Assessing the Impact of Health Insurance and Other Socioeconomic Factors on Inequality in Health Care Expenditures among Farm Households

Hisham S. El-Osta

This research uses data from the 2005–2011 Agricultural Resource Management Survey and a two-part-model regression procedure to examine the impact of health insurance and other relevant socioeconomic factors on the distribution of health care expenditures among U.S. farm households. Findings show the importance of privately acquired health insurance coverage in explaining inequality in health care expenditures. The results also reveal, among other things, a statistical positive association between health care expenditures and farm operators who fall into the baby boomer age category. A similar statistical association is found for higher income levels but not for inequality of income.

***Key Words*:** farm households, government payments, health insurance, inequality decomposition, source of insurance

Annual health care expenditures in the United States in 2011 were nearly $2.7 trillion (see Figure 1), or close to 18 percent of the nation's gross domestic product (GDP) (Centers for Medicare and Medical Services 2012), a significant increase over the $27.4 billion spent in 1960 (5 percent of GDP) and greater than any other industrialized country (Organization for Economic Cooperation and Development (OECD) 2012). On a per capita basis, the 2011 U.S. expenditures average out to about $8,606 per year, up from $147 in 1960. Newhouse (1992) attributed the significant increases to growing demand for health care services and the rising cost of those services caused by accelerated development and use of new medical technologies. Giles (2003) attributed it in large part to structural factors such as insurance contracts and particularly emphasized the share of expenses covered by third parties (e.g., insurance companies, employers, and Medicare). This dramatic increase in both raw spending and share of GDP is a major concern to policymakers because of the likely adverse impact on employment, inflation, and per capita GDP (Office of the Assistant Secretary for Planning and Evaluation 2008).[1]

[1] Despite the increases in national health care expenditures relative to GDP, the share of total health expenditures that are out-of-pocket for U.S. citizens has been steadily declining from a high of nearly 48 percent in 1960 to a low of nearly 12 percent in 2010 (Centers for Medicare and Medical Services 2012). The difference between the share of national health care expenditures relative to GDP and the share paid by citizens is picked up either by the federal government or through a tax subsidy to employers.

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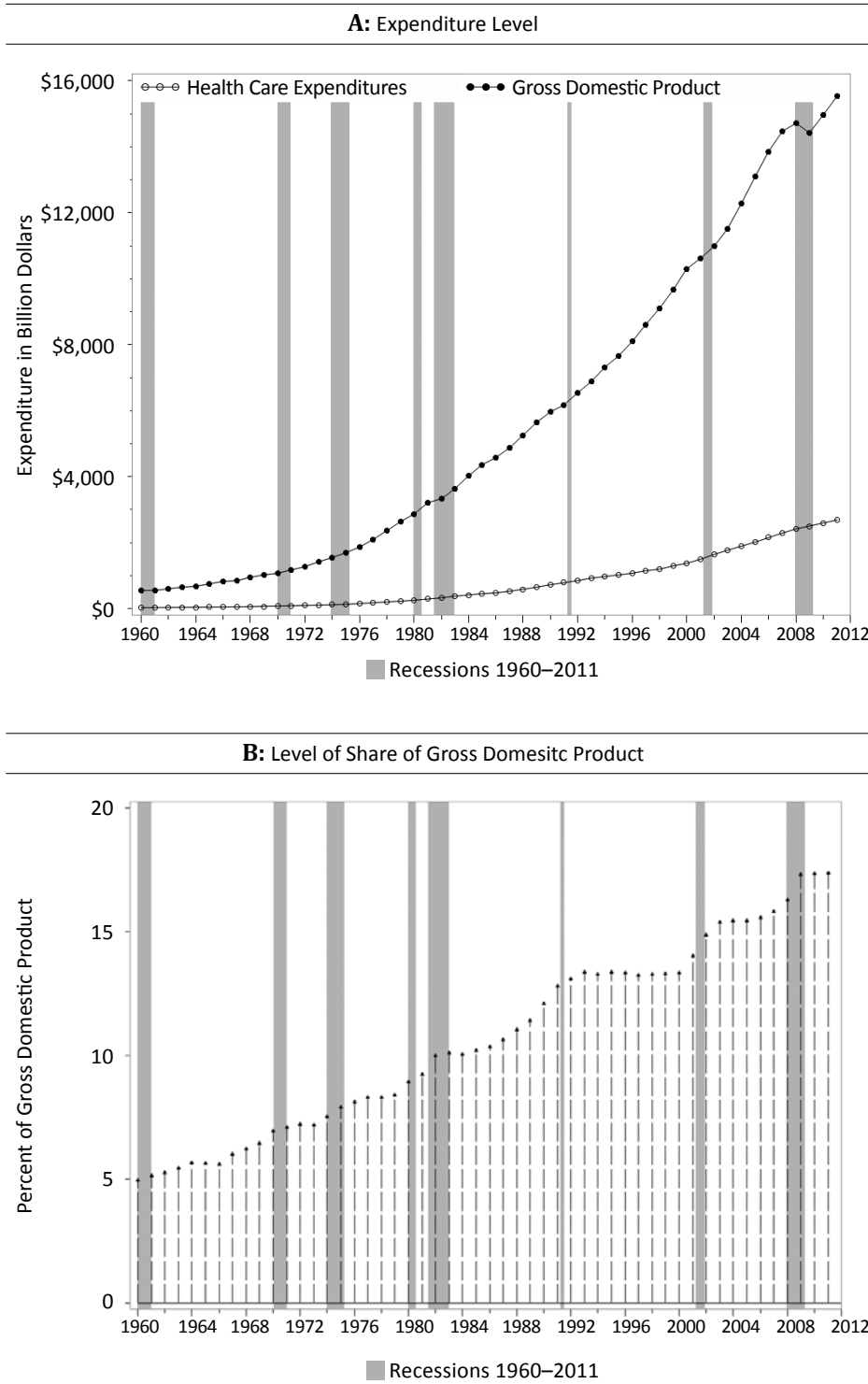


Figure 1. Health Care Expenditures and Gross Domestic Product in the United States, 1960–2011

Data source: Center for Medicare and Medical Services; Office of the Actuary, National Health Statistics Group; U.S. Department of Commerce, Bureau of Economic Analysis; and U.S. Bureau of the Census.

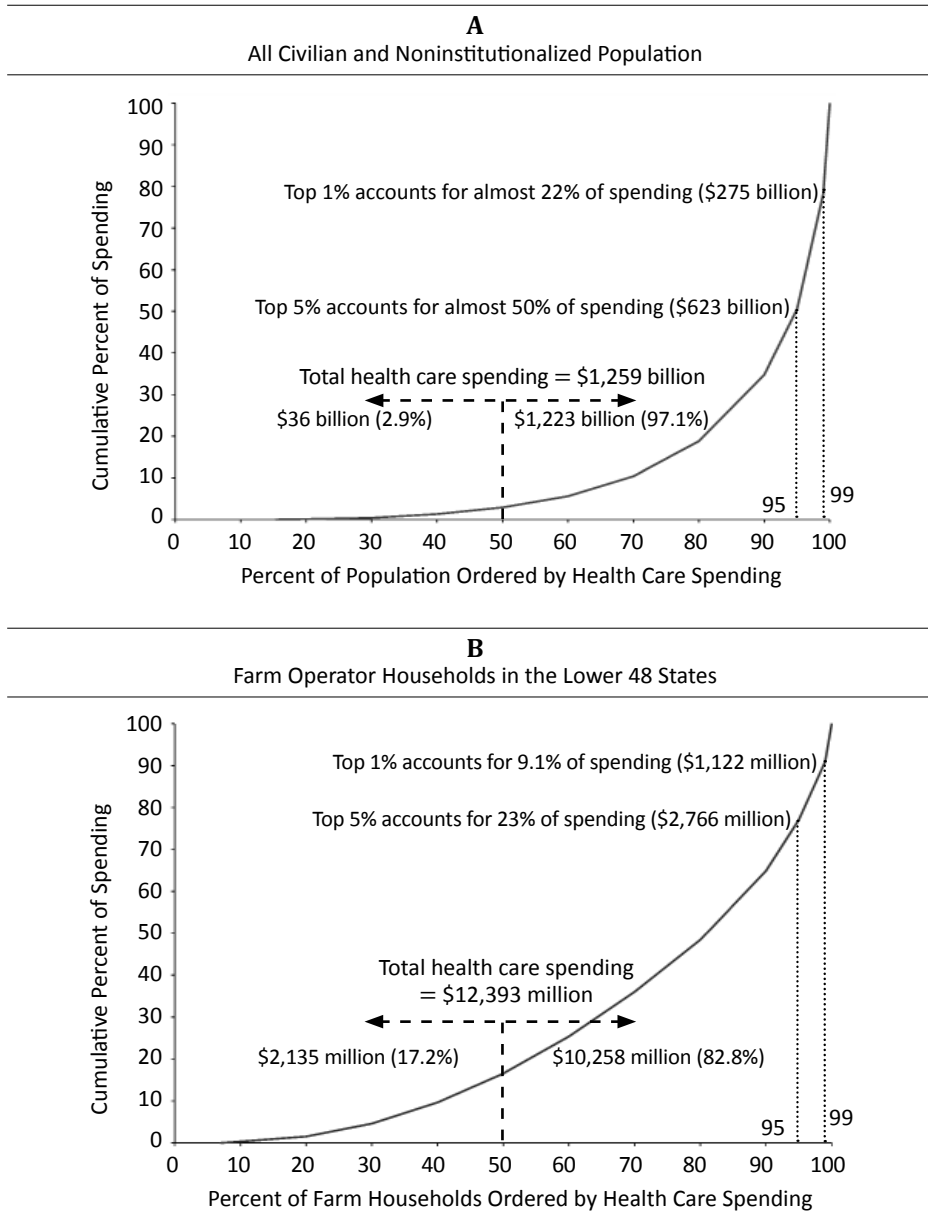


Figure 2. Cumulative Distribution of Health Care Spending for U.S. and Farm Persons in 2009

A: Adapted by the author from Schoenman (2012). B: Computed by the author based on data from the 2009 ARMS.

Another characteristic of health care expenditures in the United States is wide variation in its distribution. According to Stanton and Rutherford (2005), nearly half of the population (civilian noninstitutionalized) in 2002 spent little or nothing on health care while the top 5 percent in terms of individual expenditures accounted for nearly half of the total amount. This concentration of spending continued throughout the decade. As shown in Figure 2, the top 5 percent in 2009 were once again responsible for nearly half of all spending.

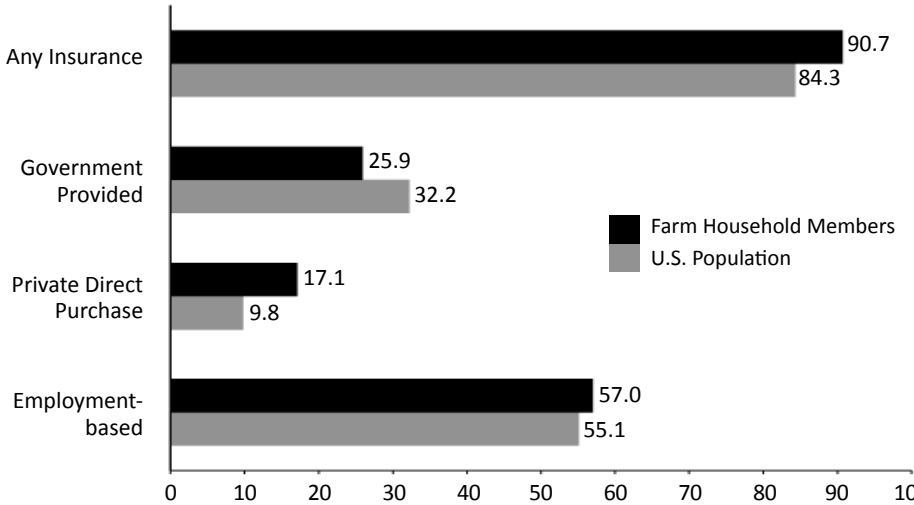


Figure 3. Source of Health Insurance Coverage for U.S. and Farm Persons in 2011

Source: Adapted by the author from Weber and Ahearn (2013).

Concentration in U.S. health care spending is further demonstrated by the 15 percent of the population with no spending and the lower half of the health care distribution, which accounts for only 3 percent of total spending.

Because they are self-employed and at a relatively high risk of injury and illness due to the nature of their work, farmers are among the most disadvantaged groups in the U.S. economy in terms of health risks (Mishra, El-Osta, and Ahearn 2012).[2] Consequently, their decisions about insurance and spending on health care are of particular interest. Figure 2 demonstrates that there is a concentration of spending on health care among farmers (excluding nonfamily corporations and cooperatives) as well, though it is less pronounced. Households representing the top 5 percent of annual health care expenses captured only a little more than one-fifth of all spending. On the low end of the spending distribution, nearly 7 percent of households had no expenditures and 50 percent of the lowest-spending farm household members represented 17 percent of total spending.

As shown in Figure 3, farmers rely more often than the population in general on individual insurance policies purchased directly from insurance providers (Mishra, El-Osta, and Ahearn 2012)—17.1 percent compared to 9.8 percent for the U.S. population as a whole. Farmers are, however, less likely than the general population to lack health insurance altogether (9.3 percent versus 15.7 percent). In addition, a smaller share of farm household members receives

[2] The rate of occupational fatalities was 3.6 per 100,000 for workers in the United States in general in 2008 and 39.5 per 100,000 for individuals whose major occupation was farming or ranching, making the fatality rate among farmers more than ten times higher than their nonfarmer counterparts (Bureau of Labor Statistics 2009). The study by Mishra, El-Osta, and Ahearn (2012) showed that 77 percent of nonelderly farm families in 2006 purchased their insurance directly from a private insurance source; 54 percent of nonelderly families in the general U.S. population bought privately provided insurance. In terms of total out-of-pocket health care expenditures, nonelderly farm families annually spent an average of $5,107 in 2006 while nonelderly families in the United States in general spent $2,396 annually.

health insurance provided by the government through Medicare or Medicaid and about the same share receives coverage through off-farm employers of the operator and/or a spouse.[3]

The primary objective of this study is to examine the inequality in the distribution of health care expenditures (both for private insurance purchases and out-of-pocket medical expenses) noted for nonelderly farm operators and determine how much of that inequality is explained by privately acquired health insurance. In addition, the study explores the impacts of income positions of farm households, the extent of income inequality among households, socioeconomic factors (such as age, educational attainment, ethnicity, and family structure), and other factors such as farm subsidies on inequality in health care spending.

A direct link between limited education and low earnings has been documented by many researchers (e.g., Blank 1997, Deavers and Hoppe 1992, Parker and Gibbs 2005, Schiller 2004). It stems from reduced incentives to enter the labor market and limited opportunities for higher earnings and stable employment without education. Consideration of education as a contributor to inequality in health care expenditures is in accordance with a 2002 United Nations report noting that households characterized by limited education are vulnerable to ill health and disability among household members, price and credit swings, and natural disasters. Furthermore, the notion of this positive association between health outcomes and schooling and its adverse impacts on health care expenditures has been examined and invariably confirmed in a number of studies (e.g., Clark and Royer 2010, Conti, Heckman, and Urzua 2010, Altindag, Cannonier, and Mocan 2011).

There is interest in whether farm subsidies affect farm households' expenditures on health care because the rationale for introducing them in federal farm policies in the 1930s was to alleviate poverty among farmers and provide farm households with a safety net. However, to the extent that eligibility for these subsidies is determined by a limited number of "program crops" and not by poverty standards, the payments end up supporting primarily large profitable farm operations rather than cash-strapped small-scale farmers (Riedl 2004). Unevenness in the distribution of farm payments was examined in Hoppe (2007), which showed, for example, that while less than half of all farms received farm program payments in 2005, the share of payments received by large family farms, which accounted for 8 percent of all farms, was 58 percent. That the design of farm policies has not focused on improving the welfare of the most financially vulnerable farm households is evident from the fact that two-thirds of the recipient farms in 2005 received less than $10,000 in payments, an amount that accounts for only 7 percent of their cash farm incomes (Hoppe 2007). The skewed distribution of federal farm subsidies provides impetus to determine empirically whether farm program payments have any impact on the level and distribution of health care expenditures by nonelderly farm households.

One important federal policy aimed at diminishing health care disparities and controlling the rising cost of health care in the United States is the Patient Protection and Affordable Care Act (ACA), which was signed into law on

[3] A study by Baily (2009) noted considerable evidence that rural household members who have employer-provided health insurance obtain a greater quantity of health care services at a lower cost than those who have privately purchased health insurance.

March 23, 2010. Through its health-insurance requirement, the act is expected to affect farmers' on-farm and off-farm labor allocation decisions (Ahearn, El-Osta, and Mishra 2013) and, consequently, the socioeconomic well-being of farm households. Of all of the factors considered in this study that are likely to impact both the level and the distribution of health care expenditures—and perhaps because of the burden of higher premiums and less comprehensive coverage when it must be purchased directly from insurance providers (Sundaram-Stukel and Deller 2009), something the ACA aims to mitigate—privately acquired health insurance coverage is the most vital.

Background

A Congressional Budget Office (2008) study pointed to wide variation in health care costs across the United States, where per capita annual health care spending in 2004 ranged from roughly $4,000 in Utah to about $7,000 in Massachusetts. Over the same period, per capita personal income was $27,000 in Utah (in the bottom five) and $42,000 in Massachusetts (among the top five) according to the Bureau of Economic Analysis (2005).[4] The positive link between high health care costs and high incomes among states is not particularly surprising since a similar link has been found for a wide variety of countries (Fosler 2012), but this type of correlation seems less likely when the underlying unit of health care expenditure is the individual. A study by Wagstaff, van Doorslaer, and Paci (1991) in fact found that health care expenditures by the poorest individuals in a country are greater than those of the wealthiest individuals. Yet another study (Marmot 2002) pointed to a positive correlation in rich countries between the income poverty of households and increased morbidity and premature mortality.[5] Those results and similar findings in the literature (Ettner 1996, Kawachi, Kennedy, and Wilkinson 1999) that were based on empirical work have galvanized the premise among researchers (known as the income inequality hypothesis) that low incomes in general and income inequality in particular have adverse impacts on health outcomes.

In light of escalating health care costs in the midst of budgetary pressures and declines in median real incomes in the preceding decade (DeNavas-Walt, Proctor, and Smith 2011), a major concern for U.S. policymakers is whether a relationship similar to the one between income and health outcomes holds for health care expenditures.[6] Examining the effects of the income position of the farm household and of income inequality on health care expenditures is important since, as shown in Figure 4, working-age households in states with

[4] The five states with the highest per capita income in 2004 (from highest to lowest) are Connecticut, Massachusetts, New Jersey, Maryland, and New York. The five states with the lowest per capita income in 2004 (from lowest to highest) are Mississippi, Arkansas, West Virginia, New Mexico, and Utah. However, Utah's placement among the states with the lowest per capita income could be associated with its ranking number one in terms of average household size at 3.01.

[5] As noted by a reviewer, the increased morbidity and premature mortality rates may be due to a positive correlation between income poverty and unhealthy behaviors (e.g., obesity resulting from consumption of relatively less healthy food, lack of exercise, smoking, and/or drinking), which has been found in some studies in the health-related literature (Meara, Seth, and Cutler 2008), or to difficulty accessing health care and/or a lack of health insurance.

[6] In 2007, federal, state, and local government personal transfer payments amounted to $1.71 trillion with medical benefits accounting for the single largest transfer payment category in both metropolitan and nonmetropolitan areas. Over nearly three decades (1979–2007), nonmetropolitan transfer payments for medical benefits increased 480 percent while metropolitan payments increase 412 percent (Parker 2009).

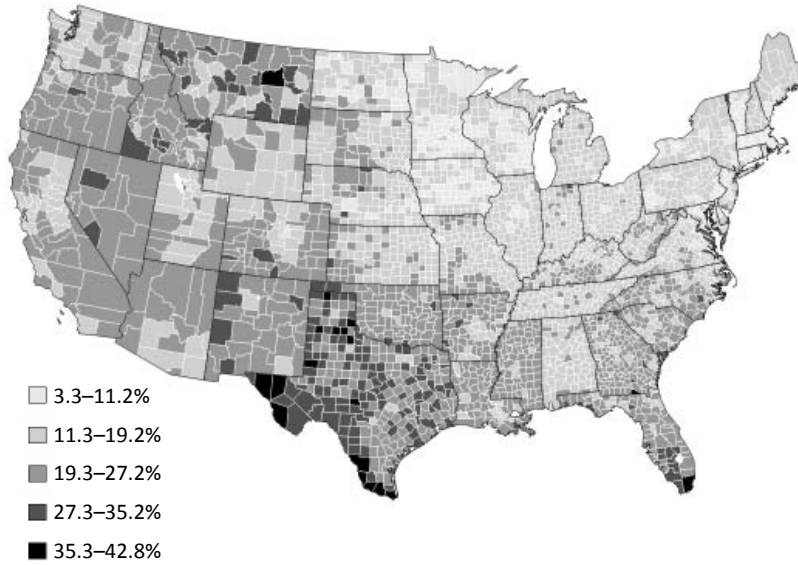


Figure 4. Percent of the Uninsured by County for 2010

Note: All races, under age 65. Data source: U.S. Census Bureau small area health insurance estimates.

relatively low per capita incomes (the south and west) tend to have little or no insurance.

Economic Framework

The conceptual model of health care used here is a close adaptation of the theoretical constructs of household behavior developed by Thomas (1990) and employed by Mwabu (2007). Specifically, for an individual farm-operator household with M members in any time period, the aim is to maximize welfare, W, which is a function of M individual utility functions, $U_1, U_2, \ldots, U_M$:

$$\text{(1)} \qquad \text{Max } W = W[\tau_1 U_1(\mathbf{Q}, \mathbf{K}, \mathbf{A}), \ldots, \tau_M U_M(\mathbf{Q}, \mathbf{K}, \mathbf{A})] = \prod_{m=1}^{M} \boldsymbol{\tau}_M U_M(\mathbf{Q}, \mathbf{K}, \mathbf{A}).$$

The maximization function is subject to a budget constraint,

$$\text{(2)} \qquad Y = \mathbf{pQ} = NFI + \sum_{m=1}^{M} w_m \mathbf{T}_m + \sum_{m=1}^{M} y_m,$$

and a production function for each particular element of $\mathbf{K}$ that includes health (H),

$$\text{(3)} \qquad H = H(\mathbf{Q}),$$

in which Y is a farm household's annual total income, $\mathbf{Q}$ is a vector of goods consumed by all M members of the household (e.g., health care (HC) and other goods, including leisure), $\mathbf{p}$ is a vector of corresponding commodity prices, NFI is the household's net farm income, $\mathbf{T}$ is a vector of work hours in off-farm wage/salary jobs and off-farm businesses, w_m is the price of off-farm work time, and y_m is unearned income (e.g., interest and dividend income, private pension

and disability payments, and social security) from all other off-farm sources for family member m, $\mathbf{K}$ is a vector of home-produced nonmarket goods (e.g., health (H), housework, and unpaid farm labor), $\mathbf{A}$ is a vector of demographic characteristics and socioeconomic factors, $U_m(\cdot)$ is a utility function, and τ_m is an indicator of the bargaining power of the mth individual, which is also known in the literature as a Pareto weight.

The first-order conditions for the preceding model provide many useful results, including the optimality conditions for the demand equations of the qth element of Q and the equation for health care, HC:

$$Q_q^* = HC^* = f(Y, H, \mathbf{A}, \tau_1, \ldots, \tau_M). \tag{4}$$

As depicted in this model, the welfare of households facing budgetary and labor-time constraints is assumed to be tied to the household members' health and consumption of other commodities. Health, as indicated by equation 3, is produced by combining various inputs of production such as medical care.[7] Accordingly, and as in Parker and Wong (1997), health care consumption as measured by total health care expenditures is, by construct, a derived demand for health. The household model is simplified by valuing the Pareto weight of the farm operator (τ_1) at 1 while valuing the Pareto weights for the rest of the members of the household at 0. Thus, to the extent that $\sum_{m=1}\tau_m = 1$, the consumption allocation decision within the household is determined by one individual and formulation of the model is consistent with Becker's (1965) unitary household model with egotistical preferences while allowing for income pooling by members of the household (Chiappori 1988). A generalization of this model, one that will allow for all members of the household to have the same preferences in terms of consumption allocation decisions, results when all of the weights for all household members are set equal to each other.[8]

Data Sources, Empirical Estimation, and Inequality Decomposition

Data

Pertinent data from the U.S. Department of Agriculture's (USDA's) Agricultural Resource Management Survey (ARMS) for 2005 through 2011 were pooled for the purpose of measuring factors that contribute to inequality in health care expenditures among U.S. farm households. The ARMS, which has a complex,

[7] As a reviewer noted, and as allowed for by equation 3, individuals' health care expenditures are likely to be correlated with their current health status, which suggests that healthy individuals will tend to spend little or no money on doctors' visits for treatment of illness-related conditions. This correlation, in and of itself, raises the question of why inequality of health care expenditures matters. The author is grateful for an added reason provided by the reviewer—inequality in health care expenditures among farm households may reflect inequality in access to any sort of preventative care or health care, perhaps because of few options for services and insurance in rural areas. It is common knowledge among health care practitioners that, along with comprehensive, affordable health insurance coverage as a prerequisite to access to high-quality providers, rural areas need access to a diverse group of providers for their communities' health care needs (Seshamani, Lambrew, and Antos 2008). For example, compared to urban areas where there was an average of 72 primary care physicians per 100,000 residents in 2005, the ratio was 55 per 100,000 in rural areas in general and 36 per 100,000 in isolated rural areas (Fordyce et al. 2007).

[8] Note that this model, which is specified in a purely economic sense due to difficulties associated with valuing leisure and other consumption components that are not based on market transactions, captures expenditures rather than consumption allocation decisions.

stratified, multi-frame design, is a national survey conducted annually by the Economic Research Service and the National Agricultural Statistics Service (for more detail, see USDA (2012)). The target population of the survey is operators of farm businesses representing agricultural production in the 48 contiguous states. In it, a farm is defined as an establishment that sold or normally would have sold at least $1,000 of agricultural products during the year. Each observation in the ARMS represents a number of similar farms; the particular number for each observation is the survey expansion factor (the inverse of the probability of a surveyed farm being selected for surveying, hereafter referred to as the survey weight), W_i ($i = 1, \ldots, n$ where n denotes sample size). For example, the initial samples considered in this analysis consist of 6,828 observations in 2005 and 9,488 observations in 2011. When properly expanded using survey weights, these samples yield a population of 2,021,236 and 2,119,693 farm operator households, respectively.[9]

The health care expenditures (*HC*) are measured as the sum of health and/or dental insurance costs (*HI*) and all of the out-of-pocket expenses for health and medical needs (*HM*). Since all individuals age 65 or older have the option of receiving Medicare insurance coverage and virtually all enroll in that program (Mishra, El-Osta, and Ahearn 2012), the seven-year pooled ARMS data are restricted to households in which both the operators and their spouses (some of whom may also be operators) are younger than 65. After restricting the pooled cross-sections of data based on the age of the operators and their spouses and after trimming (as in Sastre and Ayala (2002)) the upper 1 percent of the weighted observations in the distributions of *HI* and *HM* based on the presence of excessively high and outlying health care expenditures, the sample sizes were reduced to 5,156 for 2005 and 6,514 for 2011, and when the samples were expanded using the survey weights, the corresponding targeted populations of farm households were 1,405,508 in 2005 and 1,379,027 in 2011.[10]

Empirical Estimation

As evident from Figure 5, the distribution of *HC*, notwithstanding the presence of nearly 8.5 percent of the population that had no health care expenditures, resembles a log-normal distribution with a long and heavy right tail. The mode of the skewed distribution of *HC* is at zero, indicating that some farm households had no health care expenditures; as a result, our analysis of the distribution of expenditures required a two-part model rather than the traditional linear regression-based one-part model. The two-part model (2PM) used by health economists (e.g., Duan et al. 1983, Hay and Olsen 1984, Manning, Duan, and Rogers 1987, Mullahy 1998, Blough, Madden, and Hornbrook 1999) and by agricultural economists (e.g., Zheng and Zimmer 2008, Mishra, El-Osta, and Ahearn 2012) models the probability (Pr) of a nonzero cost for the *i*th

[9] These are the sample sizes that remained after excluding observations in which the household's farm was organized as a non-family corporation or cooperative or when none of the net income generated by the farm business was received by the household. Thus, the observations are of family farms organized as self-proprietorships, partnerships, and family C- or S-corporations.

[10] As is typical in survey data, some respondents opted to refuse to provide information for some of the components of household expenditures, including *HI* and *HM*. Rather than deleting these observations, we first converted the refused items to missing and then imputed values based on information for household expenditures from cohorts of similar ages and incomes that were used to replace the missing data.

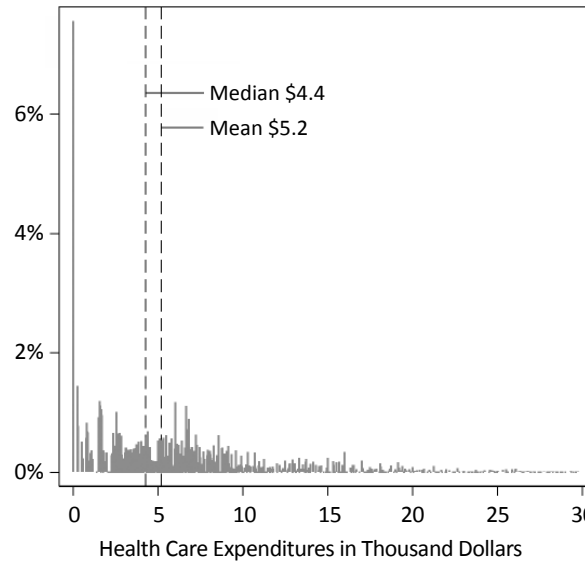


Figure 5. Histogram of Weighted Health-related Expenditures for 2005 through 2011

($i = 1, \ldots, n$) farm household separate from the level conditional on nonzero costs. The first part of such models assumes that $P(HC > 0 \mid x)$ is governed by a parametric binary model with an indicator variable, I_i, that consequently is estimated using a probit regression:

$$I_i^* = \gamma' x_i + \varepsilon \text{ and}$$

$$I_i = \begin{cases} 1, I_i^* > 0 \\ 0, I_i^* \leq 0 \end{cases} . \tag{5}$$

In the second part of the model, $E[HC \mid HC > 0, Z] = Z\hat{\alpha}$ is a linear function of Z; as such, it is estimated using ordinary least squares (OLS) or, in this case, using weighted least squares (WLS). Specifically, this part of the model, which is the econometric representation of the health care demand model previously described for the nonelderly farm households with positive health care spending, is depicted by the following dummy variable linear regression model:

$$HC = \beta_0 + \beta_1 G_Y^* + \sum_{j=2}^{k} \beta_j X_j + \sum_{l=k+1}^{J} \beta_l D_l + \varepsilon = \sum_{j=1}^{J+1} \alpha_j Z_j \tag{6}$$

where HC is total health care expenditures; G_Y^* is an income inequality indicator; X_j and D_l are a continuous and a dummy variable representing characteristics of the operator, the household, and the farm; β is a parameter to be estimated, ε is an error term; $\alpha_j = [\beta_0, \beta_1, \beta_2, \ldots, \beta_J, 1]$; and $Z_j = [1, G_1, X_2, \ldots, D_J, \varepsilon]$. Some of the dummy variables in equation 6 represent regional and year fixed effects that are used to control for region- and time-based omitted variables in the

estimation of model parameters.[11] The explained variation of the dependent variable *HC* is described by the goodness of fit measure, R^2.

Estimation of equations 5 and 6 allows for estimation of $E[HC \mid Z]$:

$$\begin{aligned}(7)\ \mathrm{E}[HC] &= [\Pr(HC > 0 \mid x)\ \mathrm{E}(HC \mid HC > 0, Z)] + \Pr(HC = 0)\ \mathrm{E}[HC \mid HC = 0] \\ &= [\Pr(HC > 0 \mid x)\ \mathrm{E}(HC \mid HC > 0, Z)] + 0 \\ &= \Phi(\hat{\gamma}'x)\hat{\alpha}'Z\end{aligned}$$

where $\Phi(\cdot)$ is the standard cumulative distribution function. The marginal effect (*ME*) for a continuous explanatory variable q that appears in both parts of the two-part model is measured, as in Dow and Norton (2003), by

$$(8) \qquad ME_q = \partial \mathrm{E}[HC] \,/\, \partial q = \hat{\alpha}_q \Phi(\hat{\gamma}'x) + \hat{\gamma}_q \phi(\hat{\gamma}'x)(\hat{\alpha}'Z).$$

The corresponding marginal effect for discrete variable Q is computed as the difference between the results of equation 7 for $Q = 1$ and $Q = 0$.

Inequality Decomposition

Following Fields (2003), the inequality in health care expenditures and the components that contribute to such inequality are assessed using a variance measure of dispersion, σ^2_{HC} (see Burt and Finely (1968) for an early derivation of this method in the agricultural economics literature):[12]

$$(9) \qquad \sigma^2_{HC} = \mathrm{cov}\left[\sum_{j=1}^{J+1} \alpha_j Z_j, HC\right] = \sum_{j=1}^{J+1} \mathrm{cov}(\alpha_j Z_j, HC).$$

The contribution of the *j*th factor toward inequality, as derived by Fields (2003), is measured by a relative factor inequality weight:

$$(10) \qquad s_j(HC) = \mathrm{cov}(\alpha_j Z_j, HC) \,/\, \sigma(HC)$$

where

$$(11a) \qquad \sum_{j=1}^{J+1} s_j(HC) = 100\%$$

[11] Since the ARMS data set is not longitudinal and each of the 2005–2011 repeated cross-sections was sampled independently, we recognize (as in Liao and Taylor (2010)) that these cross-sections may not fully eliminate averages of individual invariant regional and time effects by simply expressing all of the variables in terms of deviations from their within-sample means (i.e., deviations calculated for each individual over regions and time). Consequently, some of the potential correlation between the explanatory variables and the omitted unobservable regional and time effects is not swept from equation 6. Despite this, and because of the large number of observations in the selected sample (34,864 over the seven-year period) and use of the White/Huber method of variance estimation, the specified model in equation 6 may allow for heteroskedasticity-consistent standard errors of estimated model parameters.

[12] Expansion of equation 9 when $\alpha_j = [\beta_0, \beta_1, \ldots, \beta_j, 1]$ shows that

$$\begin{aligned}\sigma^2_{HC} &= \sigma^2(HC \mid \alpha_0, \alpha_1, \ldots, \alpha_j) \\ &= \beta_1^2\sigma_{11} + \beta_1\beta_2\,\sigma_{12} + \ldots + \beta_1\beta_j\,\sigma_{1J} + \beta_2\beta_1\sigma_{21} + \beta_2^2\sigma_{22} + \ldots + \beta_2\beta_j\,\sigma_{2J} + \\ &\quad \ldots\ldots \\ &\quad \beta_j\beta_1\sigma_{j1} + \beta_j\beta_2\sigma_{j2} + \ldots + \beta_j^2\sigma_{JJ} + \sigma^2_{\varepsilon}.\end{aligned}$$

and where, after excluding the last element of Z,

$$\sum_{j=1}^{J+1} s_j(HC) = \mathrm{R}^2(HC). \tag{11b}$$

Fields (2003) further demonstrated that the portion of the variance measure of dispersion explained by the jth factor, $p_j(HC)$, is

$$p_j(HC) = s_j(HC) \;/\; \mathrm{R}^2(HC). \tag{12}$$

Each regression in the 2PM contains a potentially endogenous explanatory variable. Suspicions regarding endogeneity arise because unobserved factors captured in the error term of each model are likely to be correlated with their observed determinants. The first such explanatory variable in the first part of the model of whether a farm household has made any health care expenditures ($\Pr(HC > 0 \mid x)$) is a dummy variable, d_i, which indicates the status of health insurance coverage ($d_i = 1$ when the household has no health insurance coverage; $d_i = 0$ otherwise). To test for the possibility that d_i is endogenous in a binary regression model and, if so, to account for it, we use the two-step instrumental variable (IV) estimation procedure proposed by Vella (1993). The first step estimates the lack-of-insurance-participation decision using a probit regression model:

$$d_i^* = \eta' z_i + u_i \text{ and}$$
$$d_{il} = \left\{ \begin{matrix} 1, d_i^* > 0 \\ 0, d_i^* \leq 0 \end{matrix} \right\} . \tag{13}$$

Estimation of this probit model, which includes a set of k exclusion-restriction variables ($\dot{x}_{ik}$) needed for model identification, results in the $\hat{\eta}$ estimated parameters, the standard cumulative distribution function $\Phi(\cdot)$, the probability density function of the standard normal, $\varphi(\cdot)$, and the generalized residuals (Gourieroux et al. 1987):

$$\hat{v}_i = (d_i - \Phi(\hat{\eta}' x_i))\varphi(\hat{\eta}' x_i)((1 - \Phi(\hat{\alpha}' x_i))^{-1}\Phi(\hat{\alpha}' x_i)^{-1}. \tag{14}$$

The next step involves estimating $\Pr(HC > 0 \mid x)$ using a probit regression with both d and $\hat{v}$ included. A t-test of the hypothesis that the coefficient of $\hat{v}$ equals zero is a test of the exogeneity of d (Smith and Blundell 1986, Vella 1993), and failing to include $\hat{v}$ in equation 5 based on rejection of this hypothesis would yield inconsistent and biased parameter estimates. This two-stage-predictor substitution method for attending to endogeneity concerns thus allows for consistent regression estimates in the probit regression model (Terza, Basu, and Rathouz 2008).

The second variable suspected to be endogenous occurs in the second part of the model described in equation 6: the censored continuous variable X_{ik}, which represents the amount of farm subsidies received by a household. The process of testing for the endogeneity of this variable is somewhat similar. We first estimate a Tobit model for government payments and then test the statistical

significance of the coefficient of the resulting vector of residuals, $\hat{\mathbf{v}}_i$, where $\hat{\mathbf{v}}_i = X_{ik} - \hat{X}_{ik}$ and $\hat{X}_{ik}$ is computed (Greene 2008) as

$$\text{(15)} \qquad E[X_{ik} \mid x] = \Phi(\hat{\gamma}'x_i / \sigma)(\hat{\gamma}'x_i + \sigma\lambda_j)$$

where $\lambda_i = \varphi(\hat{\gamma}'x_i / \sigma) / \Phi(\hat{\gamma}'x_i / \sigma)$, σ is the standard deviation of X_{ik}, $\varphi(\cdot)$ is the standard normal probability density function, and $\Phi(\cdot)$ is the standard normal cumulative density function. Here again, a statistically significant coefficient of $\hat{v}$ when equation 6 is estimated using weighted linear regression while including both X_{ik} and $\hat{v}$ would indicate endogeneity of the continuous variable X_{ik}. To mitigate the adverse impact of such endogeneity, we use its predicted value as shown in equation 15 when estimating equation 6.

A final econometric concern is the presence of heteroskedasticity in the model, which would not be surprising since the data originate from pooled cross-sections. The regression analysis at each stage addresses this problem using the Huber-White sandwich robust-variance estimator (Huber 1967, White 1980). Since one of the main objectives is to assess the impact of income inequality on disparities in health care expenditures and some of the farm households reported negative total incomes, this issue is addressed on a state-year level using the concept of the adjusted Gini coefficient, G^* (see equation 6), developed by Chen, Tsaur, and Rhai (1982). This method of measuring inequality differs from the standard Gini coefficient in that it corrects the problems associated with the presence of negative incomes by normalizing the distribution of income so that the upper bound on the Gini coefficient is unity.[13] The adjusted Gini coefficient was further developed by Berrebi and Silber (1985) and applied later by Boisvert and Ranney (1990) and El-Osta, Bernat, and Ahearn (1995) to measure income inequality among farm families. The coefficient in each of the 48 contiguous states in seven time periods is computed as

$$\text{(16)} \quad G_Y^* = \frac{(2/N)\sum_{i=1}^{n} i s_{Y_i} - \dfrac{N+1}{N}}{\left[1 + (2/N)\sum_{i=1}^{m} i s_{Y_i}\right] + (1/N)\sum_{i=1}^{m} s_{Y_i}\left[\dfrac{\sum_{i=1}^{m} s_{Y_i}}{s_{Y_{m+1}}} - (1+2m)\right]} \quad \text{where}$$

$$S_{Y_i} = Y_i / (N\overline{Y}) \text{ and } \overline{Y} = \sum_{i=1}^{n} W_i Y_i / N > 0.$$

For each time period, W_i is the survey weight of the ith household in the state, n is sample size, N is the expanded number of farm households in the state, S_{Y_i} is the corresponding weighted income share of the ith household in the state, Y_i is the household's total income where $Y_1 \leq \ldots \leq Y_n$ with some $Y_i < 0$, and m is the size of the subset of households that have a combined weighted income of zero with $Y_1 \leq \ldots \leq Y_m$. For computational purposes, m is determined where

[13] A zero value for the standard and the adjusted Gini coefficients suggests perfect equality in the distribution of income while a value of one indicates perfect inequality. However, when some of the income values are negative, the standard Gini coefficient (unlike in the case of G_Y^*) may be overstated and in some cases may exceed the upper bound of one.

the sum of incomes over the first m households is negative and the first $m + 1$ household is positive.[14]

Results

Definitions of the variables used in the regression analysis and summary statistics are presented in Table 1. A majority of the 1.3 million farm operators in the sample identified as having positive out-of-pocket health care expenditures on average over the seven-year period had high school educations, were white, and were 50 to 64 years of age.[15] About 20 percent of this population secured health insurance coverage from private sources. The average value of 0.595 for the per-year state-level adjusted Gini coefficient (G^*) of total farm household income indicates a rather concentrated distribution. Panel A of Figure 6 indicates wide variation across the 48 contiguous states in the extent of concentration in this variable; the highest levels are associated with mountain and western states. To control for any unobserved heterogeneity in the model of health care expenditures resulting from farm location, nine dummy variables represent the agricultural production regions (depicted in panel B of Figure 6).

In the first step of the IV procedure to test for and correct endogeneity in the d_i binary variable for lack of health insurance coverage in equation 5, we obtain a vector of generalized residuals by fitting a probit regression model of this variable as shown in equation 13 using a maximum-likelihood procedure. The set of k exclusion-restriction variables ($\dot{x}_{ik}$) used in the probit regression model as part of this step includes two dummy variables, one for whether the farm operator had an off-farm job in the preceding year and one for whether the spouse had an off-farm job in the preceding year.[16] Use of the IV approach to attend to endogeneity concerns and insure that the IV estimators are consistent requires satisfaction of two conditions. First, the instruments in $\dot{x}_{ik}$ must be orthogonal to the error term, ε_i; that is, $\text{cov}(\dot{x}_{ik}, \varepsilon_i) = 0$. To the extent that ε_i is unobservable, testing of this condition is impractical; consequently, the orthogonality condition is taken as a maintained assumption (Wooldridge 2002). The second condition is that the instruments in $\dot{x}_{ik}$, while conditioning on the set of exogenous variables x in equation 5, must be non-weak (i.e., $\text{cov}(\dot{x}_{ik}, d_{il}) \neq 0$). According to the results of an empirical test based on the joint significance of these instruments in the probit regression model in equation 5 for lack of insurance coverage using a likelihood ratio test ($\chi^2(2) = 194.04$; $p = 0.000$), the instruments were correlated with d_i (for more detail, see Mallar (1977), Angrist and Krueger (1991), Bound, Jaeger, and Baker (1995), Wooldridge (2002), and Stock, Wright, and Yogo (2002)).[17] In terms of the

[14] Sample sizes for some states in the Northeast in each of the seven years of the ARMS data fell slightly short of the minimum of 30 observations needed for statistical reliability of G^*.

[15] Farm operators who fell into this age category in 2010 belong to the baby boomer generation based on the likelihood that they were born between 1946 and 1964.

[16] For this method of testing for and correcting endogeneity to work, the distributional assumption of normality in the probit regression model must hold, particularly when the elements of the vectors of explanatory variables in equations 5 and 13 are the same. Because the normality assumption cannot be asserted based on a variant of the Jarque and Bera (1987) test of normality that examines the skewness and kurtosis of the distribution ($\chi^2_{df=2} = 9{,}659$; $p = 0.00$) (see Jarque and Bera (1987), D'Agostino and Belanger (1990), and Gould (1991)) and to insure identification for equation 5, an exclusion restriction is imposed (Vella 1998) in which additional variables are included in equation 13 but not in equation 5.

[17] The probit regression results from the first stage of the two-stage procedure to test for and

Table 1. Weighted Means of Survey Variables Used in the 2PM Regression Analyses of Health Care Expenditures for 2005 through 2011 ($HC > 0$)

	Mean (Standard Error)			
Variable	**Part One**		**Part Two ($HC > 0$)**	
Dependent Variables				
Health care expenditures (1 = expenditures > 0; 0 otherwise)	0.915		—	
Total out-of-pocket health care expenditures in dollars	—		5,972	(53)
Explanatory Variables				
Operator and Household Characteristics				
Age 35–49 (1, 0)	0.320		0.318	
Age 50–64 (1, 0)	0.614		0.618	
Ethnicity: white, non-Hispanic (1, 0)	0.914		0.911	
Education: high school, some college (1, 0)	0.670		0.672	
Education: college, graduate education (1, 0)	0.258		0.261	
Male with no wife (1, 0)	0.127		0.116	
Female with no husband (1, 0)	0.039		0.037	
Household size	2.94	(0.02)	2.96	(0.02)
Established nonbeginner farmer (1, 0)[a]	0.731		0.731	
Household income in highest income quartile at $t - 1$ (1, 0)	0.250		0.253	
Adjusted Gini coefficient of income	0.595		0.595	
Household has access to internet (1, 0)	0.635		0.643	
Household has no health insurance (1, 0)	0.205		—	
Health insurance from private source, fully purchased by the household (1, 0)	—		0.189	
Farm and County Characteristics				
Government payments (thousand dollars)	5.39	(0.10)	5.46	(0.11)
Miles to nearest town with a population of at least 10,000	23.38	(0.22)	23.35	(0.24)
Region: Northeast (1, 0)	0.072		0.070	
Region: Lake States (1, 0)	0.107		0.109	
Region: Corn Belt (1, 0)	0.194		0.191	
Region: Northern Plains (1, 0)	0.083		0.083	
Region: Appalachia (1, 0)	0.130		0.130	
Region: Southeast (1, 0)	0.074		0.074	
Region: Delta (1, 0)	0.056		0.055	
Region: Southern Plains (1, 0)	0.143		0.141	
Region: Mountain (1, 0)	0.074		0.073	
County unemployment rate (percent)	6.763		6.777	

continued on following page

Table 1 (continued)

Variable	Mean (Standard Error)	
	Part One	Part Two ($HC > 0$)
Year 2006	0.143	0.143
Year 2007	0.152	0.148
Year 2008 (financial crisis)	0.145	0.144
Year 2009	0.143	0.144
Year 2010	0.141	0.144
Year 2011	0.137	0.140
Sample size	34,864	32,073
Per year average sample size	4,981	4,582
Average expanded number of farm households	1,437,298	1,314,573

[a] Established, nonbeginner farm operators are defined as operators with more than ten years of farming experience (Ahearn 2011).

Notes: The data come from the 2005–2011 ARMS (Version 1, Phase III). Standard errors for the inflation-adjusted (2011 = 100% using the GDP implicit price index (Bureau of Economic Analysis 2012)) dollar-based variables and other continuous variables are shown in parentheses. The standard errors were estimated using the bootstrapping variance estimation method with 1,000 drawn samples. The sample includes households of farm operators (and their spouses when married) who were 64 or younger.

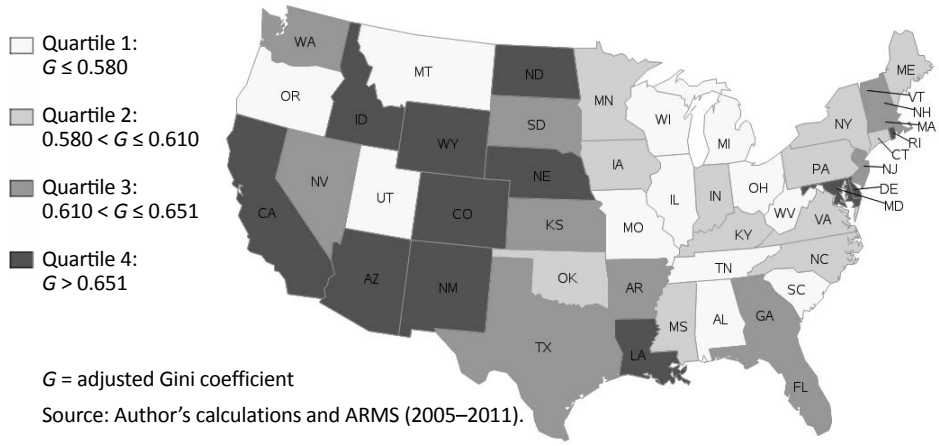
impact of each of the exclusion-restriction variables (the dummy variables that represent operators and spouses who worked off-farm in the previous year), each was statistically negatively correlated when considered separately with the likelihood of not having health insurance.

Parameter estimates from the first stage are reported in Table 2. A z-test based on the robust standard error reported in the table (0.1373) of the hypothesis that the coefficient of $\hat{v}$, the generalized residuals, in equation 5 equals zero (part of the second step of the IV procedure) showed that the coefficient was significant ($p = 0.096$). Based on this finding, $\hat{v}$ was included in the estimated model in equation 5 as an additional covariate along with the indicator of lack of health insurance (d_i) to mitigate the potential for inconsistent estimation of the probit regression model had d_i been treated as an exogenous variable.

The possibility that the variable representing government payments in equation 6 is endogenous was tested using the same two-step IV procedure previously discussed. In this case, a vector of residuals was obtained from a Tobit regression model of government payments that was estimated using a maximum-likelihood procedure. Four dummy variables were used as exclusion restrictions in the Tobit equation in the first step of the IV procedure—whether the farm operator had worked on the farm full-time (2,000 hours or more per year), whether the farm specialized in production of cash grains, whether the farm's sales exceeded $250,000, and whether the farm operation was organized as a sole proprietorship—along with one continuous variable that captured the percent of each county's income from agriculture. A test of the joint significance of the instruments in this government-expenditure Tobit regression model using an F-test ($F(5, 32{,}040) = 284.77$; $p = 0.000$) indicated

correct endogeneity of a lack of insurance coverage are not included here but are available from the author upon request.

A
Income Inequality



G = adjusted Gini coefficient

Source: Author's calculations and ARMS (2005–2011).

B
Agricultural Production Regions

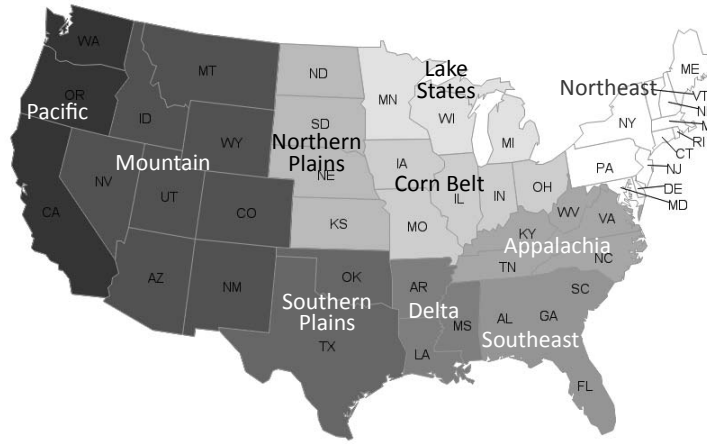


Figure 6. State-level Distribution of Income Inequality and Delineation of Regions Used in the Regression Model

that these instruments were correlated with government payments.[18] The second step involved re-estimating the regression model when including the vector of residuals as an additional explanatory variable. The exogeneity assumption for the government-payment variable was rejected based on the statistically significant coefficient of the vector of the residuals (t-ratio = –7.24; $p = 0.000$; see Smith and Blundell (1986) and Wooldridge (2002)). Accordingly, and to mitigate ill effects of the endogeneity of this variable on the estimated parameters, a vector of expected values of this variable is used instead as in equation 15 in the health-care-expenditure model.

[18] The orthogonality condition is taken as a maintained assumption here as well. The results of the Tobit regression model of government payments are available from the author upon request.

Table 2. Part One of the Two-part Model: Weighted Probit Estimates of Having Any Out-of-pocket Health Care Expenditure by Farm Household for 2005 through 2011

	$\hat{\gamma}$	Robust Standard Error	Marginal Effect	Delta-Method Std. Error
Intercept	1.0594***	0.1613	—	—
Age 35–49	0.0695	0.0722	0.0099	0.0101
Age 50–64	0.2100***	0.0707	0.0314***	0.0109
Education: high school, some college	0.3089***	0.0771	0.0482***	0.0130
Education: college, graduate education	0.2807***	0.1025	0.0375***	0.0126
Male with no wife	–0.2985***	0.0756	–0.0499***	0.0143
Female with no husband	–0.2187**	0.1093	–0.0362**	0.0204
Household size	0.0682***	0.0196	0.0099***	0.0029
Miles to nearest town with a population of at least 10,000	–0.0002	0.0008	–0.0000	0.0001
Region: Northeast	–0.2410***	0.0928	–0.0397**	0.0172
Region: Lake States	0.0022	0.0935	0.0003	0.0135
Region: Corn Belt	–0.2294***	0.0827	–0.0369**	0.0147
Region: Northern Plains	–0.1722*	0.1048	–0.0274	0.0182
Region: Appalachia	–0.1073	0.0926	–0.0164	0.0150
Region: Southeast	–0.1013	0.0876	–0.0154	0.0141
Region: Delta	–0.2202***	0.1012	–0.0359*	0.0185
Region: Southern Plains	–0.2408**	0.0906	–0.0397**	0.0168
Region: Mountain	–0.1808*	0.1024	–0.0289	0.0180
County unemployment rate (percent)	–0.0268***	0.0088	–0.0039***	0.0013
Year 2006	0.2052***	0.0768	0.0269***	0.0090
Year 2007	0.0972	0.0715	0.0134	0.0095
Year 2008 (financial crisis)	0.1874***	0.0716	0.0247***	0.0086
Year 2009	0.3925***	0.0785	0.0470***	0.0078
Year 2010	0.4692***	0.0796	0.0541***	0.0075
Year 2011	0.4266***	0.0737	0.0518***	0.0077
Household has no health insurance	–0.6179***	0.2314	–0.1147**	0.0541
Generalized residual($\hat{v}$)[a]	0.2286**	0.1373	0.0330	0.0200
Log pseudo-likelihood	–2,796,213			
Sample size	34,864			

[a] Estimated from a first-stage probit model used to deal with potential endogeneity of lack of health insurance coverage by the farm household.

Notes: Average marginal effects were obtained based on use of the *margeff* command after estimation of the probit model with the effects for dummy variables being the discrete change from the base level (Bartus 2005). Estimated *t*-statistics are based on robust standard errors. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$. The excluded categories are (i) farm operators younger than 35, (ii) farm households in which the education of the farm operator is less than a high school diploma, (iii) households of married farm couples, (iv) Pacific region, (v) year 2005, and (vi) households that have health insurance.

Table 3. Part Two of the Two-part Model: Weighted Least Square Estimates of Out-of-pocket Health Care Expenditures and Marginal Effects for 2005 through 2011

	$\hat{\beta}$	Robust Standard Error	Marginal Effect	Delta-Method Std. Error
Intercept	3,165.036***	565.10	—	—
Age 35–49	751.27***	212.99	754.54***	206.90
Age 50–64	1,412.36***	216.18	1,484.49***	206.80
Ethnicity: white, non-Hispanic	–608.07	179.50	–558.38***	164.84
Education: high school, some college	287.26	233.14	561.88**	223.25
Education: college, graduate education	725.40***	251.27	914.94***	248.04
Male with no wife	–1,562.43***	144.05	–1,699.66***	148.16
Female with no husband	–1,637.31***	213.23	–1,677.82***	212.83
Household size	111.87***	40.20	131.41	113.31
Established nonbeginner farmer	143.10	123.38	165.12***	41.12
Household income in highest income quartile at t – 1	528.93***	128.95	485.72***	118.42
Adjusted Gini coefficient of income	763.81	516.19	701.40	474.02
Household has access to internet	327.21***	114.93	300.48***	105.54
Health insurance from private source	2,942.57***	152.73	2,702.13***	140.48
Expected government payments (thousand dollars)	46.03***	4.82	42.27***	4.43
Miles to nearest town with a population of at least 10,000	–4.57**	2.17	–4.39**	2.13
Region: Northeast	–549.51**	228.22	–739.40***	226.12
Region: Lake States	–301.48	221.32	–274.92	219.17
Region: Corn Belt	–363.28*	216.90	–557.88***	212.99
Region: Northern Plains	306.75	267.99	100.82	268.52
Region: Appalachia	–285.54	218.37	–362.57*	217.91
Region: Southeast	–322.31	221.92	–390.03*	218.54
Region: Delta	37.12	271.91	–194.86	268.87
Region: Southern Plains	274.95	250.69	–9.05***	248.02
Region: Mountain	395.48*	239.88	170.35	245.90
County unemployment rate (percent)	–47.46**	22.80	–68.14***	22.42
Year 2006	209.99	220.61	366.32*	214.63
Year 2007	528.14**	209.36	574.56***	203.92
Year 2008 (financial crisis)	–176.23	198.68	–9.14	193.10
Year 2009	707.16***	225.98	966.22***	219.98
Year 2010	1,077.41***	216.22	1,367.00***	211.62
Year 2011	1,023.00***	196.31	1,294.99***	192.05
R^2	0.1102			
Sample size	32,073			

Notes: *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$. The average marginal effects were obtained based on the use of the *margins* command in STATA after estimation of the 2PM using the *tpm* routine as delineated by Belotti et al. (*forthcoming*). The excluded categories are (i) farm operators younger than 35, (ii) households of nonwhite operators, (iii) farm households in which the education of the farm operator is less than a high school diploma, (iv) households of married farm couples, (v) beginner farmers, (vi) income of the household is in the first three quartiles of the income distribution at t – 1, (vii) the household has no access to the internet, (viii) the household has no health insurance coverage or coverage is from other than a private source, (ix) Pacific region, and (x) year 2005.

Table 3 reports results of the 2PM estimation of the parameters of the WLS regression model of health care expenditures. The null hypothesis that all slope coefficients equal zero was rejected based on the result of an *F*-test of the overall fit of model specification ($F(31; 32{,}043) = 41.63$; $p < 0.000$). Despite the overall statistical significance of the model, the R^2 value of 11.02 percent indicates that the estimated WLS model provides only modest explanatory power since it explains only about 11 percent of the variation in health care expenditures. The low R^2 value, however, is typical in studies of health care use, which have generated R^2 values on the order of 20 percent or less (Newhouse et al. 1989, Diehr et al. 1999). While the distribution of out-of-pocket health care expenditures is skewed, as demonstrated in Figure 5, an auxiliary model that was estimated based on a log transformation of this dependent variable produced, on the margin, a slightly lower, practically equivalent R^2 value of 10.91 percent. A linear model is chosen for the analysis given the lack of improvement in the fit of the model that was based a log-transformation of *HC* and the fact that such a transformation substantially reduces the weight on a high-expenditure observation (Ellis et al. 2012), which results in severe overestimation at the upper tail after retransforming the logged *HC* values to their original dollar-based values. Use of the linear rather than the log-linear model offers two additional benefits: the regression coefficients are easier to interpret since they remain on a dollar-based scale and expected values of *HC* can be computed directly per equation 7 without having to retransform the linear predictions by means of exponentiation.

Of the dummy variables related to operator education, having a graduate degree appears, *ceteris paribus*, to have a statistically significant impact on health care expenditures, a result that points to greater use of health care services among the highly educated.[19]

Similarly, farm operators who are older, including baby boomers (aged 50–64), appear to expend more on health care than operators who are younger than 35, plausibly due to advantageous economic positions (El-Osta and Morehart 2009). This result is in line with patterns for the general U.S. population (Foster and Kreisler 2011, Cohen and Yu 2012, Health Care Cost Institute 2012) and is supported by evidence from the ARMS sample; the average expenditure for those older farm operators was nearly 1.4 times the average for younger operators—$5,788 versus $4,117.

The results shown in Table 3 indicate that farm householders who are white and who are unmarried (expected for single individuals (Hawk 2011)) tend to spend less on health care ($558 and about $1,700 less per annum, respectively, based on estimated marginal effects) than non-white and married householders.

Of the variables representing characteristics of the household and the farm, having a household income in the upper quartile of the distribution, an expectation of an increase in government payments, access to the internet, and health insurance from a private source all are found to have a positive

[19] This finding, which concerns farm households, contradicts the results of some studies of the general U.S. population in which relatively highly educated individuals were more likely to manage their use of health care resources efficiently and, accordingly, were able to produce a given level of health using less inputs (Yoo 2011). A study by Meara, Seth, and Cutler (2008) pointed to the likelihood of higher rates of obesity and tobacco use among poorer individuals (who also tend to be less educated), resulting in adverse impacts on health outcomes and, consequently, on health care expenditures.

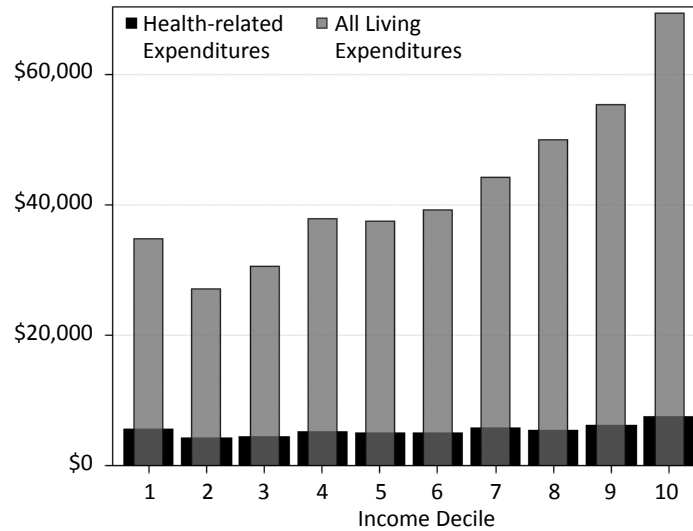


Figure 7. Health-related Expenditures and Total Living Expenditures of Farm Households by Income Decile for 2005 through 2011

association with health care expenditures. In fact, the results of estimates of the marginal effects of the explanatory variables indicate that health insurance obtained from a private source has the greatest positive impact on health care expenditures at $2,700. By comparison, negative associations are found between health care expenditures and (i) increased distance to a town with a population of 10,000 and (ii) farms located in the northeast, southeast, Appalachia, the Lake States, and the Corn Belt (Figure 6, panel B).

The positive association found in the estimated regression model between incomes in the upper quartile of the distribution and health care expenditures is supported by Figure 7, which shows a higher level of expenditures in the last two income deciles. Since the levels of health care expenditure for farm households for the second decile through the eighth decile are nearly equal, it is less surprising to see a lack of statistical significance for inequality of income measured by G^* on health care expenditures. In our models, income inequality is based on current incomes rather than on incomes with lag periods of five, fifteen, or even twenty-five years as used by Blakely et al. (2000) and Subramanian and Kawachi (2004), who cast doubt on the notion that economic inequality has an instantaneous effect on population health and, consequently, in the case of this study, on health care expenditures.

The results presented in Table 3 show an upward time trend on health care expenditures, particularly after the 2008 financial crisis.[20] All else equal, health care use by farm households appeared, on an inflation-adjusted dollar basis, to increase with the subsequent financial recovery witnessed in the general economy. This also is not surprising since the farm sector, unlike other sectors in the economy, was not severely adversely impacted by the recession (average household income between 2008 and 2009 in the agricultural sector declined

[20] The null hypothesis that all slope coefficients of the year dummies are equal to zero is rejected based on a joint F-test of significance ($F(6; 32{,}024) = 11.56$; $p < 0.000$).

by only 3.3 percent (Park et al. 2010)) and the recovery, according to higher average net cash incomes, was most evident among dairy, hog, and cattle producers (Park et al. 2010).[21]

Table 4 presents the results of decomposition of the inequality in health care expenditures per equations 10 and 12. In the first column, six covariates have the strongest influence in explaining the inequality in health care expenditures. In order by the strength of the impact on the explained variation around the mean of expenditures as measured by R^2, they are (i) health insurance coverage from a private source, (ii) expected farm program payments, (iii) a male unmarried operator, (iv) an operator 50–64 years of age, (v) an operator with a college and/or graduate education, and (vi) household income in the upper quartile of the income distribution. The second column identifies private-source health insurance as the predominant driver of inequality of health care expenditures (48 percent). The combined contribution of the demographic and family-structure characteristics represented by operators in the baby boomer age group, those who are highly educated, and unmarried male householders is 23 percent. Expectations of increased government payments and incomes in the highest quartile, which capture some characteristics of the farm and its household, contribute 16 percent.

Summary and Concluding Remarks

A data set for 2005–2011 from a national survey of farmers is used to examine how much of the inequality in the distribution of health care expenditures (for both private insurance and out-of-pocket medical expenses) among nonelderly farm operator households is explained by reliance on privately acquired health insurance. To achieve this objective and the subsidiary objective of discerning the role of other factors, a WLS regression is employed from a two-part regression model.

The results point to the importance of income (being in the upper income quartile) but not to inequality of income as impacting health care expenditures. The lack of evidence to support the income inequality hypothesis for health care expenditures is in line with results from other studies of the general U.S. population (Mellor and Milyo 2001, 2002). Thus, redistributive tax-based income policies are likely to have no immediate bearing on equalization of the distribution of health care expenditures. In contrast, policies aimed at increasing the income position of farm households will, given the statistically significant coefficient of the variable representing the upper-quartile income category, have the desired effect, although it will be a mild one at best given the marginal impact of about $500. The regression results also point to a positive but minute marginal impact on health care expenditures from increases in expected farm subsidies.

[21] An auxiliary model with interactions between the dummy variable indicating privately sourced health insurance and the time and regional dummy variables was also estimated. The results of the primary model regarding higher health care expenditures in 2010 and 2011, which were likely due to higher incomes of livestock producers, are not particularly surprising in view of the results of the auxiliary model with interactions. Specifically, the auxiliary model results show greater use of health care by households in 2010 and 2011; for households in the livestock-dominated regions of the northeast, Lake States, Appalachia, and mountain states where feed costs are relatively low and access to high-demand markets is strong; and for households that obtained health insurance primarily from a private source.

This positive association between health care expenditures and both expected government payments and farm household incomes is consistent with the general view of Cutler, Lleras-Muney, and Vogl (2008) that income and wealth improve access to health inputs (e.g., medical care and food). Improved access and attendant positive impacts on health outcomes likely allow farm operators to maintain or even further increase the earning capacity of the household since the operator can stay healthy and consequently remain

Table 4. Decomposition of Inequality in Health Care Expenditures for 2005 through 2011

	Relative Factor Inequality Weight s_j (*HC*)	Portion Explained by the jth Factor p_j(*HC*) ×100
Age 35–49	–0.3311	–3.0052
Age 50–64	1.0721	9.7291
Ethnicity: white, non-Hispanic	–0.0003	–0.0033
Education: high school, some college	–0.1206	–1.0947
Education: college, graduate education	0.4619	4.1920
Male with no wife	0.9883	8.9685
Female with no husband	0.3550	3.2215
Household size	0.0600	0.5452
Established nonbeginner farmer	0.0790	0.7169
Household income in highest income quartile at t – 1	0.4097	3.7180
Adjusted Gini coefficient of income	0.0585	0.5310
Household has access to internet	0.2460	2.2331
Health insurance from private source	5.2984	48.0798
Expected government payments (thousand dollars)	1.3648	12.3850
Miles to nearest town with a population of at least 10,000	–0.0207	–0.1879
Region: Northeast	0.0986	0.8949
Region: Lake States	0.0291	0.2645
Region: Corn Belt	0.0419	0.3803
Region: Northern Plains	0.1021	0.9267
Region: Appalachia	0.0754	0.6846
Region: Southeast	0.0382	0.3475
Region: Delta	0.0012	0.0110
Region: Southern Plains	0.0570	0.5179
Region: Mountain	0.0568	0.5159
County unemployment rate (percent)	0.0603	0.5475
Year 2006	–0.0082	–0.0751
Year 2007	0.0333	0.3022
Year 2008 (financial crisis)	0.0511	0.4641
Year 2009	0.0569	0.5169
Year 2010	0.2257	2.0485
Year 2011	0.1788	1.6227
Explained inequality	11.0201	
Residual inequality	88.9799	
Total	100.00	100.00

Notes: Results are based on authors' calculations using SAS/IML. Corresponding results were also obtained using the *ineqrbd* command in STATA (Fiorio and Jenkins 2007).

productive both on and off the farm. Many studies have asserted the strong role played by negative health shocks in predicting retirements and reduced labor force participation even among households with health insurance coverage (e.g., Smith 1999, 2004). The results support existence of a connection between improved access to health inputs and improved economic conditions in general: farm households responded to the 2008 financial crisis by first holding health care expenditures in 2008 to 2005 levels and then by increasing health care spending in 2009 when the financial crisis officially ended and post-crisis in 2010 and 2011.

The regression-based inequality decomposition technique shows that the main driver of inequality in health expenditures for farm households was purchasing health insurance from a private source rather than through an employer or public plan or having no coverage. This finding, which asserts the presence of health insurance even when it is paid for entirely by the household, agrees with the commonly accepted notion that lack of health insurance coverage is one of the major sources of disparities in both health outcomes and health expenditures. Uninsured individuals are more likely to postpone and/or go without medical care, preventative care, and prescription medicines (Tu 2004, Kaiser Commission on Medicaid and the Uninsured 2011).

Based on data from the ARMS, nearly 84 percent of the close to 1.1 million farm operators in 2011 were in the baby boomer category and will qualify for Medicare coverage in the near future. Thus, a large portion of aging farmers will be contributing to growth in the escalating cost of Medicare spending, which in 2011 was $554 billion, 21 percent of all national health expenditures (Centers for Medicare and Medical Services 2013). Like aging individuals in other segments of the population, aging farm operators (whose occupations are relatively hazardous, resulting in high rates of injury and chronic illness) are already shouldering some of the rise in the cost of health care through their premiums, cost-sharing, and other out-of-pocket costs (National Advisory Committee on Rural Health and Human Services 2008, Jones et al. 2009).

Federal policies aimed at diminishing rural health disparities and controlling the rising cost of health care will be advantageous in shrinking the escalating cost of Medicare, and such policies will also lower the share that aging farmers have to spend on their own health care. These policies are consistent with efforts to improve the economic well-being of farm households in rural areas where poverty rates are high relative to urban areas (DeNavas-Walt, Proctor, and Smith 2008). This observation of the potential for improved economic well-being among aging farmers due to policies aimed at lowering health care costs and diminishing rural health disparities is supported by the statistical significance of the coefficient of the 50–64 age category, which reflects greater health care expenditures by farmers in the baby boomer generation (by nearly $1,500 in terms of marginal impact) and the sizable contribution of these older farmers to inequality in health care expenditures (9.73 percent).

The ACA, which requires all individuals not already covered by a health care plan (employer-sponsored plans or Medicaid) to either secure coverage or pay a penalty, includes various tax credits as incentives to low-income individuals to purchase health insurance (Government Accountability Office 2012) and may foster mitigation of the spending disparity among farm households. If all states eventually implement the Medicaid expansion provision of the ACA, uninsured farm households that fall at or below 138 percent of the federal poverty line will become eligible for Medicaid (Ahearn, Black, and Williamson

2013).[22] Furthermore, farm households that have incomes at 138–400 percent of the poverty line and do not have employer-based health insurance will be eligible for premium tax credits to purchase insurance in newly created health insurance exchanges. A provision in the ACA will limit the cost of health insurance for qualified farm households that purchase insurance through the exchanges, capping the premium payment at between 2 percent and 9.5 percent of the household's income (Ahearn, Black, and Williamson 2013).

With its prohibition on charging higher rates or otherwise discriminating against farm operators and their family members who have pre-existing conditions, the law will allow for wider choices in terms of prices and coverage, thereby increasing access to health care. This, in turn, could negate the pressure to hold off-farm jobs as a means of securing health insurance (D'Antoni and Mishra 2012, Bharadwaj, Findeis, and Chintawar 2013). Under the ACA, all farmers and their family members who have been covered under a privately purchased insurance plan since October 1, 2013, are now able to purchase plans with similar coverage under their existing policies through the insurance exchanges, which, depending on how they are designed, are expected to contain costs through improved pooling of risks. A consequence of the law, then, is an improvement in the economic well-being of a sizable segment of farm households (17.1 percent, as shown in Figure 3) that previously depended on directly purchased private insurance plans that have tended to have higher premiums, deductibles, and co-pay charges than other types of health insurance plans (Ahearn 2009, Mishra, El-Osta, and Ahearn 2012).

Notions of a reduction in the disparity in health care spending and an improvement in the economic well-being of farm households because of implementation of the ACA are supported by these findings. Specifically, since private health insurance explains a large portion of the inequality of spending among farm households, the ACA may give a larger number of farm operators and their households greater access not only to health insurance but also to better insurance policy and treatment options such as ones provided by employer-sponsored plans. Furthermore, while the marginal impacts of the estimated variables indicate that farm households covered by the more expensive privately acquired plans tend to spend about $2,700 more annually on health care than farm households that have no health insurance or that obtained insurance from other sources, the potential for lower-cost coverage through the ACA is apt, on the margin, to improve the economic welfare of those households by providing additional money for discretionary spending. Such improvements in the economic position of households may come primarily from reducing the premium portion of their out-of-pocket expenditures through ACA's cost-containment measures as new policies are purchased with grandfathered levels of coverage, from premium tax credits, and from expansion of access to health care.[23]

An unintended consequence of the ACA for rural economies is the potential for additional stress on already overburdened rural health-care delivery systems (Ewing 2011). The ACA is expected to add 5 million rural Americans

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[22] In 2014, only 27 states and the District of Columbia were moving toward Medicaid expansion (see *www.advisory.com/Daily-Briefing/Resources/Primers/MedicaidMap*).

[23] A caveat is in order here. In many cases, insurance costs under the ACA may end up rising through higher premiums of privately purchased plans for households that do not qualify for federal premium subsidies because of their high incomes and because of the more generous levels of coverage required under the ACA.

to the rolls of the insured by 2019. A large percentage of those Americans will be baby boomers who will have a greater demand for medical services than younger Americans (Ewing 2011). The results of this study support the notion that aging farmers will have greater demand for health care; farmers aged 50 to 64 spent, on average, nearly $1,500 more on health care than younger farmers.

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