.

²Our findings also reveal empirically significant coefficient estimates for the country size as measured by the real GDP for all countries except for Japan (coefficient estimates for real GDP are not reported). These findings are in contrast to Campa and Goldberg (2005).

Results from other approximations are qualitatively similar and can be obtained upon request. We compute each test over $z_t = \{\Delta s_{t-d}, \frac{1}{d} \sum_{i=2}^d \Delta s_i\}$ and $z_t = \{\pi_{t-d}, \frac{1}{d} \sum_{i=2}^d \pi_i\}$ respectively for delay parameters $d \in \{1, 2, \dots, 4\}$. The reported p-values correspond to the transition variable and delay parameter for which the LM statistics have been optimized. Careful inspection of the reported findings show that residuals from the linear model have significant nonlinear components especially of the logistic form. More specifically, p-values show statistically compelling evidence in favor of logistic type nonlinearity in the residuals of linear pass-through regressions for all countries.

5.2 Nonlinearity in pass-through to import prices

We follow the specification procedures suggested by Teräsvirta (2004) and estimate both LSTP and ESTP models with transition variables given by $z_t = \{\Delta s_{t-d}, \frac{1}{d} \sum_{i=2}^d \Delta s_{t-i}\}$ and $z_t = \{\pi_{t-d}, \frac{1}{d} \sum_{i=2}^d \pi_{t-i}\}$ for delay parameters $d \in \{1, 2, \dots, \bar{d} = 4\}$ and compare alternative specifications by using several diagnostic tests. We utilize tests for serial correlation and normality in residuals, tests for parameter constancy, and tests for remaining nonlinearity of exponential and logistic forms in the residuals of estimated models suggested by Eitrehiem and Teräsvirta (1996). This procedure involves estimation and diagnostic testing of several models for each country. Findings reveal that for all import prices LSTP models constantly outperform the ESTP models for all transition variables and delay parameters considered.³ These extensive estimation and diagnostic results confirm our findings from the linearity test results reported in Table 1.

In Table 2, we report our final specifications of LSTP models for each of the countries. For each country, the first column reports the results when the transition variable is given by the past exchange rate changes and the second column reports when the tran-

³More specifically, residuals from the ESTP models show statistically significant remaining nonlinearity of logistic form and parameter non-constancy in most of the cases. Complete estimation and testing results can be obtained upon request.

sition variable is the past inflation rate. The estimation results show that the transition parameter γ (which is normalized by the standard deviation of the relevant transition variable) is statistically significantly different from zero for all countries and transition variables. The parameter estimates and reported robust standard errors reject the null of $\gamma = 0$ at conventional significance levels. Strictly speaking, testing $\gamma = 0$ by using conventional critical values is not correct as the test statistic may not have the standard asymptotic null distribution. This is because under the null, short-run and long-run pass-through parameters in the extreme regimes (i.e. $\beta_0, \sum_i \beta_i$ and $\beta_0^*, \sum_i \beta_i^*$) and the threshold parameter c are not identified and hence the asymptotic distribution of the test statistic depends on the unidentified nuisance parameters (see Davies 1987, for example). To overcome this problem, we utilize two approaches. In the first approach, we compute p-values via simulations. The simulated p-values are computed by generating data through calibrating on the parameters of the linear pass-through regressions (the null model reported in Table 1) and drawing randomly from the residuals of the model. Then the LSTP models are estimated and t statistic for $\gamma = 0$ are computed. The procedure is repeated 2000 times. Then the proportions of “simulated” t -statistics that are greater than the actual t -statistic are reported as the p-values in Table 2.

In the second approach we extend the Taylor series expansion procedure typically used in the context of STAR models (see Teräsvirta 1994) to the STP specification. To the best of our knowledge this extension is somewhat novel as the STP specifications we use in this paper differ from those used in the STAR models (for an application of a similar approach in the context of conditional volatility see Kılıç 2010). The first order Taylor series expansion of the LSTP model around the null hypothesis $\gamma = 0$ leads to

$$\Delta p_t^j = \alpha + \sum_{i=0}^p \delta_i \Delta s_{t-i}^j + \sum_{i=0}^p \delta_i^* \Delta s_{t-i}^j z_{t-d} + \sum_{i=0}^p \omega_i \Delta w_{t-i}^j + \psi \Delta gdp_t^j + \varepsilon_t^j \quad (5)$$

where $\delta_i = [\beta_i - \frac{\gamma}{4}c\lambda_i]$, $\delta_i^* = \frac{\gamma}{4}\lambda_i$ with $\lambda_i = (\beta_i^* - \beta_i)$ and ε is an error term. A test of the linear pass-through regression model against the LSTP model can be carried out by estimating (5) and testing $H'_0 : \sum_{i=0}^p \delta_i^* = 0$ against the alternative that the null is not correct by using a robust Wald test. Under the null, the Wald statistic should have a χ -squared distribution with $p + 1$ degrees of freedom. The third Taylor series approximation of the LSTP model around $\gamma = 0$ gives the following:

$$\Delta p_t^j = \alpha + \sum_{i=0}^p \delta_i \Delta s_{t-i}^j + \sum_{i=0}^p \delta_{1,i}^* \Delta s_{t-i}^j z_{t-d} + \sum_{i=0}^p \delta_{2,i}^* \Delta s_{t-i}^j z_{t-d}^2 + \sum_{i=0}^p \delta_{3,i}^* \Delta s_{t-i}^j z_{t-d}^3 + \sum_{i=0}^p \omega_i \Delta w_{t-i}^j + \psi \Delta g d p_t^j + \varepsilon_t^j \quad (6)$$

where, now, $\delta_i = [\beta_i - \lambda_i c (\frac{\gamma}{4} + \frac{\gamma^3}{48} c^2)]$, $\delta_{1,i}^* = -\lambda_i (\frac{\gamma}{c} + 3 \frac{\gamma^3}{48} c^2)$, $\delta_{2,i}^* = 3 \lambda_i \frac{\gamma^3}{48}$, and $\delta_{3,i}^* = -\lambda_i \frac{\gamma^3}{48}$. Note that testing the linear pass-through against the LSTP model becomes equivalent to testing the joint significance of the polynomial coefficients, i.e. $H''_0 : \sum_{i=0}^p \delta_{1,i}^* = \sum_{i=0}^p \delta_{2,i}^* = \sum_{i=0}^p \delta_{3,i}^* = 0$ by using a Wald test. Under the null of linearity, the proposed test should have an asymptotic χ^2 with $3(p + 1)$ degrees of freedom. In Table 2, we label the p-values from these tests as pW_1 and pW_3 respectively.⁴

Both the simulated p-values and the p-values corresponding to the Wald tests strongly reject the null hypothesis that $\gamma = 0$ which lends statistical support for the proposed nonlinear pass-through specifications. Estimated transition parameters show variation in the speed of the transition between the lower and outer regimes (i.e. between $F(\cdot) = 0$ and $F(\cdot) = 1$) across countries. For example when exchange rate change is used as the transition variable, estimated value for US import prices is about 2.7 while for Australia

⁴Note that the argument for the asymptotic critical values relies on the assumption that the LSTP specification satisfies the necessary stationarity conditions. Tests for stationarity and unit root (not reported) show that import price and exchange rate changes and other regressors are stationary. Investigating the asymptotic distribution under different assumptions for the regressors and the error term is an interesting topic of its own and beyond the scope of this paper.

it is about 8.4. Similarly, when we use the past inflation rate as the transition variable, the estimated transition parameter ranges between about 3.1 (for Japan) and 7.9 (for Australia).

Estimates of delay parameters show that for Japan, aggregate import price changes respond to exchange rate changes after one quarter while for US, Canada, and Germany response is delayed about four quarters when the transition dynamics are driven by past movements of currency. For the UK, on the other hand, average appreciation/depreciation of Pound over the previous three quarters drives the nonlinear dynamics. Using the past inflation rate as transition variable shows that for the US, the UK, Germany, and Australia the past quarter's inflation rate characterizes the smooth transition dynamics. For Canada and Japan results suggest that the import price response to exchange rate changes is delayed three quarters. Estimated threshold parameters are statistically significant for all countries at conventional significance levels. Estimates for the threshold parameters reveal differences across countries for each given transition variable. The variation in the slope of the transition functions and the estimated threshold values across countries may be due to differences in export/import market conditions including the composition of imports, market demand, and the degree of competition among other factors.

More evidence is reported for the estimated models by the several diagnostic tests reported in Table 2. Reported Sum of Squared Residuals (SSR) ratios between linear and the nonlinear models (SIC values-not reported) favor the LSTP model against the linear pass-through regressions. Moreover reported p-values from remaining nonlinearity of exponential and logistic form in the residuals of the estimated models show no evidence of remaining nonlinearity in the residuals in contrast to the linear models reported in Table 1. Further evidence in favor of LSTP models are provided by the findings from the W_1 and W_3 tests discussed above. Overall diagnostics of the estimated models are considerably

good and both diagnostic tests and other hypothesis tests, as well as information criteria (not reported), favor the LSTP models against the linear models.

We further analyze the validity of the LSTP models by estimating an ERPT regression for all the observations lying in each “extreme” regime. This seems a reasonable procedure since the residuals in LSTP regressions appear to be approximately serially uncorrelated, so that each observation in a particular regime is randomly drawn throughout the sample. In the results reported in Table 3, the so called “complete” regime is defined to be the regime at which $F(.) \leq 0.25$ when the past exchange rate change is the transition variable (and hence “incomplete” regime is defined as $F(.) \geq 0.75$). Similarly when the transition variable is the past inflation rates, the “complete” and incomplete regimes are given by $F(.) \geq 0.75$ and $F(.) \leq 0.25$ respectively. This way of defining complete and incomplete ERPT regimes is clearly arbitrary, but is necessary due to the uncertainty in the estimation of the transition function. The choice of values adopted here for the extreme regime identification was found to be consistent across countries in order to achieve a reasonable number of observations in each regime. To restrict the regime to values of the transition function to, say, 0.05 or 0.95 greatly reduces the number of observations for some countries which renders the ERPT regression very sensitive to estimation error. Also, given the well-known difficulties and uncertainty of estimating the transition functions in the context of smooth transition models (see for example Teräsvirta 1994), it seems reasonable to err on the side of inclusivity rather than exclusivity by using reasonably wide bands. Analogous results for Table 3 for a full grid of values are available from the authors on request, but are omitted for reasons of conserving space.

Results of estimating the linear ERPT regression over the “complete” and “incomplete” ERPT regimes from the LSTP model are given in Table 3. It can be seen that the estimated regressions for the “complete” regime are apparently quite supportive of

the complete pass-through, with considerably higher R^2 values despite the very small sample sizes when compared to incomplete regimes. Robust Wald tests fail to reject the complete ERPT for five of the countries when the transition variable is given by past exchange rate changes and for four of the six countries when the transition function is given by the past inflation rates. These findings are also supported by the 95% confidence intervals despite the estimated intervals' tendency to be considerably wide due to the occurrence of much smaller sample sizes in the complete ERPT regimes. These results provide further support for the LSTP models.

5.3 ERPT over exchange rate appreciations and depreciations

Findings from Table 2 show that for all countries, when the past (average) currency appreciations exceed estimated threshold levels, the relationship between pass-through to import prices and exchange rate changes approach to the lower regime where the degree of pass-through increases. Careful inspection of the reported p-values for the robust t and Wald statistics for testing the complete short-run (i.e. $\beta_0 = 1$ and $\beta_0^* = 1$) and long-run pass-through (i.e. $\sum_{i=0}^4 \beta_i = 1$ and $\sum_{i=0}^4 \beta_i^* = 1$) in the extreme regimes suggest strong evidence for the complete pass-through to import prices in the both short and long-run for all countries, except for the UK, when the local currency appreciations are far above the estimated threshold levels. On the other hand, incomplete pass-through tends to occur for small appreciations or for depreciations.

Tests for equality of estimated short-run and long-run pass-through measures in the lower and upper regimes strongly rejected by the reported p-values for the robust Wald tests (with the exception of Germany) sending further support for the nonlinear and asymmetric pass-through dynamics over large appreciations and depreciations of each country's currency.⁵ Overall tests and analysis indicate the presence of two extreme

⁵Similar to testing $\gamma = 0$, testing the equality of long run pass-through across extreme regimes

regimes one with “complete” or “near-complete” short-run and long-run pass-through and another with “incomplete” pass-through depending upon the past appreciations and depreciations of exchange rates.

To gain further insight into the time-variation and regime-dependence on the ERPT, we display estimated transition functions and short run and long run pass-through estimates over the transition variables in Figures 1 and 2.⁶ Careful inspection of the plots reveal that estimated transition functions visit both extreme regimes during our sample period and transition from the incomplete pass-through regime to the complete pass-through regime occurs for large enough appreciations of these currencies.

The displayed transition functions and ERPT estimates over the transition variables in Figure 1 show that for large enough appreciations, the difference between short-run and long-run pass-through estimates increases and for the appreciations of currencies in the neighborhood of estimated threshold levels, the difference tends to be smaller. Plots also reveal that for small appreciations and depreciations estimated transition functions approach to the incomplete pass-through regime (i.e. regime where $F(\cdot) = 1$) where both short-run and long-run pass-through tend to be relatively low. An interesting observation is that for both depreciations and appreciations of about 2-3%, transition functions move towards incomplete pass-through regime. In order for the exchange rate changes to be completely passed-through to import prices, currencies not only need to appreciating over the past d - quarters but the amount of appreciation should also be high enough (exceed the estimated threshold levels) so that a cluster of firms may adjust their mark-ups and hence the import prices respond to a given exchange rate shock. Note that this

involves the nuisance parameter problem. This is because under the null that $H_0 : \sum_i \beta_i = \sum_i \beta_i^*$ both the transition (γ) and threshold (c) parameters are not identified. We compute marginal significance levels by simulations as discussed above. Since the simulated p-values are very similar to the asymptotic p-values, we report the asymptotic p-values for Wald tests (i.e. pW_{LR}) in Table 2.

⁶Note that the short-run pass-through is given by $SRPT_t = \hat{\beta}_0(1 - F(\hat{\gamma}, \hat{c}, z_{t-d})) + \hat{\beta}_0^*F(\hat{\gamma}, \hat{c}, z_{t-d})$ while long-run pass-through is $LRPT_t = \left[\sum_{i=0}^4 \hat{\beta}_i \right] (1 - F(\hat{\gamma}, \hat{c}, z_{t-d})) + \left[\sum_{i=0}^4 \hat{\beta}_i^* \right] F(\hat{\gamma}, \hat{c}, z_{t-d})$.

indicates that import prices may respond to even smaller sized shocks once the degree of past appreciation of a currency reaches and exceeds estimated threshold levels. This may suggest that firms probably absorb (or at least partially absorb) the changes in the exchange rates before the appreciation of a currency reaches and exceeds certain threshold levels. As the currency appreciation approaches to the neighborhood of a certain threshold level, mark-ups depress and firms become increasingly forced to pass the exchange rate changes into prices.

Plots reveal that estimated transition functions tend to have considerably large numbers of realizations in the upper regime compared to the lower regime for all countries. Therefore, we observe more frequent incomplete pass-through than complete pass-through. These findings, further supported by the graphs in Figure 3 which displays long-run pass-through estimates and past movements in the transition variable, namely the past exchange rate changes over time with the estimated threshold levels superimposed. Careful inspection of the plots in Figure 3 reveals that for all countries, pass-through estimates change considerably over time when the transition dynamics are characterized by exchange rate changes. Plots show that past movements in exchange rates, for all but the Australian Dollar, were mostly above the estimated threshold levels. This suggests depreciations or appreciations that are much smaller than the estimated threshold levels during the sampling period. Consistent with this, pass-through estimates tend to be lower during most of the time periods with occasional swings towards complete pass-through region. This dynamic may explain why pass-through estimates based on linear regression specifications may suggest weak responsiveness of import prices to exchange rates.

Careful inspection of the plots also show that countries differ in terms of temporal ERPT dynamics. For example, estimates for the US stay in the neighborhood of 0.3 during most of the initial period of our sample, the second half of the 1980s, most of the

1990s, and the second half of the 2000s. The displayed plots reveal that import price elasticities approach to one during the long US Dollar appreciation of the 1980s and the appreciations of the late 1990s and early 2000s. For the UK, estimates stay mostly in the neighborhood of zero (in fact -0.15) and reach to the order of 0.6 during the late 1970s and early 1980s, with occasional swings towards 0.6 during our sampling period. For Canada estimates tend to stay in the order of 0.4 for most of the time periods with the exception of 1975-1976, the early 1980s, late 1980s, early 1990s, and most of the period after 2005. For Japan, we observe pass-through estimates to swing towards complete pass-through during the early and later parts of 1980s, the first half of the 1990s, during the late 1990s, and the early 2000s, and during the mid 2000s. Long run pass-through estimates for Germany and Australia show relatively more variation over time with a small range for Australia (in the range of about 0.4 to 0.7).

5.4 ERPT and low and high inflation rates

When past inflation is used as the transition variable, estimation and test results reported in Table 2 indicate that for all countries short-run pass-through is statistically different across lower and upper regimes, except for Canada. The larger the past inflation rates are the larger both the short-run and long-run ERPT are (only exception is the short-run pass-through for Canada). Plots in Figure 2 also reveal that once the annualized inflation rate exceeds the estimated threshold value estimates increase and we fail to reject the complete pass-through for all countries at the 5% significance level. On the other hand, when the inflation rate falls far below the estimated threshold levels, long-run pass-through estimates are statistically different from zero but less than one indicating incomplete pass-through to import prices during low inflation and disinflation periods.

Estimation results and displayed plots in Figure 2 indicate the presence of two distinct

pass-through regimes; one consistent with a complete pass-through and another with incomplete pass-through especially in the long-run. For the US, the UK, Japan, Germany, and Australia both short-run and long-run pass-through follows pretty similar dynamics as both tend to increase with inflation rates. On the other hand, for Canada short-run pass-through and long-run pass-through show somewhat differential dynamics over the inflationary periods. However, we fail to reject the null hypothesis that short-run pass-through coefficients in the lower and upper regimes are the same for Canada. We also note that for the US and the UK in the incomplete pass-through regime, short-run import price elasticities tend to frequently be in the vicinity of zero while for Japan, Canada, Germany, and Australia are above zero. This suggests very small or no pass-through to import prices in the US and the UK during low inflation periods.

Plots of long-run pass-through estimates over time and past inflation rates are displayed in Figure 4 with the estimated threshold levels of inflation rates superimposed. Careful inspection of the plots reveal that pass-through estimates tend to stay in the neighborhood of unity during the late 1970s, early 1980s and the late 2000s for the US and the UK. For Japan estimates suggest near complete pass-through during most of the 1980s, the first part of 1990s, and the late 2000s. For Canada estimates stay in the vicinity of 0.6 to 0.8 during most of the late 1970s and the 1980s with sharp decreases during the early 1990s and show considerable variation in the late 1990s and the 2000s. Long-run pass-through estimates are near the complete pass-through levels for Germany during the second half of the 1970s, the early and late 1980s, the early 1990s and the late 2000s. Estimates fall towards 0.6 especially during the second half of the 1990s and the late 2000s with occasional increases. Plots for Australia show that the long-run pass-through estimates stay near 0.7 between 1985 and the earlier part of the 1990s and then stay mostly around 0.5 afterwards with a few swings towards 0.7.

For the US and the UK, past inflation rates are below the estimated thresholds (about

4.4% and 3.0 % respectively) for most of the period after the 1980s until the late 2000s. Hence, the ERPT is far from the complete pass-through regime. For Germany and Japan inflation rates fall below the estimated thresholds (about 1.4% and 0.5% respectively) after 1995 and pass-through estimates are therefore closer to about 0.4 after 1995. Similarly for Australia past inflation rates tend to stay below the estimated threshold level for most of the period after 1990 and as a result, the estimates stay near 0.55. For Canada, past inflation rates frequently stay either above or below the estimated threshold rate of 0.36% causing more temporal variation in pass-through estimates. The sharper decreases in inflation rates at various periods after the 1990s lead ERPT to stay in the neighborhood of zero relatively more frequently in contrast to other countries. This might explain why some studies reported considerable decline in the Canadian ERPT during the late 1990s and the early 2000s.

Inspection of findings from Table 2 and Figures 2 and 4 show that long-run pass-through estimates stay in the range of about 0.35 and 1.0 for the US (very similar finding when exchange rate appreciations derive the transition dynamics) during the sample period. On the other hand, estimates range now between 0 and 1.4 and between 0 and 0.8 for Japan and Canada respectively. Long run pass-through estimates when inflation rate characterizes the nonlinear dynamics now ranges between 0.35 and 0.55 for the UK and still a very narrow range is obtained for Australia (between 0.55 and 0.75) for the entire sampling period. The range of estimates for Germany stays about the same as before.

6 Conclusion and Discussion

In this paper, we investigate empirically the ERPT over the floating period by utilizing novel econometric models. Specifically, we show that a logistic smooth transition ex-

change rate pass-through specification offers a very convenient framework in examining the relationship between the ERPT and exchange rate appreciations/depreciations and the ERPT and inflation rates.

Findings show that a complete pass-through regime is reached for large enough appreciations of currencies and an incomplete pass-through regime exists for small amounts of appreciations and depreciations. Similarly for inflation rates that exceed a threshold level, we observe complete pass-through especially in the long-run for all countries studied. On the other hand incomplete pass-through regimes occurs for inflation rates that are far below the threshold levels. Findings reveal not only the presence of complete and incomplete pass-through regimes but also of asymmetric dynamics in the relationship between import prices and exchange rate changes. For example the degree of pass-through to import prices tends to be higher during appreciation periods than during depreciation periods of a currency. Similarly the degree of pass-through and convergence towards complete pass-through occurs during high inflation periods than low inflation or disinflation periods. Our analysis also shows that since exchange rate appreciations and inflation rates stayed mostly under the threshold levels, ERPT estimates tended to be far below one frequently during the sampling period for most of the countries analyzed. The reported ERPT variation over past currency and inflationary regimes may reconcile the findings from the linear pass-through regressions reported in previous studies.

Our findings may have important implications for macroeconomic policy. The presence of complete and incomplete pass-through regimes where regimes are characterized (endogenously) by the past exchange rate appreciations/depreciations and inflation supports the arguments raised by Taylor (2000) in that ERPT may not be exogenous to macroeconomic factors and to the environment. Estimation results also reveal considerably stable pass-through dynamics over complete and incomplete pass-through regimes regardless of whether inflation rates or currency appreciations drive the regime-

dependence in the pass-through relationship. This is quite striking as it suggests that once the inflation rate or the degree of appreciation of a currency exceeds a certain threshold level, estimates tend to stay in the vicinity of complete pass-through. Therefore, policy makers should pay attention to the macroeconomic environment in terms of inflation and exchange rate appreciations when considering the impact of exchange rate movements on the import prices. Findings also imply that international transmission of shocks may also have time-variation and regime-dependence which may be worth further investigation. Moreover, firms may find it useful to know that past movements in exchange rates and inflation rates may derive the degree of pass-through to import prices as these may have implications for hedging and marketing strategies.

We should also note that the estimated transition and threshold parameters are at the aggregate level, and therefore, there may be differences for various firms in terms of threshold levels as well as the speed of transition across extreme regimes which in turn will be related to several other factors; some examples are the elasticity of demand, export market conditions, and firm specific conditions. In other words, the estimates may capture the general dynamics at the aggregate level and at the firm and industry levels there may be differences in terms of nonlinear dynamics. It is quite plausible to imagine that the decision to pass through a certain amount of appreciation to prices is made abruptly at the firm level and the amount of appreciation needed to induce such a change in the prices may differ across firms and industries. For example, at the individual firm level, different firms may respond in different degrees with different threshold levels which may merit further investigation by using the firm level data. However, at the aggregate level it is considerably intuitive to model such dynamics by a smooth transition model as different firms may make decisions at different times and at different levels of thresholds.

Generally speaking our findings also reveal some variation in the nonlinear dynamics of the ERPT relationship depending on if past exchange rate changes or inflation rate

movements are used as the transition variable. This is consistent with the idea that with the heterogeneity of firms in the export/import markets import prices may respond to exchange rate shocks at various degrees depending on if the inflation rate and/or the currency appreciation dominantly characterize the macroeconomic environment, at a given time. Despite some differences in the temporal dynamics, estimated ERPTs show considerable similarities in terms of being generally low or generally high over exchange rates and inflationary regimes. An alternative approach to the one in this paper is to estimate STP models where the transition variable is given by a combination of exchange rates, inflation rates and possibly some other macroeconomic variables. Moreover, one can also develop an indicator variable based on several economic factors which might prove useful in characterizing the ERPT. The advantage of using variables in isolation, as is done in this paper, is that it allows us to examine nonlinear dynamics where the transition between incomplete and complete regimes, if they exist with respect to the specific transition variable used, is to be investigated thoroughly. This may also be desirable for policy purposes, at a general level, as the analysis provided may foster insights on the relationship between ERPT and currency and inflationary states.

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Table 1: Estimated short-run and long-run pass-through elasticities from the linear model

Country	Elasticity		Linearity Tests			
	Short Run	Long Run	$pL(s)$	$pL(\pi)$	$pE(s)$	$pE(\pi)$
USA	0.260*†	0.417*†	0.013	0.014	0.113	0.089
UK	0.324*†	0.457*†	0.011	0.008	0.105	0.118
Japan	0.695*†	0.905*	0.016	0.017	0.091	0.077
Canada	0.622*†	0.592*†	0.010	0.009	0.128	0.130
Australia	0.620*†	0.591*†	0.003	0.064	0.044	0.031
Germany	0.568*†	0.911*	0.095	0.017	0.150	0.053

Key:* (†) implies that an elasticity is significantly different from 0 (1) at 5% level. $pL(s)$ and $pL(\pi)$ are the maximal p-values for testing linearity in the residuals of linear pass through regression model against the alternative of LSTP over $z_t = \{\Delta s_{t-d}, \frac{1}{d} \sum_{i=2}^d \pi_i\}$ and $z_t = \{\pi_{t-d}, \frac{1}{d} \sum_{i=2}^d \pi_i\}$ respectively for delay parameters $d \in \{1, 2, \dots, \bar{d}\}$. Similarly $pE(s)$ and $pE(\pi)$ are the maximal p-values for testing the linearity in the residuals of the linear pass through regression model against the alternative of the ESTP model.

Table 2: Logistic Smooth Transition Pass-through Regression Results: $\Delta p_t = \alpha + [\sum_{i=0}^4 \beta_i \Delta s_{t-i}] (1 - F(\gamma, c, z_{t-d})) + [\sum_{i=0}^4 \beta_i^* \Delta s_{t-i}] F(\gamma, c, z_{t-d}) + [\sum_{i=0}^4 \omega_i \Delta w_{t-i}] + \psi \Delta gdp_t + u_t$

	USA		UK		Japan		Canada		Germany		Australia	
z_t	Δs_{t-4}	π_{t-1}	$\frac{1}{3} \sum_{i=1}^3 \Delta s_{t-i}$	π_{t-1}	Δs_{t-1}	π_{t-3}	Δs_{t-4}	π_{t-3}	Δs_{t-4}	π_{t-1}	Δs_{t-2}	π_{t-1}
β_0	0.712 (0.154)	0.154 (0.051)	0.415 (0.050)	0.113 (0.058)	0.975 (0.164)	0.458 (0.062)	1.243 (0.271)	0.674 (0.136)	0.702 (0.102)	0.426 (0.131)	0.746 (0.055)	0.599 (0.044)
pt_1^L	0.449	0.000	0.000	0.000	0.904	0.000	0.814	0.009	0.002	0.001	0.000	0.000
β_0^*	0.147 (0.062)	0.639 (0.140)	-0.273 (0.343)	0.370 (0.101)	0.686 (0.113)	1.019 (0.142)	0.582 (0.053)	0.583 (0.049)	0.171 (0.193)	0.687 (0.144)	0.528 (0.050)	0.720 (0.041)
pt_1^U	0.000	0.033	0.000	0.000	0.003	0.552	0.000	0.000	0.000	0.016	0.000	0.000
pW_{SR}	0.000	0.000	0.000	0.001	0.017	0.040	0.014	0.527	0.014	0.180	0.003	0.063
$\sum_{i=0}^4 \beta_i$	1.158 (0.283)	0.453 (0.115)	0.672 (0.126)	0.349 (0.200)	1.275 (0.288)	0.284 (0.175)	1.258 (0.329)	-0.022 (0.322)	0.895 (0.213)	0.528 (0.206)	0.658 (0.096)	0.569 (0.073)
pW_0^L	0.000	0.001	0.000	0.080	0.000	0.105	0.000	0.917	0.000	0.001	0.000	0.000
pW_1^L	0.449	0.000	0.010	0.000	0.194	0.000	0.466	0.000	0.632	0.005	0.001	0.000
$\sum_{i=0}^4 \beta_i^*$	0.299 (0.124)	1.018 (0.288)	-0.157 (0.201)	0.557 (0.120)	0.378 (0.130)	1.365 (0.185)	0.378 (0.155)	0.753 (0.163)	0.615 (0.251)	1.246 (0.234)	0.484 (0.099)	0.729 (0.128)
pW_0^U	0.035	0.000	0.674	0.000	0.013	0.000	0.015	0.000	0.014	0.000	0.000	0.000
pW_1^U	0.000	0.941	0.017	0.097	0.000	0.174	0.000	0.130	0.122	0.356	0.000	0.003
pW_{LR}	0.001	0.000	0.000	0.000	0.000	0.000	0.020	0.002	0.416	0.018	0.228	0.203
γ	2.723 (0.874) [0.014]	4.252 (2.859) [0.042]	6.632 (1.552) [0.021]	3.647 (1.580) [0.056]	3.239 (0.365) [0.018]	3.079 (0.676) [0.007]	3.922 (1.456) [0.026]	3.216 (1.337) [0.052]	7.224 (3.308) [0.083]	7.422 (3.060) [0.087]	8.380 (1.253) [0.049]	7.860 (1.190) [0.042]
pW_1	0.043	0.057	0.071	0.060	0.058	0.049	0.073	0.074	0.121	0.091	0.080	0.017
pW_3	0.007	0.011	0.008	0.002	0.004	0.009	0.005	0.008	0.064	0.073	0.055	0.015
c	-4.366 (0.361)	4.413 (0.119)	-3.036 (0.437)	4.025 (0.356)	-6.317 (0.179)	0.504 (0.064)	-3.607 (0.208)	0.363 (0.086)	-2.422 (0.428)	1.361 (0.337)	-1.993 (1.400)	2.502 (0.161)
SSR_{ratio}	0.818	0.844	0.827	0.906	0.594	0.770	0.828	0.837	0.906	0.919	0.852	0.815
pLM_{SC}	0.244	0.223	0.141	0.185	0.326	0.238	0.156	0.239	0.187	0.109	0.279	0.729
pLM_C	0.214	0.461	0.542	0.484	0.228	0.378	0.685	0.835	0.232	0.251	0.327	0.866
$pLM_{L,s}$	0.373	0.456	0.229	0.316	0.601	0.398	0.524	0.725	0.226	0.169	0.346	0.658
$pLM_{L,\pi}$	0.351	0.433	0.269	0.282	0.210	0.401	0.126	0.154	0.227	0.367	0.318	0.236
$pLM_{E,s}$	0.361	0.698	0.603	0.540	0.515	0.286	0.407	0.848	0.299	0.600	0.577	0.654
$pLM_{E,\pi}$	0.624	0.471	0.576	0.611	0.708	0.503	0.326	0.312	0.301	0.564	0.331	0.297

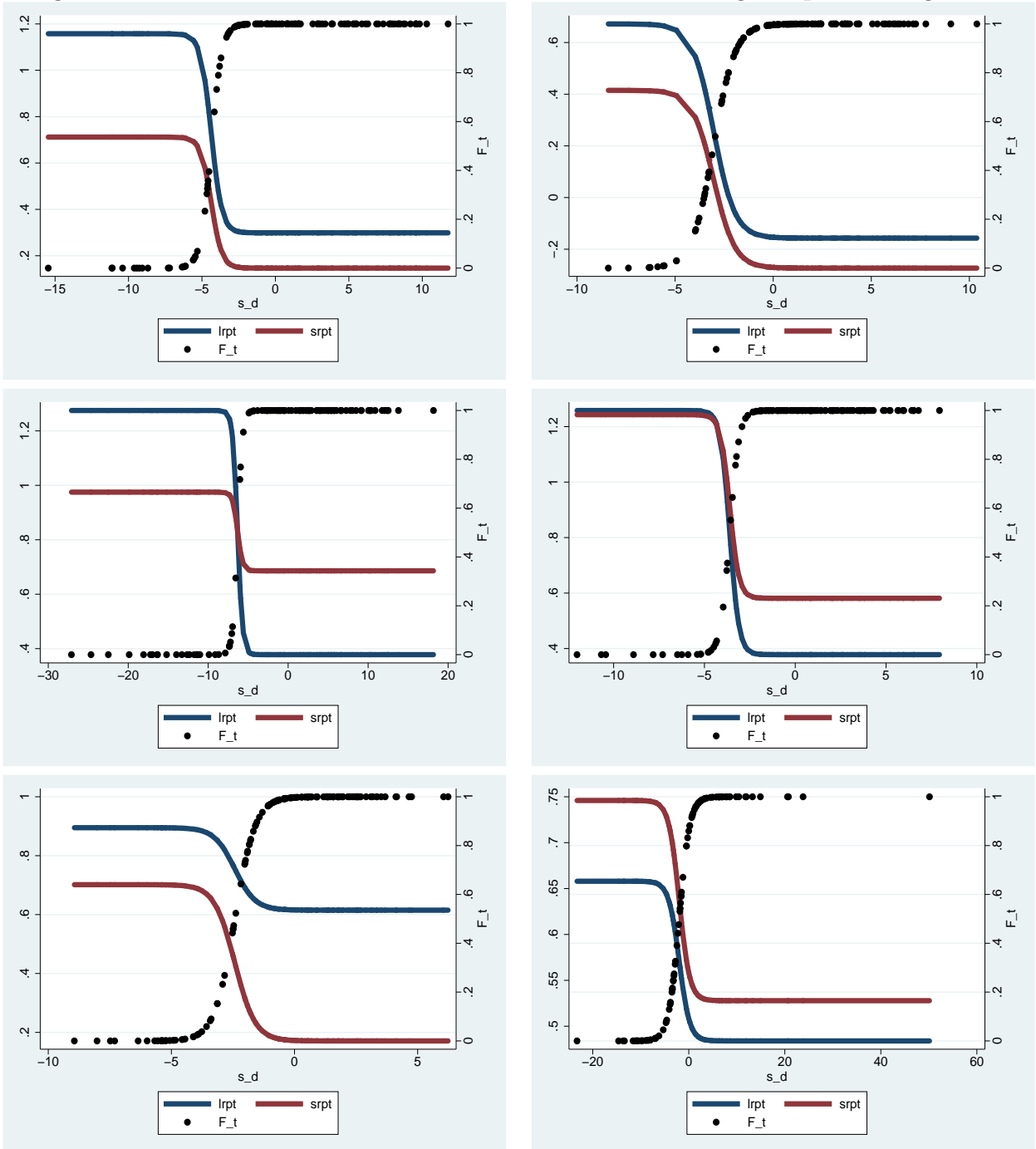
Key: Table reports elasticities of exchange rate pass-through into import prices from LSTPT model, tests, and diagnostic statistics. Values in parentheses underneath estimates are the heteroscedasticity and serial correlation robust standard errors. pt_1^L and pt_1^U are the p-values for testing short run elasticity is equal to one in lower ($F(\cdot) = 0$) and upper ($F(\cdot) = 1$) regimes respectively. pW_{SR} is the p-value for the robust Wald test for testing the null hypothesis that $\beta_0 = \beta_0^*$, (i.e., the short-run elasticities are the same across extreme regimes). pW_0^L and pW_0^U are the p-values from the robust Wald statistic for testing the null that long-run elasticities are zero for lower and upper regimes respectively. pW_1^L and pW_1^U are the p-values from the robust Wald statistic for testing the null that long-run elasticities are one for lower and upper regimes respectively. pW_{LR} is the p-value for testing the null that the long-run elasticity is the same across lower and upper regimes (i.e. $\sum_{i=0}^4 \beta_i = \sum_{i=0}^4 \beta_i^*$). SSR_{ratio} is the ratio of sum of squared residuals of linear pass-through regression to LSPT regression, pLM_{SC} and pLM_C are the p-values for testing no serial correlation in the residuals up to order four and parameter constancy of estimated models. pW_1 and pW_3 are the p-values for robust Wald tests based on first and third order Taylor series expansion of the LSPT model around $\gamma = 0$. Rows corresponding to pLM_L and pLM_E are the p-values corresponding to the maximum LM test statistic for no remaining nonlinearity of exponential and logistic form in the residuals of the estimated LSPT models for each given transition variable with delay parameter in the range 14. Note that we compute LM tests for $z_t = \{\Delta_{t-d}, \pi_{t-d}\}$ and their averages over $d = 2, 3, 4$ and report the p-value that corresponds to the largest LM statistic over these transition variables.

Table 3: ERPT Regressions estimated for different regimes identified from LSTP models

Country	Regime	$\sum_{i=0}^4 \beta_i$	$p - t_{\sum_{i=0}^4 \beta_i=0}$	$p - F_{\sum_{i=0}^4 \beta_i=1}$	95% CI	R^2	n
Panel A: Regimes identified by LSTP model with past exchange rate changes as the transition variable							
USA	Complete	0.809(0.402)	0.032	0.697	[0.221, 1.839]	0.790	22
	Incomplete	0.274(0.138)	0.025	0.000	[-0.001, 0.549]	0.159	103
UK	Complete	1.199(0.754)	0.064	0.287	[-0.401, 2.802]	0.531	16
	Incomplete	0.696(0.121)	0.000	0.014	[0.456, 0.936]	0.587	103
Japan	Complete	2.293(0.595)	0.001	0.041	[1.056, 3.531]	0.658	29
	Incomplete	0.614(0.164)	0.000	0.021	[0.288, 0.940]	0.664	85
Germany	Complete	0.274(0.352)	0.222	0.051	[-0.456, 1.004]	0.555	30
	Incomplete	0.614(0.164)	0.000	0.021	[0.288, 0.940]	0.664	85
Canada	Complete	0.493(0.279)	0.053	0.097	[-0.009, 1.017]	0.866	19
	Incomplete	0.379(0.240)	0.059	0.011	[-0.097, 0.855]	0.398	109
Australia	Complete	0.622(0.159)	0.000	0.074	[0.267, 0.953]	0.882	34
	Incomplete	0.454(0.103)	0.000	0.000	[0.139, 0.647]	0.781	50
Panel B: Regimes identified by LSTP model with past inflation rates as the transition variable							
USA	Complete	0.145(0.769)	0.427	0.298	[-1.629, 1.918]	0.733	15
	Incomplete	0.455(0.128)	0.000	0.000	[0.201, 0.709]	0.243	116
UK	Complete	0.632(0.151)	0.003	0.051	[0.263, 1.002]	0.909	14
	Incomplete	0.458(0.090)	0.000	0.000	[0.280, 0.636]	0.403	106
Japan	Complete	0.727(0.385)	0.037	0.487	[-0.080, 1.533]	0.485	28
	Complete	0.536(0.193)	0.004	0.020	[0.147, 0.924]	0.554	58
Germany	Regime C	0.470(0.276)	0.050	0.066	[-0.096, 1.037]	0.514	34
	Regime I	0.409(0.198)	0.002	0.004	[0.013, 0.804]	0.197	82
Canada	Regime C	0.530(0.165)	0.002	0.006	[0.201, 0.859]	0.403	76
	Regime I	-0.008(0.244)	0.976	0.003	[-0.559, 0.544]	0.840	17
Australia	Regime C	0.732(0.107)	0.000	0.024	[0.505, 0.959]	0.940	23
	Regime I	0.541(0.088)	0.000	0.000	[0.367, 0.716]	0.768	73

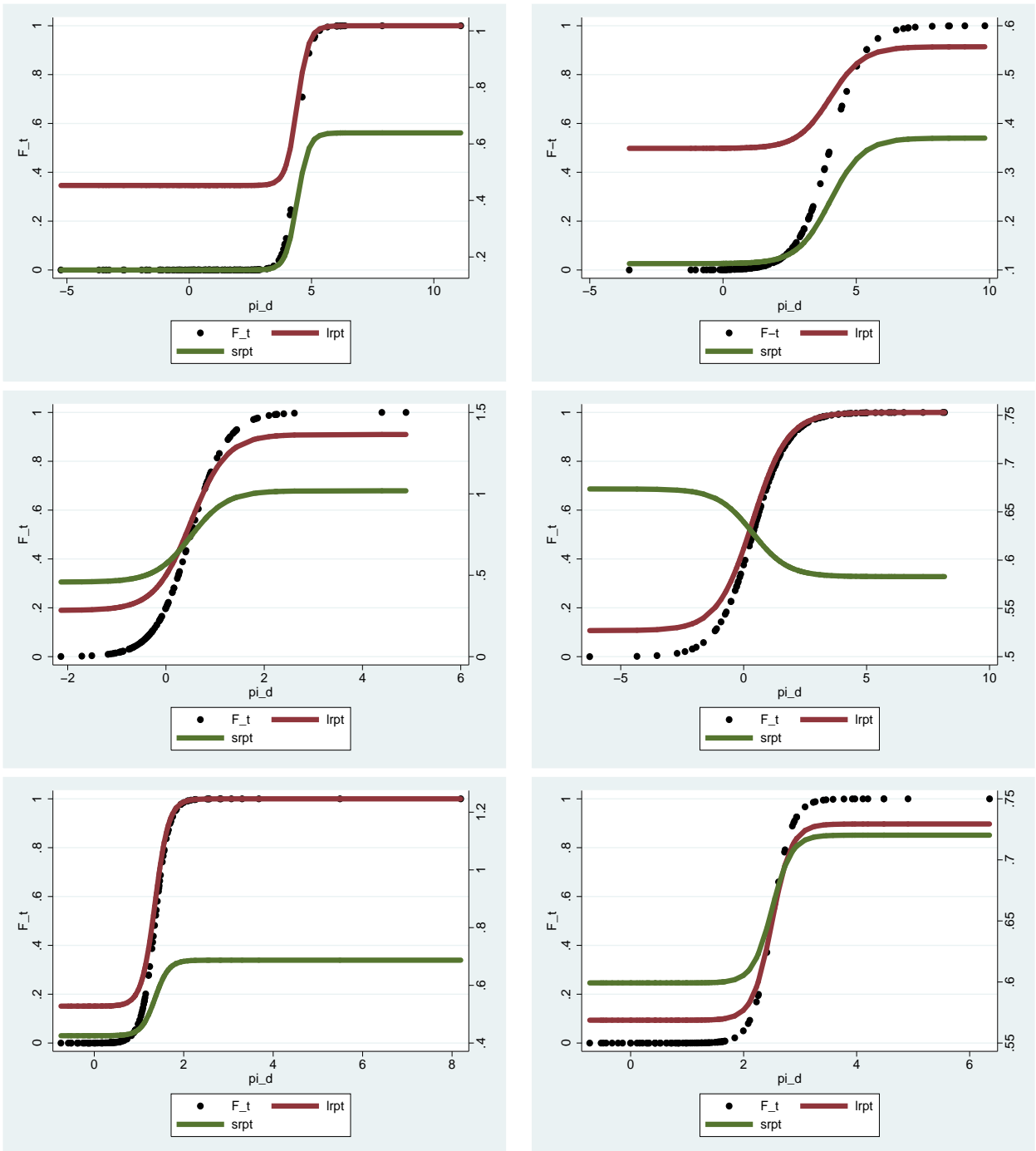
Key: Table reports the long-run estimates from the linear ERPT regressions over extreme regimes where regimes are identified from the estimated LSTP models with transition variables given by d -lagged or past d -quarter moving averages of exchange rate changes and inflation rates as reported in Table 2. The values in parentheses next to the long run elasticity estimate are the standard errors. Columns corresponding to $p - t_{\sum_{i=0}^4 \beta_i=0}$ give the p-values for testing the null that the elasticity is zero against the one-sided alternative that it is greater than zero. Similarly the columns underneath $p - F_{\sum_{i=0}^4 \beta_i=1}$ are the p-values for testing the null that the elasticity is unity against the one sided alternative that it is less than zero. 95%CI is the estimated 95% confidence interval for the long run elasticity, R^2 is the regression R-square and n is the sample size in each identified regime.

Figure 1: Estimated transition function and short-run and long-run pass-through



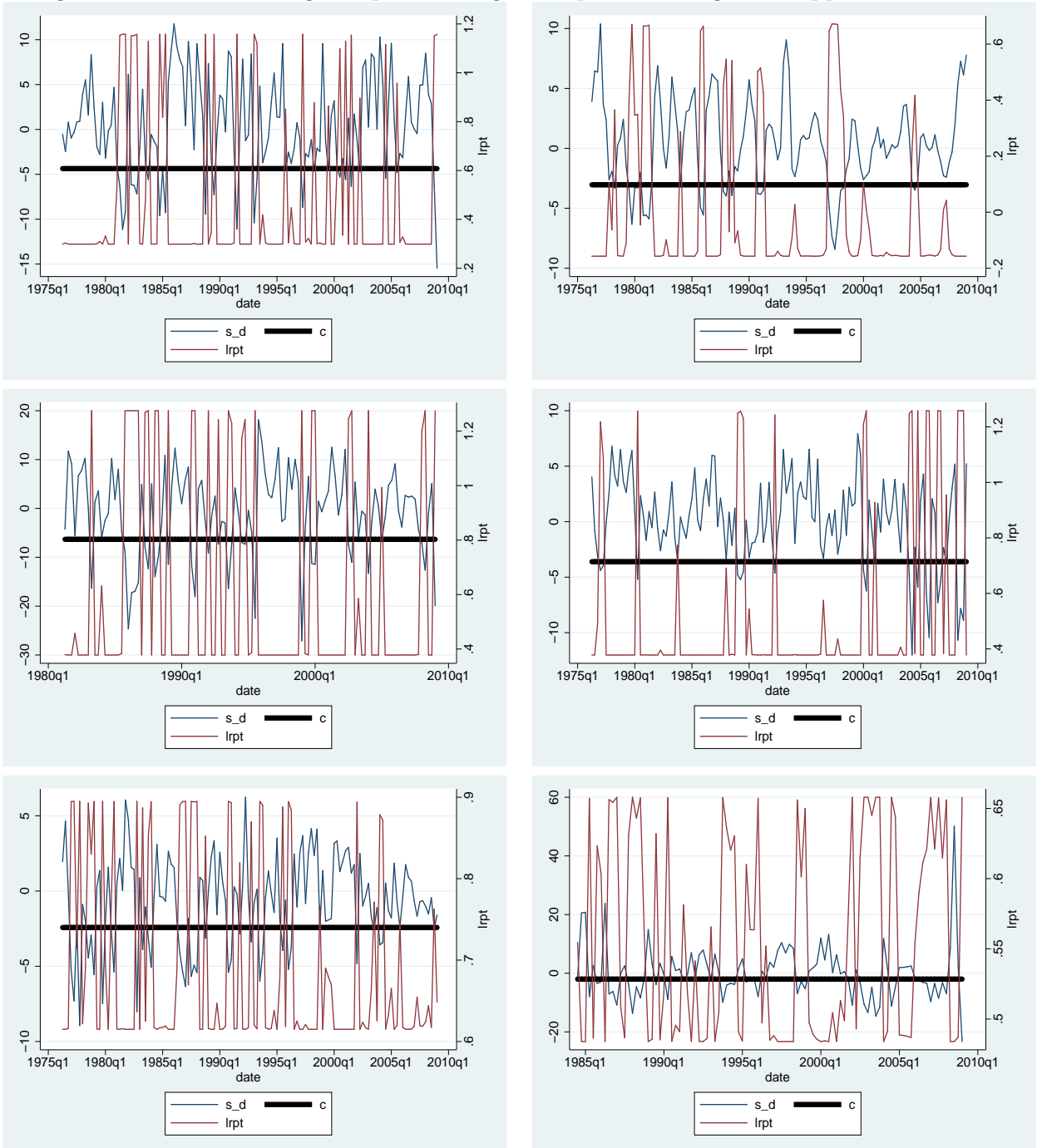
Key: The figure plots the estimated transition functions, short-run and long-run pass-through over past depreciations for the US, the UK, Japan, Canada, Germany, and Australia respectively.

Figure 2: Estimated transition function and short-run and long-run pass-through over past inflation rates



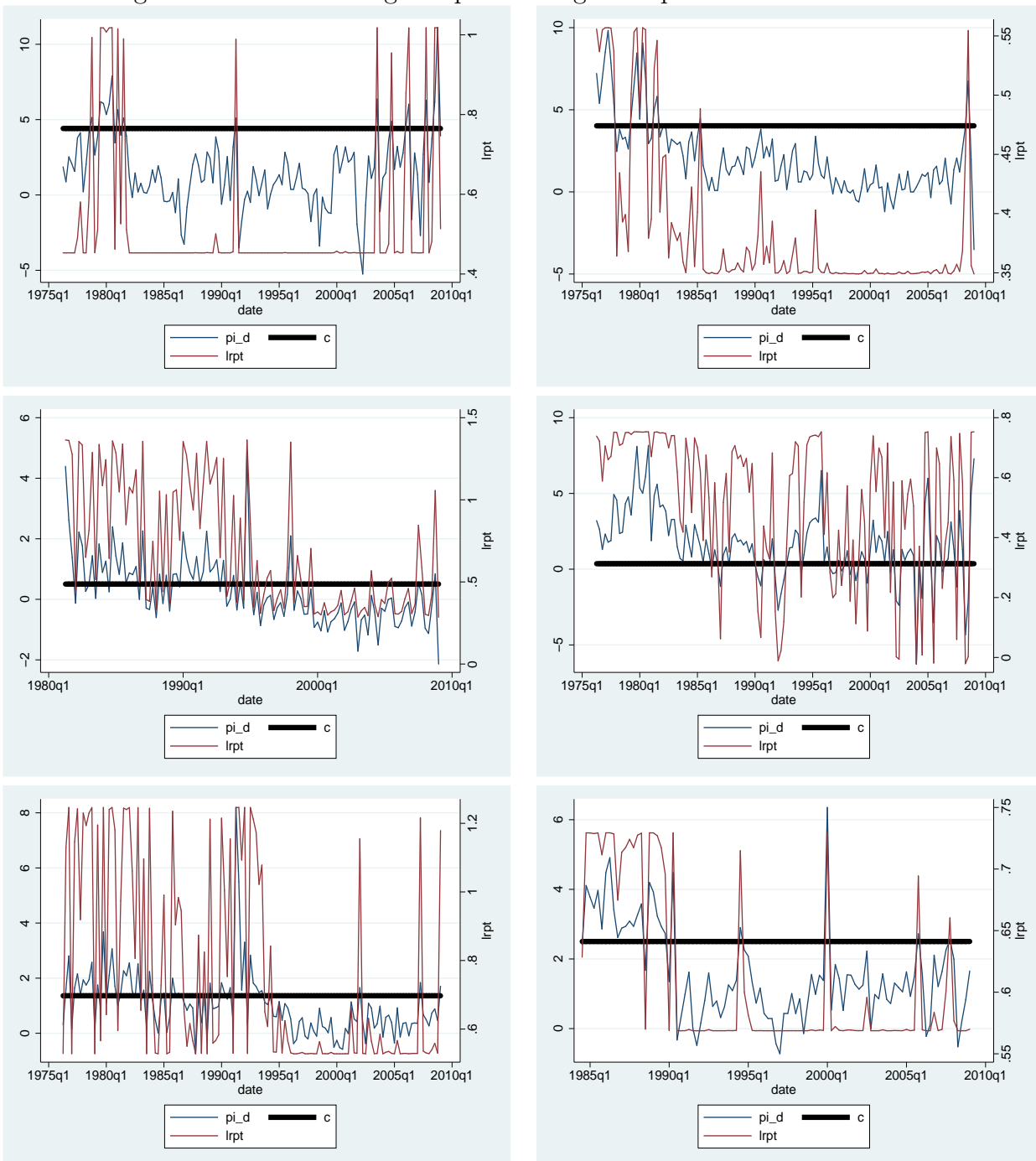
Key: The figure plots the estimated transition function, short and long-run pass-through over past inflation rates for the US, the UK, Japan, Canada, Germany, and Australia respectively.

Figure 3: Estimated long-run pass-through and past exchange rate appreciations



Key: The figure plots the estimated long-run pass-through and past appreciations of the US Dollar, the UK Pound, Japanese Yen, Canadian Dollar, German Mark, and Australian Dollar respectively.

Figure 4: Estimated long-run pass-through and past inflation rates



Key: The figure plots the estimated long-run pass-through and past inflation rates for the US, the UK, Japan, Canada, Germany, and Australia respectively.