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Pricing-to-Market Analysis: The Case of EU Wheat Exports

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Abstract[1]

The EU is a major player in the global wheat market. This paper examines the pricing behavior of EU wheat exporters using a pricing-to-market (PTM) analysis. Wheat is an exemplary product for testing PTM theories as it is widely and frequently traded, and it is largely unbranded. We estimate the relationship between export unit values and exchange rates using quarterly panel data for 11 export destinations for 2000-2013. Results show that there is a meaningful long-run relationship between export unit values and exchange rates, but there is little evidence of differential mark-ups between export markets. Belarus and Iceland are exceptions where the EU exercises local currency price stabilization.

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1. Introduction

Knetter (1993) defines pricing-to-market (PTM) as the "destination-specific adjustment of mark-ups in response to exchange-rate changes". This implies that currency changes are not fully transmitted into export prices with divergent movements in different markets (Krugman, 1986). PTM has received considerable attention as it tests whether exporters can differentiate their prices between destination markets, providing an insight into the degree to which trade is characterised by a lack of convergence in market prices across export markets (Krugman, 1986; Jin, 2008). Much of the empirical literature on PTM has focused on manufactured goods and there has been "little research on agro-food products" (Carew and Florkowski, 2003). However, policy-makers have become increasingly interested in pricing behaviour in agri-food trade (EC, 2013).

Initial investigations of PTM behaviour in both manufacturing and agri-food sectors suffered from a failure adequately to consider time-series properties and disaggregate to ensure product homogeneity and minimise measurement errors (Carew and Florkowski, 2003; Lavoie and Liu, 2007). Moreover, the literature is biased to investigations of North American markets (Pick and Park, 1991; Pick and Carter, 1994; Carew and Florkowski, 2003; Lavoie, 2005; Jin, 2008), with a paucity of studies for the EU (Fedoseeva, 2013). This paper investigates the existence of PTM behaviour in EU wheat exports. More specifically, we estimate the relationship between export units value and exchange rates using quarterly panel data from 11 countries for 2000-2013. Wheat is an exemplary product for testing PTM theories as it is a widely and frequently traded good that is largely unbranded. As far as we are aware, this paper represents the first empirical investigation of PTM behaviour for EU wheat exports.

Wheat is the world's third most important crop after rice and maize when measured by the value of production (FAO, 2014). It is a strategically important commodity for which trade is politically sensitive. Recent instability in trade and prices have raised "…a number of research questions about current and future competition and price-setting behaviour in the world wheat market" (Pall et al., 2013), but there is little evidence on pricing strategies. The absence of research on EU wheat exports is surprising given that it is a major player in the market and that

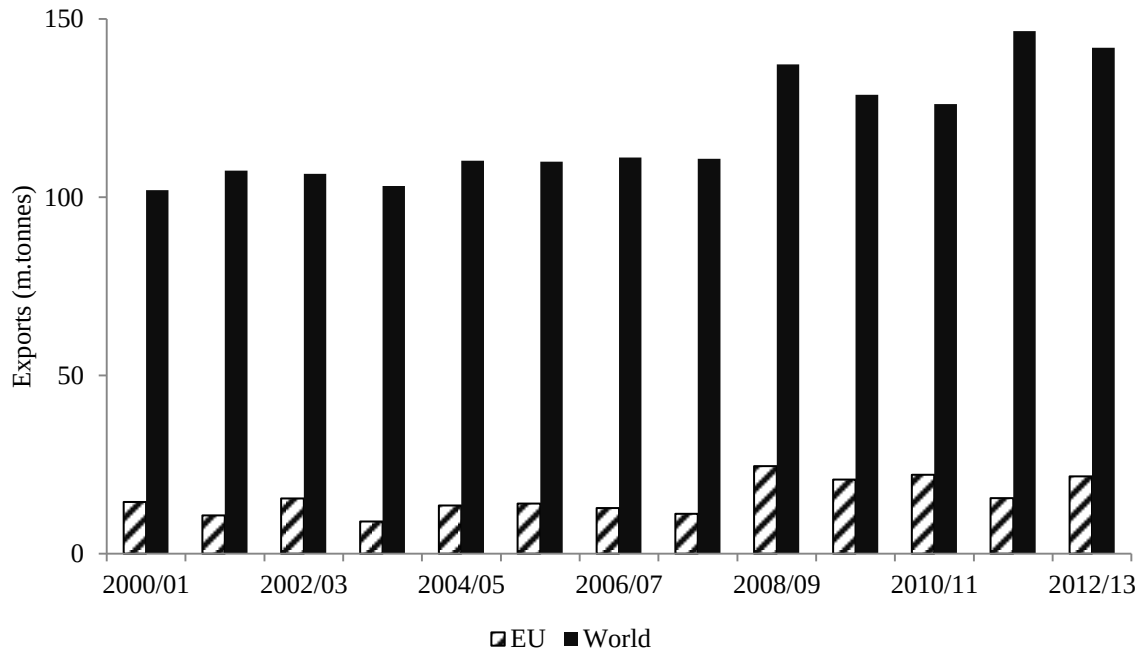
many of the EU's main destination markets (for example, Algeria, Cote D'Ivoire, Egypt, Iran, Morocco, Senegal, Switzerland, Turkey, and Tunisia) are highly dependent on imports for wheat. The paper is structured as follows. Section 2 describes briefly the EU wheat market and in particular its exports. Section 3 discusses the PTM model in the existing literature. Section 4 presents the empirical methodology. Section 5 examines the results and Section 6 concludes.

2. The EU and Wheat Exports

Global wheat production in 2013 amounted to 716m. tonnes, of which 20% was by EU member states (FAO, 2014). In the EU, wheat is the largest grown cereal accounting for almost half of total cereals output and half of cultivated area. The rest is composed of maize and barley (one-third) and other cereals such as rye and oats (Eurostat, 2014). The most important producers in the EU are France and Germany, with Poland and Romania from the new member states also significant producers (Eurostat, 2014). Almost 50% of cereals produced is used for animal feed, followed by human consumption (at around 20%) (Eurostat, 2015).

The EU is also a major player in global cereals markets. Most of EU cereals exports consist of wheat and barley, with 15% of the wheat crop exported every year, while maize and some other feed grains are imported, along with large quantities of oilseeds, oilseed meals, vegetable oils and rice (EC, 2014). In 2013, the EU was the second largest exportor of wheat with the US being the largest. France dominates wheat exports with around two-thirds of total EU exports (EC), and the main destinations are countries in North Africa (Tunisia, Algeria and Egypt), the Middle East (Saudi Arabia), and the Far East (South Korea). The EU has a weaker presence in the important export markets of East Asia, where Australia benefits from lower shipping costs stemming from closer geographical proximity compared to its main rivals, USA, Canada, and the EU. Figure 1 shows world and EU exports for crop years 2001/02-2012/13. World wheat exports range from 102m. tonnes in 2001/02 to 147m. tonnes in 2011/12 and the proportion from the EU varies between 9% in 2003/04 and 18% in each of 2008/9 and 2010/11. The EU's importance in world wheat export markets is clear.

Figure 1: Wheat Exports (m. tonnes)



Note: EU15 up to 2003/04, EU25 from 2004/05, and EU27 from 2006/07.
Source: Personal communication with International Grains Council

3. A Review of the PTM Literature

Krugman (1987) developed the idea of PTM whereby an exporter with power in multiple markets adopts price-discriminating behaviour and can either maintain or increase their export prices when the exchange rate rises. Following Knetter (1989), the profit, π, of an exporter selling to $i = 1,\ldots,N$ export markets is:

$$\pi(p_1,\cdots,p_N) = \sum_{i=1}^{N} p_i\, q_i(ER_i p_i; S_i) = C(\sum_{i=1}^{N} q_i\,(ER_i p_i; S_i), W) \qquad (1)$$

where p_i is the price for country in the exporter's currency, q_i is quantity demanded which is determined by the price in the buyer's currency, ER_i, and a demand shifter S_i, and $C(q,w)$ is the cost function where w denotes input prices. The first-order profit-maximising conditions show that the price to each export market is the product of the common marginal cost, C_q, and a destination-specific mark-up:

$$p_i = C_q\left(\frac{\eta_i}{\eta_i - 1}\right), \qquad i = 1,\ldots,N \qquad (2)$$

where η_i is the absolute value of the elasticity of demand in export market i. Thus, the firm equates the marginal revenue from sales in each export market with the common marginal cost, and the ability to adopt price-discriminating behaviour depends on the elasticity of demand in the export market and marginal cost. A change in the exchange rate affects the price in an export market because it affects marginal cost (through changes in q_i or w) or because it affects the elasticity of demand in the export market. The former affects (spills-over to) other export markets, while the latter (to which PTM refers) is destination-specific, and both determine pass-through.

The empirical counterpart of (2) with $t = 1,\ldots,T$ observations is derived by taking natural logarithms and totally differentiating:

$$\Delta lnp_{it} = \theta_t + \beta_i \Delta lnER_{it} + \varepsilon_{it}, \qquad i = 1,\ldots,N,\ t = 1,\ldots,T \tag{3}$$

where ε_{it} is an error term with the usual properties and ln is the natural logarithm. Knetter (1993) notes that (3) can be rewritten in levels:

$$lnp_{it} = \theta_t + \lambda_i + \beta_i lnER_{it} + \varepsilon_{it} \tag{4}$$

where θ_t are common time-specific effects, λ_i are country-specific effects, and β_i are the PTM-coefficients or the elasticities of the export price with respect to exchange rates. Equation (4) is sometimes preferred to (3) because it also contains information on λ_i.

Knetter (1989) distinguishes between alternative scenarios, depending on the estimated parameters, θ_t, λ_i and β_i, in (4). In the case of perfect competition, export price equal the common marginal cost, and the common time-specific effects measure the common price in each period and are an exact measure of marginal cost. Here, (4) is deterministic and there is no residual variation in prices. Thus, country-specific effects are absent so λ_i=0 and β_i=0, and θ_t measure the evolution of marginal costs over time. In integrated markets, price is equal amongst buyers; and in imperfect competition, the mark-up of price over marginal cost is not necessarily zero. Thus the time-specific effects measure the marginal cost plus an unidentifiable common mark-up which implies that country-specific effects are zero. In (4), the cases of competition and integration are indistinguishable.

Imperfect competition in export markets is characterised by market segmentation and price discrimination. If each export market has a constant elasticity of demand, the price in each is a fixed mark-up over marginal cost, and price variation has two time components: the time-specific effects, θ_t, are an exact index of marginal cost; and country-specific effects, λ_i, measure mark-ups. Since, constant elasticity of demand implies constant mark-ups, then β_i=0 for all i. By contrast, if $\lambda_i \neq 0$ and/or $\beta_i \neq 0$, market segmentation exists with a non-constant elasticity of demand and the time-specific effects, θ_t, are only a noisy measure of cost changes. Here, there is price discrimination across export markets and the price elasticity varies with changes in the exchange rate. For instance, if an importer's currency depreciates relative to an exporter's currency, the price faced by consumers in the export currency rises but if the price elasticity changes, the optimal mark-up over marginal cost also changes. The optimal mark-up by a price discriminating exporter varies across export markets ($\lambda_i \neq 0$) and with changes in bilateral exchange rates so that $\beta_i \neq 0$. If $\beta_i < 0$, local currency price stabilization exists whereby exporters adjust their mark-ups to maintain relatively constant prices in export markets, for example by absorbing part of a change the export unit value; and if $\beta_i > 0$, there is amplification of exchange rate fluctuations by the exporter in terms of their mark-up. Thus, if β_i=-0.5 then a 10% appreciation (depreciation) of the exchange rate implies a 5% fall (rise) in the mark-up of export price over marginal cost, and the exchange rate pass-through is 50%. Typically, $\beta_i < 0$ (Goldberg and Knetter, 1997).

Early empirical studies of PTM focused on manufacturing. For example, Krugman (1986) uses US-German trade data and concludes that PTM occurs but it is limited to transportation equipment and machinery industries. Subsequent work by Knetter (1993) suggests that the existence and extent of PTM varies widely between industries and exporting countries. In explaining this variation, early studies stress the roles of both supply dynamics, namely the costs of rapidly adjusting the marketing and distribution infrastructure required for selling imports, and demand dynamics stemming from firms' desires to protect and enhance reputation (Krugman, 1986). The latter would seem to be of greater relevance for branded, manufactured goods than for a commodity like wheat.

Initial studies of PTM in agricultural markets by Pick and Park (1991) and Pick and Carter (1994) examine North American wheat exports to eight destination markets using quarterly panel

data for 1978-1988. Pick and Park (1991) estimate (4) and find strong evidence of price discrimination across destination markets for US wheat exports. Pick and Carter (1994) examine both US and Canadian wheat exports and (4) is extended to include the Canadian dollar/US dollar exchange rate. Results show evidence of PTM in Canadian wheat exports and that the exchange rate plays a significant role in the export pricing decisions of both Canadian and US exporters. Carew and Florkowski (2003) also consider US and Canadian wheat exports using (4) likewise extending it to include the Canadian dollar / US dollar exchange rate. Panel annual data for 1980-1998 is used and the time-series properties of the data are additionally examined. Results show price discrimination across export markets, and there is evidence that in setting prices Canadian exporters tend to magnify the effects of exchange rate changes in export markets. Jin (2008) tests for PTM to assess whether the Canadian Wheat Board (CWB) is able to price discriminate in exports. Using (4) and annual panel data for 1988-2003, results show that the CWB exercises PTM behaviour only in some export markets, but a caveat concerns the limitations of annual data where higher frequency data may provide better goodness of fit and information of PTM behaviour. Recently, Pall et al. (2013) consider PTM behaviour of Russian wheat exporters using (4) and quarterly panel data for 2002-2010. Results show evidence of PTM for five countries (Algeria, Azerbaijan, Cyprus, India and Mongolia) and this is attributed to Russia's large share of wheat imports and/or lack of major competitors for these export markets. Like Jin (2008), Lavoie (2005) also tests for PTM to assess whether the CWB is able to price discriminate in exports. Monthly CWB contracts from 1982-1994 are used for exports to four export markets, Japan, UK, Rest-of-the-World west coast, and Rest-of-the-World east coast. Lavoie's model is not based on (4) but instead the difference in price between two export markets is a function of the difference in the values of the instruments of price discrimination namely freight rates, import duties, export subsidies, and exchange rates, quality differences between Canadian and US wheat, and difference in the Export Enhancement Programme's bonus. Results indicate that the CWB does discriminate by charging different prices to different countries, even after accounting for differences in product quality. These studies of PTM in agricultural markets are subject to two criticisms. First, Carew and Florkowski (2003) and Jin (2008) use annual data, and Fedoseeva (2013) argues that such data are insufficiently fine-tuned to examine PTM behaviour. Second, only Carew and Florkowski (2003) and Pall et al. (2013) examine the time

series properties of the panel data; perhaps surprisingly, both find that export unit values and the nominal exchange rate are stationary, and the estimation of (4) in levels is appropriate. This paper addresses both criticisms.

4. Empirical Method

Equation (4) is a two-way fixed effects (or least squares dummy variable) model with both country and time effects and most studies of PTM estimate it by ordinary least squares (OLS) (Greene, 2012, pp.399-408).[2] However, many economic series including those constructed in a panel are trended and if the trend can be removed by first-differencing, the series is integrated of order one, I(1). OLS regressions between such series are in general spurious but the exception is where two (or more) non-stationary series move together and their linear combination is stationary. Here, the series are cointegrated and a meaningful long-run equilibrium exists (Granger, 1988). To test for non-stationarity in a panel, a number of tests have been developed (Harris and Sollis, 2003, pp.191-200) of which two are applied here. The first is the test of Im et al. (2003), which is denoted as the IPS-test, where the null is that each individual (country) series in the panel contains a unit root and the alternative is that at least one of the individual series in the panel is stationary. To implement this test, an augmented Dickey-Fuller (ADF)-equation is estimated for each country (Dickey and Fuller, 1981), and heterogeneous dynamics are allowed. The IPS-statistic, which is essentially the average of the individual ADF-statistics, is adjusted to be asymptotically standard normal. The test is one-sided and the critical value at 5% is -1.645. A test-statistic in the left tail of the distribution provides evidence for rejecting the null; and non-rejection of the null implies that the individual series is I(1). Second, we use the test of Hadri (2000) where the null is that each individual series in the panel is stationarity against the alternative of a unit root in the panel. This is a generalisation of the KPSS test (Kwiatkowski et al., 1992) for a single time series. The Hadri-statistic is asymptotic standard normal and rejection of the null implies that the individual series is I(1). The test is one-sided and the critical value at 5% is 1.645. A test-statistic in the right tail of the distribution provides evidence for rejecting the

[2] By contrast to the usual two-way fixed effects model, β_i in (4) is allowed to vary between i.

null; and non-rejection of the null implies that the individual series is I(0). The existence of cointegration between $ln\ p_{it}$ and $ln\ ER_{it}$ requires that both series are I(1).

A number of panel tests for non-cointegration have been developed (Harris and Sollis, 2003, pp.200-206). We use the ADF-type t-statistic of Pedroni (1999, 2004) to test for non-cointegration in (4). The group-means estimator is used to estimate (4) so that the parameters vary across countries thereby permitting heterogeneity. The null of non-cointegration tests the non-stationarity of the residuals, which in (4) are $\hat{\varepsilon}_{it} = \rho_i \hat{\varepsilon}_{it-1} + u_{it}$, and essentially the test follows the IPS-test for unit roots. For this group ADF-statistic, the null is that $ln\ p_{it}$ and $ln\ ER_{it}$ are not cointegrated for each country in the panel (H_0: ρ_i=1) and the alternative is that cointegration exists (H_1: ρ_i<1) for a significant proportion of countries. We also use an alternative test for non-cointegration which is analogous to Phillips and Perron's (PP) t-statistic (Phillips and Perron, 1988; Pedroni, 1999, 2004). Both group ADF- and group PP-tests are one-sided and the critical value at 5% is -1.645. Test statistics in the left tail of the distribution provides evidence for rejecting the null, while non-rejection implies that the two series are not cointegrated.

If panel cointegration exists between ln pit and *ln ERit*, their long-run relationship can be estimated. Pedroni (2002) considers two panel estimators: (non-parametric) fully modified ordinary least squares (FMOLS) which deals with serial correlation using a heteroscedasticity and autocorrelation consistent estimator of the long-run covariance matrix, and (parametric) dynamic ordinary least squares (DOLS) which estimates the lags explicitly. Both correct for OLS bias induced by endogeneity, and both can be used to provide within- or between-group (group mean) estimates. Pedroni (2002) prefers the group mean estimator because it has relatively minor size distortions in small samples, and β_i need not be the same for all N countries. Moreover, estimates of the long-run cointegrating panel parameters can be interpreted as the mean values of all country cointegrating vectors. Accordingly, we use the group mean estimator. The evidence for preferring FMOLS or DOLS is not so clear; both provide asymptotically unbiased estimators which tend to be similar. The panel long-run elasticity of the export price with respect to the exchange rate is given by the of mean values of β_i.

An alternative to Pedroni's (2002, 2004) method is the pooled mean estimator of Pesaran et al. (1999). This is a panel extension of Engle and Granger (1987) whereby the long-run relationship in (4) is embedded in an error-correction model (ECM), and it facilitates the estimation of the long-run PTM-coefficient and the speed of adjustment towards long-run equilibrium following a shock to the system. Instead of estimating the long-run cointegrating relationship in (4) directly, the ECM with heterogeneous short-run dynamics is estimated. With one lag on the dynamics, this is:

$$\Delta lnp_{it} = \phi_i(lnp_{it} - \theta_t - \lambda_i - \beta lnER_{it}) - \delta_i \Delta lnER_{it} + \omega_{it} \qquad (5)$$

where ω_{it} is an error term which is allowed to vary between countries. In this pooled mean estimator, the PTM-coefficient is β which is common to all countries whereas other parameters are allowed to vary as well as the variances. It is possible to estimate (5) directly using weighted non-linear least squares but that requires a large parameter set to allow for all heterogeneous coefficients. Instead, Pesaran et al. (1999) use an iterative procedure which solves the first-order conditions for the two blocks of coefficients (heterogeneous versus homogeneous) given the other, and the log-likelihood increases at each step. The panel error-correcting speed of adjustment is the mean of ϕ_i which is expected to be negative.

5. Data and Results

The data consist of export unit values of wheat (€/tonne) and exchange rates expressed as units of the importer's currency per unit of the exporter's currency, for example, Albanian Lek/€. Monthly data from 11 of the main EU wheat export destinations for 2000(1)-2013(11) contain missing observations on export unit values. Both series are therefore converted into quarterly data using simple averages. To maximise the sample, a small number of missing observations remain and they are interpolated from nearby observations by simple averaging. The final balanced panel dataset consist of 56 observations for 2000(1)-2013(4) for Norway, Switzerland, Albania, Belarus, the Democratic Republic of Congo, Mauritania, Tunisia, Egypt, Morocco, Algeria and Iceland, and descriptive statistics are shown in the Appendix.

The fixed effects model in (4) is estimated with White's robust estimator to correct for unknown heteroscedasticity (Greene, 2012, pp.390-391, 425-427) and the results are shown in Table 1.

Estimates of the common time effects, θ_t for t = 1,…,56 are not reported. Testing the null is that θ_t are not significantly different yields χ^2_{55}=2134.71 (p-value: 0.00), the null is rejected and the common time-specific effects are not equal. The null that each country-specific effect, λ_i, is equal to the average of all intercepts is that there is a constant mark-up. The test yields χ^2_{10}=122.11 (p-value: 0.00) and the null is rejected.

Testing the null hypothesis that the β_i-effects are not significantly different yields χ^2_{10}=74.70 (p-value: 0.00), and the null is rejected. Thus, the PTM-coefficients are not equal between countries and there is evidence of PTM-behaviour. The maximum PTM-coefficient is 1.57 for Albania; the minimum is -0.33 for Algeria; and the average is 0.20. Seven countries have a positive PTM-coefficient which implies the amplification of exchange rate fluctuations by the EU. Of these however, only those for Switzerland and Belarus are significant. There are four countries - Congo, Egypt, Algeria and Iceland - with negative PTM-coefficients which indicates local currency price stabilization where the EU adjusts mark-ups to maintain relatively constant prices by absorbing part of a change the export unit value. Of these, only those for Algeria and Iceland are significant. Overall, there is evidence of both local currency price stabilization and of amplification of exchange rate fluctuations although there is no evidence of PTM-behaviour in the majority of export markets.

The results from the fixed effects model in (4) may be spurious and we test for unit roots. There is evidence of trends in some of the individual series, and they are included in both IPS and Hadri unit roots test equations. In the IPS-test, the null is that each individual series in the panel contains a unit root and the alternative that at least one of the individual series in the panel is stationary. Heterogeneous dynamics are allowed where the number of lags in each country equation, with a maximum of four, is determined from a general-to-specific method whereby the longest insignificant lag is dropped sequentially until the last lag is significant at the 10% level. IPS-statistics for $ln\ p_{it}$ and ER_{it} are -7.57 (p-value: 0.00) and -0.67 (0.25); the null for $ln\ p_{it}$ is rejected which implies stationary, while that for $ln\ ER_{it}$ is not rejected which implies non-stationary, that is I(1). In the Hadri-test, the null is that each individual series in the panel is stationarity against the alternative of a unit root. With four lags in the Bartlett window and allowing for heterogeneous serial correlation, Hadri-statistics for ln pit and $ln\ ER_{it}$ are 22.52 (p-

value: 0.00) and 78.92 (0.00) and both nulls are rejected which implies that both series are non-stationary I(1) variables. It is not uncommon for panel unit root tests to draw different conclusions. However, they cast doubt on whether the fixed effect results are meaningful. On balance, it seems sensible to proceed on the basis that both series are non-stationary, and we examine the existence of a cointegrating relationship between them. We now test the null of non-cointegration following Pedroni (1999, 2004). Common time effects are subtracted out and a trend is included in test equations. The group ADF-statistic is -15.11 (p-value: 0.00) while the group PP-statistic is -16.08 (0.00) and both imply rejection of the null of non-cointegration. Thus there is a meaningful long-run relationship exists between export unit values and exchange rates.

Following Pedroni (2002, 2004), the results of estimating (4) using FMOLS and DOLS both with four lags in the Bartlett kernel are shown in Table 1. Estimates between the two are similar as expected. Two hypothesis tests for heterogeneity are undertaken. First, we test the null that each country's intercept, λ_i, is equal to the average of all intercepts. For the FMOLS estimates, χ^2_{10}=71.86 (p-value: 0.00) and the null is rejected; and for the DOLS estimates, χ^2_{10}=7.99 (p-value: 0.63) and the null is not rejected. There is some evidence therefore that optimal mark-ups vary across export markets. Belarus from the FMOLS estimates has the highest and only significant mark-up, while Belarus and Iceland are significant in the DOLS estimates; otherwise, there is little evidence of differential mark-ups between export markets.

Second, we test the null that each country's PTM-coefficient, β_i, is equal to the average of all PTM-coefficients. For the FMOLS estimates, χ^2_{10}=20.65 (p-value: 0.02); and for the DOLS estimates, χ^2_{10}=24.26 (p-value: 0.01). We conclude therefore that PTM-coefficients are not everywhere equal. The FMOLS PTM-coefficients, β_i, range from -0.673 for Belarus to 0.165 for Mauritania while the DOLS PTM-coefficients range from -0.855 for Tunisia to 0.461 for Norway. Of the 11 countries, FMOLS results show that Norway, Switzerland, Mauritania and Egypt have positive but insignificant long-run PTM-coefficients. In addition, DOLS results indicate that that for Albania is also positive but insignificant. FMOLS results show that Albania, Belarus, Congo, Tunisia, Morocco, Algeria and Iceland have negative long-run PTM-coefficients and all but that for Belarus are insignificant. Negative PTM-coefficients from DOLS estimates are significant only for Belarus and Iceland.

The estimates of the PTM-panel parameters from FMOLS and DOLS can be interpreted as the mean values of all country cointegrating vectors, and are around -0.13; the former is significant while the latter approaches significance. Common to both estimators is the negative and significant PTM-coefficient for Belarus where there is local currency price stabilization and a 1% increase in the exchange rate leads to around a 0.68% decrease in export unit value in both cases. Iceland is the only other case of a negative and significant PTM-coefficient and the DOLS estimate implies that a 1% increase in the exchange rate leads to around a 0.38% decrease in export unit value. None of the positive PTM-coefficients from either estimator is significant. Overall, there is much evidence that the EU does not practise PTM-behaviour in its wheat export markets, although for Belarus and Iceland there is local currency price stabilization whereby the EU adjust mark-ups to maintain export prices relatively constant.

Table 1 about here

Since the equations for Belarus and Iceland yield significant DOLS PTM-coefficients, we construct a sub-sample panel dataset for these two countries. Testing the null that each PTM-coefficient is equal to their average yields χ_1^2=0.09 (p-value: 0.76) and the null is not rejected. Accordingly, we use the panel method of Pesaran *et al*. (1999) to estimate the (common) long-run PTM-coefficient, β, and the speed of adjustment, which is the mean of ϕ_i in the ECM in (5) which also includes a trend. The estimate $\hat{\beta}$=-0.827 (p-value: 0.09) which is significant at the 10% level, and its value is higher than the individual FMOLS and DOLS estimates in Table 1. The estimate of the speed of adjustment is $\hat{\phi}$=-0.325 (p-value: 0.00) and adjustment is relatively quick. In particular, 33% of the adjustment to long-run equilibrium takes place in the first year following a shock to the system, almost 70% of full adjustment takes place by the third year, and full adjustment takes around 10 years.

6. Concluding Remarks

Since the seminal studies of Krugman (1986) and Knetter (1989), a substantial literature has emerged on PTM. Initial research has been criticized for being based on insufficient disaggregation of product categories (Lavoie and Liu, 2007) and for failing adequately to examine the time-series properties of data. More recent studies investigate PTM in agricultural

trade, and examine wheat exports from Canada (Carew and Florkowski, 2003; Jin, 2008; Lavoie, 2005; Pick and Carter, 1994), the US (Pick and Park, 1991; Pick and Carter, 1994; Carew and Florkowski, 2003) and Russia (Pall et al., 2013). However, there is a lack of PTM analysis for EU wheat exports, which is surprising given its importance in this market. Accordingly, this paper investigates the existence of PTM in EU wheat exports using data for 11 export destination markets for 2000-2013.

Following Knetter (1989), PTM models are conventionally estimated by a fixed effects estimator. Results show that neither of the common time- or country-specific effects are equal. The former implies that marginal costs differ while the latter imply differential mark-ups. Similarly, the PTM-coefficients are not equal between countries and there is evidence of PTM-behaviour. However, significant amplification of exchange rate fluctuations by the EU is only evident for Switzerland and Belarus, and significant local currency price stabilization is only evident for Algeria and Iceland. These results may be spurious if export unit values or exchange rates are non-stationary. We therefore test for panel unit roots: the test of Im et al. (2003) indicates that export unit values are stationary while exchange rates are non-stationary while that of Hadri (2000) indicates that both series are non-stationary. Intuition suggests conclusions from the latter may be more appropriate and we therefore examine the existence of cointegration, and following Pedroni (1999, 2004) we find that a meaningful long-run relationship exists between export unit values and exchange rates.

We estimate the PTM-model using both fully modified ordinary least squares (FMOLS) and dynamic ordinary least squares (DOLS) following Pedroni (2002, 2004), and they produce similar results as expected. There is little evidence of differential mark-ups between export markets although those from Belarus (from FMOLS estimates) and Belarus and Iceland (from DOLS estimates) are significant. Tests also show that PTM-coefficients are not everywhere equal and they range from -0.855 to 0.461. However, none of the positive PTM-coefficients are significant. Of the negative PTM-coefficients, only that for Belarus (from FMOLS estimates) and Belarus and Iceland (from DOLS estimates) are significant. Thus, wheat export markets are integrated although Belarus and Iceland are exceptions where local currency price stabilization exists whereby EU exporters adjust their mark-ups to maintain constant prices by absorbing part of a change the export unit value. Accordingly, a panel sub-sample is formed for Belarus and

Iceland. The estimated (common) long-run PTM-coefficient is -0.827 which implies that a 1% increase in the exchange rate leads to around a 0.83% decrease in export unit value. The speed of adjustment has not been considered in previous PTM analyses of wheat markets, and the estimated value here implies that a third of the adjustment to long-run equilibrium takes place in the first year following a shock to the system, with full adjustment taking around 10 years. This is somewhat slower than in the analysis of Germany's exports of sugar confectionery of Fedoseeva (2013) where most of the adjustment occurs within a few months.

Our finding of modest evidence that EU wheat exporters adjust their prices to offset local exchange rate movements, or what Knetter (1989) refers to as a local currency price stabilisation strategy, is consistent with most previous studies of PTM in wheat markets (Pick and Carter, 1994; Jin, 2008; Pall et al., 2013). Moreover, the results show no clear distinction in export behaviour between developing and developed nations, nor any relationship with geographical proximity. The disparities between destination markets mirror the findings of Carew and Florkowski (2003) and Jin (2008) for Canadian wheat exports, who attribute these to variations in the degree of competition in different export markets, that is, exporters engage in mixed practices by differentiating between destination markets which reflects greater market power in some markets.

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Appendix: Descriptive Statistics

	Export Unit Value (€/tonne)				Exchange Rate (Importer's currency per €).			
	Mean	Std. Error	Minimum	Maximum	Mean	Std. Error	Minimum	Maximum
Norway	172.22	57.64	104.54	292.82	8.01	0.37	7.32	8.95
Switzerland	206.87	56.40	118.70	313.85	1.47	0.14	1.17	1.66
Albania	185.18	81.03	92.42	481.00	131.55	6.88	121.84	140.99
Belarus	298.32	254.95	93.17	1528.57	4052.00	3327.06	364.12	12450.00
Congo	159.44	48.17	94.31	270.09	724.79	418.58	6.36	1326.77
Mauritania	180.96	117.06	78.62	931.68	328.04	58.48	212.59	411.09
Tunisia	195.18	86.62	101.64	468.03	1.68	0.28	1.24	2.23
Egypt	164.96	54.00	94.96	289.09	6.87	1.78	3.23	9.43
Morocco	174.80	68.40	99.48	432.98	10.91	0.47	9.56	11.47
Algeria	176.19	55.41	107.72	331.04	90.41	12.07	67.26	109.70
Iceland	170.46	62.31	104.25	413.10	117.66	43.60	70.54	269.20

Sources: 1. Export unit values – Eurostat (2014).

2. Exchange rates – ECB (2014).

Table 1: Panel Results

	Fixed Effects		FMOLS			DOLS		
	Const. (λ_i)	PTM-coeff. (β_i)	Const. (λ_i)	PTM-coeff. (β_i)	Trend (γ_i)	Const. (λ_i)	PTM-coeff. (β_i)	Trend (γ_i)
Norway	4.558	0.344	0.062	0.078	0.001	0.597	0.461	0.005
	(6.35)	(1.00)	(0.32)	(0.50)	(0.21)	(1.73)	(1.80)	(1.40)
Switzerland	5.326	0.384	0.425	0.048	-0.004	0.713	0.131	-0.004
	(61.70)	(2.21)	(0.67)	(0.22)	(0.92)	(0.83)	(0.44)	(0.77)
Albania	-2.334	1.568	0.182	-0.110	-0.003	-0.325	0.289	0.000
	(0.57)	(1.86)	(0.34)	(0.32)	(0.51)	(0.40)	(0.49)	(0.03)
Belarus	3.526	0.261	2.153	-0.673	0.036	2.267	-0.694	0.035
	(8.70)	(4.86)	(3.54)	(3.87)	(6.42)	(3.35)	(3.62)	(6.22)
Congo	5.314	-0.017	-0.050	-0.004	-0.002	0.046	-0.056	-0.001
	(48.03)	(1.25)	(1.27)	(0.19)	(1.65)	(0.40)	(0.99)	(0.45)
Mauritania	4.912	0.062	-0.446	0.165	0.002	-0.526	0.198	0.003
	(5.84)	(0.43)	(0.54)	(0.42)	(0.90)	(0.38)	(0.30)	(0.97)
Tunisia	5.236	0.252	-1.736	-0.539	0.001	-2.807	-0.855	0.002
	(50.29)	(1.66)	(1.50)	(1.46)	(0.30)	(1.19)	(1.17)	(0.77)
Egypt	5.366	-0.071	0.022	0.016	-0.003	0.092	0.052	-0.003
	(34.05)	(0.99)	(0.08)	(0.12)	(2.87)	(0.29)	(0.31)	(2.70)
Morocco	5.176	0.042	-0.179	-0.116	-0.001	-0.031	-0.012	-0.001
	(5.39)	(0.10)	(0.62)	(0.42)	(0.29)	(0.06)	(0.03)	(0.29)
Algeria	6.779	-0.328	0.496	-0.458	-0.006	0.384	-0.342	-0.005
	(12.35)	(2.75)	(2.07)	(1.80)	(3.12)	(0.74)	(0.58)	(1.52)
Iceland	6.508	-0.266	0.130	-0.080	-0.004	0.256	-0.382	0.000
	(16.32)	(3.36)	(0.88)	(0.52)	(2.01)	(2.51)	(3.33)	(0.13)
Panel	-	0.049	0.096	-0.152	0.001	0.060	-0.110	0.003
		(2.12)	(1.20)	(2.20)	(1.07)	(2.35)	(1.92)	(1.12)

Source: authors' estimation;

Notes: 1. Fixed effects equation: $ln\, p_{it} = \theta_t + \lambda_i + \beta_i \ln ER_{it} + \varepsilon_{it}$.

2. FMOLS/DOLS equation: $ln\, p_{it} = \theta_t + \lambda_i + \beta_i \ln ER_{it} + \gamma_i Trend_{it} + \varepsilon_{it}$.

3. t-statistics in parentheses.