Abstract

We explore whether the U.S. exchange rate could have an influence on cash rental rates for farmland in five U.S. cornbelt states. We find that farmland cash rents have a fairly strong, positive correlation with the U.S. dollar, in terms of its real value relative to major agricultural trading partners. One explanation for the correlation is that a strong dollar lowers the price of key inputs and thus has purchasing power effects. A strong dollar may therefore be associated with higher net returns, and the payment of higher cash rents by farmers. We find support for this hypothesis through a series of econometric models.

JEL classifications: E31, F31, Q15, Q24

Keywords: Cash rents; Exchange rates; Imported inputs; Inflation

1. Introduction

We explore whether the U.S. exchange rate could have an influence on cash rental rates for farmland in five cornbelt states. The exchange rate is of interest because it is well known to affect many outcomes in agriculture (Batten and Belongia, 1986; Pick, 1990; Schuh, 1974). In the case of U.S. Midwestern agriculture, the dollar is a key determinant of foreign demand for export commodities such as corn and soybeans. Foreign demand for U.S. products increases as the dollar weakens relative to the currencies of foreign importers. This may increase the net returns from farming, and thus farmers’ demand for additional cash-rented acres. It follows that cash rental rates for farmland may have an inverse relation with the dollar.

A weaker dollar may not translate into higher net returns for farmers, however, if it causes their costs to go up. For example, a weak dollar may raise the price of imported fertilizers and agricultural chemicals. A weak dollar is also associated with a high price of oil, particularly in the short run (Amano and van Norden, 1998; Chen and Chen, 2007; Krugman, 1983). This is important because oil and related energy-based resources, such as natural gas, tend to underpin many of the inputs used in agriculture. For these reasons, a depreciating dollar may therefore reduce the net returns from renting additional acres, and reduce the demand for cash-rented farmland, and hence cash rents. To the extent that this purchasing power argument is important, it means that cash rental rates for farmland may have a positive relationship with the dollar. Based on the two above arguments, the relationship between the dollar and cash rental rates is an empirical issue.

Cash rental agreements are an increasingly common way of leasing farmland in the U.S. cornbelt. In Iowa, for example, 77% of leased farmland is done through a cash lease. This type of arrangement has become more popular than the crop share lease in which a landlord and tenant split the costs and revenues of farming. More farmland in Iowa is now cash rented (46%) than owner operated (40%) (Iowa State University Extension, 2008). The situation in other cornbelt states is similar (Ryan et al., 2001).

When one plots deflated cash rents for U.S. cornbelt states since 1975, they often trend monotonically for a number of years in one direction before changing course and trending monotonically in the other direction. The U.S. dollar exhibits similar patterns with respect to its value against other major currencies over time. Since the 1970s, there are clearly demarcated periods...
in which the U.S. dollar is monotonically increasing and then monotonically decreasing. This study asks whether these patterns have any relation, and if so, whether the relation is causal or driven by a common third process.

Looking at plots of data, we find a rough positive correlation between the real exchange rate (strength of the dollar) and real cash rents. We believe this to be an important stylized fact that has not been documented in the literature. According to the simple theory outlined above, it suggests that a weaker dollar may be reducing the purchasing power of farm operators. This factor has never been considered in the literature that examines the evolution of farmland rental rates and their determinants.

To get at this issue more systematically, we develop theoretical and econometric models to quantify the relationship between cash rents and exchange rates over time, along with other potentially important factors. Although we do not imagine that the relationship between cash rents and exchange rates over time, along with other factors of production, is quantified in the literature that examines the dollar’s effect on labor usage in the manufacturing sector. Revenga (1992), for example, finds that the manufacturing sector has addressed many topics, including exchange rate pass-through, pricing-to-market, and the effects of exchange rate uncertainty (e.g., Carter and Pick, 1989; Cho et al., 2002; Kandilov, 2008; Pick, 1990). Our contribution is to show that exchange rates also have a meaningful relationship with cash rental rates. We argue that exchange rates may matter as much on the input-cost side as on the foreign demand side. To our knowledge, this point has not been considered in the literature.

The remainder of the article is organized as follows. In the following section, we introduce our data and present a stylized fact. In the subsequent section, we develop a model that shows two channels by which the exchange rate could influence cash rental rates in farmland. We then form an econometric model from this theoretical model to test the hypotheses outlined above. We wrap up with a discussion of results before drawing our conclusions in the final section.

2. First look at the data

The data are collected from the U.S. Economic Research Service (ERS) and the National Agricultural Statistics Service (NASS) and are described in detail at the end of this article; the data themselves are available alongside the online version of the article at the publisher’s website. We focus on five U.S. corn-belt states over the 1975–2005 period: Illinois, Indiana, Iowa, Missouri, and Ohio. Since the data reflect transactions among multiple countries and currencies, we must account for multiple bilateral exchange rates at the same time. We use the ERS “U.S. markets agricultural trade” exchange rate index, which
measures the strength of the U.S. dollar, in real terms, relative to key importers of major U.S. agricultural commodities. We report some descriptive statistics regarding this series in Table 1. Looking near the top of this table, the "agricultural trade" exchange rate index has a mean of 89.9 deflated units of foreign currency per dollar. This appears to straddle ERS exchange rate indexes for corn and soybeans, which have means of 92.5 and 86.4, respectively (Table 1). Descriptive statistics regarding deflated cash rents by state are also reported in Table 1. The mean values range from $57 per acre in Missouri to $102.7 per acre in Illinois.

We plot the exchange rate index along with deflated cash rents for the five states over time in Fig. 1. There seems to be a roughly positive correlation. The correlation averages approximately 0.3 for the five states. The positive nature of this correlation can be appreciated when looking at individual time periods. Consider the 1979–1985 period, for example. During this time, the U.S. real exchange rate rose 43% against major agricultural trading partners’ currencies. During the same period, real cash rents rose by 16% in Illinois and by 9% in both Iowa and Indiana. There were slight declines after 1982, and in Ohio this was enough to make real 1985 cash rents slightly lower than in 1979 (76.3 vs. 77.1). However, the fact that there was a rise at all is somewhat surprising given that this period is well known to have had large increases in real U.S. interest rates and declines in agricultural output prices and export volume (ERS/NASS data).

Let’s next look at the 1985–1988 period (Fig. 1). During these years, the real value of the dollar fell 17%. Similarly, real cash rents fell in all five states: by 27% in Illinois, 27% in Indiana, 24% in Iowa, 12% in Missouri, and 18% in Ohio. These falls occurred despite an escalation of government supports for cornbelt farmers during this period. According to ERS/NASS data, real government payments to agriculture, per acre, rose by 142%–158% in the five states during 1985–1988 (see also Table 1 for descriptive statistics, and the data definitions at the end of this article).

Another example of the positive correlation is the 1996–2002 period, when the U.S. real exchange rate rose 22% against major trading partners’ currencies (Fig. 1). During this time, real cash rents rose by 42% in Illinois, 32% in Indiana, 41% in Iowa, 73% in Missouri, and 34% in Ohio.

### 3. Theoretical and econometric models

In cornbelt states, cash-rented farmland is generally used to produce corn (maize) and soybeans. These crops are sometimes planted in rotation due to agronomic benefits. The market for cash-rented farmland is conducted through tenant agreements that are typically renewable and renegotiated annually. For a landlord, the opportunity cost of renting a piece of ground to one farmer is likely to be the value of renting it to yet another farmer. According to Ryan et al. (2001), approximately 93% of cornbelt producers farm a mix of own and rented farmland. We assume that fixed costs, such as machinery, are covered primarily by owned land, and that the demand for additional cash-rented acres can be represented by the marginal revenue associated with farming an additional acre.

We do not propose that farmers and landowners actively consider the exchange rate when negotiating a cash rent. Rather, they will consider the expected net returns from farming additional acres. Net returns are determined in part on expected output prices (e.g., the futures price for November delivery), which are influenced by the strength of the dollar. In addition, net returns are influenced by expected input prices, which are also influenced by the strength of the dollar. Many inputs (fuel, fertilizer, chemicals) are imported or their costs linked in some way to the price of oil. In both cases, input prices are likely to be inversely related to the strength of the dollar, at least in the short run.

To state these ideas more formally, we introduce a simple model of demand for cash-rented farmland. We imagine a representative producer operating in a perfectly competitive environment with constant returns to scale production. Production...
is given by a Cobb-Douglas form:

\[ q = L^{1-\gamma}Z^{\gamma}, \quad (1) \]

where \( q \) is an index of exportable outputs, \( L \) is cash-rented land, and \( Z \) is an importable input whose price is sensitive to exchange rates. We employ this functional form because the cost share of importable inputs is simply \( \gamma \), which will be useful for our purposes below. While \( Z \) may not always be imported, we assume that its price may be affected by exchange rates. The farmer maximizes profit given by:

\[ \max_{L,Z} \pi = pL^{1-\gamma}Z^{\gamma} - wL - sZ, \quad (2) \]

where \( p \) is an output price index, \( w \) is the rental rate of cash-rented land, and \( s \) is the rental rate on \( Z \), respectively. The farmer is a price taker in input and output markets. However, output prices are a function of the strength of the U.S. dollar in units of foreign currency per dollar, denoted \( e \). Therefore, we have that \( p = p(e) \). We assume the law of one price, which means that if the price received in foreign markets is \( p^* \) (denominated, say, in Euros), then we have that \( p = ep^* \). The input price \( s \) is also a function of \( e \), implying that \( s = s(e) \). We assume full exchange rate pass-through for our theoretical model. In the econometric model, however, the responsiveness of output and input prices to exchange rate changes will be determined empirically.

Maximization of (2) gives rise to the following first-order conditions:

\[ \frac{\partial \pi}{\partial L} = p(1-\gamma)L^{-\gamma}Z^{\gamma} - w = 0, \quad (3) \]

\[ \frac{\partial \pi}{\partial Z} = p\gamma L^{1-\gamma}Z^{\gamma-1} - s = 0. \quad (4) \]

These equations show that \( L \)'s marginal revenue product for land equals its rental rate (\( w \)), and that \( Z \)'s marginal revenue product equals its rental rate (\( s \)). Division of (3) by (4) further shows that the rate of technical substitution of factors equals the ratio of input prices:

\[ \frac{(1-\gamma)Z}{\gamma L} = \frac{w}{s}. \quad (5) \]

We can then solve (5) for \( Z \):

\[ Z = s^{-1}w(1-\gamma)^{-1}L. \quad (6) \]

The land rental rate can be interpreted as a function of output and input prices if we substitute (6) into (3):

\[ w(e) = p(e)^{1/(1-\gamma)}s(e)^{-\gamma/(1-\gamma)}(1-\gamma)^{\gamma/(1-\gamma)}, \quad (7) \]

where the inclusion of \( e \) signifies that a variable is dependent on the value of the dollar. We now take a partial derivative with
respect to \( e \) to determine the effect of increases in the strength of the dollar on equilibrium land rents:

\[
\frac{\partial w}{\partial e} = \frac{\partial p}{\partial e} \left( \frac{p G}{s} \right)^{\gamma/(1-\gamma)} - \frac{\partial s}{\partial e} \left( \frac{p G}{s} \right)^{1/(1-\gamma)}.
\]  

(8)

There are several important things to observe in (8). First, the sign on this derivative is ambiguous because a strong dollar will tend to decrease output prices \( (\partial p / \partial e < 0) \) yet also decrease input prices \( (\partial s / \partial e < 0) \). These might be called an output price effect and an input price effect, respectively. Cash rents \( (w) \) will fall to the extent that the former matters and rise to the extent that the latter matters. Second, the size of each effect is affected by the cost share of importable inputs \( (\gamma) \). As \( \gamma \) increases, \( \gamma^{1/(1-\gamma)} \) rises relative to \( \gamma^{\gamma/(1-\gamma)} \), and more weight is placed on the input price effect. Cash rents are more sensitive to exchange rates when importable inputs have a larger cost share, relative to nontradable land, in the production of exportable outputs.

To test this hypothesis, we need an empirical model. We build from Eq. (7), which has observed cash rents on the left-hand side. We pursue a single-equation approach, taking the supply of cash-rented farmland as fixed. For this reason, we do not need to include \( L \) on the right-hand side of the equation or instrument for it in any way. The fixed \( L \) assumption is reasonable because farmland in the cornbelt region has few alternative uses and the market is reasonably mature. While some land in a given year may be converted to urban area or is enrolled in the Conservation Reserve Program, for example, these activities constitute only a very small percentage of overall cornbelt acreage. There is much variation in cash rents \( (w) \) relative to cash-rented acreage \( (L) \) owing to variation in the profitability of farming additional cash-rented acres.

Since output and input prices are functions of exchange rates, with some kind of net effect as represented in (8), we first include \( e \) on the right-hand side to represent this net effect. Identification is aided by the fact that it is only possible for the exchange rate \( (e) \) to affect cash rents \( (w) \), not the other way around. Bilateral exchange rates are determined at the level of entire economies, and a particular sector, such as agriculture, plays an inconsequential role in this process.

The estimate of the coefficient on \( e \) will tell us three things: whether exchange rates have an effect, the size of the effect, and the particular channel by which they have an effect. We can also learn more by including the cost share of importable inputs \( (\gamma) \) on the right-hand side of our empirical model. We include it as an interaction term \( (\gamma e) \) to capture the fact that producers in a situation with high dependence on imported inputs will benefit relatively more as the dollar appreciates. The motivation for this approach comes from Campa and Goldberg (2001).

Likewise, we include an interaction term involving the share of export sales in agricultural production \( (\chi) \). This is denoted \( \chi e \), and acknowledges that if a region’s particular output mix has a high export orientation, then these producers will be hurt more as the dollar appreciates. Similarly, we also include an interaction term involving the share of imports penetrating the domestic market \( (M) \). The idea is that producers in time periods with high import penetration in the domestic market will be hurt relatively more as the dollar appreciates, since local buyers favor the foreign product; the domestic product is disadvantaged, and therefore cash rents will fall.

In addition to the above, net returns from farming additional acres may be influenced by real government payments \( (G) \), the real interest rate \( (INT) \), or real oil prices \( (OIL) \). Since we have data on five different states, we might usefully include fixed effects for these regions \( (D) \).

Since we will be using time series data, an issue of immediate concern is the possibility of autocorrelation. We test a hypothesis of first-order autocorrelation and find that it is not rejected by a Berenblut and Webb (1973) test at the 5% level of significance. We therefore first difference the variables in our preferred specifications. Given the time-series cross-section nature of our data, we additionally carry out a Hausman (1978) test. We reject a null hypothesis that a random effects estimator is consistent and efficient, and therefore adopt a Least Squares Dummy Variable (LSDV) approach for most of the specifications.

We will also need to consider the potential for endogeneity in Eq. (9). As discussed above, the supply of cash rentable cropland in the five U.S. states is effectively fixed, so we do not worry about simultaneous equations bias in this instance. However, there may be other forms of endogeneity with respect to the interaction terms that we will investigate in a special version of the model, below.

With this background, we now turn to our explicit econometric formulation. We denote real cash rents as \( w_i' \), letting \( i \) index states, and \( t \) index time. We add state-level fixed effects \( D_j \) to the regression specification, where \( D_j \) is equal to one when \( j = i \), otherwise zero. In implementing the model, we let \( \Delta \) denote a change from period \( t-1 \) to \( t \). The primary specification that we estimate is:

\[
\Delta w_i' = \sum_{j=1}^{5} \beta_j D_j + \varepsilon'_i,
\]  

(9)

where \( \alpha_1, \ldots, \alpha_6, \beta_1, \ldots, \beta_5 \) are parameters to be estimated, and \( \varepsilon'_i \) is an error term with conventional properties except as discussed below. Construction of the variables is described at the end of this article, and descriptive statistics are provided in Table 1. Note that the specification is set up with no intercept but with five state dummy variables. This allows for a common time trend across the states.

Since \( e_t \) is measured in units of foreign currency per dollar, \( \Delta e_t > 0 \) signifies an appreciation in the U.S. dollar. The sign and significance of the coefficient on \( \Delta e_t \) is all we need to know to distinguish the output price effect from the input price effect highlighted in Eq. (8). There are three cases. First, if a strong dollar’s beneficial effect on costs (the input price effect) outweighs its disadvantageous effect on revenue (the output price effect), the coefficient is positive. Second, if the output
price effect is stronger than the input price effect, the coefficient is negative. Third, if the coefficient is statistically zero, the two effects exactly offset each other. Alternatively, there may be an additional factor that affects cash rents yet has nothing to do with exchange rates. In either latter case, we would conclude that exchange rates have not caused the variation in cash rents.

In all variants, the coefficient on the exchange rate is positive. The real interest rate induces the substitution of land for other capital inputs, cash rents will be positively affected.

### 4. Regression results

Table 2 reports the results of estimating five variants of Eq. (9). The variants show how the model performs under a range of assumptions about right-hand side variables and the estimation technique. Variant I is the most pared-down version and is estimated with the LSDV approach. Variant II adds the interaction terms to variant I. Variant III further adds the interest rate variable to variant II. Variant IV is estimated using two-stage least squares (2SLS) methods to address the possibility of endogeneity in certain explanatory variables. Variant V examines whether the price of oil should replace the exchange rate the variable of primary focus in the regression.

We start by describing results that hold generally across the variants (special features of variants IV and V will be discussed separately further down). The \( R^2 \) ranges from 0.27 to 0.39 across the five variants. The coefficients in the top row of Table 2 concern the relation between the dollar and cash rents. In all variants, the coefficient on the exchange rate is positive and statistically significant at the 1% or 5% level. It ranges from 13.69 in variant I to 20.49 in variant IV. This means that a strong dollar is associated with higher cash rents. This result

<table>
<thead>
<tr>
<th>Variable</th>
<th>Expected sign</th>
<th>I (LSDV)</th>
<th>II (LSDV)</th>
<th>III (LSDV)</th>
<th>IV (2SLS)</th>
<th>V (LSDV)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Exchange rate ( (\Delta e_t) )</td>
<td>±</td>
<td>13.69**</td>
<td>20.01***</td>
<td>20.30***</td>
<td>20.49***</td>
<td>–</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.05)</td>
<td>(0.01)</td>
<td>(0.01)</td>
<td>(&lt;0.01)</td>
<td>–</td>
</tr>
<tr>
<td>Export share × exchange rate ( (\chi^{1}_t \Delta e_t) )</td>
<td>−</td>
<td>−</td>
<td>(0.56)</td>
<td>(0.53)</td>
<td>(0.39)</td>
<td>(0.89)</td>
</tr>
<tr>
<td>Import competition × exchange rate ( (M_t \Delta e_t) )</td>
<td>−</td>
<td>−</td>
<td>−0.01</td>
<td>0.27</td>
<td>−0.93</td>
<td>0.71</td>
</tr>
<tr>
<td>Imported inputs × exchange rate ( (\gamma^{1}_t \Delta e_t) )</td>
<td>+</td>
<td>+</td>
<td>(&lt;0.01)</td>
<td>(&lt;0.01)</td>
<td>(&lt;0.01)</td>
<td>(0.09)</td>
</tr>
<tr>
<td>Government payments ( (\Delta G^I_t) )</td>
<td>+</td>
<td>+</td>
<td>(&lt;0.01)</td>
<td>(&lt;0.01)</td>
<td>(&lt;0.01)</td>
<td>(&lt;0.01)</td>
</tr>
<tr>
<td>Interest rates ( (\Delta INT_t) )</td>
<td>±</td>
<td>1.77*</td>
<td>1.46</td>
<td>–</td>
<td>–</td>
<td>–</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.09)</td>
<td>(0.17)</td>
<td>–</td>
<td>–</td>
<td>(0.07)</td>
</tr>
<tr>
<td>Oil prices ( (\Delta OIL_t) )</td>
<td>±</td>
<td>–</td>
<td>–</td>
<td>–</td>
<td>–</td>
<td>–0.09</td>
</tr>
<tr>
<td>Illinois</td>
<td>±</td>
<td>0.85</td>
<td>0.59</td>
<td>0.65</td>
<td>0.78</td>
<td></td>
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<tr>
<td></td>
<td></td>
<td>(0.27)</td>
<td>(0.44)</td>
<td>(0.39)</td>
<td>–</td>
<td>(0.31)</td>
</tr>
<tr>
<td>Indiana</td>
<td>±</td>
<td>1.58**</td>
<td>1.37*</td>
<td>1.43*</td>
<td>1.91**</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.04)</td>
<td>(0.07)</td>
<td>(0.06)</td>
<td>–</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Iowa</td>
<td>±</td>
<td>1.79**</td>
<td>1.56**</td>
<td>1.61**</td>
<td>1.65**</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.02)</td>
<td>(0.04)</td>
<td>(0.03)</td>
<td>–</td>
<td>(0.03)</td>
</tr>
<tr>
<td>Missouri</td>
<td>±</td>
<td>1.22</td>
<td>0.95</td>
<td>1.00</td>
<td>1.38*</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.11)</td>
<td>(0.21)</td>
<td>(0.18)</td>
<td>–</td>
<td>(0.07)</td>
</tr>
<tr>
<td>Ohio</td>
<td>±</td>
<td>0.54</td>
<td>0.33</td>
<td>0.40</td>
<td>0.58</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.48)</td>
<td>(0.67)</td>
<td>(0.61)</td>
<td>–</td>
<td>(0.46)</td>
</tr>
<tr>
<td>Number of observations</td>
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<td>150</td>
<td>150</td>
<td>150</td>
<td>150</td>
<td></td>
</tr>
<tr>
<td>( R^2 )</td>
<td>0.27</td>
<td>0.30</td>
<td>0.39</td>
<td>0.35</td>
<td>0.32</td>
<td></td>
</tr>
</tbody>
</table>

*Notes: All variables are deflated. p-value is in parentheses. The asterisks ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.*
corresponds to what we already saw visually in Fig. 1. The results in Table 2 are based on the ERS’s composite agricultural trade exchange rate index. However, results are very similar if we use the ERS index created just for corn, or the ERS index created just for soybeans (Table 1).

Based on our simple theoretical model, the above result suggests that there may be purchasing power benefits of a strengthened dollar. For example, while oil is priced in dollars, it tends to have an inverse relationship with the strength of the U.S. dollar, particularly in the near term. In our data, we find this correlation to be $-0.42$. This type of effect also appears to occur with respect to fertilizer prices. In supplementary regressions (not reported in the table), we find that a $1\%$ appreciation in the U.S. dollar relative to our exchange rate index results in $1.5\%$ decrease in the price of urea and a $0.72\%$ decrease in the price of ammonium nitrate, both of which are widely used in fertilizers as a source of nitrogen (USDA, 2009b). Taken together, these results suggest that a strong dollar reduces costs, inducing farmers to pay higher rents for additional, cash-rented acres.

Results regarding the interaction term coefficients are reported in rows 2–4 of Table 2. The coefficients on the export orientation interaction term ($\chi_{i,t} \Delta e_t$) and the import penetration interaction term ($\gamma_{i,t} \Delta e_t$) are not statistically different from zero in any variant. These aspects of the model, therefore, do not receive support from the data. Should we be surprised by this? The trade literature identifies a number of reasons that dollar depreciations may not have a strong, immediate effect on exports. Exchange rate pass-through may be limited due to the oligopolistic nature of international grain trade, and price transmission may be limited due to policy barriers (Krugman, 1987; Pick and Park, 1991). Prices may also be “sticky” due to menu or catalog pricing. Exports may therefore have a somewhat inelastic relationship with the dollar (Batten and Belongia, 1986).

In separate calculations, however, we find that the share of foreign sales in overall sales ($\chi_{i,t}$) does have a negative correlation with real exchange rates in four of the five states (in the case of Indiana and Ohio, for example, it is $-0.46$ and $-0.49$). Thus, we have some support that a weak dollar leads to greater foreign demand for U.S. crops. However, this effect appears to be moderated by the dollar’s potential effect on input costs. In effect, the input price effect outweighs the output price effect in Eq. (8).

We now turn to the result concerning the cost share of imported inputs, which is calculated using state agricultural balance sheets over time. The coefficient on the imported input cost share interaction term ($\gamma_{i,t} \Delta e_t$) is reported in Table 2, and is positive and statistically significant in all cases. The coefficient ranges from $16.75$ for variant V to $26.63$ for variant II. In the case of variant V, the associated elasticity is $0.98$, which means that a $1\%$ appreciation in the dollar is associated with an increase in cash rents of $0.98\%$, when measured using the average imported input cost share. A positive coefficient implies that states and time periods with higher dependence on imported inputs experience a more of an increase in cash rents as the dollar appreciates. This is consistent with our expectations for this coefficient, and reinforces the above result concerning the coefficient on $\Delta e_t$.

Other aspects of the regression are in line with expectations. The sign of the coefficient on government payments is positive and statistically significant at the $1\%$ level. This is similar to studies such as Goodwin et al. (2003) and Lence and Mishra (2003), even though our representation of government policies is relatively blunt. Variants I, III, and V include real interest rates ($\Delta INT_t$). The coefficient on $\Delta INT_t$ is positive in all three cases, but statistically significant at only the $10\%$ in two cases. Although we will refrain from interpreting this result too literally, it appears that a rise in the real interest rate may tend to induce the substitution of land for other capital inputs, leading to a rise in cash rents. We also tried our regressions using other types of interest rates, such as the prime interest rate. However, these changes make no substantive difference in our results.

We have found that the cost share of imported inputs, interacted with $e$, plays a key role in our results. However, this interaction term is a function of prices, quantities, and total operating costs, and as such may not be independent of the error term $e_t$ in Eq. (9). In this case, the LSDV estimator may no longer be consistent. We therefore carry out Hausman (1978) tests regarding the endogeneity of this variable. The results are somewhat sensitive to how we design the tests. For example, there tends to be endogeneity if we choose a significance level of $10\%$, but not if we choose a significance level of $5\%$. We conclude that there may be some correlation between this variable and the error term. For this reason, we reestimate the model using 2SLS, which forms an instrumental variable from the exogenous variables of the system. We find that use of this alternative estimator changes the results in no consequential way.

We provide a representative example of one of the estimations in Table 2, which we call variant IV. If we make comparisons to the other variants, the results are highly similar. Variant I does not even include the $\gamma$ interaction term yet has no important qualitative difference with variant IV.

We now turn to our final variant. We introduce this to illustrate the possibility that we have placed too much emphasis on the exchange rate in our previous specifications. Given that there exists an inverse relationship between the dollar and the oil price, and that many of the dollars spent in agriculture are in some way tied to oil, perhaps the more relevant relationship is between cash rents and the oil price. As with exchange rates, oil prices are exogenous to this sector, so we need not worry about the endogeneity of this variable. We have explored its potential importance through a series of empirical models involving deflated oil prices. This process did not turn out to be particularly fruitful, however. We find that oil prices are less effective at explaining the variation in cash rents, as the coefficient on this variable tends to be statistically zero in the specifications. A representative result is reported in Table 2 as variant V. We see that the coefficient on oil prices is $-0.09$ with a $p$-value of $0.21$, meaning it is statistically zero. The exchange rate not only appears to perform better, it has the advantage of allowing
for the output price effect in Eq. (8) in addition to the input price effect, although the former does not appear to outweigh the latter.

We can shed further light on our overall results through consideration of bivariate Granger causality tests. This procedure identifies whether one time series helps forecast another (Granger, 1969). A time series variable \(x\) is said to Granger-cause \(y\) if \(x\) provides statistically significant information about future values of \(y\). If the Granger test statistic is greater than the specified critical value, we reject a null hypothesis that \(x\) does not cause \(y\). The test is conducted without reference to the other explanatory variables included above.

We first test whether the dollar Granger-causes cash rents. The test statistic is 4.65 and \(p\)-value is 0.0039, implying that we reject the null hypothesis. This means that the dollar Granger-causes cash rents. If there is a genuine relationship here, however, the converse should not be true. This can be tested by swapping the variables in the test. In this case, the test statistic is 0.57 and \(p\)-value is 0.63. The implication is that the real dollar does Granger-cause real cash rents, but the converse is not true. This makes sense since the value of the dollar is determined at highly aggregated levels involving many sectors of the economy. Demand for cash-rented farmland—a tiny share of the U.S. economic system—should have no discernable influence on exchange rates. Granger causality is, of course, not a foolproof indicator of causality, since, for example, both variables could be driven by a common third process, but with a different lag. However, the results make economic sense, and corroborate the findings of the main econometric models that were described above.

5. Conclusions

Rental rates for U.S. cash-rented farmland have varied extensively over the past several decades. While existing research focuses on the importance of government support for agriculture and interest rates when explaining rents, we argue that the exchange rate is also a potential causal factor. Indeed, it would be surprising if this macroeconomic variable did not have some effect, particularly in the case of U.S. cornbelt agriculture, which is export dependent and makes extensive use of imported inputs. Simple correlations with publicly available data suggest that real cash rents for U.S. cornbelt agriculture are positively correlated with the real exchange rate for agricultural commodities. To explain this stylized fact, we work out an economic model of an agricultural producer who produces an exportable product using cash-rented land and an importable input. The prices of both are sensitive to the strength of the dollar, but in opposing ways. The relative importance of each effect is an empirical issue, and we investigate this using a variety of econometric models and methods.

Our results are consistent with the idea that the purchasing power benefits of a strong dollar reduce input costs and improve the net returns from farming additional acres, thereby increasing the demand for cash-rented acres. The strength of the dollar appears to affect farm input costs more strongly than they do farm output prices.

In conclusion, our results call into question the idea that a weak dollar would be “good” for U.S. cornbelt agriculture, which seems to have some support in the farming community. A weak dollar policy may increase foreign demand for U.S. agricultural products, and weaken U.S. demand for foreign imports, both of which may increase farm profitability and raise the demand for cash-rented acreage. However, a weakened dollar may raise input costs to an extent that the benefits from rising output prices will be overwhelmed.

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Appendix: Dataset construction and variable definitions

Data come mainly from personnel and websites of the USDA ERS, NASS, and Bureau of Labor Statistics (BLS). The data are available alongside the online version of this article for replication and extension of our results.

Cash rents: From NASS, they are collected through the annual June Agriculture Survey. Enumerators contact all agricultural producers operating land within the boundary of a randomly chosen land segment, and record the per acre cash rent paid. State estimates are based upon the total amount of cropland in each state as given by the most recent Census of Agriculture. We adjust USDA cash rents for inflation using a BLS producer price index.

Exchange rates: Calculated by the USDA as part of its Agricultural Exchange Rate Data Set. The index measures the strength of the U.S. dollar, in real terms, relative to key importers of U.S. agricultural commodities. It is calculated by first multiplying the U.S. dollar exchange rate by the ratio of consumer price indexes in the United States and a foreign country. This real rate is then divided by its 2000 exchange rate to form an index. Next, its share of commodity trade is multiplied by each country’s real exchange rate. The final step involves summing all of the weighted rates across countries to get the composite commodity’s indexed exchange rate. The results reported in the article are based on the USDA’s composite “agricultural trade” exchange rate index. Alternatives include a USDA index created just for corn, and another just for soybeans. We find that results based upon these three alternatives are essentially identical. We report results for the composite agricultural trade index because its mean lies between that of the other two (see Table 1).
Furthermore, as a more general index, it seems most representative of the exchange rate faced by a broad set of producers.

Real government payments per acre: We start with the value of total direct government payments, by state and year, as calculated by the ERS. We then adjust for inflation using a BLS producer price index.

Cost share of imported inputs: Calculated using ERS state-level agricultural balance sheets over time. We view fuel and oil as the inputs that are most consistently imported in a major way, and calculate their share of total operating costs.

Interest rates: From the website of U.S. Federal Reserve. We start with the nominal Federal funds effective rate, in percent. We then subtract the percentage change in the Consumer Price Index to determine a real interest rate. We also tried our regressions using the Prime interest rate. However, it makes essentially no difference to the results.

Real oil prices: Measured by the average annual dollar price per barrel of crude petroleum reported by the International Monetary Fund.

Export shares: Value of a state’s agricultural exports divided by the total value of production. Based upon ERS estimates.

Import penetration: Total imports of competing major grains divided by their supply, calculated as U.S. production plus imports. Based upon ERS estimates.

References


