Incomplete pass-through in import markets and permanent versus transitory exchange-rate shocks

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This paper investigates the price-setting behavior of suppliers of imports. The ad-hoc estimation of error-correction models for the U.S., Japan, Germany, France, and Italy indicates that changes in the real exchange rate are not fully passed through into import prices, not even in the long run. By modeling the dynamic behavior of import suppliers as an inter-temporal optimization problem with adjustment costs it can be shown, however, that the degree of long-run pass-through might be underestimated if one does not properly distinguish between permanent and transitory exchange rate changes. Therefore, a Blanchard-Quah decomposition is used to identify the permanent and the transitory component of the exchange rate from its joint process with the current account. Re-estimation with this additional piece of information yields substantially higher degrees of pass-through in the long run as well as in the short run. Thus, a proper distinction might be crucial when predicting import price responses to exchange rate changes.

JEL-codes: C22, C51, E31, E37, F12, F32

1. Introduction

The phenomenon that import prices do not change one-to-one with the exchange rate is a long-established fact in international macroeconomics, see, e.g. Krugman (1987), Marston (1990) and Goldberg and Knetter (1997). Most recently, it has regained attention in lively debates over an appropriate monetary and exchange rate policy, and researchers increasingly recognize its potential role in reducing the international transmission of business cycles.[1] Numerous empirical studies, like for instance those of Campa and Goldberg (2002), Otani, Shigenori and Shirota (2003) and Pollard and Coughlin (2003) strengthen the earlier evidence for a variety of alternative specifications of the relevant import price equation, different sets of explanatory variables and disaggregate as well as aggregate import prices.[2] While the primary concern of this strand of the literature is to obtain a thorough understanding

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1 See OECD (2002), p. 149.
2 For a small selection of earlier studies, see Knetter (1993), Alexius and Vredin (1999), and Gross and Schmitt (1999).

of the market characteristics that lead to a specific pricing behavior under imperfect competition, related research, like in McCarthy (2000), Hüfner and Schröder (2002), and Hahn (2003), focuses on the implications of incomplete pass-through for domestic inflation rates and monetary policy. Typically, these studies estimate VARs and analyze the transmission of exchange rate changes and import price shocks into domestic prices by the resulting impulse response functions. Irrespectively that in this context the degree of pass-through into import prices is more or less a by-product, the results are quite comparable to the more micro orientated approach.

There are three explanations for incomplete pass-through of exchange rate changes suggested in the literature: 1) the prevalence of local currency invoicing, 2) increasing marginal costs of suppliers, and 3) imperfect competition in international goods markets. The last and most compelling argument is highlighted, for instance, by Dornbusch (1987) in his extension of the Dixit-Stiglitz model for a market with differentiated products. In Dornbusch's version firms engage in an oligopolistic price competition and set their price as a mark up over marginal costs. The key element of the model is that profit maximization leads to a flexible mark up which varies with the market share of each supplier.[3] Take, for example, the case that the importing country's currency depreciates, which leads to a loss in the price competitiveness of foreign exporters. Since in turn this would reduce their market share and also total profits, the exporters in the model react by lowering their mark up. In the opposite case of an appreciation, the exporters prefer not to take full advantage of possible gains in their market share but rather increase their mark up. In sum, exchange rate changes are to some extend offset by mark-up adjustments that prevent a complete pass-through from exchange rates into import prices. Since in essence import prices are kept close to the prices of domestic competitors, the underlying pricing behavior is also called pricing to market (PTM).[4]

In the long run, however, the degree of PTM should be low since wages adjust according to the profitability of firms and entry and exit decisions usually lead to a correction of profit margins.[5] These adjustments should also eliminate incomplete pass-through resulting from the other two sources. Invoicing in local currency is likely to be unsustainable in the long run, given the huge persistent swings in nominal exchange rates, and marginal costs of suppliers, especially for aggregate imports,

3 Additional determinants of the mark up that are independent from each supplier's price relation to the average market price are the substitutability of imported and domestic goods and the degree of price interdependence between the suppliers, which can be regarded as an indicator for the market structure and the tightness of the oligopoly.

4 This terminology was introduced by Krugman (1987).

5 See Dornbusch (1987), p. 104.

should not be sensitive to the importing country's demand as world exporters have enough capacity to overcome short run bottlenecks.

The seeming inconsistency between the empirical evidence and the long-run implications of economic theory provides the grounds for the main hypothesis of the paper. It is argued that the empirical results might stem from the failure of conventional estimations to properly distinguish between permanent and transitory exchange rate changes. Therefore, the true long-run pass-through might be in fact higher or even complete. The intuition for this argument is simple. Standard estimation techniques (like error-correction models in the case of non-stationary data) estimate only a mixture of responses of import prices to permanent and transitory shocks. However, if suppliers face adjustment costs with respect to changes in their prices like, for instance, in the sticky price models of Rotemberg (1982), and Froot and Klemperer (1989), the responses to persistent exchange rate shocks should be larger than to transitory ones as soon as the agents have enough information to disentangle the two. Therefore, if the information about the persistence of shocks is not used in the estimations of the price-setting behavior, the degree of pass-through is possibly understated.

A first step to develop this argument is the unconditional estimation of error-correction models (ECMs) for the supply behavior of importers to the U.S, Japan, Germany, France and Italy. Then a Blanchard-Quah decomposition of the joint process of the exchange rates and the current account is used to recover permanent and transitory exchange rate changes from the data. Finally, the estimations of the ECMs are repeated incorporating a proper correction of the pricing rules the import suppliers would make if their responses were in fact conditional on the persistence of exchange rate changes.

The detailed structure of the paper is as follows. Section 2 briefly recapitulates the pricing to market hypothesis as it comes out of the static model of Dornbusch (1987). To characterize the dynamic behavior import suppliers the model is extended to include linear quadratic adjustment-costs and rational expectations. At this stage, the dynamic import supply function is derived without any specific assumptions about transitory and permanent components in the exchange rate. Section 3 presents the estimation of the dynamic import supply functions for the countries of this study. Unit-Root tests indicate that most of the time series involved can be regarded as integrated of order one (I(1)). Hence, to establish long run import supply relationships there has to be cointegration among the relevant time series. For most of the countries a cointegrating relationship is actually supported by the data. Since the estimated cointegration vectors for the various countries imply a substantial degree of incomplete pass-through that is in line with other studies, the results provide the benchmark for the main argument of the paper. In section 4 the dynamic model is reformulated explicitly allowing for a distinction between permanent and transitory exchange rate shocks. This leads to the proposition that if suppliers consider the persistence of exchange rate shocks in their price adjustment, this might result in a downward bias of the conventional estimates for the long run pass-through elasticity. Section 5 presents the results of the bivariate decomposition of the real exchange rate into its permanent and transitory component and investigates to what extend the estimates for the supply functions incorporating this information differ from the previously obtained results. Section 6 concludes.

2. Pricing behavior in imperfectly competitive markets

A theoretical foundation for pricing to market that draws on imperfect competition is suggested in Dornbusch's (1987) model for an oligopoly with domestic suppliers and foreign exporters. In the particular market setting, the products of each supplier are imperfect substitutes for one another. This allows each supplier of imports to set his individual price P_M^f (denoted in the importing country's currency) for the goods he offers (M). In doing so, he has to take into account the substitutability of his product and the price responses of his competitors. His costs C^f can be expressed as a function of the quantity supplied and the alternative price he can get in his home country (P_Q^a).[6] Thus the static profit maximization problem of the exporter is formulated as

$$(1) \qquad \max_{P_M^f} \ \Pi = P_M^f \cdot M(P_M^f) - C^f(M, P_Q^a).$$

The first order condition for this problem is

$$\frac{M}{P_M^f}(1+\varepsilon)[P_M^f - \frac{\varepsilon}{1+\varepsilon} C_M^f] \equiv 0,$$

where ε is the subjective price elasticity of demand for the exporter's products. As Dornbusch shows, ε, which determines the mark up over marginal costs, depends positively on the relation between the domestic competitors' prices (P_Q) and the

[6] Dornbusch (1987) expresses the unit cost of imports as unit labor costs. See p. 95.

foreign exporters' prices ($P_M^f \cdot E$), where the effective exchange rate E serves to express both prices in the domestic currency.[7] Hence, the solution to (1) is the following pricing rule for the exporter:

$$(2) \qquad P_M^f = \left[\frac{P_Q}{P_M^f \cdot E} \right]^{\psi} \cdot C_M^f(M, P_Q^a), \text{ with } \psi \geq 0 .$$

If again E is used to translate foreign export prices into domestic currency import prices (by setting $P_M = P_M^f \cdot E$), then the price equation for the *supply of imports* can be solved for the following log-linear form in terms of the relative import price $p = p_M - p_Q$ and the real exchange rate $r = e + p_Q^a - p_Q$:

$$\boxed{(1) \qquad p = \gamma_0 + (1 - \gamma_1)[r - \gamma_2 m] ,}$$

where $0 \leq \gamma_1 = \psi / (1 + \psi) \leq 1$, and $\gamma_2 \geq 0$.

The notion of static PTM is fully captured by the coefficient γ_1, which can be interpreted as a measure of the degree of PTM behavior.[8] If γ_1 is smaller than one, then neither foreign cost changes nor changes in the effective exchange rate are fully passed through into import prices. The parameter $(1 - \gamma_1)\gamma_2$ marks the inverse of the price elasticity of import supply. In most empirical studies of aggregate import markets the estimate of this elasticity turns out to be infinite, that is, the supply price does not depend on the aggregate quantity of imports of a country.[9] This can be easily rationalized for aggregates given that the supply side is potentially constituted by the whole rest of the world. From this viewpoint, it seems unlikely that the marginal cost of supply is increased by the imports of a single country. Therefore, throughout the following analysis we set the parameter γ_2 to zero. This implies that the relative import price is determined by the real exchange rate only. Furthermore, if the time series are non-stationary, we expect a cointegrating relationship between the two variables with the cointegration vector [1, -(1$-\gamma_1$)]. Finally, note that the estimate

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7 This relative price is proportional to the market share of foreign exporters. Just like in the classical Cournot oligopoly, a greater market share corresponds to a smaller price elasticity of demand for the exporters' products. In turn, this allows for a higher mark up.

8 Conversely, $1 - \gamma_1$ measure the degree of long run pass-through.

9 See Goldstein and Khan (1984).

for the parameter γ_1 should not lie too far away from zero if the arguments of Dornbusch (1987) of annihilation of PTM by labor market adjustment and market entry and exit apply.

The main assumption to formulate a dynamic version of the model is that price changes are costly for suppliers.[10] The notion of costly price changes encompasses several other models for dynamic PTM, e.g. the model of Froot and Klemperer (1989) where costs for switching to other products lead to an investment character of market share or the model of Baldwin (1988) with sunk costs of entering and exiting markets. The crucial point is that in imperfectly competitive markets price changes of suppliers are closely watched by market participants. Therefore, price changes may destroy the reputation for a stable pricing policy and ultimately result in a reduced market share that is costly to rebuild. Moreover, since the reputation of a firm is more likely affected by large price changes than by small ones, price-adjustment costs can be considered to be quadratic.[11]

Price adjustment costs lead to an inter-temporal profit maximization problem in which for all periods $t+i$ ($i = 0,1,...,\infty$) the agents have to consider the costs occurring from relative price changes ($\Delta P_{t+i} / P_{t+i-1}$) as additional term in their profit function.

$$(3) \qquad \Pi_t = E_t \sum_{i=0}^{\infty} \theta^i \left[P_{t+i} \cdot P_{Q_{t+i}} \cdot M_{t+i} - C[M_{t+i}(P_{t+i})] - \alpha (\Delta P_{t+i} / P_{t+i-1})^2 \right].$$

The parameter α captures the relative importance of the adjustment costs and θ is the discount rate. In order to facilitate the further derivation Kasa (1992) proposes to rewrite this function as a second order Taylor approximation around the long run equilibrium value $\widetilde{P}_{t+i} = P_{t-1}$. Making also use of the standard approximation $P_{t+i} / \widetilde{P}_{t+i} - 1 = p_{t+i} - \widetilde{p}_{t+i}$ and $\Delta P_{t+i} / P_t \approx p_{t+i} - p_{t+i-1}$ this leads to the following approximate profit function

$$\Pi_t \approx E_t \sum_{i=0}^{\infty} \theta^i [\lambda (p_{t+i} - \widetilde{p}_{t+i})^2 + \Delta p_{t+i}{}^2], \text{ with } \lambda = \widetilde{M}_{t+i} \cdot C_M \cdot \varepsilon / \alpha .$$

[10] The linear quadratic adjustment cost model has been previously applied in a number of studies in international trade. See e.g. Gagnon (1989), Kasa (1992), Amano and Wirjanto (1994).

[11] See Rotemberg (1982), p. 1190.

Thus, in principal the suppliers face a trade-off with respect to costs for not setting their price to the optimal long-run value $\widetilde{P}_{t+i}$ instantaneously and costs resulting from price adjustments. This setup is equivalent to the classical linear quadratic adjustment cost model of Sargent (1978). Hence, the exporter's optimal pricing policy can be easily obtained as the following standard saddle-path solution:

$$(2) \qquad p_t = \mu p_{t-1} + (1-\mu)(1-\theta\mu)\sum_{i=0}^{\infty}(\theta\mu)^i E_t\ \widetilde{p}_{t+i}\ ,^{12}$$

$$\text{with} \qquad E_t\ \widetilde{p}_{t+i} = (1-\gamma_1)r_{t+i}\,.$$

The parameter μ (μ<1) captures the stable root of the saddle-path solution. It depends positively on the relative importance of the adjustment cost of suppliers ($1/\lambda$) and negatively on the size of the discount factor (θ).

As evident from equation (4), the target path of the relative import price $\widetilde{p}_{t+i}$ is solely determined by the evolution of the real exchange rate. It is assumed that this so-called forcing variable can be predicted by the following q^{th} order VECM in the n variables x_t: $\Delta x_t = \Pi x_{t-1} + \sum_{k=1}^{q} \mathrm{A}_k \Delta x_{t-k}$, where r_t itself is the first element of x_t and the import supply function is the only cointegrating relationship, that is $rank(\Pi) = 1$. Furthermore, it is assumed that the real exchange rate r_t is weakly exogenous with respect to the parameters in the long run import supply function. This implies that the error-correction term ($ect_t = p_{t-1} - (1-\gamma_1)r_{t-1}$) does not enter the equation for Δr_t.[13] We allow, however, for feedback effects from the relative import price to the real exchange rate. Transformed into the first-order companion form, the forecasting model can be written as

$$(5) \qquad \Delta z_t = \Pi^* z_{t-1} + \mathrm{A}^* \Delta z_{t-1}\,,$$

[12] This form of price-adjustment assumes that foreign suppliers adjust their prices denominated in the domestic currency. Meurers (2003) derives a supply function that mixes price adjustment in domestic and foreign currency prices. This much more complicated function can be useful in specific simulation problems. For the purpose of this paper, however, it suffices to demonstrate the dynamic structure of the model using the simpler form.

[13] See Johansen and Juselius (1990) for a discussion of weak exogeneity in the context of a VECM.

where the $n \cdot q \times 1$ vector $z_{t-1} = [x'_{t-1}, x'_{t-2}, ..., x'_{t-q}]'$ denotes the information set of the agents, and according to the above definition of x_t, ι is the $n \cdot q \times 1$ selector vector containing the parameter $(1-\gamma_1)$ in the first line and zeros elsewhere. The $n \cdot q \times n \cdot q$ matrices Π^* and A^* are defined as follows:

$$\Pi^* = \begin{bmatrix} \underset{n\times n}{\Pi} & \underset{n\cdot(q-1)\times n\cdot(q-1)}{0} \\ \underset{n\cdot(q-1)\times n\cdot(q-1)}{0} & \underset{n\cdot(q-1)\times n\cdot(q-1)}{0} \end{bmatrix}, \qquad \mathrm{A}^* = \begin{bmatrix} A_1 & A_2 & \cdots & A_q \\ & \underset{n(q-1)\times n(q-1)}{\mathrm{I}} & & \underset{n\times n}{0} \end{bmatrix}.$$

Using the law of iterative expectations, the forecasting model can be written as

$$(6) \qquad E_t\, \tilde{p}_{t+i} = \iota'[G_i(\Pi^*, \mathrm{A}^*)\Delta z_t + \Pi^* \cdot H_i(\Pi^*, \mathrm{A}^*) + z_{t-1}],$$ [14]

where $G_i(\cdot)$ and $H_i(\cdot)$ are complex non-linear functions depending on the parameters of the VECM. Finally, note that weak exogeneity of w_t^r implies that $\iota'\Pi^* = 0$.

Rational agents should employ this model for their forecast of the target path of their decision variable p_t. Hence, plugging the expected target values of (6) into the adjustment path (5) and using the definition

$$G(\mu, \theta, \mathrm{A}^*, \mathrm{H}^*) \equiv (1-\mu)(1-\theta\mu)\sum_{i=0}^{\infty}(\theta\mu)^i G_i(\mathrm{A}^*, \mathrm{H}^*)$$

the closed form solution to (5) can be shown to have the following error-correction format:[15]

$$\boxed{(7) \qquad \Delta p_t = -(1-\mu)[p_{t-1} - (1-\gamma_1)r_{t-1}] + \iota' G(\theta, \mu, \Pi, A)\Delta z_t\,.}$$

The error correction format highlights that contemporaneous and lagged values of any variable that helps to predict the target path $\tilde{p}_{t+i}$ might enter the right hand side of the dynamic behavioral equation for the relative import price. This also includes

[14] The derivation of this prediction formula is outlined in the appendix A.1.

[15] A detailed derivation of the closed form solution to similar, and in parts more general adjustment cost problems can be found in Nickell (1985), Engsted and Haldrup (1997), and Kozicki and Tinsley (1999).

the decision variable if there are feedback effects from p_t on the predictors (e.g. in case of Granger causality for the real exchange rate).

Beyond the implied long run PTM of Dornbusch's (1987) model reflected in the error-correction term of (7), there can also be short run PTM effects via contemporaneous and lagged differences of the exchange rate, since these differences are also elements of the agents' information set which potentially improve their forecasts. However; the direction of the short-run impact of exchange rate changes on the relative import price is not clear. This is due to the fact that we do not have a priori information about the parameters in the prediction equation (6) and we therefore do not dispose of any restrictions for the parameters in $\iota' G(\theta,\mu,\Pi,A)\Delta z_t$. This is an important result of the explicit formal derivation of the dynamic supply function. It implies that when estimating (7) we cannot a priori exclude any contemporaneous and lagged differences of the variables z_t from the set of right-hand explanatory variables. Instead, we have to decide upon the inclusion of explanatory variables in equation (7) solely on the basis of their contribution to the overall goodness of fit.

The advantage of the error correction form is that it enables us to estimate the long run and the short run supply price determination in a single step. Standard errors for the parameter estimates can be obtained by estimating equation (5) using the so-called Bewley transformation.[16]

3. Conventional estimation of import supply functions

The goal of this section is to establish some benchmark results for the pricing behavior of import suppliers. By estimating the conventional error-correction form (7) for the U.S., Japan, Germany, France and Italy, we would like to verify whether the results for our sample are in line with the majority of studies that report incomplete pass-through of exchange rates in the long run.

The time series used in the estimations are taken from the *OECD's Main Economic Indicators database*. For each country we constructed the relative import price ($p_{M\,t} - p_{Q_t}$) as the difference of the logarithm of the aggregate import price index and the consumer price index. The log of each country's real exchange rate (r_t) is obtained as the inverse of the *OECD*'s real effective exchange rate on the basis of consumer prices. Campa and Goldberg (2002) use the real effective exchange rate calculated from unit labour costs whereas Pollard and Coughlin (2003) base their

[16] See Bewley and Fiebig (1990), p. 346.

cost variable on producer prices. Since both studies record a substantial degree of PTM, we regard it as sufficient to base our argument on one cost variable only. In addition, several empirical studies seem to indicate that the differences in the development of the individual competitiveness indicators are rather negligible.[17]

Preliminary estimations of the static supply relationship, which are not reported here, lent only weak support for the existence of a stable long-run relationship between the two variables in the model. A possible explanation are distorting effects of raw materials prices changes, especially of those that are associated with the two oil price shocks.[18] To a large extend, such shocks carry over into import prices. Real exchange rates, however, are less responsive to shocks in raw materials prices, since raw materials make up only a small fraction of the consumption basket, and by far the largest part of fluctuations in the real exchange rate is due to nominal exchange rate changes. Empirically, these considerations can be underpinned by the pair-wise correlations of the changes in $p_{M\,t} - p_{Q_t}$ and r_t with changes in the logarithmic oil price ($p_{Oil\,t}$) reported in the following table:

Table 1: Pair-wise correlations of oil prices with exchange rates and import prices

	U.S.	*Japan*	*Germany*	*France*	*Italy*
$Corr(\Delta r_t\ , \Delta p_{Oil\,t})$	0.00	0.10	0.04	0.17	0.08
$Corr(\Delta p_t\ , \Delta p_{Oil\,t})$	0.54	0.43	0.44	0.38	0.43

In light of the different impact of raw materials prices, the log of the oil price is included in the subsequent estimations; in the long-run as well as in the dynamic part of the model. The oil price series is taken from the *IMF's International financial Statistics* and is seasonally adjusted by the ratio to moving average method. The complete sample of common observations finally ranges from 1975:1 to 2002:2 ($T = 110$).

Before estimating the supply relations, it is worthwhile to test the order of integration of the series. If the series have stochastic trends, special care must be taken in interpreting the estimation results. Only cointegrating relationships, i.e. stationary linear combination of series, can be regarded as consistent with the implications of economic theory for the long run. Therefore, two different unit-root test are carried out prior to the estimations. The first is the conventional augmented Dickey-Fuller

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[17] See, e.g. European Central Bank (2003), pp. 75 or Deutsche Bundesbank (1998), pp. 41.

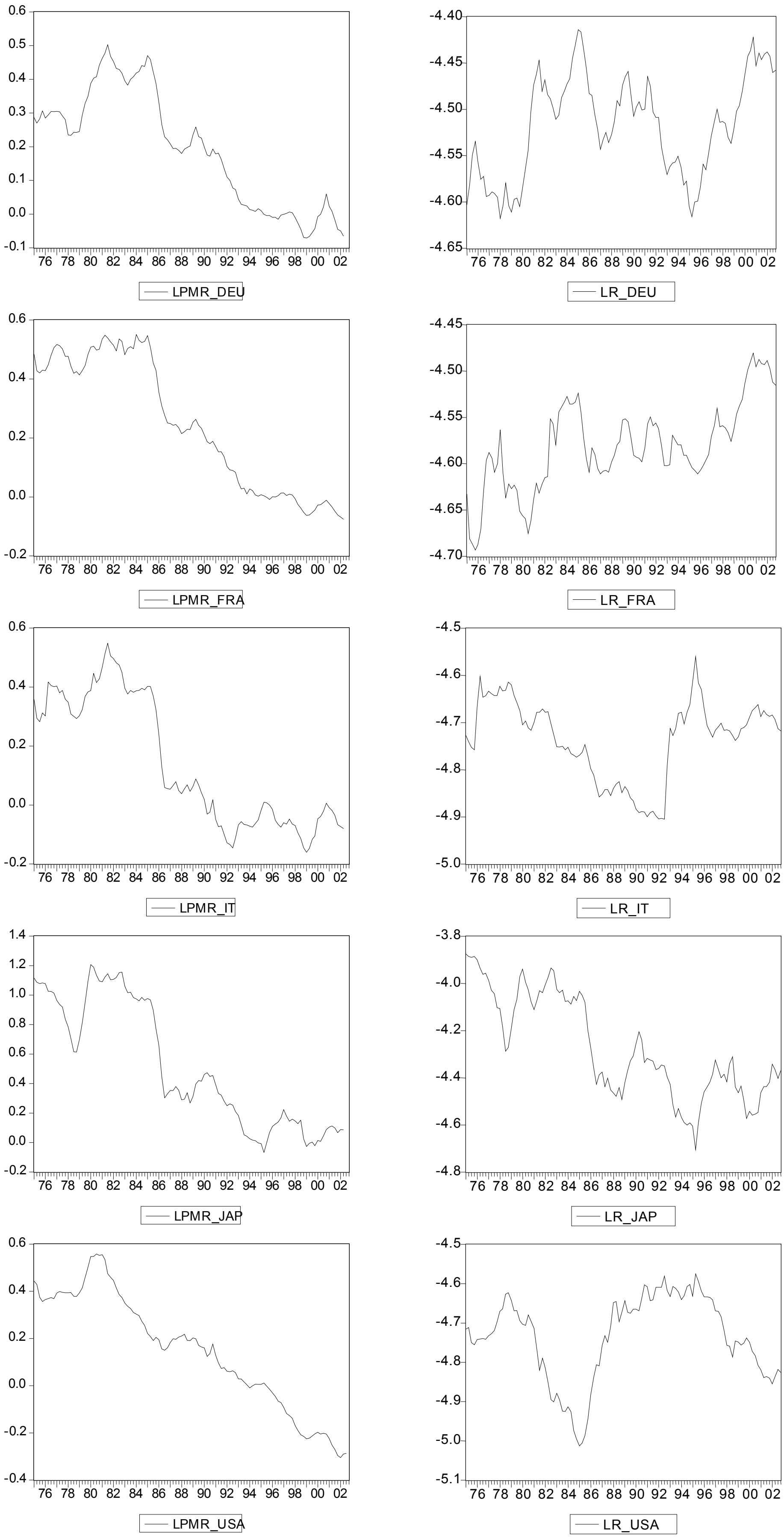
[18] Clostermann (1996) reports the same problem in his estimation of import supply functions for Germany.

(ADF) test, which test the null hypothesis of non stationarity. The second test we use was suggested by Kwiatkowski et al. (1992) (KPSS test). It proceeds the opposite way, and tests the null of stationarity against the alternative of a unit root. If the series turn out to be in fact integrated of order one (I(1)) or even of a higher order then, apart from estimating (5), we need to test the cointegration property of the hypothesized long-run relationship between $p_{M\,t} - p_{Q_t}$, r_t , and $p_{Oil\,t}$. Such a test can be based on the *t*-statistic for the loading coefficient $(1-\mu)$, where we use the critical values of Banerjee, Dolado, and Mestre (1998).

3.1 Unit root tests

The graphs for the relative import prices on the left hand side of figure 1 indicate that for all countries the time series follow a downward trend. The same is more or less true for the real exchange rates, which are displayed on the right hand side. Especially the series for France and Japan display a trending behavior, although less pronounced than the relative import prices. These observed patterns have two implications for the subsequent analysis. First, in the unit-root tests the alternative model should be specified as a trend stationary process. Second, in the later estimation of equation (5), a trend component should be included in the cointegration relationship to account for the possibility of diverging deterministic trends in the individual series. Finally, note that the two large oil price changes in our sample of observations, namely the second oil price shock in the end of 1979 and the strong fall in oil prices in 1986 are significantly reflected in the time pattern of both series. Nevertheless, the previously discussed difference in the impact on both series still justifies the inclusion of the oil price in the cointegration relationship.

Figure 1: Relative import prices and real exchange rates



The ADF test are based on the following equation, which has to be estimated for each individual series y_t by OLS:

$$(8) \qquad \Delta y_t = c_0 + (\alpha - 1) y_{t-1} + c_1 t + \sum_{i=1}^{k-1} \beta_i \Delta y_{t-i} + \varepsilon_t , \qquad \varepsilon_t \sim IID(0, \sigma^2) .$$

This test equation nests both a stochastic trend model as well as stationary alternatives. These consist either of a deterministic trend model, or in the case that c_1 equals zero, of a process that is stationary around a constant mean. Whether y_t is characterized by a unit-root process can be evaluated by a test of the null hypothesis (H_0) $\alpha = 1$ against the alternative (H_1) $\alpha < 1$. If α is not significantly smaller than one, a stochastic trend in y_t cannot be excluded and therefore, one has to take at least the first difference of the series in order to obtain a stationary process. To verify whether the time series has two unit roots and is even I(2), the ADF tests have to be repeated for the first difference of the series. In this second test, only a constant has to be included in the test equation for a correct nesting of the stationary alternative. The significance of α can be tested using the conventional *t*-statistic. Because of a non-stationary series under the null, however, inference has to be based on the tabulated asymptotic critical values of Fuller (1976). In addition, since the critical values are only applicable if the residuals of equation (6) are uncorrelated, a sufficient lag order has to be determined prior to the unit-root tests. The following general-to-specific approach was adopted for the selection of an appropriate lag order. We begin with a generous number of lags ($k = 8$) and subsequently reduce insignificant lagged differences from the test equation. As a side condition for the elimination we require that the Ljung-Box Q-statistics for serial correlation from the 1st to the 12th lag remains insignificant at the 5% level. The final choice is reported in the second column of tables 1 and 2.

The typical drawback of the ADF test as well as of other tests of the null hypothesis of a unit root, e.g. the Phillips-Perron test, is the lack of power against stationary alternatives. To be on the safe side as far as the order of integration of the series is concerned, we additionally carry out the KPSS test. It is based on the following model:

$$(9) \qquad y_t = \alpha + \beta t + d \sum_{i=1}^{t} u_i + \varepsilon_t , \qquad t = 1, \ldots, T .$$

The error terms u_i and ε_t are both covariance stationary and the parameter d is restricted to $d \in \{0,1\}$.[19] The null hypothesis of trend stationarity can then be formulated as H_0: $d = 0$. It is tested using the residuals e_t from the regression of y_t on an intercept and a time trend, which under the null should be stationary. The test statistic is given by $LM = \sum_{t=1}^{T} S_t^2 / \hat{\sigma}_\varepsilon^2$, with the partial sum process $S_t = \sum_{i=1}^{t} e_i$. The essential part of the test statistic is the estimate of the error variance from the regression of the y_t on intercept and trend. KPSS use a Bartlett Kernel to weight the estimated autocovariances in the calculation of $\hat{\sigma}_t^2$. Based on their size and power simulations for the tests statistic, they recommend a lag-truncation parameter $l = 8$ in this calculation, which is also adopted in the following tests.

Table 1: Unit root tests for the real exchange rates

variable	Lags (k)	det. comp. in test	ADF	ADF 1st diff.	KPSS (l = 8)	KPSS 1st diff. (l = 8)
R_GER	4	C,T	-3.30*	-5.03***	0.128*	0.083
R_FRA	5	C,T	-3.31*	-5.32***	0.074	0.074
R_ITA	4	C,T	-2.46	-5.55***	0.157**	0.123
R_USA	5	C,T	-2.61	-3.55***	0.105	0.050
R_JAP	6	C,T	-2.70	-4.98***	0.133*	0.199

'R' plus country extensions stands for the real exchange rate. Sample: 1971:1 – 2002:4.

Table 2: Unit root tests for the relative import prices and the oil price

variable	Lags (k)	det. comp. in test	ADF	ADF 1st diff.	KPSS (l = 8)	KPSS 1st diff. (l = 8)
PMR_GER	6	C,T	-2.25	-3.74***	0.180**	0.147
PMR_FRA	5	C,T	-2.58	-4.15***	0.157**	0.118
PMR_ITA	7	C,T	-1.81	-3.96***	0.147**	0.096
PMR_USA	7	C,T	-3.25*	-3.85***	0.136*	0.142
PMR_JAP	7	C,T	-2.53	-4.14***	0.111	0.072
POIL	1	C	-1.37	-5.23***	0.117	0.109

'PMR' plus country extensions stands for the relative import price and POIL for the oil price. Sample: 1975:1 – 2002:2, critical values are for the ADF tests are taken from Harris (1995), Table A.1, p. 156. The critical values for the KPSS tests are from Kwiatkowski et al. (1992), p. 166. ***, **, * denote significance at the 1%-, 5%- , or 10%-level.

[19] For this expostition of the KPSS test, see Hobijn, Franses, and Ooms (1998).

The ADF test statistics in table 1 and 2 do not reject the hypothesis of a unit root for most of the series at relatively large significance levels. Only for the French and German real exchange rate and the U.S. import price non-stationarity is rejected at a 10% level. On the other hand for some of the series the KPSS tests do not reject the null of trend, or mean stationarity; in some cases like for the real exchange rate of the USA and France, the relative import prices of Japan, and the oil price not even at significant levels above the conventional threshold of 10%.
The implications of these ambiguous results are open to debate. To make a final decision, we prefer a conservative synopsis of both tests. That is to consider a series as having a unit root, as long as not both tests suggest to accept an I(0) model. This implies that we would favor an I(0) model only in the case of the French real exchange rate, where both tests point to a stationary series. As far as the tests for a unit root in the first differences are concerned, both tests identically lead to convincing results against an I(2) model. In short, adopting the proposed synopsis of both tests leads us to favor an I(1) model for all the series except for the French real exchange rate. Nevertheless, since we have three series in the hypothesized long-run relationship, even in the French case we might find cointegration. This point is clarified in the following estimations where we explicitly test the hypothesis of cointegration.

3.2 Estimation of error-correction models

Now that we are clear about the order of integration of the series, we can move on to the estimation of the error-correction models (5) for the five countries of our sample. As a preliminary step, we estimate the import supply relationships by simple static OLS. The residuals of this exercise can help to identify structural breaks in the long-run relationships that, if unaccounted, lead to only weak evidence for cointegration in the ECM estimation. The obvious case for a structural shift is the German reunification that is realized in the data in the beginning of 1991. Consequently, an appropriate step dummy should be included in the estimations. Additionally, the OLS residuals reveal a structural break for all three European counties in the beginning of 1993 (see figure 2).

Figure 2: Structural shifts in the European supply functions

Germany Italy France

A likely explanation for the common structural break is the completion of the *Single European Market* in the first quarter of 1993. First, this led to an amendment in the statistical registration of trade flows within the European union, which might have triggered a change in the composition of goods in the import prices.[20] Second, the European market might have contributed to a more intense competition, could have induced fundamental price adjustments. In the estimations of the ECMs it turns out that the structural shift is best captured by a combination of a step dummy in 1993 and a broken trend.

In the following estimations we make the simplifying assumption that the set of potential predictors for the real exchange rate is restricted to the relative import price and the oil price. Thus, only significant contemporaneous and lagged differences of these variables are included as right-hand side explanatory variables apart from the error-correction term.

As already assumed in the derivation of Equation (5), the real exchange rate and the oil price are considered to be weakly exogenous with respect to the parameters of the cointegration vector. Thus OLS is an appropriate method to obtain consistent parameter estimates. The standard errors for the coefficients in the error correction term (ect_{t-1}) are obtained by a 2SLS estimation of the Bewley transformed equation using p_{t-1} as instrument for Δp_t.

The tests for cointegration of the three series in the hypothesized supply function are based on the significance the loading coefficient $(1-\mu)$. This is done by comparing the ordinary *t*-ratio with the critical values of Banerjee, Dolado, and Mestre (1998). The distribution of the test statistic (t_E) depends on the explanatory variables in the cointegration vector (two in our case) and the number of observations ($T \approx 100$).

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[20] Details of the statistical amendment are discussed in Deutsche Bundesbank (1993), p. 65.

Final issues in the estimation of equation (5) are the selection of an appropriate lag length and additional tests whether the residuals can be broadly regarded as normally distributed white-noise errors. For the lag length selection, we choose the same general-to-specific approach as in the case of the unit-root tests, this time starting from a lag length of $k = 5$. The normality of the residuals was tested using the Jarque-Bera test for excess skewness and kurtosis. If the residuals turned out to be non-normal additional, impulse dummies were included where appropriate. These are given at the bottom of the boxes with the results for the countries in table 3.

Turning to the results, for all countries except for France, the cointegration tests do reject the hypothesis of no cointegration at the 10% level. Hence, we conclude that the possible trend stationarity of the French real exchange rate makes it impossible for the supply relationship to cointegrate. We therefore consider the results for France as only indicative in the subsequent analysis. For the other countries the results support the existence of long run import supply functions in the form of equation (2). In most cases, the coefficient for the real exchange rate is below one, ranging from 0.5 for the U.S. to 0.9 for France. Except for the U.S. though, they are not significantly different from one. Nevertheless, if we limit our attention to the point estimates, the results clearly suggest pricing to market behavior even in the long run. As far as short run is concerned, the coefficients for Δr_t imply an even lower degree of pass through, ranging from 0.3 for the U.S. to 0.75 for Japan. Overall, the results for our specific sample and estimation method can be regarded as broadly consistent with those of Campa and Goldberg (2002) with an notable congruence for the U.S.[21]

As pointed out before, the implication of the results that PTM persists in the long run is puzzling. In this respect, our results motivate further efforts to clarify the underlying reasons. In this paper the hypothesis is that different responses of the agents to permanent and transitory exchange rate changes is responsible for this result. In the next sections this argument is elaborated first on theoretical grounds and is then investigated on the basis of the Blanchard-Quah decomposition of the exchange rate into a permanent and transitory component.

[21] For the U.S. they obtain a pass-through of 0.26 in the short run and 0.41 in the long-run, which significantly differ from one and zero. They also report significant incomplete pass-through in the long run for Australia, Canada, the U.K., and New Zealand.

Table 3: Conventional ECM for the import supply prices

U.S.:

$$\Delta p_t = \underset{(0.03)}{-0.128}\, ect_{t-1} + \underset{(0.07)}{0.394}\, \Delta p_{t-1} + \underset{(0.07)}{0.280}\, \Delta r_t + \underset{(0.01)}{0.065}\, \Delta p_{Oil\,t}$$

$$ect_{t-1} = p_{t-1} - \underset{(0.09)}{0.479}\, r_{t-1} - \underset{(0.03)}{0.214}\, p_{Oil\,t}$$

t_E = *-4.40, adj.* R^2: 0.74, D.W.: 2.17, $Q(12)$ = 18.07, J.B.: 1.77, n = 105

Additional impulse dummies: D91_1.

Japan:

$$\Delta p_t = \underset{(0.04)}{-0.174}\, ect_{t-1} + \underset{(0.04)}{0.148}\, \Delta p_{t-1} + \underset{(0.04)}{0.753}\, \Delta r_t + \underset{(0.02)}{0.087}\, \Delta p_{Oil\,t} + \underset{(0.02)}{0.097}\, \Delta p_{Oil\,t-1}$$

$$ect_{t-1} = p_{t-1} - \underset{(0.12)}{0.864}\, r_{t-1} - \underset{(0.06)}{0.333}\, p_{Oil\,t}$$

t_E = *-3.95, adj.* R^2: 0.89, D.W.: 1.92, $Q(12)$ = 12.59, J.B.: 0.10, n = 90

Additional impulse dummies: D91_1.

Germany:

$$\Delta p_t = \underset{(0.05)}{-0.191}\, ect_{t-1} + \underset{(0.07)}{0.214}\, \Delta p_{t-1} + \underset{(0.07)}{0.650}\, \Delta r_t + \underset{(0.01)}{0.052}\, \Delta p_{Oil\,t} + \underset{(0.01)}{0.011}\, \Delta p_{Oil\,t-1}$$

$$ect_{t-1} = p_{t-1} - \underset{(0.20)}{0.949}\, r_{t-1} - \underset{(0.02)}{0.151}\, p_{Oil\,t}$$

t_E = *-3.88, adj.* R^2: 0.70, D.W.: 2.10, $Q(12)$ = 15.47, J.B.: 0.04, n = 107

France:

$$\Delta p_t = \underset{(0.05)}{-0.148}\, ect_{t-1} + \underset{(0.07)}{0.239}\, \Delta p_{t-1} + \underset{(0.09)}{0.692}\, \Delta r_t + \underset{(0.01)}{0.060}\, \Delta p_{Oil\,t}$$

$$ect_{t-1} = p_{t-1} - \underset{(0.28)}{0.922}\, r_{t-1} - \underset{(0.04)}{0.168}\, p_{Oil\,t}$$

t_E = *-3.29, adj.* R^2: 0.64, D.W.: 1.97, $Q(12)$ = 9.58, J.B.: 0.44, n = 88

Additional impulse dummies: D84_1.

Italy:

$$\Delta p_t = \underset{(0.04)}{-0.204}\, ect_{t-1} + \underset{(0.08)}{0.357}\, \Delta p_{t-1} + \underset{(0.07)}{0.154}\, \Delta p_{t-2} + \underset{(0.08)}{0.498}\, \Delta r_t + \underset{(0.01)}{0.081}\, \Delta p_{Oil\,t}$$

$$+ \underset{(0.01)}{0.019}\, \Delta p_{Oil\,t-1} - \underset{(0.01)}{0.044}\, \Delta p_{Oil\,t-2}$$

$$ect_{t-1} = p_{t-1} - \underset{(0.28)}{0.833}\, r_{t-1} - \underset{(0.04)}{0.156}\, p_{Oil\,t}$$

t_E = *-5.37, adj.* R^2: 0.79, D.W.: 2.15, $Q(12)$ = 18.30, J.B.: 0.52, n = 88

Additional impulse dummies: D80_3, D81_4.

Coefficients for the deterministic components (intercept, trend structural breaks) are not shown here for expository purposes. Standard errors are in brackets. For T = 100 observations, the 10%, 5%, and 1% critical values of the cointegration test (t_E) are –3.66, -3.98, and -4.60 respectively. See Banerjee et al. (1998), p. 277.

4. A dynamic model accounting for permanent and transitory exchange rate changes

This section discusses the pricing behavior of agents in the case that they know how persistent are the observed exchange rate changes. To illustrate the alterations of the model presented in sections 2 and 3, we make the assumption that the exchange rate follows a simple unit-root process.[22] Furthermore, we assume that it can be decomposed into a permanent component q_t and a transitory component s_t:

(10a) $r_t = q_t + s_t$,

(10b) $q_t = q_{t-1} + u_t$, $\quad\quad (s_t, u_t) \sim \text{iid}[0,\text{diag}(\sigma_s^2, \sigma_u^2)\text{I}]$.

If the agents take into account the process of the observed r_t in their determination of the target path for their import prices $\tilde{p}_{t+i}$ in equation (4), they forecast values of r_t in future periods $t + i$ through

$$r_{t+i} = r_t - s_t .$$

This greatly simplifies the summation over the discounted expected target values $\sum_{i=0}^{\infty} (\theta\mu)^i E_t \tilde{p}_{t+i}$ in equation (2), such that in the end one obtains the following closed form solution:

(11) $\Delta p_t = -(1-\mu)[p_{t-1} - (1-\gamma_1) r_{t-1}] + (1-\mu)(1-\gamma_1)\Delta r_t - (1-\mu)\theta\mu(1-\gamma_1) s_t .$

Let us now compare this solution to the case when agents ignore the above decomposition of exchange rates and make predictions on the basis of the unconditional model $r_t = r_{t-1} + v_t$ with v_t as an unknown white noise disturbance. In this case one would obtain the import price dynamics as a special case of the closed-form solution given by equation (7). Hence, the last term involving the transitory shock s_t would simply be missing in their dynamic pricing rule.

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[22] See, e.g., Meese and Rogoff (1983) on the poor forecasting performance of structural exchange rate models in comparison to a simple random walk.

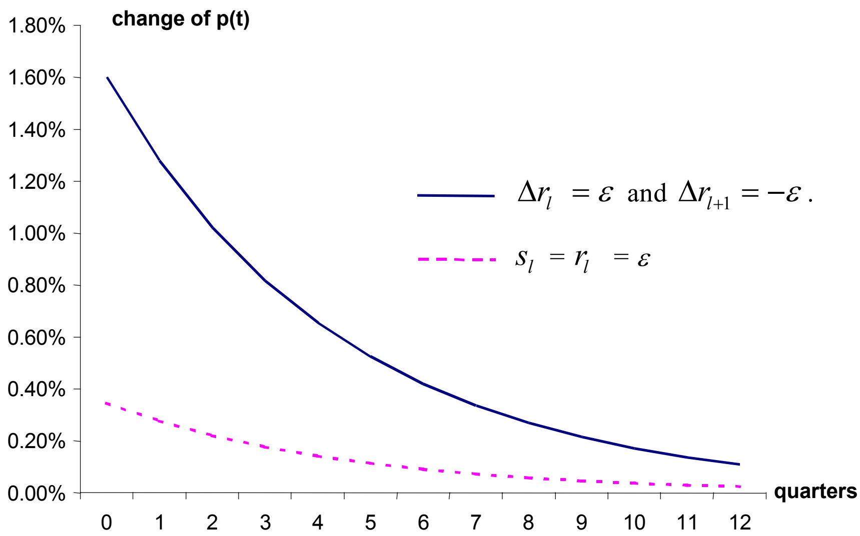
A first important observation is that the estimation of (11) excluding the transitory exchange rate shocks s_t causes a bias to the parameter estimates since the term $-(1-\mu)\theta\mu\delta s_t$ it is pushed into the residual. It is well known from textbook econometrics that the direction of the bias in the estimates depends on the covariance of the explanatory variables and their covariance with the error terms.[23] A simulation in the appendix A.2 shows that in terms of the simple exchange rate model outlined above, there is a tendency to underestimate all three coefficients in equation (11), the adjustment coefficient $(1-\mu)$ in absolute value, the short run pass-through elasticity $(1-\mu)(1-\gamma_1)$ and the long run pass-through elasticity $(1-\gamma_1)$. Especially in case of a relatively high variance of the transitory shocks s_t compared to the permanent shocks u_t and also in case that s_t are to some extend serially correlated, a significant downward bias of the estimate of long-run pass through results.

Our simulations to evaluate the bias in the estimates can be regarded as no more than a preliminary experiment, and a more detailed study allowing for more parameter variations and more general exchange rate models is certainly desirable. Notwithstanding this somewhat limited and explorative character, we can conclude that the frequently recorded incomplete pass-through in the long run might be, at least with a certain probability, simply the consequence of ignoring transitory and permanent exchange rate changes the agents take into account in their price adjustment. Moreover, the cointegration tests for the significance of the error correction term might mistakenly lead to the rejection of non-cointegration.

Even if the bias in the coefficients should turn out to be negligible, the neglect of the decomposition of exchange rates has severe consequences for the precision of forecasts. Consider the frequent exercise to predict the reaction of the relative import price to a transitory exchange rate shock. The conventional forecast for model (11) would treat Δr_t as a white noise process. A temporary shock at time $t = l$ would then enter the model only as $\Delta r_l = \varepsilon$ and $\Delta r_{l+1} = -\varepsilon$. On the other hand, if we properly separated the shocks according to (10a) and (10b), we would set both s_l and r_l equal to ε. For a ten-percent depreciation and illustrative parameter values $\mu = 0.8$, $\theta = 0.98$, and $\gamma = 0.2$ we would get the time paths of the relative import price in figure 3.

[23] See, e.g., Davidson and MacKinnon (1993), pp. 209.

Figure 3: Forecasting the impact of a transitory exchange rate shock



Evidently, the impact of the transitory shock would be drastically exaggerated if we worked with the wrong model. The reason is simply that the shock is mistaken as permanent and the downward adjustment of rational forward-looking agents by $-(1-\mu)\theta\mu\delta s_t$ is ignored.

In sum, this section suggests an extension of the ECM framework for the import supply functions of equation (7) to improve the precision of the estimates for the parameters and the validity of inference from the cointegration tests. The isolated transitory component of the exchange rate should be included as an additional explanatory variable in the estimated ECMs. Even further lags of the transitory component might be warranted since the true exchange rate model could be more complicated than the one sketched above.

5. Empirical Analysis of the extended model

To incorporate the implications of the previous section into our empirical analysis, we require a method to decompose the exchange rate into its permanent and transitory component. Several alternatives have been proposed for this exercise. First of all, we could simply use filtering methods like in Hodrick and Prescott (1997), or univariate decompositions like in Beveridge and Nelson (1981). The problem with these is that they consider only limited information contained in the series itself and

the question is whether these decompositions provide more information about the dynamics of the exchange rate than a generous inclusion of lags in model (7). Apart from this, the former methods usually make strong a priori assumption about the covariance structure of the permanent and transitory component.
A less restrictive way is the bivariate Blanchard-Quah (BQ) decomposition.[24] It offers a way to incorporate implications from economic theory for the joint process of the variable of interest with another economic variable. The general idea is that the joint process of two variables (x_t, z_t), the first non-stationary and the other stationary, is governed by two kinds of shocks. (In the original treatment of Blanchard and Quah the variables are output and unemployment). The main assumption is that only one type of shock has a long run (permanent) impact, and this only on the nonstationary variable x_t. The effect of the other shock dies off in this variable. In other words, the shock only leads to transitory changes of x_t. Moreover, since the other variable is stationary by assumption, it is unaffected by any kind of shock in the long run. This constellation, in particular the non-stationarity of the variable with long run effects of shocks, implies an identifying restriction for the vector moving average (VMA) representation of the process of (Δx_t, z_t). The coefficient matrices in the VMA have to be such that the long-run multiplier of the transitory shock on Δx_t is zero.[25] When imposing this restriction on the VMA representation of an estimated VAR, the permanent and transitory components for both variables can be recovered from the data.

A particularly interesting application of the BQ decomposition to the exchange rate is proposed in Lee and Chinn (2002). The authors show that the BQ framework can be applied to the joint process of the real exchange rate and the current account (in relation to nominal GDP).[26] Two shocks can be expected to govern the process of the two variables r_t and CA_t: 1.) real shocks that stem from shifts in the aggregate demand and supply functions of the country or in one of its trading partners, and 2.) nominal shocks that are the result of disturbances in financial markets. As pointed out by the authors, a broad range of open-economy macro models, e.g., Obstfeld and Rogoff (1996), Betts and Devreux (2000) and Chari et al. (1998), imply that nominal shocks have no or only a negligible effect on the real exchange rate in the long run. Thus the implications of economic theory for the joint process of r_t and CA_t provide

[24] The decomposition is proposed in the seminal paper of Blanchard and Quah (1989).
[25] See Blanchard and Quah (1989), p. 657.
[26] Lee and Enders (1993) use a bivariate VAR of the real and the nominal exchange rate for the decomposition.

for the required identifying assumption in a BQ decomposition. The interpretation of permanent and transitory exchange rate movements in this context is hence that the former are those movements caused mainly by real shocks, whereas the latter are mainly due to nominal or monetary policy shocks.

5.1 The Blanchard-Quah decomposition

The BQ decomposition is only applicable in the way proposed by Lee and Chinn (2002) if the real exchange rates are a non-stationary I(1) process and the current account variables are stationary.[27] The formerly conducted unit-root tests for the real exchange rate clearly support the fulfillment of the first condition with the exception of France. Additional ADF and KPPS unit root test for the current account variable lend also support for the hypothesis that this variable is stationary for the five countries.[28]

The next step is to estimate the bivariate VAR for Δr_t and CA_t. For the selection of an appropriate lag length, we primarily rely on the Akaike information criterion (detailed results are reported in the appendix A.5). If the results for the decompositions were robust against alternative specification, the lag order was reduced in the direction of the choice according to the Schwarz Bayesian criterion.

The estimated VAR is then transformed into the vector moving average (VMA) representation with coefficient matrices C(L). A Choleski factorization of the long run multiplier matrix C(1) of the errors in the VMA representation leads to the identification of the matrix C(0) that measures the contemporaneous impact of the structural shocks (real and nominal) on the variables in the model. Since these impacts coincide with the errors in the VAR, the identified C(0) helps to recover the structural errors from the VAR errors.[29] Finally, with the known structural errors in all periods of the sample it is possible to calculate two paths of exchange rate changes that sum to the original series; one for the exchange rate resulting from real and the other resulting from nominal shocks respectively.

In figure 4 these paths are graphed for the five countries in the sample. DLR, DLRP, DLRT plus the country extensions are abbreviations for the observed changes of the real exchange rate and for the changes of the permanent and transitory component.

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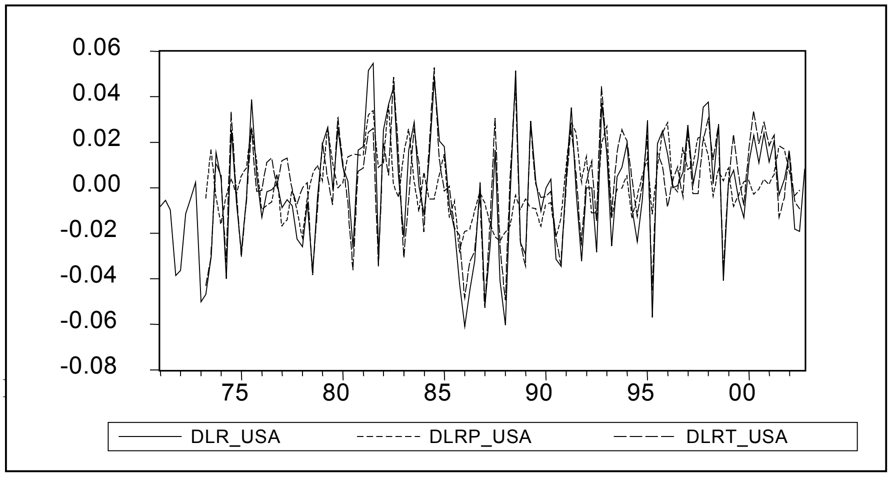
[27] The current account in relation to nominal GDP is also constructed from the *OECD's Main Economic Indicators.*

[28] Detailed results of the unit-roots are presented in the appendix A.4.

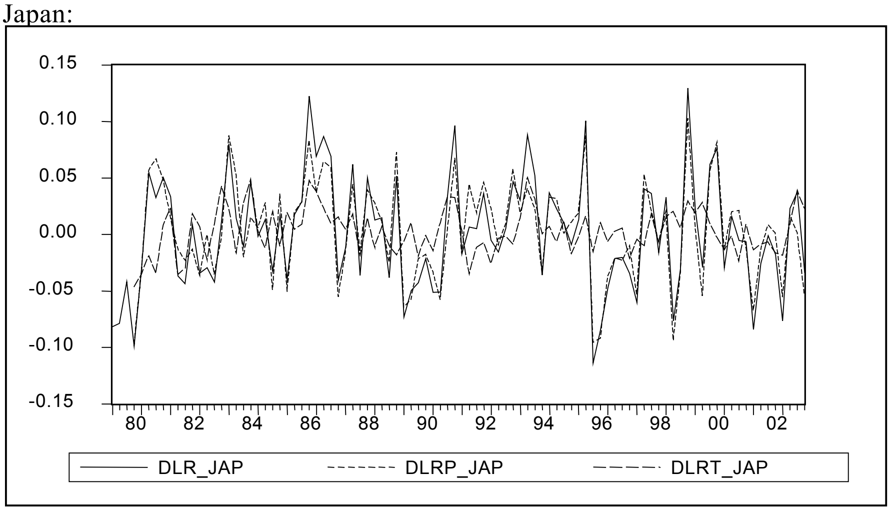
[29] For a formal treatment consult the appendix A.3.

Figure 4: Historical decomposition of exchange rate changes:

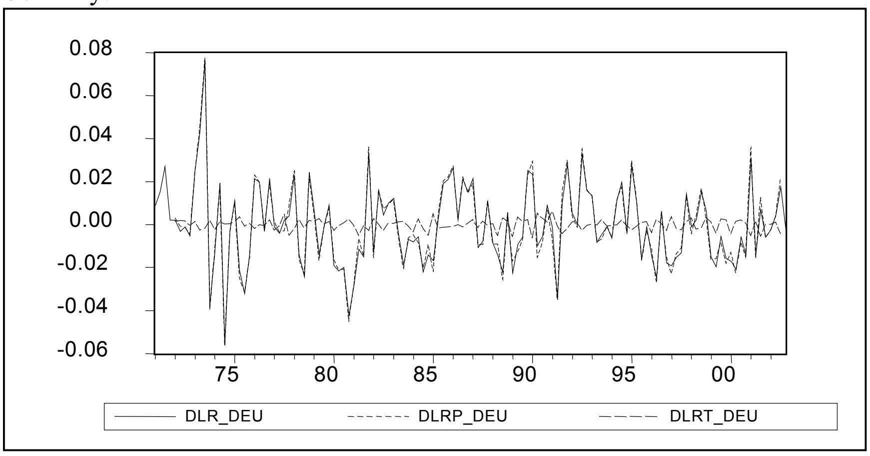
U.S.:



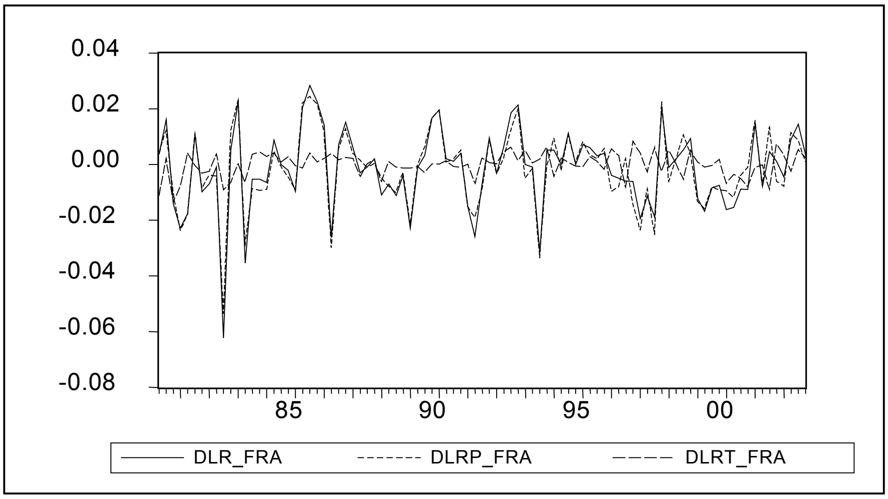
Japan:



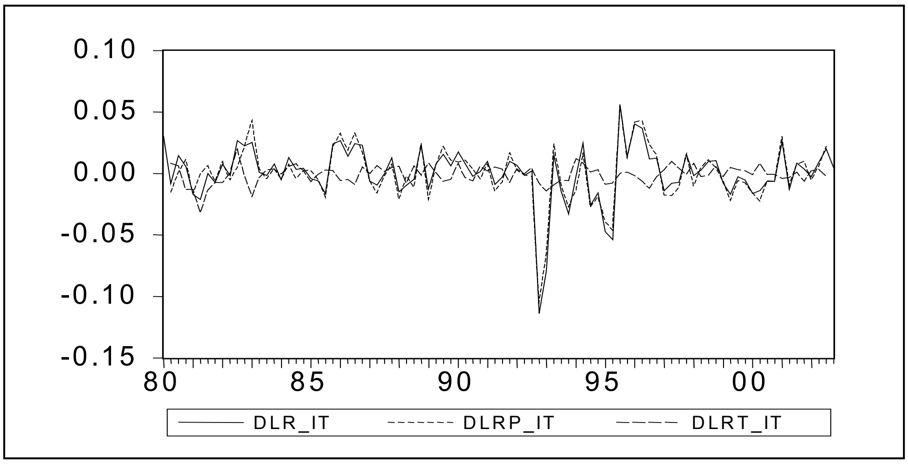
Germany:



France:



Italy:



For the four countries Japan, Germany, France, and Italy exchange rate movements seem to be dominated by the permanent component, indicating that real shocks play a prominent role. The case of the U.S. seems somewhat different, as the transitory component and hence the impact of nominal shocks is the driving force. This result is in line with the decompositions obtained by Lee and Chinn (2002). They argue that the greater dominance of nominal shocks in the U.S. real exchange rate might be due to the substantial real adjustment of the current account in the mid-1980s and due to the 1990 Gulf War transfers. Another interpretation might be that the U.S. dollar is less responsive to real shocks because of its role as the world's leading currency.

In light of the simulations in section 4, the different role of transitory exchange rate movements across countries indicate that most likely in the case of the U.S., where the transitory component dominates, an inclusion of this component into the ECM leads to an increase in the estimate of the long run pass-through elasticity. This and other implications of the simulations are reconsidered in the following interpretation of the results of the augmented estimations of the supply relationships.

5.2 Re-estimation of the supply equation

Without changing the basic specification of the ECMs, the estimations were repeated with the inclusion of one, or if significant, further lagged values of the identified transitory exchange rate change s_t. Only in the U.S. supply function a lagged transitory component turned out to be significant. For the other countries, at the most the contemporaneous transitory component matters. For Japan and Italy though, its impact on relative import prices is not significant. As table 4 shows, for all countries the transitory component enters the supply functions with the correct (negative) sign. Thus the results offer some support for the validity of models like (11), that is, agents make a downward correction to their 'regular' pricing rule in the case that a part of the observed exchange rate movement is only transitory.

The considerations of section 4 about a possible downward bias in the estimates for the long-run pass-through are also empirically confirmed. The coefficients substantially increase to values ranging from 0.8 now for the U.S. to 1.4 for France. Note that a pass-through elasticity greater than one like for France is also incompatible with long-run profit maximization in the model. However, due to the relative high standard deviation of the estimate, the outlier result for France cannot be regarded as a significant contradiction to economic reasoning. Beyond that, the poor cointegration properties of the French import supply relationship advocate some caution in interpreting the results for this country.

Most noteworthy is the strong increase of the estimated long-run pass-through for the U.S. Just like the simulations have predicted, the dominance of U.S. transitory exchange rate changes seems to be responsible for a biased estimate without the decomposition. Furthermore, the U.S. supply equation is the only one where a significant influence of the lagged transitory component could be detected. This suggests that the import suppliers to the U.S. possibly take into account higher order autocorrelation of the transitory component. As the simulations showed, this in principle can also be a contributing factor to the downward bias recorded in the foregoing estimations.

Table 4: Estimates using the decomposition:

U.S.

$$\Delta p_t = \underset{(0.03)}{-0.100}\, ect_{t-1} \underset{(0.07)}{+0.340}\, \Delta p_{t-1} \underset{(0.09)}{+0.436}\, \Delta r_t \underset{(0.01)}{+0.064}\, \Delta p_{Oil\,t} \underset{(0.10)}{-0.192}\, s_t \underset{(0.09)}{+0.166}\, s_{t-1}$$

$$ect_{t-1} = p_{t-1} \underset{(0.28)}{-0.788}\, r_{t-1} \underset{(0.04)}{-0.241}\, p_{Oil\,t}$$

$t_E = -3.80$, *adj.* R^2: 0.69, D.W.: 1.94, $Q(12) = 16.66$, J.B.: 0.24, $n = 105$

Additional impulse dummies: D91_1.

Japan:

$$\Delta p_t = \underset{(0.04)}{-\,0.176}\, ect_{t-1} \underset{(0.04)}{+\,0.148}\, \Delta p_{t-1} \underset{(0.04)}{+\,0.748}\, \Delta r_t \underset{(0.02)}{+\,0.087}\, \Delta p_{Oil\,t} \underset{(0.02)}{+\,0.098}\, \Delta p_{Oil\,t-1} \underset{(0.05)}{-0.02}\, s_t$$

$$ect_{t-1} = p_{t-1} \underset{(0.13)}{-\,0.848}\, r_{t-1} \underset{(0.06)}{-\,0.325}\, p_{Oil\,t}$$

$t_E = -3.93$, *adj.* R^2: 0.90, D.W.: 1.92, $Q(12) = 16.65$, J.B. 0.45, $n = 90$

Additional impulse dummies: D91_1.

Germany:

$$\Delta p_t = \underset{(0.05)}{-0.179}\, ect_{t-1} \underset{(0.06)}{+0.205}\, \Delta p_{t-1} \underset{(0.07)}{+0.627}\, \Delta r_t \underset{(0.01)}{+0.050}\, \Delta p_{Oil\,t} \underset{(0.01)}{+0.011}\, \Delta p_{Oil\,t-1} \underset{(0.53)}{-0.775}\, s_t$$

$$ect_{t-1} = p_{t-1} \underset{(0.21)}{-0.996}\, r_{t-1} \underset{(0.02)}{-0.137}\, p_{Oil\,t}$$

$t_E = -3.61$, *adj.* R^2: 0.71, D.W.: 2.07, $Q(12) = 14.05$, J.B.: 0.20, $n = 107$

France:

$$\Delta p_t = \underset{(0.05)}{-\,0.188}\, ect_{t-1} \underset{(0.07)}{+\,0.201}\, \Delta p_{t-1} \underset{(0.09)}{+\,0.741}\, \Delta r_t \underset{(0.01)}{+\,0.060}\, \Delta p_{Oil\,t} \underset{(0.19)}{-0.384}\, s_t$$

$$ect_{t-1} = p_{t-1} \underset{(0.28)}{-\,1.377}\, r_{t-1} \underset{(0.03)}{-\,0.157}\, p_{Oil\,t}$$

$t_E = -3.87$, *adj.* R^2: 0.66, D.W.: 1.96, $Q(12) = 10.36$, J.B.: 0.52, $n = 88$

Additional impulse dummies: D84_1.

Italy:

$$\Delta p_t = \underset{(0.04)}{-\,0.124}\, ect_{t-1} \underset{(0.08)}{+\,0.237}\, \Delta p_{t-1} \underset{(0.09)}{+\,0.489}\, \Delta r_t \underset{(0.01)}{+\,0.100}\, \Delta p_{Oil\,t} \underset{(0.13)}{-0.048}\, s_t$$

$$ect_{t-1} = p_{t-1} \underset{(0.26)}{-\,1.043}\, r_{t-1} \underset{(0.07)}{-\,0.293}\, p_{Oil\,t}$$

$t_E = -3.26$, *adj.* R^2: 0.64, D.W.: 2.14, , $Q(12) = 8.76$, J.B.: 3.55, $n = 88$

Additional impulse dummies: D80_3, D81_4.

Coefficients for the deterministic components are not shown here for expository purposes. Standard errors are in brackets. For $T = 100$ observations, the 10%, 5%, and 1% critical values of the cointegration test (t_E) are –3.66, -3.98, and -4.60 respectively. See Banerjee et al. (1998), p. 277.

Regarding the short-run pass-through elasticities (the coefficients for the contemporaneous difference in the real exchange rate), the results are broadly in line with the simulations. As predicted, the short-run pass-through for the U.S., Japan and France increases. For Germany and Italy the implied pass-through remains basically the same.

In contrast to the predictions of the simulation study, in all estimations the adjustment coefficient for the error-correction term drops in absolute value. However, due to the limited scope of the simulation exercise, we cannot really consider this as an invalidation of the previously established consistency between the bias predictions of the simulations and the actual results of the augmented estimations. It rather shows that more extensive simulations are required that allow for more general exchange rate models.

All in all, both the theoretical model and the sample in our empirical study might be too limited to find an unambiguous support for the maintained hypothesis that the distinction between permanent and transitory exchange rate changes matters in the estimation of short run and long run exchange rate pass-through. However, we consider our results as a first indication that the rationality of suppliers in international markets requires a more thorough theoretical treatment of their dynamic behavior before interpreting empirical results about the pass-through of exchange rates. Furthermore, alternative ways to separate transitory from permanent exchange rate movements should be taken into account in order to test the hypothesis that come out of dynamic optimizing behavior of import suppliers.

6. Conclusions

The benchmark analysis of this paper started with the estimation of error-correction models for the dynamic price-setting behavior of import suppliers. The underlying dynamic pricing rule was derived using a quadratic adjustment cost framework and drawing on the implications of the static oligopoly model of Dornbusch (1987). The results show that stable long-run import supply functions exist for the U.S, Japan, Germany and Italy. However, it was necessary to additionally include the oil price in order to obtain a cointegrating relationship. In line with previous studies, for most of the price functions, in particular for the U.S., the results suggest an incomplete pass-through of exchange rate changes into import prices, even in the long run.

Based on the idea that permanent and transitory components in exchange rate movements might be responsible for this result, the model for the dynamic behavior of import suppliers was extended to incorporate such a two-components model for the real exchange rate. This led to an augmented version of the previously derived dynamic import supply function that additionally takes into account a correction of the supply price for transitory exchange rate movements.

A simulation study demonstrated that the failure to account for the two components of the exchange rate might cause a bias in the conventional coefficient estimates of the import supply relationships. Above all, the simulations showed that also the long run PTM implied by several empirical studies including this one might be simply a result of this bias. Hence, the true degree of pass-through might be in fact close to one, which would be more consistent with economic reasoning that market forces should eliminate price rigidities in a long-run perspective.

Therefore, the next step consisted of a decomposition of the observed real exchange rate into a permanent and a transitory component. For this purpose, a Blanchard-Quah decomposition of the bivariate process of real exchange rate and the current account was carried out. The main implication of the decompositions is that only for the U.S. exchange rate movements are dominated by the transitory component whereas for the other countries the permanent component is crucial.

Finally, the augmented dynamic import supply functions were estimated including the obtained transitory component of the exchange rate. The most important finding is that the degree of short run and long run pass-through substantially increases in these refined estimations. In general, many of the predictions of the simulation study about the determinants of the direction and size of the bias in the estimates were confirmed. Only with respect to the adjustment coefficient of the error-correction term the augmented estimations contradicted the simulations. In turn, this underlines the indicative character of the analysis. The consideration of more general exchange rate models in both the derivation of the dynamic supply relationship and the simulations to evaluate the potential bias as well as a more comprehensive empirical analysis are desirable.

In sum, the results of the study provide a first indication that more attention should be devoted to a proper distinction between permanent and transitory exchange rate changes when estimating pass-through relationships. In this paper a decomposition of the real exchange rate is carried out that can be easily incorporated into conventional error-correction models for import supply functions and could also enrich the VAR analysis of exchange rate pass-through into domestic prices. Last but not least, possibly less biased estimates of import price relationship combined with the knowledge of the permanent and transitory component in observed exchange rate movements may contribute to improved forecasts for domestic inflation and also net external demand.

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Appendix

A.1 The prediction equation

Note first that the contemporaneous value of z_t can be also expressed as:

$$\text{(A.1)}\quad z_t = z_{t-1} + \Delta z_t \,.$$

The one-period-ahead forecast of Δz_{t+1} based on equation (5) can be written as

$$\text{(A.2)}\quad \Delta z_{t+1} = \mathrm{A}^* \Delta z_t - \Pi^* z_t \,.$$

If we substitute (A.1) into this equation, we obtain $\Delta z_{t+1} = (\mathrm{A} - \Pi^*)\Delta z_t - \Pi^* z_{t-1}$ or equivalently the one period ahead forecast for z_{t+1}:

$$\text{(A.3)}\quad z_{t+1} = z_{t-1} + (\mathrm{I} + \mathrm{A}^* - \Pi^*)\Delta z_t - \Pi^* z_{t-1} \,.$$

If we use (2) now for the prediction of z_{t+2}, from (A.3) it follows that

$$z_{t+2} = z_t + (\mathrm{A}^* - \Pi^*)\Delta z_{t+1} - \Pi^* z_t \,.$$

Employing equations (A.1) and (A.2), this expression can be again formulated in such a way that it only depends on z_{t-1}, $\Pi^* z_{t-1}$ and Δz_t.

$$\text{(A.4)}\quad z_{t+2} = z_{t-1} + (\mathrm{A}^* - \Pi^*)^2 \Delta z_t - [\mathrm{I} + (\mathrm{A}^* - \Pi^*)]\Pi^* z_{t-1} \,.$$

The forecast for z_{t+3} is obtained by shifting (A.4) one period ahead. The resulting expression will again depend on z_t und Δz_{t+1}. Like before it can transformed such that it only depends on z_{t-1}, $\Pi^* z_{t-1}$ and Δz_t. It is evident that this procedure can be continued over an arbitrary horizon. Therefore, foregoing an explicit derivation of the complex non-linear combinations that arise for further forecasting steps, for the purpose of the paper it suffices to claim that the following general forecasting equation is valid:

$$(6) \qquad E_t\ \tilde{p}_{t+i} = \iota'[G_i(\Pi^*, \mathrm{A}^*)\Delta z_t + \Pi^* \cdot H_i(\Pi^*, \mathrm{A}^*)z_{t-1} + z_{t-1}],$$

where $G_i(\cdot)$ and $H_i(\cdot)$ are complex non-linear functions depending on the parameters of the VECM that are left undefined.[30]

A.2 Bias through unaccounted transitory exchange rate changes

The simulations are based on the slightly extended model, which allows for a stationary AR(1) process of the transitory shocks.

$$r_t = q_t + s_t,$$

$$s_t = \alpha s_{t-1} + e_t, \qquad 0 \le \alpha < 1, \qquad e_t \sim iidN(0, \sigma_e^2)$$

$$q_t = q_{t-1} + u_t, \qquad u_t \sim iidN(0, \sigma_u^2)$$

If this model is considered in the derivation of the target path in the supplier's pricing rule, then the following closed form solution results:

(A.2.1)

$$\Delta p_t = -(1-\mu)[p_{t-1} - (1-\gamma)r_{t-1}] + (1-\mu)(1-\gamma)\Delta r_t - \frac{(1-\mu)(1-\gamma)\theta\mu(1-\alpha)}{1-\theta\mu}s_t.$$

It is easy to see that this collapses to equation (11) from the text if α is set to zero. The experiments are based on the following parameter values: $\mu = 0.8$, $\theta = 0.98$, $\gamma_1 = 0$ (no PTM in the long run), and values of σ_u such that together with the transitory shock the real exchange rate variable has a variance of $(2{,}5\%)^2$, which corresponds approximately to the average of the five countries in the sample. For different values of α and σ_e / σ_u 150 observations of p_t were generated according to (A.2.1). In addition, p_t is distorted by normal shocks such that the adjusted R^2 in the OLS estimation of (11) corresponds to the value of more or less 80% in the conven-

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30 Kozicki and Tinsley (1999), pp. 1307, derive dynamic decision rules for multiple decision variables with polynomial adjustment costs. In their derivations, they keep track of the complicated restrictions on the matrices that arise

tional ECMs of the sample. In the tables the triple of numbers represents the 5% quantil, the mean, and the 95% quantil of the bias in 1,000 estimates of the parameters, $1-\mu$, $1-\gamma_1$, and $(1-\mu)(1-\gamma_1)$ (the short run pass-through elasticity).

Bias in $1-\mu$:

α / σ_e/σ_u	0.0	0.5	0.9
0.5	-0.045, 0.003, 0.016	-0.036, -0.002, 0.017	-0.038, 0.000, 0.018
5	-0.159, -0.139, -0.119	-0.072, -0.052, -0.026	-0.020, -0.002, 0.016
10	-0.174, -0.157, -0.138	-0.073, -0.047, -0.012	-0.018, -0.001, 0.019

Bias in $1-\gamma$:

α / σ_e/σ_u	0.0	0.5	0.9
0.5	-0.029, -0.002, 0.023	-0.026, 0.000, 0.024	-0.024, 0.000, 0.026
5	-0.152, 0.059, 0.257	-0.166, -0.037, 0.052	-0.060, -0.015, 0.032
10	-0.257, 0.130, 0.565	-0.269, -0.101, 0.045	-0.071, -0.021, 0.028

Bias in $(1-\mu)(1-\gamma_1)$:

α / σ_e/σ_u	0.0	0.5	0.9
0.5	-0.046, -0.027, -0.008	-0.045, -0.025, -0.004	-0.033, -0.011, 0.010
5	-0.149, -0.141, -0.132	-0.128, -0.118, -0.100	-0.068, -0.051, -0.016
10	-0.157, -0.152, -0.147	-0.135, -0.124, -0.115	-0.070, -0.053, -0.034

The general tendency of the results is that all three estimates are biased downward. The biases also show a tendency to increase with the size of the variance of the transitory shock. For the estimates of $1-\mu$ and $(1-\mu)(1-\gamma_1)$ the biases tend to decrease with the size of the AR coefficient α, whereas for $1-\gamma_1$ the bias seem to peak for an intermediate value of α.

A.3 The Blanchard-Quah decomposition

The goal is to identify the real and nominal structural errors ε_t^r and ε_t^n of the following moving-average (MA) representation of the system:

(1) $\begin{bmatrix} \Delta r \\ CA \end{bmatrix}_t = C(L)\begin{bmatrix} \varepsilon^r \\ \varepsilon^n \end{bmatrix}_t$, where $C(L)$ represents a matrix polynomial of the lag operator L.

Without loss of generality, the covariance matrix of the structural errors can be normalized to $\Sigma_\varepsilon = \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix}$.

On the other hand, the estimated VAR representation is

(2) $\begin{bmatrix} \Delta r \\ CA \end{bmatrix}_t = A(L)\begin{bmatrix} \Delta r \\ CA \end{bmatrix}_{t-1} + \begin{bmatrix} e_1 \\ e_2 \end{bmatrix}_t$,

with Σ_e as the variance-covariance matrix of the reduced form errors $e_{1,t}$ and $e_{2,t}$. Thus, the relation between structural and reduced form errors is given by:

(3) $\begin{bmatrix} e_1 \\ e_2 \end{bmatrix}_t = C(0)\begin{bmatrix} \varepsilon^r \\ \varepsilon^n \end{bmatrix}_t$.

In order to identify the permanent and transitory component in both variables we need to know the elements of $C(0)$. The transformation of the VAR into the MA representation is given by:

(4) $\begin{bmatrix} \Delta r \\ CA \end{bmatrix}_t = (I - A(L)L)^{-1}\begin{bmatrix} e_1 \\ e_2 \end{bmatrix}_t$. Inserting (3) this can be reduced to

(5) $\begin{bmatrix} \Delta r \\ CA \end{bmatrix}_t = (I - A(L)L)^{-1}C(0)\begin{bmatrix} \varepsilon^r \\ \varepsilon^n \end{bmatrix}_t = C(L)\begin{bmatrix} \varepsilon^r \\ \varepsilon^n \end{bmatrix}_t$.

The identifying assumptions to recover the structural shocks results from our economic hypothesis that nominal shocks are neutral on the real exchange rate in the long run. Hence, we require the sum of all coefficients in the polynomial matrix C(L) to be lower-triangular, that is

(6) $C(1)_{1,2} = 0$.

Using (3) and (4), the restriction can be expressed as

$$\sum_{L=0}^{\infty} \left[(I - A(L)L)^{-1} C(0) \right]_{1,2} = 0 .$$

Now, first define $(I - A(1))^{-1} = \Phi(1)^{-1}$. The long run covariance matrix of the two variables Δr_t and CA_t can then be expressed as $\Phi(1)^{-1} \Sigma_e \Phi(1)^{-1}$. Alternatively, it is given by $\Phi(1)^{-1} C(0) C(0)' \Phi(1)^{-1}$, where the normalization of the covariance matrix of the structural errors comes into effect. The above restriction requires that the square root of the long run covariance matrices is lower-triangular. Thus, by a Choleski factorization of $[\Phi(1)^{-1} \Sigma_e \Phi(1)^{-1}]$ into FF' we can impose the restrictions on the MA coefficients of the system. The factored lower triangular matrix is equivalently given as $F = \Phi(1)^{-1} C(0)$.

Finally, to obtain the structural errors one has to make use of

$$\begin{bmatrix} \varepsilon^r \\ \varepsilon^n \end{bmatrix}_t = C(0)^{-1} \begin{bmatrix} e_1 \\ e_2 \end{bmatrix}_t ,$$

which can be calculated using the factored matrix F:

$$\begin{bmatrix} \varepsilon^r \\ \varepsilon^n \end{bmatrix}_t = F^{-1} \Phi(1)^{-1} \begin{bmatrix} e_1 \\ e_2 \end{bmatrix}_t .$$

A.4 Unit-root tests for the current account

	Sample	k	deterministic components	KPSS test $l = 8$	ADF test
CAY_DEU	1972:2 – 2002:3	5	C, DUM1_91	0.194	-2.53
CAY_FRA	1979:3 – 2002:4	2	C, T	0.087	-4.24***
CAY_ITA	1980:2 – 2002:3	5	C, T	0.075	-3.46***
CAY_USA	1971:1 – 2002:3	1	C, T, D83_91, D83_91T, D91_1	0.090	-3.46**
CAY_JAP	1978:3 – 2002:4	6	C	0.372*	-3.09**

An impulse dummy (D91_1) for the first quarter in 1991 is included in the test equation for the U.S. current account to capture the exceptional improvement due to international transfers for expenses during the first Gulf war. D83_91 and D93_91T allow for a mean shift and a broken trend in the years 1983 – 1991, a period that experienced extraordinary adjustments in the U.S. current account.

A.5. VARs underlying the Blanchard-Quah decomposition

Endogenous variables: (1) Δr_t , (2) rca_t										
	Germany		U.S.[a)]		France		Italy		Japan	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
adj. R^2	0.02	0.82	0.17	0.96	0.04	0.75	0.04	0.79	0.04	0.89
DW	1.96	1.99	2.07	1.96	1.98	1.92	1.98	2.15	2.08	2.00
lags	3		9		5		5		10	
AIC	-17.59		-18.73		-17.92		-15.84		-20.22	
SIC	-17.27		-17.75		-17.37		-15.28		-18.19	
Obs.	124		118		91		90		94	

[a)] The dummy variables D1_91, D83_91, and D83_91T were included in the system to guarantee stationarity of the U.S. current account.