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Income mobility and income inequality in Scottish agriculture

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Abstract: *The paper explores the distributional consequences of farm income mobility in Scotland, focusing on the extent to which farm income inequality is a chronic as opposed to a temporary phenomenon and on the nature of the dynamic processes driving changes in farm income inequality over time. The empirical results reveal that the majority of farm income inequality was long-run or structural in nature, reflecting differences in both farm business size and farm-specific factors such as land quality, managerial ability and business structures. Evidence of absolute convergence in farm incomes is explained by short-run adjustments towards equilibrium or target incomes conditional upon prices, technology and farm business size, with farm business growth conditional upon survival found to have had no significant redistributive effect.*

Keywords*: farm incomes, income mobility, income inequality, Scotland.*

JEL classifications*: D31, D63, Q18*

1. Introduction

Income mobility at the farm level is a major driver of changes in the distribution of farm incomes at the sectoral level. However, this connection has received less attention than it deserves in empirical work. The main aim of this paper is to explore the distributional implications of farm income changes in Scottish agriculture and thus address two distinct but interrelated issues that are of relevance to policy makers. Firstly, we consider the extent to which farm income inequality is a short-run phenomenon, due to transitory income shocks, as opposed to a chronic problem owing to 'permanent' differences in income between farms. If income inequality is largely transitory then this becomes less of a policy concern, although there may still be a call for intervention to provide insurance measures or compensation aid in the event of disasters. Conversely, if income inequality reflects 'permanent' differences in

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income then attention should focus more on targeting support to those deemed in need of assistance. Secondly, we consider whether the pattern of farm income growth has been systematically biased in favour of high or low income farms and the extent to which this may have been driven by changes in the farm business size structure. If income growth is concentrated among farms at the top of the income distribution, this may call for structural measures to alleviate the constraints that trap some farms in a low income condition. Even if income growth rates are higher on average on low income than high income farms, income inequality may rise due to the dispersion in individual growth rates among farms with similar initial levels of income.

We add to a relatively small body of literature that makes use of longitudinal data to analyse the micro-dynamics of farm incomes. In particular, a number of previous studies (e.g. Hegrenes *et al*., 2001, Meuwissen et. al., 2008) have provided evidence of considerable volatility in individual farm incomes, thereby emphasising the importance of using multiyear average data to draw meaningful conclusions about the living standards of individual farmers. We complement this work by proposing an alternative to the Shorrocks (1978) mobility index that measures the extent to which incomes are equalized over the longer term based on equilibrium rather than multiyear average incomes. Phimister et al. (2004) have further explored the impact of the movements of farms within the income distribution on the persistence of poverty in Scottish agriculture, building on an older tradition of modelling income mobility within agriculture using transition matrices (see e.g. Meuwissen et. al (2008) for a recent example). We more broadly characterise the process of distributional change underlying the evolution of cross-sectional inequality in farm incomes over time by adapting and extending the regression-based decomposition procedures proposed in Allanson and Petrie (2013). Specifically, we decompose changes in cross-sectional farm income inequality over time into vertical and horizontal income mobility components, where the former addresses the question of whether high or low income farms have benefited more from farm income growth and the latter captures the effects of the reranking of farms in the income distribution. We further identify the contribution of farm business size changes to vertical mobility based on a dynamic model of farm incomes that explicitly takes into account the impact of both systematic factors and transitory shocks.

The paper is structured as follows. Section 2 establishes the empirical setting, providing a brief description of the Farm Accounts Survey (FAS) dataset used in the study and reporting some basic descriptive statistics for our preferred definition of farm income. Section 3 employs the Shorrocks (1978) mobility index to provide a first look at the extent to which farm incomes are equalised over the longer term. Section 4 characterizes changes in farm income inequality over time, identifying both the extent to which vertical mobility may have been driven by farm business growth and the factors that contribute to equilibrium or structural inequality. The section also presents our alternative measure of the potential for the equalisation of farm incomes over the longer term. The final section concludes with a discussion of the empirical findings in the light of the most recent round of CAP reform.

2. Data

Longitudinal data are required to explore the distributional implications of farm income mobility. For example, if half the farms in Scotland are always poor while half are rich then it will not be possible to determine whether it is always the same farms in each category by examining changes in cross-sectional data over time. We construct an unbalanced panel of farms using data from the Scottish Farm Accounts Survey (FAS) relating to the farming years from 1995/66 to 2009/10. The study examines mobility both over the whole of the study period and for the two sub-periods defined by the introduction of the Single Farm Payment scheme in 2005/06.

The FAS is an annual survey of around 500 full-time farms carried out on behalf of the Scottish Government and provides the main source of microeconomic data on farm businesses in Scotland, with data collected on a range of physical and financial variables. The survey is conducted on an accounting year basis with a typical year-end in early March. Thus, for example, the 1995/96 FAS centres on the 1995 production and subsidy year. The FAS is based on a stratified simple random sample, with farms chosen randomly to be representative of their economic size and type. Economic size is measured in terms of standard gross margin (SGM) prior to 2003/04 and standard labour requirement thereafter, while the farm type classification is based on the relative importance of the various crop and livestock enterprises in terms of SGM.[1]

The FAS potentially provides a rich source of information for the analysis of farm income mobility since farms, once recruited, can stay in the survey for an unlimited length of time (Scottish Government, 2012a). However farms in the survey that experience significant

[1] The sampling frame excludes small farms less than 8 Economic Size Units (ESUs) prior to 2003/04 and 0.5 Standard Labour Requirements (SLRs) thereafter; specialist livestock units larger than 200 ESU prior to 2003/04; and certain minor farm types (most notably horticulture and specialist pig and poultry farms).

structural change (such as amalgamation with another farm) are assigned a new identifier and the sample is therefore subject to a 'virtual' form of selective attrition that could bias mobility estimates.[2] We address this problem by making use of information supplied by the data provider on the linkage of identifiers between years to assign a unique number to each farm so long as it remains in the sample. The FAS is not informative of farm entry and exit processes since farms recruited to the survey are not necessarily new businesses and no reasons are given as to why farms leave the survey. The analysis is based on an unbalanced panel of 933 farms, of which 174 were present over the whole 15 year period and with a median duration of 7 years.

Probability weights are used throughout our analysis with these being based on farm numbers enumerated by size and type in the annual June Agricultural Census. The weighted sample is therefore representative of the population of full-time farms in Scotland in each year, with a sampling fraction of between 3% and 4% over most of the study period.[3] Standard errors for all mean, inequality and mobility measures are generated using bootstrap procedures that reflect the sample design. In particular, bootstrap standard errors for the mobility indices are obtained by the resampling of clusters within each stratification class, where each cluster consists of all observations on a single farm, and therefore allows for the possible correlation of individual farm incomes across years.

The farm income measure used in the current study is Cash Income, which represents the cash return to the group with an entrepreneurial interest in the farm for their manual and managerial labour and on their investment in the business (Scottish Government, 2012a) and is defined as the difference between trading revenue (sales of livestock, livestock products, crops, subsidy and payments, revenue from diversified activities, etc) and trading expenditure (variable costs, general overheads, fuel, repairs, rent paid, paid labour, etc).[4] Of the various

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[2] In contrast, farms in the Farm Business Survey in England and Wales retain their unique number except in exceptional circumstances, such as the farm splitting into two units that both continue to participate in the survey, but even in this case the larger unit will retain the original number (cf. DEFRA, 2014).

[3] Farms that were directly affected by foot and mouth disease culls and compensation are excluded from the analysis, but the resultant sub-samples for 2001/02 and 2002/03 are nevertheless sufficient "to give a representative picture of full-time Scottish farm businesses" in these years (SEERAD, 2003, 2004).

[4] Note that cash income, unlike cash flow, does not take account of net investment spending.

alternative FAS indicators of farm income, Cash Income may be seen to correspond most closely to the farm income position as perceived by the farmer (cf. DEFRA, 2002, Appendix 3), but it is important to recognise that it does not include non-farm sources of income about which FAS collects only limited information. The analysis is conducted at the farm level rather than per unit of unpaid labour because of doubts concerning the relevance and reliability of data on the unpaid labour input in the UK context (see Hill 1991).

The inequality measure used throughout the study is the Gini coefficient. Let $G(y_t) = 2\,\mathrm{cov}(y_{it}, R_{it}) / \bar{y}_t$ be the Gini coefficient of incomes in year t, where y_{it} is the income of farm i (i=1,….N) in year t, $\bar{y}_t$ is average income, and R_{it} is the farm's relative rank in the year t income distribution. $G(y_t)$ is invariant to equiproportionate changes in all incomes, taking a value of zero when all individual farm incomes are identical and of one when all income accrues to one farm and all other farms receive nothing. The first two columns of Table 1 report the mean and Gini coefficient of income for each production year between 1995 and 2009. Farm incomes fell after 1996 due to a combination of factors including a strong pound, weak world commodity prices and the impact of bovine spongiform encephalopathy; and only recovered gradually following the end of the Foot and Mouth Disease outbreak in 2001. Indeed average incomes did not rise in nominal terms above the levels observed in 1996 before 2007 and were only £5000 per farm higher in 2009, despite the growth in average farm business size – as measured in terms of total SGM, based on Scottish average gross margins for the years 1998 to 2002[5] – reported in column 3. The coefficient of variation of average annual incomes was 20% over the study period, but this measure of the aggregate instability of the income distribution as a whole may tell us little about the experience of individual farms, which will also be determined by the effects of the movement of farms within the distribution due to idiosyncratic income shocks. Changes in the Gini coefficient reflect changes in both the absolute dispersion and mean level of incomes, with relative inequality generally higher in years of lower average incomes.

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[5] SGMs are representative of the level of gross margin – enterprise output less variable costs – that could be expected on an average farm under 'normal' conditions and are calculated using SGM coefficients per unit area of crops and per head of livestock. We ensure comparability over time by using the most recent set of SGM coefficients available to calculate SGMs for the entire period.

3. Is farm income inequality a transitory or chronic problem?

Income inequality, as measured by the Gini coefficient, will typically fall if income is measured over a longer period due to the reranking of farms within the income distribution. The extent of such equalization if the measurement period is extended from one to *T* years is captured by the mobility index due to Shorrocks (1978):

$$M_T = 1 - \frac{G(y_A)}{\sum_{t=1}^{T} w_t G(y_t)} \tag{1}$$

where $G(y_A) = 2\operatorname{cov}\left(y_{iA}, R_{iA}\right) \big/ \bar{y}_A$ is the Gini coefficient of individual average incomes $y_{iA} = \sum_{t=1}^{T} y_{it} \big/ T$ calculated over the *T*-year period $t=1,\ldots,T$; R_{iA} are the corresponding relative ranks; $\bar{y}_A = \sum_{i=1}^{N} \bar{y}_{iA} \big/ N$ is overall average income over the entire period; and $w_t = \bar{y}_t \big/ \bar{y}_A$ are a set of weights that sum to one by construction. $M_T = 0$ by definition if T=1. For T>1, the index will equal one when longer-term incomes are exactly equalised over the measurement period such that the *T*-year Gini coefficient is equal to zero, and will equal zero in the absence of exchange mobility such that each farm occupies the same rank in all *T* annual income distributions.

Hence if inequality is largely a short-run or temporary phenomenon due to transitory idiosyncratic income shocks then the mobility index will take a value close to one, whereas if inequality largely arises from long-term or permanent differences between farms then the index will take a value close to zero. From this perspective, greater mobility may not favour risk-averse farmers if it is associated with greater uncertainty due to more pronounced income fluctuations, even though the equalisation of long-term incomes *per se* may be seen as a socially desirable goal. Jantti and Jenkins (2015, p.814) observe that the Shorrocks index can be interpreted as a measure of risk if incomes are given as the simple sum of a fixed individual-level permanent component, approximated by *T*-year average income, and an idiosyncratic transitory component that is *ex-ante* unknown. In practice, lower values of M_T need not necessarily imply lower levels of risk due to the correlation of income shocks across the farm sector as a result of common factors such as price movements and weather conditions.

The final column of Table 1 gives values of M_T as the measurement period is extended from the base year of 1995, initially aggregating over the first two years for all farms present in both years, then the first three years and so on.[6] Thus the index value of 0.057 for T=2 implies that averaging incomes over 1995 and 1996 reduces inequality by 5.7% compared to the weighted average of the Gini coefficients for the two years. Phimister *et al.* (2004) have previously reported an 8% fall in the Gini coefficient for Scotland over the period 1988 to 1999 if cash income values are calculated using rolling two-year individual farm averages. M_T tends to increase as the length of the measurement period is extended but approaches an upper limiting value of about 12% after about 10 years, with no further equalisation once relative incomes have approached their long-term or permanent values. Thus the overwhelming bulk of cross-sectional inequality, as measured by annual Gini coefficients, would appear to have been long-term or chronic in nature inasmuch as it reflected permanent differences in incomes between farms. The choice of alternative base years produced broadly similar findings (not reported), with an average value for T=2 of 6.6% over all possible base years between 1995 and 2008, and with upper limiting values in the range of 12% to 16%.

4. Characterizing processes of inequality change

Changes in income inequality over time in a fixed population of farms are related to the pattern of income growth across the income range and the reranking of farms within the income distribution. In particular, following Jenkins and van Kerm (2006; see also Kakwani, 1984), the change in the Gini coefficient from some base year s to a final year f may be written as:

$$G(y_f) - G(y_s) = \left\{G(y_f) - CI(y_f, R_s)\right\} + \left\{CI(y_f, R_s) - G(y_s)\right\} = M_H + M_V \qquad (2)$$

where $CI(y_f, R_s) = 2\,\mathrm{cov}\left(y_{if}, R_{is}\right) \Big/ \bar{y}_f$ is defined as the concentration index (CI) of final year incomes ranked by positions in the base year income distribution; and the vertical and horizontal mobility indices, M_V and M_H respectively, are discussed further below.

M_V provides a measure of vertical mobility that addresses the question of whether the distribution of income changes favours farms with initially low or high incomes and thus provides a natural counterpart to $G(y_t)$ which addresses the distribution of income between

[6] Limiting the entire analysis to the 174 farms present in all 15 years leads to lower values of M_T for small T, though the estimate of M_T for T=15 is identical to that reported in Table 1 by construction.

low and high income farms. M_V will be zero if expected income changes are unrelated to base year income and will be negative if expected income changes are equalising in relative terms, which will be the case if low income farms in the base year experience either larger relative gains on average than high income farms or smaller relative losses. M_V can usefully be written as the product of progressivity and scale indices, P and q respectively, such that $M_V = Pq$. The progressivity of income changes is captured by the disproportionality index $P = CI\left(y_f - y_s, R_s\right) - G(y_s)$ where $CI\left(y_f - y_s, R_s\right) = 2\operatorname{cov}\left(y_{if} - y_{is}, R_{is}\right) / \left(\bar{y}_f - \bar{y}_s\right)$ is the CI of income changes $(y_{if} - y_{is})$ ranked by base year incomes. For any given P, the gross redistributive effect M_V is proportional to the relative magnitude of income changes as measured by the scale factor $q = \left(\bar{y}_f - \bar{y}_s\right)\big/\bar{y}_f$. Note that negative values of P imply that expected income changes will be equalising if incomes are growing on average, but diverging if incomes are falling.

$M_H = 2\operatorname{cov}\left(y_{if}, R_{if} - R_{is}\right)\big/\bar{y}_f$ is the Atkinson−Plotnick reranking index, which captures the effect of the movement of farms within the income distribution. M_H is non-negative by definition (see Lambert, 2001), implying that any reranking that does occur has an adverse impact on the overall redistributive effect of the income changes. Thus farm income growth will only reduce inequality if expected income changes favour lower income farms and the resultant vertical mobility effect is not swamped by any offsetting horizontal mobility effect due to the reranking of farms.

The top panel of Table 2 presents the decomposition of annual changes in the Gini coefficient into vertical and horizontal components based on (2), where the results are generated using observations on all farms present in both the base and final year, and therefore are not strictly comparable either with the annual summary statistics presented in Table 1 or between pairs of years. These results reveal three main points. First, the vertical mobility index M_V is significantly negative in all cases, indicating that expected annual income changes conditional upon initial income had an equalising effect throughout the period. Thus farms with low incomes in one year experienced on average over the following year either larger relative income gains than high income farms or smaller relative losses. However, most of the estimates of the progressivity index P and some of the estimated scale factors q are not significantly different from zero. Second, the horizontal mobility index M_H is significantly positive in all cases, reflecting the impact of idiosyncratic income shocks on the ranking of farms in the income distribution between successive years. Third, the equalising effect of expected income changes was only sufficient to outweigh the diverging impact of re-ranking in some years, with no clear trend in the level of farm income inequality over the entire period.

The finding that expected income changes were not independent of initial incomes needs to be treated with some caution as the apparent progressivity of farm income growth may simply reflect regression toward the mean if, as seems likely to some extent, individual farm incomes are subject to idiosyncratic shocks (or measurement errors) that are uncorrelated over time such that large positive (negative) shocks to income in a particular year are offset by slow (fast) income growth in subsequent years.[7] We employ a number of alternative strategies to investigate whether the observed progressivity of income growth is in fact spurious.

First we consider multiyear rather than annual changes in income inequality on the assumption that extending the measurement period is likely to reduce the importance of the transitory component in any observed change. The bottom panel of Table 2 presents the results of these multiyear decomposition analyses, which have been generated using observations on all farms present throughout the relevant measurement period. We find that vertical mobility was significantly negative for all multi-year periods and, in particular, that low income farms in 1995 experienced higher average rates of income growth over the full study period than high income farms. Nevertheless inequality rose over the first sub-period and the full period, though not over the second sub-period from 2005 to 2009.

Our other two robustness checks employ alternative techniques to mitigate the potential for bias due to transitory shocks in the estimation of vertical mobility. First we follow common practice in the mobility literature by measuring income as a three-year centred moving average to reduce the impact of transitory variability (see, e.g. Solon, 2002). Second, we employ the so-called 'IV' approach proposed by Jenkins and van Kerm (2011) to purge the rank variable of income shocks by replacing observations on ranks in the base year distribution with estimates based on ranks in the distribution of the average of one year lag

[7] For example, if incomes are given as the simple sum of a fixed farm-specific permanent component and an idiosyncratic transitory component then a farm that is subject to a positive (negative) income shock in one year can expect to receive a lower (higher) income in the following year given that the expected value of the transitory component is zero. So, despite there being no association between incomes in the two years, there is a correlation between incomes in the first year and the subsequent changes in income.

and lead incomes.[8] Table 3 presents the results of these alternative estimates of M_V, where we examine changes both over successive three-year periods and the full (truncated) study period. These show that when using smoothed income data then expected income growth conditional on initial income is still significantly equalising in most cases, but the extent of vertical mobility is typically reduced somewhat. Conversely only three of the 'IV' estimates of M_V are significantly different from zero, one negative and two positive, which might suggest that neither low nor high income farms were systematically favoured by the pattern of income changes.

In conclusion the findings provide some evidence against the hypothesis that expected income changes were independent of initial incomes although the results of the robustness tests are not unequivocal. In particular, it would appear that transitory shocks are unlikely to account for all of the observed bias of annual income growth rates in favour of lower income farms conditional upon survival.

4.1 To what extent has vertical mobility been driven by changes in farm business size?

Bakucs *et al.* (2103) observe that empirical research on the relationship between farm size and farm growth has yielded rather contradictory results, with a number of studies (e.g. Shapiro *et al.* 1987; Weiss, 1999) having found evidence that smaller farms tend to grow faster than larger ones. This sub-section extends the preceding analysis by considering the extent to which observed levels of vertical income mobility might have been driven by changes in farm business size. Specifically, we follow the empirical strategy adopted in Allanson and Petrie (2013) to identify the determinants of vertical mobility by estimating a dynamic model of individual farm incomes. We first consider the specification of the dynamic model before showing how the estimates may be used in the decomposition of the vertical mobility index M_V.

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[8] This is not a conventional instrumental variables approach, though Jenkins and van Kerm (2011) argue that it is analogous to one. Given that $M_V = CI(y_f, R_s) - G(y_s)$, $CI(y_f, R_s)$ and $G(y_s)$ can each be estimated using the 'convenient regression approach' of Kakwani *et al.* (1997) as the response coefficient from a simple regression of a normalised measure of income on base year rank, with the so-called 'IV' procedure intended to eliminate possible correlation between the 'explanatory' rank variable and the 'error term' in this regression.

Let $\pi(p_t; z_{it}; \tau_t; \mu_i)$ be the income-generating potential of farm *i* (*i*=1, … *N*) in year *t* (*t*=1, … *T*), where p_t is a vector of input and output prices that are assumed to be common across farms; z_{it} represents farm business size; τ_t the state of technology common to all farms; and μ_i captures unobservable factors - such as land quality - that may be assumed to be constant by farm over the study period. Assuming a simple linear functional form with interaction terms[9] and replacing the common time-varying factors, prices and technology, by annual dummies yields the target income function:

$$y_{it}^{*} = \pi(p_t; z_{it}; \tau_t; \mu_i) = \beta_0 + \beta_{1t} d_t + \beta_2 z_{it} + \beta_{3t} d_t z_{it} + \mu_i; \qquad t=1,\ldots T \tag{3}$$

Observed farm incomes y_{it} may be expected to diverge from target incomes y_{it}^{*} due to both farm production system adjustment costs and the influence of transitory idiosyncratic shocks. Accordingly we incorporate (3) into a first-order Error Correction Model (ECM):

$$\begin{aligned}
\Delta y_{i,t+1} &= \left(y_{i,t+1} - y_{it}\right) = \delta_1 \Delta z_{i,t+1} + \delta_{2t} d_t \Delta z_{i,t+1} + \lambda\left(y_{it}^{*} - y_{it}\right) + \varepsilon_{i,t+1} \\
&= \lambda\left(\beta_0 + \lambda\mu_i + \lambda\beta_{1t} d_t\right) + \left(\delta_1 + \delta_{2t} d_t\right)\Delta z_{i,t+1} + \lambda\left(\beta_2 + \beta_{3t} d_t\right) z_{it} - \lambda y_{it} + \varepsilon_{i,t+1} \\
&\equiv \alpha_{0it} + \alpha_{1t}\Delta z_{i,t+1} + \alpha_{2t} z_{it} - \lambda y_{it} + \varepsilon_{i,t+1}; \qquad i = 1,\ldots N;\ \ t = 1,\ldots T-1
\end{aligned} \tag{4}$$

where $\left(y_{it}^{*} - y_{it}\right)$ corresponds to the 'equilibrium error' in the current period and λ $(0 \le \lambda \le 1)$ determines the rate of adjustment to equilibrium. Hence, annual changes in income depend on the effects of contemporaneous changes in farm business size $\Delta z_{i,t+1}$, where the size of these effects may vary between years; the initial extent of any disequilibrium in income; and the size of the idiosyncratic income shock $\varepsilon_{i,t+1}$.

The ECM is obtained as a reparameterisation of the first-order autoregressive distributed lag (ARDL) model, which nests the partial adjustment, first-order autoregressive and static models as special cases. For analytical purposes, the main attraction of this dynamic specification is the clear distinction between the short-run dynamics and the implied long-run income relationship. In particular, it is possible using the ECM to identify both the short-term impact on farm income inequality due to contemporaneous changes in farm business size and also which factors contribute to equilibrium or structural inequality.

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[9] The specification could be extended to include higher-order terms in z_{it} but these proved to be insignificant in the empirical application.

With respect to the determinants of vertical mobility between consecutive periods, it is readily shown (see Allanson and Petrie, 2013) that if $f=s+1$ then M_V in (2) may be decomposed using (4) to yield:

$$\begin{aligned} M_V = Pq = P_{\Delta z} q_{\Delta z} + P_{EqE} q_{EqE} + P_e q_e &= \left(CI(\Delta z_f, R_s) - G(y_s) \right) \frac{\hat{\alpha}_{1s} \overline{\Delta z_f}}{\overline{y}_f} \\ &+ \left(CI(\hat{y}_s^* - y_s, R_s) - G(y_s) \right) \frac{\hat{\lambda}\, \overline{(\hat{y}_s^* - y_s)}}{\overline{y}_f} + \left(CI(e_f, R_s) - G(y_s) \right) \frac{\overline{\hat{\varepsilon}_f}}{\overline{y}_f} \end{aligned} \tag{5}$$

where a hat over a variable or parameter indicates that it is an estimate and a bar denotes a sample mean, such that $\overline{\Delta z_f}$, $\overline{(\hat{y}_s^* - y_s)}$ and $\overline{\hat{\varepsilon}_f}$ are respectively the average change in farm business size, mean predicted equilibrium error and regression residual, with $CI(\Delta z_f, R_s) = 2\operatorname{cov}(\Delta z_{if}, R_{is}) \,/\, \overline{\Delta z_f}$, $CI(\hat{y}_s^* - y_s, R_s) = 2\operatorname{cov}\left((\hat{y}_{is}^* - y_{is}, R_{is}) \right) / \overline{(\hat{y}_s^* - y_s)}$ and $CI(\hat{\varepsilon}_f, R_s) = 2\operatorname{cov}(\hat{\varepsilon}_{if}, R_{is}) \,/\, \overline{\hat{\varepsilon}_f}$ being the corresponding CI's ranked by base year income.

Hence M_V is given in (5) as the sum of contributions due to changes in farm business size, the predicted equilibrium error and contemporaneous income shocks, where each contribution is expressed in terms of the scale and progressivity of the income changes due to that factor. The intuitive interpretation is that a factor can only contribute to vertical mobility M_V if it is statistically associated with changes in income and the distribution of those changes among high and low income farms is different from the initial distribution of income. In particular, vertical mobility due to farm business growth $P_{\Delta z} q_{\Delta z}$ will be equalising if farm business growth is positively associated with farm income growth and a larger share of the resultant income gains accrue to low income farms than their base year share of income, such that $q_{\Delta z}$ is positive and $P_{\Delta z}$ negative. Similarly the contribution of the 'error correction' mechanism to vertical mobility $P_{EqE} q_{EqE}$ will depend on the scale and progressivity of the resultant income changes,[10] where the process of adjustment towards equilibrium or

[10] $P_{EqE} q_{EqE}$ in (5) could be further broken down using (3) to identify the 'apparent' contribution of farm business size z_s to M_V through the disequilibrium adjustment process as $P_{EqE(z)} q_{EqE(z)} = \left(CI(z_s, R_s) - G(y_s) \right) \left(\hat{\alpha}_{2s} \overline{z}_s / \overline{y}_f \right)$. But this would be misleading as the causes of the base year equilibrium error are unknown: for example, the disequilibrium may have arisen due to prior changes in prices or technology not farm business size.

target income levels may generally be expected to have a negative impact on M_V and hence reduce inequality.[11]

We extend this decomposition analysis to consider the determinants of vertical mobility over a multiyear period. Thus, if $f=s+m$ with $m\geq 1$ then income changes over this period can be expressed in terms of the dynamic income model as:

$$y_{if} - y_{is} = \Psi_i + \Lambda\left(y^*_{is} - y_{is}\right) + \Sigma_i + \Omega_i + \Theta_i \tag{6}$$

where: $\Psi_i = \alpha_{1,s+m}\Delta z_{i,s+m} + \sum_{k=1}^{m-1}\left((1-\lambda)^{m-k}\alpha_{1,s+k} + \left(\sum_{j=0}^{m-(k+1)}(1-\lambda)^j \alpha_{2s}\right)\right)\Delta z_{i,s+k}$;

$$\Lambda = \sum_{k=1}^{m}(1-\lambda)^{m-k}\lambda\,; \qquad \Sigma_i = \sum_{k=1}^{m}(1-\lambda)^{m-k}\varepsilon_{i,s+k}\,;$$

$$\Omega_i = \sum_{k=1}^{m-1}(1-\lambda)^{m-k}\left(\left(\alpha_{1,s+k} - \alpha_{1,s}\right) + \left(\alpha_{2,s+k} - \alpha_{2,s}\right)z_{is}\right)$$

$$\Theta_i = \sum_{k=1}^{m-1}\left(\sum_{j=0}^{m-(k+1)}(1-\lambda)^j\left(\alpha_{2,s+k-(j+1)} - \alpha_{2s}\right)\right)\Delta z_{i,s+k}$$

and which reduces to (4) if $m=1$ with $s\equiv t$. Hence (5) may be generalised to give:

$$\begin{aligned} M_V &= Pq = P^m_{\Delta z}q^m_{\Delta z} + P^m_{EqE}q^m_{EqE} + P^m_e q^m_e + P^m_{\Delta\Omega}q^m_{\Delta\Omega} + P^m_{\Delta z\Delta\Omega}q^m_{\Delta z\Delta\Omega} \\ &= \left(CI(\hat{\Psi}, R_s) - G(y_s)\right)\frac{\bar{\hat{\Psi}}}{\bar{y}_f} + \left(CI(\hat{y}^*_s - y_s, R_s) - G(y_s)\right)\frac{\overline{\hat{\Lambda}(\hat{y}^*_s - y_s)}}{\bar{y}_f} \\ &\quad + \left(CI(\hat{\Sigma}, R_s) - G(y_s)\right)\frac{\bar{\hat{\Sigma}}}{\bar{y}_f} + \left(CI(\hat{\Omega}, R_s) - G(y_s)\right)\frac{\bar{\hat{\Omega}}}{\bar{y}_f} + \left(CI(\hat{\Theta}, R_s) - G(y_s)\right)\frac{\bar{\hat{\Theta}}}{\bar{y}_f} \end{aligned} \tag{7}$$

where $\hat{\Psi}$, $\hat{\Lambda}$, $\hat{\Sigma}$, $\hat{\Omega}$ and $\hat{\Theta}$ are estimates of the corresponding entities in (6), with mean values denoted by bars. Therefore, vertical mobility in any given multiyear period is determined, as before, by the (cumulative) effects of changes in farm business size over the period, the equilibrium error in the base year and the sequence of idiosyncratic shocks to farm incomes. But if $m\geq 2$ there are two additional terms that also have to be taken into consideration, which

[11] It can be shown that $P_{EqE}q_{EqE} = (CI(\hat{y}^*_s, R_s) - G(y_s))(\hat{\lambda}\bar{\hat{y}}^*_s / \bar{y}_f)$ where $CI(\hat{y}^*_s, R_s)$ is the CI of predicted equilibrium income ranked by actual incomes in the base year. Typically we would expect $CI(\hat{y}^*_s, R_s) - G(y_s) < 0$ since $CI(\hat{y}^*_s, R_s) < G(\hat{y}^*_s)$ by definition (see Lambert, 2001, p.29), where $G(\hat{y}^*_s)$ is the Gini of predicted equilibrium incomes, and $G(\hat{y}^*_s) < G(y_s)$ due to the disequalising effect of transitory shocks to annual incomes.

capture the effect of changes in the target income function (3) due to changing prices and technology (as reflected in the time-varying parameters in (4)) and the interaction between farm business size and price/technology changes.

Table 4 reports the results from the estimation of the ECM, with the dependent variable being the annual change in cash income and farm business size measured in terms of total standard gross margin (SGM), based on Scottish average gross margins for the years 1998 to 2002. OLS estimates of (4) will be biased due to the correlation between lagged income and the fixed effects in the error term (see Bond, 2002, for a discussion). To overcome this problem we follow Mundlak (1978) by explicitly modelling the fixed effects as a function of farm-specific SGM averages, and further control for initial conditions in the manner of Wooldridge (2005) by including as a separate explanatory variable the level of income in the year in which a farm first entered the sample. This estimation strategy has the appeal that it provides explicit estimates of the farm-specific fixed effects, which will prove informative in the decomposition of equilibrium inequality, and avoids the further restriction of the sample that would result from the use of Generalised Method of Moments (GMM) estimators as these require higher-order lags of income to serve as instruments.[12] The preferred specification imposes the restriction that $\alpha_{1t} = \alpha_1$ for all *t*, i.e. that the immediate impact of changes in farm business size is constant over the study period.[13]

The first set of columns report the estimates of the dynamic income model (4). Thus the short-run or impact effect of a £1 increase in the SGM of a farm business was to increase cash incomes by 19.5 pence. The remainder of the dynamic income function relates to the equilibrium error, where the coefficient on lagged income provides an estimate of the adjustment parameter $\hat{\lambda}$ equal to 0.51, implying that just over half of the gap between any farm's actual and target income in one year was closed by the next year. Dividing the coefficients on the lagged determinants of income by $\hat{\lambda}$ yields the parameters of the implied

[12] The preferred estimator yields an estimate of the adjustment parameter λ between those of the downwardly biased OLS estimator and the upwardly biased within-groups estimator. In contrast, Arellano and Bond (1991) and Blundell and Bond (1998) GMM estimates of λ were both less than the OLS estimate and close to zero.

[13] The set of annual slope dummies on the farm business size change variable is only just significant at the 5% level (F=1.79; F(13,6365,5%)=1.72), unlike both the intercept dummies (F(13,6365)=3.01) and farm size slope dummies (F(13,6365)=6.33).

equilibrium or target income function (3), which are reported in the second set of columns. Taking the reference year of 1995 as an example, the implied long-run effect of a £1 increase in SGM was 45.4 pence given prevailing agroeconomic conditions, or more than twice as large as the impact effect. Long-run income effects of changes in farm business size are predicted to have been positive in all years, being on average 1.8 times the impact effect. Finally, farm-level fixed effects are not significantly affected by the average business size of the farm but there is a significant positive relationship with the level of income in the year in which the farm first entered the sample.

Table 5 expands upon the results of the decomposition analysis in Table 2, identifying the separate contributions of farm business size changes, the initial equilibrium error, the residual, price changes, and the interaction between farm size and price changes to the vertical mobility index M_V. First, changes in farm business size made contributions to annual vertical mobility that were both negligible and statistically insignificant in all years. Moreover this continues to be the case even when considering vertical mobility over multiyear periods, where these multiyear estimates take into account not only the contemporaneous effects of farm business size changes but also the lagged effects operating through the error correction mechanism. Accordingly, the results provide no evidence that the pattern of income changes due to farm business growth over the study period was biased in favour of either low or high income farms.

Second the contribution of the equilibrium error in the base year to vertical mobility is significantly negative in every year, which is consistent with our expectation based on the discussion of (5). More intuitively, the process of adjustment towards equilibrium or target incomes is equalizing with large positive (negative) shocks to income in a particular year offset by slow (fast) income growth in subsequent years. The correction of the initial equilibrium error accounts, on average, for all of the vertical mobility associated with annual income changes over the period, providing an explanation of how income growth could appear to be biased in favour of lower income farms in spite of the finding that the income effect of farm business growth was not.

Third, the contribution of the residual offsets the equalising effect of the equilibrium error in some years and reinforces it in others, though the effect is not significant in any single year and is roughly equal to zero on average over the full set of annual changes. This lack of systematic contribution to vertical mobility is to be expected given that the residual allows for the impact of idiosyncratic shocks to farm incomes after controlling for both farm-specific fixed effects and year-specific price effects. By construction the residual is uncorrelated with lagged income over the full panel.

Finally, the effect of the common time-varying factors in the multiyear decompositions was diverging in the first sub-period over which average incomes fell and significantly equalising in the second when they rose sharply (see Table 1), consistent with the earlier finding that relative inequality was generally higher in years of lower average incomes. By implication, the cash incomes of the higher income farms were disproportionately affected by changes in the economic fortunes of the agricultural sector, benefiting more in relative terms from upturns but losing more in downturns. The effect of the interaction term was trivial and insignificant throughout.

4.2 Equilibrium or structural inequality

Finally we note that the ECM implies a measure of equilibrium or target income y_{it}^* in year t (t=1,…T−1) conditional upon prices, technology and farm business size. Hence inequality in y_{it}^* may be interpreted as a measure of equilibrium or structural inequality, which may be compared with the observed inequality of annual farm incomes and used to identify which factors contribute to structural inequality.

By analogy with the Shorrocks index M_T, we propose a measure that reflects the degree to which observed annual incomes are equalized in the implied equilibrium. Specifically, if $\hat{y}_{it}^* = \hat{\beta}_0 + \hat{\beta}_{1t} d_t + \hat{\beta}_2 z_{it} + \hat{\beta}_{3t} d_t z_{it} + \hat{\mu}_i$ is the predicted level of equilibrium income in year t given the derived set of parameter estimates of (3), then we define the following set of disequilibrium adjustment mobility indices:

$$M_{DA}^t = 1 - \frac{G(\hat{y}_t^*)}{G(y_t)}; \qquad t=1,\ldots T-1 \tag{8}$$

where the Gini coefficient of predicted equilibrium incomes $G(\hat{y}_t^*) = 2\,\mathrm{cov}(\hat{y}_{it}^*, \hat{R}_{it}^*)/\overline{\hat{y}_t^*}$ is interpreted as a measure of structural inequality; $\overline{\hat{y}_t^*}$ is average predicted equilibrium income; and $\hat{R}_{it}^*$ is the individual's relative rank in the predicted equilibrium income distribution. M_{DA}^t will equal one when there is no structural inequality, in which case $G(\hat{y}_t^*) = 0$, and will equal zero if actual and equilibrium incomes are identical, in which case $G(\hat{y}_t^*) = G(y_t)$. Hence, as with the Shorrocks index M_T, the disequilibrium adjustment mobility index will take values close to one if inequality is largely a short-run phenomenon due to transitory peturbations from equilibrium, whereas if inequality largely arises from structural differences between farms then the index will take values close to zero.

Table 6 reports the equilibrium Gini coefficient estimates for each year where these are smaller than the corresponding annual Gini coefficients (repeated from Table 1) in all years except 2005, which is consistent with the finding that averaging income over a number of years typically reduces inequality. Accordingly, the disequilibrium adjustment indices M^{t}_{DA} take values in the unit interval in all years but 2005, with an average value of 29.1% over the entire period indicating that a significant proportion of the inequality in annual incomes in any year was due to the incidence and persistence of idiosyncratic shocks. This estimate of the potential for the equalisation of incomes in the longer term is more than twice as high as the limiting Shorrocks Index M_T value of 12% for the entire period, where the difference may be ascribed to the alternative definitions of "longer term" incomes employed in the construction of the two indices. Nevertheless both measures do imply that the overwhelming fraction of farm income inequality is permanent or structural in nature.

Finally, the determinants of structural inequality may be obtained using regression-based procedures (see, e.g., Morduch and Sicular, 2002) to decompose the Gini coefficient of predicted equilibrium incomes:

$$\begin{aligned} G(\hat{y}^*_t) &= 2\operatorname{cov}(\hat{y}^*_{it}, \hat{R}^*_{it}) / \bar{\hat{y}}^*_t = 2\operatorname{cov}(\hat{\beta}_0 + \hat{\beta}_{1t} d_t + \hat{\beta}_2 z_{it} + \hat{\beta}_{3t} d_t z_{it} + \hat{\mu}_i, \hat{R}^*_{it}) / \bar{\hat{y}}^*_t \\ &= (\hat{\beta}_2 + \hat{\beta}_{3t}) \frac{2\operatorname{cov}(z_{it}, \hat{R}^*_{it})}{\bar{\hat{y}}^*_t} + \frac{2\operatorname{cov}(\hat{\mu}_i, \hat{R}^*_{it})}{\bar{\hat{y}}^*_t} = \frac{(\hat{\beta}_2 + \hat{\beta}_{3t})\bar{z}_t}{\bar{\hat{y}}^*_t} CI(z_t, \hat{R}^*_t) + \frac{\bar{\hat{\mu}}}{\bar{\hat{y}}^*_t} CI(\hat{\mu}, \hat{R}^*_t) \end{aligned} \quad ; t=1,..T-1 \quad (9)$$

where $CI(z_t, \hat{R}^*_t)$ is the CI of farm business size ranked by predicted equilibrium income in year t, with corresponding average value $\bar{z}_t$; and $CI(\hat{\mu}, \hat{R}^*_t)$ and $\bar{\hat{\mu}}$ are the corresponding statistics for the farm-specific fixed-effect term. Hence the Gini coefficient is given as a weighted sum of CIs, with the weight on each CI equal to the share of predicted equilibrium income attributable to that factor where this is given by the elasticity of equilibrium income with respect to that factor evaluated at the means. The intuitive interpretation is that a factor can only contribute to structural inequality if the factor is statistically associated with equilibrium income and concentrated among either high or low income farms.

The remainder of Table 6 provides results from the analysis of the determinants of structural inequality. On average, just under two thirds (64.6%) of structural inequality in farm incomes was due to observable differences in the size of farm businesses, as measured in terms of standard gross margin: larger farm businesses tended to generate higher cash incomes so farm business size is a source of income inequality. This leaves the remaining third (35.4%) of structural inequality attributable to farm-level fixed effects, where these effects make a statistically significant contribution in some years. This may seem a surprisingly high proportion until it is remembered what the fixed effects represent. Firstly they allow for a multitude of factors - most notably land quality and managerial ability - that affect farms' financial performance but are hard to measure and therefore not explicitly controlled for in the model: empirical analyses of farm enterprise performance (e.g. Scottish Government, 2012b) provide ample evidence of the considerable variation in returns achieved by Scottish farmers. Secondly, they also allow for differences in workforce composition and land ownership between farms, which will affect the cash incomes of farms but are not taken into account in the calculation of SGMs. Thus equilibrium or structural inequality is not only due to differences in the economic size of farms as conventionally measured but also in their cash income generating performance.

5. Conclusions

This paper provides a thorough evaluation of the distributional consequences of farm income mobility in Scotland between 1995 and 2009 using a range of mobility indices to explore two distinct but interrelated issues: the extent to which farm income inequality is a chronic as opposed to a temporary phenomenon; the nature of the dynamic processes driving changes in farm income inequality over time. The empirical study is based on an unbalanced panel of FAS farm records in which each farm was assigned a unique identifier for the whole of the time it remained in the survey in order to fully capture income mobility within the sample. Cash income was chosen as the FAS farm income indicator that corresponds most closely to the farm income position as perceived by the farmer.

The empirical results reveal that farm income inequality was partly a temporary or short-run phenomenon, with the estimates of the Shorrocks and disequilibrium adjustment mobility indices implying that somewhere between 12% and 30% of inequality in annual incomes may have been due to the incidence and persistence of idiosyncratic shocks. Farm income instability would likely have been higher but for the substantial role played by Pillar 1 direct payments in reducing the exposure of farms to market and production risk (Tangermann, 2011; Hennessy, 2014). However, it remains to be shown formally that such payments would also have had the effect of reducing the degree to which farm incomes are equalised in the longer term, which remains a topic for further research. The most recent CAP reform included a new income stabilisation tool as part of a 'risk management toolkit' under Pillar 2, which would allow for the compensation of farmers who experience a severe drop in their incomes (European Commission, 2013). However the Scottish Government

(2015, p.744) has chosen not to implement this provision on the grounds that it is more appropriate for basic levels of income protection to be provided through Pillar 1 measures.

The overwhelming proportion of farm income inequality was, however, permanent or structural in nature. Roughly two thirds of the structural inequality is further shown, on average, to have been due to differences in farm business size, with the remainder due to farm-level fixed effects that represent differences in both financial performance and business structure. Within Scotland, the move from historic to area-based direct payments in the new CAP will inevitably redistribute support in future from farms with more intensive enterprises towards those with more extensive systems (see, e.g., Vosough Ahmadi et al., 2014). However the Scottish Government has sought to limit the resultant scale of farm income redistribution by adopting a regionalised model in which regional payment rates reflect the productive capacity of land (Scottish Government, 2014). The Scottish Government also chose not to adopt the redistributive payment scheme, which could potentially have done more to tackle the unequal distribution of farm income than the modulation and capping of direct payments to larger farms (cf. Matthews, 2013).

Finally, the empirical findings provide mixed evidence on whether or not the proportional rate of farm income growth was independent of farm income. The raw estimates of vertical mobility indicate that farm income growth conditional upon farm survival was higher on average on lower income farms, with a battery of robustness tests largely serving to validate this result. However, the further analysis of the determinants of vertical mobility reveal that the equalizing effect of expected income changes was almost entirely due to the process of adjustment towards equilibrium or target incomes conditional upon prices, technology and farm business size. In contrast, the decomposition results provide no evidence that relative income growth due to farm business size changes was associated with initial incomes. Further work is required on the impact of farm entry and exit processes to evaluate fully the effects of changes in the farm business size structure on the evolution of the farm income distribution.

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Table 1. Basic summary statistics and Shorrocks mobility index M_T

	Annual summary statistics			Multiyear analysis Base year =1995		
Year	$\bar{y}_t$ (£)	$G(y_t)$	*Mean total farm SGM* (£)	T	$G(y_A)$	M_T
1995	40489***	0.505 ***	55154 ***	1	0.505***	0
	1724	*0.017*	*2248*		*0.017*	-
1996	43707***	0.447 ***	56746 ***	2	0.442***	0.057***
	1372	*0.015*	*2109*		*0.014*	*0.017*
1997	27644***	0.512 ***	57942 ***	3	0.442***	0.067***
	995	*0.019*	*2250*		*0.013*	*0.010*
1998	30008***	0.495 ***	59471 ***	4	0.446***	0.065***
	1208	*0.018*	*2238*		*0.014*	*0.009*
1999	27161***	0.594 ***	61950 ***	5	0.452***	0.069***
	1507	*0.034*	*2482*		*0.015*	*0.009*
2000	28934***	0.545 ***	61666 ***	6	0.456***	0.086***
	1536	*0.023*	*3092*		*0.015*	*0.012*
2001	28874***	0.546 ***	56897 ***	7	0.457***	0.097***
	1187	*0.022*	*2185*		*0.017*	*0.012*
2002	31196***	0.502 ***	67294 ***	8	0.470***	0.089***
	1445	*0.022*	*3207*		*0.019*	*0.015*
2003	36414***	0.460 ***	63544 ***	9	0.450***	0.099***
	1343	*0.018*	*2555*		*0.029*	*0.021*
2004	36576***	0.478 ***	65239 ***	10	0.444***	0.103***
	1227	*0.018*	*2280*		*0.031*	*0.021*
2005	31654***	0.486 ***	64584 ***	11	0.411***	0.138***
	1146	*0.019*	*2495*		*0.022*	*0.022*
2006	35168***	0.539 ***	66498 ***	12	0.422***	0.105***
	1596	*0.022*	*2839*		*0.027*	*0.022*
2007	46891***	0.537 ***	66772 ***	13	0.418***	0.116***
	2040	*0.020*	*2952*		*0.027*	*0.023*
2008	47087***	0.512 ***	69687 ***	14	0.421***	0.119***
	1777	*0.016*	*2929*		*0.026*	*0.019*
2009	48935 ***	0.480***	70570 ***	15	0.417***	0.121***
	1731	*0.017*	*3689*		*0.029*	*0.021*

*Source: Authors' calculations based on (1). Annual summary statistics based on the full sample available in the relevant year. Multiyear analysis statistics are based on the sample of farms present in all T years of the relevant measurement period (e.g. 1995-2009 for T=15). Bootstrapped standard errors in italics based on 1000 replications. Statistical significance at 1%, 5% and 10% levels are denoted by ***, ** and * respectively.*

Table 2. Decomposition of changes in the annual Gini coefficient over selected periods

Period	$G(y_s)$	$G(y_f)$	Change	M_V	P	q	M_R
1995-1996	0.485***	0.452***	-0.033**	-0.123***	-1.580	0.078***	0.090***
	0.015	*0.016*	*0.015*	*0.017*	*8.696*	*0.028*	*0.021*
1996-1997	0.431***	0.492***	0.061***	-0.031**	0.054**	-0.571***	0.092***
	0.012	*0.015*	*0.016*	*0.014*	*0.025*	*0.042*	*0.009*
1997-1998	0.515***	0.511***	-0.004	-0.115***	-1.890	0.061	0.111***
	0.020	*0.020*	*0.023*	*0.028*	*164.962*	*0.038*	*0.017*
1998-1999	0.493***	0.542***	0.049***	-0.056***	0.657	-0.086**	0.105***
	0.019	*0.020*	*0.017*	*0.020*	*3.428*	*0.037*	*0.014*
1999-2000	0.616***	0.548***	-0.068***	-0.222***	-3.328	0.067	0.155***
	0.023	*0.019*	*0.025*	*0.030*	*59.292*	*0.048*	*0.026*
2000-2001	0.536***	0.554***	0.018	-0.172***	-2.553	0.067	0.190***
	0.024	*0.022*	*0.031*	*0.037*	*35.223*	*0.051*	*0.026*
2001-2002	0.556***	0.525***	-0.031***	-0.187***	4.083	-0.046	0.155***
	0.023	*0.021*	*0.024*	*0.027*	*64.129*	*0.038*	*0.019*
2002-2003	0.477***	0.465***	-0.012	-0.135**	-0.690***	0.195***	0.122***
	0.024	*0.018*	*0.020*	*0.038*	*0.191*	*0.032*	*0.027*
2003-2004	0.458***	0.465***	0.007	-0.115***	1.557	-0.074**	0.122***
	0.018	*0.020*	*0.016*	*0.021*	*421.861*	*0.035*	*0.017*
2004-2005	0.473***	0.490***	0.017	-0.106***	0.647**	-0.164***	0.123***
	0.017	*0.019*	*0.016*	*0.021*	*0.271*	*0.036*	*0.014*
2005-2006	0.487***	0.532***	0.045**	-0.093***	-0.978	0.096***	0.138***
	0.019	*0.021*	*0.020*	*0.024*	*2.089*	*0.035*	*0.016*
2006-2007	0.538***	0.528***	-0.010	-0.141***	-0.566***	0.250***	0.131***
	0.023	*0.020*	*0.024*	*0.027*	*0.147*	*0.034*	*0.016*
2007-2008	0.527***	0.512***	-0.014	-0.152***	5.250	-0.029	0.138***
	0.021	*0.017*	*0.017*	*0.021*	*95.404*	*0.038*	*0.016*
2008-2009	0.515***	0.482***	-0.034**	-0.162***	-13.515	0.012	0.129***
	0.016	*0.016*	*0.016*	*0.018*	*297.357*	*0.029*	*0.012*
1995-2009	0.472***	0.532***	0.060**	-0.112***	-1.784	0.063	0.172***
	0.022	*0.027*	*0.027*	*0.030*	*39.819*	*0.046*	*0.022*
1995-2004	0.481***	0.545***	0.064**	-0.107***	0.373***	-0.288***	0.172***
	0.017	*0.027*	*0.028*	*0.028*	*0.123*	*0.056*	*0.019*
2005-2009	0.486***	0.483***	-0.003	-0.145***	-0.471***	0.309***	0.142***
	0.022	*0.020*	*0.021*	*0.022*	*0.074*	*0.025*	*0.017*

*Source: Authors' calculations based on (2). Each statistic is based the sample of farms that are present in all years of the relevant period. Bootstrapped standard errors in italics based on 1000 replications. Statistical significance at 1%, 5% and 10% levels are denoted by ***, ** and * respectively.*

Table 3. Alternative estimates of M_V over selected periods

	Estimation technique		
Period	*Standard*	*Smoothed incomes*	*'Instrumented' ranks*
1996-1999	0.030*	-0.014	0.069***
	0.018	*0.014*	*0.021*
1997-2000	-0.103***	0.009	0.018
	0.035	*0.014*	*0.028*
1998-2001	-0.104***	-0.078***	-0.051
	0.036	*0.024*	*0.041*
1999-2002	-0.171***	-0.065***	-0.061**
	0.033	*0.020*	*0.030*
2000-2003	-0.193***	-0.080***	-0.003
	0.038	*0.021*	*0.031*
2001-2004	-0.169***	-0.072***	-0.015
	0.021	*0.012*	*0.026*
2002-2005	-0.078*	-0.022	-0.029
	0.046	*0.025*	*0.030*
2003-2006	-0.013	-0.015	0.065**
	0.030	*0.017*	*0.027*
2004-2007	-0.125***	-0.015	0.012
	0.030	*0.014*	*0.022*
2005-2008	-0.104***	-0.055***	0.000
	0.025	*0.017*	*0.020*
1996-2008	-0.005	-0.051**	0.028
	0.028	*0.026*	*0.038*

*Source: Authors' calculations based on (2). Each statistic is based on the sample of farms that are present in all years from the year before the base year to the year after the final year of each period (e.g. 1995 to 2000 for the first period 1996-1999) to allow construction of the smoothed income and 'instrumented' rank variables. The need to generate lags and leads limits the analysis to the period 1996-2008, with the three year intervals chosen to avoid overlap in the construction of base and final year measures. Bootstrapped standard errors in italics based on 1000 replications. Statistical significance at 1%, 5% and 10% levels are denoted by ***, ** and * respectively.*

Table 4. Dynamic income model results and the implied equilibrium income function

Explanatory variables		Error Correction Model (4) Coeff.		Robust Std error	Target income function (3) Coeff.		Bootstrapped Std error
Change in SGM	$\Delta z_{i,t+1}$	0.195	***	*0.074*	-		
Lagged income	y_{it}	-0.512	***	*0.028*	-		
Lagged SGM	z_{it}	0.233	***	*0.079*	0.454	***	*0.143*
Dummy1996* z_{it}		-0.184	***	*0.062*	-0.360	***	*0.111*
Dummy1997* z_{it}		-0.076		*0.063*	-0.148		*0.115*
Dummy1998* z_{it}		-0.092		*0.063*	-0.179		*0.116*
Dummy1999* z_{it}		-0.081		*0.067*	-0.158		*0.123*
Dummy2000* z_{it}		-0.024		*0.075*	-0.046		*0.130*
Dummy2001* z_{it}		-0.155	**	*0.064*	-0.303	***	*0.114*
Dummy2002* z_{it}		0.027		*0.064*	0.054		*0.114*
Dummy2003* z_{it}		-0.115		*0.074*	-0.224	*	*0.133*
Dummy2004* z_{it}		-0.150	**	*0.066*	-0.293	**	*0.118*
Dummy2005* z_{it}		0.075		*0.087*	0.147		*0.156*
Dummy2006* z_{it}		0.157		*0.113*	0.308		*0.205*
Dummy2007* z_{it}		-0.084		*0.081*	-0.163		*0.146*
Dummy2008* z_{it}		-0.041		*0.071*	-0.081		*0.129*
Intercept		7047	**	*2878*	13770	***	*5209*
Dummy1996		-6246	**	*3126*	-12205	**	*5701*
Dummy1997		-5082		*3114*	-9930	*	*5737*
Dummy1998		-7750	**	*3116*	-15144	**	*5916*
Dummy1999		-4117		*3294*	-8044		*6124*
Dummy2000		-5296		*3827*	-10349		*6692*
Dummy2001		785		*3236*	1534		*5892*
Dummy2002		-5286		*3246*	-10330	*	*5863*
Dummy2003		757		*3942*	1478		*7169*
Dummy2004		-2081		*3379*	-4066		*6084*
Dummy2005		-12068	**	*5049*	-23583	***	*9128*
Dummy2006		-8005		*7016*	-15642		*12598*
Dummy2007		2053		*4606*	4011		*8320*
Dummy2008		960		*3886*	1875		*7071*
Farm-specific average SGM	$\bar{z}_{i.}$	-0.035		*0.052*	-0.069		*0.098*
Income in sample entry year	y_i^{entry}	0.160	***	*0.021*	0.312	***	*0.035*
Sample size		6412					
R^2		0.288					
F(31,6380)		34.42	***				

*Source: Authors' estimates based on full unbalanced panel. Robust standard errors allow for heteroscedasticity and autocorrelation. Bootstrapped standard errors based on 1000 replications. Statistical significance at 1%, 5% and 10% levels are denoted by ***, ** and * respectively.*

Table 5. Decomposition of vertical mobility M_V

Period	M_V	*Change in farm business size*	*Base year equilibrium error*	*Residual*	*Change in common time-varying factors*	*Interaction term*
		Contribution to vertical mobility M_V				
1995-1996	−0.1229***	−0.0014	−0.1174***	−0.0041	-	-
	0.0169	*0.0022*	*0.0254*	*0.0183*	-	-
1996-1997	−0.0306**	−0.0003	−0.0164	−0.0139	-	-
	0.0141	*0.0024*	*0.0167*	*0.0165*	-	-
1997-1998	−0.1151***	0.0018	−0.1135***	−0.0034	-	-
	0.0283	*0.0020*	*0.0225*	*0.0250*	-	-
1998-1999	−0.0564***	−0.0027	−0.0564***	0.0026	-	-
	0.0205	*0.0054*	*0.0184*	*0.0185*	-	-
1999-2000	−0.2225***	−0.0024	−0.1907***	−0.0294	-	-
	0.0295	*0.0032*	*0.0305*	*0.0209*	-	-
2000-2001	−0.1723***	0.0095	−0.1693***	−0.0125	-	-
	0.0372	*0.0076*	*0.0374*	*0.0310*	-	-
2001-2002	−0.1866***	0.0060	−0.2216***	0.0290	-	-
	0.0273	*0.0082*	*0.0292*	*0.0287*	-	-
2002-2003	−0.1349***	−0.0009	−0.1377***	0.0037	-	-
	0.0383	*0.0021*	*0.0261*	*0.0287*	-	-
2003-2004	−0.1148***	−0.0018	−0.1484***	0.0354	-	-
	0.0211	*0.0018*	*0.0282*	*0.0255*	-	-
2004-2005	−0.1061***	0.0004	−0.1273***	0.0208	-	-
	0.0210	*0.0034*	*0.0237*	*0.0261*	-	-
2005-2006	−0.0934***	0.0009	−0.0953***	0.0009	-	-
	0.0240	*0.0027*	*0.0363*	*0.0290*	-	-
2006-2007	−0.1412***	−0.0027	−0.1431***	0.0045	-	-
	0.0273	*0.0026*	*0.0463*	*0.0384*	-	-
2007-2008	−0.1518***	0.0007	−0.1705***	0.0179	-	-
	0.0207	*0.0018*	*0.0320*	*0.0264*	-	-
2008-2009	−0.1620***	0.0026	−0.1702***	0.0055	-	-
	0.0181	*0.0023*	*0.0228*	*0.0192*	-	-
1995-2009	−0.1124***	−0.0190	−0.2175***	0.0937 *	0.0256	0.0048
	0.0303	*0.0199*	*0.0543*	*0.0537*	*0.0516*	*0.0066*
1995-2004	−0.1074***	−0.0110	−0.3134***	0.1318 **	0.0817	0.0034
	0.0280	*0.0153*	*0.0704*	*0.0594*	*0.0673*	*0.0054*
2005-2009	−0.1454***	0.0071	−0.1329**	0.1485***	−0.1665***	−0.0016
	0.0220	*0.0111*	*0.0555*	*0.0340*	*0.0512*	*0.0043*

*Source: Authors' calculations based on Eqs. (5) and (7). Sample definitions as given in Table 2. Bootstrapped standard errors in italics based on 1000 replications. Statistical significance at 1%, 5% and 10% levels are denoted by ***, ** and * respectively.*

Table 6. Disequilibrium adjustment mobility and the determinants of structural inequality

Year	Equilibrium inequality $G(y_t^*)$	Observed inequality $G(y_t)$	Disequilibrium adjustment mobility M_{DA}	Contribution to structural inequality $G(y_t^*)$	
				Farm business size	Fixed Effects
1995	0.312 ***	0.505***	0.382 ***	0.222***	0.089 *
	0.046	*0.018*	*0.090*	*0.065*	*0.054*
1996	0.417 ***	0.447***	0.065	0.126	0.291 *
	0.060	*0.015*	*0.137*	*0.163*	*0.152*
1997	0.376 ***	0.512***	0.265 ***	0.241**	0.135
	0.035	*0.019*	*0.075*	*0.094*	*0.084*
1998	0.456 ***	0.495***	0.077	0.283**	0.173
	0.044	*0.018*	*0.089*	*0.127*	*0.112*
1999	0.367 ***	0.594***	0.382 ***	0.234**	0.133
	0.044	*0.063*	*0.087*	*0.107*	*0.084*
2000	0.394 ***	0.545***	0.277 ***	0.288***	0.106
	0.048	*0.023*	*0.096*	*0.077*	*0.073*
2001	0.246 ***	0.546***	0.550 ***	0.100	0.146 **
	0.037	*0.023*	*0.070*	*0.075*	*0.067*
2002	0.408 ***	0.502***	0.186 ***	0.323***	0.085
	0.030	*0.023*	*0.061*	*0.079*	*0.067*
2003	0.262 ***	0.460***	0.430 ***	0.143*	0.119 *
	0.049	*0.019*	*0.108*	*0.081*	*0.064*
2004	0.282 ***	0.478***	0.411 ***	0.134	0.148 *
	0.046	*0.018*	*0.096*	*0.099*	*0.087*
2005	0.525 ***	0.486***	-0.079	0.447***	0.078
	0.094	*0.019*	*0.194*	*0.148*	*0.080*
2006	0.438 ***	0.539***	0.188	0.393***	0.045
	0.090	*0.022*	*0.179*	*0.092*	*0.051*
2007	0.257 ***	0.537***	0.521 ***	0.177**	0.081
	0.056	*0.020*	*0.105*	*0.085*	*0.061*
2008	0.296 ***	0.512***	0.423 ***	0.221***	0.075
	0.047	*0.016*	*0.092*	*0.081*	*0.059*
Average contribution to structural inequality				64.6%	35.4%

*Source: Authors' calculations based on Eqs. (8) and (9). All summary statistics based on the full sample available in the relevant year. Note that equilibrium income is not defined in 2009. Bootstrapped standard errors in italics based on 1000 replications. Statistical significance at 1%, 5% and 10% levels are denoted by ***, ** and * respectively.*