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Do German Exporters PTM?
Searching for Right Answers in Sugar Confectionery Exports

by

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Abstract

Pricing-to-market (PTM) evidence in German sugar confectionery exports is examined, combining the original fixed-effects model of Knetter (1989) and an error-correction specification (ECM) at three frequency levels, to assess how neglecting time-series properties and the choice of frequency affect the outcomes. In order to ensure validity of unit values as price proxies, the type of competition with every destination market is evaluated, proving price-driven competition with Canada, Sweden, the UK and the US, but not with Switzerland. Results show that fixed-effect model findings of PTM might be spurious, if time-series properties of the data are not considered. German exporters seem to exploit their market power and adjust their markups to protect market shares in strategically important expanding destinations. Local currency price stabilization (LCPS) was found for Canada and the UK, while LCPS for the exports to US in fixed-effects model turned out to be an erroneous result, as cointegration was rejected. Data of the higher frequency was suggested to be preferable for PTM studies, once measurement error due to heterogeneity is minimized. Finally, using marginal costs estimates from a fixed-effects model as cost proxies in the ECM improves the quality of the model and reveals a higher degree of PTM.

Keywords: German agri-food exports. pricing-to-market, sugar confectionery.

JEL codes: C52, F14, L11

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1 Introduction

Germany is an important participant of the sugar confectionery international trade, whose share in worldwide exports accounted for around 10 % in 2011 (UN Comtrade, 2012), both in terms of value and quantity, making Germany the leading exporter in the industry. Such position allows for assuming potential market power realization by German exporters, making pricing policy a very important instrument of strategic behavior. Then the markup adjustment might become a tool of smoothing externalities transmission into destination country's prices and thus protecting the market share of German exporters on the foreign market.

Pricing-to-market (PTM), introduced by Krugman (1987), is an example of such strategic behavior, where markup adjustment takes place as a reaction of the exporter towards the change of the exchange rate between his and the destination country's currency. Empirical literature often associates PTM with the local currency price stabilization (LCPS) mechanism, which implies that a part of the exporter's currency appreciation is being absorbed via a decrease in the markup charged by the exporter and thus the price he sets in terms of the exporter's currency. In this case only a part of the exporter's currency appreciation is transmitted into the price, paid by the importer in his local currency that keeps price level in the destination country relatively rigid.

PTM has been a subject of interest of many empirical studies in the few last decades, particularly since Knetter (1989) came up with an empirical model, combining trade theory with a new industrial organization approach in order to distinguish between changes in marginal costs and markup fluctuations in the price setting of a producer, exporting to various destinations and employing market power. Being an accessible, yet powerful tool to achieve such a goal, the original Knetter model relies on a number of assumptions, which are crucial to ensure the outcomes' reliability.

The main question this paper investigates is whether German exporters of sugar confectionery use their strategic advantages and differentiate between destination markets, adjusting prices in a way to smooth the Euro appreciation (or depreciation) effect on domestic prices of important destination markets. The only study in this direction so far was carried out by Glauben and Loy (2003), who applied a dynamic PTM model and the residual demand elasticity (RDE) approach to German exports of food and beverages, including sugar confectionery, and concluded that competitive conduct prevails on those markets.

In this paper a much larger time span is analyzed for a partly different set of destination countries, within both static and dynamic frameworks, addressing in the meantime a number of questions related to the model specification and outcomes' reliability, which often stay out of scope of PTM studies. Among those: are results of the original Knetter (1989) model reliable, once variables included into the estimation are nonstationary or are of different order of integration and how does the time frequency choice affect outcomes; how to deal with the potential heterogeneity problem and minimize the measurement error; and how to introduce marginal cost changes in a dynamic model.

Stationarity of variables under consideration is one of the implicit model assumptions, which does not always hold true, as exchange rates and prices often behave nonstationary, making the inference from the model questionable (e.g. Baffers, 1997). To assess how neglecting time-series properties and the choice of frequency affect the outcomes, the PTM evidence in German sugar confectionery exports is examined, combining the original fixed-effects model of Knetter (1989) and an error-correction specification (ECM) at three data frequency levels (monthly, quarterly and annual). As the order of variables' integration is often ambiguous as a result of different stationarity tests, which makes both conventional and time-series approaches inapplicable, the bounds testing approach (Pesaran et al., 2001) is used to overcome this problem and estimate relationships between parameters irrespective of the order of integration.

In order to ensure the product homogeneity and minimize the measurement error (e.g. Lavoie and Liu, 2007), the most disaggregated 8-digit export data is employed and the validity of unit values as price proxies is tested, following Aiginger (1997) and Gehlhar and Pick (2002), who proposed to compare the net trade flow direction with unit value differences in order to assess the driving power of the competition on the market. The pre-validation of the unit values as price proxies has not been yet applied within the PTM-framework. Finally, in order to include cost shifts into the dynamic model, time-specific estimates from the original fixed-effects model are introduced in the ECM specification as marginal costs proxies, which is also a new approach in dynamic PTM studies to the best of my knowledge.

In order to investigate if and to what extent German exporters of confectionery price-to-market, exports to five destination countries (Canada, Sweden, Switzerland, the UK and the US), which jointly account for around 30% of German exports of gum and jelly confectionery (code 17049065 in CN8 trade classification) during 1991-2011 are analyzed.

The remainder of the paper is structured as follows: Section 2 provides a literature overview on the crucial points addressed in the paper, Section 3 presents the model and empirical specifications to be estimated, Section 4 briefly describes the data, Section 5 gathers empirical results, and Section 6 provides a summary and concludes.

2 Literature overview

2.1. PTM in empirical literature

When markets are segmented and exporters have market power, markup adjustments might become an important instrument of their strategic behavior. In order to stay competitive and protect market shares exporting firms might maintain stable prices in destination markets by adjusting their markups and cutting prices in exporter's currency in order to partly absorb or to smooth the externalities transmission. Such markup adjustment due to exchange rate fluctuations became known as PTM due to Krugman (1987) and was assessed empirically in works of Knetter (see e.g. Knetter, 1989, 1993).

For an imperfectly competitive market it is plausible to assume, according to Dunn (1970), that large suppliers avoid periodic changes in domestic and foreign prices in response to short-term variations in market conditions, following that the relationship between the price in exporter's and importer's currency is a result of the estimated long-run equilibrium exchange rate, while transitory changes might be simply ignored. When a significant part of the product is exported, exchange rate volatility can be easier controlled via markup adjustment than via quantity changes, following that for important markets local currency prices are held stable by the exporter who wishes to maintain his market share (Hatemi-J and Irandoust, 2004). Froot and Klemperer (1989) also suggest that due to the possibility of price adjustments, import prices are more sensitive to the expected future rather than to the current exchange rate, and temporary fluctuations would be reflected in the change of the exporters' markups instead of the price paid by the importer. According to Knetter (1992), PTM is more pronounced when an industry allows for setting high markups over marginal cost, but the home firms have a relatively modest share of the foreign market and need to keep their market shares by markup adjustments, as they do not have much influence over the equilibrium market price on the destination market. PTM is then lower when the industry is rather competitive, but the source countries' firms dominate the destination countries market, resulting in a nearly complete pass-through.

The PTM phenomenon was intensively investigated after Knetter (1989) proposed an easy-to-implement solution for distinguishing between competitive markets and segmented markets with constant and non-constant elasticity of demand, where both country-specific markups and destination specific markups changes due to exchange rate fluctuations are taking place. His study of German and US exports (including such products as onions, bourbon, orange juice, breakfast cereal, refrigerators and switches for US exports, and fan belts, titanium dioxide pigment, small cars, large cars, beer, white wine, sparkling wine, potassium chloride, mining was and motorcycles for Germany) on quarterly data from 1977/78 to 1985/86 proved PTM adjustments for many German exports, with negative coefficients occurring three times more frequently than positive ones, implying local currency price stabilization (LCPS) in destination markets. In German exports to the US, LCPS was revealed for every product group, while American exports were found to be less sensitive to exchange rate fluctuations.

Existing evidence on the price adjustments is mixed. Knetter (1997) noticed that PTM magnitude and its persistence vary across destinations, especially between Continental Europe, the US and the UK, while within the European area they are less pronounced. The degree of PTM of German exports differs between products and destinations, and is more pronounced in chemical exports to US and Japan (Falk and Falk, 2000).

Germany, the US and Japan seem to be the source countries of the highest interest (see e.g. Knetter, 1993; Gagnon and Knetter, 1995; Feenstra et al., 1996), and PTM is widely declared to be more pronounced in German and Japanese than in American exports, where dollar prices rather adjust in a way that amplifies the effect of exchange rate fluctuations on domestic currency prices (e.g. Mann, 1986; Knetter, 1989). Goldberg and Knetter (1997) suggested nearly one half of the exchange rate effect to be a reasonable estimate for the PTM of German exports, when around half of the source country's exchange rate appreciation is being offset via markups.

Empirical literature also provides examples of less common source countries: Yumkella et al. (1993) check PTM evidence in rice exports of the US and Thailand; Brown (2001) investigates the Canadian canola market; Griffith and Mullen (2001) concentrate on rice exports and Swift (2004) estimates PTM in Australian milk and meat exports; Bowe and Saltvedt (2004) study the fishing industry of Norway; Balaguer et al. (2003) and Silvente (2005) estimate PTM and market power extension for tile ceramic products between Spanish and Italian exporters.

Agricultural products, machinery and chemicals products are the most intensively investigated sectors (see e.g. Kasa, 1992; Gagnon and Knetter, 1995). Besides the mentioned above studies in agricultural exports, alcoholic beverages and wheat are often analyzed within the PTM framework. Knetter (1989, 1992, etc.) studied, among others, exports of beer, white and sparkling wine, orange juice, onions, bourbon, olive oil, cocoa powder, and yellow corn. Kasa (1992) and Gil-Pareja (2002) also included beer and/or wine to their investigation of PTM in German exports. Wheat exports were considered by Pick and Carter (1994), Pick and Park (1991), Carew (2000) and Carew and Florkowski (2003), who also included pulse, apples and tobacco in their studies.

By now, studies of PTM in German food and agricultural sectors are rather limited. The most detailed investigations in this area were conducted by Knetter (1989, 1992, etc.) and Kasa (1992). Attempts to evaluate the exercise of market power by German food and beverage export industries on international markets were made by Glauben and Loy (2003), who focused on beer, chocolate, cocoa powder and sugar confectionery, exported to the US, Canada, France, Belgium, Italy and the UK from April 1991 to May 1998. In their paper the PTM model was applied together with a RDE approach to a monthly data, taking its time-series properties into account. To the best of my knowledge, this is the only study, which includes German sugar confectionery exports into the investigated product groups.

2.2. Unit values as price proxies

According to Lavoie and Liu (2007) most pricing-to-market studies use export unit values due to its availability for most of the markets and its relatively low costs[1]. However, unit values as price proxies have received a lot of critique, as homogeneity of products, involved into testing of the market integration/segmentation is often a questionable assumption, causing unit values to be perceived as the main source of potential measurement error in PTM models[2]. According to Falk and Falk (2000) and Silver (2010) using customs unit values while measuring price development in international trade potentially leads to significant biases. Aggregation problems and

[1] According to their research only three PTM studies were based on the product-level data.

[2] Gehlhar and Pick (2002) found that a lot of American food exports are characterized by product differentiation and non-price driven competition.

possible heterogeneity within product groups might cause unit value changes to be rather a reflection of the product composition or quality change than price shifts (see e.g. Aiginger, 1997; Gil Pareja, 2002). Aw and Roberts (1986) and Silvente (2005) name excessive volatility and effects on price due to changes in product quality over time as the most serious problems of using unit values. Dullek et al. (2005) and Fontagné et al. (2006) even defend unit values exactly as a measure of the traded goods quality, while Kinoshita (2009) concludes that customs unit values reflect both the price and quality changes and thus overestimate the price increase by the amount of the increase in quality.

When it comes to measuring the PTM effects, it should be considered that findings of price discrimination might be in some cases explained by neglecting the importance of the product differentiation (Sexton and Lavoie, 2001). While Knetter (1989, 1995) and Feenstra et al. (1996) typically argue that common quality differences will be captured by time-effects, and systematic differences in product quality across destination markets should be captured by the country-specific effect introduced into the model, Lavoie and Liu (2007) examined the incidence of spurious PTM alarms, and found false evidence of PTM even in the simulated perfect competition settings, once unit values are used as price proxies, and concluded that the magnitude of such bias depends heavily on the level of product differentiation. To overcome those problems, use of the market power by the exporter has to be a valid assumption with a solid background, while the heterogeneity problem and potential measurement error can be minimized by using the maximum disaggregated product level data available (see e.g. Lavoie and Liu, 2007; and Kinoshita, 2009) and by testing the competition type, prevailing on the market. For instance, Aiginger (1997) proposed to use unit values in order to discriminate between the price and quality competition and divided all markets in elastic and quality-dominated with subdivision according to the relative size of unit values. This approach was further adopted by Gehlhar and Pick (2002), who rearranged the four-quadrant scheme of Aiginger (1997) and came up with a taxonomy of trade flows, allowing for checking whether unit value differences for particular goods are consistent with the expected net direction of trade. This is the case once a negative unit value difference (export unit value minus import unit value) comes along with a positive trade balance (value of exports minus value of imports), which would imply that cheap prices allow exporters to sell more than they import in return. The same holds for the inverse situation, where relatively high unit values are followed by a negative trade balance. These two cases would be examples of price competition in the homogeneous goods' trade, while in the two other cases unit value difference is not consistent with expected trade direction and should be explained by other factors than price (e.g. quality). Then product homogeneity cannot be proved and unit values might be a bad proxy for the exported good price (Table 1).

This approach is adopted in Section 4 to ensure the validity of unit values as export price proxies in estimating the PTM effects. For higher security the highest order of disaggregation available in the Eurostat database is used for the analysis.

Table 1. Validity of unit values according to the taxonomy of trade flows

<table>
<tr><td colspan="2" rowspan="2"></td><td colspan="2">Trade flow balance</td></tr>
<tr><td>Positive</td><td>Negative</td></tr>
<tr><td rowspan="2">Unit value difference</td><td>Positive</td><td>Non price-driven competition</td><td>Price-driven competition market (Unit values are valid price proxy)</td></tr>
<tr><td>Negative</td><td>Price-driven competition (Unit values are valid price proxy)</td><td>Non price-driven competition</td></tr>
</table>

Source: Adapted from Gehlhar and Pick, 2002.

2.3. Data frequency and time series issues in PTM studies

Using the most disaggregated product level data is often an argument towards addressing a lower data frequency, since high frequency data might have more missing observations and seasonal variation in shipments and reporting lags, which could increase the amount of noise in the unit values (e.g. Knetter, 1992 and 1995). Following such arguments, most of the authors employ the data of the lower frequency, which keeps the sample size too short to investigate time-series properties of the data.

As prices and exchange rates are often nonstationary (e.g. Adolfson, 2001), assuming the contrary for such series might lead to improper conclusions about relationships between them. When non-cointegrated nonstationary series are included in the model, regressions could be spurious (e.g. Engle and Granger, 1987; Baffers, 1997), as application of conventional econometric technics to non-stationary data might lead to misleading results and erroneous inference.

If prices and exchange rates are indeed nonstationary and are cointegrated, the original Knetter's model in levels (e.g. Knetter, 1989) will still provide long-term elasticities, which might be the reason for estimations in levels to be prevailing among researchers (see e.g. Carew and Florkowski, 2003; Hatemi-J and Irandoust, 2004). However, only in Knetter (1992) it is explicitly stated that exchange rate coefficients should be interpreted as long-run elasticities. Besides that, such model specification allows no dynamic adjustments, which could be fixed by estimating the model in first differences (e.g. Knetter, 1993, 1995; Gil-Pareja, 2002). This often solves the nonstationarity problem, but reveals no information about the long-term relation between variables, and is regarded as a misspecification, especially for agri-food markets (Larue, 2004). To capture both short and long effects, PTM can be estimated via an ECM framework (see e.g. Feenstra et al., 1996; Glauben and Loy, 2003; Bowe and Saltvedt, 2004).

The situation becomes more complicated once variables in the estimation are of different order of integration, as conventional measures then are inapplicable due to non-stationarity of regressors, but cointegration technics are neither feasible, as they require the same order of integration of parameters[3]. It is, however, well documented, that unit root tests often suffer from poor size and low power properties, especially in small samples (see e.g. Harris, 1995; Larue, 2004), which could lead to erroneous results about the order of integration of variables.

The power of cointegration tests is also actively discussed in the literature. According to Banerjee et al. (1993, 1998) and Kremers et al. (1992), when variables are cointegrated, the single equation ECM is more powerful than the residual based Engle-Granger test, as it does not push the short-run dynamics into the residual term, does not suffer in finite samples from imposing a potentially invalid common factor restriction (see e.g. Pattichis, 1999; Tang, 2005), and is more powerful for moderate samples (Jansen, 1996). The unrestricted ECM might also be used to conduct the bound testing procedure, proposed by Pesaran et al. (2001). Monte Carlo simulations suggest that the bounds approach performs better in small samples (see e.g. Haug, 2002, or Narayan and Smyth, 2003, for argumentation), and can be applied irrespective of the order of integration of the explanatory variables (I(0), I(1) or mutually integrated). Belke and Polleit (2006) applied the bound testing approach in order to avoid spurious regression problems, caused by the mixed order of integration of used time series. Conventional cointegration technics fail to deal with such settings, as they require the same degree of integration of non-stationary series. The bound testing approach also helps to solve the problem of the low power of stationarity tests, which are fragile to various factors and introduce a further degree of uncertainty in the analysis (Pesaran et al., 2001), as pre-testing is not required. Another important advantage of the bounds test procedure is that estimation is possible even when the explanatory variables are endogenous; it is sufficient to correct for residual serial correlation (Pesaran and Shin, 1999) and it compensates for not applying the structural break unit root test (Belke and Polleit, 2006).

Sample size becomes crucial when evaluating the power of cointegration tests. Pattichis (1999) and Tang and Nair (2002) found out that Engle-Granger and Johansen tests are not reliable for small samples. Mah (2000) estimated import demand for IT products for annual data and stated that for small samples a single-equation ECM cointegration test is neither reliable. Tang (2005) could not come up with a definite answer whether Pesaran's bound test performs better than conventional tests, however it is widely used in empirical literature (see e.g. Tang, 2001; Ziramba, 2007, with 23 and 36 observations in their data sets respectively). Narayan and Smyth (2003) mentioned that the critical values can deviate substantially from those derived in Pesaran et al. (2001) for small samples, while Kremers et al.(1992) concluded that for small sample sizes no cointegration at all can be found among variables which are integrated of order one. Hakkio and Rush (1991) underlined that rejecting cointegration in a relatively small sample might be simply a result of low power of the tests.

Such limitations could lead to the idea of using larger samples, or at least data of higher frequency, when time spans are fixed. Even though the length of the time

[3] See Baffers (1997) for a deeper insight into this problem.

series is often stated to be more important than the frequency of observations (see e.g. Shiller and Perron, 1985; Lahiri and Mamingi, 1995), using high frequency data for cointegration analysis can compensate for the power loss and reduce distortion, especially for relatively short time spans. Zhou (2001) concluded, that using a small sample of annual observations instead of more observations of higher frequency data often results in significant loss of the cointegration tests power, while Glauben and Loy (2003) mention, that using aggregated data for PTM investigations does not make much sense, as most of the system's disequilibrium tends to correct within the first few months.

In order to ensure the reliability of the outcomes, all the models in this paper are estimated using monthly, quarterly and annual data, and various stationarity and cointegration tests at every frequency are applied to see whether results change considerably between different specifications or data frequencies.

3. Model and empirical specifications

The foundation of Knetter's model comes from a producer's maximization problem, where a perfectly competitive market would assure an equality between price and marginal costs, while imperfect competition allows producers to add a markup, which would be fixed in the case of constant demand elasticity in the destination market and flexible when demand elasticity of the importer is perceived as non-constant[4]:

$$p_{i,t} = c_t \frac{\epsilon_{i,t}}{\epsilon_{i,t}-1}, i = 1, \dots, N \text{ and } t = 1, \dots, T \text{ , (1)}$$

where *p* is the price in terms of the exporter's currency, *c* is the marginal cost of production and $\epsilon_{i,t}$ is the elasticity of demand with respect to the local currency price in destination market *i* in period t.

Following the line, in perfectly competitive markets, free on board (FOB) prices for homogeneous products expressed in terms of the exporter's currency should be the same for all destinations, while for segmented markets prices might differ, depending on the markup, which is determined by the elasticity of demand in the destination market.

To distinguish between the above mentioned three alternative hypotheses of the market structure (perfect competition and imperfect competition with constant/non constant elasticity of demand) a fixed-effects regression model based on German export panel data is estimated, following Knetter (1989):

$$ln\, p_{i,t} = \theta_t + \lambda_i + \tau_i \, ln\, e_{i,t} + \varepsilon_{i,t} \text{ , (2)}$$

where $p_{i,t}$ is the export unit value, expressed in terms of the exporter's currency (Euro per kilo), $e_{i,t}$ is the exchange rate (nominal or inflation-adjusted), expressed as units of destination country's currency per unit of exporter's currency (units of national currency per 1 Euro), with θ_t representing the time-specific effect, λ_i – the

[4] See Knetter (1989) for the details of the derivation.

destination-specific markup, τ_i – the PTM coefficient, and $\varepsilon_{i,t}$ - the disturbance term, assumed to be iid. One time effect and one country effect should be dropped from the estimation to avoid singularity of the regressor matrix.

From Equation (2) a case of perfect competition would imply equality between FOB prices and marginal costs across destinations, with λ and τ being equal to zero, and θ reflecting the development of marginal costs over time. Constant elasticity of demand would allow for different fixed markups, set to destination markets, when λ is no longer equal to zero, while τ still is. τ equal to zero implies the case when isoelastic demand schedules are observed, where the markup does not depend on exchange rate fluctuations and the exchange rate changes do not affect exporter prices and are fully passed through into prices in importer's currency. A τ-coefficient significantly different from zero would mean that export prices depend on exchange rates fluctuations, due to demand elasticity changes as a consequence of fluctuating local currency prices. A negative exchange rate coefficient would imply the LCPS, when exporters are adjusting their markups accordingly in order to keep the price in the destination level relatively constant (for example, by absorbing a part of the Euro appreciation). Positive exchange rate elasticity would be an indicator for amplifying the effects of the exchange rate fluctuations by the exporter, when the demand is perceived to become less elastic as the local currency price rises. According to existing literature on PTM, in the case of Germany we suppose to see rather LCPS (negative τ-coefficients), than the opposite case[5].

The second model to be estimated is the error-correction specification of the Knetter model, which allows investigating both long- and short-term dynamics[6]. The ECM is estimated using single-equation framework (Equation 3) and the two-stages Engle-Granger approach (Equations 4 and 5).

$$\Delta \ln p_{i,t} = \alpha_i + \beta_i\left(\ln p_{i,t-1} - \gamma_i \ln e_{i,t-1} - \delta_i T_{i,t}\right) + \\ + \sum_{q=1} \eta_{q,i} \Delta \ln p_{i,t-q} + \sum_{q=0} \kappa_{q,i} \Delta \ln e_{i,t-q} + \varepsilon_{i,t} \quad (3)$$

In this single-equation ECM coefficients of the first differences represent short-run dynamics, while coefficients of the lagged level variables are the long-run elasticities, where β_i is the speed of adjustment of the system to its equilibrium, γ_i is the long-term PTM coefficient (elasticity of the exporter's price with respect to exchange rate changes) and T stands for trend, which is used here as a demand shifter[7]. When the model is estimated by OLS, only a product of β_i and γ_i can be obtained, so the estimated coefficient has to be divided by -β_i to obtain the long-term elasticity. As the significance of the coefficient of the lagged exchange rate cannot be tested within the

[5] See, for example Knetter (1989), Goldberg and Knetter (1996), Glauben and Loy (2003).

[6] A simple error-correction setting is chosen as there is no reason to assume endogeneity of the exchange rate in the estimated model, once a disaggregated product group, which accounts for less than 1% of total German exports, is investigated.

[7] As GDPs among the destination countries, once adjusted with the GDP deflator and normalized to some point in time reveal similar development across countries, a simple trend line was chosen to proxy income shifts.

model using standard methods, the two stages Engel-Granger procedure is also applied and its outcomes are reported.
The two-stage Engle-Granger ECM takes the following form:

$$ln\, p_{i,t} = \alpha_i + \gamma_i\, ln\, e_{i,t} + \delta_i T_{i,t} + \upsilon_{i,t} \quad (4)$$

$$\Delta\, ln\, p_{i,t} = \alpha_i + \beta_i \upsilon_{i,t-1} + \sum\nolimits_{q=1} \eta_{q,i} \Delta ln p_{i,t-q} + \sum\nolimits_{q=0} \kappa_{q,i} \Delta\, ln\, e_{i,t-q} + \varepsilon_{i,t} \quad (5)$$

In the last equation $\upsilon_{i,t-1}$ are residuals from the 1st stage, and all the other coefficients are similar to (3).
The order of the ECMs' dynamic lag structure is chosen according to the Akaike information criterion (AIC) to avoid residual autocorrelation and to capture possible important deferred effects. Selection from up to 36 lags for monthly data and 12 lags for quarterly data was applied, which should be enough to capture the long-term relationship (e.g. Enders, 1995). For yearly data, only lag 1 was used in order to conserve degrees of freedom. However in all the cases system adjustments happen much faster.
Inference on the relationship between the export price and the exchange rate is additionally tested by means of the bounds approach in order to deal with the ambiguity of stationarity tests outcomes. The bounds testing is based on an unrestricted single equation ECM, and checks the null hypothesis of no level relationship between series of the equation through a joint significance tests on lagged variables based on the Wald test[8].
However, estimating an ECM might result in some information loss compared to the panel Knetter-type model, as it deals with a set of individual equations, and thus assigns the average country-specific effects and cost changes to a constant, while deviations from the average level accumulate in the error term. To overcome this limitation, the time-specific coefficients θ from the panel estimation (Equation 2) are introduced into dynamic ECMs (Equation 3 or Equations 4 and 5) as proxies for marginal costs, allowing for combining advantages of both approaches.
Hence, the main innovation of this study takes the following form:

$$\Delta\, ln\, p_{i,t} = \alpha_i + \beta_i \left(ln\, p_{i,t-1} - \varphi_i \theta_{i,t-1} - \gamma_i\, ln\, e_{i,t-1} - \delta_i T_{i,t}\right) + \\ + \sum\nolimits_{q=1} \eta_{q,i} \Delta ln p_{i,t-q} + \sum\nolimits_{q=0} \mu_{q,i} \Delta\, \theta_{t-q} + \sum\nolimits_{q=0} \kappa_{q,i} \Delta \ln e_{i,t-q} + \varepsilon_{i,t}, \quad (6)$$

while the two-stages version is being constructed analogous to the Model 2 (Equations 4-5).
According to Knetter (1989) including various destination markets for the product may still provide an unbiased measure of the period to period changes in marginal

[8] Under the null hypothesis of no cointegration the asymptotic distribution of the F-statistic is not standard, and the critical values for 5 different ECM specifications can be found in Pesaran et al.(2001). When computed statistics exceed the upper critical value, variables are said to have a long-run relationship, no matter of their degree of integration, when the F-statistic is lower than the lower bound, the null hypothesis of no-cointegration cannot be rejected. If the F-statistic falls between the bounds a straight forward conclusion cannot be made and other techniques should be applied.

costs, which is used in Equation 6 to implicitly account for cost shifts over time in the set of individual ECMs. All the ECMs were additionally estimated using the SUR-procedure. Results can be obtained upon request.

4 Data and descriptive statistics

The present study uses monthly data from January 1991 to December 2011 for German sugar confectionery exports (code CN8 – 17049065 - Gum confectionery and jelly confectionery, including fruit pastes in the form of sugar confectionery) to Canada, Sweden, Switzerland, the United Kingdom and the United States. Two conditions had to be fulfilled for a country to be included into the sample: the data had to be available for each monthly observation[9] and the national currency had to be different from Euro and show some variation in the exchange rate. To obtain lower frequencies the export data was aggregated by summing up over the subsequent period.

The share of destination countries in Germany's sugar confectionery exports and the share of German imports in all sugar confectionery imports of the destination markets vary significantly across countries and sample periods (Appendix 1). While the share of exports to Sweden and Switzerland in German exports remained relatively constant, Canada and the UK nearly doubled their shares. Only the share of the US experienced a drop from 18.1 to 9.4 per cent during the period of the study. The US also considerably reduced the share of imports from Germany in their sugar confectionary imports (from 14.9 to 2.8 percent).

Although total exports to sample countries on average do not exceed 30% of all sugar confectionery exports, the analysis of the importance of different source countries in imports of destination countries shows that Germany is an important trading partner for all of the countries in the sample (1st on average for Switzerland and the UK, 3d – for Canada and Sweden, and 5th for the US), which might be a reason for expecting potential market segmentation and the realization of the PTM.

Unit values, used in the model as export prices proxies, were calculated by dividing the export volume in Euros (FOB) by exported quantity in kilos, taken from Eurostat database for CN8 classification for each trading partner. Export unit values show significant variation both between destination countries (Table 2) and within the analyzed period, that might be the first indicator of the market segmentation, as FOB prices set to various destinations differ:

[9] The only exception has been made for Canada, where missing observations (less than 6% of the sample) were found in stable time periods and were interpolated as an average of the annual value.

Table 2. Average unit values (Euros/kilo) and nominal exchange rates (1991-2011)

	Canada	Sweden	Switzerland	UK	USA
Unit value	2.57 (0.51)	2.12 (0.23)	2.83 (0.12)	2.31 (0.11)	2.66 (0.58)
Exchange rate	1.53 (0.11)	9.02 (0.09)	1.57 (0.08)	0.73 (0.12)	1.21 (0.14)

Notes: Coefficients of variation are shown in parentheses.

Monthly exchange rates series are taken from the IFS and the Bundesbank and are normalized to the level of 2005 and adjusted by consumer price indices (CPI) in the destination countries to control for the impact of the inflation on foreign markets. CPI monthly series are taken from the OECD database. Both exchange rates and unit values enter all estimations in the log form (see Appendix 2 for descriptive statistics).

5 Empirical results

5.1. Validation of unit values as price proxies

Most of the destination markets in the sample can be described as price-driven (Table 3) as unit value difference is consistent with the expected trade direction, thus the unit values can be used as valid proxies for export prices.

Only in the case of Switzerland, Germany has a negative trade balance and a negative unit value difference, which implies that a price advantage does not lead to a positive trade balance, hence the traded goods cannot be regarded as homogeneous and unit values cannot be proved to be reliable export price proxies and PTM estimates for this destination should be treated with caution. However, in the next section Switzerland is chosen to be a reference country to assess the country-specific markups. As all destination markets are compared to the similar benchmark, results should not be distorted.

Table 3. Bilateral comparisons of trade flows and unit values, average from 1991-2011

Trade partner	Trade flow balance, Mio. USD	Unit value difference, USD per kg	Competition type
Canada	9.4	-4.3	price driven
Sweden	10.2	-0.8	price driven
Switzerland	-9.0	-2.2	driven by other factors
UK	25.2	-2.4	price driven
USA	37.5	-2.5	price driven

Notes: According to Gehlhar and Pick (2002).

5.2. Fixed-effects model outcomes

To avoid singularity of the regressor matrix Switzerland was chosen as a reference country and January 1991 as a time benchmark. Country-specific estimates and time effects (Figure 1) are very similar between estimations including nominal and adjusted exchange rates, so only the latter results are reported (Table 4), following Knetter (1989), who suggested to exclude the effect of the inflation on the destination market in order to observe pure PTM adjustments.

All the country-specific coefficients are highly significant, confirming the market segmentation by German exporters, who set different fixed markups to trade partners.

Output suggests Switzerland to have the highest fixed markup, as all the other country-specific coefficients are negative. A quite similar result was obtained by Falk and Falk (2000), who estimated the average German exports to Switzerland to be charged 10 per cent higher than to other European destinations.

Significant PTM-coefficients were found for most of the destination countries based on the monthly data. LCPS varies between -0.205 (the UK), -0.409 (the US) and -0.504 (Canada), and a positive coefficient is found for Sweden but only significant at 10% level. Empirical literature has already registered positive coefficients for exports to Sweden (e.g. Gagnon and Knetter, 1995), whose fixed markup is the lowest for this destination market (-0.302). No PTM is found for Switzerland.

For the yearly data no PTM could be found for any country, except for Canada. Significance of the coefficients declines with the data aggregation, suggesting rather fast price-adjustments, which could be a finding in favor of sticking to disaggregated data and estimation of the error-correction model to capture also the short-term dynamics.

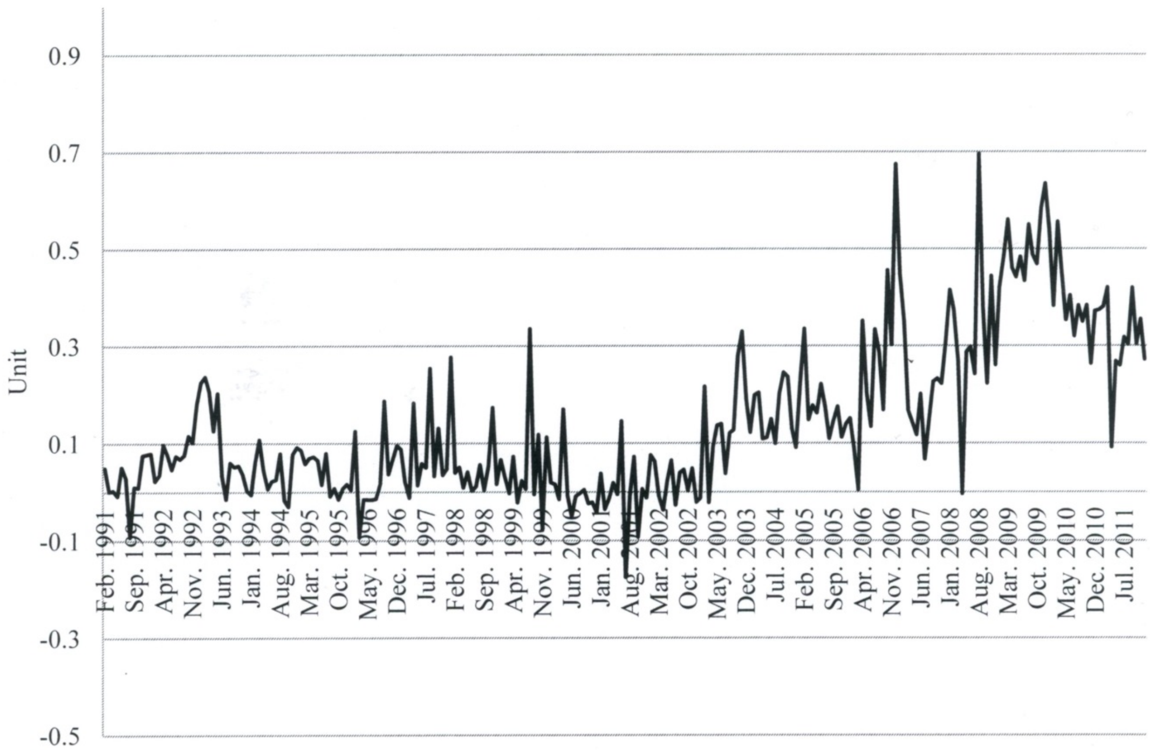

Figure 1. Estimated time-effects from the fixed-effects model
(cpi-adjusted exchange rate, monthly data)

Table 4. Results of the fixed-effects model estimation

Destination	Monthly estimates		Quarterly estimates		Yearly estimates	
	λ	τ	λ	τ	λ	τ
Canada	-0.117*** (0.025)	-0.504*** (0.120)	-0.140*** (0.033)	-0.555*** (0.160)	-0.117** 0.049	-0.580** (0.277)
Sweden	-0.302*** (0.017)	0.329* (0.198)	-0.316*** (0.024)	0.241 (0.289)	-0.314*** 0.042	0.252 (0.571)
Switzerland		0.027 (0.115)		0.029 (0.184)		0.039 (0.356)
UK	-0.174*** (0.022)	-0.205** (0.100)	-0.168*** (0.0323)	-0.297** (0.136)	-0.165*** 0.063	-0.302 (0.253)
USA	-0.141*** (0.030)	-0.409*** (0.113)	-0.142*** (0.046)	-0.442** (0.173)	-0.113 0.091	-0.485 (0.347)
R-squared		0.471		0.580		0.628

Notes: 1.Standard errors are shown in parentheses. 2. *, **,*** indicate significance at 10, 5 and 1%.

5.3. Stationarity testing

As exchange rates and prices are often nonstationary, testing for unit-roots is required. Here two tests, the Augmented Dickey-Fuller (ADF) test and the Phillips-Perron (PP) test, with the null hypothesis of a unit root, and the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test with the alternative null hypothesis of stationarity are conducted (Table 5).

Table 5. Stationarity of prices and exchange rates, 1991-2011

		Canada		Sweden		Switzerland		UK		USA	
		Ln P	Ln E	Ln P	Ln E	Ln P	Ln E	Ln P	Ln E	Ln P	Ln E
Monthly (252 obs.)	ADF	I(0)*	I(1)	I(0)*	I(1)	I(0)*	I(1)	I(0)*	I(1)	I(1)	I(1)
	PP	I(0)*	I(1)	I(0)*	I(1)	I(0)*	I(1)	I(0)*	I(1)	I(0)*	I(1)
	KPSS	I(1)*	I(0)	I(1)*	I(0)	I(0)	I(0)	I(1)*	I(1)*	I(1)*	I(1)*
Quarterly (84 obs.)	ADF	I(0)*	I(0)*	I(0)*	I(0)*	I(0)*	I(1)	I(1)	I(1)	I(1)	I(1)
	PP	I(0)*	I(1)	I(0)*	I(1)	I(0)*	I(1)	I(1)	I(1)	I(0)*	I(1)
	KPSS	I(1)*	I(0)	I(0)	I(0)	I(0)	I(0)	I(1)*	I(1)*	I(1)*	I(1)*

Notes*:* *Implies rejection of $H_{o.}$ I(d) indicates the order of integration at 5% level (with intercept and trend). I(0) points out stationary variables.

For monthly unit value series test results are controversial, in most cases ADF and PP rejected the null of a unit root, while KPSS rejected stationarity. Results for the annual data are not presented here due to a poor performance of stationarity tests in small samples; however test results for both quarterly and yearly series remain ambiguous in most of the cases. All series are I(0) after first differencing.
Due to contradictory results, an assumption that all series are I(1) is made and the estimation is conducted as planned, following the Granger's representation theorem (1983). The Engle-Granger two stages procedure and an unrestricted single-equation ECM, along with the Pesaran's bound testing approach are applied to test for level relationships irrespective of the order of the integration order of variables.

5.4. Cointegration and ECM results

Main results from the first stage Engle-Granger procedure (Equation 4) and the residual cointegration test are reported in Table 6.
Outcomes for most of the countries share the sign and the magnitude of the panel Knetter-type estimation; however, the PTM-coefficient obtained in the ECM for the US is positive and highly significant.

Table 6. Main outcomes from the first stage ECM

		Monthly estimates		Quarterly estimates		Annual estimates	
Canada	*Ln e*	-0.454	**	-0.639	**	-0.709	
		(0.214)		(0.298)		(0.541)	
	T	0.002	***	0.001	*	0.001	
		(0.001)		(0.001)		(0.001)	
	R-sq.	0.260		0.379		0.516	
	DW-stat.	1.720		1.133		0.937	
	EG test-stat.[(a)]	-13.743		-5.729		-2.471	
Sweden	*Ln e*	0.245		0.238		0.245	
		(0.170)		(0.172)		(0.275)	
	T	0.001	***	0.001	***	0.001	***
		(0.000)		(0.000)		(0.000)	
	R-sq.	0.161		0.433		0.591	
	DW-stat.	1.629		1.194		1.378	
	EG test-stat. [(a)]	-8.209		-4.443		-3.219	
Switzerland	*Ln e*	-0.022		0.022		0.659	*
		(0.072)		(0.094)		(0.334)	
	T	0.001	***	0.001	***	0.002	**
		(0.000)		(0.000)		(0.001)	
	D[(b)]	-1.406	***	-0.664	***	-0.167	*
		(0.061)		(0.046)		(0.081)	
	R-sq.	0.805		0.866		0.677	
	DW-stat.	1.496		1.151		1.675	
	EG test-stat. [(a)]	-8.067		-6.417		-3.683	
UK	*Ln e*	-0.345	***	-0.363	***	-0.405	*
		(0.058)		(0.072)		(0.127)	
	T	0.000		0.000		0.000	
		(0.000)		(0.000)		(0.000)	
	R-sq.	0.228		0.414		0.545	
	DW-stat.	1.479		0.733		1.457	
	EG test-stat. [(a)]	-4.804		-4.299		-3.267	
USA	*Ln e*	0.946	***	0.983	***	1.102	**
		(0.148)		(0.223)		(0.413)	
	T	0.005	***	0.006	***	0.006	***
		(0.000)		(0.001)		(0.001)	
	R-sq.	0.522		0.605		0.706	
	DW-stat.	1.022		0.871		1.500	
	EG test-stat. [(a)]	9.255		-3.104		-3.512	

Notes: 1.Standard errors are shown in parentheses. 2.*, **, *** indicate significance at 10, 5 and 1%. Valid only if no cointegration is rejected. [(a)] Critical value from Davidson and Mackinnon (1993) at 5% is -3.78. [(b)] A dummy was added to control for unexplained outlier in the price data (2007M7, 2007Q3, 2007).

The single-equation ECM cointegration test was conducted, as it is believed to perform better than Engle-Granger residual test in finite samples. The null-hypothesis of no cointegration was rejected for monthly data for all countries except US (Table 7). For the annual data H_o could not be rejected for any destination by means of both tests.

Table 7. Single-equation ECM cointegration test results

	Monthly estimates		Quarterly estimates		Annual estimates	
	Ln P(-1) t-stat.	Critical value	Ln P(-1) t-stat.	Critical value	Ln P(-1) t-stat.	Critical value
Canada	-5.457	-3.75	-3.672	-3.78	-2.272	-3.89
Sweden	-8.178	-3.75	-3.899	-3.78	-2.980	-3.89
Switzerland	-10.784	-3.75	-4.683	-3.78	-2.872	-3.89
UK	-4.613	-3.75	-4.667	-3.78	-3.509	-3.89
USA	-2.720	-3.75	-3.562	-3.78	-2.812	-3.89

Notes: Critical value of the ECM-test for the case with constant and trend and one regressor at the 5% level for T=100, 50, 25 from Banerjee et al. (1998). H_o: no cointegration between variables.

Pesaran's bounds testing approach (Table 8) also rejected the null-hypothesis of no long-term relation between the level parameters for monthly data for all destinations except for the US, what implies that there is no cointegration between prices and exchange rates in German exports to US and thus this regression outcome of original Knetter model as well as of the ECM might be spurious.

Table 8. Results of the bound-testing approach

	F-statistic			Critical bond at 5% CI (iv), k=1	
	Monthly estimates	Quarterly estimates	Annual estimates	Lower bound	Upper bound
Canada	9.975	4.606	1.744	4.68	5.15
Sweden	22.336	5.274	3.013	4.68	5.15
Switzerland	39.297	7.369	2.756	4.68	5.15
UK	7.299	7.659	5.999	4.68	5.15
US	2.575	4.327	4.943	4.68	5.15

Notes: Critical values taken from Pesaran et al. (2001:301), CI (iv).

As the speed of error-correction is relatively high, only results for the monthly frequency data from the single-equation and the second stage ECM are reported (Table 9). Long-run PTM coefficients cannot be interpreted straight from the equations' outcomes and are thus neglected here. However, their sign and magnitude are in line with the estimates from the first stage Engle-Granger procedure.

Table 9. Error correction term from the single-equation ECM and second stage ECM (monthly data)

Country	Single equation ECM			Second stage EG		
	Ln P(-1)	R-sq.	DW-stat.	Resid1 (-1)	R-sq.	DW-stat.
Canada	-0.576 (0.106)	0.466	2.017	-0.576 (0.105)	0.466	2.016
Sweden	-0.648 (0.079)	0.450	2.045	-0.647 (0.079)	0.449	2.043
Switzerland	-0.725 (0.067)	0.844	2.046	-0.746 (0.066)	0.842	2.075
UK	-0.377 (0.082)	0.498	2.009	-0.378 (0.077)	0.497	2.008
USA	-0.238 (0.088)	0.356	1.996	-0.236 (0.077)	0.356	1.995

Notes: Standard errors are shown in parentheses.

The speed of adjustment toward the long-run equilibrium varies between -0.725 and -0.377, what implies that most of the system's disequilibrium is being corrected within few months, which supports the use of the high frequency data. ECM outcomes reveal LCPS in German monthly exports to Canada (-0.454) and the UK (-0.345), and share the signs and the magnitude with results obtained from the original fixed-effects model (-0.504 and -0.205 respectively). The long run-coefficient for Switzerland is also negative, but very close to zero (-0.022) and not significant, which can be explained by the stationarity of Swiss unit values. The long-term elasticity of the price with respect to the exchange rate in Sweden has a positive sign (0.245), and is not significant, however the contemporaneous adjustment seems to take place in Swedish exports[10]. As cointegration tests could not reject the null-hypothesis of no cointegration between unit values and exchange rates in exports to the US, results of the fixed-effect model and also of the first stage of Engle-Granger ECM might be spurious.

5.5. Introducing cost shifters into dynamic model

Augmenting the ECM by the time effects θ, estimated in Equation (2) with cpi-adjusted exchange rates, leads to Equation (6) or its two-stage substitute. Both single-equation and two-step ECMs reveal similar results, thus only two-stage estimates are reported (Table 10), as long-term PTM-coefficients and their standard errors can be assessed there directly. Results are in line with previously obtained outcomes, however PTM coefficients are somewhat higher, once marginal costs are introduced into the model. For Canadian exports the long-run PTM-coefficient is -0.626, most of the disequilibrium corrects within the next month, short-term dynamics of the exchange rate are not significant. In British exports, the long-term relationship т-coefficient was found to be -0.458 and the speed of correction of roughly 41 %. For Sweden, the long-term coefficient is positive, but not significant, there is no long-term relation between export prices and exchange rates, however the immediate reaction remains significant and positive. For the case of Switzerland no PTM was detected, neither in the long, nor in the short run.

[10] Short term adjustments are not reported in the paper, however Δe for Sweden is significant and positive and remains once the model is estimated in differences. Results can be provided by request.

Table 10. Main estimates from the ECM with marginal cost proxies (monthly data)

	Canada		Sweden		Switzerland		UK		US	
					1 stage					
θ	1.943 (0.113)	***	0.660 (0.081)	***	0.042 (0.034)		0.210 (0.060)	***	1.955 (0.147)	***
Ln E	-0.626 (0.145)	***	0.165 (0.151)		-0.054 (0.076)		-0.458 (0.065)	***	0.037 (0.132)	
T	-0.002 (0.000)	***	0.000 (0.000)		0.001 (0.000)	***	-0.000 (0.000)	**	0.001 (0.000)	*
R-sq.	0.662		0.338		0.806		0.263		0.722	
DW-stat.	1.890		1.782		1.499		1.516		1.506	
EG-residual test [(a)]										
t-stat.	-14.521		-14.173		-8.067		-5.068		4.939	
					2 stage					
Resid1(-1)	-0.855 (0.056)	***	-0.770 (0.084)	***	-0.744 (0.066)	***	-0.409 (0.079)	***	-0.339 (0.100)	***
d(θ)	2.406 (0.179)	***	0.637 (0.096)	***	0.059 (0.037)		0.166 (0.059)	**	1.548 (0.151)	***
R-sq.	0.754		0.543		0.844		0.515		0.567	
DW-stat.	2.013		2.009		2.072		1.998		2.009	
Single-equation ECM cointegration test [(b)]										
t-stat.	-16.086		-9.105		-10.665		-5.166		-3.538	

Notes: 1.Standard errors are shown in parentheses. 2. *, **, *** indicate significance at 10, 5 and 1%. [(a)]Critical value from Davidson and Mackinnon (1993) at 5% is -3.78. [(b)]Critical value of the ECM-test is 3.98 for the case with constant, trend and two regressors at the 5% level for T=100 from Banerjee et al. (1998)

No PTM was revealed for German exports to the US[11] even for the monthly data, once marginal costs were added, supporting cointegration tests results and suggesting that the outcomes of the fixed-effects model, revealing LCPS in American exports, are a result of a spurious regression of non-stationary non-cointegrated variables. This leads to the idea, that simply neglecting time-series properties is not a good solution, even when applied to low frequency data samples. PTM seem to be a ‘long-term’ phenomenon (besides Sweden), yet the speed of error correction is rather high, and most of the disequilibrium is being corrected within the following months, which also has to be considered when choosing the frequency of the data.

[11] Similar results were achieved in Glauben and Loy (2003), whose destination countries included the UK and the US. PTM was found for the UK, but not for US exports.

In all estimations, decreasing time frequency resulted in a lower significance of estimated parameters. Furthermore as the sample size shrinks, most of stationarity and cointegration tests lose their power. Quarterly and annual estimates, though, not presented here, can be obtained from the author upon request.

6 Summary

Results suggest that German exporters of sugar confectionery differentiate between destination markets and adjust their pricing behavior accordingly, trying to keep price levels relatively stable on some markets, while allowing for a complete pass-through on the other markets.

LCPS was found for those destinations, whose share in German sugar confectionery exports was growing over the sample period. For those destinations (Canada and the UK) results stayed robust throughout different model specifications (Table 11) and revealed the absorption of the Euro appreciation (depreciation) via price adjustment between -0.454 and -0.626 for Canada and between -0.205 and -0.458 for the UK. For Sweden and Switzerland, whose share of German exports remained relatively constant over time, no evidence of PTM was found.

Outcomes reveal that fixed-effect model findings of PTM might be spurious, if time-series properties of the data are not considered. That was the case for the exports to the US, where the Knetter-type model revealed LCPS adjustment, which turned out to be a spurious outcome, as cointegration was rejected. The share of exports to the US in German total confectionery exports was constantly declining over time, and such loss of interest towards the American market could be the explanation of complete pass-through of exchange rate fluctuations to the price[12], as revealed by the ECM with marginal cost proxy.

Data of higher frequency was found to be preferable for PTM studies, once the measurement error due to heterogeneity is minimized. In order to assure the validity of unit values, the prevailing type of competition for every destination market was tested, using the trade flow taxonomy be Gehlhar and Pick (2002) and the most disaggregated product level data (8-digit CN) available on bilateral basis. For most of the destination markets price-driven competition could be proved.

[12] Causality might also go the other way round.

Table 11. Comparison of the long-term price elasticities with respect to exchange rate fluctuations between the models (monthly data)

Destination	Knetter-type panel estimation		ECM (first stage)		ECM with added cost shifters	
Canada	-0.504 (-4.190)	***	-0.454 (-2.120)	**	-0.626 (-4.309)	***
Sweden	0.329 (1.660)	*	0.245 (1.447)		0.165 (1.092)	
Switzerland	0.027 (0.240)		-0.023 (-0.318)		-0.054 (-0.704)	
UK	-0.205 (-2.060)	**	-0.345 (-5.912)	***	-0.458 (-6.979)	***
US	-0.409 (-3.640)	***(*)	0.946 (6.376)	***(*)	0.038 (0.284)	(*)

Notes: 1.Standard errors are shown in parentheses. 2.***, **, * denotes significance at 1, 5 and 10% level and is only valid if H_0 of no cointegration is rejected by means of a single-equation error correction cointegration test (Banerjee et al., 1998). (*) denotes impossibility to reject H_0.

In all estimations, decreasing the data frequency resulted in lower significance levels of the estimated parameters, furthermore as the sample size shrinks, most of stationarity and cointegration tests lose their power. Relatively high speed of the system's correction towards the equilibrium is another reason to stick to the data of higher frequency.

Finally, using marginal costs estimates from a fixed effect model as cost proxies in the ECM improves the quality of the dynamic model and reveals a higher degree of PTM for the UK and Canada. No relationship between export prices to the US and exchange rate fluctuations could be found, once the marginal cost proxy was introduced to the model, proving that the outcomes from the original model and its ECM specification were spurious.

Glauben and Loy (2003) reported quite similar outcomes for the matching countries, with even higher PTM coefficient for the UK (-0.63) and not significant positive coefficient for the US (0.07). Results of the RDE approach, however, led them to a conclusion that competitive conduct seem to prevail on the markets.

This paper' outcomes somewhat support this finding, as PTM was found to be used as a form of the market power realization, where markup is adjusted as a response to Euro appreciation (depreciation) in order to keep prices paid by importers in local currency relatively stable and to ensure German exporters' competitiveness on important markets, protecting thus their market shares.

Despite robustness of the outcomes and their general resemblance to the existing literature, it cannot be ruled out, that changes in unit values are caused by the other factors, besides PTM as a response to exchange rate changes (see e.g. Abbot et al., 1993, or Glauben and Loy, 2003), which stay out of the focus of this study and should be considered in future investigations.

Appendix 1. Share of destination markets in German exports/Share of imports from Germany in destination markets' imports

	Share of the destination country in German sugar confectionery exports					Share of imports from Germany in imports of the destination country				
Period	CAN	SWE	SWI	UK	USA	CAN	SWE	SWI	UK	USA
1991	0.81	1.42	3.81	6.73	18.10	2.36	4.40	15.04	13.60	na
1992	0.61	1.17	3.10	7.01	25.78	3.53	na	15.09	12.95	14.86
1993	0.69	1.89	3.70	9.34	17.83	1.81	7.48	14.60	18.95	11.79
1994	2.18	1.75	3.20	8.21	14.46	6.32	7.28	13.66	17.70	10.02
1995	2.80	2.58	3.68	8.46	13.23	7.77	11.23	14.61	20.51	6.97
1996	1.99	3.22	3.00	7.75	9.21	6.74	11.12	14.00	18.57	5.41
1997	1.72	2.48	3.49	9.01	7.45	4.58	12.76	16.40	17.80	4.27
1998	1.71	3.46	3.25	9.14	8.61	4.22	14.32	16.03	18.68	3.77
1999	1.79	2.90	3.51	9.96	8.39	4.69	9.99	17.25	19.50	3.42
2000	2.32	3.42	3.69	10.66	8.01	5.29	11.14	16.57	17.19	2.90
2001	2.45	3.14	3.34	11.14	6.57	3.77	11.90	16.19	16.27	2.61
2002	1.51	3.08	3.40	11.14	5.57	3.48	8.66	17.67	17.98	2.02
2003	1.72	3.19	3.04	11.07	4.41	3.54	10.65	16.13	15.24	1.57
2004	2.22	3.17	2.83	10.21	6.57	2.82	12.22	10.99	15.90	1.51
2005	1.57	3.69	3.32	10.09	4.98	3.80	14.56	33.73	14.14	2.24
2006	1.76	2.77	3.61	10.36	5.70	3.47	10.58	35.67	13.46	2.35
2007	1.62	2.39	4.11	10.70	5.23	3.84	9.84	38.67	13.98	2.45
2008	1.55	2.05	3.63	9.84	4.50	4.10	8.68	35.64	16.22	2.45
2009	1.88	1.84	4.33	9.67	6.80	4.29	7.31	38.20	16.59	2.70
2010	1.95	1.81	4.03	11.45	6.18	4.00	8.02	40.85	18.61	2.84
Average	1.74	2.57	3.50	9.58	9.38	4.22	10.11	21.85	16.69	4.53

Source: FAOSTAT.

Notes: CAN-Canada, SWE-Sweden, SWI-Switzerland, UK-The United Kingdom, US – The United States.

Appendix 2. Descriptive statistic

	Ln p					Ln E (cpi adjusted)				
	CAN	SWE	SWI	UK	USA	CAN	SWE	SWI	UK	USA
Monthly data										
Mean	0.864	0.730	1.030	0.832	0.862	0.095	0.010	0.050	0.112	0.063
Median	0.806	0.694	1.021	0.831	0.668	0.081	0.001	0.038	0.065	0.033
Maximum	2.398	1.731	1.288	1.107	2.230	0.464	0.208	0.293	0.409	0.431
Minimum	0.091	0.431	-0.284	0.481	-0.479	-0.258	-0.149	-0.364	-0.082	-0.285
Std. Dev.	0.366	0.187	0.136	0.115	0.450	0.172	0.071	0.128	0.143	0.172
Obs.	252	252	252	252	252	252	252	252	252	252
Quarterly data										
Mean	0.832	0.712	1.027	0.824	0.856	0.095	0.010	0.050	0.112	0.063
Median	0.803	0.692	1.028	0.843	0.665	0.078	0.003	0.039	0.065	0.037
Maximum	1.934	1.062	1.214	0.991	1.926	0.441	0.190	0.285	0.399	0.398
Minimum	0.310	0.505	0.465	0.620	0.485	-0.227	-0.140	-0.325	-0.063	-0.259
Std. Dev.	0.312	0.129	0.118	0.093	0.422	0.171	0.070	0.129	0.143	0.171
Obs.	84	84	84	84	84	84	84	84	84	84
Annual data										
Mean	0.848	0.710	1.023	0.822	0.868	0.095	0.010	0.050	0.112	0.063
Median	0.798	0.701	1.007	0.840	0.661	0.087	0.000	0.042	0.068	0.032
Maximum	1.409	0.921	1.173	0.950	1.846	0.417	0.143	0.267	0.386	0.341
Minimum	0.522	0.569	0.808	0.699	0.530	-0.206	-0.133	-0.273	-0.045	-0.232
Std. Dev.	0.282	0.109	0.098	0.087	0.417	0.170	0.067	0.130	0.143	0.170
Obs.	21	21	21	21	21	21	21	21	21	21

Source: Own calculations based on the data from Eurostat, IFS, OECD and Bundesbank.
Notes: CAN-Canada, SWE-Sweden, SWI-Switzerland, UK-The United Kingdom, US – The United States.

References

Adolfson M. Export price responses to exogenous exchange rate movements. Economic Letters 2001, 71, 91-6.

Aiginger K. The use of unit values to discriminate between price and quality competition. Cambridge Journal of Economics 1997, 21, 571-92.

Aw BY, Roberts MJ. Measuring quality change in quota-constrained import markets. The case of U.S. footwear. Journal of International Economics 1986, 21, 45-60.

Baffes J. Explaining stationary variables with non-stationary regressors. Applied Economic Letters 1997, 4, 69-75.

Balaguer J, Orts V, Uriel E. Pricing to market behavior in öligopolistic competition: A microeconometric approach. Instituto Valenciano de Investigaciones Economicas, Working Paper WP-EC 2003-06.

Banerjee A, Dolado JJ, Galbraith JW, Hendry DF. Cointegration, error correction, and the econometric analysis of non-stationary data. Oxford University Press, Oxford, 1993.

Banerjee AJ, Dolando JJ, Mestre R. Error-correction mechanism tests for cointegration in a single equation framework. Journal of Time-Series Analysis 1998, 19, 267-85.

Belke A, Polleit T. Monetary policy and dividend growth in Germany: Long-run structural modeling versus bounds testing approach. Applied Economics 2006, 38, 1409-23.

Bowe M, Saltvedt TM. Currency invoicing practices, exchange rate volatility and pricing-to-market: Evidence from product level data. International Business Review 2004, 13, 281-308.

Brown J. Price discrimination and pricing to market behavior of Canadian canola exporters. American Journal of Agricultural Economics 2001, 83, 1343-9.

Carew R. Pricing to market behavior: Evidence from selected Canadian and U.S. agri-food exports. Journal of Agricultural and Resource Economics 2000, 25, 578-95.

Carew R, Florkowski WJ. Pricing to market behavior by Canadian and U.S. agri-food exporters: Evidence from wheat, pulse and apples. Canadian Journal of Agricultural Economics 2003, 51, 139–59.

Davidson R, MacKinnon JG. Estimation and inference in econometrics. Oxford University Press, Oxford, 1993.

Dullek U, Foster N, Stehrer R, Woerz J. Dimensions of quality upgrading. Economics of Transition 2005, 13, 51-76.

Dunn RM, Jr. Flexible exchange rates and oligopoly pricing: A study of Canadian markets. Journal of Political Economy 1970, 78, 140-51.

Enders W. Applied econometric time series: User's guide. John Wiley & Sons Inc, New York, US, 1995.

Engle RF, Granger CWG. Co-integration and error correction: Representation, estimation, and testing. Econometrica 1987, 55, 251–76.

Falk M, Falk R. Pricing to market of German exporters: Evidence from panel data. Empirica 2000, 27, 21-46.

FAOSTAT: Detailed World Agricultural Trade Flows, 2012. (Accessed November 02, 2012, at http://faostat.fao.org/DesktopModules/Faostat/WATFDetailed2/watf.aspx?PageID=536)

Feenstra RC, Gagnon JE, Knetter MM. Market share and exchange rate pass-through in world automobile trade. Journal of International Economics 1996, 40, 187-207.

Fontagné L, Freudenberg M, Gaulier G. A systematic decomposition of World Trade into horizontal and vertical IIT. Review of World Economics 2006, 142, 459-75.

Froot KA, Klemperer PD. Exchange rate pass-through when market share matters. The American Economic Review 1989, 79, 637-54.

Gagnon JE, Knetter MM. Markup adjustment and exchange rate fluctuations: Evidence from panel data on automobile exports. Journal of International Money and Finance 1995, 14, 289-310.

Gehlhar MJ, Pick DH. Food trade balances and unit values: What can they reveal about price competition? Agribusiness 2002, 18, 61-79.

Gil-Pareja S. Export price discrimination in Europe and exchange rates. Review of International Economics 2002, 10, 299–312.

Glauben T, Loy JP. Pricing-to-market versus residual demand elasticity analysis of imperfect competition in food exports: Evidence from Germany. Journal of Agricultural and Food Industrial Organization 2003, 1, article 3.

Goldberg PK, Knetter MM. Goods prices and exchange rates: What have we learned? Journal of Economic Literature 1997, 35, 1243-72.

Griffith G, Mullen J. Pricing-to-market in NSW rice export markets. The Australian Journal of Agricultural and Resource Economics 2001, 45, 323–34.

Hakkio CS, Rush M. Cointegration: How short is long run? Journal of International Money and Finance 1991, 10, 571-81.

Harris R. Using cointegration analysis in econometric modeling. Prentice Hall/Harvester Wheatsheaf, London, UK, 1995.

Hatemi-J A, Irandoust M. Is pricing to market behavior a long-run phenomenon? A non-stationary panel analysis. Empirica 2004, 31, 55-67.

Haug AA. Temporal aggregation and the power of cointegration tests: A Monte Carlo study. Oxford Bulletin of Economics and Statistics 2002, 64, 399-412.

Jansen WJ. Estimating saving-investment correlations: Evidence for OECD countries based on an error correction model. Journal of International Money and Finance 1996, 15, 749-81.

Kasa K. Adjustment costs and pricing-to-market: Theory and evidence. Journal of International Economics 1992, 32, 1-30.

Kinoshita S. The effect of product classifications on the formulation of export unit value indices: A comparison of export unit value indices based on SITC and HS. Institute of Developing Economies, IDE 2009, Discussion Paper 213.

Knetter MM. Price discrimination by U.S. and German exporters. The American Economic Review 1989, 79, 198-210.

Knetter MM. Exchange rates and corporate pricing strategies. NBER 1992. Working Paper No. 4151.

Knetter MM. International comparisons of pricing-to-market behavior. The American Economic Review 1993, 83, 473-86.

Knetter MM. Pricing to market in response to unobservable and observable shocks. International Economic Journal 1995, 9, 1-25.

Knetter MM. The segmentation of international markets: Evidence from The Economist. NBER 1997. Working Paper No. 5878.

Kremers JJM, Ericsson NR, Dolado JJ. The power of cointegration tests. Oxford Bulletin of Economics and Statistics 1992, 54, 325-48.

Krugman P. Pricing to market when the exchange rate changes. In: Arndt S.W., Richardson J.D., eds. Real-financial linkages among open economies. MIT Press, Cambridge. 1987, 49-70.

Lahiri K, Mamingi N. Testing for cointegration: Power versus frequency of observation – another view. Economics Letters 1995, 49, 121-4.

Larue B. Pricing-to-market: Simple theoretical Insights, formidable econometric challenges. Canadian Journal of Agricultural Economics 2004, 52, 387-98.

Lavoie N, Liu Q. Pricing-to-market: Price discrimination or product differentiation? American Journal of Agricultural Economics 2007, 89, 571-81.

Mah JS. An empirical examination of the disaggregated import demand of Korea – the case of information technology products. Journal of Asian Economics 2000, 11, 237-44.

Mann CL. Prices, profit margins, and exchange rates. Federal Reserve Bulletin 1986, 72, 366-79.

Narayan P, Smyth R. Dead man walking: An empirical reassessment of the deterrent effect of capital punishment using the bounds testing approach to cointegration. Monach University Australia, 2003. Discussion paper No.10/03.

Pattichis CA. Price and income elasticities of disaggregated import demand: Results from UECMs and an application. Applied Economics 1999, 31, 1061-71.

Pesaran MH, Shin Y. An autoregressive distributed lag modelling approach to cointegration analysis. In Strom S., Diamond P., eds. Econometrics and Economic Theory in the 20th Century: The Ragnar Frisch Centennial Symposium. Cambridge University Press, Cambridge, 1999, 371-413.

Pesaran MH, Shin Y, Smith RJ. Bounds testing approaches to the analysis of level relationships. Journal of Applied Econometrics 2001, 16, 289-326.

Pick DH, Carter CA. Pricing to market with transactions denominated in a common currency. American Journal of Agricultural Economics 1994, 76, 55-60.

Pick DH, Park TA. The competitive structure of U.S. agricultural exports. American Journal of Agricultural Economics 1991, 73, 133-41.

Sexton RJ, Lavoie N. Food processing and distribution: An industrial organization approach. In Gardner B., Rausser GC., eds. Handbook of Agricultural Economics. Amsterdam, North-Holland, 2001, 863-932

Shiller RJ, Perron P. Testing for random walk hypothesis: Power versus frequency of observation. Economic Letters 1985, 18, 381-6.

Silvente FR. Price discrimination and market power in export markets: The case of the ceramic tile industry. Journal of Applied Economics 2005, 8, 347-70.

Silver M. The wrongs and rights of unit value indices. Review of Income and Wealth 2010, 56, 206–23.

Swift R. The pass-through of exchange rate changes to the prices of Australian exports of dairy and livestock products. Australian Journal of Agricultural and Resource Economics 2004, 48, 159-185.

Tang TC. Revisiting South Korea's import demand behavior: A cointegration analysis. Asian Economic Journal 2005, 19, 29-50.

Tang TC. Bank lending and inflation in Malaysia: Assessment from unrestricted error-correction models. Asian Economic Journal 2001, 15 275-89.

Tang TC, Nair M. A cointegration analysis of Malaysian import demand function: Reassessment from the bounds test. Applied Economics Letters 2002, 9, 293-6.

UN Comtrade: United Nations Commodity Trade Statistics Database, 2012. (Accessed November 2, 2012, at http://comtrade.un.org/db/.)

Yumkella KK, Unnevehr LJ, Garcia P. Noncompetitive pricing and exchange rate pass-through in selected U.S. and Thai rice markets. Journal of Agricultural and Applied Economics 1994, 26, 406-16.

Zhou S. The power of cointegration tests versus data frequency and time spans. Southern Economic Journal 2001, 67, 906-21.

Ziramba E. Demand for money and expenditure components in South Africa: Assessment from unrestricted error-correction models. South African Journal of Economics 2007, 75, 412-24.

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