Welfare Decomposition in the Context of the Life Cycle of Farm Operators: What Does a National Survey Reveal?

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This paper examines the role of the life cycle in impacting the distribution of a combined income and wealth measure using data from the 2001 and 2006 Agricultural Resource Management Survey. Such an assessment is made using both graphical representation of the distribution of the well-being measure along with utilization of the social welfare decomposition procedure. Results show a mild yet statistically insignificant improvement in the distribution of the economic measure over the five-year period. Contribution to social welfare is found highest among the cohort where the age of the head of household is between 45 and 54 years. Targeted programs are found to enhance social welfare if they are aimed towards cohorts where the age of the head of household is younger than 35 years or where the age of the head of household is in the 35-to-44 age group, depending on whether the analysis is based on a per-farm household or on a per-capita basis.

Key Words: ARMS, economic well-being, Gini coefficient, Lorenz curve, welfare decomposition

An important aspect of the well-documented diversity of U.S. farm households is the strong influence of operators' life-cycle stages.[1] Household resource allocation decisions are, in part, determined by a person's age and family circumstances—the factors that shape their life cycle. These consumption, savings, labor allocation, and production choices are ultimately reflected in two key financial measures: household earnings and wealth. Farm policy has, for many years, advanced as a goal the notion of maintaining incomes of farmers due to the highly risky nature of agriculture production. As a result, most assessments of farm household well-being and resulting policy prescriptions have relied on the distribution of farm household income and comparisons with other household groups. Since the inception of price and income support programs over 70 years ago, a key stimulus for legislative action has been disparity between the income of farm and non-farm households (Gardner 1992, Houthakker 1967).

Recent evidence from the USDA's 2006 Agricultural Resource Management Survey (ARMS) shows that, on average, the incomes of farm households ($76,224), except for those whose operators are 65 years or older, tend to exceed by nearly 15 percent the incomes of all U.S. households ($66,570). This trend of incomes of farm households exceeding the incomes of all U.S. households was established in the early 1990s, with the incomes of households operated by either farmers aged 35 years or younger or by farmers older than 65 being the exception. The fact that the farm households are no longer economically disadvantaged in comparison to their counterpart in the general population as evidenced by the closing of their income gap, with the exception of those farm households operated by either younger or older operators, has implications for targeting policies aimed at income stabilization as well as redistribution.

The life-cycle hypothesis of savings proposes that household savings and consumption reflect

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[1] The term "life cycle" refers to the various stages through which a person passes during a lifetime. The concept provides a coherent linkage between an individual's consumption patterns and expectations on income and savings as the individual passes from childhood, through education, training, labor participation, and retirement. For farm operators, it can trace the stages of the farm business from entry into farming, growth of the farm, consolidation, and retirement and transfer of the business to the next generation.

the life-cycle stage of the household, and that consumption is a linear function of available cash and the discounted value of future income (Ando and Modigliani 1963, Modigliani 1986). A study by Kotlikoff and Summers (1981), however, challenged the traditional view that life-cycle saving is the dominant source of capital accumulation by noting that the bulk of accumulated wealth is due to bequests and intergenerational transfers instead. Assuming that income will increase during working years and decline at retirement (see Paglin 1975, Kearl and Pope 1983, Ahearn, Perry, and El-Osta 1993, Pudney 1993), farm households tend to borrow when they are young, save during middle age, and spend down during retirement. So at any point in time, income and wealth differences among farm households can be attributed to their relative position in the life cycle.

It has been the contention of many economists that agricultural policy would be better informed if the assessment of economic well-being considered both income and wealth in the context of the life cycle (Mishra et al. 2002). A study by Hill (2002) points out that wealth is important not only because it generates income in a variety of forms, but also because it provides security, freedom to maneuver resources, and economic and political power. Until recently, data limitations precluded this type of comprehensive approach. ARMS has evolved to become the most extensive annual national resource for farm business and household economics along with capturing key demographic information pertaining to the farm operator household. Using two time periods (2001 and 2006) that represent the latter years of a farm bill, and data from ARMS, this paper seeks primarily to examine the inequality in the distribution of a composite measure of economic well-being (CWB) that combines annualized wealth and the money income of U.S. farm households.[2] Because the goal of policymakers is to increase both equity and efficiency (see OECD 1998), the analysis also aims for the decomposition of economic disparity and of the social welfare function in the context of the life cycle. As farmers begin to advance in age, they tend to become less dependent on farm earnings for their livelihood (Mishra et al. 2002). In addition, they begin to shift to less labor- and capital-intensive production, and in fact may also begin to reduce the size of their operation by renting or selling part of their assets to younger, more productive operators (Gale 1994). Because of this, older farmers tend to have different income-generating strategies and different portfolios of farm assets than younger farmers, which, in and by itself, furthers the need for incorporating the life-cycle effects when examining the disparity in economic well-being among farm households.

Literature Review

Assessing how best to model the size distributions of economic units, both theoretically and empirically, has received much attention in the literature since the work of early economists and of social statisticians. Dagum (1980, 1999), for example, refers to the early 1800s work by David Ricardo, which dealt with, among other things, income distributions among owners of factors of production, while Kleiber and Kotz (2003) point to the work of Vilfredo Pareto on the modeling of personal income and of wealth distributions over 100 years ago. A graphical representation of inequality of income and/or wealth distributions along with the introduction of a summary statistic based on this work were introduced by, respectively, the American economist Max Lorenz and the Italian statistician Corrado Gini.

Since the pioneering work of these classical economists and social scientists, and particularly over the last fifty years, studies on the distribution of income and wealth have continued to be introduced. While some of these studies had a similar yet more advanced utilization of the concept of inequality, others had a notable shift in emphasis. Examples of research with a broader economic and statistical implementation of the concept of inequality are those by Singh and Maddala (1976), Dagum (1977), Atkinson and Harrison (1978), Champernowne and Cowell (1998), and Sen (1997). Biancotti (2006) points to investigations by Kuznets (1955) and Kaldor (1956) where the focus of the research was on within-country inequality and its link to growth, either assessing

[2] The two farm bills that underlie this paper in terms of their impact on the income and asset values of farm households are the 1996 Federal Agriculture Improvement and Reform (FAIR) Act and the Farm Security and Rural Investment Act (FSRIA). For more specific information on the associated wealth effects of government payments, see Orden (2003) and Burfisher and Hopkins (2004).

how economic disparities impact development, or explaining the pattern of inequality in any single economy that may result as a consequence of progress. The specific proposition of Kuznets' model, which was shown to be more suitable for developing economies, is that economic growth can push up inequality before ultimately causing it to fall. A model by Okun (1975), which is deemed more adequate for developed economies (Burtless 2002), points to a potential sacrifice in efficiency when such economies enact regulatory schemes or income distribution programs tailored at ensuring equity in the final distribution of income.

While many researchers have examined the inequality in the distribution of either income or wealth of agricultural households using the concept of the Gini coefficient, the literature is scant when it comes to similar work for a combined income and wealth measure. Studies by Larson and Carlin (1974) and by Ahearn, Johnson, and Strickland (1985), for example, examined the economic well-being of farm households in the United States by assessing the extent of disparity in the size distributions of income and/or wealth. Findeis and Reddy (1987) and Reddy, Findeis, and Hallberg (1988) measured the inequality in the distribution of income for farm families, by region, using the 1995 Current Population Survey. Boisvert and Ranney (1990) measured income inequality for New York dairy farmers using the concept of the Gini coefficient with adjustment for the presence of negative incomes. Gould and Saupe (1990) assessed the disparity of a combined income and wealth measure for farmers using Wisconsin panel data. El-Osta, Bernat, and Ahearn (1995) extended the previous studies by exploring specifically the impact of government payments and off-farm work on income inequality using data from the 1991 Farm Costs and Returns Survey. Other studies by El-Osta and Morehart (2002) and by El-Osta and Mishra (2005) examined, using ARMS data, the role of value judgments concerning society's level of aversion to wealth concentration and the role of farm subsidies on wealth dispersion among farm households.

For the purpose of this study, the disparity in the economic well-being among farm households is first examined using the whole distribution of the combined income and wealth index, which is done in the form of distributing the shares of this index based on its quintiles and by time period. Next, the corresponding Lorenz curves (both ordinary and generalized) are plotted and the concept of the Gini coefficient is utilized in order to, respectively, provide graphical representation of inequality and provide a corresponding single summary statistic of dispersion for each of the selected groups of households.[3] Resulting statistics are then used to derive social welfare functions, which, in turn, are disaggregated in order to allow for the discernment of the contribution of five age cohorts towards overall welfare disparity. This method of decomposing welfare disparity, which was developed by Podder (1993) and later applied by Mukhopadhaya (2002), can help policymakers by identifying which age cohort might benefit the most from a targeted economic policy aimed at reducing disparity in the overall distribution of economic well-being.

Data and Measurement

Pertinent data from the 2001 and 2006 ARMS were used to create the samples of farm operator households.[4] ARMS is a national survey conducted annually by the Economic Research Service and the National Agricultural Statistics Service. Each observation in the ARMS' sample of size $\tilde{n}$ represents a number of similar farms, the particular number being the survey expansion factor [or the inverse of the probability of the surveyed farm being selected for surveying, also known as survey weight (W_j, where $j = 1, \ldots, \tilde{n}$)]. As Table 1 shows, the expanded number of

[3] Using income as an example, the ordinary Lorenz (OL) curve is obtained by plotting the cumulative proportion of the income of the population (y axis) against the cumulative proportion of the population (x axis), ranked by income. The generalized Lorenz curve (GL), due independently to Kakwani (1980) and Shorrocks (1983), is obtained by scaling OL by the mean (μ) of the income distribution, causing each value on the y axis to represent the levels of income, and the highest point on this axis to be μ. The Gini coefficient, which is a statistical construct, is related to the ordinary Lorenz curve in that it is equal to twice the area between the curve itself and a 45-degree line, also referred to as the "egalitarian line," in a graph that has axes y and x. Among two ordinary Lorenz curves, the one curve which is everywhere closer to the 45-degree line than another, its corresponding distribution is judged unambiguously less unequal.

[4] The households in the samples are those of the principal farm operators who make the day-to-day management decisions. What are not included in these samples are those farm households whose farms are organized as non-family corporations or cooperatives, and farms where the majority of the farm ownership is not held by family members.

Table 1. Summary Statistics by Age Categories (2001 and 2006)

Item	Younger than 35	35 to 44	45 to 54	55 to 64	65 or older	All
	Age of operator (years)					
			2001			
Sample size	282	996	1,715	1,289	1,157	5,439
Population	140,852	396,679	582,961	455,333	516,094	2,091,919
Population share (%)	6.7	19.0	27.9	21.8	24.7	100.0
Averages ($)						
Household income	62,384	81,944	78,839	79,036	56,841	72,936
Marketable wealth	4,078	6,722*	11,278	20,477	34,626	17,692
CWB measure	66,462	88,667	90,117	99,513	91,467	90,628
Shares (%)						
Household income	93.9	92.4	87.5	79.4	62.1	80.5
Marketable wealth	6.1	7.6	12.5	20.6	37.9	19.5
CWB measure	100/0	100/0	100/0	100/0	100/0	100/0
CWB share (%)	4.9	18.6	27.7	23.9	24.9	100.0
Average age	29	40	49	59	73	54
Average household size	2.90	3.96	3.06	2.26	2.01	2.73
			2006			
Sample size	295	857	1,829	1,917	1,559	6,457
Population	91,764	235,381	540,984	592,709	561,698	2,022,535
Population share (%)	4.5	11.6	26.7	29.3	27.8	100.0
Averages ($)						
Household income	76,915	82,092	84,901	80,276	61,018	76,224
Marketable wealth	9,613	16,078	20,096	33,142	55,050	32,683
CWB measure	86,528	98,170	104,997	113,418	116,068	108,907
Shares (%)						
Household income	88.9	83.6	80.9	70.8	52.6	70.0
Marketable wealth	11.1	16.4	19.1	29.2	47.4	30.0
CWB measure	100.0	100.0	100.0	100.0	100.0	100.0
CWB share (%)	3.6	10.5	25.8	30.5	29.6	100.0
Average age	30	40	50	59	73	57
Average household size	3.24	3.96	3.15	2.29	1.95	2.67

Source: Authors' calculations and Agricultural and Resource Management Survey (2001 and 2006).

Note: * indicates that the coefficient of variation is greater than 25 and less than or equal to 50. "CWB" is composite measure of economic well-being.

farm households in 2001 and 2006 totaled 2.09 million and 2.02 million, respectively. The table provides evidence that the loss of about 70,000 farm households, in the span of 5 years, may be attributable, to a large extent, to the shrinkage in the proportion of households operated by farmers younger than 55 years old when entry rates into farming by young farmers is on the decline (see Hoppe 2002). A similar trend was found by Gale (2003), who reported a steady decline between

1978 and 1997 in both the number of young farmers entering farming and the number of older farmers exiting farming. The table also shows that about 47 percent of the farmers in 2001 are 55 years old or older, with the proportion of older farmers rising in 2006 to 57 percent, along with the increase in the average age of farmers between the two time periods from 54 to 57, pointing to the continued trend, as was noted by Gale (2002), of the graying of the farm population. In terms of the relative economic position in the life cycle, while Table 1 shows an inverted U-shaped income-age profile, it nevertheless shows an increasing linear trend in marketable wealth across all age sub-groups in both 2001 and 2006.

Economic Well-Being Measure

Using income alone as a measure of well-being is problematic since two individuals with the same income but with different amounts of wealth will have different consumption potential, and different ability to weather unexpected financial shocks (see Burkhauser, Butler, and Wilkinson 1985, Crystal and Shea 1989, Wolff 1995, Johnson and Wahl 2004). This paper adjusts for the limitation of considering only income in examining the inequality in economic well-being of farm operator households by using a combined income-wealth measure—similar to what has been proposed initially by Weisbrod and Hansen (1968) and reaffirmed later by Hill (2002). Accordingly, let the following define this composite measure of economic well-being for each household:

$$(1) \quad CWB_j = HHMI_j + (tc\ MNW_j)\frac{r}{[1-(1+r)^{-l_j}]},$$

where tc represents a proportional adjustment factor reflecting transaction costs, l is the life expectancy of the unit, r is an assumed interest rate set at 4 percent for this paper, and $HHMI_j$ and MNW_j represent the jth total household's money income and marketable net worth, respectively.[5]

Next, an adjusted per-capita equivalent concept of composite measure of well-being (CWB_j^*) is constructed as in the following:

$$(2) \quad CWB_j^* = \frac{CWB_j}{S^{\varepsilon}},$$

where S and ε ($0 \le \varepsilon \le 1$) are, respectively, the number of household members and the elasticity of household "need" (also referred to as the elasticity of the scale rate) with respect to household size (see Burkhauser, Smeeding, and Merz 1996, Daly and Royer 2000).

When ε is 1, this is the per-capita notion of CWB_j^* and it indicates the presence of no economies of scale. This value of ε assumes that a household of two members requires twice as much in combined income and annualized wealth as a household with just one to be as equally well off. When ε is 0, this is the per-household notion of CWB_j^* where economies of scale are assumed perfect, and where no adjustment according to the size of the household is made. This level of elasticity assumes that a household with two, three, or even an infinite number of individuals can live equally well off the income-wealth combination as can a single person household with no increase in *CWB*. When ε is between 0 and 1, this implies a certain level of economies of scale, which grows smaller as ε increases.

In this paper, ε is valued at 0, 0.5, and 1, with these levels being used to provide a sort of sensitivity analysis of the main results to the choice of ε. In addition, setting $\varepsilon = 0.5$ is done here in accordance with other studies, aimed at allowing for comparisons of economic well-being of house-

[5] Total household's money income ($HHMI_j$) is the sum of earnings of the farm household from farming and from off-farm activities. Farm earnings is the sum of the operator household's share of net farm business income (less depreciation) and wages paid to the operator. Off-farm earnings include incomes from wages and salaries, off-farm self employment, interest, dividends, private pensions, Social Security, veterans' benefits, and other public programs. Marketable wealth (*MNW*), a term used by Wolff (1995, p. 59), is defined as the current market value of all fungible assets (farm and non-farm) less the current value of (farm and non-farm) debts. Assets would include owner-occupied housing, bank accounts and certificates of deposit, corporate stocks, and other types of financial assets. In the case of farm assets, they would include those that were marketable in the short run and that would not alter the expected returns from farming. The sale of land as a productive asset would change the production possibilities and potential income to the household from farming and thus was not included as part of marketable farm assets. Marketable farm inventories of crops and livestock would also be included. This measure therefore considers only those assets that are easily converted to cash and purposely exclude farm production assets and household durable goods. An additional consideration in determining the income stream from marketable wealth is transaction-cost adjustment, *tc*. To more accurately portray the annuity, some allowance (or subtraction from *MNW*) must be made for costs incurred in the disposal of assets, set here at 2 percent, which gives factor *tc* a value of 0.98.

holds with different sizes (see Gottschalk 1993, Atkinson, Rainwater, and Smeeding 1995).

Methods

A popular way to measure inequality is to utilize the concept of "size distribution," where the population and the underlying economic variate are typically divided into quintiles or deciles. In this paper, a graphical representation of the size distribution of CWB_j^* is presented for a farm household (i.e., $\varepsilon = 0$) over the 2001 and 2006 time periods based on weighted CWB_j^* quintiles. A limiting aspect of this method of measuring inequality is that it can only provide information about points in the CWB_j^* distribution rather than capture inequality based on the entire distribution. Use of a graphical yet "ordinal" representation of the whole distribution based on the concept of the Lorenz curve ("ordinary" or "generalized"), or a summary statistic based on the concept of the Gini coefficient particularly when measured in conjunction with the Lorenz curve, provides the means to remedy this shortcoming.

Let $x = x_1, x_2, \cdots, x_{\tilde{n}}$ describe the finite distribution of CWB^* where $x_i \geq 0$ and where $x_1 \leq x_2$, ...,$\leq x_{\tilde{n}}$.[6] Let also $u = k/n$ ($k = 0,1,\cdots,\tilde{n}$), and let the share of total CWB^* received by the $k/\tilde{n} \times$ 100 percent poorest households be represented by $L(u)$. Accordingly, the ordinary Lorenz curve of CWB_j^* is described by the following (Kleiber 2005):

$$(3) \qquad L_x(u) = \frac{1}{\mu_x}\int_0^u \varphi_x^{-1}(t)dt, \qquad u \in [0,1],$$

where φ_x^{-1} is the quantile function defined as the inverse of the cumulative distribution function, φ_x, and where μ_x is the sample's expected value of x.

Among two nonintersecting Lorenz curves of distributions A and B, for example, if the one depicting distribution A is everywhere above that of B, the distribution corresponding to the upper curve A is said to Lorenz "strongly" dominate the distribution representing the lower curve B [Bishop, Formby, and Thistle 1992, Deaton 2000 (p. 159)]. The distribution of A is said to Lorenz-dominate the distribution of B if its corresponding Lorenz curve lies everywhere above the corresponding Lorenz curve of B, except for portions of both A's and B's where both of the corresponding Lorenz curves touch. In both situations, distribution A is unambiguously less unequal than distribution B. However, Lorenz domination does not allow for a complete ranking of distributions A and B if their corresponding Lorenz curves do intersect. This limitation of the concept of the Lorenz curve to provide full ranking of distributions is circumvented by a scaling of the distributions by their corresponding means, which results in yet a more useful concept known as the "generalized" Lorenz curve, $GL_X(u)$, as in the following (see Kleiber 2005):

$$(4) \quad GL_x(u) = \mu_x . L_x(u) = \int_0^u \varphi_x^{-1}(t)dt \quad u \in = [0,1].$$

Unlike the concept of the ordinary Lorenz curve, which allows for ranking of distributions only in terms of inequality, $GL_x(u)$, which was originally yet independently introduced by Kakwani (1980) and by Shorrocks (1983), allows for social welfare ranking of distributions. If the Lorenz curves of distributions C and D, with means μ_C and μ_D, are intersecting, and if μ_C is greater than μ_D and is large enough to "lift" its corresponding distribution higher and clear of that of D, then distribution C is said to dominate distribution D according to the generalized Lorenz criterion, but not according to the Lorenz criterion (Deaton 2000, p. 159). Accordingly, distribution C is considered more favorable than distribution D by any equality-preferring social welfare function (SWF), where SWF is required to be increasing and concave (for additional detail, see Maasoumi and Heshmati 2000). However, if this is not the case (i.e., both the generalized Lorenz curves of distributions C and D continue to cross even when μ_C is greater than μ_D, which can occur if μ_C is not large enough), then the social welfare ranking of C and D will ultimately depend on the precise specification of the social welfare function. In other words, social welfare ranking be-

[6] While the concept of the Lorenz curve imposes a non-negativity restriction on the economic variable of interest, many studies (see Amiel, Cowell, and Polovin 1996, Jenkins and Jantti 2005) nevertheless assert the appropriateness to continue with the required monotonicity assumption in the presence of negative values. In this study, while the presence of negative values of CWB_j^* is not considered excessive, it is however considerably more prevalent in 2001 than in 2006 (4.55 percent and 3.25 percent of all the weighted samples, respectively).

tween the two distributions will be determined based on the trade-off between more "equality" in the distribution of *D* and the higher "mean" in *C*.

The third tool used to assess the disparity in economic well-being among farm households is that of the standard Gini coefficient, which has the added benefit of allowing for the "cardinal" measurement and, consequently, for the disaggregation of *SWF* by five age cohorts.[7] To accomplish this while attending to policymakers' joint concern regarding both equity and efficiency as noted by Mukhopadhaya (2001, 2002, and 2003), assume the following "Bergson-Samuelson" *SWF*:

$$(5) \qquad W = W(X, \theta),$$

where, in the context of this paper, X and θ are both functions of CWB^* (for in-depth detail regarding the properties of this function in particular, and for a discussion on distributional and welfare axioms in general, see Inada 1971, Samuelson 1977, Pollak 1979, and Cowell 2007). Specifically, the argument X in (5) is an indicator for total CWB_j^* representing efficiency, while θ denotes a measure of dispersion that represents inequity. Accordingly, this augmented *SWF* must satisfy the following:

$$(6) \qquad \frac{\partial W}{\partial X} > 0 \quad \text{and} \quad \frac{\partial W}{\partial \theta} < 0,$$

which indicates that a rise in efficiency will increase social welfare, while a rise in inequality will decrease social welfare. Since many social welfare functions satisfy the conditions set by (6), Sen (1974) highlighted, based on the assumption that social marginal utility is inversely related to income rank and based on widely accepted axioms, the benefit of using a specialized form of the Bergson-Samuelson class of *SWF*'s, as in

$$(7) \qquad W = \mu_x (1 - G),$$

where G is the Gini coefficient.[8] The rate of substitution between "equity" and "efficiency" at a constant welfare level is given by setting the total differential of equation (7) to zero:

$$(8) \qquad \begin{aligned} dW &= (1-G)d\mu_x + \mu_x d(1-G) \\ &= (1-G)d\mu_x - \mu_x dG = 0, \end{aligned}$$

which, in turn, after collecting terms yields the slope of an indifference curve of *SWF* in a space that spans both inequality and average CWB^*:

$$(9) \qquad \left.\frac{d\mu_x}{dG}\right|_W = \frac{\mu_x}{1-G}.$$

Next is the decomposition of *SWF* by age subgroups, which is done here following Podder (1993) and Mukhopadhaya (2002). The method starts by dividing the ordered vector $x = x_1, x_2, ...,$ $x_{\tilde{n}}$ into five vectors of size n, with each vector populated by the values x_{ik} corresponding to its kth age group while maintaining the initial placement of these values as they were in vector x, with zeros being placed everywhere else. Take the example of the first vector, x^1, which reflects the values of CWB_j^* for those households whose heads are aged 35 years or younger, and where the number of households in this group is denoted by n^1. Vector x^1 now has n^1 values of CWB_j^* and $n - n^1$ zeros. Vectors $x^2 - x^5$ are defined in similar fashion, which, when combined with vector x^1 allows for the recapturing of vector x as in the following:

$$(10) \qquad x = \sum_{k=1}^{5} x^k .$$

[7] This section draws heavily on the works of Johansson (1991, pp. 22–35), Podder (1993), and Mukhopadhaya (2001, 2002, and 2003).

[8] The social welfare function given in equation (7) has a direct relationship with the ordinal generalized Lorenz curve in equation (4) since it is equal to twice the area under the curve itself (see Chatterjee and Podder 2002, p. 4). Note that equation (7) could be rewritten as $1 - G = W/\mu_x$, such that $1 - G$ can be viewed as a measure of equality. In the context of this paper, the implication of equation (7) for a given value of μ_x is that a society that is averse to inequality would prefer a distribution of CWB^* that has a lower Gini coefficient (G) over one that has a higher Gini coefficient. It is important to note, thus, that inequality measures other than the Gini coefficient (e.g., Theil index, coefficient of variation, etc.) may have very different social welfare implications (for more detail, see Xu 2004). As was pointed out by a reviewer, the "cardinal" measure of social welfare captured by equation (7) permits the quantification of improvements in the distribution of CWB^* over the time period considered, but only after the form of the chosen *SWF* as described by W [see equation (6)], in comparison to other forms, is presumed to be acceptable.

Let X^T and X^k denote the total sum of CWB_j^* for the whole population and the total sum of CWB_j^* for the kth age sub-group, respectively. This allows for the following:

$$(11) \qquad X^T = \sum_{k=1}^{5} X^k .$$

Another needed step in the process of decomposing *SWF* by age sub-groups is the construction of the concentration coefficient for each of the corresponding k vectors (k = 1,…, 5), x^k, denoted as C_k. This is in addition to computing the Gini coefficient of vector x, which contains the individual observations of CWB_j^* for the whole population of farm households. In this study, C_k is computed in the same way as in the case of the Gini coefficient (see Pyatt, Chen, and Fei 1980, and Lerman and Yitzhaki 1985):

$$(12) \qquad C_k = 2\,\mathrm{cov}\left[x^k, F(x^k)\right] / \bar{x}^k ,$$

where $F(x^k)$ is the cumulative distribution of x^k (ranked in a non-decreasing order) and $\bar{x}^k$ is the weighted sample mean of x^k, and *cov* is a covariance indicator [for detail regarding the method used to estimate $F(x^k)$, see Lerman and Yitzhaki (1989)].

The Gini coefficient of all elements of CWB^* (see Podder 1993 and Mukhopadhaya 2002), apportioned based on the contribution to inequality by the five age sub-groups, is described by

$$(13) \qquad G = \sum_{k=1}^{5} \frac{X^k}{X^T} C_k .$$

To allow for the weighted decomposition of social welfare among farm households by age cohorts, X^T and X^k in equation (12) need to be presented as follows:

$$(14) \qquad X^T = N\bar{x} \quad \text{and} \quad X^k = N^k \bar{x}^k ,$$

where $\bar{x}$ is the weighted mean of the whole population of farm households,

$$N = \sum_{j=1}^{\tilde{n}} W_j$$

is the weighted total number of all farm households, and N^k is a subset of N reflecting the weighted number of households in the kth age cohort. Based on this new formulation of X^T and X^k, and by collecting terms, equation (12) can be rewritten as

$$(15) \qquad \bar{x}G = \sum_{k=1}^{5} \frac{N^k}{N} \bar{x}^k C_k .$$

Replacing the sample unweighted mean μ_x with the weighted sample mean $\bar{x}$, along with utilizing the expressions in equations (11) and (15), the social welfare function delineated in (7) can now be written as

$$(16) \quad W = \mu_x(1-G) = \bar{x}(1-G) = \sum_{k=1}^{5} \frac{N^k \bar{x}^k}{N}(1-C_k) = \sum_{k=1}^{5} \frac{N^k}{N}[\bar{x}^k(1-C_k)] = \sum_{k=1}^{5} \varphi_k ,$$

where φ_k is considered as the absolute share of the kth age cohort in total social welfare. Accordingly, the relative share of the kth age cohort in total social welfare can be computed as

$$(17) \qquad \varpi_k = \frac{\varphi_k}{W} = \frac{\varphi_k}{\bar{x}\ (1-G)} .$$

From a policy perspective, the effectiveness of a dollar rise in the total CWB^* of the kth age cohort can be assessed based on the following expression, which denotes the marginal effect of such a change in the income-wealth measure on social welfare:

$$(18) \qquad \xi_k = \frac{1-C_k}{1-G} .$$

Specifically, if the value of ξ_k is greater (less) than unity, this indicates that a dollar increase in the total CWB^* of the kth age cohort will result in more (less) social welfare than the effect of the dollar rise being spread over the entire society (see Mukhopadhaya 2002).

Results

The economic unit targeted primarily in the analysis is the farm operator household (i.e., ε =

0, or when economies of scale are assumed infinite), which is similar to what researchers in the U.S. Department of Agriculture use to officially report the economic well-being status of farm families (see Green, Ahearn, and Parker 2008). Figure 1 presents an assessment of the extent in the disparity in the distribution of CWB^* for farm households over the 2001 and 2006 time periods. The figure shows, and in both periods, a large gap in the shares of CWB^* accrued to households in the lowest and highest quintiles, which points to a concentrated income-wealth distribution. For example, while households in the lowest quintiles of the samples held one percent or less of CWB^* depending on the time period considered, households in the highest quintiles held disproportionate shares, ranging from about 53 percent to 55 percent. The distribution in 2001 of CWB^* appears slightly more concentrated than the distribution in 2006. Evidence of this is the 55 percent of CWB^* found held by the upper 20 percent of farm households in 2001 compared to the 53 percent held by their farm household counterparts in 2006.

Figure 2 shows the Lorenz curves depicting the distributions of CWB^* in 2001 and 2006. The fact that the first panel of the figure shows the Lorenz curve of the CWB^* distribution in 2006 falling above its counterpart in 2001 at all ordinates alludes to its "strong" Lorenz dominance, in addition to demonstrating an apparent slight improvement in the distribution of CWB^* over the two time periods.[9] The second panel of Figure 2 shows the plots of the generalized Lorenz curves over the 2001 and 2006 time periods. Because the generalized Lorenz curve corresponding to the distribution of CWB^* in 2006 is measurably above that in 2001 (except in the lower portion of the plot where the ordinates of both curves are nearly identical), the implication of this is that at each percentile (particularly from the 20th onward to the 100th) of the distribution, farm households have more in terms of the combined income-wealth resource than in 2001.[10] This finding indicates that, based on a social welfare function with a preference towards equity, the distribution of CWB^* in 2006, with its statistically higher weighted mean (\$108,907 versus \$90,628), is considered more desirable than the corresponding CWB^* distribution for farm households in 2001. Both panels of Figure 2 show, thus, the distribution of CWB^* in 2006 to be welfare-superior in comparison to its 2001 counterpart, based on both Lorenz and generalized dominance criterion.[11]

The disparity in the combined income-wealth measure of economic well-being on a per-farm household basis is captured in the upper panel of Table 2, where the evidence, based on the concept of the Gini coefficient (G), points to a highly concentrated yet mildly improving distribution of CWB^*. For example, between 2001 and 2006, the Gini coefficient of CWB^* for farm households declined by a value of 0.029, from a high of 0.538 to 0.509, which indicates a decline in inequality, although lacking statistical significance, by 5.4 percent.[12] Over the same time period, the decline in inequality was even much milder when

[9] It should be noted, however, that the difference in the ordinates of the 2001 and 2006 Lorenz curves, as visual inspection of these curves reveals, is not likely to be statistically different from zero. Beach and Richmond (1985) provided statistical methods to allow for the discernment of strong dominance among Lorenz curves. Such methods were not pursued here since they are beyond the scope of this paper.

[10] Careful inspection of the ordinates of these two curves reveals a tiny yet barely observable crossing at some lower CWB^* percentiles (i.e., those with abscissa values of less than 5 percent). This, however, should be interpreted as an anomaly driven by the presence of negative values in the distributions of the composite income-wealth measure in both 2001 and 2006, with the effect of such presence magnified due to the fact that the observations in both samples are weighted.

[11] As one reviewer correctly pointed out, the fact that the distribution of CWB^* in 2006 is found to Lorenz-dominate its corresponding distribution in 2001 makes it unnecessary to consider any further point estimates, including the Gini coefficient, to assess the order of inequality in these distributions. This is because inequality measures satisfying the transfer principle [i.e., when inequality increases (decreases) due to a mean-preserving transfer from a poorer (richer) household to a richer (poorer) household], as was demonstrated by Atkinson's seminal work (Atkinson 1970), will yield the same ranking as one would have achieved based on Lorenz dominance results. That the paper pursues measurement of inequality using the concept of the Gini coefficient beyond the use of the Lorenz dominance results is due to the fact that the main objective of the paper is to conduct welfare decomposition of the joint income-wealth distribution in the context of the life cycle.

[12] Statistical significance of difference among all measured statistics across the 2001 and 2006 time period is measured at the 95 percent confidence level. The estimated 2001 and 2006 Gini values, for example, are considered statistically not different if their corresponding confidence bands overlap with each other. A study by Xu (2000) employed an iterative-bootstrap method in its computation of confidence bands for the generalized Gini indices that were estimated in order to evaluate changes of income inequality in the United States over time. In this study, standard deviations and corresponding confidence intervals for all of the measured dispersion statistics are computed, and because of the complex multistage design of ARMS, based on the jackknife variance estimation method similar to earlier applications by Milanovic (2002) and Giles (2004) [for more detail, see also Sandstrom, Wretman, and Waldman (1988), Dubman (2000)].

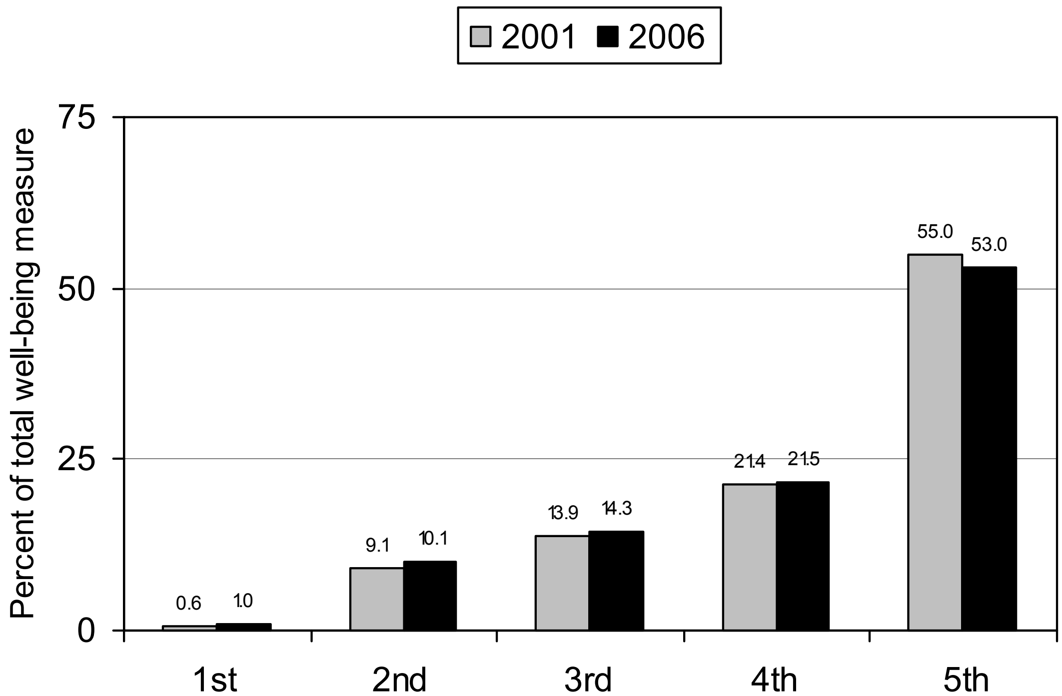

Figure 1. Share of Household's Composite Measure of Economic Well-Being by Quintiles (2001 and 2006)

Source: Authors' calculations and Agricultural Resource Management Survey (2001 and 2006).

the economies of scales were allowed to increase by means of increasing the elasticity of the scale rate ε from a value of 0 to 0.5, and then to 1. For example, when ε = 0.5, the decline in the Gini coefficient between 2001 and 2006 is measured at 2.97 percent, with the decline becoming even less profound, at only 0.709, when ε = 1.

Table 2 also presents the 2001 and 2006 weighted averages of CWB^* for the whole population of farm households and for five age cohorts, with results pointing to statistically larger mean values of CWB^* in 2006 than in 2001 across all age groups. These averages, along with the corresponding population and CWB^* shares and concentration ratios, are statistics needed for the apportionment of social welfare based on the age of the household head. Findings indicate that while the concentration coefficient in both time periods is highest when the age of the household head is between 55 and 64 years, the mean CWB^* is highest for this age group only in 2001 and for those in the 65-years-or-older category in 2006. A nearly identical trend, in both time periods, in terms of higher concentration and higher means of CWB^* for households operated by older farmers, is exhibited at the 0.5 and 1.00 values of the elasticity of the scale rate ε.

Table 3 presents the findings of welfare decomposition based on five age cohorts. Evidence points to a statistically significant increase of 27.6 percent in total social welfare, from $41,886 in 2001 to $53,455 in 2006. The improvement in social welfare over the period appears to be driven mainly by a rise in efficiency in the distribution of CWB^* and not necessarily by a rise in equity [see equation (7)]. The rise in efficiency is evidenced in the larger mean value of CWB^* in 2006 than in 2001, while the muted rise in equity over the period is due to the insignificant decline in the corresponding computed Gini coefficients.

For farm households with the age of the head in the 45- to 54-year-old category in 2001 and for those in the 65-or-older category in 2006, both absolute and relative shares of social welfare are highest compared to when the age of the head falls in the other age groups. An explanation of

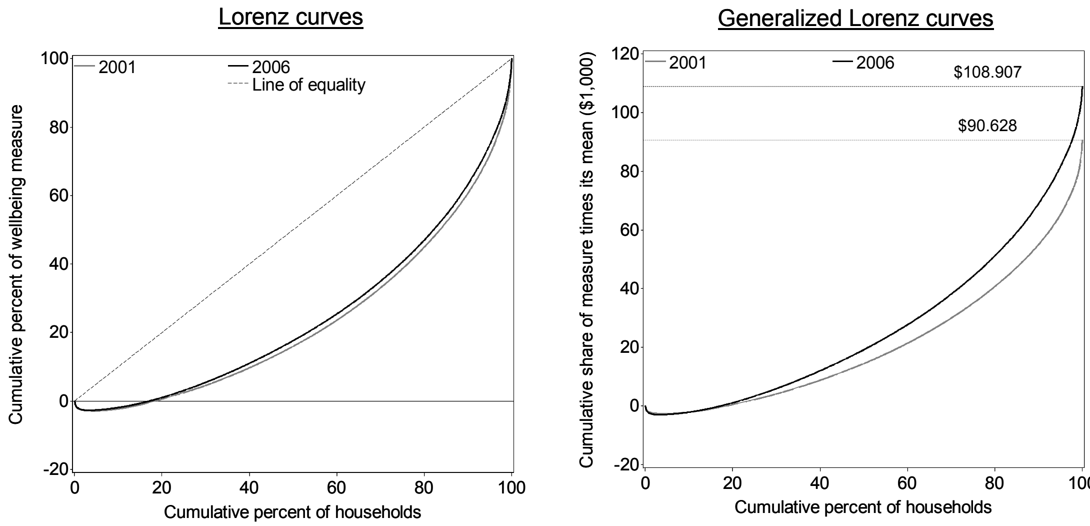

Figure 2. Ordinary and Generalized Lorenz Curves for Composite Measure of Economic Well-Being (2001 and 2006)

Source: Authors' calculations and Agricultural Resource Management Survey (2001 and 2006).

this finding, particularly for 2001, is the economic importance of these two groups of households in relation to the other groups. Specifically, as indicated in Tables 1 and 2, the group of households in the 45- to 54-year-old age category in 2001 accounted for the highest shares (nearly 28 percent) of both the total population of farm households and the total amount of CWB^*. For households in the 65-or-older age category in 2006, the population and the CWB^* shares, while not the highest among all of the age groups (at 27.8 percent and 29.6 percent, respectively), were next to the highest, the highest being for the group of households in the 55- to 64-year-old age category (at 29.3 percent and 30.5 percent, respectively). What caused both the absolute and relative shares of social welfare for households in the 65-or-older category in 2006 to exceed, although by a small margin, their corresponding values for households in the 55- to 64-year-old age category is perhaps the higher mean of CWB^* and the lower value of the concentration coefficient [see equations (16) and (17)]. Results further show an exact trend in terms of the impact of the life cycle of the operator on the absolute and the relative shares of welfare, in 2001 and 2006, when the analysis is based on the presence of only some economies of scale (i.e., $\varepsilon = 0.5$) or based on the absence of any economies of scale (i.e., $\varepsilon = 1.0$).

The effectiveness of a dollar rise in the total amount of CWB^* for any of the five age cohorts is captured in Table 3 by values of ξ_k exceeding unity. Findings indicate that such a rise in CWB^* of three groups of households in 2001 (e.g., younger than 35, 35–44, or 65 or older) and of three groups of households in 2006 (e.g., younger than 35, 45–54, or 65 or older) will result in more social welfare than the effect of such an increase being spread over the whole society. Of these age cohorts, and in both 2001 and 2006, the cohort where the age of the head is younger than 35 years old has the highest value of ξ_k, indicating the relevance of this group of farm households over all others for a policy aimed at improving the economic well-being of its members while concurrently improving the social welfare of the whole population. However, when allowing for the size of the farm household to enter into the analysis (e.g., when $\varepsilon = 1$), the marginal effects of a dollar increase in total CWB^* indicate an increase in social welfare, in both periods, if such an increase occurred for the three age cohorts where the age of the head is younger than 55 years. The largest marginal effects are found for both 2001 and 2006 when the analysis is on a per-

Table 2. Averages and Shares of Composite Measure of Economic Well-Being, Shares of Population, and Concentration Coefficients by Age Categories and by Various Levels of ε (2001 and 2006)

Age Categories	CWB^* ($\bar{x}_k$)[a] 2001		CWB^* ($\bar{x}_k$)[a] 2006		Population shares [N_k/N (%)] 2001		Population shares [N_k/N (%)] 2006	
				ε = 0				
Younger than 35	66,462	(7,865)	86,528	(11,192)	6.73	(1.44)	4.54	(0.59)
35 to 44	88,667	(8,593)	98,170	(7,782)	18.96	(4.42)	11.64	(1.03)
45 to 54	90,117	(9,736)	104,997	(5,174)	27.87	(1.97)	26.75	(1.42)
55 to 64	99,513	(16,626)	113,418	(5,542)	21.77	(3.46)	29.31	(1.58)
65 or older	91,467	(11,408)	116,068	(4,843)	24.67	(2.84)	27.77	(1.14)
Total	90,628 [b]	(7,482)	108.907 [b]	(2,021)	100.00	(0.00)	100.00	(0.00)
				ε = 0.5				
Younger than 35	43,793	(8,126)	52,326	(6,513)	6.73	(1.44)	4.54	(0.59)
35 to 44	47,617	(3,652)	53,942	(4,528)	18.96	(4.42)	11.64	(1.03)
45 to 54	54,559	(6,185)	62,784	(2,878)	27.87	(1.97)	26.75	(1.42)
55 to 64	67,805	(10,604)	77,445	(3,813)	21.77	(3.46)	29.31	(1.58)
65 or older	66,235	(7,852)	86,867	(3,812)	24.67	(2.84)	27.77	(1.14)
Total	58,281 [b]	(5,101)	72,265 [b]	(1,402)	100.00	(0.00)	100.00	(0.00)
				ε = 1				
Younger than 35	31,716	(8,785)	33,971	(4,188)	6.73	(1.44)	4.54	(0.59)
35 to 44	26,855	(1,781)	31,696	(3,223)	18.96	(4.42)	11.64	(1.03)
45 to 54	35,155	(3,932)	39,596	(2,015)	27.87	(1.97)	26.75	(1.42)
55 to 64	47,002	(6,998)	54,417	(2,642)	21.77	(3.46)	29.31	(1.58)
65 or older	49,200	(5,498)	66,996	(3,255)	24.67	(2.84)	27.77	(1.14)
Total	39,393 [b]	(3,576)	50,374 [b]	(1,144)	100.00	(0.00)	100.00	(0.00)
	CWB^* shares [X^k/X^T (%)]				Concentration coefficients (C_k)			
				ε = 0				
Younger than 35	4.94	(0.82)	3.60	(0.67)	0.419	(0.121)	0.421	(0.085)
35 to 44	18.55	(5.23)	10.49	(0.96)	0.516	(0.128)	0.511	(0.055)
45 to 54	27.71	(2.17)	25.79	(1.71)	0.541	(0.022)	0.478	(0.030)
55 to 64	23.90	(3.14)	30.52	(2.12)	0.600	(0.085)	0.552	(0.043)
65 or older	24.90	(2.96)	29.60	(1.22)	0.514	(0.071)	0.503	(0.021)
Total	100.00	(0.00)	100.00	(0.00)	0.538 [c]	(0.021) [c]	0.509	(0.013) [c]
				ε = 0.5				
Younger than 35	5.06	(1.14)	3.29	(0.60)	0.441	(0.163)	0.370	(0.086)
35 to 44	15.49	(4.59)	8.69	(0.83)	0.396	(0.123)	0.404	(0.052)
45 to 54	26.09	(2.21)	23.24	(1.50)	0.508	(0.023)	0.437	(0.030)
55 to 64	25.32	(2.96)	31.41	(2.11)	0.632	(0.078)	0.578	(0.041)
65 or older	28.04	(3.12)	33.38	(1.38)	0.579	(0.067)	0.578	(0.019)
Total	100.00	(0.00)	100.00	(0.00)	0.539 [c]	(0.020) [c]	0.523	(0.011) [c]
				ε = 1				
Younger than 35	5.42	(1.65)	3.06	(0.55)	0.518	(0.197)	0.399	(0.076)
35 to 44	12.93	(3.98)	7.32	(0.82)	0.311	(0.137)	0.373	(0.055)
45 to 54	24.87	(2.25)	21.02	(1.37)	0.514	(0.024)	0.441	(0.037)
55 to 64	25.97	(2.83)	31.66	(2.13)	0.660	(0.068)	0.604	(0.038)
65 or older	30.81	(3.40)	36.94	(1.63)	0.636	(0.061)	0.641	(0.019)
Total	100.00	(0.00)	100.00	(0.00)	0.564 [c]	(0.018) [c]	0.560	(0.010) [c]

[a] At 2006 price levels. Also, standard deviations for all estimates are in parentheses and are computed based on the jackknife variance estimation method.
[b] These are the weighted means ($\bar{x}$) for the whole population.
[c] These are the Gini coefficients of the composite measure of economic well-being for the total population.

Note: "CWB" is composite measure of economic well-being.

Table 3. Decomposition of Welfare for Farm Operator Households by Age Categories and by Various Levels of ε (2001 and 2006)[a]

Age Categories	Absolute share of total welfare [φ_k (2003 dollars)]				Relative share of total welfare (ϖ_k)				Marginal effect (ξ_k)			
	2001		2006		2001		2006		2001		2006	
					ε = 0							
Younger than 35	2,602	(626)	2,274	(404)	0.062	(0.013)	0.043	(0.008)	1.258	(0.288)	1.180	(0.177)
35 to 44	8,141	(3,682)	5,590	(809)	0.194	(0.091)	0.105	(0.015)	1.048	(0.247)	0.997	(0.108)
45 to 54	11,525	(1,120)	14,663	(1,075)	0.275	(0.024)	0.274	(0.017)	0.993	(0.045)	1.064	(0.071)
55 to 64	8,660	(2,138)	14,896	(1,242)	0.207	(0.050)	0.279	(0.019)	0.865	(0.163)	0.913	(0.068)
65 or older	10,958	(3,042)	16,032	(813)	0.262	(0.064)	0.300	(0.014)	1.051	(0.183)	1.013	(0.050)
Total	41,886	(2,148)	53,455	(1,904)	1.000	(0.000)	1.000	(0.000)	1.000	(0.000)	1.000	(0.000)
					ε = 0.5							
Younger than 35	1,650	(377)	1,495	(254)	0.061	(0.012)	0.043	(0.008)	1.213	(0.381)	1.321	(0.179)
35 to 44	5,456	(2,197)	3,741	(451)	0.203	(0.088)	0.109	(0.013)	1.310	(0.234)	1.250	(0.112)
45 to 54	7,475	(986)	9,459	(660)	0.278	(0.031)	0.275	(0.016)	1.066	(0.060)	1.182	(0.070)
55 to 64	5,425	(1,361)	9,572	(837)	0.202	(0.051)	0.278	(0.019)	0.797	(0.153)	0.885	(0.067)
65 or older	6,880	(1,978)	10,172	(525)	0.256	(0.062)	0.295	(0.014)	0.913	(0.166)	0.885	(0.047)
Total	26,886	(1,417)	34,439	(1,200)	1.000	(0.000)	1.000	(0.000)	1.000	(0.000)	1.000	(0.000)
					ε = 1							
Younger than 35	1,029	(224)	927	(153)	0.060	(0.011)	0.042	(0.007)	1.105	(0.477)	1.367	(0.170)
35 to 44	3,507	(1,410)	2,315	(257)	0.204	(0.090)	0.104	(0.012)	1.578	(0.278)	1.426	(0.130)
45 to 54	4,760	(767)	5,916	(421)	0.277	(0.034)	0.267	(0.017)	1.114	(0.075)	1.270	(0.084)
55 to 64	3,475	(889)	6,316	(567)	0.202	(0.051)	0.285	(0.018)	0.779	(0.143)	0.900	(0.070)
65 or older	4,414	(1,309)	6,685	(371)	0.257	(0.062)	0.302	(0.016)	0.834	(0.155)	0.817	(0.050)
Total	17,185	(1,060)	22,158	(824)	1.000	(0.000)	1.000	(0.000)	1.000	(0.000)	1.000	(0.000)

[a] $\varphi_k = \frac{N_k}{N}\bar{x}_k(1-C_k); \varpi = \frac{\varphi_k}{x(1-G)}; \xi_k = \frac{1-C_k}{1-G}$.

Note: Standard deviations for all estimates are in parentheses and are computed based on the jackknife variance estimation method.

capita basis (i.e., when ε = 1), and when the age of the operator is in the 35- to 44-year-old category. This finding is not surprising due to the lower disparity of CWB^* for households in this age group, as indicated by the lower value of C_k [see equation (18)]. In addition, households in this age group tend to have larger family sizes (see Table 2), with their attendant higher economic needs.

Summary and Conclusions

This paper has, first, examined the disparity in economic well-being among farm households over the 2001 and 2006 time periods, and second, has measured and decomposed total social welfare of these households based on five age cohorts. In accomplishing this, both ordinal (i.e., ordinary and generalized curves) as well as "cardinal" (i.e., social welfare functions) methods were used. The underlying financial and demographic information needed for the analysis was obtained from the Agricultural Resource Management Survey (ARMS).

Findings based on frequency density and on Gini coefficient analyses that allow for the partial ranking of the distributions of the combined income-wealth measure (CWB^*) for farm households revealed declining inequality between 2001 and 2006, although the decline was not found to be statistically significant. Use of the concepts of the "Lorenz" and of the "generalized Lorenz" curves showed the 2006 distribution of CWB^* to be more socially preferred than the corresponding distribution in 2001.

Results based on Podder's (1993) method of welfare valuation and decomposition revealed an improvement in social welfare over the 2001 to 2006 time period, which appears to be driven mainly by a rise in efficiency in the distribution of CWB^* and not necessarily by a rise in equity. Findings also indicated the importance of the 45- to 54-age-group in terms of its contribution towards the total social welfare for farm households in 2001, and the importance of the group of farm households in the 65-years-or-older age group in 2006. For both periods, a one-dollar increase in the total combined income-wealth of those farm households with heads in the younger-than-35-years age sub-group, in contrast to those in the other age sub-groups, is found to produce a larger increase in social welfare than the effect of the same dollar increase being spread over the entire society. When the analysis is based on a per-capita concept of economic well-being, rather than on a per-household basis, a marginal one-dollar increase in the total combined income-wealth produced in both 2001 and 2006 the largest increase in social welfare, but only when the age of the operator is in the 35-to-44 age category.

Findings in this paper can be used as a guide to determine which group of households could be targeted in terms of, among other policy options, direct transfers, loan guarantees, redistributive tax schemes, or enhanced business opportunities to produce economic benefit to the households in this targeted group itself while concurrently benefiting society at large. The group of farm households where the age of the head is younger than 35 years is the appropriate target when the analysis is done on a per-farm household basis, and the group of households where the age of the operator is in the 35-to-44 age category is the appropriate target when the analysis is carried on a per-capita basis.

For farm households, there are policies in place now that provide assistance to those on the extreme of the age cohort. For example, the Farm Service Agency (FSA) provides direct and guaranteed loans to beginning farmers and ranchers who tend to be young and frequently tend to be unable to obtain financing from commercial credit sources. Among the provisions of the 2008 farm bill (i.e., the Food, Conservation, and Energy Act) is an increase in the inflation-indexed loan limit for an individual beginning farmer from $250,000 to $450,000, and competitive grants for education, extension, and outreach initiatives. Findings of this paper suggest that future efforts to address policies providing assistance to those farmers in the 35-to-44 age group, who tend to have already-formed and growing families, with their attendant higher livelihood costs, may be the most efficient way to maintain a viable farm economy. Given the high relative social welfare impacts found for the younger and the middle-age farm age cohorts, policies that directly address farm-entry and farm-preservation issues would likely have the largest payoff.

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