The Pass-Through of Exchange Rate in the Context of the European Sovereign Debt Crisis

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Abstract

This paper investigates whether the exchange rate pass-through (ERPT) to CPI inflation is a nonlinear phenomenon for five heavily indebted euro area (EA) countries, namely the so-called GIIPS group (Greece, Ireland, Italy, Portugal, and Spain). Using logistic smooth transition models, we explore the existence of nonlinearity with respect to sovereign bond yield spreads (versus German) as an indicator of confidence crisis/macroeconomic instability. Our results provide strong evidence that the extent of ERPT is higher in periods of macroeconomic distress, i.e. when sovereign bond yield spreads exceed some threshold. For all the GIIPS countries, we reveal that the increasing of macroeconomic instability and the loss of confidence during the recent sovereign debt crisis has entailed a higher sensibility of CPI inflation to exchange rate movements.

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1 Introduction

The financial and economic crisis of 2007-2008 has had a serious impact the euro area (EA) government bond market and turned into a sovereign debt crisis by the beginning of 2010. Due to the general weakness of fiscal fundamentals in the so-called GIIPS countries, i.e. Greece, Italy, Ireland, Portugal and Spain, financial markets have been highly affected by deeply felt concerns on the solvency of this group of countries. There was a change in the markets' assessment of sovereign debt risks which have caused yield spreads to German bonds to rebound to levels exceeding those observed in the early years of the third stage of Economic and Monetary Union (EMU). As a matter of fact, the introduction of the single currency in 1999 has eliminated the normal market reactions towards highly indebted EA countries and there has been a phase of pronounced government bond yield convergence as the euro is regarded as a safe haven. Also, the improvement in general macroeconomic environment brought back by the formation of the monetary union has changed the general risk perception towards EA economies with high debt ratios, narrowing interest rate differentials significantly.

In this paper, we test whether this changing in macroeconomic conditions, caused by the crisis of confidence on sovereign debt, could influence the extent of exchange rate pass-through (ERPT). The exporter's decision on how much to pass-through exchange rate movements into prices may depend on the perceptions about the importing country's macroeconomic stability. When the economy faces a financial or a confidence crisis, foreign firms may decide to pass-through a larger proportion of exchange rate changes in view of the increased likelihood of default from the importer. In this case, the ERPT would be higher since exporters tend to set prices in their own currency (*producer-currency pricing* or PCP strategy). However, in periods of good macroeconomic conditions, exporting firms would absorb currency fluctuations within markup by setting prices in the currency of the stable importing country (*local-currency pricing* or LCP setting). Consequently, ERPT is expected to be higher in times of confidence crisis than during macroeconomic stability episodes. As the European sovereign debt crisis has spilled over into a strongly deteriorating macroeconomic environment, we investigate whether it has entailed a changing behaviour in the rates of pass-through.

Several studies has put forth the role of macroeconomic environment in determining the degree to which currency changes are transmitted into domestic prices (TAYLOR, 2000; CHOUDHRI and HAKURA, 2006; GAGNON and IHRIG, 2004, to name but a few). A common drawback of most of this literature is to assume a linear connection between pass-through and macroeconomic factors (such as inflation environment or credibility levels of monetary policy) rather than testing it. However, as pointed out by BUSSIÈRE (2012), there are several sources of nonlinearities in the ERPT mechanism, and the relationship between macroeconomic variables and pass-through can be potentially nonlinear. In spite of its policy relevance, studies dealing with nonlinearities in pass-through mechanism are still relatively scarce. Most of the few empirical works dealing with nonlinearities have used a nonlinear threshold models where the transition across regimes is abrupt (see e.g. CORREA and MINELLA, 2006). Nevertheless, a threshold framework would be more appropriate at the microeconomic level where a single foreign firm is setting prices. In fact, an exporter can change pricing behaviour sharply with respect to macroeconomic conditions in the importing country. While, at the aggregate level, there is a great diversity across firms when forming opinions about the macroeconomic environment, hence assuming abrupt transition from one regime to the other is unrealistic. Therefore, the transition across regimes is rather gradual since there is some heterogeneity across firms in their attitude towards the state of the importer's macroeconomic environment. To overcome this shortcoming, a very recent empirical literature proposes to use another class of nonlinear regime-switching models, namely the smooth transition regression (STR) model where the transition between states is rather smooth.

To the best of our knowledge, there are very few empirical studies testing for nonlinearities in ERPT using a smooth nonlinear regression. SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013) estimated the ERPT to US domestic prices with respect to inflation regime. They found that periods of low ERPT would be associated with low inflation levels. In a similar vein, BEN CHEIKH (2012) has investigated for the presence of nonlinear mechanism in pass-through for the EA case. The author found a strong evidence of nonlinearities with respect to inflation environment in 8 out of 12 EA countries, that is, when inflation rate surpasses some threshold, the transmission of exchange rate becomes higher in some European countries. For the Mexican case, NOGUEIRA JR. and LEON-LEDESMA (2011) examine the possibility of nonlinear pass-through with respect to macroeconomic instability. The authors conclude that under bad macroeconomic conditions, as in periods of financial or confidence crisis, ERPT to consumer prices will be higher than in periods of macroeconomic stability. As a measure of macroeconomic instability, NOGUEIRA JR. and LEON-LEDESMA (2011) used the real interest rate differential of Mexico with respect to the U.S., which corresponds to the transition variable in their smooth transition model.[1]

In this paper, we follow NOGUEIRA JR. and LEON-LEDESMA (2011) approach by using a STR framework to estimate the nonlinear behaviour of ERPT with respect to macroeconomic instability. Unlike the authors, we focus on the five GIIPS EA countries due to the recent context of the European sovereign debt crisis. The perception that some EA countries are reaching an unsustainable fiscal

[1] HERZBERG, KAPETANIOS, and PRICE (2003) have analyzed the ERPT into UK import prices using a STR model but did not find any evidence of nonlinearity.

situation means a worsening of the macroeconomic fundamentals. Therefore, we propose that exporting firms may tend to change their pricing behaviour following the widening of sovereign yield spreads to German bonds as an indicator of deteriorating macroeconomic environment. Thus, the advantage of the STR models is to capture this changing behaviour in a nonlinear fashion. To our knowledge, there is no other study has applied a nonlinear STR estimation approach in this context.

The remainder of the paper is organized as follows. Section 2 gives some arguments that justify the existence of nonlinear ERPT with respect to macroeconomic environment and discusses the analytical framework that underlies this dynamic behavior. The empirical approach is provided in the Section 3. Section 4 gives the main empirical results and Section 5 concludes.

2 Pass-Through and Nonlinearities

2.1 ERPT in time of crisis

In the recent years, there has been a revival of interest in the role of macroeconomic factors as important determinants of the extent of pass-through. This strand of literature highlights the role of a stable macroeconomic environment and in particular the shift towards credible monetary policy regime in explaining the observed decline in the degree of ERPT. One of the first to put forward this argument was TAYLOR (2000) who stipulates that countries with low relative inflation variability or stable monetary policies are more likely to have their currencies chosen for transaction invoicing, and hence more likely to have low pass-through to domestic prices. In other words, a stable inflation environment in the importing country may lead exporters to adopt *local-currency pricing* (LCP) strategy. Firms can absorb currency changes within markup, leading to lesser extent of pass-through. By contrast, when the importer experience higher inflation levels, exporter would change their pricing decision by adopting *producer-currency pricing* (PCP) strategy, by transmitting exchange rate variations to the price in the importer currency. Drawing on this intuition, one can think that the degree of pass-through may depend on the importing country's general macroeconomic conditions.

Using nonlinear STR framework, NOGUEIRA JR. and LEON-LEDESMA (2011) argued that the exporter's decision on how much to pass-through exchange rate movements into prices depends on the perceptions about the importing country's macroeconomic stability. When the economy faces a financial or a confidence crisis, foreign firms may decide to pass-through a larger proportion of its cost changes in view of the increased likelihood of default from the importer. However, in periods of good macroeconomic conditions, prices will become more insulated from exchange rate changes since foreign firms are willing to adopt LCP setting. The authors conclude that under bad economic conditions ERPT to consumer prices will be higher than in periods of macroeconomic stability for the Mexican case. As a measure of macroeconomic instability, NOGUEIRA JR. and LEON-LEDESMA (2011) used the real interest rate differential of Mexico with respect to the U.S., which corresponds to the transition variable in their smooth transition model.

In line with this intuition, one can think that the recent sovereign debt crisis experienced by some EA countries would influence the extent to which the exchange rate changes are transmitted to prices. In fact, the start of the economic and financial crisis in summer 2007 and its intensification in 2008 (in the aftermath of the collapse of Lehman Brothers), has had a serious impact on the EA government bond market and marked the beginning of financial stress for the so-called GIIPS countries, i.e. Greece, Ireland, Italy, Portugal, and Spain.[2] As shown in Figure 1, after ten years of stability at very low levels, the long-term government bond yields relative to the German Bund have started rising since the beginning of 2010. Due to the unsatisfactory performance of the GIIPS countries group, the spreads was well above those of emerging market countries, such as South Korea and Brazil.

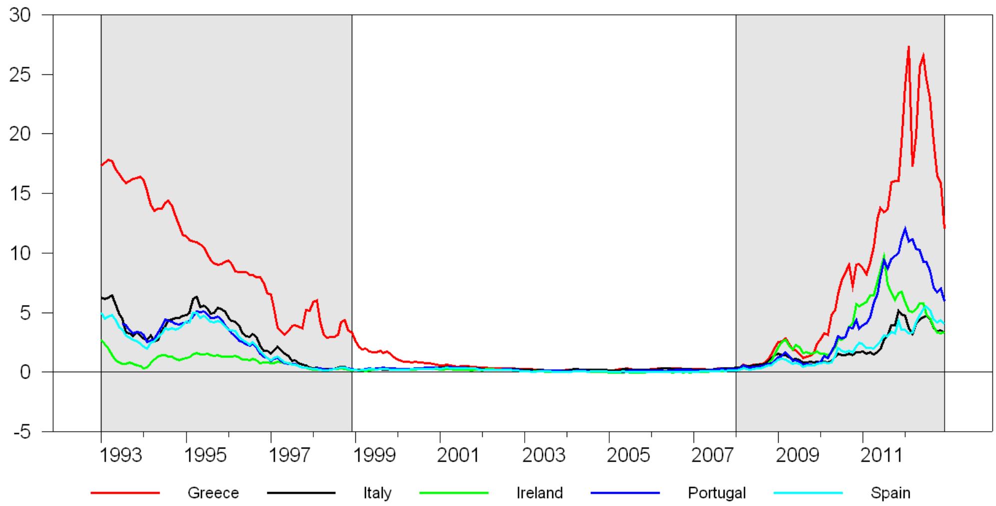
In accordance with the above arguments, we propose that the exporting firms may change their attitude during the recent context of European sovereign crisis due to the general weakness of macroeconomic fundamentals in the GIIPS group. With the perception of relatively high sovereign debt default risk in the fiscally distressed EA countries, firms will have no incentive to absorb the single currency fluctuations within their margins. After a period of stable macroeconomic environment, where the interest rate spreads of EA 10-year government bonds against the German benchmark have declined dramatically (see Figure 1), the fiscal vulnerabilities and the risk of default heightened since 2010 has widened the sovereign bond yield spreads to levels exceeding those observed in the early years of the monetary union. As a result, with weaker fiscal fundamentals in the importing country, exporters may tend to modify pricing strategy, shifting from importer's currency pricing (LCP strategy) to exporter's currency invoicing (PCP strategy), leading to higher rates of ERPT.[3] Therefore, in line with the view of NOGUEIRA JR. and LEON-LEDESMA (2011), we expect that for the GIIPS group the sensibility of consumer

[2] These Eurozone members are called "peripheral " EA countries. Since the European sovereign debt crisis, the term "GIIPS" is used to refer to this group of countries as a label for heavily-indebted economies.

[3] As mentioned in the literature, the introduction of the euro has entailed a decline in the extent of pass-through for the most of the EA members, since the process of monetary union has entailed some convergence towards more stable macroeconomic conditions. Foreign firms are choosing the euro as the currency of denomination of their exports (LCP strategy) and the European prices has become more insulated from exchange rate variations (see e.g. DEVEREUX, ENGEL, and TILLE, 2003, among others).

prices will be higher during periods of macroeconomic instability or confidence crisis such as the recent sovereign debt crisis. As an indicator of macroeconomic instability, we use the 10-year government bond yield spreads to the German Bund.

Figure 1: Spreads of 10-year government benchmark bonds to German Bund



Source: European Central Bank.

2.2 Theoretical Considerations

The pricing behavior of a representative foreign exporting firm is a useful starting point for understanding the dynamic of the pass through of exchange rate changes into prices. As commonly in the literature, we consider a simple static profit-maximization problem faced by the foreign firm. The latter exports its product and has some degree of pricing power in the domestic importing country. Under monopolistic competition, the first-order conditions for exporter profit maximization yield the following expression at time t:

$$P_t = E_t \mu_t W_t^* \tag{1}$$

where P_t is the price of the good (denominated in the importing country currency), E_t is the exchange rate (in units of the importer currency per unit of the foreign currency), μ_t is the markup of price over exporter marginal cost W_t^*. The markup is defined as $\mu_t \equiv \eta_t/(1-\eta_t)$, where η_t is the price elasticity of demand in the importing country. As in BAILLIU and FUJII (2004), μ_t is assumed to depend essentially on demand pressures in the importing market: $\mu_t = \mu(Y_t)$, with Y_t is the income (expenditures) level in the importing country.

The log-linear form of equation (1) gives the standard pass-through regression traditionally tested throughout the literature (see e.g. GOLDBERG and KNETTER, 1997; CAMPA and GOLDBERG, 2005):

$$p_t = \alpha + \beta e_t + \psi y_t + \delta w_t^* + \varepsilon_t, \tag{2}$$

where lowercase letters denote logarithmic values of the variables.

From equation (2), the ERPT coefficient is given by coefficient β and is expected to be bounded between 0 and 1. If $\beta = 1$, exporter markup will not respond to fluctuations of the exchange rates, hence price is set in foreign country currency (PCP) and pass-through is complete. If $\beta = 0$, the ERPT is zero since foreign firm decide not to vary the prices in the destination country currency and absorb the fluctuations within the markup. This is a purely LCP strategy.

In fact, pricing strategies of firms depend not solely on demand conditions in the importing market. As is well-known, foreign firm may adjust price after exchange rate movements with respect to some macroeconomic factors. Following NOGUEIRA JR. and LEON-LEDESMA (2011), we assume that foreign firms' markup may also depend on the importing country's general macroeconomic stability. The exporter's decision on how much to pass-through exchange rate movements into prices depends on its view on the importer's macroeconomic conditions. Under bad economic conditions, firms have no incentive to absorb exchange rate fluctuations in their margins which thus leads to higher ERPT (PCP setting). However, countries with a stable macroeconomic environment would have their currencies chosen for transaction invoicing. As a result, prices will become more insulated from exchange rate changes as foreign firms are willing to adopt local pricing strategy (LCP).

Following this argument, we extend the basic model (2) by assuming that the macroeconomic environment stability of importing country will determine the adjustment of exporter markup in a nonlinear way. Given that exporting firm sets prices for several periods in advance, markup responds more to exchange rate changes if the economy faces a financial or a confidence crisis. Hence, bad macroeconomic conditions tend to increase the extent of pass-through. We then consider $\kappa(M)$ as a function that depicts the nonlinear markup response to the general macroeconomic condition. This macroeconomic dependence is seen as a firms' strategic decision on how much to translate exchange rate changes given different macroeconomic scenarios in the importing country. Taking into account this factor, we can re-write exporting firm markup as follows:

$$\mu_t = \mu\left(Y_t, E_t^{\kappa(M)}\right), \tag{3}$$

We can capture the arguments of equations (1) and (3) through a log-linear regression specification to obtain the following nonlinear ERPT equation:

$$
\begin{aligned}
p_t &= \alpha + \beta e_t + \psi y_t + \kappa(M) e_t + \delta w_t^* + \varepsilon_t \\
&= \alpha + [\beta + \kappa(M)]\, e_t + \psi y_t + \delta w_t^* + \varepsilon_t,
\end{aligned}
\tag{4}
$$

According to the function $\kappa(M)$, there is an indirect channel of pass-through which depends on the general macroeconomic stability. M would actually be a measure of macroeconomic instability (or confidence crisis) and is modelled in a way that high values imply a bad macroeconomic environment. Assuming that macroeconomic instability affect firm's markup in a nonlinear way, we consequently consider that there is some threshold M^* which provide two macroeconomic regimes, namely a bad and a good macroeconomic environment.

$$
\kappa(M) = \begin{cases} 0 & \text{for } M \leq M^* \\ \phi & \text{for } M > M^* \end{cases} \tag{5}
$$

According to (4) and (5), the degree of pass-through would be different and depends on whether the measure of macroeconomic instability M is above or below a threshold level. For a small value of M, i.e. the measure of macroeconomic instability is below the threshold value $(M < M^*)$, the importing country is in period of stable macroeconomic environment and ERPT is equal to β. For a high value of M, i.e. the measure of macroeconomic instability is above some threshold $(M > M^*)$, the importing country is facing a bad macroeconomic environment and ERPT becomes equal to $(\beta + \phi)$. The equations (4) and (5) have the advantage to describe this changing behavior in pass-through mechanism in a nonlinear fashion. Thus, we expect the transmission of exchange rate changes would be higher in periods of macroeconomic distress than in periods of good macroeconomic conditions. Finally, it should be noted that the transition from one regime to the other is assumed to be smooth. At the aggregate level, there is a great diversity across firms when forming opinions about the macroeconomic environment, thus assuming smooth transition from one regime to the other is more realistic.

3 Empirical approach

3.1 Smooth transition regression models

To capture nonlinearity in the exchange rate transmission, we use a class of smooth transition regression (STR) models as a tool. A STR model is defined as follows:

$$
y_t = \beta' \mathbf{z}_t + \phi' \mathbf{z}_t G(s_t; \gamma, c) + u_t \tag{6}
$$

where $u_t \sim \text{iid}(0, \sigma^2)$, $\mathbf{z}_t = (\mathbf{w}_t^{'}, \mathbf{x}_t^{'})^{'}$ is an $((m+1) \times 1)$ vector of explanatory variables with $\mathbf{w}_t^{'} = (y_{t-1}, ..., y_{t-d})^{'}$ and $\mathbf{x}_t^{'} = (x_{1t}, ..., x_{kt})^{'}$. $\beta = (\beta_0, \beta_1, ..., \beta_m)^{'}$ and $\phi = (\phi_0, \phi_1, ..., \phi_m)^{'}$ are the parameter vectors of the linear and the nonlinear part, respectively. $G(s_t; \gamma, c)$ is the transition function bounded between 0 and 1, and depends upon the transition variable s_t, the slope parameter γ and the location parameter c.[4] The transition variable s_t is an element of $\mathbf{z}_t$, and then is assumed to be a lagged endogenous variable ($s_t = y_{t-d}$) or an exogenous variable ($s_t = x_{kt}$). A popular choice for the transition function is the logistic specification that is given by:[5]

$$G(s_t; \gamma, c) = [1 + \exp\{-\gamma(s_t - c)\}]^{-1} \tag{7}$$

where the parameter c can be interpreted as the threshold level between two extremes regimes: $G(s_t; \gamma, c) = 0$ is called the lower regime and $G(s_t; \gamma, c) = 1$ is the upper regime. Equations (6) and (7) jointly define the logistic STR (LSTR) model. In this latter, the nonlinear coefficients would take different values depending on whether the transition variable is below or above the threshold. Thus, the parameters $[\beta + \phi G(s_t; \gamma, c)]$ changes monotonically as a function of s_t from β to $(\beta + \phi)$. In this sense, as $(s_t - c) \to -\infty$, $G(s_t; \gamma, c) \to 0$ and coefficients correspond to β; if $(s_t - c) \to +\infty$, then $G(s_t; \gamma, c) \to 1$ and coefficients become $(\beta + \phi)$; and if $s_t = c$, then $G(s_t; \gamma, c) = 1/2$ and coefficients will be $(\beta + \phi/2)$.[6] LSTR model is pertinent in describing asymmetric dynamic behavior between negative or positive deviations of the transition variable s_t from the threshold level c. As mentioned in the STR literature (see e.g. VAN DIJK, TERÄSVIRTA, and FRANSES, 2002), when modeling business cycle, LSTR can describe processes whose dynamic properties are different in expansions from what they are in recessions. For example, if the transition variable s_t is a business cycle indicator (such as output growth), and if $c \simeq 0$, the model distinguishes between periods of positive and negative growth, that is, between expansions and contractions.

As discussed in TERÄSVIRTA (1994), the modelling strategy of STR models is consisting of three stages: specification, estimation, and evaluation. The first stage consists in testing for nonlinearity and choosing the appropriate threshold variable s_t and the most suitable form of the transition function, i.e. logistic or exponential specification.[7] In the second stage, the parameters of the STR model are estimated

[4] The parameter γ is also called the speed of transition which determines the smoothness of the switching from one regime to the other.
[5] An alternative transition function which is often used in the literature is the exponential specification: $G(s_t; \gamma, c) = 1 - \exp\{-\gamma(s_t - c)^2\}$.
[6] It should be noted that LSTR model would follow the same pattern as the threshold model described in the theoretical model (equation (5)) assuming a smooth adjustment across regimes.
[7] More details for linearity tests in Appendix A.

by nonlinear least squares (NLS) estimation technique which provides estimators that are consistent and asymptotically normal. Finding good starting values is crucial in this procedure. Thus, STR literature suggests to construct a grid search for estimating γ and c. The values for the grid search for γ were set between 0 and 100 for increments of 1, whereas c was estimated for all the ranked values of the transition variable s_t. For each value of γ and c the residual sum of squares is computed. The values that correspond to the minimum of that sum are taken as starting values into the NLS procedure. This procedure increases the precision of the estimates and ensures faster convergence of the NLS algorithm.[8] In the final stage, evaluation stage, the quality of the estimated STR model should be checked against misspecification as in the case of linear models. Several misspecification tests are used in the STR literature, such as LM test of no error autocorrelation, LM-type test of no ARCH and Jarque-Bera normality test. Also, EITRHEIM and TERÄSVIRTA (1996) suggested two additional LM-type misspecification tests, namely an LM test of no remaining nonlinearity and LM-type test of parameter constancy.

3.2 Model specification and data

In this section, we try to define a *LSTR pass-through equation* that enables us to test the presence of a nonlinear ERPT mechanism with respect to macroeconomic instability. It is important to note that the theoretical model (4) presented above is designed to the ERPT to "import prices", while our paper instead examines the transmission of exchange rate changes to "consumer prices", an issue which is most relevant for monetary policy in the EA. In this case, as recommended in the literature, typical pass-through equation must be adjusted in order to have all the elements of a backward-looking Phillips curve. Mainly, two issues are considered here: first, the inertial behavior of inflation by including lags of inflation (π_{t-j}) as explanatory variables in the empirical specification (backward-looking inflation). Second, a proxy for changes in domestic demand conditions by using the changes in real GDP (Δy_t) to capture this effect.[9]

[8] It should also be noted that when constructing the grid, γ is not a scale-free. The transition parameter γ is therefore standardized by dividing it by the sample standard deviation of the transition variable s_t.

[9] Also, we can use the output gap computed as the difference between actual and potential output (constructed with a Hodrick-Prescott filter) instead of real output growth. This does not alter the estimates of pass-through.

Once these two elements have been considered, our LSTR pass-through equation can be described as a nonlinear backward-looking Phillips curve as follows:

$$\begin{aligned}\pi_t = \alpha + \sum_{j=1}^{N}\lambda_j\pi_{t-j} + \sum_{j=0}^{N}\psi_j\Delta y_{t-j} + \sum_{j=0}^{N}\delta_j\Delta w^*_{t-j} \\ + \sum_{j=0}^{N}\beta_j\Delta e_{t-j} + \left(\sum_{j=0}^{N}\phi_j\Delta e_{t-j}\right)G(s_t;\gamma,c) + \varepsilon_t,\end{aligned} \tag{8}$$

where π_t is the CPI inflation rate, Δw^*_t is the changes in foreign producer cost, Δy_t is the output growth and Δe_t is the rate of depreciation of the nominal effective exchange rate. $G(s_t;\gamma,c)$ is the logistic transition function driving the nonlinear dynamic. The transition variables used as a measure of macroeconomic instability is the 10-year government bond yield spreads (versus Germany) $s_t = bys_{t-j}$. In our analysis, we focus on the long-run exchange rate pass-through (LR ERPT) which is given by the following long-run time-varying coefficients:

$$\text{LR ERPT} = \left[\sum_{j=0}^{N}\beta_j + \sum_{j=0}^{N}\phi_j G(s_t;\gamma,c)\right] \Big/ \left[1 - \sum_{j=1}^{N}\lambda_j\right] \tag{9}$$

Due to the features of logistic STR models, long-run ERPT coefficient is expected to take different values depending on whether the transition variable, i.e. government bond yield spreads, is below or above the threshold. If the measure of macroeconomic instability is below the threshold, i.e. $(s_t - c) \to -\infty$, then the importing country experiences a stable macroeconomic environment and pass-through elasticity is equal to: LR ERPT$= \sum_{j=0}^{N}\beta_j/(1-\sum_{j=1}^{N}\lambda_j)$. If the measure of macroeconomic instability is above the threshold value, i.e. $(s_t - c) \to +\infty$, then the economy is facing a confidence crisis and pass-through coefficient becomes: LR ERPT$= (\sum_{j=0}^{N}\beta_j + \sum_{j=0}^{N}\phi_j)/(1-\sum_{j=1}^{N}\lambda_j)$.

The LSTR pass-through equation (8) is estimated for the five highly indebted EA countries i.e. Greece, Italy, Ireland, Portugal and Spain, also called the GIIPS group. We use monthly data spanning the period 1993:01 to 2012:12 in order to cover the changing behaviour in the pass-through dynamics during the EA sovereign debt crisis. All the data we use are taken from the OECD's *Main Economic Indicators* database, except for exchange rate series which are obtained from *International Financial Statistics* (IFS) of the International Monetary Fund (IMF) and the 10-year government bond yield spreads are from the European Central Bank (ECB) statistics. Inflation rates series represents the monthly change in consumer prices index (CPI). Output growth is constructed using the rate of growth of the Industrial Production Index. The nominal exchange rate is defined as domestic currency units per unit of foreign currencies, which implies that an increase represents a depreciation for home country. Finally, to capture changes in foreign costs, we follow BAILLIU and FUJII (2004) by constructing an exporter partners' cost proxy. In logarithms, this latter is measured as follow: $w_t^* \equiv q_t + ulc_t - e_t$, where q_t is the unit labor cost (ULC) based real effective exchange rate, ulc_t is the ULC in domestic country and e_t is the nominal effective exchange rate.[10] To determine the lag length of the variables, we follow VAN DIJK, TERÄSVIRTA, and FRANSES (2002) by adopting a general-to-specific approach to select the final specification. We start with a model with maximum lag length of $N = 4$, and then dropping sequentially the lagged variables for which the t-statistic of the corresponding parameter is less than 1.0 in absolute value.

4 Empirical Results

In this section we investigate empirically whether the ERPT responds nonlinearly to the macroeconomic instability in the GIIPS countries. We assume that foreign firms' markup responds nonlinearly to the importing country's general macroeconomic conditions. When the economy faces a financial or a confidence crisis, foreign firms may decide to pass-through a larger proportion of currency changes in view of the increased likelihood of default from the importer. However, in periods of good macroeconomic conditions, foreign firms would prefer local currency price-setting, leading to lower ERPT rates due to a better macroeconomic management.

Therefore, in our empirical analysis we must look for a suitable proxy for macroeconomic instability/confidence crisis. In their LSTR model, NOGUEIRA JR. and LEON-LEDESMA (2011) used the real interest rate differential of Mexico with respect to the U.S. as measures of macroeconomic instability, which is the transition variable in their nonlinear smooth transition model.[11] The use of real interest rate spread as a proxy of macroeconomic instability, and particularly as a leading indicator of confidence crises, has been advocated by KAMINSKY, LIZONDO, and REINHART (1998) among others. In our study, we propose an alternative indicator of macroeconomic instability due the recent context of the European sovereign debt crisis. Hence, we use the sovereign yield spreads to German bonds as an indicator of macroeconomic instability. We expect that this variable would provide

[10] We have checked the possibility of cointegrating relationship among our variables in ERPT equation (4). Individual series in level are non-stationary but do not appear to be cointegrated according to Engle-Granger tests. As a result, log differences of the variables are used in the estimation the STR pass-through equation as shown in equation (8). Augmented Dickey Fuller (ADF) tests suggest that variables in differences are appropriately described as stationary series.

[11] To obtain real interest rate differential, NOGUEIRA JR. and LEON-LEDESMA (2011) used data on money market interest rates for Mexico and for the United States. CPI inflation was used to obtain the real interest rates from the nominal interest rates collected.

some proxy of the risks perceived by foreign firms with respect to the economy under consideration. The widening of sovereign bond yield differentials would indicate the increasing of macroeconomic instability and the loss of confidence in a given economy. In such a case, exporters are willing to shift from LCP strategy to set prices in their own currencies (PCP setting), leading to higher extent of ERPT.

Using LSTR model, we assume exporter's markup to depend nonlinearly on the importing country's sovereign bond yield differential, that is, when the economy faces a confidence crisis, ERPT becomes higher. The transition variables used as a measure of macroeconomic instability in the nonlinear framework is the 10-year government bond yield spreads to the German Bund (bys_{t-j}). The data is obtained from the European Central Bank (ECB) statistics. When the transition variable $s_t = bys_{t-j}$ is exceeding an estimated threshold, these can be interpreted as a period of confidence crisis/macroeconomic instability. The choice of the adequate lagged bond yield spread bys_{t-j} as a transition variable by means of linearity tests is reported in Table 2 in Appendix A. The linearity tests are conducted for each lagged bond yield spread bys_{t-j} with $j = 1,2,3,4$. According to linearity tests, there is a strong evidence of presence of nonlinearities in the five peripheral EA countries. LSTR model is found to be the best specification to capture the nonlinearity with respect to sovereign bond yield differential.

Estimation results from LSTR pass-through equation are summarized in Table 1.[12] In addition to the estimated threshold values and the speed of transition, we report LR ERPT coefficients for the two extremes regimes, i.e. under good macroeconomic conditions $(G(s_t;\gamma,c) = 0)$ and bad macroeconomic environment $(G(s_t;\gamma,c) = 1)$ as defined in (9). Moreover, we compute sum of squared residuals ratio (SSR_{ratio}) between LSTR model and the linear specification which suggests a better fit for the nonlinear model. We also check the quality of the estimated LSTR models by conducting several misspecification tests. In most of cases, the selected LSTR models pass the main diagnostic tests, i.e. no error autocorrelation, no conditional heteroscedasticity, parameters constancy and non remaining nonlinearity.

According to Table 1, threshold values of bond spreads are strongly significant and vary across the GIIPS countries, ranging from 0.67% in Ireland to 2.72% in Greece. Regarding speed of transition γ, our results indicate relatively moderate values which is a proof of smooth transition between good and bad macroeconomic regimes.[13] Concerning long-run ERPT, we note that for the five highly indebted EA countries we find significant nonlinear response of CPI inflation to exchange rate movements with respect to macroeconomic instability. The GIIPS group shows

[12] Full results of NLS estimates of our LSTR models are presented in the Table 3 in Appendix C.

[13] According to VAN DIJK, TERÄSVIRTA, and FRANSES (2002) estimates of γ may appear to be insignificant. This should not be interpreted as evidence of weak nonlinearity.

significant positive relationship between the widening of bond yield spreads and the extent of pass-through. For example, when the Greek bond yield spread (versus Germany) exceeds 2.7%, the rate of pass-through increases from 0.30% (when $G = 0$) to about 0.61% (when $G = 1$). Similarly, for Portugal, the extent of ERPT differs strongly in periods of confidence crisis. The ERPT is equal to 0.32% when the Portuguese yield differential is below 2.14%, but beyond this threshold level, ERPT becomes higher and reaches 0.73%.

Table 1: Estimated ERPT elasticities from the LSTR model with $s_t = bys_{t-i}$

	Greece	**Ireland**	**Italy**	**Portugal**	**Spain**
Transition variable (s_t)	bys_{t-4}	bys_{t-4}	bys_{t-2}	bys_{t-1}	bys_{t-1}
Threshold (c)	2,720	0,670	2,088	2,137	1,098
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
Speed of transition (γ)	28,632	14,187	9,084	10,203	20,264
	(0,348)	(0,352)	(0,326)	(0,468)	(0,318)
Linear Part : G=0					
LR ERPT	0,305	0,071	0,036	0,325	0,203
	(0,002)	(0,318)	(0,553)	(0,003)	(0,153)
Nonlinear Part : G=1					
LR ERPT	0,614	0,782	0,151	0,736	0,472
	(0,001)	(0,140)	(0,029)	(0,071)	(0,093)
R^2	0,947	0,788	0,657	0,694	0,737
SSR_{ratio}	0,588	0,676	0,655	0,670	0,796
AIC	-8,531	-8,857	-10,189	-7,267	-8,859
pJB	0,005	0,134	0,628	0,000	0,187
$pLM_{AR(4)}$	0,760	0,922	0,934	0,513	0,439
$pLM_{ARCH(4)}$	0,511	0,878	0,914	0,946	0,184
pLM_C	0,490	0,797	0,198	0,594	0,275
pLM_{RNL}	0,688	0,473	0,363	0,204	0,347

Note: Table reports elasticities of exchange rate pass-through into CPI inflation from LSTR models. Numbers in parentheses are p-values of estimates. R^2 denotes the coefficient of determination, SSR_{ratio} is the ratio of sum of squared residuals between LSTR model and the linear specification, and *AIC* is the Akaike Information Criterion. The following rows corresponds to the misspecification tests: *pJB* is the p-values of Jarque-Bera normality test, $pLM_{AR(4)}$ is the p-values of the LM test of no error autocorrelation up to forth order, $pLM_{ARCH(4)}$ is the p-values of the LM test of no ARCH effects up to forth order, pLM_C is the p-values of the LM test of parameter constancy and pLM_{RNL} is the p-values of the LM test of no remaining nonlinearity.

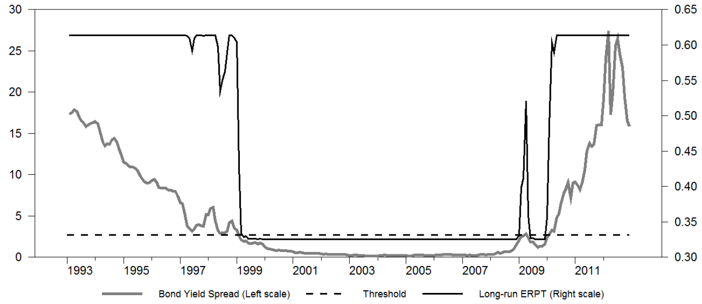
Thereafter, we plot the estimated logistic transition functions and the ERPT as a function of the transition variable $s_t = bys_{t-j}$ (see Figure 3 in Appendix B). It is clear that the transition between both extreme regimes, i.e. $G = 0$ and $G = 1$, is smooth in most of cases. Also, plots reveal the regime dependence of ERPT to macroeconomic conditions. The positive connection between the degree of the ERPT and macroeconomic instability is quite clear for all GIIPS countries.

Our results suggest that there is an important effect of macroeconomic instability on the ERPT. Under bad economic conditions, firms have no incentive to absorb exchange rate movements in their margins which thus leads to higher

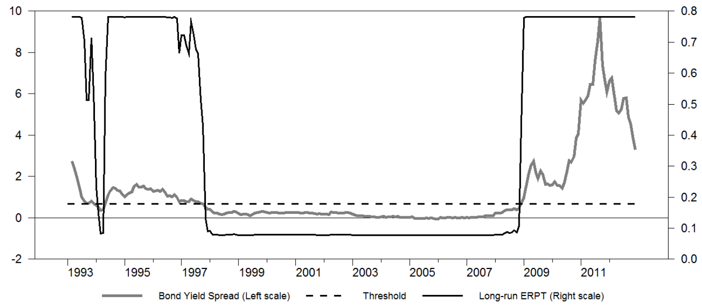
ERPT, in opposition with periods of macroeconomic stability when ERPT would be expected to decline. This is in line with NOGUEIRA JR. and LEON-LEDESMA (2011) who found that the sensibility of CPI inflation is higher when Mexican economy faces financial or a confidence crisis. To gain further insight into the role of crisis in determining the degree of pass-through, plots of long-run ERPT estimates over time and past yield spreads are displayed together in Figure 2 with the estimated threshold level superimposed. The displayed plots reveal that, since the eruption of the sovereign debt crisis in the beginning of 2010, the transmission of the single currency movements becomes higher after ten years of stability at very low levels. The loss of confidence in GIIPS markets has entailed a higher ERPT rates. This effect might result from foreign firms recognizing that those countries are themselves fundamentally in severe trouble. Indeed, the EA sovereign debt crisis would force exporters to follow PCP strategy due to the general weakness of macroeconomic fundamentals in GIIPS group.

Moreover, a very interesting result is that the 10-year yield spreads versus Germany was very low during the first ten years the third stage of the EMU. During this period, there was a small rate of ERPT throughout our GIIPS EA countries. However, during the pre-EA era, the yield differentials were more pronounced with higher degree of exchange rate transmission. It is plausible that the credibility gained from the adoption of the single currency was responsible for the tightening of bond yield spreads and to some extent to the decline in the rates of ERPT. This conclusion reinforces the argument that the introduction of a set of policies that boost market confidence in the economy can indeed lead to lower ERPT. The adoption of sounder policies may be an effective tool for reducing ERPT. Of course, we do not pretend that all the gain in terms of lower ERPT rates are due to better macroeconomic management or the only source of nonlinearity, but it is an important finding for the EA countries with historically poor macroeconomic policies. Furthermore, in this context of sovereign debt crisis more attention must be paid to the impact of the euro fluctuations on the CPI inflation. We see that more macroeconomic instability can give rise to higher ERPT, which can be a serious threat to price stability for the Eurozone members. This conclusion has strong policy implications. European monetary authorities must take into account the nonlinear mechanism of ERPT in periods of financial crisis and how it affects inflation dynamics.

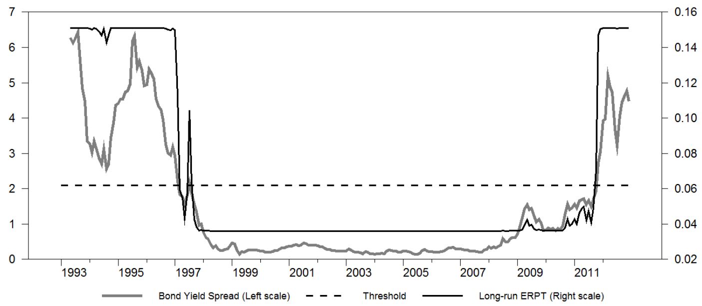
Figure 2: Time-varying long-run ERPT and past bond yield spread



Greece

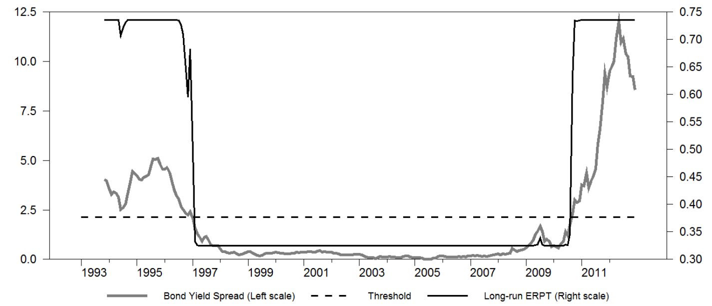


Ireland

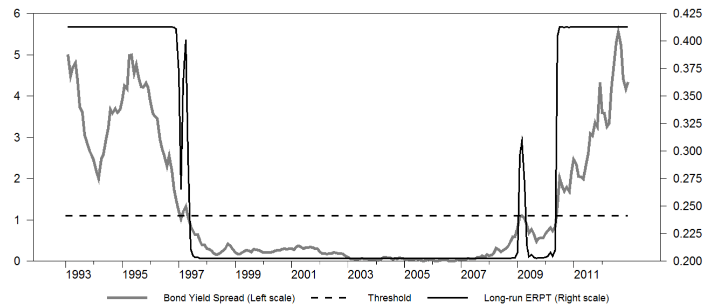


Italy

Figure 2: *Continued*



Portugal



Spain

Note: Time-varying long-run ERPT and past bond yield spread during 1993-2012. Results are from LSTR model with $s_t = bys_{t-i}$.

5 Conclusion

This paper examines the presence of nonlinear mechanism in the extent of pass-through for five heavily indebted EA countries. As the European sovereign debt crisis has spilled over into a strongly deteriorating macroeconomic environment, we investigate whether it has entailed a changing behaviour in the rates of pass-through in a nonlinear way. Using logistic smooth transition models, we explore the existence of nonlinearity with respect to sovereign bond yield spreads (versus German) as an indicator of confidence crisis/macroeconomic instability. Using monthly data spanning from 1993 to 2012, we find that, for our highly indebted EA economies, i.e. the GIIPS group (Greece, Ireland, Italy, Portugal, and Spain), the extent of ERPT would be higher in periods of macroeconomic distress. In other words, the increasing of macroeconomic instability and the loss of confidence would entail a higher sensibility of CPI inflation to exchange rate movements in the GIIPS countries. The perception of relatively high sovereign debt default risk in the fiscally distressed EA countries may lead exporting firms to change pricing behaviour, shifting from LCP to PCP setting which lead to higher degree of ERPT.

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Appendix A. Linearity test

In order to derive a linearity test, TERÄSVIRTA (1994, 1998) suggested to approximate the logistic function (7) in (6) by a third-order Taylor expansion around the null hypothesis $\gamma = 0$. The resulting test has power against both the logistic STR (LSTR) and exponential STR (ESTR) models. Assuming that the transition variable s_t is an element in $\mathbf{z}_t$ and let $\mathbf{z}_t = (1, \tilde{\mathbf{z}}_t^{'})^{'}$, where $\tilde{\mathbf{z}}_t^{'}$ is an $(m \times 1)$. Taylor approximation yields the following auxiliary regression:

$$y_t = \alpha_0^{'} \mathbf{z}_t + \sum_{j=1}^{3} \alpha_j^{'} \tilde{\mathbf{z}}_t s_t^j + u_t^*, \quad t = 1, ..., T, \tag{10}$$

Where $u_t^* = u_t + R_3(\gamma, c, s_t)\theta^{'} \mathbf{z}_t$, with $R_3(\gamma, c, s_t)$ the residual of Taylor expansion. The null hypothesis of linearity is $H_0 : \alpha_1 = \alpha_2 = \alpha_3 = 0$. VAN DIJK, TERÄSVIRTA, and FRANSES (2002) suggest the use of F-versions of Lagrange Multiplier (LM) test statistic, which has an approximate F-distribution with $3m$ and $T - 4m - 1$ degrees of freedom under H_0. Linearity tests are executed for each of the candidates potential transition variables, which are lagged bond yield spreads in our case. Once linearity has been rejected, one has to choose whether logistic or exponential function should be specified. The choice between these two types of models is based on the auxiliary regression (10). TERÄSVIRTA (1994, 1998) suggested that this choice can be based on testing the following sequence of nested null hypotheses:

1. Test $H_{04} : \alpha_3 = 0$
2. Test $H_{03} : \alpha_2 = 0 | \alpha_3 = 0$
3. Test $H_{02} : \alpha_1 = 0 | \alpha_2 = \alpha_3 = 0$

According to TERÄSVIRTA (1994), the decision rule is the following: if the test of H_{03} yields the strongest rejection measured in the p-value, choose the ESTR model. Otherwise, select the LSTR model. Table (2) provides the p-values of the F version of the LM test with the different lags for the bond yield spreads. The third column reports the test of the null hypothesis of linearity against the alternative of STR nonlinear model. The following columns give the sequence of null hypotheses for choosing the LSTR or the ESTR model.

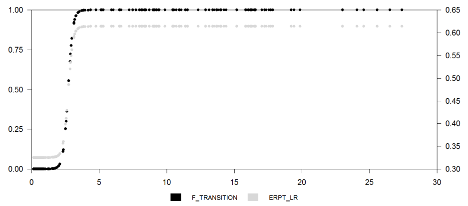
Table 2: Linearity tests against STR model with $s_t = bys_{t-j}$

Country	Transition Variable	H_0	H_{04}	H_{03}	H_{01}	Specification
Greece	bys_{t-1}	0,026	0,720	0,139	0,000	LSTR
	bys_{t-2}	0,128	0,738	0,386	0,000	Linear
	bys_{t-3}	0,079	0,685	0,296	0,000	Linear
	bys_{t-4}	0,047	0,080	0,210	0,256	LSTR
Ireland	bys_{t-1}	0,058	0,185	0,265	0,063	Linear
	bys_{t-2}	0,009	0,121	0,073	0,026	LSTR
	bys_{t-3}	0,122	0,678	0,198	0,014	Linear
	bys_{t-4}	0,002	0,571	0,143	0,000	LSTR
Italy	bys_{t-1}	0,000	0,042	0,002	0,040	ESTR
	bys_{t-2}	0,000	0,001	0,013	0,023	LSTR
	bys_{t-3}	0,000	0,032	0,006	0,011	ESTR
	bys_{t-4}	0,000	0,057	0,007	0,005	LSTR
Portugal	bys_{t-1}	0,002	0,017	0,081	0,077	LSTR
	bys_{t-2}	0,058	0,185	0,265	0,063	Linear
	bys_{t-3}	0,026	0,035	0,563	0,047	LSTR
	bys_{t-4}	0,031	0,035	0,803	0,018	LSTR
Spain	bys_{t-1}	0,003	0,049	0,092	0,015	LSTR
	bys_{t-2}	0,003	0,043	0,165	0,006	LSTR
	bys_{t-3}	0,004	0,077	0,177	0,003	LSTR
	bys_{t-4}	0,006	0,112	0,180	0,003	LSTR

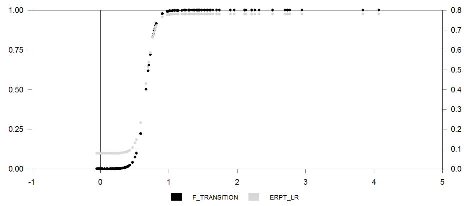
Note: The numbers are p-values of F versions of the LM linearity tests. Third columns show the test of linearity against the alternative of STR nonlinearity. In the forth columns until the sixth are the p-values of the sequential test for choosing the adequate transition function. The decision rule is the following: if the test of H_{03} yields the strongest rejection of null hypothesis, we choose the ESTR model. Otherwise, we select the LSTR model. The last row gives the selected model.

Appendix B. Plots from LSTR pass-through equation

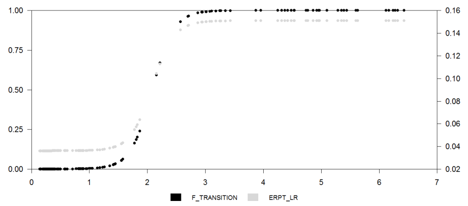
Figure 3: Logistic functions and long-run ERPT as a function of past yield spread



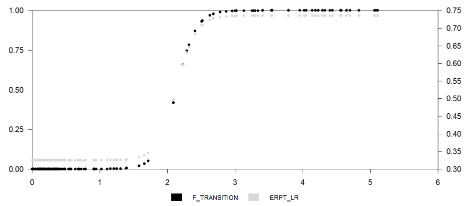
Greece



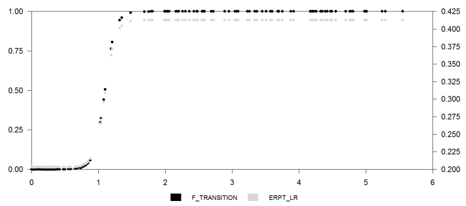
Ireland



Italy



Portugal



Spain

Note: Estimated transition functions and long-run ERPT as function of past bond yield spread. Results are from LSTR model with $s_t = bys_{t-i}$.

Appendix C. Full Results from LSTR models

Table 3: Estimation results from LSTR pass-through equation

	Greece	Ireland	Italy	Portugal	Spain
Transition variable (s_t)	bs_{t-3}	bs_{t-2}	bs_{t-2}	bs_{t-1}	bs_{t-4}
Threshold (c)	2,720	0,670	2,088	2,137	1,098
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
Speed of transition (γ)	28,632	14,187	9,084	10,203	20,264
	(0,348)	(0,352)	(0,326)	(0,468)	(0,318)
Linear Part: G=0					
Constant	-0,005	-0,006	0,002	0,006	-0,002
	(0,043)	(0,000)	(0,000)	(0,002)	(0,005)
π_{t-1}	-0,401	0,150		0,366	0,287
	(0,000)	(0,144)		(0,000)	(0,001)
π_{t-2}			0,143		
			(0,109)		
π_{t-3}		0,147			
		(0,141)			
π_{t-4}	0,120		0,283		0,197
	(0,231)		(0,003)		(0,019)
Δe_t	0,243	0,100	0,012	0,163	0,039
	(0,004)	(0,010)	(0,588)	(0,000)	(0,009)
Δe_{t-1}			0,041		
			(0,074)		
Δe_{t-2}	0,174				0,022
	(0,037)				(0,704)
Δe_{t-3}			-0,033		
			(0,153)		
Δe_{t-4}		-0,045		0,043	
		(0,207)		(0,320)	
Δw_t^*	0,059	0,070	0,030	0,375	0,024
	(0,014)	(0,005)	(0,006)	(0,000)	(0,223)
Δw_{t-1}^*	0,040		0,020	0,200	
	(0,125)		(0,064)	(0,005)	
Δw_{t-2}^*	0,062		0,011		
	(0,014)		(0,312)		
Δw_{t-3}^*		-0,045		-0,115	0,024
		(0,125)		(0,112)	(0,275)
Δw_{t-4}^*			-0,014		
			(0,194)		
Δy_t				-0,030	
				(0,166)	
Δy_{t-1}					0,012
					(0,445)
Δy_{t-2}		-0,007	0,009		0,019
		(0,226)	(0,261)		(0,208)
Δy_{t-3}	-0,023				
	(0,188)				
Δy_{t-4}	-0,026	0,006			
	(0,072)	(0,232)			
Nonlinear Part: G=1					
Δe_t	0,189	0,283	0,021	0,100	0,068
	(0,103)	(0,429)	(0,485)	(0,615)	(0,387)
Δe_{t-1}		0,213	-0,039		0,041
		(0,120)	(0,233)		(0,412)
Δe_{t-2}	0,180			-0,138	0,029
	(0,101)			(0,118)	(0,572)
Δe_{t-3}			0,045		
			(0,139)		
Δe_{t-4}			0,039	0,299	
			(0,184)	(0,014)	

Key: Table reports estimates of LSTR pass-through equation. Numbers in parentheses are p-values.