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# Is U.S. Agricultural Productivity Growth Slowing?

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**ABSTRACT:** Tests for a structural break in a time series typically involves partitioning the data into two sub-periods and compare the resulting mean rates of growth. A more formal approach involves estimating the regression parameters for each sub-period and testing the equality of the two sets of parameters. An important limitation of both approaches is that the breakdate must be known a priori. The researcher must either pick an arbitrary breakdate or pick a breakdate based on some known feature of the data. The test results can be 'uninformative' because they can miss the true breakdate, or can be 'misleading' because the breakdate is endogenous and the test can indicate a break when none in fact exists. This paper tests for slower productivity growth in agriculture using techniques that allow for unknown structural breaks jointly with a possible unit root that can have either or both stochastic and deterministic components.

# Is U.S. Agricultural Productivity Growth Slowing?

## 1. Introduction

The notion that U.S. agricultural productivity growth is slowing is implicit in the recent policy debate. Alston and Pardey (2009) link the decline in real food prices over the last 50 years to growth in agricultural productivity, but they question if the pattern of growth over the next 50 years will mirror that of the last 50 years. They answer this question in James et al. (2009) by describing a significant slowdown in productivity growth after 1990, a phenomenon "…that is not a temporary cyclical event but more structural and sustained." Slower productivity growth, they argue, will lead to an erosion in international competitiveness and to a widening gap between global supply and demand for agricultural products. The implications for food security, especially among the world's poor, are decidedly negative.

The above assessment reveals considerable pessimism regarding the prospects for continued strong productivity growth. But other authors project a less pessimistic tone. Ball et al. (2010) addressed the international competitiveness of agriculture in the European Union (EU) and the United States. They found that the United States maintained a competitive advantage throughout the period 1973-2002, except for the years 1983 to 1985. Further, they attributed United States competitiveness to relatively rapid productivity growth.[1]

More recently, Wang, Schimmelpfennig, and Fuglie (2012) looked at productivity growth rates for the twelve countries included in the Ball et al. (2010) study to determine if any of these countries experienced a significant slowing in the rate of growth. France and Sweden exhibited a statistically significant slowdown. Fuglie (2008) used FAO data to assess globally the slowdown

[1] They provide a formal definition of the concept of competitiveness and relate it to the more conventional concept of relative productivity growth.

hypothesis. He found no evidence of a systematic slowdown in the rate of sector-wide productivity growth.

In this paper, we revisit the slowdown hypothesis. We believe there are a number of issues that were not adequately addressed in previous studies. First, productivity growth has often been measured by a single factor, such as output per acre. Alston, Beddow, and Pardey (2009) point to declining crop yields as the proximate cause of recent food price increases. But growth in land productivity depends on changes in the levels of other inputs (e.g., fertilizers) per acres, as well as the pace of technical innovations. Total factor productivity (TFP), on the other hand, measures the contribution to output of innovation alone and is a more meaningful indicator of sustainability. Second, arbitrary or convenient sub-periods have often been chosen when investigating the slowdown hypothesis, but the timing of transitory shocks and the occurrence of statistically significant structural breaks might fall in more irregular time periods. Those exogenous shocks that have briefly but forcefully influenced productivity growth must be identified if unmeasured factors influencing changes in productivity are to be considered, as Thirtle, Holding and Jenkins (2004) argue they should. And third, even though the time series properties of the productivity data might influence the characterization of a possible slowdown, this avenue of investigation has not been pursued. Phillips (1986; 1998; 2010) steadfastly argues that those time series tests that make rigorous conclusions possible must be part of the investigation.

Our investigation of the slowdown hypothesis focuses on statistics on agricultural productivity published by the U.S. Department of Agriculture (USDA). The USDA has long been concerned with sectoral productivity growth. In fact, the USDA in 1960 was the first agency to introduce a multifactor productivity measure into the Federal statistical program.

Today, the Department's Economic Research Service (ERS) bases its official productivity statistics on a sophisticated system of production accounts. The model of productivity growth is based on the translog transformation frontier. It relates the growth rates of multiple outputs to the cost-share weighted growth rates of labor, capital, and intermediate goods.[2]

The applied USDA model is quite detailed. The changing demographic character of the agricultural labor force is used to construct a quality-adjusted index of labor input.[3] Similarly, much asset specific detail underlies the measure of capital input. Construction of the measure of capital input begins with estimating the capital stock for each component of capital input. For depreciable assets, the capital stocks are the cumulation of past investments adjusted for discards of worn-out assets and loss of efficiency of assets over their service life.[4] The capital stocks of land and inventories are measured as implicit quantities derived from balance sheet data. The index of capital input is formed by aggregating over the various capital assets using cost-share weights based on asset-specific rental prices. As is the case for labor input, the resulting measure of capital input is adjusted for changes in asset quality. The contributions of feed and seed, energy, and agricultural chemicals to output growth are captured in the index of intermediate inputs. An important innovation is the use of hedonic price indexes in constructing measures of fertilizers and pesticides consumption. The result is a time series of productivity indexes spanning the years 1948 to 2009.

## 2. Patterns of Productivity Growth

Before addressing formally the hypothesis of slower productivity growth, we provide an overview of the patterns of growth over the postwar period. Table 1 reports average annual rates

[2] A complete description of the USDA model can be found in Ball et al. (1999).
[3] See Jorgenson and Griliches (1967) for a discussion of input quality.
[4] See Ball et al. (2008) for further discussion.

of productivity growth for the full 1948-2009 period and for twelve peak-to-peak sub-periods.[5] Over the 1948-2009 period, productivity growth averaged 1.52% per year. Growth in productivity was negative in only one sub-period (1948-53), and fell below 1% only then and in three other sub-periods. The average rate of growth in productivity exceeded 2% per year in six of twelve sub-periods.

The USDA measure of output is equated with gross production. As a result, inputs of intermediate goods, obviously crucial to agricultural production, are treated symmetrically with labor and capital inputs. From table 1, we see that growth in intermediate inputs averaged 1.43% per year over the 1948-2009 period. Energy use increased less than 1% per year, but the growth rate in chemicals input exceeded 2.5% annually. Inputs of farm origin increased an average of 1.15% per year.

Labor input in agriculture contracted at an average annual rate of 2.51% over the postwar period. Moreover, this pattern persisted through all twelve sub-periods.

Capital input increased dramatically in the immediate postwar period. Service flows from durable equipment increased at an 11.1% average annual rate over the 1948-53 period, reflecting rapid mechanization of agriculture. But the average rate of growth over the full 1948-2009 period was less than 1% per year. Other capital inputs, including non-residential structures and inventories, exhibited similar patterns. Land input, however, declined at a 0.52% average annual rate. Overall, capital input declined 0.21% per year.

In spite of the declines in capital and labor inputs and the relatively modest increase in intermediate inputs, growth in farm sector output averaged 1.63% per year over the 1948-2009

[5] Annual data are available electronically at http://www.ers.usda.gov/data/agproductivity/

period. These results leave little doubt that productivity growth was the principal factor responsible for economic growth in postwar agriculture. In the late 1970s, however, the capacity of the sector to sustain historical rates of productivity growth came into question. In 1972, the subsidized retirement of land totaled 60 million acres. This figure was less than 20 million acres in 1973 and declined to zero in 1974. In the aftermath of a rapid expansion in land use and total agricultural output, such indicators as crop output per acre of land lent some credence to the hypothesis of slower productivity growth and pending economic scarcity.

Between 1973 and 1979, the average annual rate of growth in output was 2.26%. But nearly three-fourths of the growth in output over this period was accounted for by growth in inputs. In contrast, input growth accounted for only 0.46 percentage points of the 2.65% annual growth rate achieved during the 1969-1973 period. The contraction in labor input continued during the 1973-79 period, but at a much slower pace than during the previous six sub-periods. The rate of growth in capital input over the 1973-79 period was second only to that of the 1948-53 period. Service flows from durable equipment increased more than 3.5% per year. Despite a threefold increase in the price of petroleum fuels following the 1973 oil embargo, the rate of growth in energy consumption exceeded 4% per year. Consumption of agricultural chemicals increased at a 3.29% annual rate. The 0.62% average annual rate of growth in total factor productivity was less than one-third the growth rate over the 1969-73 period and was well below trend.

The ‘specter of scarcity’ of the 1970s gave way to expectations of chronic economic surpluses in the 1980s. In 1982, the land area set aside totaled some 10 million acres. This total exceeded 80 million acres in 1983.[6] Output growth slowed to just over 1.5% per year in the 1979-81 sub-period and less than 1% per year over the succeeding 1981-90 sub-period. But total input growth was negative in both periods. Growth in total factor productivity rebounded. In fact, the 3.39% rate of growth in productivity during the 1979-81 sub-period was second only to the 3.53% average annual rate of growth over the period 1957 to 1960.

The early 2000s saw renewed concerns about a possible slowing of productivity growth. The average annual rate of growth in productivity was less than 1% per year over the 2000-07 period. In this way, the situation was similar to that in the 1970s. During the period 1973-79, growth in output did not deviate much from trend. Durable equipment, energy, and chemicals inputs, however, exhibited significant positive growth at rates far in excess of the incremental growth in output, accounting for the measured decline in productivity growth.

During the 2000-07 sub-period, we saw the emergence of bio-fuels as a major source of demand for grains and oilseeds. Corn used in ethanol production in 2007 accounted for roughly one-fourth of total demand. In response, producers abandoned usual crop rotations, opting instead for continuous planting of corn. The land area planted to corn increased some 15 million acres in 2007. Despite increased consumption of fertilizers and pesticides, yields per acre were essentially unchanged.

Overall, the gains in agricultural productivity have been quite impressive. However, it remains to be determined if slower productivity growth during the 1970s and more recently

[6] In January of 1983, the Reagan administration implemented a program that was to make sharp cuts in production and reduce Government stocks at the same time by paying farmers not to produce, with payment to be made in the form of commodities from Government stocks. As was the case in earlier legislation, farmers had to idle land to be in the price support program. A total of 82 million acres were pledged for diversion to conserving uses under the payment-in-kind (PIK) and related acreage reduction programs, the largest amount of land ever taken out of production in the United States. This meant that over a third of the land normally planted to wheat, rice, upland cotton, and corn and sorghum grain was idled.

represent a shift in the long term path of productivity growth. This is the focus of the following section.

## 3. The Slowdown Hypothesis

Tests for a productivity slowdown typically involve partitioning the sample into two sub-periods and comparing the resulting mean rates of growth. A more formal approach, attributed to Chow (1960) involves estimating regression parameters for each sub-period and testing the equality of the two sets of parameters using an $F$-test. An important limitation of both approaches is that the breakdate must be known a priori. The researcher must either pick an arbitrary candidate breakdate or pick a breakdate based on some known feature of the data. As Hansen (2001) points out, the test results can be 'uninformative' because they can miss the true breakdate, or can be 'misleading' because the breakdate is endogenous and the test can indicate a break when none in fact exists. Therefore, we propose to test for a productivity slowdown using techniques that allow for unknown structural breaks jointly with a possible unit root that can have either or both stochastic and deterministic components.[7]

### *Analytical Framework*

To address the slowdown hypothesis, we posit a simple trend model

(1) $$\ln TFP = c_0 + \tau_0 t + \varepsilon_t,$$

[7] Time series filters such as the Hodrick-Prescott filter that separate a time series into trend and cyclical components are not useful here. Their assumption of a stationary cyclical component driven by stochastic cycles assumes away much of the variability that we wish to analyze.

and test the null hypothesis of a stable linear model against the alternative of ‘breaks’ in the parameters in the trend regression. The simple comparative static of (1) with respect to time $t$ yields

$$\frac{d \ln TFP}{d\ t} = \tau_0$$

where $\tau_0$ is the trend rate of productivity growth over time.

***Structural Breaks***

To determine if there are structural breaks, we first conduct the Elliott and Müller (2006) ‘quasi-Local Level’ (qLL) test. The qLL test allows for a large class of breaking trends including many or relatively few breaks, breaks that occur close to one another, breaks that occur at regular intervals, and smooth transitions in $\tau_0$ that statistically indicate a break in trend. This allows us to determine if the time series has been free of structural breaks over the study period. The qLL test accomplishes this by assessing the general persistence of the time variation in the regression coefficients in (1). The first row of table 2 reports the qLL test results. We can reject the null hypothesis of fixed coefficients, both intercept and trend, at the usual confidence level. However, this qLL test does not provide information on the timing of the structural breaks. Additional break tests are required to determine the number of breaks and their form over time.

***Unit Root Tests without Structural Breaks***

Our next objective is to determine if there is also a stochastic component to the time series and if this result is sensitive to the structural breaks just identified. As Perron (1989) shows, standard unit root tests that ignore possible structural breaks will be biased in favor of failing to reject the null of a unit root. Accordingly, we conduct unit root tests with and without structural breaks.

We conduct three fundamentally different unit root tests without structural breaks. The augmented Dickey and Fuller (1979, 1981) unit root test and the autocorrelation robust Phillips and Perron (1988) test are the most widely used tests for stationarity, but these tests have low power, increasing the likelihood of type II errors (Schmidt and Phillips, 1992). It is possible that these first two unit root tests could fail to reject the null hypothesis of a unit root when the series are actually stationary. For this reason, we compare these test results with those from the Kwiatkowski, Phillips, Schmidt, and Shin (1992) test that posits a null of no unit root. Reliability is greatly enhanced if all three unit root tests give results that are in agreement. The results reported in the first panel of table 3 shows that this is in fact the case, indicating that the TFP series is stationary with a deterministic trend.

***Unit Root Tests with Structural Breaks***

The result of the qLL test indicates the presence of one or more structural breaks in the time series, and we test if the absence of a stochastic trend and a unit root are influenced by possible structural breaks. We employ two types of unit root tests that allow for possible structural breaks, the Clemente, Montanes and Reyes (CMR) (1998) test and the Zivot and Andrews (1992) test. While both tests maintain a unit root under the null hypothesis, the former has as the alternative hypothesis stationarity when only a shift in the intercept is present and the latter when there is trend stationarity with trend breaks.

The CMR test allows us to determine if the time series is stationary with possibly two intercept shifts, but it does not allow for a trend break. There are two models under the CMR test. One is an ‘additive outliers’ (AO) model, which can capture a sudden change in the series. The other is an ‘innovational outliers’ (IO) model which captures a more gradual shift in the intercept of the series. Table 3 shows the test results for both the AO and IO models, which indicate non-stationarity with either one intercept shift or a double-intercept shift.

The Zivot and Andrews test for a unit root has as the alternative hypothesis that the time series is trend stationary with a one-time break in the level or trend or both occurring at an unknown point in time. They choose the breakpoint that minimizes the one-sided *t*-statistic under the null of a unit root (i.e., $\alpha = 1$ in equation (2) below):

(2) $$y_t = \mu + \alpha y_{t-1} + \varepsilon_t$$

The trend-stationary alternative hypotheses include:

(3) $$y_t = \mu_1 + \beta t + (\mu_2 - \mu_1) DU_t + \varepsilon_t$$

(4) $$y_t = \mu + \beta_1 t + (\beta_2 - \beta_1) DT_t^* + \varepsilon_t$$

(5) $$y_t = \mu_1 + \beta_1 t + (\mu_2 - \mu_1) DU_t + (\beta_2 - \beta_1) DT_t^* + \varepsilon_t$$

where $DU_t = 1$ if $t > T_B$ and zero otherwise and $T_B$ is the breakpoint, and $DT_t^* = t - T_B$ if $t > T_B$ and zero otherwise. Equation (3) allows for an intercept shift, while equation (4) allows for a trend break. Equation (5) combines both an intercept shift and a trend break. We note, however, that the Zivot and Zndrews tests only allow for one breakdate. The results presented in table 3 indicate that the TFP series is stationary when an intercept shift or both an intercept shift and a trend break are included. The Zivot and Andrews tests place an optimal breakpoint at 1985.

The evidence suggests that the TFP series is trend stationary with structural breaks. The CMR test indicates the presence of a unit root. But this test does not include a possible break in trend under the alternative hypothesis. As noted by Perron's (1989), unit root tests that ignore possible structural breaks will be biased in favor of failing to reject the null of a unit root. We conclude that it is necessary to include a deterministic trend in a regression model with TFP as the dependent variable to ensure stationarity, but additional tests are required to fully characterize the nature and timing of the structural change.

***Estimating Structural Breaks with Unknown Timing***

The next step we take is to estimate the breakdate or breakdates that best depict the trend rate of growth of TFP over the study period. Hansen (2001) suggests that one candidate for a breakdate is the date that yields the largest test statistic in a sequence of Chow tests. The problem with this approach is that the Chow test is designed for a break that separates the entire sample into two subperiods; only one breakdate is considered. What if the breakdates are different for an intercept shift and a break in trend? Although Bia and Perron (1998) develop a test for multiple structural changes, their test is appropriate only when there are no trending regressors. Similarly, the CMR test does not test for a break in trend. We could, however, use the Zivot and Andrews test results as a first step in determining optimal breakdates.[8] For this purpose, we rewrite equation (1) to allow for multiple breaks in intercept and trend as follows:

(6) $$\ln TFP = c_0 + c_1 D_{B_1} + \tau_0 t + \varepsilon_t$$

(7) $$\ln TFP = c_0 + c_1 D_{B_1} + \tau_0 t + \tau_1 D_{B_2} t + \varepsilon_t$$

where $D_{B_1} = 1$ if $t > TB_1$ and zero otherwise, $D_{B_2} = 1$ if $t > TB_2$ and zero otherwise, and $TB_i, i = 1{,}2$, represent possibly different breakdates. We note that additional breakdates may be necessary to completely characterize the breaking activity. We further note that breaks in trend and intercept need not be contemporaneous.

[8] Lee and Strazicich (2001) evaluate the interaction between this unit root test and optimal breakdates.

In estimating the parameters in equations (6) and (7)—including the breakdates and their coefficients—we use least squares. The least squares breakdate is that which minimizes the full sample sum of squared errors (equivalently, minimizes the residual variance). We conduct qLL tests on sub-sample periods before and after sequential breakdate candidates until no further breakdates are detected.

Given the results of the qLL test (see table 2, rows 2 through 5) and the Zivot and Andews test (see table 3), we estimate equation (6) using 1985 as the intercept breakdate. The results are presented in table 4. The qLL test based on this estimation shows that there is at least one additional structural break. Next, we estimate equation (7) with both an intercept shift and a trend break in 1985. The results reported in table 4 show that the break in 1985 shifted the intercept to a higher level, but the trend break in 1985 was not significant. This suggests that there may exist different breakdates for intercept and trend. Accordingly, we estimate equation (7) with an intercept shift in 1985 and a trend break in different years sequentially. We plot the residual variances as a function of individual trend breakdates in figure 1. The sum of squared errors will have a minimum near the true breakdate. According to figure 1, we can find two well defined minima—a global minimum in 1974 and the second minimum in 1978 for the trend breakdates. We then conduct the qLL test using 1985 as the intercept shift and 1974 as the trend break. The results reported in table 4 indicate that there are no further breaks.

Having independently verified the possibility of breaks in intercept and trend, we present three sets of results in tables 4 and 5. These include a trend break in 1985, both an intercept shift and a trend break in 1985, and a trend break in 1974 and an intercept shift in 1985. The intercept shift is positive and statistically significant, indicating a shift to a higher level of productivity in 1985. The trend break in 1985 is positive and significant when we allow only a trend break, but the trend break becomes negative in both 1974 and 1985 when an intercept shift is included. The results are statistically significant only when the breakdate is 1974. These results imply a permanent slowdown in the rate of productivity growth beginning in 1974. Prior to 1974, TFP grew at an annual rate of 1.71 percent, but this rate of growth slowed to 1.56 percent after 1974. Moreover, this slower rate of productivity growth persisted after the intercept shift in 1985, implying a slower rate of growth but from a higher absolute level of productivity. This outcome is depicted graphically in figure 2.

***Breakdates and Exogenous Shocks***

Those statistical tests that allow the appropriate characterization of trends in TFP also identify structural breaks in the time series. This section concludes with a discussion of shocks that were exogenous to the sector that provide anecdotal evidence in support of the choice of breakdates, as well as suggesting possible directions for future research. The break in trend in 1974 was coincident with the 1973 oil embargo that resulted in a rapid and unexpected rise in energy prices. It seems plausible to argue that the rise in energy prices also accelerated the rate of obsolescence of the stock of physical capital in agriculture.[9] But since conventional measures of capital stock do not capture changes in the rate of obsolescence, conventional measures of productivity growth will fail to identify this effect. Instead, it will be suppressed into a residual estimate of productivity change.

If there is a significant link between energy prices and obsolescence, it should be revealed in the price of used assets. Indeed, if rising energy costs did render older, energy-inefficient capital obsolete, the price of used assets would decline.[10] An examination of the prices

[9] Empirical support for the obsolescence hypothesis can be found in Ball et al. (2001).
[10] Note that a decline in an asset's value is the definition of obsolescence.

of used assets before and after the energy price shock should reveal the sign and magnitude of the obsolescence effect. This is a subject for future research.

The intercept shift in 1985 was more likely linked to macroeconomic developments and to changes in farm policy.[11] During the 1970s, monetary policy largely accommodated inflation, ostensibly to erode real wages and lower unemployment. But this changed in the early 1980s. The adoption of a restrictive monetary policy by the Federal Reserve pushed up interest rates sharply. The dollar appreciated on foreign exchange markets, and world export prices started to fall. This had the short-run effect of undermining United States competitiveness.

Agriculture's competitive position was further eroded when the Agricultural and Food Act of 1981 was enacted. In the 1981 Act, the tie between target prices and inflation was broken, and specific levels of price support were mandated for each year between 1982 and 1985. More important, the levels of support were determined based on the expectation of continued high rates of inflation. As a result, prices in the United States remained at or near record levels long after the momentum of inflation was broken.

Equally significant was the accumulation of stocks in the Farmer Owned Reserve (FOR) program. The FOR provided storage subsidies to farmers to store grain under loan for 3 to 5 years.[12] Producers agreed not to market grain from the Reserve for this period unless the price reached a specified release level. The longer duration of the FOR program, combined with high release prices needed for grain to exit the Reserve, effectively isolated a large amount of grain from the market. By 1982, corn held in the FOR rose to almost 2 billion bushels, or 25 percent of annual use. Wheat in the Reserve exceeded 1 billion bushels or some 40 percent of annual use.

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[11] Costello (1993) found that a substantial fraction of annual changes in productivity can be attributed to macroeconomic factors that are common across industries.
[12] The marketing loan was a 9-month program.

Thus high price supports along with the isolation of FOR stocks from the market resulted in a significant policy effect on prices and production.

Changes to the price support program in the 1985 farm legislation reduced the impact of that program on price determination. These changes were part of a more market oriented farm policy. Significant were changes in the level of support; loan rates were adjusted downward. The volume of grain permitted into the FOR was sharply curtailed. And a new policy instrument, generic certificates, made grain in the Reserve available to the market by allowing access to that grain before its FOR contract expired.[13] As a consequence, since 1985, commodity prices have largely been based on market supply and demand with reduced influence of government price support and commodity stockholding programs.

## 5. Concluding Remarks

A number of recent studies have concluded that productivity growth in the U.S agricultural sector has slowed (Alston and Pardey, 2009; James et al., 2009). In arriving at this conclusion, the authors partitioned the time series into two sub-periods and compared the resulting mean rates of growth. A more formal approach, attributed to Chow (1960), involves estimating the regression parameters for each sub-period and testing the equality of the two sets of parameters using an $F$-test. An important limitation of both approaches is that the breakdate must be known a priori. The researcher must either pick an arbitrary candidate breakdate or pick a breakdate based on some known feature of the data. As Hansen (2001) points out, the test results can be ‘uninformative’ because they can miss the true breakdate, or can be ‘misleading’ because the

[13] Generic certificates were dollar denominated negotiable certificates that were issued by the USDA in lieu of cash payments to program participants. Generic certificates did not specify any particular commodity. They could be used to acquire stocks held as collateral on government loans (marketing loans or FOR loans) or stocks owned by the Commodity Credit Corporation.

breakdate is endogenous and the test can indicate a break when none in fact exists. We test for a productivity slowdown using techniques that allow for unknown structural breaks jointly with a possible unit root that can have either or both stochastic and deterministic components.

Our empirical results show that two types of structural change have occurred over the postwar period, and both have influenced the long term path of productivity growth in the agricultural sector. First, we observe a break in trend in 1974. Prior to 1974, productivity grew at an average annual rate of 1.71 percent, but this rate of growth slowed to 1.56 percent per year after 1974. The break in trend was coincident with the 1973 oil embargo that resulted in a rapid and unexpected rise in energy prices. We find it plausible that the rise in energy prices accelerated the rate of obsolescence of the stock of physical capital in agriculture. But since the measure of capital stock does not capture changes in the rate of obsolescence, conventional measures of productivity growth will fail to identify this effect. Instead, it will be suppressed into a residual estimate of productivity change.

A different type of structural change occurred in 1985. Following a significant liberalization of farm policy in that year, we observed a one-time upward shift in the level of productivity. The slower rate of productivity growth persisted after the intercept shift in 1985, implying a slower rate of growth but from a higher absolute level of productivity.

## References


Alston, Julian M., Jason M. Beddow and Philip G. Pardey (2009) "Agricultural Research, Productivity, and Food Prices in the Long Run," *Science* 325: 1209-10, September 4.

Alston, Julian M. and Philip G. Pardey, 2009, "Theme Overview: Agricultural Productivity and Global Food Security in the Long Run." *Choices* 24(4) 4th Quarter.

Alston, Julian M., Matthew A. Andersen, Jennifer S. James, Philip G. Pardey, 2010, *Persistence Pays: U.S. Agricultural Productivity Growth and the Benefits from Public R&D Spending*, Springer.

Bai, Jushan and Pierre Perron, 1998, "Estimating and Testing Linear Models with Multiple Structural Changes," *Econometrica*, 66 (1), pp. 47–78.

Ball, V. Eldon, Jean-Christophe Bureau, Jean-Pierre Butault, and Richard Nehring, 2001, "Levels of Farm Sector Productivity: An International Comparison," *Journal of Productivity Analysis* 15:5-29.

Ball, V. Eldon, Jean-Pierre Butault, Carlos San Juan, and Ricardo Mora. 2010. "Productivity and International Competitiveness of Agriculture in the European Union and the United States," *Agricultural Economics* 41 (6): 611-627.

Ball, V. Eldon, Frank Gollop, Alison Kelly-Hawke, 1999. "Patterns of Productivity Growth in the U.S. Farm Sector: Linking State and Aggregate Models," *American Journal of Agricultural Economics* 81 (6): 164-179.

Chow, Gregory C. (1960). "Tests of Equality Between Sets of Coefficients in Two Linear Regressions." *Econometrica* 28 (3):591-605.

Clemente, Jesus, Antonio Montanes, Marcelo Reyes, 1998. "Testing for a Unit Root in Variables with a Double Change in the Mean", *Economics Letters* 59:175-182.

Costello, Donna M. 1993. "A Cross-Country, Cross-Industry Comparison of Productivity Growth." *Journal of Political Economy* 101 (2): 207-222.

Dickey, David A., and Wayne A. Fuller. 1981. "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root." *Econometrica* 49, no. 4: 1057-1072.

Dicky, D. and W.A. Fuller, (1979), "Distribution of the estimates for autoregressive time series with a unit root, *Journal of the American Statistical Association*, 74, 427-431.

Elliot, Graham, Ulrich K. Muller, 2006. "Efficient Tests for General Persistent Time Variation in Regression Coefficients", *Review of Economic Studies*, 73(4): 907-40.

Fuglie Keith O., 2008. "Is a slowdown in agricultural productivity growth contributing to the rise in commodity prices?" *Agricultural Economics* 39 supplement 431–441.

Hansen, Bruce E. 2001. "The New Econometrics of Structural Change: Dating breaks in U.S. Labor Productivity." *Journal of Economic Perspectives* 15(4): 117-28.

James, Jennifer S., Julian M. Alston, Philip G. Pardey, and Matthew A. Andersen (2009), "Structural changes in U.S. agricultural production and productivity", *Choices* 24 (4)

Jorgenson, D., and Z. Griliches, 1967. "The Explanation of Productivity Change," *Review of Economic Studies* 34: 249-83.

Kwiatkowski, Denis, Peter C.B. Phillips, Peter Schmidt and Yongcheol Shin, 1992. "Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root", *Journal of Econometrics* 54: 159-178.

Lee, Junsoo, and Mark C. Strazicich. 2001. "Break Point Estimation and Spurious Rejections with Endogenous Unit Root Tests." *Oxford Bulletin of Economics and Statistics* 63, no. 5: 535-558.

Perron, Pierre, 1989, "The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis,"*Econometrica*, 57 (6), pp. 1361–1401.

Phillips, Peter C.B., 2010. "The Mysteries of Trend." *Cowles Foundation Discussion Paper* No. 1771 (September). Accessed Oct. 7, 2010 at http://ssrn.com/Abstract=1676216

Phillips, Peter C.B., 1998. "New Tools for Understanding Spurious Regressions," *Econometrica*, 66:1299-1326.

Phillips, Peter C.B., 1986. "Understanding Spurious Regressions in Econometrics," *J. of Econometrics*, 33:311-340.

Phillips, P.C.B. and P. Perron, 1988. "Testing for a Unit Root in Time Series Regression", *Biometrika*, 75(2), 335-346.

Schmidt, P. and P. Phillips, 1992, "LM Tests for a Unit Root in the Presence of Deterministic Trends." *Oxford Bulletin of Economics and Statistics* 54, 257-287.

Thirtle, C, Lin, L, Holding, J and Jenkins, L, 2004, Explaining the Decline in UK Agricultural Productivity Growth, *Journal of Agricultural Economics*, 55 (2), 343-366, July.

Wang, Sun Ling, D. Schimmelpfennig, and Keith O. Fuglie. 2012. "Is Agricultural Productivity Growth Slowing in Western Europe?" in *Productivity Growth in Agriculture: An International Perspective*. Fuglie, Wang and Ball, editors, Oxfordshire, UK: CAB International (forthcoming).

Zivot, Eric, and Donald W. K. Andrews. 1992. "Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit-Root Hypothesis." *Journal of Business And Economic Statistics* 10, no. 3: 251-270.

Table 1. Average annual rates of growth (percent), 1948-2009 (2005=1)

| Period | Total output | Total farm input | Capital | Labor | Intermediate inputs | | | | | TFP |
|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | All | Farm origin | Energy | Agricultural chemicals | Purchased services | |
| 1948-2009 | 1.63 | 0.11 | -0.21 | -2.51 | 1.43 | 1.15 | 0.85 | 2.54 | 1.15 | 1.52 |
| 1948-1953 | 1.18 | 1.34 | 1.75 | -3.34 | 3.72 | 2.23 | 4.61 | 2.87 | 2.40 | -0.16 |
| 1953-1957 | 0.96 | 0.28 | -0.10 | -4.58 | 2.86 | 3.73 | 0.15 | 1.35 | 1.88 | 0.68 |
| 1957-1960 | 4.03 | 0.50 | -0.43 | -3.74 | 2.95 | 2.58 | 0.06 | 5.95 | 5.42 | 3.53 |
| 1960-1966 | 1.21 | 0.05 | 0.05 | -3.75 | 1.72 | 1.73 | 1.65 | 5.54 | -0.67 | 1.16 |
| 1966-1969 | 2.24 | -0.08 | 0.34 | -2.78 | 0.88 | 2.13 | 0.43 | -2.99 | -0.75 | 2.32 |
| 1969-1973 | 2.65 | 0.46 | -0.44 | -1.84 | 2.02 | 1.70 | -0.42 | 8.22 | 0.40 | 2.19 |
| 1973-1979 | 2.26 | 1.64 | 1.04 | -1.06 | 2.76 | 1.83 | 4.09 | 3.29 | 4.84 | 0.62 |
| 1979-1981 | 1.53 | -1.85 | 0.39 | -1.39 | -3.05 | -2.68 | -3.35 | 2.54 | -7.50 | 3.39 |
| 1981-1990 | 0.96 | -1.22 | -2.16 | -2.79 | -0.13 | -0.19 | -1.59 | -0.75 | 0.39 | 2.19 |
| 1990-2000 | 1.84 | 0.31 | -0.75 | -1.63 | 1.64 | 1.20 | 0.96 | 2.85 | 2.38 | 1.53 |
| 2000-2007 | 0.77 | 0.14 | -0.16 | -1.56 | 0.88 | 0.48 | -0.74 | 1.55 | 1.18 | 0.63 |
| 2007-2009 | 1.88 | -1.80 | 0.88 | -3.69 | -2.41 | -3.27 | 5.24 | 1.32 | -5.19 | 3.68 |

Note: The subperiods are measured from cyclical peak to peak in aggregate economic activity as defined by the National Bureau of Economic Research (see http://www.nber.org/cycles.html).

**Table 2 Elliott-Muller qLL test results**

| Breaks | qLL statistics[1] | |
|---|---|---|
| without break | -17.41 | ** |
| with intercept break at 1985 only | -16.32 | ** |
| with trend break at 1985 only | -22.11 | *** |
| with intercept and trend break at 1985 | -13.44 | * |
| with intercept break at 1985 and trend break at 1974 | -9.05 | |

Note 1: The null hypothesis is coefficients are fixed over the sample period
Note 2: '*' indicates significant at 5% level; '**' indicates significant at 10% level, '***' indicates significant at 1% level.

**Table 3 Results of unit root tests**

| Tests[1] | Test result |
|---|---|
| **Unit root test without break** | |
| ADF unit root test | |
| without trend | nonstationary |
| with trend | stationary |
| KPSS unit root test | |
| without trend | nonstationary |
| with trend | stationary |
| PP unit root test | |
| without trend | nonstationary |
| with trend | stationary |
| **Unit root test with break** | |
| CMR unit root test | |
| AO1--with one intercept shift | nonstationary |
| AO2--with double intercept shifts | nonstationary |
| IO1--with one intercept shift | nonstationary |
| IO2--with double intercept shifts | nonstationary |
| Z-Andrews unit root test with trend | |
| both trend and intercept shift with break at 1985 | stationary |

Note1: the significance level used for these tests is 5%.

**Table 4. Results of OLS regression model estimation**

| Models | Model A[1] coefficients | t statistics[4] | | Model B[2] coefficients | t statistics | | Model C[3] coefficients | t statistics | |
|---|---|---|---|---|---|---|---|---|---|
| Intercept | 3.6528 | 270.00 | *** | 3.6461 | 325.62 | *** | 3.6317 | 307.50 | *** |
| t | 0.0153 | 25.63 | *** | 0.0156 | 30.88 | *** | 0.0171 | 22.30 | *** |
| d1985 | | | | 0.0920 | 2.02 | ** | 0.0718 | 4.68 | *** |
| tD1985 | 0.0013 | 3.11 | *** | -0.0007 | -0.68 | | | | |
| tD1974 | | | | | | | -0.0015 | -2.56 | ** |
| D-W statistics | 2.1135 | | | 2.2666 | | | 1.9846 | | |
| Pr < D-W | 0.5732 | | | 0.7437 | | | 0.3224 | | |
| Pr > D-W | 0.4268 | | | 0.2563 | | | 0.6776 | | |
| $R^2$ | 0.9877 | | | 0.9897 | | | 0.9965 | | |
| adjusted $R^2$ | 0.9873 | | | 0.9892 | | | 0.9917 | | |

Note 1: Model with trend break at 1985 only
Note 2: Model with intercept shift and trend break at 1985
Note 3: Model with intercept shift at 1985 and trend break at 1974
Note 4: '***' indicates significant at 1% level; '**' indicates significant at 5% level; ''*' indicates significant at 10% level.

Table 5: Test Results of Productivity Slowdown

| | Annual productivity growth rate (%) | | Difference | |
|---|---|---|---|---|
| Models with alternative breaks | First period (A) | Second period (B) | (B-A) | |
| A. With trend break at 1985 only | 1.54 | 1.67 | 0.12 | *** |
| B. With both intercept and trend breaks at 1985 | 1.57 | 1.50 | -0.07 | |
| C. With intercept break at 1985 and trend break at 1974 | 1.71 | 1.57 | -0.14 | ** |

Note 1: '***' indicates significant at 1% level; '**' indicates significant at 5% level

Figure 1. Least Squares Breakpoint Estimation

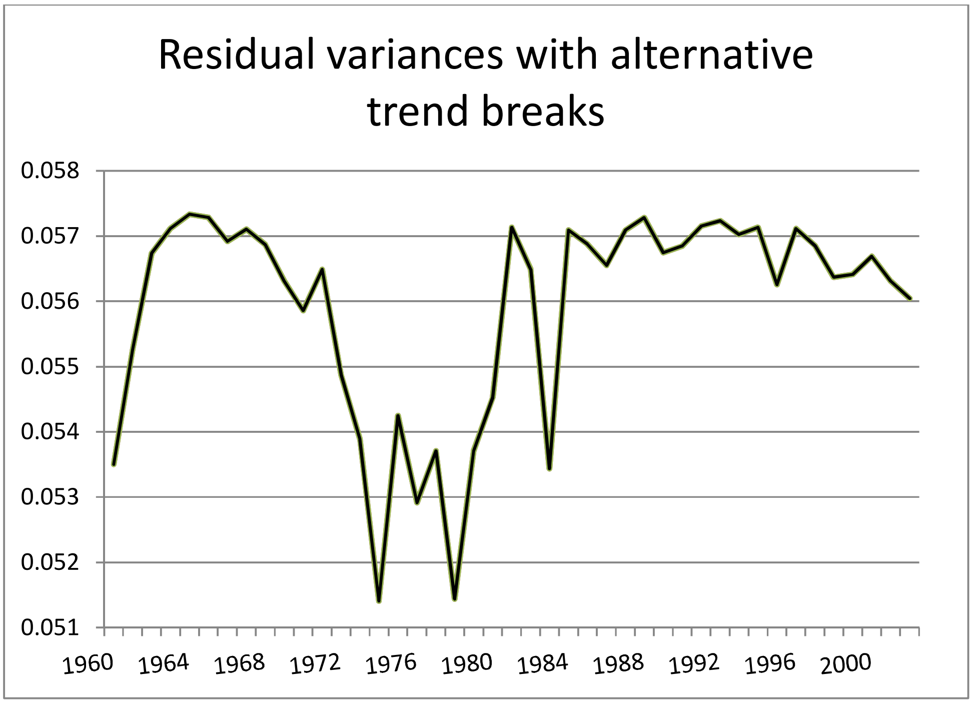
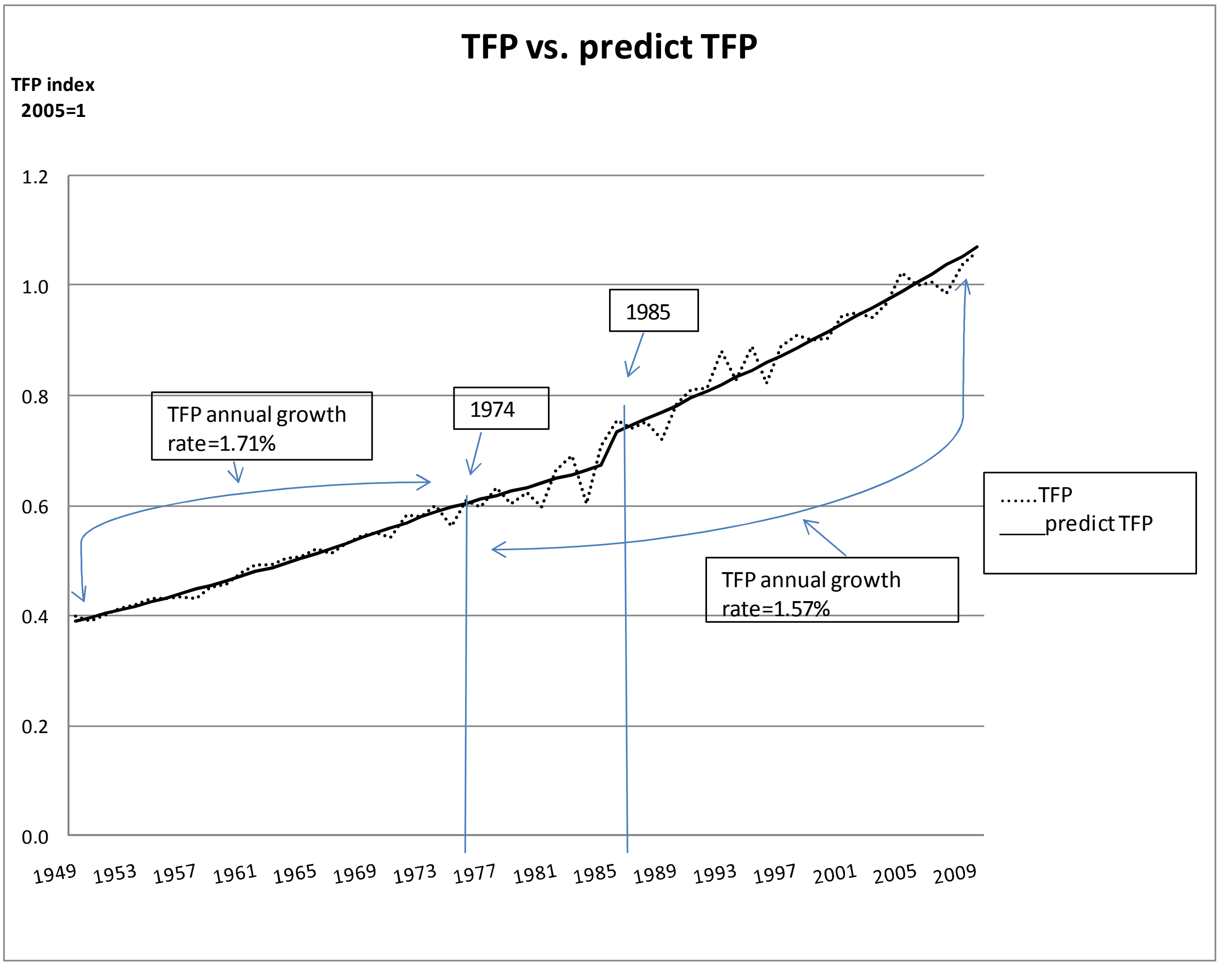


Figure 2


TFP vs. predict TFP
TFP index
2005=1
1.2
1.0
0.8
0.6
0.4
0.2
0.0
1949
1953
1957
1961
1965
1969
1973
1977
1981
1985
1989
1993
1997
2001
2005
2009
TFP annual growth rate=1.71%
1974
1985
TFP annual growth rate=1.57%
......TFP
_____predict TFP