FARM GROWTH IN HUNGARY, SLOVENIA AND FRANCE

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ABSTRACT

The article investigates the validity of Gibrat's Law for French, Hungarian and Slovenian farms with FADN data and Heckman selection models, quantiles regressions and panel unit root tests. The contribution to the literature is threefold. First, we compare farm growth in countries with rather different farm structures. Second, we apply two different testing techniques. Finally, we focus on specialised crop and dairy farms rather than all farms, avoiding biases due to heterogeneous structures across the agricultural sector. Results reject the Gibrat's Law for crop farms in France (except for one sub-period) and Hungary but confirm it for French and Slovenian dairy farms.

KEYWORDS: farm growth, Gibrat's Law, panel unit root, quintile regression

1. INTRODUCTION

Different approaches have been developed in firm/farm level analyses to test whether Gibrat's Law holds (Gibrat, 1931), that is to say whether the rate of growth of a firm/farm is independent from its size (Goddard et al., 2002; Harris and Trainor, 2005; Goddard et al., 2006; Bakucs and Fertő, 2009). Most often cross-section tests, panel tests, and alternative panel unit root tests have been applied to test the relationship between firm/farm growth and the measures of firm/farm size. The empirical research yielded rather contradictory results. Some studies (Weiss, 1999; Shapiro et al., 1987) rejected Gibrat's Law for farm growth, finding that small farms tend to grow faster than large ones. Other studies (Upton and Haworth, 1987; Kostov et al., 2005) found no evidence (except for the small farms in the case of Kostov et al., 2005) to reject Gibrat's Law. Previous research on Hungarian agriculture shows that the growth trajectory of family and corporate farms is similar (Fertő and Bakucs, 2009).

The historical development and the evolution of farms in Europe vary by countries, not only between Eastern and Western Europe, but also inside both regions. grouping Eastern Europe, these differentials in farm size and its growth are caused by the initial conditions that are linked to the agricultural history during the previous communist system and later institutional and policy reforms, while in Western Europe they are caused by the long-term institutional and policy evolutionary factors and market conditions. To test the validity of the Gibrat's Law, we associated this test to three countries differing in the initial conditions, institutional and policy reforms, and farm structures. The analysis is based on Hungarian, Slovenian and French Farm Accountancy Data Network (FADN) including farms above two European Size Units (ESUs). During the communist system Hungarian agriculture was collectivised and the average farm size has been all the time among the largest in Europe. In Slovenia the collectivisation failed and small-scale farm structure has remained among the smallest in Europe. In France farm structure has developed under market conditions and policy support (Piet et al., 2010). While its farms are among the largest in Western Europe, they are smaller than in Hungary. Therefore, our comparative analysis includes three countries with different historical-institutional developments and different farm structures: small-scale farms in Slovenia, medium-sized farms in France, and large-scale farms in Hungary. The proportion of small farms in Slovenian agriculture is much higher than in Hungary.

This rest of the paper is structured as follows: section 2 presents the methodology applied, section 3 describes the data. The empirical analysis is presented in section 4, and then section 5 concludes.

2. METHODOLOGY

Equation (1) represents the stochastic process underlying Gibrat's Law:

$$\frac{S_{i,t}}{S_{i,t-1}} = \alpha S_{i,t-1}^{\beta_1 - 1} \varepsilon_{i,t} \tag{1}$$

where $S_{i,t}$ is the size of farm i in year t. $S_{i,t-1}$ is the size of the i^{th} farm in the previous period, $\varepsilon_{i,t}$ being the disturbance, independent from $S_{i,t-1}$. α is the common growth rate of all farms, whilst β_1 measures the effect of the initial size upon the given farm's growth rate. If $\beta_1 = 1$, then growth rate and initial size are independently distributed and Gibrat's Law holds. If the coefficient is less than one, it follows that small farms tend to grow faster than large farms, while the opposite is the case if β_1 is greater than unity. Rewriting equation (1) into the form represented by equation (2), allows testing the significance of the β_1 coefficient:

$$\log S_{i,t} = \beta_0 + \beta_1 \log S_{i,t-1} + \mu_{i.t} \tag{2}$$

where β_0 = logα and $\mu_{i,t}$ = log$\varepsilon_{i,t}$. Following Ward and McKillop (2005), if β_1=*1*, i.e. Gibrat's Law holds, then positive (negative) values of β_0 indicate a growth (decrease) in the average farm size. If however β_1<*1*, i.e. smaller farms tend to grow faster than larger ones, then the long-run mean size of the farm population is given by ($\frac{-\beta_0}{\beta_1 - 1}$).

The empirical analysis faces several econometric issues to test Gibrat's Law. The first concern is the heteroskedasticity issue which may occur when the Gibrat's Law is not confirmed (if small farms grow faster than their larger counterparts, the variance of growth should tend to decrease with size). The second traditional problem is that when there is serial correlation in growth rates, ordinary least square (OLS) estimators are inconsistent even though estimation proceeds using cross-sectional data (Chesher, 1979). An important issue in the empirical analysis is the sample selection problem. Since growth rate is only possible to be measured for surviving farms (still operating in period *t*), and since slow growing farms are most likely to exit, it is easy to see that small, fast growing farms can easily be overrepresented in the sample, thus introducing biases in the results. This problem is of a particular importance in the present paper, since the proportion of small farms in transition economies in general, and in Slovenia in particular, is much higher than in developed economies. Heckman (1979) introduced a two-step procedure to control for the selection problem. In step one, a farm survival model for the full sample (both surviving and exiting farms) is estimated, using a probit regression. This equation is used to obtain the inverse of Mill's Ratio for each observation (equation (3)):

$$P(f_i = 1) = F(\delta + \gamma \log S_{i,t-1}) + \mu \tag{3}$$

where f_i = 1 denotes survivor, f_i = 0 exit, and μ is the disturbance.

The inverse Mill's Ratio derived from equation (3) is then introduced in step two, which is equation (2). A significant coefficient for the inverse Mill's Ratio would then indicate that the sample selection problem is present.

In the OLS regression estimation, error terms are assumed to follow the same distribution irrespectively of the value taken by the explanatory variables. Since we can only analyze surviving farms, estimations are conditional on survival (*conditional objects*, see Lotti et al., 2003). Therefore, in this paper we use the more robust and more informative quantile regression estimation technique. Following Lotti at al. (2003), the θ^{th} sample quantile, where $0 < \theta < 1$, can be defined as:

$$\min_{b \in R} \left[\sum_{i \in \{i: y_i \geq b\}} \theta |y_i - b| + \sum_{i \in \{i: y_i < b\}} (1-\theta)|y_i - b| \right] \tag{4}$$

For a linear model such as $y_i = \beta' x_i + \varepsilon_i$, the θ^{th} regression quantile is the solution of the minimization problem, similar to equation (4):

$$\min_{b \in R^k} \left[\sum_{i \in \{i: y_i \geq x_i b\}} \theta |y_i - x_i b| + \sum_{i \in \{i: y_i < x_i b\}} (1-\theta)|y_i - x_i b| \right] \tag{5}$$

Solving equation (5) for b results a robust estimate of β. To obtain unbiased error terms, we use bootstrap methodology to estimate the variance-covariance matrix.

In their seminal paper, Goddard et al. (2002) showed that the above presented cross-sectional methodology results in biased parameter estimates and test statistics suffer of low power if there are heterogeneous individual farm effects. If there are heterogeneous individual effects, α in equation (1) is not constant, i.e. it should be correctly represented in the equation as α_i. Since the most common way of testing farm growth (equation (2)) is very close to the auxiliary regression used in Dickey-Fuller type unit root tests, Goddard et al. (2002) present an alternative farm growth testing methodology, using time series econometrics, suitable for longer panel datasets. Gibrat's Law is satisfied if logged size variables for individual farms are non-stationary (i.e. in unit roots) and it is rejected if size variables are stationary. Oliveira and Fortunato (2006) apply the method to test Gibrat's Law amongst Portuguese manufacturing firms, Goddard et al. (2006) to test the firm size, profit rate and growth of 96 large farms in the United Kingdom (UK) using a 31-year long panel, and Harris and Trainor (2005) to analyse whether the Law of Proportionate Effects holds amongst UK manufacturing industries during the 1973-1998 period.

Panel unit root tests are similar, but not identical, to unit root tests run on individual series. Consider equation (6):

$$y_{i,t} = \rho y_{i,t-1} + X_{i,t}\delta_i + \varepsilon_{i,t} \tag{6}$$

where i = 1,2,…,N are cross-section units and t=1,2,…,T the observed periods, X_{it} possible exogenous variables, ρ_i the autoregressive coefficients, and the errors $\varepsilon_{i,t}$ are assumed to be mutually independent idiosyncratic disturbance terms. If $|\rho_i| < 1$, y_i is considered stationary, while if $|\rho_i| = 1$, the process contains a unit root. With panel unit root tests, there are two assumptions regarding ρ. First, the persistence parameters are common across cross-sections, that is to say $\rho_i = \rho$, for all i. Second, ρ_i can freely vary across cross-sections. There are a number of panel unit root tests assuming one of the above assumptions. Considering the well known low power properties of unit root tests, in this paper we employ a battery of unit root tests: Levin et al. (2002) method (common unit root process), Im et al. (2003) method (assuming individual unit root processes), ADF-Chi square and PP-Chi square.

3. DATA

The analysis is based on Hungarian, Slovenian and French FADN including farms above two ESUs (one ESU is equivalent to 2,200 euros of gross margin). The analysis is performed for two farm specialisations: dairy farms and fieldcrop farms, based on their European type of farming (TF) classification. The European classification into a specific TF is based on which production farms derive at least 75 percent of their gross margin from. Dairy farms are classified as TF41 and fieldcrop farms as TF1. The time span used for analysis is 2001-2007 for Hungarian and French farms, and 2004-2006 for Slovenian farms based on data availability.

In agriculture there is no single measure of farm size. The proxy mainly depends on farms' production specialisation and technology. Although statistics on farm size often refer to utilised agricultural area (UAA), this measure is often irrelevant for livestock farms. Therefore, in this paper UAA is used as a farm size proxy for crop farms, while livestock units (LSUs, that is to say the total number of livestock heads on the farm aggregated with European standard weight coefficients) are employed for dairy farms' size. Moreover, within a specific specialisation, technology (such as capital or land intensity) may be different and may thus render the comparison with UAA or livestock units difficult. For this reason here farm size is also measured with the labour force range: the number of full-time equivalent workers per year on the farm (Annual Working Units, AWU), both family and hired, is used for dairy and crop farms.

Table 1 presents some descriptive statistics of the data used.

<< Table 1 >>

These summary statistics clearly indicate the size differences of dairy and crop farms between Hungary, France and Slovenia. The Hungarian samples present the largest farms on average, while the Slovenian samples the smallest farms. This can be explained by the different historical trajectories of both countries as explained above: while the Hungarian farming sector had been almost fully collectivised during the communist time, this was not the case for Yugoslavia, including Slovenia, where small family farms prevailed. The Slovenian farms use more labour on average than French farms, despite their UAA (for crop farms) or number of LSUs (for dairy farms) being much lower than those of French farms.

In Slovenia the maximum size of the dairy farm in LSU is less than the minimum size in Hungary. The Hungarian dairy farm is approximately 38 times greater than in France or 85 times greater than in Slovenia. Even much greater differentials in terms of size of dairy farms between the analysed countries are seen in terms of labour in AWU: the average Hungarian dairy farm has 4,714 workers, which is approximately 2,619 times greater than in France or 1,924 times greater than in Slovenia. This implies that in Hungary the ratio between number of LSU and number of AWU is less than one, while it is 49.4 for France and 15.8 for Slovenia. It is worth mentioning that one AWU is equal 1,800 hours annual full time employment in Slovenia and 2,200 hours in France and Hungary. Farm size differentials between the analysed countries are also seen for crop farms. The average crop farm size in Hungary is 3,318.4 hectares (ha), which is around 25 times greater than in France or around 164 times greater than in Slovenia. The Hungarian crop farm uses on average 2,904.4 workers, which is around 1,571 times more than in France or 1,351 times more than in Slovenia. On average, one AWU on a Hungarian crop farm cultivates 1.14 ha of land, against 72.3 ha in France and 9.5 ha in Slovenia.

This confirms that using a single farm size measure for Gibrat Law's analysis may give unreal results.

The agricultural sector in the three countries had to face changes in their economic and policy environment during the period studied. Most notably, Hungary and Slovenia have entered the European Union (EU) in 2004. For this reason, in addition to analysing the Gibrat's Law over the full period, two sub-periods are used for Hungary, 2001-2003 and 2004-2007, to test for the influence of EU accession. Unfortunately, the time span for Slovenian data (2004-2006) is not long enough to analyse such effect. Regarding France, the agricultural sector has experienced the 2003 reform of the Common Agricultural Policy (CAP), which introduced the new decoupled instrument of Single Farm Payments (SFP), that is to say payments given to farms on a per hectare basis regardless of their production level and type on the area. The reform was implemented in France in 2006. Therefore, the two sub-periods used for this country are 2001-2005 and 2006-2007.

4. EMPIRICAL RESULTS

We present our econometric results by farm type (crop and dairy farms) separately. Results of the estimation for crop farms are shown in Table 2. Interestingly we do not find evidence for the selection bias, the inverse Mill's Ratios are insignificant for all specifications. Our estimations suggest that we can reject the Gibrat's Law for crop farms in Hungary and France (except for the 2006–2007 period in France if UAA is used as size measure), irrespective of the methodology. The coefficients are usually less than one implying that small farms grew faster than large farms during the periods studied. Note that coefficients are very close to, or even equal to, one for France. Slovenia shows a rather different and mixed picture. The results suggest that Gibrat's Law can be rejected for the Heckman selection model using labour as the measure of farm size. In addition, we find that coefficients are larger than one if using land as measure of farm size, providing evidence for faster growth of larger farms.

The estimations results for dairy farms are presented in Table 3. They show that we can reject the Gibrat's Law for Hungarian dairy farms independently from the period, farm size measure and econometric approach. We do not find evidence for selection bias. Results are rather different for French dairy farms. Again, all coefficients are very close to one, but Gibrat's Law is confirmed only for the full period 2001-2007 and for the sub-period 2001-2003 using LSUs as the size measure with the Heckman selection model. Results indicate that Gibrat's Law is valid for Slovenia for dairy farms using LSUs as the size measure. In addition, the Mill's Ratio provides evidence for selection bias when using labour size measure.

<< Table 2 >>

<< Table 3 >>

Table 4 presents panel unit root test results for specialised crop farms. Four unit root tests were applied. Figures in this table represent significances for the unit root null hypothesis against the alternative of stationary processes. Since FADN data for Slovenia are only available for the 2004-2006 period, the time span is too short and thus unit root tests were run for Hungary and France only.

Regardless of the specification or unit root test employed, the null hypothesis of non-stationary process of both land and labour size variables is strongly rejected for Hungary. For

French crop farms, land is stationary, whilst for the labour variable only the Levin, Lin and Chu (LLC) test does not reject the unit root null with intercept and trend specification.

Table 5 presents unit root test results for specialised dairy farms. Results are very similar to those obtained in Table 4 for specialised crop farms. With the exception of LLC for French labour variable with intercept and trend deterministic specification, the presence of the unit root null is strongly rejected for all other variables.

<< Table 4 >>

<< Table 5 >>

5. CONCLUSIONS

The article investigated the validity of Gibrat's Law for French, Hungarian and Slovenian farms using FADN data and employing Heckman selection models, quantiles regressions and panel unit root tests. The contribution to the literature is threefold. First, we analyse and compare farm growth in three countries with rather different farm structures: Hungary as one New Member State (NMS) of the EU that experienced agricultural collectivisation, Slovenia as another NMS of the EU where collectivisation did not happen, and France as an Old Member State of the EU. Second, we contribute to the methodology by applying two different testing techniques, one rooted in cross sectional econometrics, and one in panel time series econometrics. Finally, contrary to most studies analysing farm growth rate, we focus on specialised crop and dairy farms rather than the whole farm sample, thus eliminating possible biases due to heterogeneous farm structures across the agricultural sector.

Our results strongly reject the validity of the Gibrat's Law for crop farms in France (with one exception) and Hungary, providing evidence that smaller farms grew faster than larger ones over the period studied. The proportion of small crop farms in Slovenian agriculture is much higher than in France or Hungary and empirical results for Slovenia suggests that the rate of growth of crop farm is independent from its size. Similarly, estimations confirm the validity of the Gibrat's Law for French and Slovenian dairy farms.

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Table 1. Descriptive statistics

		Dairy farms		Crop farms	
		Livestock units	**Labour in AWU**	**Arable land in hectares**	**Labour in AWU**
Hungary	Number of obs.	692	692	5482	5482
	Mean	3300.66	4713.89	3318.35	2905.37
	St. Dev.	1759.29	2210.02	1871.98	2304.78
	Min	255	63	50	1
	Max	6,169	8376	6517	8436
France	Number of obs.	7598	7598	13403	13403
	Mean	88.97	1.80	133.82	1.85
	St. Dev.	51.50	0.84	82.44	1.47
	Min	12.33	0.8	2	0.75
	Max	658.59	8.19	774.42	41
Slovenia	Number of obs.	726	726	174	174
	Mean	38.69	2.45	20.33	2.15
	St. Dev.	31.89	0.87	38.77	1.59
	Min	3.86	0.38	2.07	0.21
	Max	236.03	6.75	325.62	11.93

Note: 1 AWU is equivalent to 2,200 hours full time labour in France and Hungary, and 1,800 hours in Slovenia.

Table 2. Heckmann and quantile regression (q50) estimates for crop farms

	2001 – 2007				**2001 – 2003 (2001 – 2005)[a]**				**2004 – 2007 (2006 – 2007)[a]**			
	Heckmann		**Quantile**		**Heckmann**		**Quantile**		**Heckmann**		**Quantile**	
	land	**lab**	**land**	**lab**	**land**	**lab**	**land**	**lab**	**land**	**lab**	**land**	**lab**
					Hungary							
Size	0.55*	0.25*	0.51*	0.74*	0.35*	0.49‡	0.74*	0.64*	0.6*	0.88*	0.92*	0.92*
cons	3.47*	5.82*	4.00*	2.07*	5.06*	3.87*	2.15‡	2.94*	3.04*	0.82*	0.66	0.59◊
Mills λ	0.00	0.00	-	-	0.00	0.00	-	-	0.00	0.00	-	-
Wald1	0.00	0.00	-	-	0.00	0.00	-	-	0.00	0.00	-	-
Wald2	-	-	0.02	0.00	-	-	0.00	0.00	-	-	0.00	0.00
Wald3	14.89*	194.4*	-	-	33.43*	172.6*	-	-	174.1*	681.5*	-	-
Pseudo R^2	-	-	0.08	0.31	-	-	0.16	0.31	-	-	0.49	0.54
N surv	240	240	240	240	248	248	248	248	295	295	295	295
N total	272	272	-	-	272	272	-	-	330	330	-	-
					France							
Size	0.97*	0.83*	0.99*	1.00*	0.98*	0.85*	0.99*	1.00*	0.99*	0.97*	1.00*	1.00*
cons	0.17*	0.09*	0.02◊	0.00	0.11*	010*	0.01◊	0.00	0.01	0.01*	0.00	0.00
Mills λ	0.00	0.00	-	-	0.00	0.00	-	-	0.00	0.00	-	-
Wald1	0.00	0.00	-	-	0.00	0.00	-	-	0.64	0.00	-	-
Wald2	-	-	0.00	0.00	-	-	0.00	0.00	-	-	0.00	0.00
Wald3	11081*	1620*	-	-	19618*	2446*	-	-	10144*	1998*	-	-
Pseudo R^2	-	-	0.80	0.52	-	-	0.85	0.57	-	-	0.94	0.86
N surv	975	975	975	975	1277	1277	1277	1277	1571	1571	1571	1571
N total	2061	2061	-	-	2061	2061	-	-	1838	1838	-	-
					Slovenia[b]							
Size	-	-	-	-	-	-	-	-	1.07*	0.81*	1.04*	0.97*
cons	-	-	-	-	-	-	-	-	-0.16	-0.38	-0.11	0.02
Mills λ	-	-	-	-	-	-	-	-	-0.02	0.79	-	-
Wald1	-	-	-	-	-	-	-	-	0.03	0.37	-	-
Wald2	-	-	-	-	-	-	-	-	-	-	0.56	0.88
Wald3	-	-	-	-	-	-	-	-	1038*	15.89*	-	-
Pseudo R^2	-	-	-	-	-	-	-	-	-	-	0.84	0.51
N surv	-	-	-	-	-	-	-	-	27	27	27	27
N total	-	-	-	-	-	-	-	-	48	48	-	-

Notes: land = UAA (ha), lab = labour (AWU), Mills λ = probability (significance of the inverse Mill's Ratio)

* significant at 1%

‡ significant at 5%

◊ significant at 10%

Wald1: test of H_0: size at the beginning of period (2001, 2004 or 2006) = 1 (probability)

Wald2: test of H_0: equality of the coefficients from quintile regression when: q = 0.10, q = 0.25, q = 0.50, q = 0.75, and q = 0.90. (probability)

Wald3 - regression statistic, H_0: all coefficients equal 0. (χ^2 statistic)

N surv: number of surviving farms

N total: total number of observations used

[a] Sub-periods for France

[b] Period for Slovenia is 2004 and 2006.

Table 3. Heckmann and quantile regression (q50) estimates for specialised dairy farms

	2001 - 2007				2001 – 2003 (2001 – 2005)[a]				2004 – 2007 (2006 – 2007)[a]			
	Heckmann		Quantile		Heckmann		Quantile		Heckmann		Quantile	
	liv	lab	liv	lab	liv	lab	liv	lab	liv	lab	liv	lab
Hungary												
Size	0.48*	0.53*	0.73*	0.76*	0.78*	0.69*	0.91*	0.66*	0.47*	0.7*	0.81*	0.92*
cons	4.1*	4.04*	2.21	2.01	1.62$^{\Diamond}$	2.6*	0.69	2.92*	4.22*	2.51*	1.57	0.60
Mills λ	0.00	0.00	-	-	0.00	0.00	-	-	0.00	0.00	-	-
Wald1	0.00	0.00	-	-	0.07	0.00	-	-	0.00	0.00	-	-
Wald2	-	-	0.07	0.1	-	-	0.09	0.18	-	-	0.12	0.00
Wald3	10.95*	32.15*	-	-	43.29*	68.43*	-	-	19.3*	68.7*	-	-
Pseudo R^2	-	-	0.29	0.28	-	-	0.43	0.47	-	-	0.43	0.47
N surv	26	26	26	26	41	41	41	41	42	42	42	42
N total	108	108	-	-	108	108	-	-	84	84	-	-
France												
Size	0.99*	0.83*	1.00*	1.00*	0.98*	0.90*	0.99*	1.00*	1.00*	0.95*	1.00*	1.00*
cons	0.01	0.08‡	-0.01	0.00	0.05	0.04‡	0.00	0.00	-0.02	0.01*	0.00	0.00
Mills λ	0.00	0.00	-	-	0.00	0.00	-	-	0.00	0.00	-	-
Wald1	0.81	0.00	-	-	0.26	0.00	-	-	0.17	0.00	-	-
Wald2	-	-	0.83	0.00	-	-	0.55	0.00	-	-	0.03	0.00
Wald3	2965*	731*	-	-	7240*	1308*	-	-	25471*	8955*	-	-
Pseudo R^2	-	-	0.70	0.54	-	-	0.75	0.61	-	-	0.86	0.87
N surv	417	417	417	417	601	601	601	601	761	761	761	761
N total	1267	1267	-	-	1267	1267	-	-	973	973	-	-
Slovenia[b]												
Size	-	-	-	-	-	-	-	-	1.00*	0.63*	1.00*	0.92*
cons	-	-	-	-	-	-	-	-	0.02	0.28*	0.00	0.05
Mills λ	-	-	-	-	-	-	-	-	0.00	-0.37*	-	-
Wald1	-	-	-	-	-	-	-	-	0.68	0.00	-	-
Wald2	-	-	-	-	-	-	-	-	-	-	0.99	0.01
Wald3	-	-	-	-	-	-	-	-	5338*	76.01*	-	-
Pseudo R^2	-	-	-	-	-	-	-	-	-	-	0.83	0.44
N surv	-	-	-	-	-	-	-	-	180	180	180	180
N total	-	-	-	-	-	-	-	-	217	217	-	-

Notes: liv = livestock units, lab = labour (AWU), Mills λ = probability (significance of the inverse Mill's Ratio)

* significant at 1%

‡ significant at 5%

$^{\Diamond}$ significant at 10%

Wald1: test of H_0: size at the beginning of period (2001, 2004 or 2006) = 1 (probability)

Wald2: test of H_0: equality of the coefficients from quintile regression when: q = 0.10, q = 0.25, q = 0.50, q = 0.75, and q = 0.90. (probability)

Wald3 – regression statistic, H_0: all coefficients equal 0. (χ^2 statistic)

N surv: number of surviving farms

N total: total number of observations used

[a] Sub-periods for France

[b] Period for Slovenia is 2004 and 2006.

Table 4. Panel unit root tests for crop farms

	Land				Labour			
Specification	**LLC**	**IPS**	**ADF**	**PP**	**LLC**	**IPS**	**ADF**	**PP**
				Hungary				
intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
intercept, trend	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
				France				
intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
intercept, trend	0.00	0.00	0.00	0.00	1.00	0.00	0.00	0.00

Note: LLC = Levin, Lin and Chu test (probability, assumes common unit root process)

IPS= Im, Pesaran and Shin test (probability, individual unit root process)

ADF= ADF Fisher Chi square (probability, individual unit root process)

PP = PP Fisher Chi square (probability, individual unit root process)

Lag length 0 selected by Schwarz Bayesian Criterion

Table 5. Panel unit root tests for dairy farms

	Livestock				Labour			
Specification	**LLC**	**IPS**	**ADF**	**PP**	**LLC**	**IPS**	**ADF**	**PP**
				Hungary				
intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
intercept, trend	0.00	0.03	0.02	0.00	0.00	0.00	0.00	0.00
				France				
intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
intercept, trend	0.00	0.00	0.00	0.00	0.50	0.00	0.00	0.00

Note: LLC = Levin, Lin and Chu test (probability, assumes common unit root process)

IPS= Im, Pesaran and Shin test (probability, individual unit root process)

ADF= ADF Fisher Chi square (probability, individual unit root process)

PP = PP Fisher Chi square (probability, individual unit root process)

Lag length 0 selected by Schwarz Bayesian Criterion