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Pricing to market analysis: the case of EU wheat exports

Philip Dawson, Matthew Gorton, Carmen Hubbard and Lionel Hubbard¹

Abstract

The EU is a major player in the global wheat market. This paper examines the pricing behaviour 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 EU 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 EU export markets. Belarus and Iceland are exceptions where exporters from the EU appear to exercise local currency price stabilisation.

Key Words: pricing-to-market, PTM analysis, wheat exports, EU, trade, panel data, cointegration, local currency price stabilization

JEL Classifications: C23, F14, F15, Q17

¹ Philip Dawson, Carmen Hubbard and Lionel Hubbard are in the School of Agriculture, Food and Rural Development, Newcastle University, UK. Matthew Gorton is in the Business School, Newcastle University, UK. Carmen Hubbard is the corresponding author (carmen.hubbard@ncl.ac.uk). We would like to thank the two anonymous referees for their helpful comments that undoubtedly led to the improvement of this paper. This paper was generated as part of the COMPETE Project, Grant Agreement No. 312029 funded by the European Community under the 7th Framework Programme. The authors gratefully acknowledge this financial support.

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 (i.e. the import price fails to change proportionately to the exchange currency change) (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). It thus provides insight into whether price discrimination occurs in international trade. While attracting considerable academic interest, much of the initial empirical literature on PTM focussed on manufactured goods (e.g. Knetter, 1989; Knetter, 1993; Falk and Falk, 2000) 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 (OECD, 2013).

Initial investigations of PTM behaviour in both manufacturing and agri-food sectors were criticised for insufficient disaggregation of product categories (to ensure homogeneity and minimise measurement errors) and for failing adequately to examine the time-series properties of data (Carew and Florkowski, 2003; Lavoie and Liu, 2007). Furthermore, the literature is biased to investigations of North American markets (Pick and Park, 1991; Carew and Florkowski, 2003; Lavoie, 2005; Jin, 2008; Jin and MilijkoVIC, 2008; No *et al.*, 2015). In contrast, notwithstanding some notable exceptions (e.g. Glauben and Loy, 2003 and Fedoseeva, 2013), few studies are related to the European Union (EU). This paper investigates the existence of PTM behaviour in EU wheat exports. More specifically, it

estimates the relationship between export unit values and exchange rates using quarterly panel data from 11 countries (including the largest importers of wheat from the EU measured by value) 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 and exported in bulk (hence, it minimises product heterogeneity). As far as we are aware, this paper represents the first empirical investigation of PTM behaviour for EU wheat exports. Additionally, the paper aims to address some of the criticisms of past studies which employed annual data (Carew and Florkowski, 2003; Jin, 2008). Carew and Florkowski (2003) and Pall *et al.* (2013) are amongst the few studies that examine the time series properties of panel data. This study adopts a PTM approach given the interest in whether EU wheat exporters discriminate across a diverse range of export markets. While alternatives such as Residual Demand Elasticity (RDE) models can give a more sophisticated picture regarding the magnitude of market power (Pall *et al.*, 2014), they require data on cost and demand shifters which were not available for most of the countries examined in this analysis. Adopting an RDE approach, for example, would have thus narrowed greatly the scope of the research, in terms of country coverage.

Wheat is the world's third most important crop after rice and maize when measured by the value of production (FAO, 2015) and around 20% of global production is traded internationally (IGC, 2015). The total value of wheat exports in 2011 amounted to \$47.4 billion (IGC, 2015). Wheat is thus a strategically important commodity for which trade is politically sensitive, particularly for those importing countries where it is a staple food, and price volatility may make them more vulnerable to social unrest (Friebel *et al.*, 2015). Price discrimination may add to this, and the concentration of world wheat trading in the hands of a

few multinational firms (e.g. the ABCD companies²) and State Trading Enterprises ~~(an~~ oligopolistic market structure) increases the probability of discrimination between destination markets (Scoppola, 2007; Friebe *et al.*, 2015). Recent instability in trade and prices has 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 international trade of this commodity and the EU's main export markets (i.e. North Africa and the Middle East) are highly dependent on imports of 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 describes the empirical methodology. Section 5 presents the results and Section 6 summarises and concludes.

2. The EU and Wheat Exports

Global wheat production in 2013 amounted to 716 metric tonnes, of which 20% was by EU member states (FAO, 2015). In the EU, wheat is the largest grown cereal accounting for almost half of both total cereals output and cultivated area (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). The EU is also a major player in global cereals markets. Wheat and barley are the most important cereals exported, when measured by value, with 15% of the wheat crop exported every year. (European Commission, 2014). In 2013, the EU was the second largest exporter of wheat after the United States (US).

²Archer Daniels Midland (ADM), Bunge, Cargill and (Louis) Dreyfus are four major firms that dominate the global grain trade.

In international trade, there are four main types of wheat exporters. The most important actors are four large, global commodity traders (Archer Daniels Midland [ADM], Bunge, Cargill and Louis Dreyfus), collectively known as the ABCD companies. These companies account for approximately three quarters of the global grain trade (Productivity Commission, 2010). All of the ABCD companies have operations in the EU; for instance ADM exports wheat via Constanța in Romania and Bunge operates port terminal facilities in Poland. The other actors are State Trading Enterprises, some of which have been privatised (e.g. Cargill bought the newly privatised Australian Wheat Board), regional trading groups such as Wilmar International and COFCO (China National Cereals, Oils and Foodstuffs Corporation) group, which lack of the global scope of the ABCD companies, and other specialist or localised actors. However, it is difficult to obtain detailed data on the market share, and the operations of many these actors (Murphy *et al.*, 2012).

Table 1 presents recent information on the 11 EU wheat export markets considered in the PTM analysis. The EU has four main types of wheat export markets. The first group comprises *regular, high volume importers*, such as Algeria, Morocco and Egypt. These are the most important wheat export markets for the EU, measured by both volume and value. Second, there are *regular, lower volume destination markets* such as Albania and Iceland for which demand is smaller but relatively stable. Third, there are *variable volume importers* such as Belarus. For instance, in the latter case trade is often very thin but in some periods sales are substantial, reflecting instabilities in supply from Russia and Ukraine (Belarus's main wheat suppliers) due to weather related factors and intermittent trade restrictions (Gafarova *et al.*, 2015). Finally, there are *intermittent importers* where there are periods for which the EU records no sales and so are excluded from the time-series analysis conducted here. This group includes countries in East Asia, where the EU has a weaker presence

generally with Australian supplies benefiting from lower shipping costs and closer geographical proximity. Collectively the 11 countries studied, in recent years, accounted for between 46 and 67% of EU wheat exports, when measured by volume.

Table 1 about here

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, maintaining or even increasing their export prices to a foreign market when the currency of the import market appreciates. Following Knetter (1989), the profit, π , of an exporter selling to $i = 1, \dots, N$ export markets is:

$$\pi(P_1, \dots, P_N) = \sum_{i=1}^N P_i q_i(ER_i P_i; S_i) - C(\sum_{i=1}^N q_i(ER_i P_i; S_i), w) \quad (1)$$

where P_i is the price for country i in the exporter's currency, q_i is quantity demanded which is determined by the price in the buyer's currency, $ER_i P_i$, 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:

$$P_i = C_q \left(\frac{\eta_i}{\eta_i - 1} \right), \quad i = 1, \dots, N \quad (2)$$

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 $t = 1, \dots, T$ observations is derived by taking natural logarithms and totally differentiating:

$$\Delta \ln P_{it} = \theta_t + \beta_i \Delta \ln ER_{it} + \varepsilon_{it}, \quad i = 1, \dots, N, t = 1, \dots, T \quad (3)$$

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

$$\ln P_{it} = \theta_t + \lambda_i + \beta_i \ln ER_{it} + \varepsilon_{it} \quad (4)$$

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 prices 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, the export price equals 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 $\lambda_i=0$ and $\beta_i=0$, and θ_t measures 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 $\beta_i=0$ 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 $\beta_i = -0.5$ then a 10% appreciation (depreciation) of the importing country's currency 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-Germany 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 exports, 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 equation (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 export pricing decisions. 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 are used and the time-series properties of the data examined. Results show price discrimination across export markets, with an amplification of the effects of exchange rate changes in export markets. Lavoie (2005) tested for PTM to assess whether the Canadian Wheat Board (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 discriminated by charging different prices to different countries, even after accounting for differences in product quality. Like Lavoie, Jin (2008) also tested for PTM behaviour within CWB. By employing equation (4) and annual panel data for 1988-2003, results show that the CWB did exercise PTM behaviour, though 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.

In recent years, the increased role of Russia and other ex-soviet countries as key players on the world wheat market, led to several new PTM studies such as Pall *et al.* (2013), Friebel *et al.* (2015) and Gafarova *et al.* (2015). 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) where this is attributed to Russia's large share of wheat imports and/or lack of major competitors for these export markets. Similarly, Friebel *et al.* (2015) based their PTM analysis of Russian wheat exports on Knetter's panel model, conducting the analysis for two time periods, i.e. 2002-2011 and 2006-2011 (with high world wheat prices) using a firm-level dataset. Their results show a high degree of PTM with price discriminating behaviour by Russian firms in 25 of 61 destination markets during the longer period and 14 of 49 export markets for the shorter period. Gafarova *et al.* (2015), who examine price discriminating behaviour of Kazakhstan, Russia and Ukraine exporters in response to bilateral exchange rate fluctuations

during 1996-2012, find that there is evidence of PTM behaviour, though this is the exception rather than the norm. The authors attribute this to the relative weakness of ex-Soviet countries in international wheat markets, stemming from production being highly weather dependent, of undifferentiated quality and periodically affected by export restriction policies.

Notwithstanding differences in approaches, the PTM literature on agricultural markets as a whole is subject to two criticisms. First, amongst others, Carew and Florkowski (2003) and Jin (2008) use annual data, which are insufficiently fine-tuned to examine PTM behaviour (Fedoseeva, 2013). Second, only a minority of studies, such as Carew and Florkowski (2003) and Pall *et al.* (2013) examine the time series properties of the panel data. The former find mixed unit root results while the latter, perhaps surprisingly, find that export unit values and the nominal exchange rate are stationary, and both estimate (4) in both levels. We address both these 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).³ 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

³ By contrast to the usual two-way fixed effects model, β_i in (4) is allowed to vary between i .

equilibrium exists (Granger, 1988). To test for non-stationarity in a panel, a number of tests have been developed (Harris and Sollis, 2003) 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, the paper employs the test of Hadri (2000) where the null is that each individual series in the panel is stationary 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 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). 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 ($H_0: \rho_i=1$) and the alternative is that cointegration exists ($H_1: \rho_i<1$) 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; Pedroni, 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 $\ln P_{it}$ and $\ln ER_{it}$, their long-run relationship can be estimated. Pedroni (2000) 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 (2000) 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 are 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 . Parameter significance of both individual and panel elasticities is assessed by t-statistics with the usual confidence levels.

An alternative to Pedroni's (2000, 2001) method is the pooled mean estimator of Pesaran, Shin *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:

$$\Delta \ln P_{it} = \phi_i (\ln P_{it} - \theta_t - \lambda_i - \beta \cdot \ln ER_{it}) - \delta_i \Delta \ln ER_{it} + \omega_{it} \quad (5)$$

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 (€/tonne of wheat, f.o.b. prices)⁴ and exchange rates, obtained for 11 EU wheat export destinations (that regularly import EU wheat) for the period 2000-2013. Exchange rates are expressed as units of the importer's currency per unit of the

⁴ The f.o.b price data exclude any export subsidies. Export subsidies have not been used for cereals by the EU since 2006.

exporter's currency, for example, Albanian Lek/€. However, as the monthly data for these destinations contained missing observations on export unit values, both series are 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. Descriptive statistics are shown in the on-line Appendix.

We first test for panel unit roots in both export unit values and exchange rates using IPS and Hadri unit root tests. There is evidence of trends in some of the individual series, and they are included in unit root 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 $\ln ER_{it}$ 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 stationary against the alternative of a unit root. With four lags in the Bartlett window and allowing for heterogeneous serial correlation, Hadri-statistics for $\ln P_{it}$ 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 and 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.⁵ We concluded that there is a meaningful long-run relationship between export unit values and exchange rates.

Following Pedroni (2000, 2001), the results of estimating equation (4) using FMOLS and DOLS both with four lags in the Bartlett kernel are shown in Table 2. 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, $\chi^2_{10}=71.86$ (p-value: 0.00) and the null is rejected; and for the DOLS estimates, $\chi^2_{10}=7.99$ (p-value: 0.63) and the null is not rejected. There is some evidence therefore that optimal mark-ups, λ_i , 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, $\chi^2_{10}=20.65$ (p-value: 0.02); and for the DOLS estimates, $\chi^2_{10}=24.26$ (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 for Albania is also positive but

⁵ The existence of cointegration also implies that export unit values and exchange rates are non-stationary.

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 are 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, indicating local currency price stabilisation with a 1% increase in the exchange rate leading to 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 a 0.38% decrease in the export unit value. None of the positive PTM-coefficients from either estimator is significant. Overall, the evidence suggests that exports of wheat from the EU are not subject to PTM-behaviour, although for Belarus and Iceland there is local currency price stabilisation whereby exporters adjust mark-ups to maintain export prices relatively constant.

Table 2 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 $\chi^2_1=0.09$ (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 $\hat{\beta}=-0.83$ (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 2. The estimate of the speed of adjustment is $\hat{\phi}=-0.32$ (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.

Finally, we compare results from the commonly-estimated fixed effects model with our FMOLS/DOLS results (Table 2). 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). Estimates of the common time effects, θ_t for $t = 1, \dots, 56$ are not reported.⁶ The null is that θ_t are not significant 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 (i.e. that there is a constant mark-up), is rejected. The null that the β_i -effects are not significantly different is rejected and there is evidence of PTM-behaviour. The PTM-coefficients range from -0.33 to 1.57, that of the panel under the restriction that $\beta_i=\beta=0.05$ for $i=1,2,\dots,11$. There are clear differences between the results from the fixed effects model and those using FMOLS/DOLS, and the former appear spurious.

Reflecting on the FMOLS/DOLS results indicates that there is no significant evidence of PTM for EU wheat exports to *regular, high volume destination markets*, such as Algeria, Morocco and Egypt. These countries are substantial net importers of wheat and they are attractive markets also for major competitors such as Canada, Russia, Ukraine and US, limiting opportunities to price discriminate. In this regard, our results on the absence of PTM

⁶ In the fixed effects model, one country is taken as the reference, with other countries represented by dummies. The reference fixed effect is identified in the constant (b_1), and the individual country i fixed effects (reported here) are identified as $b_1 + b_i$. In this case, both the constants and slope coefficients differ by country as in (4), and this presentation facilitates easy comparison between the methods.

in EU wheat exports for such destination markets mirror those for ex-Soviet rivals (Pall *et al.*, 2014; Gafarova *et al.*, 2015). Overall, there is also little evidence of PTM in exports to *regular, lower volume destination markets* such as Albania which already imports most of its wheat from Russia, Serbia and Montenegro (ACIT, 2015). The DOLS estimations, but not FMOLS, indicate that in the case of Iceland EU exporters employ a local currency price stabilisation strategy, but in this instance most wheat is imported from North America and the incomplete exchange rate pass-through may reflect a desire to preserve market share. The one case of PTM identified in both the FMOLS and DOLS results is that of Belarus, where volumes of EU wheat imports are highly variable. In this case the estimates, by employing export unit values, are sensitive to the EU's intermittently changing status from a bulk, high volume exporter to niche, high mark-up but very low volume supplier. As a result, changes in the quality composition of exports and the effects of economies of scale may not be fully captured in our analysis.

6. Summary and Concluding Remarks

Since the seminal studies of Krugman (1986) and Knetter (1989), a substantial literature has emerged on PTM. Initial research has been criticised 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 (Pick and Carter, 1994; Carew and Florkowski, 2003; Lavoie, 2005; Jin, 2008), the US (Pick and Park, 1991; Pick and Carter, 1994; Carew and Florkowski, 2003) and Russia or other ex-soviet countries (Pall *et al.*, 2013; Friebe *et al.*, 2015; Gafarova *et al.*, 2015). 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 major export destination markets for 2000-2013.

In our empirical method, we first 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, which also implies that both series are $I(1)$.

We estimate the PTM-model using both fully modified ordinary least squares (FMOLS) and dynamic ordinary least squares (DOLS) following Pedroni (2000, 2001), and both methods produce similar results. There is little evidence of differential mark-ups between export markets although those of 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 stabilisation is implied whereby EU exporters adjust their mark-ups to maintain constant prices by absorbing part of a change of 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.

There are clear differences between these FMOLS/DOLS results and those from a fixed effects model, which is commonly estimated in the literature for PTM-models that use panel data. The fixed effects model appears to indicate more widespread PTM. However, there is some evidence that both export unit values and exchange rates are non-stationary; in which case the fixed effects results may well be spurious and empirical PTM-analyses should examine the time series properties of the variables.

Reassuringly for policy makers and competition authorities, the PTM analysis identifies little evidence of price discrimination across markets for EU wheat exports. In particular neither the FMOLS nor DOLS results identify PTM behaviour for the EU's regular, high volume destination markets, which are highly dependent on wheat imports. The (re)emergence of Kazakhstan, Ukraine and Russia as significant wheat exporters, along with enduring presence of established competitors like Canada and US restricts the scope for price discrimination. There is no evidence of discrimination against developing countries. The results for Belarus, where both the FMOLS and DOLS results suggest PTM behaviour should be treated with caution given that trade volumes have been erratic and in some time periods very thin.

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Table 1: Wheat export unit values and tonnes exported by the EU to 11 destination markets included in the PTM analysis

Country	Unit value of exports (€ per tonne)			Tonnes exported to country from EU			Share of total EU wheat exports by volume (%)		
	2011	2012	2013	2011	2012	2013	2011	2012	2013
Norway	264	235	270	328,137	347,908	372,905	1.96	2.18	1.70
Switzerland	263	240	258	339,702	311,676	358,628	2.03	1.95	1.63
Albania	243	244	298	108,989	29,468	42,050	0.65	0.18	0.19
Belarus	330	350	379	29	119	14,660	0.00	0.00	0.07
Congo D.R.	250	222	247	71,700	46,847	117,342	0.43	0.29	0.53
Mauritania	281	261	276	147,116	88,616	273,797	0.88	0.55	1.25
Tunisia	232	225	274	598,653	254,545	832,834	3.58	1.59	3.80
Egypt	240	233	252	1,120,565	1,453,067	1,440,341	6.70	9.09	6.56
Morocco	230	234	246	2,502,482	1,285,404	1,645,751	14.96	8.04	7.50
Algeria	250	234	238	5,982,180	3,940,228	4,879,694	35.76	24.64	22.24
Iceland	247	228	249	33,699	39,664	35,936	0.20	0.25	0.16
Average or Total values	257	246	272	17,530,172	15,989,655	21,943,666	67.14	48.77	45.63

Source: own calculations based on Eurostat (2014)

Table 2: Panel Results

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

Notes: 1. FMOLS/DOLS equations: $\ln P_{it} = \lambda_i + \beta_i \ln ER_{it} + \gamma_i \text{Trend}_{it} + \varepsilon_{it}$.

2. Fixed effects equation: $\ln P_{it} = \theta_t + \lambda_i + \beta_i \ln ER_{it} + \varepsilon_{it}$.

3. t-statistics in parentheses.

Pricing to market analysis: the case of EU wheat exports

Philip Dawson, Matthew Gorton, Carmen Hubbard and Lionel Hubbard

On-Line 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 D.R.	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).