THE PRODUCTIVITY OF FAMILY AND HIRED LABOUR IN EU ARABLE FARMING

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Abstract

This paper investigates the impact of labour force composition on productivity in EU arable farming. We test the heterogeneity of family and hired labour for a set of eight EU member states. To this end, we estimate augmented production functions using FADN data for the years 2001-2008. The results reject the notion that hired labour is generally less productive than family workers. In fact, farms with a higher share of hired workers are more productive than pure family farms in countries traditionally characterised by family labour, namely France, West Germany and Poland. Here, an increase in reliance on hired labour or the shift of family labour to more productive tasks could raise productivity. This finding calls into question a main pillar of the received family farm theory. In about half the countries, there are no statistically different effects of both types of labour. For the United Kingdom, we find the classical case with family farms being more productive than those relying on hired labour. In this situation supervision by family members could increase productivity. As a side result, we find little evidence of non-constant technical returns to scale.

Keywords

Labour productivity, production function estimation, European Union, FADN

1 Introduction[1]

According to a widely accepted view, large-scale farming operations involving many workers under a centralised management authority are economically inferior to smaller family-run businesses, at least in the temperate zones (HAYAMI, 2010). The two maintained hypotheses of the underlying "family farm theory" are that (1) technological scale economies are typically exhausted before farm size exceeds the labour capacity of a family and that (2) growth of the labour force beyond family members is inhibited by rising supervision costs. These hypotheses used to be supported by a large body of empirical literature from developed and developing countries (BREWSTER, 1950; SCHMITT, 1991; HAYAMI and OTSUKA, 1993; ALLEN and LUECK, 1998; EASTWOOD et al., 2010). For many decades after World War II, the economic and social superiority of family farms over agriculture based on hired labour was a widely held notion among researchers, governments and international organizations.

However, even in agricultural regions traditionally dominated by small to medium family farm operations, such as Western Europe or the US, farm sizes have been growing and, more importantly, the share of hired workers in total labour force has been steadily increasing (BLANC et al., 2008). According to latest figures by the EUROPEAN COMMISSION (2012), regularly employed non-family members on average contributed 14.7 per cent of the total agricultural workload in the EU-27 in 2010, whereas irregularly employed non-family members contributed another 7.7 per cent. This share has been on the rise for years, a fact that calls into question the validity of hypothesis (2) outlined before.

The typical argument for different productivities of these two types of labour is based on the idea that both have diverging incentives. Hired labour is usually no residual claimant and their effort cannot commonly be observed because of the idiosyncracy of agricultural production

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(e.g. seasonality, weather effects). Therefore, hired labourers have incentives to "shirk", resulting in effort levels that are only a fraction of those achieved by family labour. As a result, both kinds of labour are not easily substituted. This perceived problem can be mitigated by hired labour supervision. Hence, transaction costs in the form of supervision costs arise, making farm production based on hired labour more expensive. On the other hand, however, the following argument in favour of hired labour is often overlooked: growing farms with a larger stock of workers may allow more specialisation and the division of labour into distinct tasks (ALLEN and LUECK, 1998; KIMHI, 2009). For example, family members might concentrate on management and/or supervision tasks, while hired labourers specialise in non-managerial tasks. To the extent that modern farming technologies allow such specialisation benefits, the productivity of hired labour may well exceed that of a family member who is a "jack of all trades but the master of none".

Given these conflicting views, the present study aims to revisit the relative superiority of family over hired labour by confronting the accepted wisdom with new empirical evidence. In exploring the relative productivity of family versus hired labour, we follow BARDHAN (1973), DEOLALIKAR and VIJVERBERG (1987) and FRISVOLD (1994) who investigated this question for the developing country context of India. Whereas these authors found evidence in favour of both arguments presented before, we are primarily interested in their methodological approach. We follow these authors in using a parametric production function specification that accounts for heterogeneous labour impacts. This approach focuses on a single parameter of relative labour productivity and thus allows straightforward interpretation. Yet, our estimation technique goes beyond the received estimators used by the previous authors in tackling potential endogeneity problems. DEOLALIKAR and VIJVERBERG as well as FRISVOLD resorted to traditional household/farm fixed effects approaches. While we also report results for such models, we focus on state of the art estimators introduced by LEVINSOHN and PETRIN (2003) and WOOLDRIDGE (2009). Our database is a panel originating from the Farm Accountancy Data Network (FADN) of eight EU member states: Denmark, France, Germany, Italy, Poland, Slovakia, Spain, and the United Kingdom. We split Germany into East and West. Data is available for the years 2001-2008. It includes arable farms in member states with very different farm structure, both with traditional family-type farming (e.g., France and Italy) and a high share of hired labour (e.g., East Germany, Slovakia). We provide a comparison of results obtained by the received ordinary least squares (OLS) and fixed effects approaches. To our knowledge, there exist no comparable studies for EU agriculture in the area of labour force heterogeneity to date.

The results reject the notion that hired labour is generally less productive than family workers. As a most striking outcome, hired labour is more productive than family members in countries traditionally characterised by family farms, namely France, West Germany and Poland. In about half the countries, there are no statistically different effects of both types of labour. Only in the United Kingdom do we find the classical case with family labour being more productive than hired labour. As a side result, we find little evidence of non-constant technical returns to scale. Farm growth in Europe may thus indeed be increasingly driven by scale neutral technologies which allow the realisation of gains from labour specialisation.

The study proceeds as follows. In section 2 we discuss the theoretical framework to measure labour heterogeneity. Section 3 briefly discusses the empirical strategy together with the estimation methods. In section 4 we present the data. Section 5 discusses the results. Section 6 concludes.

2 Theoretical framework

Estimation frameworks for production analysis in agriculture can be grouped into dual and primal approaches (MUNDLAK, 2001). The dual approach to labour productivity was recently used by D'ANTONI et al. (2011) and relies solely on an evaluation of the partial substitution elasticity between hired and family labour. This involves the estimation of a cost function; usually Translog jointly with its cost share equations. The flexible nature of the Translog functional form makes no prior assumptions on the substitution relationship between hired and family labour. However, the direct impact of hired and family labour on productivity can only be assessed by the primal approach, in which a production function is estimated. Suppose production can be described by the following generalised function:

$$y_{it} = f(A_{it}, E_{it}, K_{it}, M_{it}) + \omega_{it} + \varepsilon_{it}, \tag{1}$$

where y_{it} the natural logarithm of output Y, A_{it} is land use, E_{it} is the effective labour input, K_{it} fixed capital, M_{it} materials (working capital) and i and t are farm and time indices. ω_{it} are farm- and time-specific factors known by the farmer but unobserved by the analyst. ε_{it} are the remaining independent and identically distributed errors. Previous studies on labour force heterogeneity focused on a Cobb-Douglas functional form for (1). In this case, we arrive at:

$$y_{it} = \alpha^A a_{it} + \alpha^E e_{it} + \alpha^K k_{it} + \alpha^M m_{it} + \omega_{it} + \varepsilon_{it}, \tag{2}$$

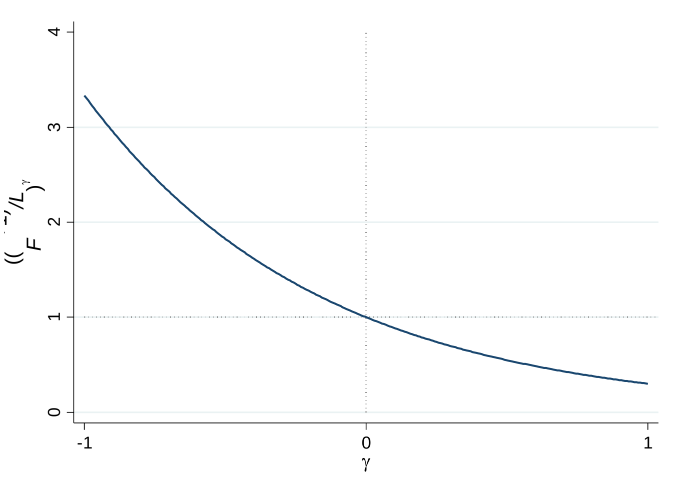
where lower case letters denote the natural logarithm of the inputs, the α^X are parameters to be estimated, and X refers to the production factors $X \in \{A, E, K, M\}$.

Next, we need a specification for the effective labour function E. Prior research proposed several functional forms. DEOLALIKAR and VIJVERBERG (1987) experimented with a generalised quadratic effective labour function, while BARDHAN (1973) and FRISVOLD (1994) employed exponential specifications. Here, we want to use the specification introduced by FRISVOLD:

$$E = L\left(\frac{F+1}{L}\right)^{\gamma}, \tag{3}$$

where E is the effective labour input in efficiency units, L is total labour time, i.e. the sum of hired and family labour time, F is family labour time, and γ is a parameter to be estimated measuring effective labour effort.

Figure 1. Effective labour as a function of $\boldsymbol{\gamma}$.



Notes: The ratio $((F+1)/L) \in [0,1]$ has been set to 0.3.

Source: Authors.

As equation (3) shows, the exponential expression $((F+1)/L)^{\gamma}$ acts as a scaling parameter for total labour time input. Following this model, the productivity of each hour of labour supplied to the farm depends on the share of family labour in total labour input and the parameter γ. If $\gamma > 0$, a higher share of family labour increases total labour productivity. If $\gamma < 0$, total labour productivity is decreased by a higher share of family labour. A given ratio of family to hired labour can decrease or increase total labour productivity, depending on whether γ is positive or negative (Figure 1). If $\gamma = 0$, there are no effects of labour force composition. We add a '1' in the numerator to allow for farms run entirely by either family or hired workers. Note that the exponential form of (3) allows for direct estimation of γ in the framework of a Cobb-Douglas function. Applying basic logarithm rules to (3) and inserting it into (2) gives:

$$y_{it} = \alpha^{A} a_{it} + \alpha^{E} l_{it} + \theta r_{it} + \alpha^{K} k_{it} + \alpha^{M} m_{it} + \omega_{it} + \varepsilon_{it}, \qquad (4)$$

where r and l are the natural logarithm of $R = ((F+1)/L)$ and L, respectively, and $\theta = \alpha^{E}\gamma$. Given this formulation, γ is equal to θ/α^{E}.

3 Empirical implementation

Factor use across firms is usually under control of the farmer. Therefore, the inputs in (4) are subject to an endogeneity problem. As a result, the ω_{it} will likely be correlated with the other input choices. The standard OLS estimator will produce biased estimates of output elasticities as it neglects the presence of ω_{it}. A typical outcome in empirical practice are upward biased elasticities for variable inputs (e.g. materials). To tackle the endogeneity problem, we need to control for ω_{it}. Several methods have been proposed to solve the problem.

A first strategy is the 'within' or fixed effects approach (MUNDLAK, 1961). Suppose we can further decompose ω_{it} in:

$$\omega_{it} = \lambda_{t_t} + \eta_i + v_{it},$$

where λ_t is a time-specific shock identical for all farms in *t*, η_i is a farm-specific fixed effect that is constant over time, and v_{it} is the remaining farm- and time-specific productivity shock. The usual approach then is to purge the fixed effects (η_i) by the so called within transformation. To do so, we substract the farm-specific means from all the variables. The λ_t are usually controlled for by incorporating time dummies into the model. However, the question remains whether the assumption of time constant fixed effects is plausible. If η_i represents factors such as management or soil quality they can be considered as time-varying over a sufficiently long period. Therefore, this assumption is likely to hold only for panels that cover rather short periods of time. Furthermore, the within transformation is known for removing too much variance from variables that exhibit little variation over time, such as land, labour and fixed capital, resulting in downward biased estimates for these factors (GRILICHES and MAIRESSE, 1998: 180-5). Especially with the effective labour function in mind, this can potentially lead to wrong conclusions.

Another promising strategy to production function estimation in agriculture that differs from the 'within' approach is to use adjustment costs as identifying information (PETRICK and KLOSS, 2013). OLLEY and PAKES (1996) were the first to assume that ω_{it} evolves with observed firm characteristics. Given a suitable proxy we can control for ω_{it}. LEVINSOHN and PETRIN (2003) propose materials as a canditate. If we further assume that materials is monotonically increasing in ω_{it} and that factor adjustment is completed within one period, we can recover the production function coefficients in two stages. First, we estimate the output elasticities of land and the parameters of the effective labour function by controlling ω_{it} with a function of materials and fixed capital. Second, we recover the output elasticities for materials and fixed capital from additional timing assumptions which are used to form appropriate orthogonality conditions. PETRICK and KLOSS (2013) note that this approach solves the endogeneity problem if the control function fully captures ω_{it}. They also give arguments where assumptions made by the control function identification strategies are somewhat questionable. Sometimes the productivity enhancing reaction to shocks in agriculture is less input use violating the assumption that materials is monotonically increasing in ω_{it}. Furthermore, the assumption of single period factor adjustment seems problematic for slowly evolving factors. Despite these theoretical issues, PETRICK and KLOSS (2013) demonstrate that the control function approach behaves robustly in empirical practice making it a plausible alternative.

A final issue issue in production function estimation is the collinearity problem (BOND and SÖDERBOM, 2005; ACKERBERG et al., 2007). If variable and intermediate inputs are chosen simultaneously, factor use across farms varies only with ω_{it} which leaves output elasticities for variable inputs unidentified. One strategy to deal with this problem again refers to heterogeneous adjustment costs for different factors and is due to WOOLDRIDGE (2009). He proposes a procedure that borrows the identification strategy from OLLEY/PAKES and LEVINSOHN/PETRIN and modifies as well as extends the moment conditions to overcome the collinearity problem. Estimation is then conducted within an instrumental variable framework using lagged variables as instruments. In the following, we present results for an estimator set that involves the 'within' estimator as a benchmark, the LEVINSOHN/PETRIN estimator and the WOOLDRIDGE (2009) modification of the latter.

4 Data

The EU's Farm Accountancy Data Network (FADN) provides a stratified farm level data set that holds accountancy data for 25 of the 27 EU member states. The stratification criteria are region, economic size and type of farming. In the present study, we only use field crop farms (TF1), to justify the assumption of a homogenous state of technology across farms. Output is measured as the total farm output in euros. The total utilized agricultural area is our land input in ha. It includes owned and rented land, and land in sharecropping. Material or working capital input is proxied by total intermediate consumption in euros. It consists of total specific costs and overheads arising from production in the accounting year. Fixed capital is approximated by using the opening valuation of assets. In this case, we took the asset value of machinery and buildings from the FADN data. In order to estimate the effective labour function (1) within a production function framework, i.e. estimating (4), we need information on hired and family labour working time separately in addition to the total labour hours. Table 1 gives definitions of the additional variables needed as well as their FADN codes. Having this data readily available we can construct the additional covariate r. To this end, we calculate $R = ((F + 1)/L)$ and take its natural logarithm. The sample of countries is selected to reflect the diverse farm sizes and structures. The range is from small-scale family farms in Italy, Poland, Spain and West Germany to medium-sized commercial farms in Denmark, France and the UK to large-scale and mostly corporate farms in East Germany and Slovakia (EUROPEAN COMMISSION, 2012).

Table 1. Variables used to calculate effective labour input.

FADN code	Variable description
SE011	Total labour input (hours)
SE016	Unpaid labour input, generally family (hours)
SE021	Paid labour input (hours)

Source: Authors, EUROPEAN COMMISSION (2011).

For every country (region in the case of Germany), we constructed a panel data set covering the years from 2001 up to 2008. The panels for Poland and Slovakia cover only five years as FADN data collection for these countries started only in 2004. Moreover, the effective panel length is reduced to four years as we use the opening valuation of fixed assets which is taken from the previous year of observations as our capital proxy. Therefore, our European database consists of 35, 296 observations. In order to be included in the estimating sample, farms had to be present for at least four years in a row (three years for Poland and Slovakia). Similar to PETRICK and KLOSS (2013), outlier analysis was performed on the basis of the fixed capital productivity per farm. Observations were excluded from the estimation if their value exceeded the median ± 1.5 the interquartile range (IQR). Table 2 summarises the number of farms for every country in our sample, the labour force composition (average percentage of family labour), and the other variable means. Our data sample covers a total of 6, 647 farms. According to the table, the dominant type of labour in EU farming is family labour. Only Slovakia displays numbers well below 50 per cent. Furthermore, there are farms entirely run on hired or family labour (e.g. in Germany, Italy and Poland).

Table 2. Sample size and variable means.

Country	Farms	Family in % of total labour	Total labour (ths hours)	Output (ths €)	Land (ha)	Materials (ths €)	Capital (ths €)
Denmark	209	84.52	2.8	180.4	122.7	110.0	840.0
France	1031	84.11	3.2	155.8	143.5	104.7	160.1
Germany (East)	292	55.86	15.6	545.6	538.9	385.0	519.4
Germany (West)	573	84.70	4.2	150.9	92.3	96.9	153.9
Italy	1362	88.55	3.6	60.6	44.7	28.4	125.2
Poland	1535	87.07	4.7	39.8	48.6	23.7	78.9
Slovakia	56	28.85	39.1	514.0	768.9	385.3	940.2
Spain	1400	90.06	2.7	40.4	72.7	19.6	31.1
United Kingdom	189	64.76	6.3	278.1	248.7	184.8	239.3

Source: Authors based on FADN data.

5 Results

To infer about the effective labour effort parameter γ, we estimate (4) employing four estimators per country. These are 1) OLS as a baseline, 2) fixed effects regression as well as the control function approaches by 3) LEVINSOHN/PETRIN (2003) (LP) and 4) WOOLDRIDGE (2009), hereafter WOOLDRIDGE/LEVINSOHN/PETRIN (WLP). All estimations were performed with Stata 12. For the LP estimator we used the user-written command `levpet` (PETRIN et al., 2004). To implement the WLP procedure we employed the `ivreg2` routine by BAUM et al. (2007) as shown in PETRIN and LEVINSOHN (2012).

The WLP estimation procedure incorporates lags up to the second order which reduces the panel length for every country by two years. For Poland and Slovakia the panel size is reduced to two years. Therefore, the WLP results for these countries should be treated with caution. Especially in the case of Slovakia an already small sample is reduced even further. To recover the standard error of γ, we use the 'delta method' (STATACORP, 2011: 1200-10).

In Table 3 we report the sample size, the point estimate of γ as well as its standard error per country and estimator. It allows for cross country and estimator comparisons. Our preferred estimator is the WLP estimator. On theoretical grounds, it corrects the biases induced by the endogeneity and collinearity problems present in production function estimation. Empirically, the results look very plausible. Compared to the LP estimator the differences in the parameter estimates of γ are negligibly small for most of the countries. This, and the fact that conclusions drawn from its test of significance do not differ in all cases makes us very confident in the control function identification approach. Furthermore, the WLP estimator is occasionally more successful in identifying the capital coefficient (not shown; detailed results tables are available upon request). Finally, it resembles more what we know about returns to scale in European field crop farming (cf. PETRICK and KLOSS, 2013). The assumption of constant technical returns to scale is rejected at the 1 percent significance level only for Italy and Spain. Returns to scale was measured as the sum of the direct production elasticities of labour, land, materials, and capital (Table A1).

Surprisingly, several countries display negative parameter estimates for land in conjunction with relatively high materials coefficients – particularly in France, Germany (East and West), Italy, Poland and Slovakia (Table A1). Variance inflation factors (not shown) suggest that there is a slight degree of multicollinearity between the land and materials input in France, West Germany, Italy, Poland and Slovakia as well as considerable multicollinearity between these inputs in East Germany. From a statistical point of view it seems that much variation of the land input is captured by the materials input. However, for farms low on materials a ceteris paribus increase in land should have a positive output effect. We therefore re-estimated (4) for the 10% sample of farms with the lowest materials intensity per ha and find a statistically significant, positive output elasticity of land in all countries but East Germany and Slovakia. These range from 0.259 (Italy) to 0.346 (France). In East Germany, the coefficient is not significantly different from zero. The negative land coefficient thus appears to be an artefact of multicollinearity between materials and land.

Regarding the significance of γ in the different member states and regions, the following picture unfolds. In Denmark, East Germany, Italy, Slovakia and Spain the coefficient of γ is not significantly different from zero, meaning that hired and family labour are perfect substitutes. The results for Slovakia are based on only 90 observations. Even though the sign changes by moving over to the LP estimate from negative to positive, the test result is not changed. The OLS and 'within' estimates also confirm this picture. Generally, labour seems to be no scarce factor in Slovakia and East Germany as their labour coefficients are not statistically different from zero (Table A1). Both exhibit large-scale farming structures. The small- and medium scale agricultural structures of West Germany, Poland and France exhibit negative and significant γ's. This means that hired labour is more productive than family labour but this productivity differential decreases as the farm operation increasingly relies on hired labour. It is probably here where hired labour specialises on high productivity tasks and/or family labour focuses on low productivity tasks. The size of the parameter for West Germany suggests that hired labour is much more productive than family labour. The classical case with family members being more productive than hired labour is only observed for the United Kingdom. Here, we have an argument for labour supervision. Finally, the distribution of labour force heterogeneity across the sample countries suggests that mainly small- to medium- scaled agrarian structures (with the exception of Italy) display differing effects on productivity for the two types labour

Compared to the WLP approach, the fixed effects regression results detect labour force heterogeneity only in one case, namely Poland. A possible reason could be that after transforming the variables, there is not much left to explain, i.e. too much variance was removed from the variables in the effective labour effort function. Furthermore, with regards to the OLS estimator, it seems that labour force heterogeneity was found in too many cases. As this estimator neglects the presence of endogeneity the estimates are most likely biased. To sum up, the choice of estimator matters a lot in in inferring about labour force heterogeneity. Therefore, we should resort to methods that can mitigate problems immanent in production function estimation.

Table 3. Effective labour effort parameter (γ) in comparison.

Country	**OLS**			**'Within'**		
	N	γ	SE	N	γ	SE
Denmark	1027	0.329***	0.057	1027	0.272	0.220
France	6361	-0.436***	0.090	6361	-0.649	0.401
Germany (East)	1740	0.491	0.441	1740	-4.828	17.709
Germany (West)	3603	-1.370***	0.264	3603	-3.955	4.434
Italy	6415	-0.393***	0.113	6415	-0.369	0.395
Poland	5635	0.002	0.049	5635	-0.931*	0.541
Slovakia	202	0.237**	0.111	202	0.861	0.634
Spain	9317	-0.024	0.016	9317	-0.074	0.060
United Kingdom	996	0.175***	0.040	996	0.111	0.165
Country	**Levinsohn/Petrin**			**Wooldridge/Levinsohn/Petrin**		
	N	γ	SE	N	γ	SE
Denmark	1027	0.145	0.096	609	0.132	0.099
France	6361	-0.411***	0.131	4299	-0.457**	0.208
Germany (East)	1740	0.240	0.214	1156	0.147	0.216
Germany (West)	3603	-1.322***	0.444	2457	-1.339***	0.499
Italy	6415	-0.379	0.248	3691	-0.177	0.181
Poland	5635	-0.292***	0.089	2565	-0.498**	0.204
Slovakia	202	1.150	4.900	90	-0.287	0.455
Spain	9317	-0.033	0.021	6518	0.028	0.031
United Kingdom	996	0.203***	0.057	618	0.242**	0.095

Notes: Year dummies included in all models. *** (**, *) significant at the 1% (5%, 10%) level

Source: Authors.

6 Conclusions

In this study we assessed the heterogeneity of family and hired labour in European agriculture. To this end, we took a sample of eight EU countries and estimated augmented production functions that allow testing for labour force heterogeneity using farm-level FADN data. The results unveil a diverse picture.

Contrary to the received wisdom, we find that hired labour is more productive than family labour in the small- and medium-scale agrarian structures of France, West Germany and Poland. According to our estimates, hired labour performs the high productivity tasks in these countries. In such a situation, an increase in reliance on hired labour or the shift of family labour to more productive tasks could raise productivity. In the majority of countries we found no evidence for labour force heterogeneity. Amongst the countries in this group are also large-scale farming structures such as East Germany and Slovakia. For the United Kingdom, we observe that total labour productivity is higher when there are more family members in the labour force. In this case, supervision by family members apparently does increase productivity. We regard it an interesting question for future research to find out why hired labour in arable farming is so productive in France, West Germany and Poland, three countries traditionally characterised by family farms. One possible explanation is that farm technology, e.g. modern tractors and other field machinery using precision farming methods, has reached such levels of sophistication that benefits from labour specialisation can be reaped. Farming in the UK, on the other hand, traditionally displays higher levels of hired labour. This pool of workers may to a larger extent consist of lower qualified personnel subject to the classical incentive problems.

The results have implications for future theoretical and empirical work. Most importantly, our results call into question the general validity of one of the received family farm theory's main pillars, i.e. the dominant effect of supervision costs on hired labour productivity. Countries regarded as traditional strongholds of the family farm have apparently crossed a technological threshold where specialisation of hired labour overcompensates the negative effects of workers' moral hazard. Factors such as the increasing importance of non-traditional and non-agricultural sources of farm household income are likely reinforcing this trend. On the other hand, the assumption of constant technical returns to scale is confirmed.

In classical production function estimation, labour input is measured as the sum of both, hired and family workers. Given the evidence on labour force heterogeneity in some countries with different effects on productivity, their heterogeneity should not be ignored. Such a treatment will improve model fit and avoid misspecification.

Finally, this work is also a plea for refined methods that control for the problems in production function estimation. Endogeneity and collinearity problems potentially lead to wrong inference. The OLS estimator neglecting the presence of endogeneity seems to detect labour force heterogeneity in too many cases. The fixed effects regression, while dealing with endogeneity, seems to find evidence for labour force heterogeneity in too few situations. A possible reason could be that the 'within 'transformation of the variables does not leave enough variance in the data. Therefore, the control function framework introduced by OLLEY/PAKES and then further refined by LEVINSOHN/PETRIN and WOOLDRIDGE are a promising alternative to traditional estimators. The very similar results that we obtained from the LEVISOHN/PETRIN and WOOLDRIDGE/LEVINSOHN/PETRIN approaches seem to strengthen their validity on empirical grounds, besides being plausible in the theoretical domain.

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Appendix

Table A1. Results of production function estimations for the Wooldridge/Levinsohn/Petrin estimator per country.

	Denmark		**France**		**Germany (East)**		**Germany (West)**		**Italy**	
	Coeff	SE	Coeff	SE	Coeff	SE	Coeff	SE	Coeff	SE
Labour	0.548***	0.082	0.129***	0.022	0.038	0.051	0.129***	0.031	0.326***	0.026
Ratio (r)	0.072	0.058	-0.059***	0.020	0.006	0.008	-0.173***	0.030	-0.058	0.057
Land	0.145*	0.075	-0.059***	0.017	-0.231***	0.058	-0.071***	0.020	-0.045***	0.017
Materials	0.468*	0.274	0.921***	0.079	1.417***	0.243	0.807***	0.095	0.530***	0.091
Capital	0.111*	0.068	0.109***	0.014	0.062	0.055	0.129***	0.031	0.013	0.030
N	609		4299		1156		2457		3691	
Elasticity of scale	1.272***	0.231	1.100***	0.073	1.286***	0.202	0.946	0.089	0.825***	0.088
p-value const. ret. to scale	0.239		0.173		0.156		0.548		0.005	
p-value coeff. jointly zero	<0.001		<0.001		<0.001		<0.001		<0.001	

Table A1 continued.

	Poland		**Slovakia**		**Spain**		**United Kingdom**	
	Coeff	SE	Coeff	SE	Coeff	SE	Coeff	SE
Labour	0.143***	0.022	-0.105	0.132	0.449***	0.027	0.188***	0.046
Ratio (r)	-0.071***	0.024	0.030*	0.015	0.013	0.014	0.045**	0.018
Land	-0.067***	0.021	-0.328**	0.160	0.042**	0.016	0.092	0.062
Materials	1.077***	0.100	1.914***	0.560	0.760***	0.061	0.888*	0.470
Capital	0.021	0.043	-0.111	0.187	0.078***	0.027	0.112*	0.063
N	2565		90		6518		618	
Elasticity of scale	1.174***	0.083	1.370***	0.342	1.328***	0.066	1.280***	0.455
p-value const. ret. to scale	0.036		0.280		<0.001		0.539	
p-value coeff. jointly zero	<0.001		<0.001		<0.001		<0.001	

Notes: Year dummies included in all models. *** (**, *) significantly different from zero at the 1% (5%, 10%) level, based on standard errors robust to clustering in groups.

Source: Authors' calculations.