Asymmetric Price Transmission in the Israeli Citrus Export Sector in the Aftermath of Liberalization

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Abstract:
The Israeli citrus export sector was liberalized in 1991 with the aim to increase citrus growers' income and to improve overall efficiency of the international citrus marketing channel. However, the former government export monopoly's activities were mainly taken over by four large companies accounting for over 90% of total Israeli citrus market exports. In addition, citrus exporters maintained the monopoly's consignment system, substantially limiting transparency on how the grower price is determined. This lead the government to intervene in the newly liberalized market by implementing a minimum price agreement in the 1994/95 season to protect citrus growers against exporters' abuse of market power.
In this paper we analyze whether market power was exerted by exporting companies over citrus growers in the form of asymmetric price transmission. Our study is unique in that it investigates vertical price transmission across international borders, i.e. in the context of Israeli grapefruit exports to France. We explicitly account for possible changes in exporters' pricing behaviour in the post-liberalization period.
We apply an error correction model (ECM) to disaggregated firm-level Israeli grower price and French import price data. An ECM is estimated individually for each of the major exporting companies within a seemingly unrelated regression (SUR) framework. We find asymmetric price transmission in the first years after liberalisation, but symmetry in the second half of the 1990s. Our results indicate that growers' losses due to asymmetry amounted to as much as 2.5% of their total revenues. Our results suggest that liberalization improved the efficiency of the Israeli citrus international marketing channel, but that this took time and was probably accelerated by government intervention.

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1 Introduction

Prior to 1991, Israeli fresh citrus fruits were exported exclusively by the governmental Citrus Marketing Board of Israel (CMBI), which was established in 1948. During the eighties, Israeli citrus export quantity and citrus grower prices decreased significantly, creating political pressure to abolish the CMBI's monopoly. The goal of liberalizing the Israeli citrus sector in 1991 was to increase the citrus growers' income and to strengthen the efficiency of the Israeli citrus export marketing channel by establishing competition between the exporting companies.

Despite liberalisation however, the citrus export market remained dominated by four companies: Agrexco, Mehadrin, Pardess and Tnuport. In the first 10 years after liberalization, these companies accounted for over 90% of all Israeli citrus exports. In contrast, Israel's citrus production is less concentrated with about 630 citrus growers accounting for roughly 80% of the citrus growing area. In addition, the consignment system of the former monopoly has been maintained, substantially limiting the transparency of grower price determination. Growers supply produce to exporting companies on consignment and are not provided with information on the grower price until after the produce has been sold in the export market. This induced the government, in 1994/95, to intervene in the newly liberalized market by establishing a minimum price agreement to protect growers against the abuse of market power by exporting companies. According to this agreement, exporters qualified for a government subsidy only if they signed a written, standardized contract with growers, guaranteeing a minimum grower price and stating the timetable of payments and conditions triggering additional payments to the growers (KACHEL, 2003).

This paper studies whether there is evidence that exporting companies have exerted market power over citrus growers by transmitting price changes on export markets asymmetrically. Market power is considered to be one of the primary potential causes of vertical asymmetric price transmission (MEYER AND VON CRAMON-TAUBADEL, 2004). Asymmetric vertical price transmission implies that price changes at one level of the marketing chain (in this case the export market for Israeli citrus)

that squeeze a firm's margin are transmitted faster and more completely to another level of the marketing stage (in this case the grower price in Israel) than price changes that stretch a firm's margin (VON CRAMON-TAUBADEL, 1998). As a result, the firm which transmits prices asymmetrically increases its profit temporarily.

This study is unique in two important regards. First, it investigates vertical price transmission across international borders[a]. In particular, we analyze vertical price transmission from the export market for Israeli grapefruit in France to Israeli grapefruit producers. With growing agricultural trade and the globalisation of agriculture, understanding international price transmission is of increasing importance. International agricultural trade is especially susceptible to the abuse of market power since price transparency is often limited due to restricted availability of price data, difference in currency units, etc.

Second, in contrast to most studies of vertical price transmission, this analysis is not restricted to aggregated data. We apply an error correction model (ECM) to disaggregated firm-level Israeli grower prices and French import price data. An ECM is estimated individually for each of the exporting companies within a seemingly unrelated regression (SUR) framework. This is supplemented by an estimation based on aggregated grower price data to cast further light on the issue of aggregation and its impact on the study of price transmission (VON CRAMON ET AL, 2006).

Our empirical model explicitly accounts for the hypothesis that the exporters' pricing behaviour may have changed in the post-liberalization market due to two common external factors – the enforcement of the minimum price agreement in 1994/95, and a substantial increase in sea transport costs in the 1990s. This is accomplished by allowing for a structural break in the cointegration regression. In addition, we distinguish a heterogeneous, volatile phase directly after liberalization from a more homogeneous, calm phase some years later, and estimate the ECM for those two regimes separately.

The analyses based on the disaggregated as well as aggregated data both find that exporters transmitted changes in the French import market to Israeli growers asymmetrically in the volatile

phase directly after liberalization, but symmetrically in the calm phase thereafter. Further, we find the measured asymmetry in price transmission to be economically significant. In particular, the growers' losses amounted up to 2.5% of growers' total revenue of one season. Overall, our study demonstrates that liberalization improved the efficiency of Israel's international citrus marketing channel, although this took time and was probably accelerated by government market intervention.

The paper is structured as follows. Section two explains characteristic features of the data set and how they are accounted for in the empirical specification. The methodological background of this study is explained in section 3 and empirical model results are presented in section 4. Chapter 5 concludes.

2 Dataset and critical issues

The analysis is based on weekly firm-level grower price data[b] from each of the four major Israeli exporting companies, and the corresponding French import price for red 'Sunrise' grapefruits in the seasons 1991/92 to 1999/00[c]. Both data series are stated in New Israeli Shekel (NIS) per ton and deflated with the Israeli monthly consumer price index (2000=100; CBS Israel). Beside these disaggregated, firm-level data, aggregated grower price data are used as well. The aggregated grower price is calculated as the average of the individual grower prices of the four largest exporters – Tnuport, Mehadrin, Agrexco and Pardess – weighted by the respective exporter's export quantity (Figure 1). Data from the 1991/92, 1994/95 and 1996/97 seasons are excluded from the aggregated data set since they are incomplete. Thus, the aggregated data set comprises 205 observations from the 92/93, 94/94, 95/96, 97/98, 98/90 and 90/91 seasons, each season consisting of 32 to 37 observations. The analysis of disaggregated, firm level Israeli grapefruit grower price is done using a balanced data set for the three largest exporting companies Agrexco, Mehadrin and Tnuport, i.e. including only those weeks for which grower price data is available for all three exporters[d]. This data set contains altogether 7 seasons with a total of 205 observations[e] (Figure 2).

We identify three empirical particularities which are explicitly accounted for in the model approach:

1. It is highly probable that the pricing behaviour of the citrus exporting companies changed in the

post-liberalization market. Exporters might have adjusted their long-run pricing strategy in response to the government's minimum price agreement in 1994/95. Also, Israeli shipping costs increased significantly over the period of this analysis. The index of Israeli shipping services (1990=100; source: Central Bureau of Statistics Israel) increased from 120 in 1992 to 160 in 1997 before decreasing again to 140 in 2000. All exporters will have attempted to pass increased shipping costs on to the growers, but firm-specific strategies and the scope for passing on costs may have varied depending *inter alia* on each firms market power.

Figure 1: Aggregated Israeli grower prices and French import prices for grapefruits, 1991/92 to 1999/00 (real NIS/t; sources: Citrus Growers' Association of Israel, CMBI)

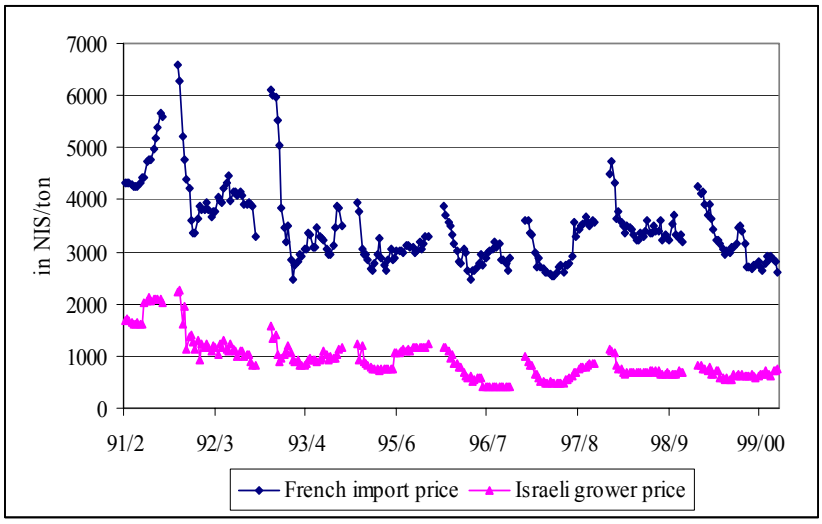
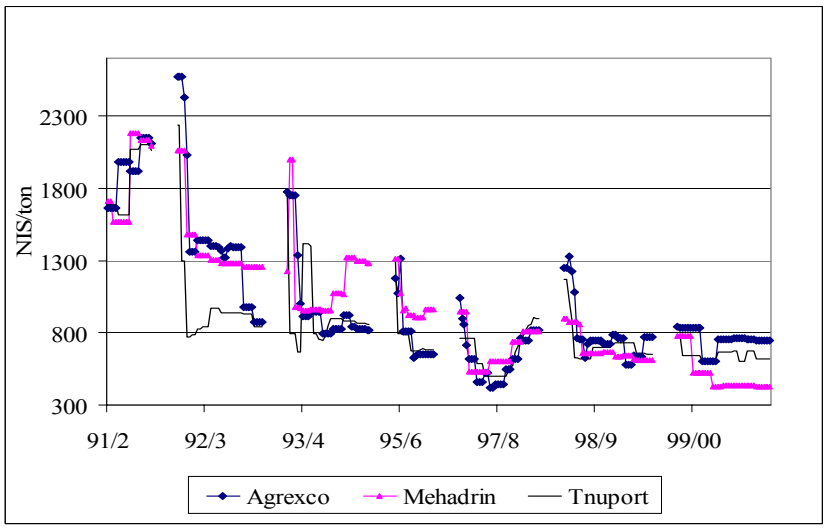


Figure 2: Firm-level Israeli grower prices for the three largest Israeli exporting companies 1991/92 to 1999/00 (real NIS/t; source: CMBI)



In addition, data indicates that homogeneity of grower prices increased over time, which suggests that market competition strengthened. In particular, the coefficient of variation of the grower price

of the three major exporters is particularly high at the beginning of season 92/93 and decreases thereafter (Figure 3). It remains especially low in seasons 97/98 to 98/99. Although it increases again towards the end of season 98/99, exporters' grower prices almost do not vary but remain on a particular level. Further, looking at the aggregated grower price in % of the French import price (Figure 4) we can clearly identify two sub periods, i.e. seasons 92/93 to 95/96 and seasons 97/98 to 99/00, which differ in variance and level. The substantial decrease in the level can be attributed to the increase in transport costs. The decrease in variance indicates that price volatility decreased.

Figure 3: Coefficient of variation of the grower prices of Agrexco, Mehadrin and Tnuport

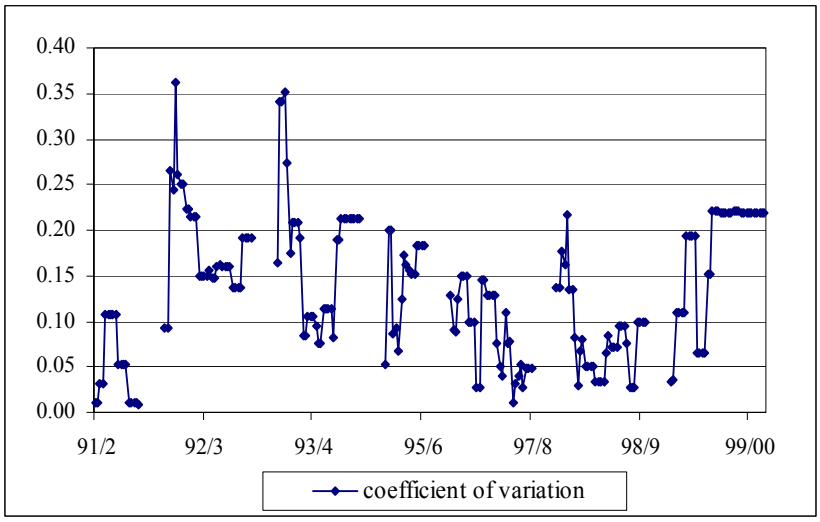
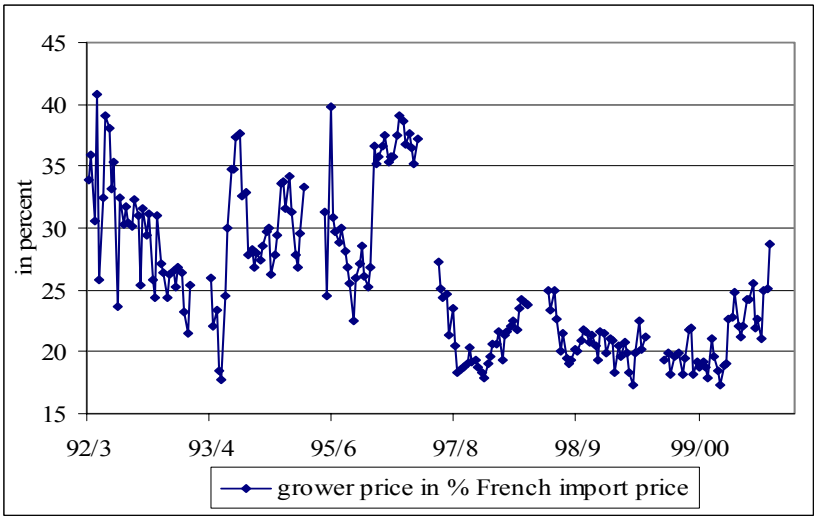


Figure 4: Aggregated Israeli grower price for grapefruits in percent of French import price



We take this into account in the empirical analysis by allowing for individual structural breaks in the cointegration regressions. In addition, we distinguish a heterogeneous, volatile phase in 92/93, 93/94 and 95/96 from a more homogeneous, calm phase in 97/98, 98/99 and 99/00, and estimate the

ECMs for those two phases separately. In the analysis we refer to the former as SUBSET 1 and the latter as SUBSET 2.

2. The empirical approach recognizes that the time series are characterized by gaps resulting from the seasonality of grapefruit production and trade. This implies that for the first observation of each season, no information on the prior time period is present and e.g. meaningful lagged variables can not be created. Therefore, we omit those observations for which the required data is not available[f].

3. The empirical model recognizes the lag between the period (week) to which the grower price corresponds and the period at which the French corresponding import price is determined. The grower price represents the value of the produce at the point of time of its delivery to the packing station whereas the French import price is determined at the harbour of Marseille. According to the consignment system, the grower price is determined *ex post*, i.e. after the produce is sold in the French import market. The minimum time lag between those two points in the transport chain amounts 7 to 9 days. Since delays may occur at several points, this lag is stochastic. Simplifying, we assume a transport lag of two weeks for all models.

3 Methods

3.1 Identifying asymmetry in price transmission

We utilize an ECM model approach (ENGLE AND GRANGER, 1987), which requires that the time series are cointegrated, i.e. a long-run equilibrium exists. First, the long-run equilibrium relationship between the data series p_{it} and p_{jt} is estimated as

$$p_{it} = \alpha_0 + \alpha_1 * p_{jt} + \upsilon_t \quad \text{with t = 1,…,T} \tag{1}$$

The residual vector υ_t represents the short-run deviations from the long-run equilibrium. It is lagged by one period and enters the ECM as the error correction term ECT ($ECT_t = p_{it-1} - \alpha_0 - \alpha_1 * p_{jt-1}$):

$$\Delta p_{it} = \beta_0 + \sum_{n=1}^{K} \beta_{1n} \Delta p_{jt-n+1} + \sum_{m=1}^{L} \beta_{2m} \Delta p_{it-m} + \phi ECT_t + \varepsilon_t \tag{2}$$

To allow for asymmetry in price transmission, change and equilibrium effects caused by price

increases are distinguished from those caused by price decreases by including additional dummy variables in the model:

$$\Delta p_{it} = \beta_0 + \sum_{n_1=1}^{K_1} \beta_{1n_1}^{+} D_{1t}^{+} \Delta p_{jt-n_1+1} + \sum_{n_2=1}^{K_2} \beta_{1n_2}^{-} D_{1t}^{-} \Delta p_{jt-n_2+1} + \sum_{m=1}^{L} \beta_{2m} \Delta p_{it-m} + \phi_1 D_{2t}^{+} ECT_t + \phi_2 D_{2t}^{-} ECT_t + \varepsilon_t \quad (3)$$

with $D_{1t}^{+} = 1$ if Δp_{jt-n_1+1}>0 and 0 otherwise, $D_{1t}^{-} = 1$ if Δp_{jt-n_2+1}<0 and 0 otherwise,

$D_{2t}^{+} = 1$ if ECT_t>0 and 0 otherwise, and $D_{2t}^{-} = 1$ if ECT_t<0 and 0 otherwise.

Asymmetry in price transmission is present if the null hypothesis that the estimated coefficients of the respective positive and negative variable are equal is rejected by an F-test.

3.2 Tests on structural break in a cointegration regression

Standard tests for cointegration (e.g. the residual-based ENGLE AND GRANGER (1987) test) require that the cointegrating vector is time-invariant. If the cointegrating vector changes during the sample period, the results of these tests might be misleading (GREGORY AND HANSEN, 1996). The test of GREGORY AND HANSEN (1996) allows not only to identify a structural break in the cointegration relationship but also to specify its timing and type. In this test, the null hypothesis of no cointegration is tested against the alternative hypothesis of cointegration in the presence of a regime shift within three model frameworks, i.e. a) level shift (intercept changes only), b) a level shift with trend and c) a regime shift:

$$p_{it} = \alpha_{01} + \alpha_{02} * \varphi_{t\tau} + \alpha_{11} * p_{jt} + \alpha_{12} * p_{jt} * \varphi_{t\tau} + \upsilon_t \quad (4)$$

In other words, the null hypothesis is that the standard cointegration model as given by (1) holds. In this test, the best suited model for the cointegration regression is selected and estimated for all possible breakpoints. Next, the residuals of all individual cointegration regressions, which are estimated for all different breakpoints, are tested for the existence of a unit root by an ADF-test. The 'true' structural break corresponds to the cointegration regression with a break point for which the residuals do not have a unit root and the null hypothesis of no cointegration can be rejected.

4 Empirical results

4.1 Asymmetric price transmission analysis with disaggregated data

Price transmission based on the disaggregated Israeli grower price for the three major exporting companies Agrexco, Mehadrin and Tnuport is analyzed within a SUR model. In the presence of unobserved common external factors, SUR is thus more efficient than individual OLS estimation for each exporting company.

The order of integration of the data series is determined by the DICKEY-FULLER (1981) test and the KPSS test of KWIATOWSKI ET AL (1992). The French import price (p_{jt}) is I(0) according to the ADF, but I(1) according to the KPSS. The Israeli grower price series are I(1) according to KPSS as well as the ADF for all three exporters.

The cointegration regressions are estimated for each of the three exporters. The consignment system strongly suggests that the Israeli grower price (p_{it}) is the dependent variable and the French import price (p_{jt}) the independent variable. We utilize the residual-based test by ENGLE AND GRANGER (1987) to test for cointegration. The results point to cointegration between the Israeli grower price and the French import price only for Agrexco (5% significance level). The failure to find cointegration for the other firms may be due to structural breaks as outlined above. Hence, we next test for cointegration allowing for the existence of a structural breaks using the Gregory-Hansen test. For Agrexco, Mehadrin and Tnuport, respectively, structural breaks are identified at the 1% level of significance on March 1993 (week 11), October 1992 (week 41) and October 1997 (week 42). It is striking that the structural break is earliest for Tnuport, the exporter with the largest market share and thus probably the largest degree of market power. The identified break-points of the disaggregated cointegration regressions for Agrexco, Mehadrin and Tnuport are accounted for in the estimation of the cointegration residuals, which enter the ECM as ECT terms.

The estimated coefficients of the long-run equilibrium regression according to equation (4) for each exporter are presented in Table 1. Note that α_{11} is higher than α_{12} for each exporter. Interestingly, α_{12} is by far lowest for Tnuport, the largest exporter with the potentially largest market power. This

decrease in the slope coefficient may be attributed to increasing transport costs, resulting in higher fixed costs and reducing the share of the Israeli grower price in the French import price.

Table 1: Estimated coefficients of the cointegration regression

Exporting firm	α_{01}	α_{02}	α_{11}	α_{12}
Agrexco	-261.29	-184.11	0.456	0.295
Mehadrin	-27.054	-54.74	0.378	0.258
Tnuport	-1157.0	287.0	0.577	0.138

Next, the ECM (equation 3) is estimated for each exporter within a SUR model taking into account the specified break point in the cointegration regression. We find substantial correlation (coefficient = 0.163) between the residuals of the equations for Mehadrin and Tnuport. Lag-lengths K_1 and K_2 are chosen according to the Akaike Information Criteria (AIC) and Bayesian Information Criteria (BIC). Lag-length L accounts for autocorrelation, which is detected by the Ljung-box statistic. Results are presented in Table 2 under COMPLETE. McElroy's R-squared for the SUR is 0.17. Asymmetry in price transmission is identified for Agrexco and Tnuport regarding long-run equilibrium and short-run price transmission at the 1% and 5% significance levels, respectively. Both findings of asymmetry are of the kind that it is beneficial to exporters and bad to growers. For example, when the import price falls, squeezing Agrexco's margin, the grower price is reduced, but 44% of this "error" is corrected immediately. If, on the other hand, the import price increases and Agrexco's margin is stretched, the grower price does not at all increase, which is indicated by the statistically insignificance of the estimated coefficient ϕ_2. For Mehadrin, price transmission is found to be symmetric. To test whether the exporters' price transmission behaviour may have changed, we estimate separate ECMs for the phase with relatively heterogeneous grower prices in the first years after liberalization (SUBSET 1) and the later phase with more homogeneous grower prices (SUBSET 2). Again, the break points in the individual cointegration regressions are taken into account. We detect significant correlations (>0.1) between the residuals of the individual equations for Agrexco and Tnuport (-0.11) as well as Mehadrin and Tnuport (0.16) in SUBSET 1. The estimated coefficients and test statistics are presented in Table 2. In concordance with the COMPLETE estimation, we find asymmetry in long-run price transmission for Agrexco and short-

Table 2: Estimated ECM coefficients based on disaggregated data for the complete data set and data subsets

	COMPLETE						SUBSET 1						SUBSET 2					
	Agrexco		Mehadrin		Tnuport		Agrexco		Mehadrin		Tnuport		Agrexco		Mehadrin		Tnuport	
Coef./ test stat.	Estim. value	t-value	Estim. value	t-value	Estim. value	t-value	Estim. value	t-value	Estim. value	t-value	Estim. value	t-value	Estim. value	t-value	Estim. value	t-value	Estim. value	t-value
β_0	12.79	1.015	19.34	1.470	30.22	1.969	33.69	1.528	45.27	1.560	52.18	1.678	-21.15	-1.552	5.372	0.442	1.294	0.137
β_{11}^{+}	0.069	1.039	-0.013	-0.162	-0.022	-0.268	0.100	0.897	-0.059	-0.411	-0.045	-0.31	-0.001	-0.01	0.017	0.259	0.015	0.302
β_{11}^{-}	0.080	2.080	0.114	2.439	0.205	4.410	0.085	1.517	0.124	1.71	0.331	4.342	0.095	1.772	0.088	1.681	0.131	3.131
β_{12}^{+}	na	na	na	na	na	na	na	na	na	na	0.074	0.539	na	na	na	na	na	na
β_{12}^{-}	na	na	na	na	na	na	na	na	na	na	-0.266	-3.225	na	na	na	na	na	na
β_{21}	na	na	na	na	na	na	na	na	na	na	na	na	na	na	na	na	0.169	1.769
ϕ_1	-0.444	-6.504	-0.089	-1.487	-0.169	-3.109	-0.533	-5.659	-0.153	-1.62	-0.338	-3.944	-0.109	-0.732	-0.018	-0.157	-0.002	-0.026
ϕ_2	-0.027	-0.289	0.159	2.218	0.023	0.249	0.107	0.668	0.23	2.039	-0.112	-0.527	-0.311	-3.052	0.046	0.794	-0.030	-0.648
Adj. R^2	0.245		0.053		0.149		0.28		0.05		0.24		0.15		0.023		0.143	
p-value (Ljung-Box(2))	0.221		0.124		0.182		0.210		0.157		0.469		0.653		0.765		0.510	
p-value (Jarque Bera)	<2.2e-16		<2.2e-16		<2.2e-16		<2.2e-16		<2.2e-16		<2.2e-16		<2.2e-16		<2.2e-16		<2.2e-16	
F-value (short-r. sym.) $\beta_{11}^{+} = \beta_{11}^{-}$	0.016<3.892[a]		1.592<3.892[a]		4.655>3.892[a]		0.012<3.949[c]		1.031<3.949[c]		4.24>3.952[e]		0.878<3.943[f]		1.237<3.943[f]		1.943<3.945[g]	
$\beta_{12}^{+} = \beta_{12}^{-}$	na		na		na		na		na		2.535<3.952[e]		na		na		na	
F-value (long-r. sym.)	9.006>6.773[b]		3.688<3.892[a]		2.950<3.892[a]		8.587>6.939[d]		3.469> 3.949[c]		0.008<3.952[e]		0.462<3.943[f]		1.584<3.943[f]		0.047<3.945[g]	

[a]=Ftab (0.95,1,186); [b]=Ftab (0.99,1,186); [c]=Ftab (0.95,1,88); [d]=Ftab (0.99,1,88); [e]Ftab (0.95,1,86); [f]Ftab(0.95,1,93); [g]Ftab(0.95,1,92)

run price transmission for Tnuport, and symmetry for Mehadrin in SUBSET 1. In contrast, price transmission seems to be symmetric for all three exporters in SUBSET 2. McElroy's R-squared for SUBSET 1 amounts 0.25 and thus exceeds its value for the complete data set. For SUBSET 2, McElroy's R-squared is 0.15 and is thus slightly lower than for the complete data set.

4.2 Asymmetric price transmission analysis with aggregated data

The ECM is also estimated based on the aggregated Israeli grower price. The ADF as well as the KPSS test both find the aggregated Israeli grower price to be I(1). As explained in the previous chapter, the French import prices are either I(0) or I(1). The residuals-based test on cointegration fails to prove that the Israeli grower price and the French import price are cointegrated even at the 10% significance level. Therefore, the GREGORY-HANSEN test for cointegration in the presence of a structural break is conducted (equation 4). The ADF statistic exceeds the 1% critical value of |-5.47| for two breakpoints at the very beginning of season 97/98 (October, 1997, weeks 42 (ADF=|-5.54|) and 43 (ADF=|-5.51|)). In those two cases, the null hypothesis is rejected and it follows that the two data series are cointegrated if the cointegration regression considers a regime shift at the end of season 95/96. Our estimates for the long-run equilibrium regression according to equation (4) are α_{01}=279.890, α_{02}=-22.470, α_{11}=0.222 and α_{12}=0.217. Similar to the analysis with the disaggregated data, the estimated slope coefficient α_{11} is (slightly) higher than α_{12}.

The residuals of this cointegration regression enter the ECM as the ECT term as indicated by equation (2). The asymmetric ECM model in equation (3) and the symmetric ECM in equation (2) are each specified for SUBSET 1 and SUBSET 2, corresponding to the "volatile" seasons 92/93, 93/94 and 95/96 and the "calmer" seasons 97/98, 98/99 and 99/00, respectively (Tables 3a and 3b). Since the Breusch-Pagan-test indicates heteroscedasticity, the t-values for the model variables are estimated based on White's heteroscedasticity-consistent standard errors. The F-value of the test on short-run symmetry is very low, indicating that price transmission is symmetric in the short run for SUBSET 1 as well as SUBSET 2.

Tables 3a and 3b: Estimated ECM of asymmetric (Table 3a) and symmetric (Table 3b) price adjustment for data sets SUBSET 1 and SUBSET 2

Table 3a	SUBSET 1		SUBSET 2	
Coef../ test stat.	Estim. value	T-value	Estim. value	T-value
β_{11}^{+}	0.109	1.112	0.045	1.382
β_{11}^{-}	0.045	0.458	0.136	2.096
β_{21}	-0.402	-4.078	na	na
ϕ_1	-0.512	-2.712	-0.269	-3.050
ϕ_2	0.035	0.206	-0.283	-2.220
AIC/BIC	1245.6/1261.2		1044.5/1057.5	
p-value (Ljung-Box(2))	0.3729		0.293	
p-value (Breusch-Pagan)	9.498e-05		1.621e-05	
p-value (Jarque Bera)	0.023		0.1393	
F-value (short-run sym.)	0.168<3.937[a]		1.835<3.936[b]	
F-value (long-run sym.)	6.062>3.937[a]		3.077<3.936[b]	

[a]=Ftab (0.95,1,92); [b]Ftab(0.95,1,100)

Table 3b	SUBSET 1		SUBSET 2	
Coef./ test stat.	Estim. value	T-value	Estim. value	T-value
β_{11}	0.151	3.20	0.104	3.190
β_{21}	-0.378	-4.30	na	na
ϕ	-0.280	-3.68	-0.277	-3.145
AIC/BIC	1250.8/1261.2		1043.6/1051.4	
p-value (Ljung-Box(2))	0.409		0.215	
p-value (Breusch-Pagan)	6.75e-05		6.327e-06	
p-value (Jarque Bera)	0.002		0.032	

The value of the F-statistic of the test on long-run asymmetry in SUBSET 1 exceeds the theoretical F-value at the 5% significance level. In contrast, the value of the F-statistic of the test on long-run asymmetry is very low for SUBSET 2. Thus, long-run price transmission is identified as asymmetric for SUBSET 1, but symmetric for SUBSET 2. Overall, this is also reflected by the value of AIC and BIC of the asymmetric compared to the symmetric model.

4.3 The impact of price asymmetry

Simplifying, we use our results derived from the aggregated data to quantify the economic implications of price asymmetry for the citrus growers. The grower's losses due to asymmetry are equal to the sum over the differences between the grower prices resulting from the asymmetric and the symmetric models.

The estimated coefficients of the asymmetric ECM are utilized to calculate the grower price assuming the specified type of asymmetric price transmission (p_{it}^{as}), with $ECT_t = p_{it-1} - \alpha_0 - \alpha_1 * p_{jt-1}$

as follows: $p_{it}^{as} = -\alpha_0\phi_1 D_{2t}^{+} - \alpha_0\phi_2 D_{2t}^{-} + \beta_{11} p_{jt} + (1 + \beta_{21} + \phi_1 D_{2t}^{+} + \phi_2 D_{2t}^{-}) p_{it-1} +$

$$(-\beta_{11} - \alpha_1\phi_1 D_{2t}^{+} - \alpha_1\phi_2 D_{2t}^{-}) p_{jt-1} - \beta_{21} p_{it-2} \quad (5)$$

To calculate the grower price (p_{it}^{s}) under the assumption of symmetric price transmission, we assume that the speed of adjustment for positive and negative price changes is equal ($\phi_1 = \phi_2$). The estimated coefficient for ϕ_1 is utilized since this represents the maximum speed of adjustment. The quantitative effect of price asymmetry (qe^{as}) for one season with t=v and t=w corresponding to the beginning and the end of a season, respectively, equals:

$$qe^{as} = \sum_{t=v}^{t=w} p_{it}^{s} - p_{it}^{as} \quad (6)$$

The estimated values for the growers' losses[g] are presented in Table 4. The losses are highest in 1993/94 and lowest in 1992/93. Growers' losses add up to as much as 2.5% of growers' total revenue or 24.7 NIS/ton. As growers' profits are presumably only a small proportion of their total revenues, the loss in grower profits due to asymmetry is likely to be quite important.

Table 4: Growers' losses due to asymmetry in price transmission

Season	92/93	93/94	95/96
Total value (in NIS)	293,608	866,995	667,002
Value (in % of growers' total revenue)	0.7	2.5	1.4
Mean loss (in NIS per ton)	7.7	24.7	14.0
Av. mean loss (in NIS per ton)	15.5		

5 Conclusions

The analyses of price transmission based on disaggregated as well as aggregated grower price data both suggest that price transmission behaviour of Israelis citrus exporters changed in the post-liberalization period after 1991. We attribute those changes to two common external factors, i.e. the government market intervention in favour of the citrus growers, and the substantial increase in sea transport costs. In particular, both analyses find that exporters transmitted grapefruit price changes in the French import market asymmetrically to Israeli citrus growers in the phase with

heterogeneous pricing in the first years after liberalization. The identified asymmetry was beneficial to exporters and bad to growers. However, price transmission became more symmetric in the subsequent phase (second half of the 1990s) which was characterized by more homogeneous pricing.

The results derived with disaggregated, firm-level data make it possible to draw a more differentiated picture. We identify asymmetry in long-run price transmission for Agrexco, immediate short-run price transmission for Tnuport and symmetry for Mehadrin in SUBSET 1. For SUBSET 2, price transmission is symmetric for all three exporters.

We also find that the specified asymmetry in price transmission is economically significant. Our results indicate that growers' losses due to asymmetry amounted to as much as 2.5% of citrus growers' total revenues, and hence presumably a much larger share of their profits. This study demonstrates that while liberalization improved the efficiency of the Israeli citrus international marketing channel, this improvement took time and was probably accelerated by government intervention.

The analysis based on the disaggregated data could be improved further by accounting for heteroscedasticity in the SUR model. In addition, specifying a SUR model for an unbalanced data set would make it possible to include Pardess in the analysis with the disaggregated data. And a model that explicitly incorporates the available data on transport costs might produce sharper insights. Finally, since VON CRAMON-TAUBADEL ET AL (2006) have demonstrated that aggregation can distort the results of tests for asymmetric price transmission, the results reported in this paper for aggregated and disaggregated data should be contrasted and subjected to closer study.

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[a] In contrast, previous studies on asymmetric price transmission in the fresh fruits and vegetables sector (e.g. WARD, 1982, PICK, KARRENBROCK AND CARMAN, 1990, BROOKER, EASTWOOD, CARVER AND GRAY 1997, WILLETT, HANSMIRE AND BERNARD, 1997, GIRAPUNTHONG, VANSICKLE AND RENWICK, 2003) focus on the analysis of price transmission within the national marketing channel.

[b] We are grateful to Yael Kachel for making this data available.

[c] The Israeli firm-level grower prices were surveyed by the Citrus Growers' Association of Israel. They are averages of the prices for different fruit sizes weighted with a size distribution characteristic for each season. The French import prices were collected by a large French fruit import company by a telephone survey of the major fruit importers in France.

[d] Over the time period of this analysis, Tnuport was the largest exporter with a market share of 38%, followed by Mehadrin (28%), Agrexco (26%) and Pardess (8%). For Pardess only 178 grower price observations are available. Hence, Pardess is excluded from the disaggregated data set to avoid substantial reduction of the length of the balanced data set and thus loss of information and degrees of freedom.

[e] This data set is not exactly equivalent to that utilized for the analysis with the aggregated data; e.g. observations of season 91/92 are only included in the balanced disaggregated data set.

[f] In contrast, Ward (1982) introduces additional dummy variables in the model ensuring that each observation is included in the model for the estimation of those variables only for which the required information on the previous observations is present. Our approach is characterized by a loss of data observation where as Ward's approach implies an estimation bias.

[g] This is based on the assumption of a uniform distribution of sales over time, so that each price difference from a specific point in time has the same weight. If most sales take place in weeks with lower (higher) price differences, then growers' losses would be lower (higher) than indicated by these calculations.