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Income mobility and income inequality in Scottish agriculture

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Income mobility and income inequality in Scottish agriculture

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Abstract: The paper proposes the use of a range of alternative measures to provide a rounded evaluation of the distributional consequences of farm income mobility, where this multifaceted approach is designed to shed light both on the extent to which farm income inequality is a short-run phenomenon due to transitory shocks rather than a chronic or persistent problem due to structural factors, and on the nature of the dynamic processes driving changes in farm income inequality over time. An illustrative empirical study of Scottish agriculture using Farm Accounts Survey data reveals that the majority of farm income inequality was structural in nature despite a substantial degree of income risk due to the volatility of agricultural incomes. Results on the micro-dynamics of inequality change have to be interpreted with caution due to the particular rules governing the assignment of farm identifiers in the survey.

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

JEL classifications: D31, D63, Q18

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1.   Introduction

One of the enduring goals of the Common Agricultural Policy (CAP) has been "*to ensure a fair standard of living for the agricultural community, in particular by increasing the individual earnings of persons engaged in agriculture*" (Treaty of Rome, 1957), where not only the average level but also the distribution and stability of farm incomes have generally been viewed as being of policy concern (Hill, 2012). Thus successive reforms since 1992 have continued to provide income support and safety net mechanisms for producers while increasing the market orientation of the agricultural sector. The most recent reform package seeks to ensure a more equitable distribution of direct payments among countries and farmers, while also including a new reserve to secure the financial resources needed in case of crisis and a new income stabilisation tool as part of a package of risk management measures (see European Commission, 2013).

Analyses of individual agricultural incomes typically focus either on the static distributional consequences of support (e.g. Keeney, 2000; Allanson, 2008: Morreddu. 2011) or on income instability issues (e.g. Hegrenes *et al*., 2001; Mishra and Sandretto, 2002; Finger and El Benni, 2014). The main aim of this paper is to explore the distributional implications of farm income mobility in Scottish agriculture and thereby address three distinct but interrelated issues. Firstly we consider the extent to which inequality is a short-run phenomenon due to transitory shocks as opposed to a chronic or persistent problem due to structural factors. Income inequality may be less of a policy concern if it is largely a transitory phenomenon such that farm incomes are equalised over the longer term. Second we seek to investigate the impact of systematic changes in the size structure of the farm sector on the observed level of inequality over time. In particular we consider whether structural changes have been distributionally neutral in relative terms, which will be the case if expected income growth rates are independent of size and therefore consistent with Gibrat's Law (Gibrat, 1931). Finally we provide estimates of a measure of income risk based on a dynamic

model of farm incomes that explicitly takes into account the impact of both transitory shocks and structural change.

The paper adds 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 extend this work by providing complementary measures of income risk that measure transitory shocks about expected equilibrium incomes rather than about multiyear average incomes. Phimister et al. (2004) further explore 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 mobility within agriculture using transition matrices (see e.g. Meuwissen et. al (2008) for a recent example). We more broadly consider the impact of structural mobility across the whole of the income distribution using a range of methods that have been developed in the broader economics literature but not previously applied to agriculture.

The paper is structured as follows. The next section provides an overview of the methods that are employed to explore the distributional implications of farm income mobility. Section 3 presents the empirical study of income mobility in Scottish agriculture, which is based on an unbalanced panel of farms drawn from the Scottish Farm Accounts Survey over the production years 1995-2009. The final section concludes with a discussion of the empirical findings in the light of the most recent round of CAP reform.

2. Methodology

The multifaceted nature of income mobility has resulted in the emergence of several distinct strands in the literature characterising the distributional implications of the phenomenon. Jantti and Jenkins (2015) in their recent handbook article identify four main aspects of income mobility that may be of

normative significance: reduction of longer-term inequality, individual income growth, positional change and income risk. In this section, we outline a range of methods that allow us to capture all four of these distinct dimensions and thereby provide a rounded evaluation of the distributional consequences of farm income mobility.

2.1. Choice of inequality measure

The inequality measure used throughout the subsequent analysis is the Gini coefficient. Let $G(y_t)$ be the Gini coefficient of incomes in year t, which can be written as:

$$G(y_t) = \frac{2}{\bar{y}_t} \text{cov}\left(y_{it}, R_{it}\right) \tag{1}$$

where y_{it} is the income of individual i ($i=1,….N$) in year t, $\bar{y}_t$ is average income, and R_{it} is the individual'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 incomes are identical and of one when all income accrues to one individual and everybody else receives nothing. The choice of a relative inequality measure provides a natural benchmark to evaluate farm income mobility: if the expected income growth process satisfies Gibrat's Law of proportionate effect then it will be distributionally neutral in relative terms.[1]

2.2 Reduction of longer-term inequality

Income mobility may be seen to be socially desirable to the extent that it reduces inequality over the longer term. This aspect of mobility is captured by the mobility index due to Shorrocks (1978a), which measures the extent to which equalization occurs when the measurement period is extended from one to T years:

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[1] In this context, farm income is treated as a size measure. As Kostov *et al*. (2005) point out, a wide range of different variables has been used as size measure in papers studying Gibrat's law, including farmed acreage, livestock numbers, net worth, gross sales, total gross margins and net income (Allanson, 1992; Clarke *et al*. 1992; and, Shapiro *et al*. 1987).

$$M_T = 1 - \frac{G(y_A)}{\sum_{t=1}^{T} w_t G(y_t)} \qquad (2)$$

where $G(y_A) = 2\operatorname{cov}(y_{iA}, R_{iA})/\bar{y}_A$ is the Gini coefficient of individual average incomes $y_{iA} = \sum_{t=1}^{T} y_{it}/T$ calculated over the T-year period $t=1,\ldots,T$; R_{iA} are the corresponding relative ranks; $\bar{y}_A = \sum_{i=1}^{N} \bar{y}_{iA}/N$ is overall average income over the entire period; and $w_t = \bar{y}_t/\bar{y}_A$ 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 when the relative incomes of all individual farms remains constant through time in which case the Gini coefficient for each year and for the measurement period as a whole will be the same. Hence if inequality is largely a short-run phenomenon due to transitory income shocks then the index will take a value close to one whereas if inequality largely arises from long-term differences between farms then the index will take a value close to zero.

2.3 Individual income growth and positional change

The change in inequality between any two years may be decomposed into elements due to individual income growth and positional change,[2] thereby serving to characterise the nature of the transition process from the initial to the final distribution of incomes. 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:

$$G(y_f) - G(y_s) = \big(G(y_f) - CI(y_f, R_s)\big) + \big(CI(y_f, R_s) - G(y_s)\big) = M_H + M_V \qquad (3)$$

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[2] These two elements are called 'structural' and 'exchange' mobility respectively in the sociology literature: the former may be identified with changes in the location and shape of the marginal income distribution and the latter with the permutation of a fixed set of income opportunities among individuals.

where $CI(y_f, R_s) = 2\operatorname{cov}(y_{if}, R_{is})/\bar{y}_f$ 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.

$M_V = 2\left\{\operatorname{cov}(y_{if}, R_{is})/\bar{y}_f - \operatorname{cov}(y_{is}, R_{is})/\bar{y}_s\right\}$ provides a measure of vertical mobility that addresses the question of whether the distribution of income changes favour those with initially low or high incomes and thereby provides a natural counterpart to $G(y_t)$ which addresses the distribution of income between the poor or rich. M_V will be zero if income changes are unrelated to base year income rank and will be negative if income changes are equalising in relative terms, which will be the case if the initially poor experience either larger relative gains on average than the rich or smaller relative losses. M_V in turn depends on the progressivity and scale of income changes, such that $M_V = Pq$. Progressivity is captured by the disproportionality index

$P = CI(y_f - y_s, R_s) - G(y_s)$ where $CI(y_f - y_s, R_s) = 2\operatorname{cov}(y_{if} - y_{is}, R_{is})/(\bar{y}_f - \bar{y}_s)$ 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 $q = (\bar{y}_f - \bar{y}_s)/\bar{y}_f$. Note that negative values of P imply that income changes will be equalising if incomes are growing on average but disequalising if incomes are falling.

$M_H = 2\left\{\operatorname{cov}(y_{if}, R_{if})/\mu_f - \operatorname{cov}(y_{if}, R_{is})/\mu_f\right\}$ is the reranking index proposed by Atkinson (1980) and Plotnick (1981), which captures the effect of the reshuffling of individuals within the income distribution. M_H is non-negative by definition (see Lambert, 2001), implying that any reranking that does occur has a negative impact on the overall redistributive effect of the income changes. Thus income growth will only reduce inequality if it is both pro-poor in nature and the

resultant vertical mobility effect is not swamped by any offsetting horizontal mobility effect due to the reranking of individuals.

2.4 Income risk and the determinants of structural inequality and vertical mobility

Interest in mobility is not only concerned with movement but also predictability, such that greater mobility may no longer be regarded as socially desirable if it is associated with more pronounced fluctuations and more uncertainty (Shorrocks, 1978a, 1978b): individuals with a preference for income stability and an aversion to risk may choose an income stream with a lower present value if it is both less volatile and more certain. From this perspective, Jantti and Jenkins (2015) note that the Shorrocks mobility measure M_T may be re-interpreted as a measure of income 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, at least some part of the change in individual incomes over time is likely to be predictable in which case the identification of income risk requires the specification of a model of the income generation process, with this approach also providing the basis for analyses of the determinants of structural inequality and vertical mobility.

For this purpose we assume the existence over time of a stable dynamic income function and specify a first-order autoregressive distributed lag (ARDL) model with fixed effects, which allows for both current and lagged effects of income determinants x_j (*j*=1,…*J*), income persistence and individual heterogeneity:

$$y_{i,t+1} = \alpha_0 + \sum_{j=1}^{J} \delta_j x_{ji,t+1} + \sum_{j=1}^{J} \alpha_j x_{jit} + (1-\lambda)y_{it} + \eta_{i,t+1} \; ; \quad i=1, \ldots N; \; t=1, \ldots T\text{-}1 \qquad (4)$$

where $\eta_{i,t+1} = \lambda\mu_i + \varepsilon_{i,t+1}$ is an error term, composed of a fixed individual effect $\lambda\mu_i$ and a period specific disturbance $\eta_{i,t+1}$. Eq.(4) may also usefully be expressed in the form of an Error Correction Model (ECM) of the annual change in income:

$$\Delta y_{i,t+1} = \left(y_{i,t+1} - y_{it} \right) = \sum_{j=1}^{J} \delta_j \Delta x_{ji,t+1} + \lambda \left(y_{it}^* - y_{it} \right) + \varepsilon_{i,t+1} \; ; \qquad (5)$$

$$\text{where } y_{it}^* = \beta_0 + \sum_{j=1}^{J} \beta_j x_{jit} + \mu_i \qquad (6)$$

may be interpreted as a long-run steady-state or equilibrium income function with parameters

$\beta_0 = \alpha_0 / \lambda$ and $\beta_j = \left(\alpha_j + \delta_j \right) / \lambda$, such that $\left(y_{it}^* - y_{it} \right)$ corresponds to the 'equilibrium error' in the

current period and $\lambda \left(0 \le \lambda \le 1 \right)$ determines the rate of adjustment to equilibrium. Hence, the

change in income depends on the effects of contemporaneous changes in the determinants, the initial

extent of any disequilibrium in income and the size of the idiosyncratic income shock. For our

analytical purposes, the main attraction of this representation 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 income inequality due to contemporaneous

changes in income determinants and also how these factors contribute to chronic or persistent

inequality. Eq.(4) collapses to the static model $y_{i,t+1} = y_{i,t+1}^* + \varepsilon_{i,t+1}$ if there are no lagged effects of

income determinants and full adjustment/no persistence in incomes (i.e. if the α_j's ($j=1,...J$) all

equal zero and $\lambda = 1$).

2.4.1 Income risk and the analysis of structural income inequality

We consider a measure of income risk that reflects transitory shocks about long-run or equilibrium

incomes due to both equilibrium errors and contemporaneous income shocks, implicitly identifying

y_{it}^* as the target level of income in year t.[3] Let $\hat{y}_{it}^* = \hat{\beta}_0 + \sum_{j=1}^{J} \hat{\beta}_j x_{jit} + \hat{\mu}_i$ where $\hat{\beta}_0$, $\hat{\beta}_j$ and $\hat{\mu}_i$ are

estimates of the corresponding parameters in the equilibrium income function (6). An index of

income risk in year t may then be defined as follows:

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[3] In the empirical study it is found that the contribution of structural change to vertical mobility is

negligible, which may be taken to indicate that the equilibrium error largely arises due to the

persistence of past income shocks.

$$M_R = 1 - \frac{G\left(\hat{y}_t^*\right)}{G\left(y_t\right)} \tag{7}$$

where the Gini coefficient of predicted equilibrium incomes $G\left(\hat{y}_t^*\right) = 2\,\mathrm{cov}\left(\hat{y}_{it}^*, \hat{R}_{it}^*\right)/\overline{\hat{y}_t^*}$ is interpreted as a measure of chronic or 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. The index will equal one when there is no long-term or structural inequality in which case $G\left(\hat{y}_t^*\right) = 0$, and will equal zero if actual and target incomes are identical in which case $G\left(\hat{y}_t^*\right) = G\left(y_t\right)$. Hence, as with the Shorrocks index M_T, the index will take a value close to one if inequality is largely a short-run phenomenon due to transitory factors whereas if inequality largely arises from long-term differences between farms then the index will take a value close to zero. But note that the normative interpretations of the two indices differ, with higher values of M_R unambiguously less desirable if individuals are averse to instability and risk.

Further information on 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:

$$G\left(\hat{y}_t^*\right) = \sum_{j=1}^{J} 2\hat{\beta}_j\,\mathrm{cov}\left(x_{jit}, \hat{R}_{it}^*\right)/\overline{\hat{y}_t^*} + 2\,\mathrm{cov}\left(\hat{\mu}_i, \hat{R}_{it}^*\right)/\overline{\hat{y}_t^*} = \sum_{j=1}^{J} \frac{\hat{\beta}_j \overline{x}_{jt}}{\overline{\hat{y}_t^*}} CI(x_{jt}, \hat{R}_t^*) + \frac{\overline{\hat{\mu}}}{\overline{\hat{y}_t^*}} CI(\hat{\mu}, \hat{R}_t^*) \tag{8}$$

where $CI(x_{jt}, \hat{R}_t^*)$ is the CI of income determinant x_j ranked by predicted equilibrium income in year t, with corresponding average value $\overline{x}_{jt}$; and $CI(\hat{\mu}, \hat{R}_t^*)$ and $\overline{\hat{\mu}}$ are the corresponding statistics for the individual-specific fixed-effect term. Hence the Gini coefficient is given as a weighted sum of the CIs of the determinants (including the fixed effect) of predicted equilibrium income, 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.

2.4.2 Determinants of vertical mobility

Following Allanson and Petrie (2013), Eq.(5) may also be used to explore the determinants of vertical mobility between consecutive periods. Specifically, if $f=s+1$ then M_V in (3) may be decomposed using (5) to yield:

$$M_V = Pq = \left\{ \sum_{j=1}^{J} \left(CI(\Delta x_{jf}, R_s) - G(y_s) - \right) \frac{\hat{\delta}_j \overline{\Delta x}_{jf}}{\overline{\Delta y}_f} + \left(CI(\hat{y}_s^* - y_s, R_s) - G(y_s) \right) \frac{\hat{\lambda} \overline{(\hat{y}_s^* - y_s)}}{\overline{\Delta y}_f} \right.$$
$$\left. + \left(CI(e_f, R_s) - G(y_s) \right) \frac{\overline{e}_f}{\overline{\Delta y}_f} \right\} \left(\frac{\overline{\Delta y}_f}{\overline{y}_f} \right) \tag{9}$$

where $\overline{\Delta x}_{jf}$ is the average change in income determinant j, with $CI(\Delta x_{jf}, R_s)$ being the corresponding CI ranked by base year income; $\hat{\delta}_j$'s, and $\hat{\lambda}$ are estimates of the corresponding parameters of the dynamic income function (5); $\overline{(\hat{y}_s^* - y_s)}$ is the mean predicted equilibrium error in the base year, with $CI(\hat{y}_s^* - y_s, R_s) = 2\text{cov}\left((\hat{y}_{is}^* - y_{is}, R_{is}) \right) / \overline{(\hat{y}_s^* - y_s)}$ the corresponding CI ranked by base year income; and $e_f = \hat{\varepsilon}_f$ are the regression residuals with mean $\overline{e}_f$ and CI ranked by base year income $CI(e_f, R_s)$. It follows immediately that:

$$M_V = Pq = \sum_{j=1}^{J} \left(CI(\Delta x_{jf}, R_s) - G(y_s) \right) \frac{\hat{\delta}_j \overline{\Delta x}_{jf}}{\overline{y}_f}$$
$$+ \left(CI(\hat{y}_s^* - y_s, R_s) - G(y_s) \right) \frac{\hat{\lambda} \overline{(\hat{y}_s^* - y_s)}}{\overline{y}_f} + \left(CI(e_f, R_s) - G(y_s) \right) \frac{\overline{e}_{fs}}{\overline{y}_f} \tag{10}$$
$$= \sum_{j=1}^{J} P_{\Delta j} q_{\Delta j} + P_{EqE} q_{EqE} + P_e q_e$$

where M_V is given as the sum of contributions due to changes in the J income determinants, the predicted equilibrium error and contemporaneous income shocks in (5). Each term in (10) is expressed in terms of the scale and progressivity of the income changes due to that element, with this further decomposition revealing how the average level of income changes and their distribution across base year income ranks respectively impact on vertical mobility. For example, a positive scale index $q_{\Delta j}$ implies a positive average income impact due to changes in the jth income

determinant and if the poor enjoy a larger share of these income gains than their base year share of income then the progressivity index $P_{\Delta j}$ will be negative giving rise to a reduction in inequality and hence a negative impact on M_V ceteris paribus.

The interpretation of P_{EqE} and q_{EqE} are similar in terms of the impacts of income changes due to the process of adjustment towards the equilibrium levels of income implied by individuals' conditions in the base period,[4] where this process may generally be expected to have a negative impact on M_V and hence reduce inequality. To see this point, note that the contribution of the equilibrium error to M_V can be expressed as:

$$P_{EqE}q_{EqE} = \left\{ \frac{\overline{\hat{y}_s^*}\, CI(\hat{y}_s^*, R_s) - \overline{y}_s G(y_s)}{\left(\overline{\hat{y}_s^* - y_s}\right)} - G(y_s) \right\} \frac{\hat{\lambda}\left(\overline{\hat{y}_s^* - y_s}\right)}{\overline{y}_f} = \left(CI(\hat{y}_s^*, R_s) - G(y_s) \right) \frac{\hat{\lambda}\,\overline{\hat{y}_s^*}}{\overline{y}_f} \quad (11)$$

where $CI(\hat{y}_s^*, R_s)$ is the CI of predicted equilibrium income ranked by actual incomes in the base year. Interpreting $G(\hat{y}_s^*)$ as a measure of chronic or long-run inequality then one would typically 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) and $G(\hat{y}_s^*) < G(y_s)$ in the light of the empirical evidence that measured inequality is lower if farm incomes are averaged over a number of years.

The preceding analysis may readily be extended to consider the determinants of vertical mobility over a multiyear period. Thus, if $f=s+m$ with $m\geq1$ then income changes over this period can be expressed in terms of the dynamic income model as:

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[4] The contribution of the equilibrium error $P_{EqE}q_{EqE}$ in (10) could be further broken down using (6) to identify the 'apparent' contribution of each equilibrium income determinant to M_V through the adjustment process, but this is misleading inasmuch as the causes of the disequilibrium in the base year are unknown. For example, the decomposition would identify the contribution of the jth determinant as $P_{EqE(j)}q_{EqE(j)} = \left(CI(x_{js}, R_s) - G(y_s) \right)\left(\hat{\lambda}\hat{\beta}_j \overline{x}_{js} / \overline{y}_f \right)$, irrespective of whether the base year equilibrium error had arisen due to past changes in that determinant or not.

$$y_{if} - y_{is} = \sum_{j=1}^{J} \Psi_{ij} + \Lambda\left(y_{is}^{*} - y_{is}\right) + \Sigma_i \tag{12}$$

where: $\Psi_{ij} = \sum_{j=1}^{J}\left(\delta_j \Delta x_{ji,s+m} + \left(\sum_{k=1}^{m-1}\left((1-\lambda)^{m-k}\delta_j + \left(\sum_{l=0}^{m-(k+1)} (1-\lambda)^l \right)\lambda\beta_j \right)\Delta x_{ji,s+k} \right) \right);$

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

and which reduces to (5) if $m=1$ with $s\equiv t$. Hence (10) may be generalised to give:

$$M_V = Pq = \sum_{j=1}^{J}\sum_{k=1}^{m}\left(CI(\hat{\Psi}_j, R_s) - G(y_s) \right)\frac{\overline{\hat{\Psi}_j}}{\overline{y}_f}$$

$$+ \left(CI(\hat{y}_s^{*} - y_s, R_s) - G(y_s) \right)\frac{\hat{\Lambda}\overline{(\hat{y}_s^{*} - y_s)}}{\overline{y}_f} + \left(CI(\hat{\Sigma}, R_s) - G(y_s) \right)\frac{\overline{\hat{\Sigma}}}{\overline{y}_f} \tag{13}$$

$$= \sum_{j=1}^{J} P_{\Delta j}^{m} q_{\Delta j}^{m} + P_{EqE}^{m} q_{EqE}^{m} + P_e^{m} q_e^{m}$$

where the $\hat{\Psi}_j$'s, $\hat{\Lambda}$ and $\hat{\Sigma}$ are estimates of the corresponding entities in (12), with corresponding mean values $\overline{\hat{\Psi}}_j$'s, $\overline{\hat{\Lambda}}$ and $\overline{\hat{\Sigma}}$. Therefore vertical mobility in any given multiyear period is the net result of the cumulative effects of changes in income determinants over the period, the equilibrium error in the base year, and the sequence of idiosyncratic shocks to farm incomes.

3. Empirical analysis

We explore the distributional implication of farm income mobility in Scotland over the years 1995 to 2009 using data on an unbalanced panel of farms drawn from the Scottish Farm Accounts Survey (FAS). The study examines mobility over both the whole of the study period and, where of interest, for the two sub-periods defined by the introduction of the Single Farm Payment scheme in 2005.

3.1   Data and variable definitions

The FAS is an annual survey of about 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. The survey is conducted on an accounting year basis with a typical year-end in early March so, for example, the 1995/96 FAS centres on the 1995 production and subsidy year. The farms in the

survey are chosen randomly to be representative of their size and type, where the economic size of the business is measured in terms of standard gross margin prior to 2003/04 and standard labour requirement thereafter, and the farm type classification is based on the relative importance of the various crop and livestock enterprises in terms of standard gross margin.[5]

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, 2013). The analysis is based on an unbalanced panel of 942 farms, of which 151 were present over the whole 15 year period and with a median duration of 7 years. Sample weights are used throughout the analysis with these being based on the number of farms enumerated by size and type in the June Agricultural Census. The results are therefore representative of the population of full-time farms in Scotland, with a sampling fraction of between 3% and 4% over most of the study period.[6] Standard errors for all mean, inequality and mobility measures are generated using a bootstrap procedure that reflects the sample design, with re-sampling carried out at the individual farm level within each stratification class. In particular, bootstrap standard errors for the mobility indices take into account that the income of individual farms will be correlated across years.

Farm income is measured by 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 all their investment in the business (Department of the Environment Food and Rural Affairs (DEFRA),

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[5] 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 pigs and poultry farms).

[6] 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, 2004a).

2002) and is defined as the difference between total receipts and total expenditure. The measure is seen as corresponding closely to the income position as perceived by the farmer, 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 only determinants of farm income included in the ARDL model specification, other than year slope dummies, are crop and livestock standard gross margins (SGMs). 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 consistency over time by using the most recent set of SGM coefficients available, based on Scottish averages for the years 1998 to 2002, to calculate SGMs for the entire period.

3.2 Reduction of longer-term inequality

The first two columns of Table 1 indicate the considerable level of fluctuation in farm incomes over the production years 1995 to 2009, with the coefficient of variation of average annual cash incomes equal to 20% over the period. 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. Changes in the Gini coefficient reflect changes in both the absolute dispersion and mean level of cash incomes, with relative inequality generally higher in years of lower average incomes.

Turning to the instability of individual farm incomes then M_T approaches an asymptotic value of about 13% as the measurement period is extended from the base year of 1995. Two tentative conclusions may be drawn from this finding. First most income changes reflect short to

medium term variability due to transitory shocks, with no further equalisation once relative incomes have approached their long-term or 'permanent' values after about 10 years. Second, the overwhelming bulk of cross-sectional inequality, as measured by annual Gini coefficients, may be thought of as being structural in nature inasmuch as it reflects persistent differences in incomes between farms. Choosing 2005 as the base year leads to higher values of M_T for any given $T>1$, but not too much should be read into this finding as the index is sensitive to the choice of base year for small T. 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.

3.3   Individual income growth and positional change

The top panel of Table 2 presents the decomposition of annual changes in the Gini coefficient into vertical and horizontal components based on Eq. (3), where the results are generated using observations on all farms present in both the base and final year, and are therefore not strictly comparable either with the annual summary statistics presented in Table 1 or between pairs of years. These results reveal three main points of interest. First, the vertical mobility index M_V is significantly negative in all cases, indicating that expected annual income changes had an equalising effect throughout the period. Thus farms with low incomes in one year experienced either larger relative income gains on average than richer farms over the following year or smaller relative losses, although most of the progressivity index estimates and some of the scale factors are not significantly different from zero. Second, the horizontal mobility index M_R 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 progressive effect of expected income changes was only sufficient to outweigh the disequalising 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 are not independent of base year incomes may be interpreted as a rejection of Gibrat's Law of proportionate effect. However this conclusion needs to be treated with some caution as the apparent progressivity of farm income growth may simply reflect regression to the mean: if individual incomes are subject to idiosyncratic shocks (or measurement errors) that are uncorrelated over time then the expected income increase of a farm with below-average income will be positive while that of farm with above-average income will be negative. 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 multiperiod decomposition analyses, which have been generated using observations on all farms present throughout the relevant measurement period. As before, we find that vertical mobility was significantly negative, with the incomes of poorer farms appearing to grow faster than richer ones even over the full study period. Nevertheless inequality rose over the first sub-period and the full period but not over the second sub-period from 2004 to 2009.

Our other two robustness checks employ income smoothing and instrumental variable (IV) 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 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 and lead incomes.[7] 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 the analysis is done using smoothed income data then income growth is still significantly pro-poor in most cases but the extent of vertical mobility is typically reduced somewhat. Conversely only two of the IV estimates of M_V are significantly different from zero, one negative and one positive, which might suggest that neither poorer nor richer farms are favoured by structural developments in the industry. In conclusion, the results provide at best only weak evidence against Gibrat's law, with transitory shocks likely to account for at least some and maybe all of the observed bias of annual income growth rates in favour of poorer farms.

3.4 Income risk and the determinants of structural inequality and vertical mobility

The further analysis of mobility is based on estimates of the dynamic income model (5). OLS estimates of (5) 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 crop and livestock SGM averages, and further control for initial conditions in the manner of Wooldridge (2005) by including the level of income in the year in which a farm first entered the sample as a separate explanatory variable. 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 structural

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[7] Noting 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 suggested IV procedure intended to eliminate possible correlation between the 'explanatory' rank variable and the 'error term' in this regression.

inequality, and avoids the further restriction of the sample that would result from the use of GMM estimators as these require higher-order lags of income to serve as instruments.[8]

The first set of columns in Table 4 report estimation results for the Error Correction Model (5) with the dependent variable being the annual change in cash income. The first two regression coefficients show the short-run impact of changes in the size of the crop and livestock enterprises, where these impacts are of very similar size (0.123 and 0.119) and imply that a £1 increase in the economic size of the farm as measured in terms of standard gross margins will result in an immediate increase in cash incomes of roughly 12 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 0.494, implying that just under 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 equilibrium income function (6), which are also reported in the Table. Taking the reference year of 1996 as an example, the implied long-run effects of changes in the size of cropping and livestock enterprises (0.224 and 0.590 respectively) were two and five times larger than the corresponding contemporaneous effects given prevailing agroeconomic conditions. Long-run income effects of changes in enterprise sizes are predicted to have been positive in nearly all years, being on average 1.7 and 2.6 times the corresponding impact effects for cropping and livestock respectively. Finally, farm-level fixed effects are not significantly affected by farm-average enterprises sizes but there is a significant positive relationship with the level of income in the year in which the farm first entered the sample.

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[8] Alternative estimators lead to differing estimates of the adjustment parameter λ, with the preferred estimator yielding a value between the downwardly biased OLS estimator and the upwardly biased within-groups estimator, but broadly similar estimates of the other model parameters. See Flannery and Hankins (2013) on the relative performance of alternative dynamic panel model estimators under differing assumptions about dataset characteristics.

3.4.1 Income risk and the analysis of structural income inequality

Table 5 reports the long-run or equilibrium Gini coefficient estimates for each year in the study period where these are smaller than the corresponding annual Gini coefficients (repeated from Table 1) except in 1996, which is consistent with the finding that averaging income over a number of years typically reduces inequality. Accordingly the income risk index M_R takes values in the unit interval in all but 1996, with an average value of 24% over the entire period indicating a substantial degree of income risk due to the volatility of agricultural incomes. This figure is somewhat higher than the asymptotic value of about 13% for the Shorrocks Index M_T, though both measures imply that the overwhelming bulk of farm income inequality is structural in nature.

The remainder of the Table provides results from the analysis of the determinants of equilibrium or structural inequality. On average, just over half (51.8%) of equilibrium inequality in farm incomes was due to observable differences in the size of farm businesses, as measured by economic size units: larger cropping and livestock enterprises tend to be located on farms generating higher cash incomes so the typically positive contributions of these enterprises to farm income is a source of inequality. Cropping enterprises were somewhat more unequally distributed than livestock enterprises relative to farm income in most years and therefore contribute somewhat more to equilibrium inequality than might be inferred from their relative importance in Scottish agriculture.

This leaves the remaining half (48.2%) of structural inequality attributable to farm-level fixed effects, where these effects make a highly significant contribution in all 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, weather 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, 2012) 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 structure between farms, which will affect farms' cash incomes but are not taken into account in the calculation of SGMs. Thus chronic inequality is as much due to differences in the cash income generating performance of farms as in their economic size as conventionally measured.

3.4.2 Determinants of vertical mobility

Table 6 expands upon the results of the decomposition analysis in Table 2, identifying the separate contributions of changes in cropping and livestock enterprise sizes, the equilibrium error and the residual to the vertical mobility index M_V.[9]

First, changes in the economic size of enterprises 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 the entire study period. Further investigation reveals that this is because those farms that identifiably remained in the FAS tended to be ones that did not change in size appreciably rather than because any such size changes as did occur among these farms were distributionally neutral. Thus the average growth in the economic size of farms between 1995 and 2010 is estimated to have been just 3.6% (from £54225 to £56248) based on the 151 farms that identifiably remained in the FAS throughout the entire period whereas the corresponding estimate based on the full sample is 41.7% (from £55212 to £81226). This discrepancy reflects the impossibility within FAS of tracking farms that are subject to significant structural change, such as amalgamation with another farm, because such farms are assigned a new identifier if they continue to participate in the survey (SEERAD, 2004b). Thus, although the FAS is by design representative of the structure of Scottish agriculture over time, it is

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[9] The contribution of changes in the parameters of the equilibrium income function (6) due to the year slope dummies is subsumed within that of the equilibrium error in the multiyear change analyses.

not suited to the analysis of the pattern of individual farm size movements driving structural changes in farm income inequality.[10]

The contribution of the equilibrium error to vertical mobility is significantly negative in every year, which is consistent with our expectations based on the discussion of Eq. (11). More intuitively, adjustment towards equilibrium is equalizing since the income correction of a farm with below-equilibrium income due to a previous negative idiosyncratic shock will be positive while that of farm with above-equilibrium income will be negative. This results in income growth appearing to be pro-poor despite the lack of evidence of structural change provided by the FAS panel, with the equilibrium error accounting on average for all of the apparent progressivity of annual income changes over the period.

Finally, 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 only significant in a few years 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 crop and livestock enterprise slope dummies. By construction the residual is uncorrelated with lagged income over the full panel.

4. Conclusions

Income mobility is an inherently multidimensional concept with this paper seeking to provide a rounded evaluation of the distributional consequences of farm income mobility by considering a range of alternative measures. In particular, this multifaceted approach reveals both the temporal character of farm income inequality − whether it is a short-run phenomenon due to transitory

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[10] The principal objective of FAS is to monitor changes in farm performance from year to year based on an "identical sample" of farms (see, e.g., SEERAD, 2004a, 2004b).

shocks, as opposed to a chronic or persistent problem due to structural factors − and the nature of the dynamic processes driving changes in farm income inequality over time. It also highlights the trade-off between the perceived value on the one hand of the reduction of longer-term inequality due to the averaging out of transitory shocks over time and the increase in instability and risk on the other due to the unpredictability of incomes, where both phenomena may be seen to arise from the stochastic nature of the income growth process (see Allanson, 2012, for further discussion).

The empirical analysis of farm income mobility in Scottish agriculture is based on data from the Farm Accounts Survey (FAS), which is well suited in principle to the task since farms, once recruited, can remain in the survey for an unlimited length of time. However the sample is subject to a 'virtual' form of selective attrition whereby farms in the survey that experience significant structural change are assigned a new identifier and thereby become untraceable through time.[11] As a result, balanced panels of farms constructed from the FAS will exhibit lower levels of structural change than the sector as a whole, casting doubt on the use of the survey to investigate the nature of the dynamic processes driving changes in cross-sectional inequality in terms of vertical and horizontal mobility components. Moreover estimates of the Shorrocks mobility index, which require the calculation of multiyear average incomes, will largely capture the impact of transitory income shocks *per se* due to the limited confounding by structural changes. Finally measures based on the estimation of the Error Correction Model using the full unbalanced panel of survey farms, namely the income risk index and the determinants of structural inequality, should be unbiased though less precise than would have otherwise been the case.

The empirical results reveal that farm income inequality is partly a temporary or short-run phenomenon, with the estimates of the Shorrocks mobility and income risk indices implying that

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[11] 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).

somewhere between one eighth and one quarter of inequality in annual incomes may be due to transitory 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). The most recent CAP reform includes 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.

Nevertheless the overwhelming bulk of farm income inequality is shown to have been structural in nature. The subsequent decomposition analysis reveals that just over half of this long-run inequality was, on average, due to differences in the economic size of farm businesses, with the remainder due to farm-level fixed effects that represent differences in both financial performance and business structure. The move in Scotland 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), with the Scottish Government seeking to limit the resultant scale of farm income redistribution by adopting a regionalised model in which regional payment rates reflect the productive capacity of the land (Scottish Government, 2014).

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Table 1. Basic summary statistics and Shorrocks mobility index M_T

| Year | Annual summary statistics | | | Multiyear analysis | | | | | |
| | | | | Base year =1995 | | | Base year =2005 | | |
	$\bar{y}_t$	$G(y_t)$	T	$G(y_A)$	M_T	T	$G(y_A)$	M_T
1995	40489***	0.505 ***	1	0.505***	0.000***		-	
	1693	*0.017*		*0.017*	*0.000*			
1996	43707***	0.447 ***	2	0.442***	0.057***		-	
	1415	*0.015*		*0.015*	*0.018*			
1997	27644***	0.512 ***	3	0.442***	0.067***		-	
	1008	*0.020*		*0.014*	*0.010*			
1998	29318***	0.511 ***	4	0.448***	0.067***		-	
	1170	*0.023*		*0.016*	*0.009*			
1999	27069***	0.543 ***	5	0.455***	0.067***		-	
	1361	*0.021*		*0.017*	*0.010*			
2000	28641***	0.550 ***	6	0.458***	0.083***		-	
	1610	*0.023*		*0.017*	*0.012*			
2001	29523***	0.554 ***	7	0.459***	0.088***		-	
	1424	*0.024*		*0.018*	*0.012*			
2002	27610***	0.518 ***	8	0.470***	0.087***		-	
	1138	*0.022*		*0.020*	*0.013*			
2003	36570***	0.452 ***	9	0.452***	0.103***		-	
	1489	*0.018*		*0.032*	*0.023*			
2004	36327***	0.490 ***	10	0.448***	0.106***		-	
	1274	*0.020*		*0.032*	*0.022*			
2005	31263***	0.494 ***	11	0.417***	0.129***	1	0.494 ***	0.000 ***
	1172	*0.020*		*0.023*	*0.022*		*0.020*	*0.000*
2006	34245***	0.525 ***	12	0.395***	0.130***	2	0.475 ***	0.070 ***
	1417	*0.020*		*0.025*	*0.023*		*0.018*	*0.012*
2007	44568***	0.536 ***	13	0.385***	0.139***	3	0.467 ***	0.091 ***
	1757	*0.021*		*0.027*	*0.027*		*0.019*	*0.014*
2008	46834***	0.513 ***	14	0.390***	0.136***	4	0.449 ***	0.117 ***
	1726	*0.016*		*0.028*	*0.027*		*0.021*	*0.018*
2009	48935 ***	0.480***	15	0.390***	0.134***	5	0.434 ***	0.136 ***
	1820	*0.017*		*0.030*	*0.028*		*0.022*	*0.017*

*Source: Authors' calculations based on Eqs.(1) and (2). 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. 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***	−1.58	0.078***	0.090***
	0.015	*0.016*	*0.016*	*0.018*	*4.57*	*0.029*	*0.022*
1996-1997	0.431***	0.492***	0.061***	−0.031**	0.05**	−0.571***	0.092***
	0.012	*0.016*	*0.016*	*0.015*	*0.03*	*0.042*	*0.009*
1997-1998	0.515***	0.523***	0.008	−0.105***	−2.09	0.050	0.113***
	0.020	*0.026*	*0.026*	*0.030*	*500.30*	*0.040*	*0.017*
1998-1999	0.510***	0.544***	0.034	−0.073**	0.94	−0.078*	0.107***
	0.025	*0.019*	*0.025*	*0.029*	*9.52*	*0.040*	*0.013*
1999-2000	0.561***	0.514***	−0.047**	−0.173***	−2.24	0.077*	0.126***
	0.023	*0.018*	*0.022*	*0.026*	*16.48*	*0.043*	*0.017*
2000-2001	0.523***	0.533***	0.009	−0.162***	−1.66	0.098**	0.172***
	0.018	*0.023*	*0.024*	*0.029*	*31.84*	*0.043*	*0.026*
2001-2002	0.562***	0.510***	−0.052**	−0.236***	−15.91	0.015	0.184***
	0.025	*0.023*	*0.025*	*0.055*	*102.22*	*0.054*	*0.043*
2002-2003	0.497***	0.512***	0.015	−0.074**	−0.63**	0.118***	0.089***
	0.032	*0.040*	*0.029*	*0.029*	*0.30*	*0.031*	*0.012*
2003-2004	0.452***	0.464***	0.012	−0.120***	1.71	−0.070*	0.132***
	0.019	*0.022*	*0.019*	*0.023*	*131.63*	*0.041*	*0.019*
2004-2005	0.484***	0.492***	0.008	−0.116***	0.79**	−0.147***	0.124***
	0.020	*0.019*	*0.019*	*0.023*	*0.34*	*0.036*	*0.013*
2005-2006	0.494***	0.531***	0.037*	−0.104***	−1.32	0.079**	0.141***
	0.021	*0.020*	*0.021*	*0.029*	*11.80*	*0.031*	*0.017*
2006-2007	0.527***	0.525***	−0.002	−0.138***	−0.55***	0.251***	0.136***
	0.022	*0.020*	*0.020*	*0.027*	*0.09*	*0.028*	*0.019*
2007-2008	0.527***	0.510***	−0.017	−0.167***	6.01	−0.028	0.151***
	0.022	*0.019*	*0.019*	*0.024*	*75.07*	*0.038*	*0.019*
2008-2009	0.517***	0.481***	−0.036**	−0.163***	−9.18	0.018	0.127***
	0.017	*0.016*	*0.016*	*0.017*	*550.36*	*0.029*	*0.012*
1995-2009	0.459***	0.505***	0.046*	−0.114***	−1.07	0.106***	0.160***
	0.026	*0.026*	*0.024*	*0.030*	*2.30*	*0.041*	*0.023*
1995-2004	0.482***	0.535***	0.053*	−0.124***	0.39***	−0.318***	0.177***
	0.019	*0.027*	*0.029*	*0.030*	*0.12*	*0.060*	*0.020*
2004-2009	0.495***	0.486***	−0.008	−0.152***	−0.49***	0.310***	0.144***
	0.022	*0.021*	*0.022*	*0.022*	*0.08*	*0.025*	*0.016*

Source: Authors' calculations based on Eq.(3). 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

Period	Estimation technique		
	Standard	*Smoothed incomes*	*Instrumented ranks*
1996-1999	0.031	-0.004	0.081***
	0.019	*0.014*	*0.023*
1997-2000	-0.093***	0.003	0.021
	0.035	*0.020*	*0.027*
1998-2001	-0.171**	-0.096***	-0.100**
	0.078	*0.035*	*0.050*
1999-2002	-0.149***	-0.127*	-0.015
	0.038	*0.069*	*0.032*
2000-2003	-0.162***	-0.071***	0.011
	0.035	*0.021*	*0.028*
2001-2004	-0.172***	-0.061***	-0.010
	0.039	*0.019*	*0.032*
2002-2005	-0.065*	-0.021	0.042
	0.034	*0.019*	*0.026*
2003-2006	-0.069*	-0.018	0.043
	0.039	*0.019*	*0.028*
2004-2007	-0.135***	-0.018	0.017
	0.040	*0.016*	*0.026*
2005-2008	-0.114***	-0.051***	0.006
	0.027	*0.017*	*0.025*
1996-2008	-0.017	-0.068***	0.039
	0.026	*0.020*	*0.032*

Source: Authors' calculations based on Eq.(3). Each statistic is based on the sample of farms that are present not only in the base and final years of each period but also in the years immediately before and afterwards 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 function

Explanatory variables		Error Correction Model (5)		Equilibrium function (6)	
		Coeff.	*Robust Std error*	Coeff.	*Bootsrapped Std error*
Change in cropping SGM	$\Delta x_{Ci,t+1}$	0.123*	*0.065*	-	
Change in livestock SGM	$\Delta x_{Li,t+1}$	0.119*	*0.068*	-	
Lagged income	y_{it}	-0.494***	*0.022*	-	
Lagged cropping SGM	x_{Cit}	0.111	*0.110*	0.224	*0.215*
Dummy1996* x_{Cit}		-0.233**	*0.092*	-0.471 ***	*0.179*
Dummy1997* x_{Cit}		0.006	*0.092*	0.013	*0.183*
Dummy1998* x_{Cit}		0.039	*0.089*	0.079	*0.176*
Dummy1999* x_{Cit}		-0.036	*0.092*	-0.074	*0.183*
Dummy2000* x_{Cit}		-0.089	*0.093*	-0.181	*0.182*
Dummy2001* x_{Cit}		-0.029	*0.092*	-0.058	*0.177*
Dummy2002* x_{Cit}		0.052	*0.088*	0.105	*0.174*
Dummy2003* x_{Cit}		-0.087	*0.093*	-0.176	*0.178*
Dummy2004* x_{Cit}		-0.093	*0.089*	-0.187	*0.169*
Dummy2005* x_{Cit}		0.148	*0.098*	0.299	*0.189*
Dummy2006* x_{Cit}		0.336***	*0.091*	0.680 ***	*0.182*
Dummy2007* x_{Cit}		-0.142	*0.107*	-0.287	*0.214*
Dummy2008* x_{Cit}		0.048	*0.096*	0.098	*0.188*
Lagged livestock SGM	x_{Lit}	0.291***	*0.058*	0.590 ***	*0.118*
Dummy1996* x_{Lit}		-0.262***	*0.028*	-0.530 ***	*0.063*
Dummy1997* x_{Lit}		-0.209***	*0.028*	-0.423 ***	*0.058*
Dummy1998* x_{Lit}		-0.296***	*0.026*	-0.600 ***	*0.053*
Dummy1999* x_{Lit}		-0.174***	*0.035*	-0.352 ***	*0.072*
Dummy2000* x_{Lit}		-0.010	*0.035*	-0.020	*0.067*
Dummy2001* x_{Lit}		-0.222***	*0.035*	-0.450 ***	*0.064*
Dummy2002* x_{Lit}		-0.111***	*0.033*	-0.225 ***	*0.066*
Dummy2003* x_{Lit}		-0.094***	*0.031*	-0.190 ***	*0.062*
Dummy2004* x_{Lit}		-0.217***	*0.032*	-0.440 ***	*0.066*
Dummy2005* x_{Lit}		-0.169***	*0.030*	-0.342 ***	*0.061*
Dummy2006* x_{Lit}		-0.056	*0.036*	-0.114 *	*0.068*
Dummy2007* x_{Lit}		-0.042	*0.037*	-0.086	*0.076*
Dummy2008* x_{Lit}		-0.087**	*0.042*	-0.176 **	*0.079*
Average cropping SGM	$\bar{x}_{Ci.}$	-0.007	*0.066*	-0.014	*0.127*
Average livestock SGM	$\bar{x}_{Li.}$	0.004	*0.054*	0.009	*0.107*
Income in sample entry year	y_i^{entry}	0.178***	*0.019*	0.360 ***	*0.032*
Constant	$\hat{\beta}_0$	2300.301***	*516.871*	4659.105 ***	*1007.455*
Sample size		6045			
R^2		0.23119			
F(34,6010)		35.88-			

*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. Income risk and the determinants of structural inequality

| Year | Equilibrium inequality $G(y_t^*)$ | Observed inequality $G(y_t)$ | Income Risk M_R | Contribution to equilibrium inequality $G(y_t^*)$ | | |
				Cropping SGM	Livestock SGM	Fixed Effects
1995	0.379 ***	0.505***	0.250***	0.051	0.176***	0.152 ***
	0.018	*0.017*	*0.032*	*0.047*	*0.053*	*0.044*
1996	0.455 ***	0.447***	-0.019	0.086	0.050	0.319 ***
	0.125	*0.015*	*0.279*	*0.144*	*0.103*	*0.120*
1997	0.376 ***	0.512***	0.267***	0.113**	0.047	0.216 ***
	0.024	*0.020*	*0.051*	*0.057*	*0.040*	*0.060*
1998	0.427 ***	0.511***	0.164**	0.194**	-0.002	0.235 ***
	0.032	*0.023*	*0.065*	*0.087*	*0.021*	*0.077*
1999	0.358 ***	0.543***	0.340***	0.063	0.083	0.212 ***
	0.021	*0.021*	*0.040*	*0.061*	*0.051*	*0.064*
2000	0.381 ***	0.550***	0.307***	0.007	0.208***	0.166 ***
	0.023	*0.023*	*0.047*	*0.022*	*0.057*	*0.047*
2001	0.386 ***	0.554***	0.302***	0.093	0.045	0.249 ***
	0.030	*0.024*	*0.055*	*0.072*	*0.045*	*0.069*
2002	0.376 ***	0.518***	0.274***	0.102**	0.108***	0.166 ***
	0.019	*0.022*	*0.039*	*0.046*	*0.041*	*0.049*
2003	0.337 ***	0.452***	0.254***	0.012	0.130***	0.195 ***
	0.023	*0.018*	*0.051*	*0.035*	*0.048*	*0.046*
2004	0.353 ***	0.490***	0.278***	0.010	0.083	0.261 ***
	0.029	*0.020*	*0.056*	*0.032*	*0.074*	*0.070*
2005	0.383 ***	0.494***	0.225***	0.195***	0.059*	0.129 ***
	0.025	*0.020*	*0.053*	*0.065*	*0.036*	*0.050*
2006	0.409 ***	0.525***	0.220***	0.240***	0.088***	0.081 **
	0.016	*0.020*	*0.035*	*0.040*	*0.030*	*0.035*
2007	0.413 ***	0.536***	0.231**	0.004	0.268***	0.141 **
	0.058	*0.021*	*0.108*	*0.037*	*0.079*	*0.062*
2008	0.374 ***	0.513***	0.272***	0.106**	0.128**	0.140 ***
	0.019	*0.016*	*0.037*	*0.050*	*0.052*	*0.048*
2009	0.372 ***	0.480***	0.225***	0.058	0.193***	0.122 ***
	0.024	*0.017*	*0.049*	*0.057*	*0.065*	*0.044*
Average contribution to equilibrium inequality				22.6%	29.2%	48.2%

*Source: Authors' calculations based on Eqs. (7) and (8). All summary statistics based on the full sample available in the relevant year. Bootstrapped standard errors in italics based on 1000 replications. Statistical significance at 1%, 5% and 10% levels are denoted by ***, ** and * respectively.*

Table 6. Decomposition of vertical mobility M_V

		Contribution to vertical mobility M_V			
Period	M_V	*Change in cropping SGM*	*Change in livestock SGM*	*Equilibrium error*	*Residual*
1995-1996	−0.1229***	−0.0005	−0.0004	−0.0778***	−0.0442***
	0.0182	*0.0011*	*0.0009*	*0.0105*	*0.0169*
1996-1997	−0.0306**	0.0006	−0.0007	−0.0443***	0.0138
	0.0146	*0.0014*	*0.0008*	*0.0120*	*0.0193*
1997-1998	−0.1048***	0.0007	0.0005	−0.1109***	0.0049
	0.0299	*0.0010*	*0.0008*	*0.0181*	*0.0284*
1998-1999	−0.0727**	−0.0033	0.0008	−0.0936***	0.0234
	0.0286	*0.0049*	*0.0009*	*0.0197*	*0.0232*
1999-2000	−0.1729***	−0.0014	0.0012	−0.1488***	−0.0239
	0.0258	*0.0017*	*0.0009*	*0.0209*	*0.0244*
2000-2001	−0.1623***	0.0033	0.0013*	−0.1749***	0.0080
	0.0287	*0.0030*	*0.0008*	*0.0230*	*0.0274*
2001-2002	−0.2360***	−0.0007	0.0004	−0.1391***	−0.0966**
	0.0548	*0.0026*	*0.0007*	*0.0300*	*0.0493*
2002-2003	−0.0741**	−0.0006	−0.0005	−0.1399***	0.0669*
	0.0290	*0.0022*	*0.0009*	*0.0320*	*0.0400*
2003-2004	−0.1199***	−0.0013	−0.0004	−0.1035***	−0.0147
	0.0233	*0.0013*	*0.0006*	*0.0155*	*0.0244*
2004-2005	−0.1160***	0.0006	−0.0001	−0.0866***	−0.0299
	0.0226	*0.0018*	*0.0005*	*0.0143*	*0.0246*
2005-2006	−0.1041***	0.0011	0.0003	−0.1711***	0.0656***
	0.0289	*0.0012*	*0.0015*	*0.0271*	*0.0205*
2006-2007	−0.1378***	0.0004	−0.0005	−0.1637***	0.0261
	0.0267	*0.0007*	*0.0009*	*0.0177*	*0.0258*
2007-2008	−0.1674***	−0.0004	0.0010	−0.1334***	−0.0347
	0.0244	*0.0011*	*0.0010*	*0.0172*	*0.0224*
2008-2009	−0.1627***	0.0005	0.0012	−0.1252***	−0.0390**
	0.0174	*0.0010*	*0.0012*	*0.0169*	*0.0181*
1995-2009	−0.1140***	−0.0067	−0.0075	−0.1269***	0.0270
	0.0299	*0.0097*	*0.0218*	*0.0281*	*0.0435*
1995-2004	−0.1239***	−0.0080	0.0025	−0.1638***	0.0454
	0.0302	*0.0097*	*0.0105*	*0.0307*	*0.0461*
2004-2009	−0.1525***	0.0007	0.0083	−0.2466***	0.0851***
	0.0224	*0.0036*	*0.0087*	*0.0355*	*0.0289*

Source: Authors' calculations based on Eqs. (10) and (13). 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.