The impact of market deregulation on milk price: A dynamic panel data approach

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Abstract

The scope of this paper is to investigate the impact of market deregulation on the competitiveness of raw milk producers in Greece along the suggested lines of OECD (OECD, 2014). The study uses a dynamic panel data approach, to assess changes in the relative competitiveness of milk producers as a result of certain deregulation policies imposed by the Greek government in two phases (May 2014 and September 2015). In order to account for the presence of cross-sectional dependence and non-stationarity, the empirical analysis implements novel panel econometric methodology namely Common Correlated Effects (CCE) and Augmented Mean Group estimators (AMG). Our sample uses micro level data drawn from the 45 Greek regions spanning the period from January 2010 to October 2017. By comparing the wholesale prices of milk affected by regulation before and after the policy changes, we infer that abolishing regulation led to an increase in the prices of the wholesalers and thus in their profitability levels. Moreover, we argue that the openness of the relevant milk market segment had significant implications to the level of competition in the sector. Lastly, our empirical findings which confirm the OECD competition guidelines in the milk sector remain rather robust under different empirical methodologies and sample splitting, providing a focal point to policy makers and government officials for the ex-post evaluation of the deregulation strategies.

Keywords: Deregulation, Competition; Milk price, Dynamic panel models, OECD
JEL codes: L1; L51; L52; C23

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1. Introduction

Regulation is often regarded as a means of state interventionism in imperfect competitive markets to increase producer and consumer surplus and thus social welfare (Viscusi et al, 2005). Since the scope of regulation is set it is important to wonder how deregulation policies might actually affect the overall performance of an industry/sector and change *inter alia*the level of wholesale prices (Genakos, et al, 2018). This study tries to investigate the relevant research question and attempt to fill the gap in the empirical literature regarding the post-evaluation of certain deregulatory measures implemented in an oligopolistic market such as milk industry.

It is worth emphasizing that the ex-post evaluation literature on the impact of deregulation policies on prices and profits is rather limited. In an influential study Genakos et al, (2018) by using a Difference-in-Difference methodology investigate the effect of maximum wholesale and retail markup regulation in certain Greek oligopolistic sectors. They argue that deregulation led to significant price decreases, corresponding to an estimated €256 million yearly decrease in consumer expenditure. Their study provides indirect but consistent evidence that the most likely explanation for this outcome was collusion by the wholesalers. In a similar study, Hendel el al, (2017) examine the effect of a consumer boycott on cottage cheese in Israel started back in 2011. They claim that the boycott led to an immediate decline inprices, which remain low even six years later. Moreover, the empirical findings reveal that own and cross-price elasticities increased substantially after the boycott, while post-boycott prices were substantially below thelevels implied by the post-boycott demand elasticities. This study highlights the consumers' unilateral effects such as activism and boycotts in order to mitigate the market power in a heavily regulated oligopolistic market.

Regulatory barriers in specific food sectors, such as the dairy industry that hinder the level of effective competition, have been already identified by the national governments or world organizations and institutions (EU, OECD, IMF, etc). Obstacles are often created by the fact that the legislation frequently provides restrictive definitions of certain foodproducts or their components, or activities which do not take into consideration recent technological developments or the practice in other EU countries (OECD, 2014).To give an example, one might consider that legislation concerning the shelf life of milk protects local markets from internal and foreign competition to the detriment of consumer choice and welfare.

Along this point of skepticism, it is important to stress that the Greek pasteurized retail milk sector was characterized by a strict legislative regime concerning the shelf life of milk until 2015. This year the shelf price of pasteurized milk was fully deregulated, while prior to 2015 the Greek government partially deregulated the shelf life of it from five to seven days (April of 2014). The restrictive legislative regime of shelf life of pasteurized milk prior to April of 2014 led to the following consequences: a) higher prices of pasteurized milk than the EU average, b) the exclusion of raw milk imports in the Greek market, c) the limitation of competitive pressure at all stages of production, d) the limitation of direct imports of final products from retailers and importing companies from other countries, e) the distortion of product availability and consumer choice, especially for consumers in Greek islands and mountain villages and finally to the discouragement of new small farmers to supply new products in the market (OECD, 2014).

However, Greek farmers ("*producers*") do not share the above considerations. On the contrary, they tend to believe that deregulation of shelf life of pasteurized milk in 2014 and 2015 have led to the reduction of wholesale prices, the reduction of quality of milk supplied to the final consumers and the increase of raw milk imports in the Greek market. Figure 1 shows that selling prices of raw cow's milk in five EU countries (Greece, Malta, Finland, Italy and Sweden) and EU average from 2005 to 2016. It is evident from Figure 1 that Greece continues to be among the countries with the highest wholesale prices in EU, but with a downward trend from 2013 onwards. Malta continues to exhibit the highest prices for almost the whole period, while in Finland, Italy, Sweden and EU average, the selling prices of raw cow's milk indicate a sharp declining tendency the last two years (2015 – 2016). As it is evident, the effect of market deregulation and openness on the evolution of wholesale milk prices is an important exercise that need to be carefully executed.

<Insert Figure 1 about here>

In this paper we use an unbalanced panel dataset of 4,230 monthly observations for 45 Greek regions, spanning the period from January 2010 to October 2017, to determine the driving forces of wholesale price of milk in Greece. For this purposes we apply two dynamic GMM estimators (DIF-GMM and SYS-GMM), the panel DOLS methodology and we supplement our analysis with econometric methodologies which allow for heterogeneous slope coefficients across group members, nonstationarity and cross-section dependence (Mean Group estimator, Common Correlated Effects Mean Group estimator and the Augmented Mean Group estimator).

This paper contributes the literature in many fronts. First and foremost, this is the first study to the best of our knowledge that tries to assess the impact of deregulation and market openness of the Greek milk industry on wholesale prices. In this way, we infer if changes in the relative competitiveness of milk producers are linked with certain deregulation policies gradually imposed by the Greek government on behalf of the OECD competition guidelines. Therefore, we put the OECD policy recommendations for the deregulation of the milk sector in Greece (i.e extension of the shelf life for fresh milk, lifting barriers to entry, etc) into empirical scrutiny by performing a post evaluation analysis targeted at the wholesale milk prices. The later constitutes an additional novelty of this study, acting an important companion to policy makers and government officials in their effort to further enhance the level of competition in the specific sector. Lastly, from the econometric perspective, this study goes beyond the existing literature in that it uses a battery of dynamic panel data techniques (i.e GMM, DOLS and Mean Group estimators) in order to capture the dynamic interactions between the sample variables to further explore the validity of our findings. This kind of analysis is generally new in the empirical literature and may help practitioners and government officials in their attempts to understand the driving forces of milk sector.

Our article is related to the literature that broadly studies the effect of market deregulation on the competitiveness of a certain industry (i.e milk sector).Even though the topic might be considered narrow, the literature on it is broad in its relevance, being pertinent to economic theory, dynamic competition, pricing, pass-through, collusion (Albæk et al, 1997), market power and the link between information and consumer demand, among others. One would expect that this volume of work would have led to some robust conclusions on the way that deregulation policies pursue industry competitiveness. Though a number of important contributions have been made (see among others Genakos et al, 2018; Hendel et al, 2017; Abito, et al, 2016; Carranza et al, 2015; Ofek, 2012; Kahal, 2011; Davis and Kilian, 2011) the basic question of how large and persistent are the deregulation measures, has drawn widely different answers. We try to shed some light on this debate.

The results that emerge from the empirical analysis clearly show that abolishing regulation led to a significant increase in the level of wholesale prices. Moreover, we argue that the openness of the relevant milk market segment had significant implications to the level of competition in the sector. It is noteworthy that our empirical findings remain rather robust under different empirical methodologies and sample splitting, providing a focal point to policy makers and government officials for the ex-post evaluation of the deregulation strategies

The remaining of the paper is organized as follows. Section 2 describes the milk sector in Greece with all its regulatory interventions. Section 3, presents the data and the empirical framework used in the analysis, while Section 4 discusses the different econometric methodologies employed. Section 5 presents the results along with the necessary preliminary testing and the robustness checks. Finally, Section 6 concludes the paper, while offering some policy recommendations.

2. The Greek milk sector

Greek regulation regarding pasteurised milk took a conservative approach regarding its regulation and definition during the period from 1959 to 2014 (see Table 1). Specifically, Royal decree 4/1959 on Veterinary and hygiene check of milk determined that the shelf life of pasteurised (fresh) milk cannot exceed 2 days, while it identified two types of milk, fresh and pasteurised. Presidential decree 430/1981 extended the shelf life of pasteurised (fresh) milk to 3 days, while presidential decree 104/1988 determined that the shelf life of pasteurised (fresh) milk cannot exceed 4 days. Presidential Decree 113/1999 determined that the shelf life of pasteurised milk cannot exceed 5 days. It also identified two types of milk two types of pasteurisation procedures: the low temperature pasteurization (*at least +71.7^o C for 15 seconds*), where the shelf life is defined at maximum five (5) days and the high temperature pasteurization (*between 85^o - 127^o C*), where the maximum shelf life is at the discretion of the manufacturer.

<Insert Table 1 about here>

The first attempt of Greek legislator to deregulate the restrictive Greek regime for the shelf life of pasteurised milk was on 2014. Law N. 4254 of 4th April 2014 determined that the shelf life of pasteurised milk cannot exceed 7 days. As it concerns the types of milk the Law inserted one difference beside the Presidential Decree 113/1999: it explicitly defined the notion of pasteurised milk (*at least +71.7° C for 15 seconds or at least +63° C for 30 minutes*). Full deregulation of the restrictive Greek regime for the shelf life of pasteurised milk was adopted on 14th of August 2015. Law N. 4336 determines that the shelf life of pasteurised milk and the maximum shelf life is at the discretion of the manufacturer.

Relevant EU legislation takes a less conservative approach regarding the definition of pasteurised milk. According to it, the manufacturer of milk is the only responsible to guarantee the safety of its product and specify the date of minimum durability up to which pasteurised milk may be consumed (Regulation (EC) No 852/2004[2]). Moreover, Regulation (EC) No 852/2004[3] defines two types of non-condensed milk: pasteurized milk, which is processed at low/high temperatures[4] and ultra high temperature (UHT) milk, which is processed at higher temperatures and needs not be refrigerated[5].

From the above it turns out that, on the retail level, four types of drinking milk may be found in the Greek market: a) pasteurised milk, b) high-pasteurised milk, c) UHT milk, and d) condensed milk. According to EU law, both the pasteurised and the high-pasteurised types of milk

[2]See http://eur-lex.europa.eu/LexUriServ/LexUriServ.do?uri=OJ:L:2004:139:0001:0054:en:PDF.
[3]See http://eur-lex.europa.eu/LexUriServ/LexUriServ.do?uri=OJ:L:2004:139:0055:0205:en:PDF.
[4] "*Pasteurisation is achieved by a treatment involving: i) a high temperature for a short time (at least 72°C for C for 15 seconds); ii) a low temperature for a longtime (at least 63°C for C for 30 minutes);or iii) any other combination of time-temperatureconditions to obtain an equivalent effect*."
[5]"*Ultra high temperature (UHT) treatment is achieved by a treatment: i) involving a continuous flow of heat at a high temperature for a short time (not less than 135°C ...., and ii) sufficient to ensure that the products remain microbiologically stable after incubating for 15 days at 30°C in closed containers or for seven days at 55°C in closed containers or after any other method demonstrating that the appropriate heat treatment has been applied.*"

would consist of a single category and any differences would be reflected only in the shelf life printed on the product packaging.

Figure 2 plots the evolution of consumed milk quantity (log prices measured in tones) for the top four regions in Greece (Thessaloniki, Larissa, Xanthi, Serres). The top four regions possess the 55% of milk quantity demanded in Greece during the period from January 2010 to October 2017. Figure 1 indicates that in Thessaloniki and Xanthi the demand for milk follows a negative path, while in Larissa and Serres the demand for milk follows a positive path.

Around the partial deregulation of the restrictive Greek regime for the shelf life of pasteurized milk (April 2014) consumption of milk depicts a decline until September 2014 and an immediate increase onwards until May 2015. However, the decline of milk consumption begins prior to April 2014 (January 2014), following by a huge increase of consumption from almost May 2013. This is more evident for regions 16 and 25 than in the other two regions of Figure 2.

<Insert Figure 2 about here>

Around the full openness of the restrictive Greek regime for the shelf life of pasteurized milk (August 2015) consumption of milk depicts a decline until the end of the year 2015, following by a sharp increase onwards until March 2016. However, the decline of milk consumption begins prior to August 2015 (May 2015), following by a huge increase of consumption from almost September 2014. Overall, putting the two periods together (prior and after the deregulation), Figure 1 depicts that consumed milk quantity prices follow a yearly cycle.

Figure 3 plots the wholesale prices of pasteurized milk regarding the four most productive regions (big "*four*") in Greece namely Thessaloniki, Larissa, Xanthi and Serres during the deregulated period (May 2014-. It is evident that after the full openness of the shelf life of the pasteurized milk on August 2015 the wholesale prices depicted an increase in all of the four regions. Specifically, in Thessaloniki (region 16) the average wholesale price increase of milk is about 2.73% in January 2016, after an immediate average decrease of about -1.02% in September 2015. In Larissa (region 25) the average wholesale price increase of milk is about 4.83% in January 2016, after an immediate average decrease of about -1.39% in September 2015. The corresponding values for the other two regions (regions 31 and 38 accounting for Xanthi and Serres respectively) in January 2016 are about 1.91 and 1.40%, while in September 2015 the average wholesale price decreases of pasteurized milk are about -0.97 and -1.06 respectively.

<Insert Figure 3 about here>

During the period from the deregulation to the full openness of the shelf life of pasteurized milk (May 2014 – August 2015), its wholesale price exhibits a marginal decrease. For instance, in Thessaloniki (region 16), the average decrease is about -0.22% (the lowest value), while in Larissa (region 25), the average decrease is about -0.25% (the highest value). However, during the period from May 2014 until January 2015, that is, eight months after the deregulation on April 2014, in two regions, Thessaloniki and Larissa, the average wholesale price of pasteurized milk exhibits an increase of about 0.06% and 0.24% correspondingly.

Lastly but not least, from January 2017 until October 2017 the average wholesale price of pasteurized milk exhibits an increase in the four regions. That is, in Xanthi the average wholesale price increase is about 0.62% (the highest value), while, in Serres, the corresponding increase is about o.33% (the lowest value).

Figure 4 shows the histogram of the log wholesale milk constant prices (horizontal axis) over the entire period and regions under scrutiny scaled to density (vertical axis). It is evident from the following figure that the histogram is close to symmetric, which means that the mean and median are close to each other. That is, the data is fairly balanced on each side.

<Insert Figure 4 about here>

3. Data and empirical modelling

Our econometric analysis is based on an unbalanced panel dataset of 4,230 monthly observations, spanning the period from January 2010 to October 2017 (N = 45 and T =94).The selected sample includes 45 Greek regions, with five regions being omitted (Zakynthos, Kastoria, Kefallinia, Korinthia and Messenia). The starting date for the study was dictated strongly by data availability, while the final date observation (October 2017), represents the last month for which data mostly regarding the wholesale milk prices were available at the time the research was conducted.

Specifically, the empirical analysis followed in this study employs real wholesale milk price as the dependent variable drawn from the Hellenic Agricultural Organization (HAO) "*DIMITRA*".[6] On the right hand side (RHS) of the equations used in our dynamic models we include several independent variables (covariates) to capture the main determinants of the wholesale milk price variations. Specifically, we used the consumed milk quantity per region measured in tones. Moreover, we include the level of nominal milk income expenditure of a specific region which is converted to its real income expenditure form, using the Harmonised Consumer Price Index (HCPI), given that consistent estimates of regional disposable per capita income are not available for the sample period under examination. While data on the level of quantity milk consumed per region and nominal income expenditure for each of the 45 regions are obtained from the HAO, HCPI data are obtained directly from Eurostat (Economy and finance

[6]For the years 2010-2015 all the data for prices and milk quantities are drawn from the Hellenic Milk and Meat Organization.

database). To the empirical ends of the analysis, we also include the total number of milk producers per regions a proxy for the level of market structure.

The inclusion of two dummy (dichotomous) variables and their interactions with the RHS covariates supplements our empirical modelling. The first binomial variable takes the value one when the introduction of the seven day fresh milk was legally implemented (from May 2014 onwards) and zero otherwise (up to April 2014). This variable measures the level of market deregulation on pasteurized milk. We also include a second dummy variable taking the value one when full market openness was introduced (from September 2015 onwards) and zero otherwise (up to August 2015). In addition, we focus on the 45 Greek regions, while we drop five regions from the empirical analysis since data for the milk quantity demanded were not available over the whole sample period or completely missing.

Similarly to Katsoulacos et, al, (2014) we assume after imposing zero homogeneity, the basic following log-linear inverse demand function of the form:

$$p = \beta_o + \beta_1 q + \beta_2 pm + \beta_3 f + \varepsilon \tag{1}$$

where: $p = \log(P)$, $q = \log(Q)$, $pm = \log\left(\frac{P}{M}\right)$, $f = \log(F)$, P is the price, Q is the quantity, $M = P \times Q$ is income (total expenditure) and F is the number of producers in the 45 regions. Finally ε is a random shock with $E(\varepsilon | q, pm, f) = 0$.

The following table summarizes the main descriptive statistics. For the empirical exercise presented in the next section, all data (except for the dummy variables) are presented in a logarithmic form. From the careful inspection of Table 2, it is evident that the data are well behaved, showing limited variability in relation to the mean of the population since the coefficient of variation does not exceed 50% in all of the cases. In contrast, the variables as expected do not follow the normal distribution, since the relative values of the skewness and kurtosis measures are not zero and three respectively.

<Insert Table 2 about here>

4. Empirical methodology

In this study, we apply panel methods which take into account both cross-section and time dimensions of the data. However, when the errors of a panel regression are cross-sectionally correlated then standard estimation methods can lead to inconsistent estimates and incorrect inference (Phillips and Sul, 2003; Apergis 2016).

In order to take into account possible endogeneity issues (reverse causality) arising from the inclusion of income expenditure as an independent variable in tandem with other RHS variables such as the milk quantity demanded per region, we apply two dynamic GMM estimators (DIF-GMM and SYS-GMM) and the panel DOLS methodology fully developed by Kao and Chiang (2000). Lastly, we supplement our analysis with novel econometric methodologies which allow for heterogeneous slope coefficients across group members, nonstationarity and correlation across panel members (cross-section dependence) as in our case. Specifically, we implement three different estimators namely the Pesaran and Smith (1995) Mean Group estimator (MG), the Pesaran (2006) Common Correlated Effects Mean Group estimator (CCE) and the Augmented Mean Group estimator (AMG), introduced by Bond and Eberhardt (2009) and developed by Eberhardt and Teal (2010).

4.1 Dynamic GMM estimators

With the intention to examine the dynamic aspects we use dynamic panel data techniques such as Difference Generalised Method of Moments (DIF-GMM) estimators attributed to Arellano and Bond, (1991) and System Generalised Method of Moments (SYS-GMM) estimators proposed by Arellano and Bover (1995) and Blundell and Bond (1998) respectively. The use of the latter is mainly justified as it improves significantly the estimates' accuracy and enlarges efficiency when the lagged dependent variables are considered as poor instruments as in the first-differenced regressors (Greene, 2003, Baltagi, 2002, Halkos and Polemis, 2017; Abid, 2017). As a consequence, the SYS-GMM gives more robust results than the first-differenced GLS and GMM estimation methods (Bond et al., 2001).

In our case and in modelling dynamic effects we have the lagged dependent among the independent variables in the following form:

$$Y_{it} = X'_{it}\beta + \delta Y_{i,t-1} + \alpha_t + u_{it} \qquad \text{i=1,2,…,N , t=1,2,…,T} \tag{2}$$

where δ being a scalar, X'_{it} 1xK and β Kx1 and u_{it} follow a one-way error component model ($u_{it}=\mu_i+v_{it}$); with $\mu_i \sim IID(0,\sigma_\mu^2)$ and $v_{it} \sim IID(0,\sigma_v^2)$ are independent of each other and between them. As Y_{it} is a function of μ_i then $Y_{i,t-1}$ is also a function of μ_i and it is correlated with the error term. We have used panel data methods to estimate the above equation. Then the first difference GMM estimation is given as

$$\hat{\delta}_{GMM} = \frac{\left(\sum_{i=1}^{N} \Delta y_{i,t-1} Z_i\right) W_N \left(\sum_{i=1}^{N} Z'_i \Delta y_i\right)}{\left(\sum_{i=1}^{N} \Delta y_{i,t-1} Z_i\right) W_N \left(\sum_{i=1}^{N} Z'_i \Delta y_{i,t-1}\right)} \tag{3}$$

With the choice of W_N being important with the first-step consistent estimator of d being

$$W_N^* = \frac{1}{\left(\frac{1}{N}\sum_{i=1}^{N} Z'_i \Delta\hat{\varepsilon}_i Z_t\right)} \tag{4}$$

In (4) if X_{it} are predetermined with current and lagged X_{it}s uncorrelated with current term then $E(X_{ij}u_{is})=0$ for $s \geq t$. A combination of strictly exogenous and predetermined X variables may more realistic compared to the two extreme cases with matrix Z_i adjusted according to each case. Arelano and Bover (1995) integrated this approach with the instrumental variables of Hansen and Taylor (1981) with individual series being highly persistent and δbeing near to one. Estimation details, of the DIF-GMM and SYS-GMM estimators are provided in the Appendix.

Based on the above, the dynamic specifications of the models are given by the following reduced form equations:

$$Log(P_{it}) = n_i + \gamma_t + b_o + \sum_{l=1}^{L} d_{l,i}\log(P_{it-l}) + \sum_{m=0}^{M} b_1 \log(Q_{it-m}) + \sum_{n=0}^{N} b_2 \log(M_{it-n}) + \sum_{r=0}^{R} b_3 \log(Firms_{it-r}) + c_1 D7 + c_2 D7 \times \log(Q_{it}) + c_3 D7 \times \log(M_{it}) + c_4 D7 \times \log(Firms_{it}) + u_{it} \quad (5)$$

$$Log(P_{it}) = n_i + \gamma_t + b_o + \sum_{l=1}^{L} d_{l,i}\log(P_{it-l}) + \sum_{m=0}^{M} b_1 \log(Q_{it-m}) + \sum_{n=0}^{N} b_2 \log(M_{it-n}) + \sum_{r=0}^{R} b_3 \log(Firms_{it-r}) + c_1 Dopen + c_2 Dopen \times \log(Q_{it}) + c_3 Dopen \times \log(M_{it}) + c_4 Dopen \times \log(Firms_{it}) + u_{it} \quad (6)$$

Wherei = 1,2,…45, t = 1,2,…94 and l is the time lag operator for the dependent variable.

The interpretation of the variables comes as follows. P stands for the wholesale milk prices per region deflated by the Harmonised Consumer Price Index (2015=100). Q is the milk quantity demanded per region in tones. M is the income expenditure per region deflated by the Harmonised Consumer Price Index (2015=100). The variable Firms stands for the total number of milk producers per region. D7 is the dummy variable taking the value one when the introduction of the seven day fresh milk was legally implemented (May 2014) and zero otherwise. Dopen denotes the dummy variable taking the value one when full market openness was introduced (September 2015) and zero otherwise. The relevant equations (Eq. 5 and 6) include also the interactions (cross terms) of the two dummy variables (D7 and Dopen) with the rest covariates capturing the possibility of non-linear effects in the formulation of wholesale milk price.

Moreover, n_i is the unit-specific residual that differs between regions but remains constant for any particular region (unobserved region level effect); while γ_t captures the time effect and therefore differs across years but is constant for all regions in a particular year. Moreover, we have also used sector fixed effects in our model. L, M, N and R are the lag operators determined by the minimisation of the Bayesian Information Criterion (BIC). Lastly, u_{it} denotes the i.i.d error term.

4.2 The dynamic OLS methodology

Stock and Watson (1993) provide a parametric approach for estimating long-run equilibria in systems which may involve variables integrated of different orders but still cointegrated. The potential of simultaneity and small-sample bias among the regressors is dealt with by the inclusion of lagged and led values of the change in the regressors (Phillips and Loretan, 1991; Saikkonen, 1991). More precisely, DynamicOLS methodology(DOLS) is employed to estimate the single cointegrating vector that characterizes the long-run relationship among the variables under scrutiny. It is simply a regression of one of the variables onto contemporaneous levels of other variables, a constant and leads and lags of their first differences using the methodology of Ordinary Least Squares.

The empirical model to be estimated under the DOLS methodology is specified in the following way:

$$Y_{it} = \beta_0 + \vec{\beta}X + \sum_{j=q}^{p} \vec{d}_j X_{t-j} + u_t \qquad (7)$$

where Y is the dependent variable and X is a matrix of explanatory variables, $\vec{\beta}$ is the cointegrating vector and q and p are the leads and lags correspondingly. The cointegrating vector represent the long – run cumulative multipliers, that is, the long – run effect of a change in X's on Y. Lead and lag terms included in Equation (7) have the purpose of making its stochastic error term

(u_t) independent of all past innovations in stochastic regressors (i.i.d). Unit root tests are performed on the residuals of the estimated DOLS regression, in order to test whether its stochastic error is unit-root nonstationary (Choi et al, 2008). In this study, we apply the panel DOLS estimator developed by Kao and Chiang (2000) suitable for cointegrated panel data with homogeneous long-run covariance structure across cross-sectional units.

4.3 The mean group estimators

Pesaran and Smith (1995) propose fitting separate regression for each country and calculate a simple arithmetic average of the coefficients. However, the Mean Group estimator (MG) does not concern with cross - section dependence. More precisely, Pesaran and Smith (1995) propose the estimation of equation (1) for each panel member (country) with the intercept to capture the fixed effects and a linear trend to capture time – variant unobservables. Finally, coefficients b_j are averaged across countries.

The Common Correlated Effects (CCE) methodology, recommended by Pesaran (2006) and Kapetanios et al. (2011), takes into account that the errors of a panel regression are cross-sectionally correlated and heteroskedastic (cross-sectional dependence).

Lastly, the Augmented Mean Group estimator (AMG) was developed in Bond and Eberhardt, (2009) and Eberhardt and Teal, (2010) as an alternative to Pesaran's CCE approach. The AMG approach is implemented in three levels: in the first level is performed the "*common dynamic process*" (an OLS estimation of augmented pooled regression model with the use of year dummies). In the second level the group specific regression model is augmented with the estimated coefficients of year dummies. Like the Pesaran and Smith (1995) methodology the regression model includes an intercept, which captures time – invariant fixed effects. Lastly, in the third level the group specific model parameters are averaged across countries.[7]

5. Results and discussion

In this section we present our empirical findings obtained by the implementation of standard dynamic panel methodologies (DIF and SYS-GMM, DOLS) as well as the two main mean group estimators (CCE and AMG) which serve as robustness checks to ensure the validity of the econometric results. Our empirical analysis also supplemented by some necessary preliminary tests accounting for the investigation of cross-section dependence and stationarity along with cointegration usually overlooked by similar studies.

5.1. Preliminary testing

5.1.1 Cross-section dependence

One of the additional complications that arise when dealing with panel data compared to the pure time-series case, is the possibility that the variables or the random disturbances are correlated across the panel dimension. The early literature on unit root and cointegration tests adopted the assumption of no cross-sectional dependence. However, it is common for micro-level data to violate this assumption which will result in low power and size distortions of tests that assume cross-section independence. For example, cross-section dependence in our data may arise due to common unobserved effects due to changes in milk legislation. Therefore, before proceeding to unit root and cointegration tests we test for cross-section dependence. We use the cross-section dependence test (CD test) developed by Pesaran (2004). As it is evident from Table 3 the relevant tests strongly rejects the null hypothesis of cross-section independence. In face of

[7]Estimation details, of the three Mean Group Estimators are provided in the Appendix.

this evidence we proceed to test for unit roots using tests that are robust to cross-section dependence (i.e second generation tests for unit roots in panel data).

<Insert Table 3 about here>

5.1.2 Unit root tests and cointegration

To examine the stationarity properties of the variables in our models we use two second generation unit root tests namely the Fisher type Augmented Dickey Fuller (ADF) test developed by Choi (2001) and Pesharan (ADF) both suitable for unbalanced panel data set and cross-section dependence. The former combines the p-values from N independent unit root tests, as developed by Maddala and Wu (1999). Based on the p-values of individual unit root tests, Fisher's test assumes that all series are non-stationary under the null hypothesis against the alternative that at least one series in the panel is stationary. Unlike the Im-Pesaran-Shin (2003) test, Fisher's test does not require a balanced panel as in this case (Merryman, 2004). The second test runs the t-test for unit roots in heterogenous panels with cross-section dependence, proposed by Pesaran (2003). Parallel to Im-Pesaran-Shin (2003) test, it is based on the mean of individual DF (or ADF) t-statistics of each unit in the panel. Null hypothesis assumes that all series are non-stationary. To eliminate the cross dependence, the standard DF (or ADF) regressions are augmented with the cross section averages of lagged levels and first-differences of the individual series (Lewandowski, 2006). As it is evident from the inspection of Table 4, both tests support the presence of a unit root across all four variables under consideration.

<Insert Table 4 about here>

In order to investigate whether a long-run equilibrium relationship exists among the sample variables we implement Pedroni's (1999) ADF-based and PP-based cointegration tests as well as

Kao's (1999) ADF-based tests. All these tests allow for cross-section dependence and are suitable for an unbalanced panel data (Pedroni, 2000; 2001). This is the reason for not using the error-correction-based panel cointegration tests proposed by Westerlund (2007) broadly applied in similar empirical studies. The latter also allow for cross-section dependence and represent an error-correction approach to testing for cointegration that are based on the statistical significance of the error correction term. The intuition behind this approach is that if a long run relationship between the variables in our model, we can write a regression that allows us to estimate the error-correcting terms which reflect the response of the system to random shocks that "*pushes*" the system towards its long-run equilibrium point. If the error-correction terms are significantly different from zero across sections, then there is evidence in favor of the existence of a long-run relation.

The results of the tests are presented in Table 5. All seven tests suggest the rejection of the null hypothesis of nocointegration null at any reasonable significance level.This means that cointegration statistics provide sufficient evidence to support the existence of a structural relationship between the wholesale milk price and of long run relations between the variables of the inverse demand function.

<Insert Table 5 about here>

5.2 Empirical findings

In the previous section we found evidence in favor of cointegration. Hence, our next step is to estimate the long-run equilibrium relationships.Tables 6 and 7 present the empirical results from dynamic GMM and OLS methodologies (eq. 5-7)

Table 6 presents the dynamic panel estimators of DIF/SYS GMM and OLS methodologies and of the whole sample under scrutiny. It is evident from Table 6 that the estimates are all highly statistically significant and robust given that eq. (5), (6) and (7) represent structural and not spurious long-run relations.[8] In every econometric model employed we obtain plausible signs of the estimated coefficients. The estimations of coefficient b_0 in eq.5 and 6 and the estimated coefficient of X_{t-j}in eq. 7 are always statistically significant and smaller than 1 for all the dependent variables employed within the 45 regions under scrutiny. For instance, the highest significant estimate is 0.632 under SYS-GMM methodology. We also estimate the dependent variable with two lags in the right hand side of eq. 5and with three lags in the right hand side of eq. 6 since it is found to be (highly) statistical significant in all the empirical models employed. This result strengthens the importance of the inclusion of the lagged dependent variable in the right hand side of eq. 5, 6 and 7 (Fotis and Polemis 2018; Fotis et al. 2017).

<Insert Table 6about here>

In every econometric model employed we obtain plausible signs of the estimated coefficients. The demand for pasteurized milk in Greece is price inelastic since the estimated coefficients of electricity price has the appropriate sign and is below unity (Bouamra-Mechemache et al. 2008; Wilson and Thompson 1967). For instance, a 1% percent increase of quantity of milk supplied will cause a fall of wholesale price by almost 0,15% under SYS-GMM methodology and almost 0,5% under DOLS methodology. Therefore, the inversedemand function of pasteurized milkis more sensitive under the latter rather than under the former methodology, but in both cases is price inelastic.[9]

[8]We have also estimated the relevant models using time dummies to control for seasonal effects. The estimation results are qualitatively similar in all cases and are available upon request.
[9]However, when we estimate the econometric model with SYS methodology, the estimated parameter of electricity is almost 1 indicating unit elasticity of demand.

The effect of total expenditure (M) is positive and highly statistical significant under all econometric methods employed. The highest effect of total expenditure on wholesale price of pasteurized milk is under SYS-GMM methodology where the estimated coefficient is marginal above one. The estimated coefficients of all other methodologies is below one, but always positive. The effect of the number of producers on wholesale price of milk is negative and highly statistical significant under SYS and DIF GMM econometric methods.[10] This result indicates that the higher the number of milk suppliers the lower the wholesale price of pasteurized milk.

When we introduce the effect of market deregulation on pasteurized milk (binomial variable D7) interesting results emerge. First and most interesting, the effect of deregulation on wholesale prices is positive and highly statistical significant except from the estimated coefficient of DOLS methodology, which is negative, but not statistical significant. This result coincides with the result we get from Figure 3 above, in which eight months after the deregulation the average wholesale price of pasteurized milk in Thessaloniki and Larissa exhibits an increase of about 0.06% and 0.24% correspondingly. Besides, the demand for pasteurized milk in Greece continues to be price inelastic and the product under scrutiny is a normal good. The effect of the number of producers on wholesale price is positive and statistically significant, except from the empirical result under DOLS methodology, which is negative, but statistically insignificant. However, the positive effect of the number of producers on wholesale price of pasteurized milk from May 2014 onwards, that is, after the deregulation of the pasteurized milk sector in Greece, is minor since the highly significant effect under DIF-GMM methodology is marginal above zero (0,0181).

[10]Under DOLS methodology the estimated coefficient is positive, but statistically insignificant.

We get almost the same results when we introduce the effect of market openness on pasteurized milk ((binomial variable Dopen). Under all methods of estimation the effect of deregulation on wholesale prices is positive and highly statistical significant. Recall from the analysis of figure 3 above that from January 2017 until October 2017 the average wholesale price of pasteurized milk exhibits an increase in the top four regions (Thessaloniki, Larissa, Xanthi and Serres). The demand for pasteurized milk becomes more price inelastic and the effect of total expenditure on wholesale price of milk is less pronounced (the highly statistical significant estimation is 0,00813 under DOLS methodology) than before September 2015. Even though the effect of the number of producers on wholesale price is positive and statistically significant, it seems to be also less pronounced than before the full openness of the market of pasteurized milk since its highly statistical estimation under DOLS methodology is less (0,00277) than the equivalent estimation after the deregulation of the market on April 2014.

In Table 7 we present the results from SYS-GMM estimators under sample splitting. As in Table 6 the estimations of coefficient b_0 in eq.5 and 6 and the estimated coefficient of X_{t-j} in eq. 7 are always highly statistical significant and smaller than 1 for all the dependent variables employed within the 45 regions under scrutiny. This result strengthens the importance of the inclusion of the lagged dependent variable in the right hand side of eq. 5, 6 and 7. The demand for pasteurized milk becomes less price inelastic from September 2015 onwards, but its absolute value continues to be below one. Moreover, the pasteurized milk is always a normal good, while the effect of the increase of the number of producers on wholesale price is negative, but statistically insignificant for both periods under scrutiny (from May 2014 and September 2015 onwards). The estimated coefficient of the latter effect is the same and statistically significant for the whole period

(Jan. 2010 – October 2017) and the period until the full openness of the milk sector in Greece (Jan. 2010 – August 2015).

<Insert Table 7about here>

5.3 Robustness checks

Using simple OLS to estimate the cointegrating relation will lead to bias in the estimated coefficients unless all of the explanatory variables are strongly exogenous. Moreover, its standardized distribution is dependent on nuisance parameters that are linked with the dynamics underlying the data generating processes of variables (Katsoulacos et al, 2014).Furthermore, other OLS estimators that remove the endogeneity bias such as the Fully-Modified OLS (Pedroni, 2000) or the Dynamic OLS (Kao and Chiang, 2000) are inadequate for our data since they assume cross-section independence. As Pesaran and Smith (1995) point out, other traditional methods for estimating pooled models such as the Fixed Effects and the Instrumental Variables estimators proposed by Arellano and Bond (1991) can produce very misleading estimates of the average values of the parameters in dynamic panel data models unless the slope coefficients are in fact identical. Furthermore, the Arellano and Bond (1991) method performs well for $N > T$ which is not the case in our sample.

For the above reasons, we estimate an Autoregressive Distributed Lag (ARDL) model for panel data to examine the long-run equilibrium and the short run dynamics of our models, and specifically the Pesaran and Smith (1995) Mean Group (MG) estimator. In order to account for the presence of nonstationarity and cross-section dependencewe employed two other mean group estimators such as the Common Correlated Effects (CCE) and Augmented Mean Group estimators (AMG). The benefits of using this approach is that the latter account strongly for the existence of non-stationarity and cross-section dependence in the panel sample while eliminate the asymptotical bias in the estimators due to the endogeneity of the regressors (Pesaran, 2015).

In Table 8 we present the results from Mean group estimation for market deregulation. As in Table 6 and 7 the estimations of coefficient b_0 in eq.5 and 6 the estimated coefficient of X_{t-j} in eq. 7 are always highly statistical significant and smaller than 1 for all the dependent variables employed within the 45 regions under scrutiny. This result strengthens the importance of the inclusion of the lagged dependent variable in the right hand side of eq. 5, 6 and 7. Under MG and CCE estimators market deregulation positively affects wholesale prices of pasteurized milk. The demand for pasteurized milk during the whole period under scrutiny is unit elastic and this result is the same with the one that emerges from the estimated coefficient under SYS – GMM methodology in Table 6. However, the deregulation of the market on April 2014 indicates that the market demand becomes price inelastic. Indeed, the absolute values of demand elasticity under all the empirical models employed are lower than the corresponding values derived from the dynamic estimations.

<Insert Table 8about here>

The effect of the number of producers on wholesale price is positive and statistically significant, except from the empirical result under MG-CCE methodology, which is positive, but statistically insignificant. However, the positive effect of the number of producers on wholesale price of pasteurized milk from May 2014 onwards, is minor since the highly significant effect is marginal above zero (0,016). We get the same results when we introduce the effect of market openness on pasteurized milk (binomial variable Dopen) in our analysis.[11]

[11]However, the majority of the empirical results from the introduction of binomial variable Dopen are statistical insignificant and therefore their presentation here is not useful for our analysis.

<Insert Table 9about here>

6. Conclusions and policy implications

The scope of this paper is to investigate the impact of market deregulation on the competitiveness of raw milk producers in Greece along the suggested lines of OECD (OECD, 2014). The study uses a dynamic panel data approach, to assess changes in the relative competitiveness of milk producers as a result of certain deregulation policies imposed by the Greek government in two phases (May 2014 and September 2015). For this purpose we use micro level data drawn from the 45 Greek regions spanning the period from January 2010 to October 2017.

The empirical results of the study reveal that the effect of market deregulation on April 2014 and its subsequent full openness on August 2015 positively affect the wholesale prices of pasteurized milk. The inverse demand function of pasteurized milk in Greece is price inelastic. Total expenditure positively affects the price of pasteurized milk, while the absolute value of its elasticity is below unity indicating that elasticity is income inelastic. Based on our empirical findings, it is argued that the higher the number of milk suppliers the lower the wholesale price of pasteurized milk. The demand for pasteurized milk during the deregulated period stated on April 2014 continues to be price inelastic, while after the full openness of the market (May 2015) the demand seems to be more price inelastic than before. Besides, total expenditure for pasteurised milk in Greece becomes more inelastic after the abolishment of regulation.

When we split the sample into different sub-samples (before and after the deregulation and the full openness of the market), the empirical results reveal that the demand for pasteurized milk becomes less price inelastic from September 2015 onwards, but its absolute value is estimated bellow unity. Moreover, demand continues to be income inelastic. The same result is confirmed when we account for the presence of nonstationarity and cross-section dependence, while demand is more income elastic, but bellow unity. The majority of the empirical findings indicate that consumers, continue to buy pasteurized milk, even in the cases where price increases. This result reflects the importance of pasteurized milk in the daily consumption of Greek consumers and signify that suppliers of pasteurised milk continue to possess market power against their customers from downstream market (i.e. the market for packaging and distribution of pasteurised milk).

Deregulation and full openness of the milk market firstly appeared in 2014 and 2015 enhances the level of competition in the local market. Specifically, we claim that prices fell down since local suppliers (farmers) face competitive pressure from other sources of competition, such as imports from abroad. The abolition of restrictive shelf life of pasteurised milk offers the opportunity of foreign suppliers to compete Greek local farmers more easily and effective than before. Direct imports lower the cost of raw materials and therefore the transportation cost of the pasteurised milk supplied into the Greek market. Therefore, competition is further enhanced since retailers, after the deregulation and the full openness of the market, import cheaper pasteurised milk from abroad. This result lends support to the increase of the competition intensity among the market players (i.e suppliers), leading to lower prices for the consumers. The substitution of part of Greek pasteurised milk with imported milk from abroad, leads to cost reductions and finally to lower consumer prices.

The overall effect on pasteurised milk prices is negative as a result of liberalisation. On the one hand there is a positive pressure on prices due to the fact that Greek suppliers continue to possess market power against their customers and their competitors from abroad. On the other hand, deregulation and full openness of the local market for pasteurised milk drive wholesale prices down in favour of final consumers.

Greek government must continue to provide incentives in order to attract more imports of milk from abroad. This policy needs to take into account the strategic development of Greek suppliers. Along these lines, policy makers and government officials should provide the local suppliers incentives to increase their productivity by investing in new technology and developing new business strategies. For example R&D investments will boost effectiveness among farmers and enable them to tackle efficiently the competition generated from abroad.

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APPENDIX

A DYNAMIC GMM ESTIMATORS

The GMM estimators rely on moments of the form:

$$h(\beta) = \sum_{i=1}^{N} h_i(\beta) = \sum_{i=1}^{N} \Psi_i' \, u_i' \, (\beta) \tag{A.1}$$

where Ψ_i is a $T_i xp$ matrix of instruments for cross section i and $u_i(\beta) = \left(Y_i - f\left(X_{it}, \beta\right)\right)$.

Specifically, GMM minimizes the following quadratic form with respect to β

$$M(\beta) = \left(\sum_{i=1}^{N} \Psi_i' \, u_i(\beta)\right) W \left(\sum_{i=1}^{N} \Psi_i' \, u_i(\beta)\right) = \zeta(\beta)' W \zeta(\beta) \tag{A.2}$$

where W is a pxp weighting matrix.

The coefficient covariance matrix is estimated as

$$V\left(\hat{\beta}\right) = \left(G'WG\right)^{-1} \left(G' \, W \Xi W G\right) \left(G'WG\right)^{-1} \tag{A.3}$$

Where Ξ is estimated as

$$E\left(\zeta_i(\beta)\zeta_i(\beta)'\right) = E\left(\Psi_i' \, u_i(\beta) u_i(\beta)' \, \Psi_i\right) \tag{A.4}$$

And G is a $T_i xk$ matrix given as:

$$G(\beta) = \left(-\sum_{i=1}^{N} \Psi_i' \nabla f_i(\beta)\right) \tag{A.5}$$

The weighting of matrix W may be calculated using the White robust covariances the coefficient covariance estimates are given as

$$\left(\frac{M^*}{M^*-k^*}\right)\left(\sum_t X_t' \, X_t\right)^{-1}\left(\sum_t X_t' \, \hat{u}_t \hat{u}_t' \, X_t\right)\left(\sum_t X_t' \, X_t\right)^{-1} \tag{A.6}$$

The first parenthesis is an adjustment to the degrees of freedom relying on the total number of observations; M^* is the total number of stacked observations and k^* the number of estimated parameters. The general form of the equation estimated with panel data dynamic models is one with individual effects like the following:

$$Y_{it} = \lambda_t + \eta_i + \sum_{k=1}^{p} \alpha_k Y_{i(t-k)} + \beta'(L) X_{it} + v_{it} \qquad \text{i = 1,2,...,T} \tag{A.7}$$

where λ_t and η_i correspond to specific and individual effects, X_{it} is a vector of explanatory variables, β(L) is a vector of associated polynomials in the lag operator and q is the maximum lag length. Identification of the model requires restrictions on the serial correlation of the error term v_{it} and on the properties of the independent variables X_{it} allowing only for MA or white noise errors. If the error term was originally autoregressive, the model is transformed. The weighting of matrix W may be calculated using the White robust covariance the coefficient covariance estimates are given as:

$$\left(\frac{M^*}{M^*-k^*}\right)\left(\sum_t X_t' \, X_t\right)^{-1}\left(\sum_t X_t' \, \hat{u}_t \hat{u}_t' \, X_t\right)\left(\sum_t X_t' \, X_t\right)^{-1} \tag{A.8}$$

The first parenthesis is an adjustment to the degrees of freedom relying on the total number of observations; M^* is the total number of stacked observations and k^* the number of estimated parameters. Orthogonal deviations as proposed by Arellano and Bond (1988) express each observation as the deviation from the average of future observations in the sample and weight each deviation to standardize the variance:

$$x_{it}^{*} = \left[x_{it} - (x_{i(t+1)} + ... + x_{iT}) / (T-t)\right]\sqrt{(T-t)} / \sqrt{T-t+1} \quad \text{t=1,...T-1} \tag{A.9}$$

The (T_i –q) equations for individual unit i can be written as:

$$Y_i = \delta w_i + d_i \eta_i + v_i \tag{A.10}$$

where δ is a parameter vector including α_κ's, β's and λ' s; and w_i is a data matrix containing the time series of the lagged endogenous variables, the x' s, and the time dummies. d_i is a (T_i-q) x1 vector of ones. Following Arellano and Bond (1998), linear GMM estimators of δ may be computed by the following expression:

$$\widehat{\delta} = \left[\left(\sum_i w_i^{*\prime} Z_i\right) \cdot \frac{1}{\frac{1}{N}\sum_i Z_i' H_i Z_i} \cdot \left(\sum Z_i' w_i^{*}\right)\right]^{-1} \cdot \left(\sum_i w_i^{*\prime} Z_i\right) \frac{1}{\frac{1}{N}\sum_i Z_i' H_i Z_i} \left(\sum_i Z_i' Y_t^{*}\right) \tag{A.11}$$

where w_i^* and Y_i^* denote some transformation of w_i and Y_i such as first differences, orthogonal deviations or levels. Z_i is the matrix of instrumental variables and H_i is an individual specific weighting matrix. We may have one-step estimates, which use some known matrix as the choice for H_i. For a first - difference procedure, the one-step estimator uses H_i,while for orthogonal deviations or for a levels procedure the one-step estimator sets H_i to an identity matrix. If the v_{it} are heteroskedastic, a two-step estimator is used.

B. MEAN GROUP ESTIMATORS

Consider the following multifactor residual model:

$$Y_{jt} = a_j + b_j X_{jt} + e_{jt} \tag{B.1}$$

wherejt is the jth cross section observation at time t, for t = 1,2,.........T, j = 1,2, ...N.

$$\mathrm{e_{jt}} = \mathrm{a_{1j}} + \lambda_{\mathrm{j}}^{'}\Phi_{\mathrm{t}} + \mathrm{u_{jt}} \tag{B.2}$$

whereΦ_t is a mx1 vector of unobserved common factors and $\lambda_j^{'}$ a heterogeneous factor loading. Both Y_{jt} and X_{jt} are observables and $b_j$are country-specific slopes on the observable regressors. a_{1j}capture time-invariant heterogeneity across groups, i.e. countries, and $\lambda_j^{'}\Phi_t$ capture time-variant heterogeneity and cross - section dependence.

$$X_{jt} = a_{1j} + \lambda_j^{'}\Phi_t + c_j\Gamma_t + \varepsilon_{jt} \tag{B.3}$$

$\Phi_t$andΓ_t are mx1 vectors of unobserved common factors and $\lambda_j^{'}$ and c_j are heterogeneous factors loading Even though Φ_t is present in equations (B.1) and (B.2), we assume that the regressors are not driven by the same common factors as the observables.

Pesaran (2006) adopts the following multifactor residual model:

$$Y_{jt} = a_j + b_j X + e_{jt} \tag{B.4}$$

where X is a matrix of explanatory variables, *jt* is the *jth* cross section observation at time t, for t = 1,2,.........T, *j* = 1,2, ...N and $e_{jt} = \lambda_j'\Phi_t + u_{jt}$. $\Phi_t$is a mx1 vector of unobserved common factors and λ_j' a heterogeneous factor loading.

Even though Pesaran (2006) considers the case of weakly stationary factors, Kapetanios et al. (2011) show that Pesaran's CCE approach continues to yield consistent estimation and valid inference even when common factors are unit root processes. To deal with the residual cross section dependence Pesaran(2006) uses cross sectional averagesas observable proxies for commonfactorsΦ_t. The cross sectional averages are $\bar{Y}_t = \frac{1}{N}\sum_{j=1}^{N} Y_{jt}$and $\bar{X}_t = \frac{1}{N}\sum_{j=1}^{N} X_{jt}$.

Given the previous means Equation (B.4) becomes:

$$Y_{jt} = a_j + b_j X + c\bar{Y}_t + d\bar{X}_t + e_{jt} \quad \text{(B.5)}$$

Estimation of equation (B.5) with OLS provide, under strict exogeneity, the OLS estimators $\hat{B}_j$ of the individual specific slope coefficients (b) as the CCE estimators. The Common Correlated Effects Mean Group (CCEMG) estimator is the average of the individual CCE estimators ($\hat{B} = \sum_{j=1}^{N} \hat{B}_j$) and follows asymptotically the standard normal distribution. Pesaran (2006) and Kapetanios et al. (2011) show that the CCE estimators have the correct size, and they haveshownthatsmall-sample properties of the CCE estimators do not seem to be much affected by the residual serial correlation of the errors.The dynamic version of equation (B.5) is given as follows:

$$Y_{jt} = a_j + \beta_j Y_{j,t-1} + b_j X + e_{jt},\, e_{jt} = \lambda_j' \Phi_t + u_{jt} \quad \text{(B.6)}$$

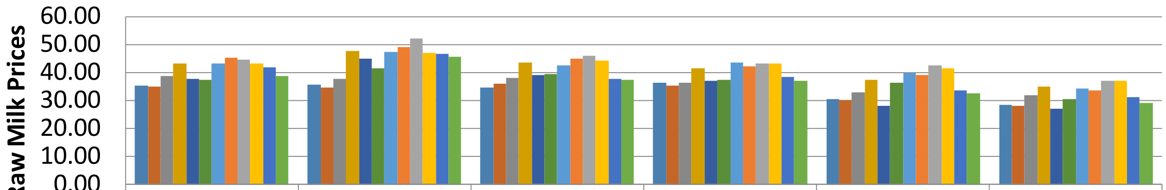
Since the lagged dependent variable is not strictly exogenous the OLS estimator of the above equation becomes inconsistent. Chudik and Pesaran (2015) show that the OLS estimator gains consistency if$\sqrt[3]{T}$ are added in equation (13). Therefore, the latter becomes:

$$Y_{jt} = a_j + \beta_j Y_{j,t-1} + b_j X + \sum_{J=0}^{\sqrt[3]{T}} \delta_{jJ}' \bar{z}_{t-I} + e_{jt},\, e_{jt} = \lambda_j' \Phi_t + u_{jt} \quad \text{(B.7)}$$

where $\bar{z}_t = (\bar{Y}_t,\ \bar{Y}_{t-1}, \bar{X}_t)$. The Mean Group Estimates are$\frac{1}{N}\sum_{j=1}^{N} \hat{\pi}_j$, where $\hat{\pi}_j = (\beta_j, b_j)$.

List of Figures and Tables

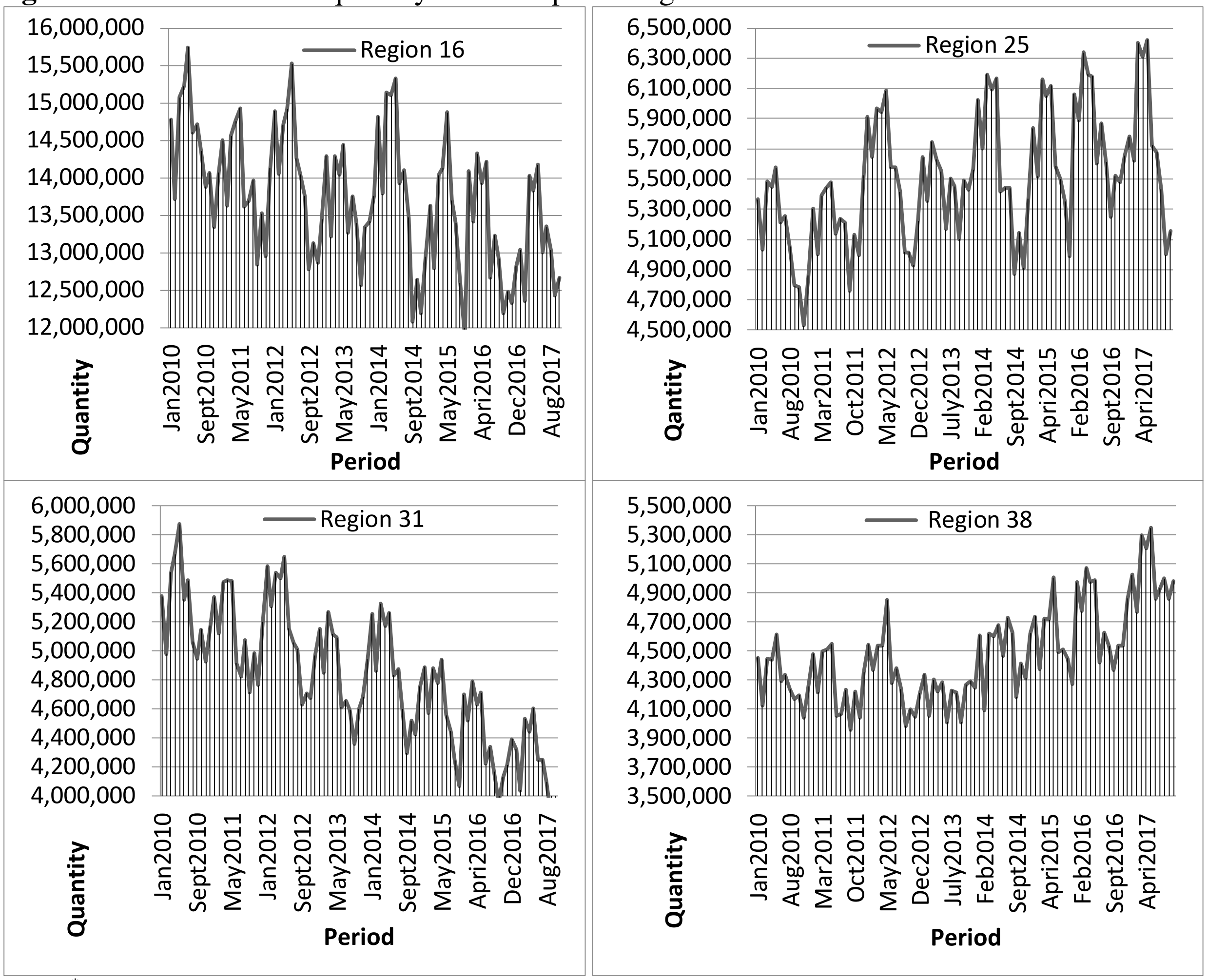
Figure 1. Selling raw cow's milkprices in selected EU countries (2005-2016)



	Greece	Malta	Finland	Italy	Sweden	EU Average
2005	35.28	35.74	34.51	36.18	30.50	28.24
2006	34.95	34.41	35.83	35.10	30.16	27.88
2007	38.67	37.51	37.91	36.41	32.87	31.98
2008	43.20	47.50	43.49	41.47	37.23	34.90
2009	37.69	44.86	38.94	37.07	28.16	27.16
2010	37.33	41.34	39.41	37.46	36.17	30.61
2011	43.16	47.18	42.63	43.38	39.65	34.36
2012	45.08	49.00	44.91	42.26	38.95	33.48
2013	44.42	52.19	45.90	43.19	42.42	36.93
2014	43.16	46.84	44.27	43.22	41.33	36.84
2015	41.84	46.66	37.58	38.23	33.50	31.05
2016	38.65	45.61	37.20	37.13	32.56	29.20

Notes: The absolute prices give information on the levels of the producer prices of raw cow's milk. Prices are net of VAT.
Source: Authors' elaboration (Data from Eurostat: http://ec.europa.eu/eurostat/tgm/table.do?tab=table&init=1&plugin=1&language=en&pcode=tag00070

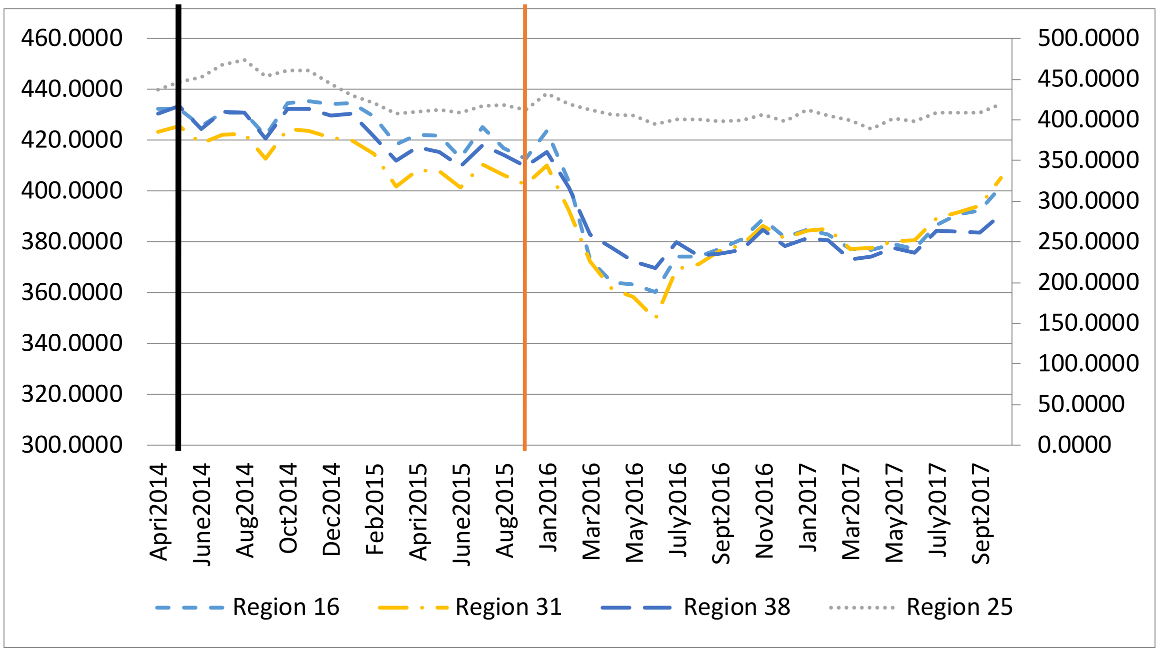
Figure 2. Consumed milk quantity for the top four regions in Greece



Notes: * in logs (measured in tones); Region 16: Thessaloniki, Region 25: Larissa, Region 31: Xanthi, Region 38: Serres

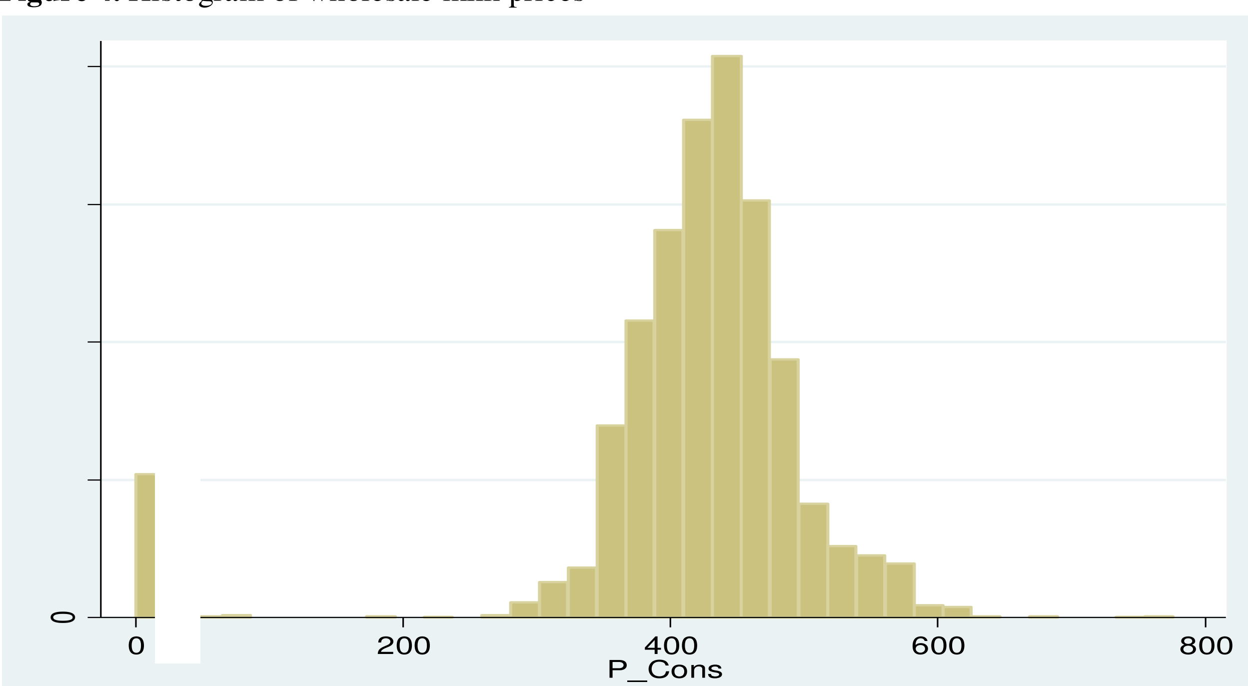
Source: Authors' calculations based on data from the Hellenic Agricultural Organization (HAO) "*DIMITRA*"

Figure 3. Prices of pasteurised milk for the big "*four*" regions in Greece



Notes: Prices are in real terms and reported here in logarithmic scale. Region 16: Thessaloniki, Region 25: Larissa, Region 31: Xanthi, Region 38: Serres. The black vertical line indicates the extension of the shelf life of milk from five to seven days (May 2014), while the orange vertical line represents the full market openness (September 2015).
Source: Authors' calculations based on data from the Hellenic Agricultural Organization (HAO) "*DIMITRA*"

Figure 4. Histogram of wholesale milk prices



Notes: P_Cons denotes the log wholesale milk constant prices.
Source: Authors' calculations based on data from the Hellenic Agricultural Organization (HAO) "*DIMITRA*"

Table 1: Legislation of Pasteurized milk in Greece (1959–2015)

Official Gazette	Maximum shelf life of pasteurized milk	Types of milk
No. 89/16.5.1959*(Royal Decree4)	2 days [article 12(6)]	Fresh and Pasteurised [article 2]
No. 113/4.5.1981* (Presidential Decree430)	3 days [article 5(6)]	-
No. 46/16.3.1988* (Presidential Decree 104)	4 days [article 1]	-
No. 115/9.6.1999*(Presidential Decree 113)	5 days [article 1]	Fresh/Pasteurised and High Pasteurised (HPM)** [article 1]
No. 85/7.4.2014*** (Law N. 4254)	7 days [article 2(a)]	Pasteurised**** and High Pasteurised (HPM)** [article 2(a)]
No. 94/14.8.2015***** (Law N. 4336)	at the discretion of the manufacturer	Pasteurised and High Pasteurised (HPM)** [article 1]

Notes: *On Veterinary and hygiene check of milk, ** HPM is a unique Greek term, not to be confused with Ultra High Temperature (UHT) pasteurised milk. The maximum shelf life is at the discretion of the manufacturer, ***Sub-paragraph F8: Removal of barriers to Competition in milk market – Regulation of Dairy products, ****The term fresh milk is not used any more on the packaging of the final product,*****Part B: Article 2, Sub-paragraph A3: Regulation of other issues

Table 2: Summary statistics

Variables	Observations	Mean	Standard Deviation	Minimum	Maximum
Log (P)	4,040	6.061	0.160	3.319	6.654
Log (Q)	4,040	12.50	2.139	3.798	16.57
Log (M)	4,040	11.65	2.122	3.304	15.77
Log (Firms)	4,040	3.002	1.719	0.000	6.301
D7	4,230	0.447	0.497	0.000	1.000
Dopen	3,780	3.920	4.997	0.000	1.000
D7 × log(Q)	4,230	0.277	0.447	0.000	1.000
D7 × log (M)	4,040	5.303	6.305	0.000	16.55
D7 × log (Firms)	4,040	4.939	5.887	0.000	15.70
Dopen × log (Q)	4,040	3.121	5.492	0.000	16.48
Dopen × log (M)	4,040	2.901	5.116	0.000	15.59
Dopen × log (Firms)	4,040	0.726	1.506	0.000	5.930

Notes: Authors' estimations. P stands for the wholesale milk prices per region deflated by the Harmonised Consumer Price Index (2015=100). Q is the milk quantity demanded per region in tones. M is the income expenditure per region deflated by the Harmonised Consumer Price Index (2015=100). The variable Firms stands for the total number of milk producers per region. D7 is the dummy variable taking the value one when the introduction of the seven day fresh milk was legally implemented and zero otherwise. Dopen denotes the dummy variable taking the value one when market openness was introduced and zero otherwise. Standard errors are 5% on each side generated by Monte-Carlo with 500 repetitions.

Table 3: Cross-section dependence test

Variable	CD test	P-value	Correlation	Absolute (correlation)
Log (P)	241.57***	0.000	0.570	0.571
Log (Q)	414.20***	0.000	0.977	0.977
Log (M)	412.09***	0.000	0.972	0.972
Log (Firms)	417.86***	0.000	0.986	0.986

Notes: Under the null hypothesis of cross-sectional independence the CD statistic is distributed as a two-tailed standard normal. Results are based on the test of Pesaran (2004). The p-values are for a one-sided test based on the normal distribution. Correlation and Absolute (correlation) are the average (absolute) value of the off-diagonal elements of the cross-sectional correlation matrix of residuals. P stands for the wholesale milk prices per region deflated by the Harmonised Consumer Price Index (2015=100). Q is the milk quantity demanded per region in tones. M is the income expenditure per region deflated by the Harmonised Consumer Price Index (2015=100). The variable Firms stands for the total number of milk producers per region. All variables are expressed in natural logarithms. Significant at ***1% level of statistical significance.

Table 4: Panel unit root tests

Variable	Fisher type ADF	Pesaran ADF
Log (P)	1,122***	-30.097***
Log (Q)	1,580***	-26.005***
Log (M)	1,626***	-25.213***
Log (Firms)	1,910***	-30.970***

Notes: The number of lags has been set to one according to BIC. The Augmented Dickey Fuller test is used rather than Phillips-Perron test (see Phillips and Perron, 1988). The null hypothesis assumes that the variable contains unit root. P stands for the wholesale milk prices per region deflated by the Harmonised Consumer Price Index (2015=100). Q is the milk quantity demanded per region in tones. M is the income expenditure per region deflated by the Harmonised Consumer Price Index (2015=100). The variable Firms stands for the total number of milk producers per region. Significant at ***1%.

Table 5: Pedronipanel cointegration test results

Within Dimension Test Statistics		Between Dimension Test Statistics	
Panel v-statistic	8.203*** (0.000)	Group ρ-statistic	-53.8*** (0.000)
Panel ρ-statistic	-55.96*** (0.000)	Group PP-statistic	-69.43*** (0.000)
Panel PP-statistic	-63.44*** (0.000)	Group ADF-statistic	-66.72*** (0.000)
Panel ADF-statistic	-61.45*** (0.000)		

Notes: Of the seven tests, the panel v-statistic is a one-sided test where large positive values reject the null hypothesis of no cointegration, whereas large negative values for the remaining test statistics reject the null hypothesis of no cointegration. The seven tests follow asymptotically a standard normal distribution. The first three non-parametric tests correct for serial correlation. These tests comprise of a non-parametric variance ratio statistic, a test analogous to the Phillips and Perron (PP) rho-statistic and, a test analogous to the PP t-statistic. These panel statistics are based on pooling the data along the within dimension of the panel. The fourth parametric test is similar to the ADF-type test. The other three panel cointegration statistics are based on a group mean approach. The first two of the group-mean panel cointegration statistics are panel versions of the Phillips and Perron rho and t-statistics, respectively. The third is a group-mean ADF test analogous to the Im, Pesaran, and Shin (2003) panel unit root test (Katsoulacos, et al, 2014). The critical values were created using a bootstrapping method. Significant at ***1% and **5% respectively.

Table 6: Dynamic panel estimators under different methodologies

Control variables	(1) SYS-GMM	(2) SYS-GMM	(3) DIF-GMM	(4) DIF-GMM	(5) DOLS	(6) DOLS
LogP(-1)	0.606*** (0.0982)	0.632*** (0.0930)	0.0233*** (0.00491)	0.0128*** (0.00431)	1.304*** (0.000802)	1.155*** (0.000815)
LogP(-2)	0.257*** (0.0889)	0.269*** (0.0906)	-0.0768*** (0.0142)	-0.0707*** (0.0132)	-1.192*** (0.000828)	-1.081*** (0.000853)
LogP(-3)	-	-	0.0398*** (0.00702)	0.0395*** (0.00686)	0.393*** (0.000679)	0.382*** (0.000699)
Log (Q)	-0.0847** (0.0383)	-0.154** (0.0753)	-1.017*** (0.00700)	-1.026*** (0.00799)	-0.497*** (0.000812)	-0.545*** (0.000758)
Log(M)	0.191*** (0.0518)	0.240*** (0.0815)	1.024*** (0.00645)	1.028*** (0.00751)	0.497*** (0.000761)	0.545*** (0.000691)
Log(Firms)	-0.113** (0.0574)	-0.0916** (0.0426)	-0.0174*** (0.00373)	-0.0137*** (0.00293)	9.75e-05 (0.000451)	6.78e-05 (0.000450)
D7	1.787** (0.751)	-	0.126*** (0.0262)	-	-0.00303 (0.00274)	-
D7 × log(Q)	-0.617*** (0.143)	-	-0.0367** (0.0156)	-	0.00200* (0.00104)	-
D7 × log (M)	0.459*** (0.0881)	-	0.0240 (0.0154)	-	-0.00191* (0.000999)	-
D7 × log (Firms)	0.198* (0.105)	-	0.0181*** (0.00403)	-	-1.45e-05 (0.000302)	-
Dopen	-	1.356** (0.594)	-	0.0293 (0.0220)	-	0.0101*** (0.00286)
Dopen × log (Q)	-	-0.315** (0.142)	-	0.0117 (0.0123)	-	-0.00887*** (0.00208)
Dopen × log (M)	-	0.173 (0.157)	-	-0.0182 (0.0120)	-	0.00813*** (0.00211)
Dopen × log (Firms)	-	0.194* (0.106)	-	0.00819* (0.00424)	-	0.00277*** (0.000325)
Diagnostics						
Observations	3,935	3,935	3,790	3,790	2,698	2,546
Regions	45	45	45	45	45	45
Fixed effects	Yes	Yes	Yes	Yes	No	No

Wald/F test	1.33e+06*** [0.000]	1.21e+06*** [0.000]	25,118.35*** [0.000]	30,667.88*** [0.000]	7.77e+06*** [0.000]	6.64e+06*** [0.000]
AR(1)	-1.61* [0.108]	-1.49 [0.137]	-5.27*** [0.000]	-5.33*** [0.000]	-	-
AR(2)	-0.27 [0.784]	-0.28 [0.782]	-4.46*** [0.000]	-5.47*** [0.000]	-	-
Hansen test	44.97 [1.000]	44.89 [1.000]	44.98 [1.000]	44.99 [1.000]	-	-

Notes: The number of lags has been determined according to BIC. P stands for the wholesale milk prices per region deflated by the Harmonised Consumer Price Index (2015=100). Q is the milk quantity demanded per region in tones. M is the income expenditure per region deflated by the Harmonised Consumer Price Index (2015=100). The variable Firms stands for the total number of milk producers per region. D7 is the dummy variable taking the value one when the introduction of the seven day fresh milk was legally implemented and zero otherwise. Dopen denotes the dummy variable taking the value one when market openness was introduced and zero otherwise. SYS-GMM is the system GMM estimator and DIF-GMM denotes the difference GMM estimator. DOLS denote the dynamic OLS estimators. Robust standard errors are in parentheses. The numbers in square brackets denote the p-values. AR(1) and AR(2) are tests for first and second order serial autocorrelation. F and Wald tests denote the joint statistical significance of all the covariates. Hansen denotes the test of over identifying restrictions of the instruments. Significant at ***1%, **5% and *10% respectively.

Table 7: System GMM estimators under sample splitting

Control variables	(1) *[Jan 2010-Oct 2017]*	(2) *[Jan 2010 – Apr 2014]*	(3) *[May 2014-Oct 2017]*	(4) *[Jan 2010 – Aug 2015]*	(5) *[Sep 2015-Oct 2017]*
LogP(-1)	0.655*** (0.0939)	0.538*** (0.0279)	0.554*** (0.0770)	0.638*** (0.0944)	0.748*** (0.162)
LogP(-2)	0.275*** (0.0913)	0.338*** (0.0949)	0.405*** (0.0183)	0.277*** (0.0939)	0.283* (0.154)
Log(Q)	-0.175** (0.0785)	-0.120*** (0.0272)	-0.461*** (0.0304)	-0.163** (0.0731)	-0.554*** (0.102)
Log(M)	0.239*** (0.0868)	0.215*** (0.0553)	0.529*** (0.0420)	0.236*** (0.0792)	0.580*** (0.0921)
Log(Firms)	-0.0592* (0.0311)	-0.0908 (0.0568)	-0.0533 (0.0457)	-0.0689** (0.0349)	-0.00365 (0.0303)
Diagnostics					
Observations	3,935	2,307	1,648	3,012	923
Regions	45	45	45	45	45
Fixed effects	Yes	Yes	Yes	Yes	Yes
Wald/F test	295,028.82*** [0.000]	225826.48*** [0.000]	284,506.78*** [0.000]	231,312.94*** [0.000]	441,989.55*** [0.000]
AR(1)	-1.48 [0.140]	-1.50 [0.134]	-1.03 [0.303]	-1.45 [0.148]	-0.99 [0.322]
AR(2)	-0.28 [0.777]	-0.68 [0.499]	-1.38 [0.167]	-0.26 [0.798]	-1.17 [0.242]
Hansen test	44.96 [1.000]	44.90 [1.000]	44.94 [1.000]	44.97 [1.000]	43.99 [0.233]

Notes: The number of lags has been determined according to BIC. P stands for the wholesale milk prices per region deflated by the Harmonised Consumer Price Index (2015=100). Q is the milk quantity demanded per region in tones. M is the income expenditure per region deflated by the Harmonised Consumer Price Index (2015=100). The variable Firms stands for the total number of milk producers per region. Robust standard errors are in parentheses. The numbers in square brackets denote the p-values. AR(1) and AR(2) are tests for first and second order serial autocorrelation. F and Wald tests denote the joint statistical significance of all the covariates. Hansen denotes the test of over identifying restrictions of the instruments. Significant at ***1%, **5% and *10% respectively.

Table 8: Mean group estimation resultsfor market deregulation

Control variables	(1) MG	(2) MG-T	(3) MG-ORM	(4) MG-CCE	(5) MG-AUG	(6) MG-AUGT	(7) MG-AUGCDP
Constant	7.111*** (0.0679)	7.155*** (0.0787)	7.116*** (0.0767)	4.35e-05*** (7.91e-06)	6.907*** (0.000744)	6.907*** (0.00147)	6.906*** (0.00146)
Trend	-	-2.19e-05 (3.30e-05)	-5.99e-06 (3.31e-05)	-	-	1.05e-05*** (1.48e-06)	1.05e-05*** (1.48e-06)
Log P(-1)	-0.0157** (0.00688)	-0.0163** (0.00708)	-0.0157** (0.00665)	6.94e-07 (5.40e-07)	-9.01e-05** (3.91e-05)	-0.000181** (9.15e-05)	-0.000180** (8.99e-05)
Log P(-2)	-0.0456*** (0.00991)	-0.0461*** (0.0100)	-0.0425*** (0.0106)	6.67e-07 (6.45e-07)	1.88e-05 (4.42e-05)	-0.000147* (8.55e-05)	-0.000140* (7.99e-05)
Log P(-3)	0.0244*** (0.00681)	0.0237*** (0.00673)	0.0224*** (0.00722)	-6.47e-07* (4.09e-07)	-	-0.000129 (8.06e-05)	-0.000133 (8.10e-05)
Log (Q)	-1.031*** (0.0101)	-1.030*** (0.0109)	-1.025*** (0.0109)	-1.000*** (4.23e-06)	-1.000*** (9.76e-05)	-1.000*** (0.000404)	-1.000*** (0.000407)
Log (M)	1.040*** (0.0104)	1.039*** (0.0111)	1.031*** (0.0110)	1.000*** (4.22e-06)	1.000*** (9.97e-05)	1.000*** (0.000398)	1.000*** (0.000401)
Log (Firms)	-0.0146*** (0.00422)	-0.0208*** (0.00682)	-0.0170*** (0.00491)	-3.04e-07 (2.87e-07)	0.000570 (0.000578)	0.00128* (0.000772)	0.00128* (0.000775)
D7	0.0124*** (0.00367)	0.00794* (0.00454)	0.0525* (0.0294)	1.04e-05** (4.16e-06)	0.0125*** (0.00124)	-0.000194 (0.00253)	-0.000232 (0.00253)
D7 × log(Q)	-0.0211*** (0.00663)	-0.0227** (0.00964)	-0.0135* (0.00842)	1.02e-06 (9.26e-07)	0.00264*** (0.000501)	0.00522*** (0.000958)	0.00522*** (0.000961)
D7 × log (M)	0.0158** (0.00657)	0.0173* (0.00953)	0.00897 (0.00810)	-1.24e-06 (9.53e-07)	-0.00257*** (0.000499)	-0.00511*** (0.000969)	-0.00512*** (0.000973)
D7 × log (Firms)	0.0160*** (0.00531)	0.0160*** (0.00618)	0.0165*** (0.00522)	3.02e-07 (4.81e-07)	-0.000915* (0.000595)	-0.00183** (0.000864)	-0.00183** (0.000867)
Diagnostics							
Observations	3,884	3,884	3,884	3,884	3,935	3,884	3,884
Wald test	233,535.67*** [0.000]	96,832.87*** [0.000]	17,731.91*** [0.000]	7.27e+10*** [0.000]	1.33e+08*** [0.000]	9.68e+06*** [0.000]	9.15e+06*** [0.000]
RMSE	0.0115	0.0114	0.0114	0.0000	0.0001	0.0001	0.0001

Notes: The number of lags has been determined according to BIC. P stands for the wholesale milk prices per region deflated by the Harmonised Consumer Price Index (2015=100). Q is the milk quantity demanded per region in tones. M is the income expenditure per region deflated by the Harmonised Consumer Price Index (2015=100). The variable Firms stands for the total number of milk producers per region. D7 is the dummy variable taking the value one when the introduction of the seven day fresh milk was legally implemented and zero otherwise. MG, MG-T, and MG-ORM stand for standard Mean Group (Pesaran and Smith, 1995), Group Mean with a linear trend and GroupMeanwith outlier robust means.MG, MG-T and MG-ORM assume cross section independence. MG-CCE refers to the Common Correlated Effects Mean Group estimation and inference method (Pesaran, 2006) and allows for cross sectional dependence.MG-AUG, MG-AUGT and MG-AUGCDP denote the Augmented Mean Group estimator (Eberhardt and Teal, 2010; Bond and Eberhardt, 2009), Augmented Mean Group with a linear trend and Augmented Mean Group characterised by a common dynamic process with a unit coefficient. RMSE stands for the Root Mean Squared Error. Robust standard errors are in parentheses. The numbers in square brackets denote the p-values. Significant at ***1%, **5% and *10% respectively.

Table 9: Mean group estimation results for market openness

Control variables	(1) MG	(2) MG-T	(3) MG-ORM	(4) MG-CCE	(5) MG-AUG	(6) MG-AUGT	(7) MG-AUGCDP
Constant	7.115*** (0.0612)	7.146*** (0.0691)	7.096*** (0.0695)	2.42e-05*** (6.58e-06)	6.900*** (0.00318)	6.907*** (0.000250)	6.907*** (0.000250)
Trend	-	6.91e-05*** (1.99e-05)	8.36e-05*** (1.77e-05)	-	-	7.12e-07*** (1.65e-07)	7.16e-07*** (1.65e-07)
Log P(-1)	-0.0161** (0.00650)	-0.0177*** (0.00678)	-0.0169** (0.00657)	9.23e-07* (5.28e-07)	0.00111** (0.000443)	5.49e-06 (1.29e-05)	4.87e-06 (1.30e-05)
Log P(-2)	-0.0433*** (0.0102)	-0.0440*** (0.0103)	-0.0386*** (0.0111)	-5.74e-07 (4.60e-07)	-	-2.42e-05* (1.58e-05)	-2.81e-05* (1.63e-05)
Log P(-3)	0.0251*** (0.00701)	0.0223*** (0.00659)	0.0217*** (0.00710)	-	-	-	-
Log (Q)	-1.033*** (0.00921)	-1.034*** (0.00978)	-1.023*** (0.00946)	-1.000*** (3.76e-06)	-0.999*** (0.000419)	-1.000*** (5.03e-05)	-1.000*** (5.47e-05)
Log (M)	1.039*** (0.00965)	1.041*** (0.0100)	1.025*** (0.00973)	1.000*** (3.76e-06)	0.999*** (0.000416)	1.000*** (4.97e-05)	1.000*** (5.41e-05)
Log (Firms)	-0.0111*** (0.00326)	-0.0115** (0.00482)	-0.00823** (0.00330)	-1.23e-07 (1.58e-07)	1.40e-05 (5.18e-05)	8.88e-05*** (3.34e-05)	8.82e-05*** (3.34e-05)
Dopen	-0.00705 (0.0505)	0.00532*** (0.0510)	0.0335*** (0.00338)	5.59e-06** (2.91e-06)	0.0415** (0.0206)	0.0125*** (0.000875)	0.0126*** (0.000876)
Dopen × log(Q)	0.0428*** (0.0115)	0.0361*** (0.0119)	0.0359*** (0.0116)	-5.02e-06*** (1.80e-06)	0.0199** (0.00931)	0.00106** (0.000464)	0.00107** (0.000464)
Dopen × log (M)	-0.0484*** (0.0121)	-0.0425*** (0.0122)	-0.0412*** (0.0112)	4.61e-06** (1.93e-06)	-0.0173* (0.00908)	-0.000965** (0.000463)	-0.000968** (0.000463)
Dopen × log (Firms)	0.00647 (0.00664)	0.00575 (0.00677)	0.000156 (0.00538)	4.30e-07 (6.37e-07)	-0.00137 (0.00314)	-0.000198 (0.000146)	-0.000196 (0.000146)
Diagnostics							
Observations	3,884	3,884	3,884	3,935	3,987	3,935	3,935
Wald/F test	222965.67*** [0.000]	2177,49.18*** [0.000]	22,846.54*** [0.000]	8.37e+10*** [0.000]	4.95e+08*** [0.000]	4.80e+08*** [0.000]	4.10e+08*** [0.000]
RMSE	0.0114	0.0114	0.0114	0.0000	0.0009	0.0000	0.0000

Notes: The number of lags has been determined according to BIC. P stands for the wholesale milk prices per region deflated by the Harmonised Consumer Price Index (2015=100). Q is the milk quantity demanded per region in tones. M is the income expenditure per region deflated by the Harmonised Consumer Price Index (2015=100). The variable Firms stands for the total number of milk producers per region. Dopen denotes the dummy variable taking the value one when market openness was introduced and zero otherwise. MG, MG-T, and MG-ORM stand for standard Mean Group (Pesaran and Smith, 1995), Group Mean with a linear trend and Group Mean with outlier robust means. MG, MG-T and MG-ORM assume cross section independence. MG-CCE refers to the Common Correlated Effects Mean Group estimation and inference method (Pesaran, 2006) and allows for cross sectional dependence. MG-AUG, MG-AUGT and MG-AUGCDP denote the Augmented Mean Group estimator (Eberhardt and Teal, 2010; Bond and Eberhardt, 2009), Augmented Mean Group with a linear trend and Augmented Mean Group characterised by a common dynamic process with a unit coefficient. RMSE stands for the Root Mean Squared Error. Robust standard errors are in parentheses. The numbers in square brackets denote the p-values. Significant at ***1%, **5% and *10% respectively.