Long Memory In Import and Export Price Inflation and Persistence of Shocks to the Terms of Trade

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Abstract

Long memory models have been successfully used to investigate the dynamic time-series behavior of inflation rates based on the CPI and WPI. However, almost no attention has been paid to import and export price inflation, nor to the terms of trade which they make up. This article investigates the dynamics of the terms of trade by focusing first on the time-series characteristics of these price series. It tests for long memory in export and import price inflation series and estimates the fractional differencing parameter using a number of approaches. To give a better idea of the degree of persistence of each series, estimates of the impulse responses are computed which take into account possible fractional integration. The dynamic behavior in changes in the terms of trade is then related to the long memory behavior of the import and export price inflation series. In a sample of eleven economies for which data is available, evidence of long memory in import and export price inflation occurs in about half the cases. Granger (1980) points out that the natural occurrence of long memory may be attributed to aggregation in macroeconomic series. Our analysis provides evidence of an alternative explanation, namely that long-memory may result from the differencing of a linear relationship between non-cointegrating variables. Specifically, the results from our analysis of eleven economies reveal that shocks to the terms of trade will persist if the constituent price inflation series are not cointegrated.

Keywords: long memory, terms of trade, imported inflation

JEL: C22, E31, F18

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1. Introduction

In international economics, the terms of trade represents the relative advantage which may accrue to the parties involved in the exchange of goods. The notion is usually implemented using the ratio of an index of export prices to a corresponding index of import prices. The classic reasoning supposes that a country benefits from an improved terms of trade in the sense that it has to give up less exports in exchange for the imports. Formal statement of this reasoning is often attributed to Harberger (1950) and Laursen and Metzler (1950). They hypothesize that a country's real income level would increase following an improvement in its terms of trade. Since the increase in real income will also have an impact on saving, they argue that improvement also extends to the trade balance. The first of these hypotheses is investigated by Obstfeld (1982), while Svensson and Razin (1983) and Persson and Svensson (1985) study the second. Separately, Ostry (1987) and Edwards (1989) employ very similar approaches to analyzing how temporary shocks to the terms of trade affect the path of real exchange rates and the real account. Within a general equilibrium, intertemporal setting of a small open economy with optimizing consumers and producers, they consider changes in the internal terms of trade (arising from tariff changes) as well as changes to the external terms of trade.

In a recent study on 42 sub-Saharan African countries using annual data from 1960-96, Cashin, McDermott and Pattillo (2004, hereafter CMP) measures the persistence of shocks to the terms of trade using estimates of the half-lives based on median-unbiased estimation (Andrews, 1993). While the majority of cases yield estimates implying shocks to the terms of trade to be impermanent, in about a third of the cases, the half-life estimates imply long-lived shocks. In motivating their focus on measures of persistence, CMP mentions the policy challenges that African countries face in dealing with shocks to the net barter terms of trade (NBTT), including the potential adoption of policy rules to make simple adjustments in the face of price shocks. The aim of any such policy rules would be to deflect the transmission of commodity price shocks through the NBTT to domestic economic performance.

While we agree that that the issue of shocks to the NBTT is important aspect for trade policy considerations, we believe a natural question that should be addressed is how such shocks may have been transmitted from the import and export price indexes which constitute the NBTT. Policy rules could simply be rendered impotent if either the NBTT is mistakenly considered to be stationary when it is in fact not so. To inform any policy rules, it would also be helpful to know about the connection to shocks to the constituent price indexes. However, we have not come across any study which has given any consideration to the hypothesis that price factors may underlie the dynamic responses of terms of trade to shocks.

Since the seventies, inflation issues have loomed large in empirical macroeconomics. Research into the properties of univariate inflation time-series has been extensive. To a certain extent, such results tend to characterize the inflationary experiences of some countries as being simple artifacts of the time-series properties. However, the experience of chronic inflation does not necessarily manifest itself in the same way in all countries. Depending on the structure of the domestic economy, inflation may be more of an issue at the retail level than at the wholesale level. The number of countries covered in Baum et al (1995), and Hassler and Wolters (1995) provides evidence of the pervasiveness of long memory in inflation. Certain sectors within the economy may be affected only. In this article, we focus on the trade sector. Despite the importance of import price inflation, there has been virtually no attempt to model its long-memory characteristics.

Countries which have known histories of severe inflation (and the studies which deal with them) include the US (Nelson and Schwert, 1977, Ball and Cechetti, 1990, Kim, 1993), UK (Franses and Ooms, 1997), and Spain (Delgado and Robinson, 1994). On the other hand. the record for economies such as Singapore, Hong Kong SAR-PRC (hereafter, HK), and

Japan has been mixed. Rather, the latter cases are often considered to be on the receiving end of a transmission process. The channel of transmission is trade, and the issue for these economies is one of imported inflation. Unlike domestic inflation, import price inflation is an issue for virtually all economies. The only possible exceptions are when trade occurs mainly in a common currency. A possible recent exception is the experience of countries like Belgium in the recent Euro-zone arrangement.

Barring such exceptions, post Bretton-Woods, import price inflation has been an issue for all trading countries. This is especially so with the move toward removing barriers to trade. With the momentum gained through the GATT talks and gathering force within the WTO, the reality is that increased trade has led to increased exposure to non-domestic price pressures.

Many countries have begun to pay serious attention to their competitiveness in the face of increasing globalization. The celebratory mood that followed the conclusion of the Uruguay round of trade talks that culminated in the WTO has begun to dissipate. In its wake, several countries are looking at or have invoked the dispute settlement mechanism of the WTO to deal with increasingly difficult issues.

One of the key questions many countries seek to deal with is the exposure that the dismantling of trade barriers places their domestic economies under. The main aspect of this exposure may be found in the terms of trade. While measures such as the current account balance signal trade competitiveness and have always had a much more positive portrayal in economics, the terms of trade has been the focus of attention in debates about the inability of poorer countries to command fairer prices for their products. The importance of the issue is seen in the fact that, since 1964, UNCTAD has devoted a forum to focus on the terms of trade as a measure of improvement in trying to achieve its aim of helping developing and under-developed nations maintain improve and maintain living standards. According to a recent

UNCTAD report (Economic Development in Africa: Trade Performance and Commodities Dependence, 26 Feb 2004), the majority of African countries are boxed into a trading structure that subjects them to secular terms-of-trade losses and volatile foreign exchange earnings.

There are many variants of the basic terms of trade definition. Instead of using the income terms of trade which is easier to obtain and would have yielded a larger useable sample, the definition employed throughout this article is the NBTT. This is the ratio of an index of export prices to an index of import prices. This definition proves to be useful for linking measures with domestic economic and trade significance, especially when interest centers on the rate of change of the terms of trade. In particular, taking logs of the ratio gives

$$TT = EX - IM \,, \tag{1}$$

Here, we let *TT*, *EX* and *IM* denote the logged values each of the terms of trade, export price index and import price index. The rate of change in the terms of trade is the difference in export price inflation and import price inflation i.e.

$$\Delta TT = \Delta EX - \Delta IM \,, \tag{2}$$

Clearly (1) is simply a definition of the NBTT and nominally devoid of model character. In the special case that *TT* is stationary while *EX* and *IM* are I(1) (integrated of order 1), (1) implies that EX and IM are cointegrated with cointegrating vector (1,−1). While this need not hold in general, we will provide evidence that such a special cointegrating relation holds in certain cases.

Under the maintained hypothesis of stationarity, Persson and Terasvirta (2003) use a smooth transition autoregressive model to model the NBTT. A number of countries seem to suffer from persistently poor terms of trade. CMP focuses on the fact that those in sub-Saharan seem helpless to reverse such a disadvantageous position. One possible explanation, entertained by CMP, is that the problem centers on the way shocks play out for such countries. One way of unraveling the apparently contradictory representation in what the terms of trade in open economies represents for the rich-poor divide is to consider the amount of time it takes for shocks to dissipate.

Most notably, time series analysis of the NBTT has apparently not touched on the potential effects for long-memory. If the constituent price inflation series were to exhibit long-memory, it is possible for the rate of change of the terms of trade, ΔTT, to have inherited some of this persistence. In that case, one would expect shocks to the NBTT to last for a considerable length of time. Our aim is to form a profile of those cases where long-memory may occur. We do this by first analyzing the potential cointegrating relation between the price index series, *EX* and *IM*. The results are then compared to measures of persistence estimated from models of long memory.

Baillie's (1996) authoritative survey describes how price series often display signs of non-stationarity, while appearing over-differenced if the unit root is imposed. In such situations, ARFIMA models have been shown to work. Indeed, since non-stationary and stationary models can be represented as special cases of the ARFIMA specification, ARFIMA models are particularly convenient for bridging the gulf between the two. For example, although the relevant asymptotic theory usually applies for only a certain range of the fractional order of differencing, d (e.g. Robinson's (1995) result requires $-\frac{1}{2} < d < \frac{1}{2}$), this is not as restrictive as it might appear because it accommodates both antipersistence and long-memory. In this article, we will fit ARFIMA models to the terms of trade and its constituents, the import and export price indexes.

In the next section, we review the methodology of ARFIMA modeling. In Section 3, we discuss the estimation techniques that are employed. In Section 4, empirical estimates are presented. Concluding remarks may be found in Section 5.

2. Methodology

In this paper, we focus on the fact that under certain conditions, potential long memory in ΔEX and ΔIM could have predictable conclusions for the long memory character of ΔTT. To begin, we give an overview to clarify the distinction between long memory, short memory and nonstationarity.

Let x_t denote a time series on which a sample of observations for $t = 1,2,\ldots,n$ is available. In this paper, x_t could represent the logarithms of import prices, export prices, or NBTT, or the inflation series derived from import prices, export prices or the NBTT.

Assuming absolute continuity of the spectral distribution function for a scalar covariance stationary series x_t, the autocovariances γ_k $(k = 0,\pm1,\pm2,\ldots)$ can be defined in terms of the spectral density at frequency $-\pi < \lambda \le \pi$ as (Robinson, 1995, p.2)

$$\gamma_k = \operatorname{cov}(x_t, x_{t-k}) = \int_{-\pi}^{\pi} f(\lambda)\cos(k\lambda)d\lambda . \qquad (3)$$

A long memory process x_t arises if the spectral density is assumed to also satisfy

$$f(\lambda) \sim C\lambda^{-2d} \quad \text{as } \lambda \to 0^+ \qquad (4)$$

where "~" means that the ratio of the left- and right-hand sides tends to 1, $0 < C < \infty$, and $-\frac{1}{2} < d < \frac{1}{2}$. The autocovariances corresponding to (4) follow

$$\gamma_k \sim Ck^{2d-1} \quad \text{as } k \to \infty . \qquad (5)$$

The original Box-Jenkins class of ARIMA(*p*,*d*,*q*) models considered only integer values of *d*. Realizing that non-integer values of *d* are potentially valid on statistical or mathematical grounds, Granger (1980), Granger and Joyeux (1980) and Hosking (1981) show that a class of fractionally integrated ARMA processes can be defined by raising the differencing operator to non-integer powers. Let B denote the backward shift (or lag) operator whose action is defined by $Bx_t = x_{t-1}$. For real values of $d > -1$, the fractional difference operator Δ^d is defined in the natural way, by a binomial series by:

$$\Delta^d = (1-B)^d = \sum_{k=0}^{\infty} \binom{d}{k} (-B)^k = 1 - dB - \tfrac{1}{2} d(1-d)B^2 - \tfrac{1}{6} d(1-d)(2-d)B^3 - \ldots$$

or in more concise notation,

$$\Delta^d = (1-B)^d = \sum_{k=0}^{\infty} (-1)^k \binom{d}{k} B^k . \qquad (6)$$

For $d > 0$, the expansion (6) can also be represented in terms of the hypergeometric function $F(-d,1,1;B)$ (Baillie, 1996, eq. 15). With this, a time series with p=q=0 has the representation

$$(1-B)^d x_t = \sum_{k=0}^{\infty} \frac{\Gamma(k-d)}{\Gamma(-d)\Gamma(k+1)} x_{t-k} .$$

The simplest model in this class is the ARFIMA(0,d,0) model $(1-B)^d (x_t - \mu) = \varepsilon_t$.

Granger and Joyeux (1980) and Hosking (1981) identify the fundamental behavior of the time series according to whether $-\frac{1}{2} < d < \frac{1}{2}$. Within this range, x_t is stationary and invertible. When d = 0, the model reduces to the case of an ARMA(p,q) model, for the spectral density of x_t is finite and positive at $\lambda = 0$ and autocorrelations are summable. In the case $0 < d < \frac{1}{2}$, x_t is said to have long memory, while in the case $-\frac{1}{2} < d < 0$, the series is said to be antipersistent.

When $d < \frac{1}{2}$, the series has finite variance, while for $d = \frac{1}{2}$, the series has infinite variance. In the latter case, the Box-Jenkins approach will suggest the need for differencing. The differenced series will then have a spectrum whose value at zero frequency is zero, provided that $d < 1$. When d = 1, the time-series process x_t is said to have a unit root.

A statistically significant estimate of the fractional order of differencing, d, is often taken as an indication of the degree of long memory in a time series. Instead of confining ourselves to this measure, we follow Balcilar's (2002) suggestion and derive estimates of the length of time a unit shock takes to dissipate. The impulse response $\xi_k = \partial x_{t+k} / \partial \varepsilon_t$ measures the effect on x_t due to a unit shock to the innovation ε_t at time *t*. To compare with the half-life measures that CMP use, we compute the number of periods, *k*, it takes for 100α% of the effect of a unit shock to x_t to dissipate, using

$$\tau_\alpha = \sup_k \left\{ \left| \frac{\partial x_{t+k}}{\partial \varepsilon_t} \right| \leq 1 - \alpha \right\}$$

These values are reported in Tables B5, B11, C5 and D5.

The impulse responses can be obtained by writing any process in the form of a pure (possibly infinite) moving average representation and reading off the resulting coefficients. For a stationary ARFIMA model the impulse responses are given by the coefficients of

$$\Xi(B) = (1-B)^{-d} \Phi(B)\Theta(B) = 1 + \xi_1 B + \xi_2 B^2 + \cdots \tag{7}$$

Stationarity implies that the impulse responses are square-summable $\sum_{k=1}^{\infty} \xi_k^2 < \infty$. For an ARFIMA(0,*d*,0) process, it can be shown that

$$\xi_k = \frac{\Gamma(k+d)}{\Gamma(k+1)\Gamma(d)} \tag{8}$$

This implies that $\xi_k \sim k^{d-1} / \Gamma(d)$ as $k \to \infty$. It follows that for any $l > k$, $\xi_l / \xi_k = (l/k)^{d-1}$ as $k \to \infty$. For a fractionally integrated process where *d* is near to 1, the ratio ξ_l / ξ_k tends to unity in the limit. In Tables B6, B12, C6 and D6, confidence intervals for the impulse responses ξ_k are reported from parametric bootstraps.

3. Estimation Techniques

The key issue is how to obtain a reliable estimate of the degree of fractional differencing. In this section, we outline the main options which have been developed within the frequency-domain.

Geweke and Porter-Hudak's (1983) estimator is based on the suggestion of a semiparametric approach by Granger and Joyeux (1980). For an ARFIMA(0,*d*,0) model $(1-B)^d x_t = \varepsilon_t$ where $-\frac{1}{2} < d < \frac{1}{2}$ and $E(\varepsilon_t \varepsilon_{t-j}) = 0 \; \forall j > 0$, Geweke and Porter-Hudak (1983) propose estimating

$$\log(I(\lambda_{jt})) = \log \frac{\sigma^2 f_\varepsilon(0)}{2\pi} - d \log\left[4\sin^2(\tfrac{1}{2}\lambda_{jt})\right] + \log\left[\frac{I(\lambda_{jt})}{f_\varepsilon(\lambda_{jt})}\right]. \qquad (7)$$

Here $f_\varepsilon(\lambda_{jt})$ is the spectral density of $\varepsilon_t = (1-B)^d x_t$ at *t*, and $I(\lambda_{jt})$ is the periodogram at the harmonic coordinates $\lambda_{jt} = 2\pi j/n$ of each point *j* = 0, 1, …, *n* – 1, *n* of the time-domain sample span. (7) has the form of a regression equation where the intercept is just the first term on the right-hand side, the slope coefficient is the fractional differencing parameter *d*, and the disturbance term is the last term on the right. The implementation of (7) as a regression simply involves computing the required quantities as observable functions of the harmonic coordinates from the original data set. (7) requires that λ_j be close to zero such that $\lambda_j < \lambda_m$, where λ_m is small and depends on the number of values available for estimation, *m*. Geweke and Porter-Hudak (1983) show that *m* is a function of the sample size *n* such that

$$\frac{m}{n} \to 0 \text{ as } n \to \infty .$$

They argue for the existence of a sequence *m* such that

$$\frac{(\log n)^2}{m} \to 0 \text{ as } n \to \infty$$

under which the estimate of *d* has an asymptotic normal distribution.

Ooms and Hassler (1997) discuss a problem with Geweke and Porter-Hudak's (1983) method when seasonal dummies are used to deseasonalize seasonal dummies. They show that the estimator will contain singularities, and recommend that in order to avoid the problem, the data set should be extended to full calendar years by padding it with zeros, after which the periodogram ordinates corresponding to seasonal frequencies are removed.

Robinson (1995) has proposed an alternative Gaussian semiparametric estimator that is also based on the periodogram. Beginning with just the parametric specification of the spectral density in (4). and setting $d = H - \frac{1}{2}$ where $0 < H < 1$ is the so-called *Hurst parameter*, Robinson (1995) shows that the estimate of H is given by minimizing the criterion

$$\log\left[\frac{1}{m}\sum_{j=1}^{m}\lambda_j^{2H-1}I(\lambda_j)\right]-\frac{(2H-1)}{m}\sum_{j=1}^{m}\log(\lambda_j) \tag{8}$$

where *m*, usually taken to be less than or equal to $(n-1)/2$, is chosen such that $\frac{1}{m}+\frac{m}{n}\rightarrow 0$ as $n\rightarrow\infty$. Robinson (1995) shows that $\sqrt{m}(\hat{d}-d)$ tends in distribution to $N(0,1/4)$ as $n\rightarrow\infty$.

Frequency-domain techniques such as these work on the basis that signals of higher frequency should be ignored. Given their association with short-term memory characteristics of the time-series (see, for example, Granger, 1966), they are considered unhelpful in estimating the long-memory characteristics. In practice, however, this opens up a number of issues especially the fact that the resulting estimates are sensitive to bandwidth choice.

The choice of bandwidth involves balancing the desirability of eliminating high-frequency signals against estimator precision. The estimate of *d* will be biased by the failure to exclude medium or high order periodogram ordinates. While a smaller bandwidth may allow sweep out more of the high-frequency signals, it ends up reducing the size of the useable sample and increasing sampling variability. As a case in point, for Geweke and Porter-Hudak's (1983) log-periodogram estimator, we follow most researchers in determining the number of periodogram ordinates available to estimate *d* by $[n^{\alpha}]$.

There is a large body of theory focusing on the choice of optimal bandwidth, exemplified by the discussions in Delgado and Robinson (1996), Henry and Robinson (1996), and Hurvich et al (1998). In light of views such as those of Geweke (1998), we believe that one way to circumvent the difficulties associated with apparently optimal selection methods is to present our results for a range of bandwidths.

4. Empirical Results

Data is obtained from Datastream on the monthly NBTT for twelve economies. Except for the United Kingdom, data on export and import price indexes are also available. The span of each data for each economy is different depending on data availability: Spain, 70M1-03M12, 408 months; Japan, 60M1-04M2, 530 months; South Korea (hereafter, Korea), 71M1-04M2, 398 months, Singapore, 78M1-04M1, 313 months; Taiwan, 76M1-04M2, 338 months, the United States, 88M1-03M3, 183 months; Mexico, 70M1-03M12, 408 months; Belgium, 93M1-03M12, 132 months, Brazil, 78M1-04M2, 314 months; Finland, 85M1-04M2, 230 months; Hong Kong Special Autonomous Region of the PRC, 231 months; and the United Kingdom (no export or import price indexes), 80M1-04M1, 289 months.

Before carrying out the ARFIMA modeling, the existence of unit roots is established in the logs of the terms of trade, the logs of import prices and the logs of export prices. This is carried out using several improved test statistics of Ng and Perron (2001) that built on the modifications first advanced in Elliot, Rothenberg and Stock (1996). The results are reported in Table A1. We are also interested in the distinction between series which are I(0) and those which are I(d), where the latter denotes fractional differencing with order d (d<1). Appropriate tests which can distinguish between long memory and short memory include the Kwiatkowski et al (1992) KPSS test, and Lobato and Robinson's (1998) LM test. In both cases, the testing pits a stationary, short memory null against a long-memory alternative. Lee and Schmidt (1992) reiterate that the KPSS test is consistent against the alternative that the series in question is I(d) where d<1. We only report the values of the Lobato-Robinson LM test statistic in Table A2. The test is applied to the logged values of all series in both levels and differences. Given that we have already established the I(1) status of the logged values of terms of trade, import prices and export prices, the test of the series in levels based on the LM statistic is only provided for comparison. We are more interested in whether evidence of long-memory exists for the differenced form of the logged series. For instance, Table A2 shows that for 5 economies (Japan, Korea, Singapore, Taiwan and HK), I(0) in import price inflation is rejected in favor of long-memory. It also shows that for 6 economies (Japan, Korea, Singapore, Mexico, Belgium and HK), I(0) in export price inflation is rejected in favor of long-memory.

To give a better accounting of the relationships underlying the nonstationary variables revealed by the unit root tests, it seems reasonable to think in terms of the possible relationships between *EX* and *IM* that can be determined by tests for cointegration. Details of the cointegration analysis are given in Table A3. Overall, the results of Johansen's (1995) test for cointegration depict three possible relationships between *TT*, *EX* and *IM*.

First, Johansen's (1995) trace and $\lambda_{\max}$ tests give values which are not significant at the 5% level for Japan, Korea and Taiwan. This implies that in these three cases, *TT* is non-stationary. In two other cases, namely US and HK, the estimated cointegrating relationship agrees with the definition of NBTT in (1). It follows that for these two, *TT* is stationary. In the remaining six cases, cointegrating relationships are found that differ from the definition of NBTT in (1). The question that we proceed to address is how what each of these three types of relationships imply for the long-memory characteristics of ΔTT , ΔEX and ΔIM .

Tables B1-B12 present the estimation results for the levels and first differences of the terms of trade, Tables C1-C6 present the results for the import price inflation rates and Tables D1-D6 contain results for the export price inflation rates. The data is deseasonalised and appropriate treatment to account for this follows the recommendation in Ooms and Hassler (1997).Model estimation is carried using the arfima package 1.01 for Ox (Doornik, 2001).

Results are given in Tables B1-B2, B7-B8, C1-C2, and D1-D2 on estimation using the methods of Geweke and Porter-Hudak (1983) and Robinson (1995). Results are also given on the fractional Gaussian noise (FGN) model of Mandelbrot (1963) estimated using the Whittle approximate maximum likelihood method (Fox and Taqqu, 1986). We also estimate a number of ARFIMA(p,d,q) specifications using the approximate Whittle estimator. Tables B3, B9, C3 and D3 report the AICs (Akaike information criteria) from the successful trials. We used orders of up to 3 each for p and q, and found that the models with high orders invariably fail to converge. Thus, in all cases, the best model chosen on the basis of the AIC is the ARFIMA(0,d,0).

Tables B1 and B2 present the estimates of the order of fractional integration of *TT*. Estimates of the fractional differencing parameter are presented for a number of bandwidth choices. In the case of Korea, Belgium, Brazil and HK, the estimates appear to be relatively more sensitive to bandwidth. Recall that the test results in Table A1 basically show that almost all of the *TT* series are nonstationary, excepting possibly Taiwan, US and HK. The estimates of d here imply that long memory is a feature of the *TT* series in all cases.

The estimates of d in Tables B7 and B8 apply to the ΔTT series. These appear more sensitive to bandwidth choice.

Instead of focusing on the estimate of d, we find it more convenient to use the estimates of the impulse responses and τ_α to assess the degree of persistence. In Section 3, we described how persistence is to be evaluated using the estimated impulse responses. The 95% confidence intervals of the estimated impulse responses from the best model – always the ARFIMA(0,d,0) in our computations – are obtained using 5000 bootstrap replications. Thus in Tables B6, B12, C6 and D6, we spot persistence by relying on the fact that long memory processes would have impulse responses which differ from zero even after 1200 months or a century. Of particular interest, the estimates of τ_α for the ΔTT, ΔEX and ΔIM series of Japan, Korea, Taiwan and Mexico show patterns consistent with long-memory. Except for Mexico, the other three cases are the ones which show up in Table A3 to have no cointegrating relations of any sort.

5. Concluding Remarks

For four of the economies (Japan, Korea, Taiwan) for which data on all three variables are available, we find that all of the growth series (i.e. export price inflation, import price inflation series and growth in terms of trade) have long memory. This is seen from results of the cointegration analysis in Table A3 and corroborated by the estimated values of fractional differencing as well as from the impulse response estimates. Also, from the results of the cointegration analysis in Table A3, the long memory features of ΔTT for Japan, Korea and Taiwan look to have resulted from differencing of a non-cointegrating relationship. In two cases, HK and US, the cointegrating relationship we have estimated is almost indistinguishable from the definition of the terms of trade. Finally, the six remaining cases yield cointegrating relationships which are quite different from the definition of the terms of trade.

Thus, in our sample, it is true to say that shocks to the terms of trade will persist when the constituent import and export price indexes are not cointegrated. Aside from this generalization, other features are worth noting.

For Singapore, Hong Kong, the U.S., and Finland, the import and export price inflation series appear to have long memory on the basis of both, the estimates of d and the impulse responses. However, this is not true of the growth rate series ΔTT. In fact, the evidence points to the growth rate of the terms of trade in these cases being stationary. Thus, we see that even though *EX* and *IM* have a cointegrating vector different from (1,-1) in these cases, the mere fact of being cointegrated appears sufficient to remove long-memory from ΔTT.

There are also some exceptions in the cases of Spain, Brazil and Belgium. The growth rates of the terms of trade and import and export price inflation series of these three countries display signs of over-differencing. For these three countries, long memory presents itself in the levels of the series instead.

In conclusion, we attempt to validate the above estimates by referring to the impact that trade price pressures may play in some of the economies concerned. As mentioned earlier, for better or worse, all trading economies in a global setting are subject to imported inflation. Much of the world first took notice of this in the oil price hikes of the seventies. Yet, even without energy-cost factors, there are more and louder reminders of the role that import prices play in channeling inflationary pressures to domestic economies. The main reason is the increased volatility of global financial markets, coupled with several isolated but significant incidents of exchange rate weakness.

The economies in our sample are chosen on the basis of data availability. Nonetheless, they exhibit a wide range of characteristics in trade which enables a better understanding of the empirical estimates presented later on.

Belgium is a member of the EU and more recently, one of the founding entrants of the Euro-zone common currency arrangement. Belgium is well-known for its specialization on intermediate goods and would be insulated from trade price shocks as a result of its activities occurring mainly within Europe. The basic determinant of fluctuations in this regard comes from inventory cycles, rather than currency or external price shocks. The antipersistence of the growth rate of the terms of trade possibly reflects the lack of import price inflationary pressures.

A large proportion of the trading that HK does occurs mainly with the U.S. and mainland PRC. By virtue of the controlled currency arrangements within this setting, exchange rate fluctuations would be expected to be less likely to be the source of import inflationary pressures. Domestically, HK-PRC has been trapped in a deflationary cycle in recent years. By being pegged to the U.S. dollar, it would be subject to the same kinds of import price inflationary pressures as the U.S. and the HK-PRC. That said, it is very much insulated. Given the size and scope of its links with the U.S. and PRC, one could conjecture that the natural cointegration that we see in Table A10 is very much a reflection of the peg in its exchange rate.

Singapore has often pursued an exchange rate policy aimed at keeping import price pressures at bay. The open nature of its economy and its trade performance also depends on export prices be kept at competitive levels. It appears that the policy has been helpful in this regard, if the picture of balance is anything to go by.

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7. Tables

Table A1: Unit root tests for logs each of the terms of trade (TOT), import price index (IMP), and export price index (EXP).

	TOT			IMP			EXP		
	ADF	ADF^{GLS}	$\overline{MZ}_{\alpha}^{GLS}$	*ADF*	ADF^{GLS}	$\overline{MZ}_{\alpha}^{GLS}$	*ADF*	ADF^{GLS}	$\overline{MZ}_{\alpha}^{GLS}$
Spain	−1.9327	−0.7906	−1.3296	−2.884*	0.85238	0.61913	−2.79*	1.8529	0.87195
Japan	−1.7102	−0.14619	−0.20404	−1.9772	−0.81279	−1.5975	−1.4705	−1.4584	−5.0467
Korea	−2.1677	0.66681	0.8766	−3.4544**	1.483	0.78633	−3.2195**	1.2333	0.6071
Singapore	1.698	2.0704	2.474	−1.6671	−1.0636	−2.3738	−0.39476	−0.31221	−0.51397
Taiwan	−1.7168	−1.7509*	−6.8371*	−2.2382	−0.033219	−0.041853	−2.5952*	0.058341	0.065719
U. S.	−4.0856***	−2.3013**	−11.187**	−2.6745*	−0.87103	−2.1432	−1.6947	−0.36695	−1.5963
Mexico	−1.2002	−1.0329	−2.5394	−3.1774**	1.4328	0.78135	−2.5394	0.14723	0.16885
Belgium	−1.885	−0.96503	−1.9623	−1.5743	−0.42291	−0.88439	−1.4774	−0.15694	−0.29656
Brazil	−2.1726	−0.75518	−1.359	−2.8328*	−0.70764	−1.3051	−2.7859*	−1.8026*	−6.4392*
Finland	−0.26688	−0.41384	−0.83987	−1.0119	0.071507	0.093523	−1.8878	−0.79496	−1.5473
HK, PRC	−4.0401***	−1.3194	−4.6312	−2.9832**	−0.13299	−0.08339	−1.8442	−0.10705	0.12529

Note: *,**,*** indicate statistics are significant at the 10%, 5% and 1% levels respectively. ADF refers to the augmented Dickey-Fuller test of Said and Dickey (1984). ADF^{GLS} is Ng and Perron's (2001) modified form of the standard ADF test, based on local GLS detrending. $\overline{MZ}_{\alpha}^{GLS}$ is Ng and Perron's (2001) test that is a form of Phillips and Perron's (1988) Z_{α} test statistic, modified to have good size and power properties. For a one-sided (left-tailed) test of the null of nonstationarity, critical values at 1, 5 and 10 percent levels are taken as -3.46, -2.91 and -2.59 for *ADF*, -2.58, -1.98 and -1.62 for ADF^{GLS}, and -13.8, -8.1 and -5.7 for $\overline{MZ}_{\alpha}^{GLS}$.

Table A2. Values of the Lobato and Robinson (1998) LM test statistic for levels and differences of logarithms of series.

		Levels						First-Differences					
	m	10		20		30		10		20		30	
		stat	p-val	stat	p-val	stat	p-val	stat	p-val	Stat	p-val	Stat	p-val
Spain	TOT	−3.5788	0.0003	−7.4504	0	−10.962	0	1.4057	0.1598	1.8782	0.0604	2.1967	0.028
	IMP	−3.5948	0.0003	−7.4922	0	−10.9921	0	0.2423	0.8085	0.289	0.7726	0.9581	0.338
	EXP	−3.5512	0.0004	−7.4089	0	−10.8897	0	−0.7963	0.4258	0.8127	0.4164	1.2203	0.2224
Japan	TOT	−3.4865	0.0005	−7.3574	0	−10.8814	0	−1.1427	0.2532	−2.5057	0.0122	−3.6359	0.0003
	IMP	−3.8159	0.0001	−7.8729	0	−11.5353	0	−1.5578	0.1193	−2.5694	0.0102	−4.2737	0
	EXP	−3.6979	0.0002	−7.6906	0	−11.3035	0	0.0641	0.9489	−0.8829	0.3773	−2.0967	0.036
Korea	TOT	−4.2577	0	−8.6073	0	−12.4812	0	−1.2023	0.2293	−2.2653	0.0235	−2.7382	0.0062
	IMP	−3.5035	0.0005	−7.3692	0	−10.8634	0	−0.1308	0.8959	−1.5855	0.1128	−3.2543	0.0011
	EXP	−3.6087	0.0003	−7.5253	0	−11.0666	0	0.5096	0.6104	−0.8079	0.4192	−2.2329	0.0256
Sgp	TOT	−1.8474	0.0647	−4.4008	0	−6.5675	0	−0.298	0.7657	0.9839	0.3252	0.9736	0.3303
	IMP	−3.5191	0.0004	−7.3297	0	−10.6757	0	−1.0798	0.2802	−2.0962	0.0361	−3.4667	0.0005
	EXP	−3.5172	0.0004	−7.3247	0	−10.6689	0	−0.8495	0.3956	−1.5016	0.1332	−3.0617	0.0022
Taiwan	TOT	−3.508	0.0005	−7.3401	0	−10.7578	0	−0.364	0.7158	−1.1085	0.2676	−2.0494	0.0404
	IMP	−3.7288	0.0002	−7.6757	0	−11.1738	0	−0.8485	0.3961	−1.9065	0.0566	−2.1663	0.0303
	EXP	−3.6464	0.0003	−7.5453	0	−11.0106	0	−0.5284	0.5972	0.1938	0.8463	0.0899	0.9284
US	TOT	−3.5689	0.0004	−7.0584	0	4.2034	0	−0.1414	0.8876	−1.8318	0.067	0.0166	0.9868
	IMP	−3.4993	0.0005	−6.9443	0	4.2473	0	−0.7414	0.4585	−2.5969	0.0094	−0.0393	0.9686
	EXP	−3.5292	0.0004	−6.992	0	4.2339	0	−0.7954	0.4264	−0.5268	0.5984	0.0076	0.994
Mexico	TOT	−3.5279	0.0004	−7.3669	0	−10.8512	0	0.1587	0.8739	−0.8137	0.4158	−1.2333	0.2175
	IMP	−3.6366	0.0003	−7.5691	0	−11.1078	0	−1.3162	0.1881	−2.5864	0.0097	−1.0312	0.3024
	EXP	−3.5663	0.0004	−7.4369	0	−10.9414	0	−0.2171	0.8281	−1.4598	0.1444	−1.8798	0.0601
Belgium	TOT	−3.4655	0.0005	−1.6482	0.0993	3.0069	0.0026	1.1388	0.2548	−0.6455	0.5186	−0.2921	0.7702
	IMP	−3.4184	0.0006	−1.624	0.1044	3.0088	0.0026	1.2967	0.1947	−0.1424	0.8868	−1.0387	0.299
	EXP	−3.4415	0.0006	−1.6359	0.1019	3.0083	0.0026	1.5697	0.1165	1.1214	0.2621	1.7202	0.0854
Brazil	TOT	−1.5406	0.1234	−2.9964	0.0027	−5.1141	0	1.1341	0.2567	1.2352	0.2168	0.4498	0.6529
	IMP	−3.4656	0.0005	−7.2701	0	−10.6193	0	0.7701	0.4413	1.0899	0.2758	0.5549	0.5789
	EXP	−3.5072	0.0005	−7.3047	0	−10.649	0	0.0704	0.9439	−0.4198	0.6747	−0.4508	0.6521
Finland	TOT	−3.9791	0.0001	−8.034	0	−10.6737	0	0.2644	0.7915	0.2099	0.8338	0.4204	0.6742
	IMP	−3.6307	0.0003	−7.4086	0	−10.1013	0	−0.3029	0.762	−0.8525	0.3939	−1.0228	0.3064
	EXP	−3.5734	0.0004	−7.3183	0	−9.997	0	−0.589	0.5558	−0.606	0.5445	−0.3298	0.7415
HK	TOT	−3.5684	0.0004	−7.3055	0	−9.9909	0	0.5467	0.5846	1.069	0.2851	0.7106	0.4773
	IMP	−3.5466	0.0004	−7.2805	0	−9.9719	0	−0.0874	0.9304	−1.4636	0.1433	−1.9846	0.0472
	EXP	−3.5384	0.0004	−7.2643	0	−9.9503	0	−1.1257	0.2603	−2.8002	0.0051	−4.1621	0

Note: P-values computed based on asymptotic normal distribution following Lobato and Robinson (1998. Further, *m* determines choice of bandwidth for computation. TOT, IMP and EXP indicate logarithms of terms of trade, import prices and export prices respectively. Figures for UK omitted.

Table A3. Johansen's (1995) tests of cointegration.

	Trace	λ_{max}	Cointegrating relation (std err in brackets)	Test options
Spain	34.53701**	28.57959**	$EX - \underset{(0.06253)}{1.122485} - \underset{(0.26832)}{0.787055}\, IM$	II
Japan	8.839929	6.578768	None	V
Korea	11.91748	10.40785	None	V
Singapore	38.52084**	32.25672**	$EX - \underset{(0.17679)}{0.449179}\, IM + \underset{(0.00020)}{0.002077}\, t$	IV
Taiwan	11.31540	6.732403	None	V
U. S.	14.52024*	13.59680*	$EX - \underset{(0.00090)}{1.006940}\, IM$	I
Mexico	16.54136*	15.48275*	$EX - \underset{(0.01662)}{0.839541}\, IM$	I
Belgium	18.84771*	15.58300*	$EX - \underset{(0.01325)}{0.819815}\, IM$	III
Brazil	29.33596	20.77962	$EX - \underset{(3.41821)}{15.06905}\, IM - \underset{(0.00536)}{0.018474}\, t$	IV
Finland	27.30779*	17.66697	$EX - \underset{(0.23622)}{1.117262}\, IM + \underset{(0.00051)}{0.001683}\, t$	IV
HK, PRC	16.27926*	15.91562*	$EX - \underset{(0.00050)}{1.000930}\, IM$	I

Notes:

1. *(**) denotes rejection of the hypothesis at the 5%(1%) level. Decisions indicated are based on the 5% and 1% critical values (Osterwald-Lenum, 1992) for λ_{max} and the trace statistic.
2. Johansen's (1995) test may be carried out with five possible choice of options governing the possible inclusion of intercept and trends in the cointegrating equation (CE) and test VAR. These are: I=no intercept or trend in CE or VAR, II=intercept (no trend) in CE and no intercept in VAR, III=intercept (no trend) in CE and VAR, IV=intercept and trend in CE and no trend in VAR, and V=intercept and trend in CE and linear trend in VAR. *t* denotes a month-based trend dummy.
3. In the case of Japan, Korea and Taiwan, the tests are reported from option V but insignificant results are obtained with any choice of test option.

Table B1: Modified Log-Periodogram (Geweke and Porter-Hudak, 1983) Estimates of *d* for logged Terms of Trade Series

	$\alpha = 0.50$	$\alpha = 0.55$	α=0.60	α=0.65	α=0.70	α=0.75	α=0.80
Spain	1.178	1.255	1.315	1.067	0.972	0.931	0.885
70:01-03:12	*0.181*	*0.15*	*0.128*	*0.105*	*0.088*	*0.075*	*0.065*
Japan	1.046	1.075	1.069	1.072	1.105	1.057	1.056
60:01-04:02	*0.166*	*0.137*	*0.113*	*0.095*	*0.079*	*0.067*	*0.057*
Korea	0.594	0.739	0.898	0.871	0.921	0.947	0.939
71:01-04:02	*0.187*	*0.153*	*0.126*	*0.106*	*0.089*	*0.076*	*0.065*
Singapore	1.048	1.048	1.131	1.058	1.042	1.006	0.993
78:01-04:01	*0.202*	*0.166*	*0.138*	*0.117*	*0.099*	*0.084*	*0.073*
Taiwan	0.957	0.898	0.971	1.109	1.196	1.183	1.138
76:01-04:02	*0.194*	*0.161*	*0.135*	*0.112*	*0.096*	*0.082*	*0.07*
United States	0.922	1.071	1.036	1.046	1.048	1.034	1.001
88:01-03:03	*0.181*	*0.15*	*0.124*	*0.104*	*0.087*	*0.074*	*0.064*
Mexico	0.958	0.994	0.983	0.928	1.069	1.108	1.14
70:01-03:12	*0.181*	*0.15*	*0.128*	*0.105*	*0.088*	*0.075*	*0.065*
Belgium	1.021	0.814	0.96	0.909	0.937	0.981	0.859
93:01-03:12	*0.293*	*0.237*	*0.197*	*0.173*	*0.145*	*0.128*	*0.113*
Brazil	1.096	1.003	1.006	0.822	0.827	0.878	0.802
78:01-04:02	*0.202*	*0.166*	*0.138*	*0.117*	*0.099*	*0.084*	*0.073*
Finland	1.029	1.047	0.986	1.042	1.095	1.077	1.032
85:01-04:02	*0.22*	*0.188*	*0.154*	*0.131*	*0.113*	*0.096*	*0.084*
HK, PRC	0.625	0.724	0.624	0.676	0.908	0.959	0.939
83:01-02:03	*0.22*	*0.188*	*0.154*	*0.131*	*0.112*	*0.096*	*0.085*
UK	1.033	1.111	1.021	0.953	0.882	0.909	0.917
80:01-04:01	*0.202*	*0.171*	*0.144*	*0.12*	*0.102*	*0.087*	*0.076*

Note: Standard errors provided in italics below the corresponding estimate.

Table B2: Robust Gaussian Semiparametric (Robinson, 1995) Estimates of *d* for logarithms of Terms of Trade Series

	$\alpha = 0.50$	$\alpha = 0.55$	α=0.60	α=0.65	α=0.70	α=0.75	α=0.80
Spain	1.133	1.167	0.956	0.92	0.915	0.879	0.853
70:01-03:12	*0.112*	*0.096*	*0.085*	*0.072*	*0.062*	*0.053*	*0.046*
Japan	0.999	1.049	1.091	1.063	1.072	1.057	1.07
60:01-04:02	*0.104*	*0.09*	*0.076*	*0.066*	*0.056*	*0.048*	*0.041*
Korea	0.826	0.961	1.063	0.997	1.025	1.015	0.99
71:01-04:02	*0.115*	*0.098*	*0.083*	*0.072*	*0.062*	*0.053*	*0.046*
Singapore	1.034	1.02	1.01	0.992	0.989	0.966	0.938
78:01-04:01	*0.121*	*0.104*	*0.09*	*0.078*	*0.067*	*0.058*	*0.05*
Taiwan	0.976	0.965	0.936	0.992	1.084	1.087	1.071
76:01-04:02	*0.118*	*0.102*	*0.088*	*0.075*	*0.066*	*0.057*	*0.049*
United States	0.873	0.969	0.945	0.96	1.003	1.012	0.977
88:01-03:03	*0.112*	*0.096*	*0.082*	*0.071*	*0.061*	*0.052*	*0.045*
Mexico	0.907	0.96	0.958	0.918	0.979	1.047	1.05
70:01-03:12	*0.112*	*0.096*	*0.085*	*0.072*	*0.062*	*0.053*	*0.046*
Belgium	0.94	0.651	0.761	0.761	0.849	0.898	0.708
93:01-03:12	*0.158*	*0.139*	*0.121*	*0.109*	*0.094*	*0.085*	*0.075*
Brazil	1	1.004	0.911	0.817	0.808	0.848	0.797
78:01-04:02	*0.121*	*0.104*	*0.09*	*0.078*	*0.067*	*0.058*	*0.05*
Finland	1.072	1.079	1.039	1.06	1.042	1.035	0.952
85:01-04:02	*0.129*	*0.115*	*0.098*	*0.086*	*0.075*	*0.065*	*0.057*
HK, PRC	0.777	0.973	0.815	0.857	1.025	1.012	0.991
83:01-02:03	*0.129*	*0.115*	*0.098*	*0.086*	*0.075*	*0.065*	*0.057*
UK	0.923	1.057	0.885	0.815	0.785	0.834	0.84
80:01-04:01	*0.121*	*0.107*	*0.093*	*0.08*	*0.069*	*0.06*	*0.052*

Note: Standard errors provided in italics below the corresponding estimate.

Table B3: AIC of fitted FGN and ARFIMA Models for logs of Terms of Trade Series

	FGN	ARFIMA								
		(0,d,0)	(1,d,0)	(2,d,0)	(0,d,1)	(1,d,1)	(2,d,1)	(0,d,2)	(1,d,2)	(2,d,2)
Spain	2.091	2.086	4.084	6.084	4.086	6.084	8.084	6.083	8.084	10.084
Japan	2.026	2.026	4.026	6.026	4.026	6.026	8.026	6.026	8	10
Korea	2.136	2.137	4.137	6	4.134	6.134	8.134	6.134	8	10
Singapore	2.066	2.067	4.067	6	4.066	6.066	8.066	6.066	8	10
Taiwan	2.057	2.058	4.058	6.058	4.057	6.057	8.057	6.057	8.057	10.057
U. S.	2.08	2.08	4.08	6	4.079	6.079	8.079	6.079	8.078	10.079
Mexico	2.028	2.028	4.028	6	4.028	6.028	8.027	6.028	8.028	10
Belgium	2.549	2.495	4.485	6	4.488	6.475	8.438	6.46	8.451	10.437
Brazil	2.154	2.16	4.156	6	4.158	6.155	8.153	6.157	8.156	10.153
Finland	2.059	2.059	4.059	6	4.059	6.058	8.058	6.058	8.058	10
HK, PRC	2.261	2.271	4.27	6.268	4.258	6.258	8.257	6.247	8.247	10.247
UK	2.238	2.242	4.241	6.239	4.238	6.238	8.237	6.233	8.232	10.229

Table B4: Estimates for FGN and ARFIMA(0,*d*,0) Model for logs of Terms of Trade Series

	FGN(*d*)		ARFIMA(0,*d*,0)	
	d	s.e.	*d*	s.e.
Spain	0.701	0.034	0.811	0.039
Japan	0.948	0.03	1.11	0.034
Korea	0.88	0.035	1.031	0.039
Singapore	0.841	0.039	0.971	0.044
Taiwan	0.992	0.038	1.17	0.043
U. S.	0.877	0.034	1.027	0.038
Mexico	0.979	0.034	1.162	0.039
Belgium	0.476	0.059	0.686	0.068
Brazil	0.78	0.039	0.884	0.044
Finland	0.871	0.046	1.001	0.052
HK, PRC	0.885	0.046	1.055	0.052
UK	0.737	0.041	0.866	0.046

Table B5: Estimates of τ_α for logs of Terms of Trade Series

	0.3	0.5	0.8	0.9	0.95	0.99
Spain	3	19	>100 yrs			
Japan	>100yrs					
Korea	>100yrs					
Singapore	>100yrs					
Taiwan	>100yrs					
U. S.	>100yrs					
Mexico	>100yrs					
Belgium	1	4	70	631	>100 yrs.	.
Brazil	11	201	>100 yrs	.	.	.
Finland	>100 yrs					
HK, PRC	>100 yrs					
UK	7	87	>100 yrs			

Note: Values indicate the number of months required for the proportion α of the effect of a unit shock to dissipate.

Table B6: Bootstrap Confidence Intervals of the Impulse Response Functions of ARFIMA(0,*d*,0) Models for logs of Terms of Trade Series

	Level	1	3	6	12	24	48	96	192	384	504	744	984	1200
Spain	2.5%	0.728	0.572	0.482	0.402	0.334	0.278	0.23	0.191	0.158	0.147	0.132	0.122	0.116
70:01-03:12	97.5%	0.917	0.854	0.811	0.768	0.726	0.686	0.648	0.612	0.577	0.565	0.547	0.534	0.525
Japan	2.5%	0.978	0.96	0.947	0.934	0.92	0.906	0.893	0.879	0.866	0.861	0.854	0.848	0.845
60:01-04:02	97.5%	1.139	1.275	1.387	1.518	1.667	1.832	2.016	2.219	2.444	2.538	2.679	2.785	2.863
Korea	2.5%	0.932	0.88	0.844	0.807	0.771	0.736	0.703	0.671	0.64	0.628	0.612	0.6	0.592
71:01-04:02	97.5%	1.084	1.16	1.221	1.289	1.363	1.443	1.529	1.619	1.716	1.755	1.813	1.856	1.887
Singapore	2.5%	0.87	0.779	0.718	0.659	0.604	0.552	0.505	0.462	0.422	0.408	0.388	0.374	0.364
78:01-04:01	97.5%	1.046	1.087	1.118	1.152	1.188	1.226	1.266	1.307	1.349	1.366	1.391	1.409	1.422
Taiwan	2.5%	0.97	0.947	0.929	0.912	0.894	0.876	0.858	0.841	0.824	0.817	0.808	0.801	0.797
76:01-04:02	97.5%	1.194	1.395	1.569	1.778	2.025	2.311	2.641	3.02	3.454	3.641	3.926	4.145	4.308
U.S.	2.5%	0.934	0.884	0.849	0.813	0.778	0.744	0.711	0.68	0.649	0.638	0.622	0.611	0.603
88:01-03:03	97.5%	1.079	1.152	1.209	1.273	1.343	1.418	1.498	1.582	1.671	1.708	1.761	1.801	1.829
Mexico	2.5%	0.977	0.958	0.945	0.931	0.917	0.902	0.888	0.874	0.86	0.855	0.847	0.842	0.838
70:01-03:12	97.5%	1.186	1.377	1.54	1.737	1.967	2.233	2.537	2.885	3.281	3.45	3.709	3.907	4.054
Belgium	2.5%	0.508	0.32	0.232	0.167	0.119	0.085	0.06	0.043	0.031	0.027	0.022	0.019	0.017
93:01-03:12	97.5%	0.832	0.72	0.648	0.58	0.518	0.462	0.411	0.366	0.326	0.312	0.292	0.279	0.27
Brazil	2.5%	0.786	0.653	0.57	0.495	0.429	0.37	0.32	0.276	0.238	0.224	0.207	0.195	0.187
78:01-04:02	97.5%	0.987	0.976	0.967	0.959	0.95	0.942	0.933	0.924	0.916	0.913	0.908	0.904	0.902
Finland	2.5%	0.874	0.784	0.725	0.667	0.613	0.562	0.515	0.472	0.433	0.418	0.398	0.385	0.375
85:01-04:02	97.5%	1.078	1.149	1.205	1.267	1.335	1.408	1.486	1.568	1.654	1.69	1.742	1.78	1.808
HK, PRC	2.5%	0.911	0.844	0.799	0.753	0.709	0.667	0.628	0.59	0.555	0.542	0.523	0.51	0.501
83:01-02:03	97.5%	1.117	1.229	1.32	1.424	1.541	1.669	1.809	1.961	2.127	2.196	2.298	2.375	2.431
UK	2.5%	0.762	0.618	0.532	0.454	0.387	0.329	0.279	0.237	0.201	0.188	0.171	0.16	0.153
80:01-04:01	97.5%	0.969	0.944	0.927	0.908	0.89	0.871	0.853	0.835	0.817	0.81	0.801	0.794	0.789

Note: All bootstrap results (incl those in later tables) based on 5000 replications. The heading row gives the number of months for which the impulse responses apply.

Table B7: Modified Log-Periodogram (Geweke and Porter-Hudak, 1983) Estimates of *d* for growth rate of Terms of Trade Series

	$\alpha = 0.50$	$\alpha = 0.55$	α=0.60	α=0.65	α=0.70	α=0.75	α=0.80
Spain	0.255	0.26	0.267	0.063	0.016	-0.011	-0.078
70:01-03:12	*0.181*	*0.15*	*0.126*	*0.105*	*0.088*	*0.075*	*0.065*
Japan	0.116	0.12	0.201	0.226	0.275	0.332	0.389
60:01-04:02	*0.166*	*0.137*	*0.113*	*0.095*	*0.079*	*0.067*	*0.057*
Korea	0.146	0.134	0.183	0.124	0.147	0.174	0.262
71:01-04:02	*0.187*	*0.153*	*0.126*	*0.106*	*0.09*	*0.076*	*0.066*
Singapore	0.384	0.422	0.166	0.164	0.097	0.031	-0.033
78:01-04:01	*0.202*	*0.166*	*0.14*	*0.117*	*0.1*	*0.085*	*0.074*
Taiwan	0.022	0.029	0.08	0.126	0.226	0.217	0.205
76:01-04:02	*0.194*	*0.161*	*0.135*	*0.113*	*0.096*	*0.082*	*0.07*
United States	-0.073	0.05	0.011	0.071	0.117	0.125	0.122
88:01-03:03	*0.181*	*0.15*	*0.124*	*0.104*	*0.087*	*0.074*	*0.064*
Mexico	-0.211	-0.151	-0.038	-0.132	-0.051	0.076	0.099
70:01-03:12	*0.181*	*0.15*	*0.126*	*0.105*	*0.088*	*0.075*	*0.065*
Belgium	0.244	-0.094	-0.072	-0.07	-0.085	-0.018	-0.183
93:01-03:12	*0.273*	*0.231*	*0.196*	*0.168*	*0.144*	*0.126*	*0.111*
Brazil	-0.026	-0.015	0.022	-0.175	-0.152	-0.145	-0.195
78:01-04:02	*0.202*	*0.166*	*0.138*	*0.117*	*0.099*	*0.084*	*0.073*
Finland	0.303	0.312	0.079	0.032	0.098	0.147	0.072
85:01-04:02	*0.22*	*0.188*	*0.154*	*0.131*	*0.113*	*0.097*	*0.085*
HK, PRC	-0.066	-0.055	-0.134	0.012	0.05	0.13	0.072
83:01-02:03	*0.22*	*0.188*	*0.154*	*0.131*	*0.113*	*0.096*	*0.084*
UK	0.055	0.116	-0.065	-0.085	-0.135	-0.076	-0.045
80:01-04:01	*0.21*	*0.171*	*0.146*	*0.121*	*0.103*	*0.088*	*0.076*

Note: Standard errors provided in italics below the corresponding estimate. Growth rate of terms of trade refers to the difference of the logarithms of the series.

Table B8: Robust Gaussian Semiparametric (Robinson, 1995) Estimates of *d* for growth rate of Terms of Trade Series

	$\alpha = 0.50$	$\alpha = 0.55$	α=0.60	α=0.65	α=0.70	α=0.75	α=0.80
Spain	0.198	0.226	0.018	-0.022	-0.036	-0.074	-0.101
70:01-03:12	*0.112*	*0.096*	*0.083*	*0.071*	*0.061*	*0.053*	*0.045*
Japan	0.024	0.094	0.212	0.201	0.243	0.305	0.388
60:01-04:02	*0.104*	*0.09*	*0.076*	*0.066*	*0.056*	*0.048*	*0.041*
Korea	0.13	0.201	0.209	0.16	0.169	0.194	0.25
71:01-04:02	*0.115*	*0.098*	*0.083*	*0.072*	*0.062*	*0.053*	*0.046*
Singapore	0.239	0.244	0.098	0.087	0.089	-0.014	-0.039
78:01-04:01	*0.121*	*0.104*	*0.091*	*0.079*	*0.069*	*0.059*	*0.051*
Taiwan	0.043	0.063	-0.002	0.038	0.117	0.136	0.134
76:01-04:02	*0.118*	*0.102*	*0.088*	*0.076*	*0.066*	*0.057*	*0.049*
United States	-0.135	-0.05	-0.07	-0.034	0.078	0.119	0.108
88:01-03:03	*0.112*	*0.096*	*0.082*	*0.071*	*0.061*	*0.052*	*0.045*
Mexico	-0.174	-0.109	-0.067	-0.131	-0.05	0.048	0.088
70:01-03:12	*0.112*	*0.096*	*0.083*	*0.071*	*0.061*	*0.053*	*0.045*
Belgium	0.002	-0.375	-0.294	-0.269	-0.187	-0.118	-0.261
93:01-03:12	*0.151*	*0.134*	*0.118*	*0.104*	*0.091*	*0.081*	*0.071*
Brazil	0.031	0.045	-0.036	-0.133	-0.132	-0.103	-0.144
78:01-04:02	*0.121*	*0.104*	*0.09*	*0.078*	*0.067*	*0.058*	*0.05*
Finland	0.042	0.098	-0.007	0.008	0.076	0.117	0.048
85:01-04:02	*0.129*	*0.115*	*0.098*	*0.086*	*0.075*	*0.066*	*0.057*
HK, PRC	-0.078	0.018	-0.077	-0.006	0.091	0.12	0.057
83:01-02:03	*0.129*	*0.115*	*0.098*	*0.086*	*0.075*	*0.065*	*0.057*
UK	-0.052	0.098	-0.119	-0.163	-0.195	-0.132	-0.113
80:01-04:01	*0.125*	*0.107*	*0.094*	*0.081*	*0.071*	*0.061*	*0.053*

Note: Standard errors provided in italics below the corresponding estimate.

Table B9: AIC of fitted FGN and ARFIMA Models for growth rate of Terms of Trade Series

	FGN	ARFIMA								
		(0,d,0)	(1,d,0)	(2,d,0)	(0,d,1)	(1,d,1)	(2,d,1)	(0,d,2)	(1,d,2)	(2,d,2)
Spain	3.885	3.89	5.89	7.869	5.89	7.891	9.869	7.868	9.868	11.868
Japan	3.364	3.403	5.354	7.345	5.337	7.334	9.334	7.337	9.333	11.333
Korea	3.782	3.795	5.782	7.776	5.799	7.8	9.8	7.786	9.786	11.786
Singapore	4.03	4.006	6.004	8.005	6.005	8.004	10.005	8.006	10.005	12.005
Taiwan	3.866	3.882	5.849	7.848	5.864	7.864	9.864	7.846	9.858	11.858
U. S.	3.999	3.993	5.993	7.986	5.993	7.993	9.986	7.983	9.983	11.984
Mexico	3.883	3.908	5.841	7.822	5.812	7.812	9.812	7.812	9.812	11.812
Belgium	3.725	3.679	5.679	7.537	5.679	7.68	9.537	7.58	9.58	11.537
Brazil	4.004	3.982	5.93	7.93	5.955	7.93	9.93	7.956	9.93	11.931
Finland	4.014	3.999	5.999	7.997	5.999	7.999	9.997	8.012	10.012	11.997
HK, PRC	4.035	4.006	6.005	7.999	6.032	8.032	10.018	7.955	9.955	11.955
UK	4.006	3.97	5.965	7.948	5.964	7.964	9.948	7.958	9.952	11.948

Table B10: Estimates for FGN and ARFIMA(0,*d*,0) Model for growth rate of Terms of Trade Series

	FGN(*d*)		ARFIMA(0,*d*,0)	
	d	s.e.	*d*	s.e.
Spain	-0.149	0.028	-0.179	0.039
Japan	0.409	0.029	0.49	0.034
Korea	0.216	0.033	0.255	0.039
Singapore	0.002	0.035	-0.007	0.044
Taiwan	0.184	0.036	0.207	0.043
U. S.	0.043	0.031	0.049	0.038
Mexico	0.187	0.032	0.205	0.039
Belgium	-0.259	0.045	-0.378	0.068
Brazil	-0.044	0.035	-0.084	0.044
Finland	0.022	0.041	0.015	0.052
HK, PRC	0.036	0.042	0.034	0.052
UK	-0.079	0.035	-0.123	0.046

Table B11: Estimates of τ_{α} for growth rate of Terms of Trade Series

	0.3	0.5	0.8	0.9	0.95	0.99
Spain	1	1	1	2	3	11
Japan	1	1	8	29	111	>100 yrs
Korea	1	1	2	4	11	89
Singapore	1	1	1	1	1	1
Taiwan	1	1	2	3	7	51
U. S.	1	1	1	1	1	6
Mexico	1	1	2	3	7	50
Belgium	1	1	2	3	4	11
Brazil	1	1	1	1	2	7
Finland	1	1	1	1	1	2
HK, PRC	1	1	1	1	1	4
UK	1	1	1	2	3	9

Note: Values indicate the number of months required for the proportion α of the effect of a unit shock to dissipate.

Table B12: Bootstrap Confidence Intervals of the Impulse Response Functions of ARFIMA(0,*d*,0) Model for growth rate of Terms of Trade Series

	Level	1	3	6	12	24	48	96	192	384	504	744	984	1200
Spain	2.5%	-0.27	-0.057	-0.023	-0.009	-0.004	-0.002	-0.001	0	0	0	0	0	0
70:01-03:12	97.5%	-0.116	-0.032	-0.015	-0.007	-0.003	-0.001	-0.001	0	0	0	0	0	0
Japan	2.5%	0.41	0.232	0.157	0.105	0.07	0.047	0.031	0.021	0.014	0.012	0.009	0.008	0.007
60:01-04:02	97.5%	0.499	0.312	0.225	0.16	0.114	0.081	0.057	0.04	0.029	0.025	0.021	0.018	0.016
Korea	2.5%	0.142	0.058	0.032	0.018	0.01	0.005	0.003	0.002	0.001	0.001	0.001	0	0
71:01-04:02	97.5%	0.351	0.186	0.121	0.078	0.05	0.032	0.02	0.013	0.008	0.007	0.005	0.004	0.004
Singapore	2.5%	-0.129	-0.035	-0.016	-0.007	-0.003	-0.001	-0.001	0	0	0	0	0	0
78:01-04:01	97.5%	0.068	0.025	0.013	0.007	0.004	0.002	0.001	0.001	0	0	0	0	0
Taiwan	2.5%	0.101	0.039	0.021	0.011	0.006	0.003	0.002	0.001	0.001	0	0	0	0
76:01-04:02	97.5%	0.291	0.144	0.089	0.055	0.034	0.021	0.013	0.008	0.005	0.004	0.003	0.002	0.002
U.S.	2.5%	-0.04	-0.013	-0.006	-0.003	-0.001	-0.001	0	0	0	0	0	0	0
88:01-03:03	97.5%	0.108	0.042	0.023	0.012	0.007	0.004	0.002	0.001	0.001	0	0	0	0
Mexico	2.5%	0.112	0.044	0.024	0.013	0.007	0.004	0.002	0.001	0.001	0	0	0	0
70:01-03:12	97.5%	0.269	0.129	0.079	0.048	0.029	0.017	0.011	0.006	0.004	0.003	0.002	0.002	0.002
Belgium	2.5%	-0.499	-0.064	-0.023	-0.009	-0.004	-0.002	-0.001	0	0	0	0	0	0
93:01-03:12	97.5%	-0.259	-0.056	-0.021	-0.007	-0.002	-0.001	0	0	0	0	0	0	0
Brazil	2.5%	-0.209	-0.049	-0.021	-0.009	-0.004	-0.002	-0.001	0	0	0	0	0	0
78:01-04:02	97.5%	-0.017	-0.006	-0.003	-0.001	-0.001	0	0	0	0	0	0	0	0
Finland	2.5%	-0.12	-0.033	-0.015	-0.007	-0.003	-0.001	-0.001	0	0	0	0	0	0
85:01-04:02	97.5%	0.106	0.041	0.022	0.012	0.007	0.004	0.002	0.001	0.001	0	0	0	0
HK, PRC	2.5%	-0.11	-0.031	-0.014	-0.007	-0.003	-0.001	-0.001	0	0	0	0	0	0
83:01-02:03	97.5%	0.134	0.054	0.03	0.017	0.009	0.005	0.003	0.002	0.001	0.001	0	0	0
UK	2.5%	-0.238	-0.053	-0.022	-0.009	-0.004	-0.002	-0.001	0	0	0	0	0	0
80:01-04:01	97.5%	-0.047	-0.015	-0.007	-0.003	-0.002	-0.001	0	0	0	0	0	0	0

Note: The heading row gives the number of months for which the impulse responses apply.

Table C1: Modified Log-Periodogram (Geweke and Porter-Hudak, 1983) Estimates of *d* for Import Price Inflation Series

	$\alpha = 0.50$	$\alpha = 0.55$	α=0.60	α=0.65	α=0.70	α=0.75	α=0.80
Spain	0.337	0.435	0.381	0.347	0.296	0.198	0.165
70:01-03:12	(0.181)	(0.150)	(0.126)	(0.105)	(0.088)	(0.075)	(0.065)
Japan	0.183	0.226	0.259	0.272	0.289	0.360	0.367
60:01-04:02	(0.166)	(0.137)	(0.113)	(0.095)	(0.079)	(0.067)	(0.057)
Korea	0.307	0.250	0.342	0.167	0.154	0.135	0.258
71:01-04:02	(0.187)	(0.153)	(0.126)	(0.106)	(0.090)	(0.076)	(0.066)
Singapore	0.315	0.307	0.191	0.205	0.210	0.167	0.167
78:01-04:01	(0.202)	(0.166)	(0.140)	(0.117)	(0.100)	(0.085)	(0.074)
Taiwan	0.134	0.208	0.292	0.201	0.246	0.239	0.226
76:01-04:02	(0.194)	(0.161)	(0.135)	(0.113)	(0.096)	(0.082)	(0.070)
United States	0.107	0.069	0.007	0.084	0.039	0.185	0.219
88:01-03:03	(0.243)	(0.202)	(0.172)	(0.145)	(0.124)	(0.108)	(0.094)
Mexico	0.266	0.334	0.430	0.452	0.434	0.377	0.251
70:01-03:12	(0.181)	(0.150)	(0.126)	(0.105)	(0.088)	(0.075)	(0.065)
Belgium	0.755	0.519	0.278	0.119	0.176	0.142	0.010
93:01-03:12	(0.273)	(0.231)	(0.196)	(0.168)	(0.144)	(0.126)	(0.111)
Brazil	0.292	0.242	0.425	0.176	0.042	-0.094	-0.166
78:01-04:02	(0.202)	(0.166)	(0.138)	(0.117)	(0.099)	(0.084)	(0.073)
Finland	0.540	0.672	0.328	0.220	0.301	0.345	0.228
85:01-04:02	(0.220)	(0.188)	(0.154)	(0.131)	(0.113)	(0.097)	(0.085)
HK, PRC	0.441	0.576	0.442	0.498	0.518	0.476	0.402
83:01-02:03	(0.220)	(0.188)	(0.154)	(0.131)	(0.113)	(0.096)	(0.084)

Note: Standard errors provided in parentheses below the corresponding estimate.

Table C2: Robust Gaussian Semiparametric (Robinson, 1995) Estimates of d for Import Price Inflation Series

	$\alpha = 0.50$	$\alpha = 0.55$	α=0.60	α=0.65	α=0.70	α=0.75	α=0.80
Spain	0.299	0.404	0.318	0.244	0.191	0.128	0.121
70:01-03:12	(0.112)	(0.096)	(0.083)	(0.071)	(0.061)	(0.053)	(0.045)
Japan	0.120	0.203	0.307	0.277	0.316	0.360	0.351
60:01-04:02	(0.104)	(0.090)	(0.076)	(0.066)	(0.056)	(0.048)	(0.041)
Korea	0.219	0.172	0.211	0.099	0.113	0.120	0.227
71:01-04:02	(0.115)	(0.098)	(0.083)	(0.072)	(0.062)	(0.053)	(0.046)
Singapore	0.235	0.217	0.124	0.097	0.153	0.147	0.165
78:01-04:01	(0.121)	(0.104)	(0.091)	(0.079)	(0.069)	(0.059)	(0.051)
Taiwan	0.112	0.199	0.199	0.160	0.158	0.207	0.219
76:01-04:02	(0.118)	(0.102)	(0.088)	(0.076)	(0.066)	(0.057)	(0.049)
United States	0.150	-0.075	-0.069	0.062	0.074	0.180	0.253
88:01-03:03	(0.139)	(0.121)	(0.107)	(0.093)	(0.081)	(0.071)	(0.063)
Mexico	0.459	0.489	0.431	0.444	0.451	0.434	0.279
70:01-03:12	(0.112)	(0.096)	(0.083)	(0.071)	(0.061)	(0.053)	(0.045)
Belgium	0.389	0.189	0.115	0.079	0.172	0.189	0.009
93:01-03:12	(0.151)	(0.134)	(0.118)	(0.104)	(0.091)	(0.081)	(0.071)
Brazil	0.175	0.212	0.114	0.034	-0.014	-0.061	-0.107
78:01-04:02	(0.121)	(0.104)	(0.090)	(0.078)	(0.067)	(0.058)	(0.050)
Finland	0.379	0.520	0.170	0.117	0.216	0.245	0.181
85:01-04:02	(0.129)	(0.115)	(0.098)	(0.086)	(0.075)	(0.066)	(0.057)
HK, PRC	0.381	0.650	0.407	0.427	0.450	0.422	0.344
83:01-02:03	(0.129)	(0.115)	(0.098)	(0.086)	(0.075)	(0.065)	(0.057)

Note: Standard errors provided in parentheses below the corresponding estimate.

Table C3: AIC of fitted FGN and ARFIMA Models for Import Price Inflation Series

	FGN	ARFIMA								
		(0,d,0)	(1,d,0)	(2,d,0)	(0,d,1)	(1,d,1)	(2,d,1)	(0,d,2)	(1,d,2)	(2,d,2)
Spain	3.997	3.986	5.951	7.951	5.98	7.951	8	7.944	9.938	10
Japan	3.362	3.397	5.346	7.344	5.362	7.333	9.327	7.345	8	10
Korea	3.615	3.678	5.522	7.521	5.559	7.521	9.521	7.551	8	11.513
Singapore	3.724	3.758	5.604	7.598	5.67	7.597	9.583	7.656	9.583	11.583
Taiwan	3.793	3.804	5.795	7.79	5.783	7.79	9.789	7.778	9.777	10
U. S.	3.693	3.738	5.556	7.541	5.576	7.507	9.503	7.531	9.501	11.468
Mexico	3.869	3.836	5.684	7.64	5.833	7.674	9.628	7.688	9.647	11.618
Belgium	3.995	3.97	5.881	7.746	5.825	7.824	9.743	7.778	9.777	11.712
Brazil	3.999	3.982	5.974	7.834	5.98	6	8	7.847	9.805	10
Finland	3.883	3.884	5.883	7.882	5.883	7.856	9.853	7.856	9.851	11.845
HK, PRC	3.465	3.466	5.452	6	5.434	7.43	9.406	7.413	9.431	10

Table C4: Estimates for FGN and ARFIMA(0,*d*,0) for Import Price Inflation Series

	FGN		ARFIMA(0,*d*,0)	
	d	s.e.	*d*	s.e.
Spain	0.041	0.031	0.056	0.039
Japan	0.391	0.029	0.458	0.034
Korea	0.339	0.034	0.37	0.039
Singapore	0.282	0.038	0.308	0.044
Taiwan	0.207	0.036	0.243	0.043
U. S.	0.337	0.05	0.379	0.058
Mexico	0.12	0.032	0.158	0.039
Belgium	-0.073	0.053	-0.092	0.068
Brazil	-0.052	0.034	-0.076	0.044
Finland	0.16	0.043	0.185	0.052
HK, PRC	0.338	0.044	0.405	0.052

Table C5: Estimates of τ_α for Import Price Inflation Series

	0.3	0.5	0.8	0.9	0.95	0.99
Spain	1	1	1	1	2	7
Japan	1	1	6	21	74	>100 yrs
Korea	1	1	4	10	29	372
Singapore	1	1	3	6	17	166
Taiwan	1	1	2	4	10	77
U. S.	1	1	4	11	32	420
Mexico	1	1	1	2	5	29
Belgium	1	1	1	1	2	8
Brazil	1	1	1	1	2	7
Finland	1	1	1	3	6	40
HK, PRC	1	1	4	13	42	617

Note: Values indicate the number of months required for the proportion α of the effect of a unit shock to dissipate.

Table C6: Bootstrap Confidence Intervals of the Impulse Response Functions of ARFIMA(0,*d*,0) Models for Import Price Inflation Series

	Level	1	3	6	12	24	48	96	192	384	504	744	984	1200
Spain	2.5%	-0.044	-0.014	-0.007	-0.003	-0.002	-0.001	0	0	0	0	0	0	0
70:01-03:12	97.5%	0.119	0.047	0.026	0.014	0.008	0.004	0.002	0.001	0.001	0.001	0	0	0
Japan	2.5%	0.378	0.206	0.137	0.09	0.059	0.038	0.025	0.016	0.01	0.009	0.007	0.006	0.005
60:01-04:02	97.5%	0.523	0.335	0.246	0.178	0.129	0.093	0.067	0.048	0.035	0.03	0.025	0.022	0.02
Korea	2.5%	0.275	0.133	0.082	0.05	0.03	0.018	0.011	0.007	0.004	0.003	0.003	0.002	0.002
71:01-04:02	97.5%	0.442	0.259	0.18	0.123	0.084	0.057	0.039	0.026	0.018	0.015	0.012	0.011	0.01
Singapore	2.5%	0.194	0.085	0.049	0.028	0.016	0.009	0.005	0.003	0.002	0.001	0.001	0.001	0.001
78:01-04:01	97.5%	0.385	0.212	0.141	0.093	0.061	0.04	0.026	0.017	0.011	0.009	0.007	0.006	0.006
Taiwan	2.5%	0.139	0.056	0.031	0.017	0.01	0.005	0.003	0.002	0.001	0.001	0	0	0
76:01-04:02	97.5%	0.314	0.159	0.101	0.063	0.04	0.025	0.015	0.01	0.006	0.005	0.004	0.003	0.003
U.S.	2.5%	0.219	0.099	0.058	0.034	0.02	0.012	0.007	0.004	0.002	0.002	0.001	0.001	0.001
88:01-03:03	97.5%	0.485	0.298	0.213	0.151	0.106	0.074	0.052	0.037	0.026	0.022	0.018	0.016	0.014
Mexico	2.5%	0.061	0.022	0.012	0.006	0.003	0.002	0.001	0	0	0	0	0	0
70:01-03:12	97.5%	0.227	0.104	0.062	0.036	0.021	0.012	0.007	0.004	0.003	0.002	0.002	0.001	0.001
Belgium	2.5%	-0.295	-0.059	-0.023	-0.009	-0.004	-0.002	-0.001	0	0	0	0	0	0
93:01-03:12	97.5%	0.021	0.007	0.004	0.002	0.001	0	0	0	0	0	0	0	0
Brazil	2.5%	-0.191	-0.047	-0.02	-0.009	-0.004	-0.002	-0.001	0	0	0	0	0	0
78:01-04:02	97.5%	-0.005	-0.002	-0.001	0	0	0	0	0	0	0	0	0	0
Finland	2.5%	0.046	0.016	0.008	0.004	0.002	0.001	0.001	0	0	0	0	0	0
85:01-04:02	97.5%	0.269	0.129	0.079	0.048	0.029	0.018	0.011	0.006	0.004	0.003	0.002	0.002	0.002
HK, PRC	2.5%	0.267	0.128	0.078	0.048	0.029	0.017	0.01	0.006	0.004	0.003	0.002	0.002	0.002
83:01-02:03	97.5%	0.499	0.312	0.225	0.161	0.114	0.081	0.057	0.04	0.029	0.025	0.021	0.018	0.016

Note: The heading row gives the number of months for which the impulse responses apply.

Table D1: Modified Log-Periodogram (Geweke and Porter-Hudak, 1983) Estimates of *d* for Export Price Inflation Series

	$\alpha = 0.50$	$\alpha = 0.55$	α=0.60	α=0.65	α=0.70	α=0.75	α=0.80
Spain	0.578	0.391	0.324	0.179	0.087	0.034	-0.091
70:01-03:12	(0.181)	(0.15)	(0.126)	(0.105)	(0.088)	(0.075)	(0.065)
Japan	0.137	0.128	0.112	0.107	0.111	0.221	0.17
60:01-04:02	(0.166)	(0.137)	(0.113)	(0.095)	(0.079)	(0.067)	(0.057)
Korea	0.092	0.016	0.066	-0.018	0.003	0.038	0.193
71:01-04:02	(0.187)	(0.153)	(0.126)	(0.106)	(0.09)	(0.076)	(0.066)
Singapore	0.431	0.369	0.346	0.24	0.277	0.329	0.306
78:01-04:01	(0.202)	(0.166)	(0.14)	(0.117)	(0.1)	(0.085)	(0.074)
Taiwan	0.022	0.057	0.147	0.079	0.068	0.061	0.14
76:01-04:02	(0.194)	(0.161)	(0.135)	(0.113)	(0.096)	(0.082)	(0.07)
United States	0.391	0.551	0.622	0.588	0.568	0.442	0.331
88:01-03:03	(0.243)	(0.202)	(0.172)	(0.145)	(0.124)	(0.108)	(0.094)
Mexico	0.037	0.067	0.203	0.054	0.099	0.2	0.226
70:01-03:12	(0.181)	(0.15)	(0.126)	(0.105)	(0.088)	(0.075)	(0.065)
Belgium	0.706	0.547	0.287	0.239	0.19	0.227	0.066
93:01-03:12	(0.273)	(0.231)	(0.196)	(0.168)	(0.144)	(0.126)	(0.111)
Brazil	-0.02	0.166	0.178	0.086	0.143	0.104	0.088
78:01-04:02	(0.202)	(0.166)	(0.138)	(0.117)	(0.099)	(0.084)	(0.073)
Finland	0.066	0.315	0.31	0.423	0.332	0.327	0.292
85:01-04:02	(0.22)	(0.188)	(0.154)	(0.131)	(0.114)	(0.098)	(0.085)
HK, PRC	0.772	0.831	0.637	0.641	0.574	0.446	0.445
83:01-02:03	(0.22)	(0.188)	(0.154)	(0.131)	(0.113)	(0.096)	(0.084)

Note: Figures in parentheses are standard errors.

Table D2: Robust Gaussian Semiparametric (Robinson, 1995) Estimates of *d* for Export Price Inflation Series

	α=0.50	$\alpha = 0.55$	α=0.60	α=0.65	$\alpha = 0.70$	α=0.75	α= 0.80
Spain	0.483	0.369	0.21	0.113	0.071	0.044	-0.013
70:01-03:12	*0.112*	*0.096*	*0.083*	*0.071*	*0.061*	*0.053*	*0.045*
Japan	0.07	0.151	0.166	0.11	0.124	0.223	0.208
60:01-04:02	*0.104*	*0.09*	*0.076*	*0.066*	*0.056*	*0.048*	*0.041*
Korea	0.1	0.06	0.076	-0.015	-0.018	0.001	0.116
71:01-04:02	*0.115*	*0.098*	*0.083*	*0.072*	*0.062*	*0.053*	*0.046*
Singapore	0.311	0.334	0.256	0.175	0.218	0.284	0.284
78:01-04:01	*0.121*	*0.104*	*0.091*	*0.079*	*0.069*	*0.059*	*0.051*
Taiwan	0.12	0.172	0.236	0.145	0.071	0.098	0.138
76:01-04:02	*0.118*	*0.102*	*0.088*	*0.076*	*0.066*	*0.057*	*0.049*
United States	0.399	0.417	0.493	0.467	0.418	0.365	0.206
88:01-03:03	*0.139*	*0.121*	*0.107*	*0.093*	*0.081*	*0.071*	*0.063*
Mexico	-0.018	0.037	0.069	-0.034	0.031	0.115	0.16
70:01-03:12	*0.112*	*0.096*	*0.083*	*0.071*	*0.061*	*0.053*	*0.045*
Belgium	0.321	0.336	0.148	0.159	0.167	0.2	0.006
93:01-03:12	*0.151*	*0.134*	*0.118*	*0.104*	*0.091*	*0.081*	*0.071*
Brazil	0.069	0.246	0.179	0.031	0.072	0.061	0.076
78:01-04:02	*0.121*	*0.104*	*0.09*	*0.078*	*0.067*	*0.058*	*0.05*
Finland	0.088	0.249	0.314	0.474	0.338	0.31	0.26
85:01-04:02	*0.129*	*0.115*	*0.098*	*0.086*	*0.076*	*0.067*	*0.058*
HK, PRC	0.66	0.861	0.503	0.589	0.52	0.421	0.434
83:01-02:03	*0.129*	*0.115*	*0.098*	*0.086*	*0.075*	*0.065*	*0.057*

Note: Figures in italics are standard errors.

Table D3: AIC of fitted FGN and ARFIMA Models for Export Price Inflation Series

	FGN	ARFIMA								
		(0,d,0)	(1,d,0)	(2,d,0)	(0,d,1)	(1,d,1)	(2,d,1)	(0,d,2)	(1,d,2)	(2,d,2)
Spain	3.918	3.932	5.747	7.745	5.795	7.745	8	7.75	9.75	11.743
Japan	3.7	3.729	5.677	7.677	5.696	7.666	8	7.672	9.663	11.667
Korea	3.756	3.822	5.618	7.606	5.657	7.607	9.605	7.651	9.589	11.588
Singapore	3.629	3.64	5.613	7.613	5.608	7.594	9.593	7.596	8	10
Taiwan	3.918	3.924	5.91	7.908	5.911	7.902	8	7.896	9.902	11.902
U. S.	3.827	3.788	5.757	7.757	5.765	7.757	9.716	7.735	9.667	10
Mexico	3.804	3.834	5.774	7.763	5.756	7.756	9.756	7.749	9.749	10
Belgium	3.977	3.956	5.832	7.695	5.747	7.746	9.693	7.7	9.693	11.657
Brazil	4.003	3.994	5.993	7.98	5.993	7.952	9.976	7.951	9.938	11.938
Finland	3.792	3.787	5.785	7.785	5.752	7.735	9.734	7.74	9.734	10
HK, PRC	3.513	3.47	5.435	7.428	5.455	7.405	9.4	7.398	9.397	11.393

Table D4:. Estimates for FGN and ARFIMA(0,*d*,0) Model for Export Price Inflation Series

	FGN		ARFIMA(0,*d*,0)	
	d	s.e.	*d*	s.e.
Spain	-0.113	0.029	-0.121	0.039
Japan	0.27	0.029	0.312	0.034
Korea	0.294	0.033	0.302	0.039
Singapore	0.289	0.038	0.342	0.044
Taiwan	0.137	0.035	0.151	0.043
U. S.	0.184	0.049	0.228	0.058
Mexico	0.231	0.033	0.262	0.039
Belgium	-0.09	0.052	-0.11	0.068
Brazil	0.041	0.036	0.045	0.044
Finland	0.202	0.043	0.246	0.052
HK, PRC	0.272	0.044	0.344	0.052

Table D5: Estimates of τ_α for Export Price Inflation Series

	0.3	0.5	0.8	0.9	0.95	0.99
Spain	1	1	1	2	3	9
Japan	1	1	3	6	17	175
Korea	1	1	2	6	16	154
Singapore	1	1	3	8	22	256
Taiwan	1	1	1	2	4	27
U. S.	1	1	2	4	8	65
Mexico	1	1	2	5	11	96
Belgium	1	1	1	2	2	9
Brazil	1	1	1	1	1	5
Finland	1	1	2	4	10	80
HK, PRC	1	1	3	8	23	263

Note: Values indicate the number of months required for the proportion α of the effect of a unit shock to dissipate.

Table D6: Bootstrap Confidence Intervals of the Impulse Response Functions of ARFIMA(0,*d*,0) Models for Export Price Inflation Series

	Level	1	3	6	12	24	48	96	192	384	504	744	984	1200
Spain	2.5%	-0.217	-0.05	-0.021	-0.009	-0.004	-0.002	-0.001	0	0	0	0	0	0
70:01-03:12	97.5%	-0.057	-0.017	-0.008	-0.004	-0.002	-0.001	0	0	0	0	0	0	0
Japan	2.5%	0.231	0.105	0.063	0.037	0.022	0.013	0.008	0.004	0.003	0.002	0.002	0.001	0.001
60:01-04:02	97.5%	0.374	0.203	0.134	0.088	0.057	0.037	0.024	0.016	0.01	0.009	0.007	0.006	0.005
Korea	2.5%	0.203	0.09	0.052	0.03	0.018	0.01	0.006	0.003	0.002	0.002	0.001	0.001	0.001
71:01-04:02	97.5%	0.371	0.201	0.133	0.087	0.056	0.037	0.024	0.015	0.01	0.008	0.007	0.005	0.005
Singapore	2.5%	0.232	0.106	0.063	0.037	0.022	0.013	0.008	0.004	0.003	0.002	0.002	0.001	0.001
78:01-04:01	97.5%	0.421	0.241	0.165	0.111	0.075	0.05	0.034	0.023	0.015	0.013	0.01	0.009	0.008
Taiwan	2.5%	0.043	0.015	0.008	0.004	0.002	0.001	0.001	0	0	0	0	0	0
76:01-04:02	97.5%	0.223	0.101	0.06	0.035	0.021	0.012	0.007	0.004	0.002	0.002	0.001	0.001	0.001
U.S.	2.5%	0.066	0.024	0.013	0.007	0.003	0.002	0.001	0.001	0	0	0	0	0
88:01-03:03	97.5%	0.324	0.166	0.106	0.067	0.042	0.026	0.016	0.01	0.006	0.005	0.004	0.003	0.003
Mexico	2.5%	0.165	0.069	0.039	0.022	0.012	0.007	0.004	0.002	0.001	0.001	0.001	0.001	0
70:01-03:12	97.5%	0.326	0.167	0.107	0.068	0.043	0.027	0.017	0.011	0.007	0.005	0.004	0.003	0.003
Belgium	2.5%	-0.314	-0.061	-0.023	-0.009	-0.004	-0.002	-0.001	0	0	0	0	0	0
93:01-03:12	97.5%	0.001	0	0	0	0	0	0	0	0	0	0	0	0
Brazil	2.5%	-0.066	-0.02	-0.009	-0.004	-0.002	-0.001	0	0	0	0	0	0	0
78:01-04:02	97.5%	0.118	0.047	0.026	0.014	0.008	0.004	0.002	0.001	0.001	0.001	0	0	0
Finland	2.5%	0.106	0.041	0.022	0.012	0.006	0.003	0.002	0.001	0.001	0	0	0	0
85:01-04:02	97.5%	0.337	0.175	0.113	0.072	0.046	0.029	0.018	0.012	0.007	0.006	0.005	0.004	0.003
HK, PRC	2.5%	0.213	0.095	0.056	0.033	0.019	0.011	0.006	0.004	0.002	0.002	0.001	0.001	0.001
83:01-02:03	97.5%	0.439	0.257	0.178	0.122	0.083	0.056	0.038	0.026	0.018	0.015	0.012	0.01	0.009

Note: The heading row gives the number of months for which the impulse responses apply.