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### **Growth Complementarity Between Agriculture and Industry: Evidence from a Panel of Developing Countries**

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# Growth Complementarity Between Agriculture and Industry: Evidence from a Panel of Developing Countries.

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## Abstract

Using dynamic panel models with data for 62 developing countries, this paper examines whether growth in agriculture elicits growth in manufacturing. For identification, I use population-weighted, average temperature as an instrument for growth in agriculture. I identify large short-run effects: An increase in growth in agriculture by one percentage point is estimated to raise contemporaneous growth in manufacturing by between 0.47 and 0.56 percentage points. The baseline models also imply sizable long-run effects of permanent increases in growth in agriculture. Extensions of the empirical model suggest that growth in agriculture benefits the manufacturing sector by improving its domestic terms of trade, by increasing the share of investment and saving in GDP, and by increasing the capacity to import industrial inputs.

The paper makes two main contributions. First, it joins a growing literature using climate data to identify supply shocks in agriculture, establishing a robust empirical relation between these shocks and growth in manufacturing. Second, it includes a stylized two-sector model to illuminate the macroeconomic channels behind this complementarity. Together, these contributions lend support to the notion that agriculture plays key macroeconomic roles in the industrialization of developing countries by relieving saving, aggregate demand, fiscal, and foreign exchange constraints on the industrial sector.

**Keywords:** Agricultural Productivity, Industrialization, Multisector Growth.

**JEL Classification Codes:** O130, O140, O410.

## 1 Introduction

There is little doubt that the expansion of industrial activities and their ancillary services characterizes sustained episodes of economic growth in developing countries.<sup>1</sup> But the initial stages of industrialization almost invariably impinge on societies where agriculture accounts for a large share of output and employment.

Several recent studies have underscored the role of agriculture in generating favorable initial conditions for modern economic growth. Their dominant theme is the historical pattern of land

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<sup>1</sup>For recent evidence on the relation between structural change and economic development, see Timmer and Vries (2009), Ocampo et al. (2009), and McMillan and Rodrik (2011).

ownership and its lasting influence on the distribution of income and education, on the incidence of social conflict, and on the development of institutions of economic and political governance.<sup>2</sup>

And yet, beyond the lasting political-economic influence of agrarian structure, agriculture also plays macroeconomic roles in industrialization. They include providing saving and foreign exchange to finance capital accumulation, as well as a home market for industry (Johnston and Mellor, 1961). Their fulfillment is a key ingredient of successful industrialization, as recognized by Alice Amsden in regard to Taiwan's post-war experience:

Agriculture managed to produce a food supply sufficient to meet minimum domestic consumption requirements as well as a residual for export. [...] Good rice harvests have been a major factor behind price (and real wage) stability. [...] Agriculture also managed to provide an important source of demand for Taiwan's industrial output, particularly chemicals and tools, and a mass market for consumption goods. [...] In summary, agriculture in Taiwan gave industrial capital a labor force, a surplus, and foreign exchange. (Amsden, 1979, p. 363)

This paper estimates whether growth in agriculture elicits growth in manufacturing, providing reduced-form evidence of macroeconomic linkages between the two sectors. Using average temperature to identify changes in agricultural supply in 62 developing countries, I estimate that a one percentage point increase in growth in agriculture raises contemporaneous growth in manufacturing by between 0.47 and 0.56 percentage points.

As discussed below, annual variation in temperature is best suited to identify short-run effects. Still, the implied long-run multipliers show that if the average country in the sample were to permanently increase growth in agriculture to 4.4%/yr (the average in China during 1961-2006), growth in manufacturing would eventually increase by between 0.95 and 1 percentage points. This effect is substantial, as the sample mean of growth in manufacturing is 4.5%/yr.

Estimating the effect of growth in agriculture using country-level data is challenging for two main reasons. First, countries differ along time-constant dimensions, such as natural conditions, that are correlated with growth in agriculture. To address this problem, I control for country fixed effects, using only relative variation within countries to identify the coefficients.

Second, regressions relating growth in the two sectors are likely to run afoul of bias due to reverse causality and omitted time-varying variables. To address these problems, I control for the previous dynamics of growth in manufacturing, and use a population-weighted measure of average annual temperature (from Dell et al., 2012a) as an instrument for growth in agriculture. I therefore assume that annual variation in average temperature within countries, while exogenously shifting agricultural supply, does not directly affect growth in manufacturing. Section 6 below further discusses the appropriateness of this assumption.

Besides estimating the reduced-form effect of growth in agriculture on growth in manufacturing, I make two additional contributions. First, to illuminate the main findings, I present a concise model of macroeconomic complementarity between the two sectors. The model examines how agriculture can relieve constraints on industrial accumulation that have featured prominently in the growth literature — namely saving, demand, foreign exchange, and fiscal constraints.

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<sup>2</sup>For recent studies linking historical inequality in land ownership to measures of broad-based educational attainment, see Engerman and Sokoloff (2002), Frankema (2009), Wegenast (2009), and Galor et al. (2009). For a study of the link between agrarian structure and urban income inequality, see Oyvatt (2013). Recent studies of the link between patterns of colonization — including policies concerning land ownership and land tenure — and the development of institutions of economic and political governance include Acemoglu et al. (2000, 2002, 2005), Engerman and Sokoloff (2002), and Rodrik et al. (2004). For earlier examples from the Latin American historiography, see Prado Junior (1967), and Furtado (1963, 1976).

Second, I use the same identification strategy to explore a number of proximate indicators of these channels of complementarity. I find that growth in agriculture improves the domestic terms of trade of the non-agricultural sector, increases the share of investment and saving in GDP, increases the capacity to import industrial inputs, and reallocates workers to activities with higher average productivity.

In sum, this paper joins a growing literature using climate data to identify supply shocks in agriculture (see Dell et al., 2013, for a broad review). It is particularly related to papers that have used this identification strategy to establish causal relations between agricultural growth and broader economic outcomes, such as local urban activity (Henderson et al., 2009), patterns of migration (Brückner, 2012), and industrial growth (Shifa, 2014, to whose empirical strategy this paper is closest). It also summarizes key propositions of the historical and theoretical literatures on macroeconomic relations between the two sectors, encapsulating them in a concise, though stylized, two-sector model. Together, these two contributions lend support to the notion that agriculture plays key macroeconomic roles in the industrialization of developing countries by relieving saving, aggregate demand, fiscal, and foreign exchange constraints on the industrial sector. Agricultural development should thus be a key ingredient of industrial development policies.

The paper is organized as follows. Section 2 motivates the question in relation to the existing empirical and theoretical literature. Section 3 describes the dataset and introduces the empirical model. Section 5 presents the main empirical results of the paper: the effect of growth in agriculture on growth in manufacturing using temperature as an instrumental variable. Section 6 examines the effects of controlling for cross-country heterogeneity along several dimensions — such as the share of agriculture in GDP and the degree of openness to international trade —, as well as other robustness checks; Section 7 examines the impact of agricultural growth on potential channels through which it would enhance industrial growth. Section 8 illustrates the main macroeconomic channels of complementarity between the two sectors by means of a stylized model. Section 9 summarizes the main findings and implications of the paper, and an Appendix provides model proofs and detailed variable definitions.

## 2 Agricultural Development and Industrialization

Most of the recent empirical literature consists of reduced-form tests of whether output or productivity growth in agriculture bolster their counterparts in other sectors. To address reverse causality and omitted variable bias, most authors have deployed time series techniques or instrumental variables.

A group of studies has used cointegration and error correction models to estimate long-run sectoral balance relations, followed by an examination of sectoral responses to deviations from this equilibrium. Studies of individual countries have yielded mixed results: Gemmell et al. (2000) found that manufacturing output and productivity in Malaysia were exogenous (in the sense of Granger) to increases in their counterparts in agriculture. By contrast, Kanwar (2000), and Chebbi and Lachaal (2007) found that they responded positively in India and Tunisia. A study of panel cointegration using a sample of 85 countries, however, confirmed the finding of positive responses for the majority of countries in the sample (Tiffin and Irz, 2006).

By contrast, this paper follows an alternative, but increasingly common strategy: the identification of exogenous shifts in agricultural value added from variation in climate variables, such as rainfall and temperature (Dell et al., 2013). Some of these studies examine the regional impact of growth in agriculture. For example, Henderson et al. (2009) find that the intensity of city lights at night increases in years of favorable rainfall in adjacent rural areas. Their results, derived from satellite images of 541 cities in 18 African countries, indicate substantial local complementarity between the rural and urban economies.

But macroeconomic complementarity can be better discerned at a higher level of aggregation. To that end, Dell et al. (2012a) build average measures of nationwide temperature and rainfall, using local population as weights. They find that GDP growth declines in poor countries when temperature is higher than the historical average.

Changes in agricultural yields are the channel most likely to explain these reduced-form relations between climate variables and economic growth. Several authors have therefore used country-level climate variables as instruments for output and productivity in agriculture. For example, using precipitation as an instrumental variable (along with international commodity prices), Brückner (2012) finds that lower value added in agriculture leads to distress migration and the expansion of urban informal activities. In turn, using temperature and precipitation as instrumental variables, Shifa (2014) finds that higher growth in agriculture elicits sizable short-run increases in growth in manufacturing in a large sample of countries.

The empirical findings above raise a natural question: what macroeconomic mechanisms explain the observed complementarity between agricultural and industrial development? The answer, formalized in the simple model of section 8, is that agriculture can ease saving, demand, foreign exchange, and fiscal constraints on industrial accumulation. These roles of agriculture have been explored in a number of previous contributions.

Traditional development theory, for example, saw the availability of domestic saving as the main constraint on the rate of capital accumulation. Many authors thus called on agriculture to elicit higher saving from the private non-agricultural sector, in particular industry. This role is predicated on three notions: first, that most saving originates in retained profits and other non-wage incomes; second, that unit costs are a key determinant of real industrial saving in terms of its own output; and, third, that in dual economies agricultural labor productivity is a key determinant of industrial unit costs.

Arthur Lewis laid out the classical model along these lines. By linking money wages in industry to the value of the average product in agriculture, Lewis suggested that agricultural development will raise industrial accumulation if the fall in agriculture's domestic terms of trade dominates the increase in industrial wages in terms of food (Lewis, 1954, p. 173-176). The presence of Engel's Law in final demand — the proposition that the share of primary goods in total expenditure falls as income increases — is crucial to ensure this result (Jorgenson, 1961, 1967).

An exception may arise if the terms of trade are unresponsive to agricultural growth — consider, for example, a small open economy in which both sectors produce tradable goods. In this case, higher labor incomes in agriculture will directly raise unit labor costs in terms of the industrial good (Skott and Larudee, 1998; Bustos et al., 2012). Detailed case studies of intersectoral resource flows in developing countries, however, show that periods of technical dynamism in agriculture often correlated with endogenous declines in the sector's domestic terms of trade (Mellor, 1973; Karshenas, 1995). By contrast, where industrial growth occurred alongside a stagnant agriculture, such as in India during 1950-1965, net resources flowed out of the non-agricultural sector due to adverse movements in the terms of trade, potentially hindering the ability to finance capital accumulation in industry (Mody et al., 1982).<sup>3</sup>

The notion that private saving is the binding constraint on industrial growth has been challenged on three fronts: by those emphasizing fiscal constraints on complementary public investment, by those emphasizing balance-of-payments constraints on the expansion of domestic demand,

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<sup>3</sup>To be sure, agriculture can also contribute saving to non-agriculture directly by accumulating net assets. But if Engel's effects are strong, technical progress in agriculture may reduce the real value of net lending in terms of the industrial good, due to adverse movements in the terms of trade. In other words, a trade-off may exist between agriculture's ability to directly contribute net saving to non-agriculture, and its ability to indirectly elicit higher saving from non-agriculture (see section 8 for a formal example). In a review of episodes of industrial growth in Asia, Karshenas (1995) shows that most growth accelerations were financed with *ex-post* increases in saving from non-agriculture.

and by those emphasizing insufficient domestic demand even in the absence of balance-of-payments constraints.

Public infrastructure and private investment may be bound by direct technical complementarity in industrializing economies. Expected demand may justify private projects, yet investors may fail to undertake them in the absence of complementary public investment in energy or transportation.<sup>4</sup> In addition, infrastructure is often subject to increasing returns to scale and market failures, and it is largely non-tradable. These characteristics hinder the ability of private investors to sustain the required level of infrastructural investment (Skott and Ros, 1997).

Agriculture, as it turns out, has often been a prominent source of fiscal revenues in industrializing economies. For example, the direct taxation of agricultural rents was key for funding infrastructure investment during the first several decades of Japan's industrialization (Ohkawa and Rosovsky, 1960). But due to technical or political constraints, in most post-war episodes of industrialization governments resorted to instruments of indirect taxation, such as trade tariffs and quotas, multiple exchange rates, and domestic marketing boards. As a result, wedges between the actual terms of trade of agriculture and border prices were common, with governments often capturing the implied transfers as fiscal or quasi-fiscal revenues (Peterson, 1979; Oliveira, 1985; Rao, 1989b; Schiff and Valdés, 1998). The importance of these mechanisms in generating fiscal revenues became evident in the wake of recent episodes of market liberalization across the developing world, which often worsened fiscal constraints (Khattry and Rao, 2002; Baunsgaard and Keen, 2010).

Prescriptions for raising private domestic saving or fiscal revenues, however, may be ineffective if industrial accumulation is constrained by insufficient foreign exchange earnings (Chenery and Bruno, 1962; Bacha, 1990; Taylor, 1994). Developing countries have limited scope for running persistent current account deficits, and yet expanding their incipient industrial sectors requires imported capital goods and intermediate inputs. Under a foreign exchange constraint, agriculture can bolster industrial accumulation by producing crops for export or by reducing food imports.

Finally, domestic demand may fail to justify private investment projects, and international trade may afford scarce possibilities to offset this shortfall, especially in the short-run. Private investment may thus be deficient even without *ex ante* saving or foreign exchange constraints, causing the economy to come under a Keynesian aggregate demand constraint.

A number of two-sector models in the post-Keynesian and Structuralist traditions have shown that technical progress in agriculture can relieve demand constraints on industrial accumulation. In common, they posit that firms make independent investment decisions with an eye to expected profitability, and that mechanisms — such as changes in functional distribution or capacity utilization — exist to endogenously bring *ex-post* saving into line with desired investment. Agriculture can be a source of autonomous demand for industry by purchasing industrial inputs and, in the presence of Engel's effects, by reducing the cost to workers of meeting their income-inelastic demand for food (see, e.g.: Taylor, 1982, 1983; Dutt, 1992; Rao, 1993; Skott, 1999; Rada, 2007).

The model in section 8 further examines these four channels of macroeconomic complementarity — the easing of saving, demand, foreign exchange, and fiscal constraints — in a formal setting. It shows that supply shocks to agriculture can quickly affect domestic demand, industrial profitability, foreign exchange receipts, and government revenues. These shocks can thus sway the output growth and capital accumulation decisions of industrial firms. The model therefore casts light on key reasons behind the main reduced-form empirical findings of this paper: that higher growth in agriculture elicits higher growth in manufacturing. The following sections describe these empirical findings.

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<sup>4</sup>For empirical evidence concerning developing countries, see Belloc and Vertova (2006), Romp and De Haan (2007), and Canning and Pedroni (2008).

### 3 Data and Empirical Model

My empirical model uses sectoral value added data from the World Development Indicators (WDI). I exclude developed and transition economies, as well as countries with less than one million inhabitants. I also exclude countries with less than 25 consecutive observations, to mitigate inference distortions caused by short panels (I relax some of these constraints to check the robustness of the results, see section 6). The resulting unbalanced panel boasts 62 countries, spans the 1960-2006 period, and has on average 36 observations per country<sup>5</sup>.

As shown in table 2, the sample mean of growth in agriculture is quite low: 2.6%/yr. Agricultural growth is also volatile, with an overall standard deviation of 8.6%. At 4.5%/yr, the sample mean of growth in manufacturing is higher, but also volatile, with an overall standard deviation of 8.5%. Perhaps surprisingly, a standard decomposition attributes most of this volatility to variation in growth within countries, as opposed to variation in growth across countries. In other words, most countries achieve high growth in both sectors, but few sustain it over time. Growth volatility is especially pronounced in Sub-Saharan Africa.

By contrast, population-weighted average temperature varies substantially across countries, but not within countries over time. The ‘between’ component of the standard deviation is nearly 4.5°C, while the ‘within’ component is only 0.5°C.

The regressions estimated in this paper follow a dynamic panel specification which allows for the propagation of short-run variation in agricultural growth over time

$$\Delta \ln(\text{VA}_M)_{i,t} = \beta_0 + \sum_{n=1}^p \alpha_n \Delta \ln(\text{VA}_M)_{i,t-n} + \sum_{j=0}^q \beta_j \Delta \ln(\text{VA}_A)_{i,t-j} + \gamma Z_{i,t} + \epsilon_i + \epsilon_t + \epsilon_{i,t} \quad (1)$$

where  $\text{VA}_M$  and  $\text{VA}_A$  denote real value added in manufacturing and agriculture. The subscripts  $i$  and  $t$  index countries and years,  $\Delta$  denotes first differences, and  $Z_{i,t}$  denotes a vector of controls. As is standard in panel data analysis, the baseline model decomposes the unobserved residual into a time-constant portion that is specific to each country ( $\epsilon_i$ ), a time-varying portion that is common to all countries ( $\epsilon_t$ ), and a time-varying portion that is specific to each country ( $\epsilon_{i,t}$ ).

By including lags of the dependent variable, the specification in (1) controls for the past dynamics of growth in manufacturing. Estimating it using the standard ‘within’ estimator thus addresses two sources of bias. First, it controls for idiosyncratic country characteristics correlated with the performance of agriculture; second, it ensures that growth in agriculture is conditionally uncorrelated with past growth in manufacturing.

Still, a causal interpretation of the coefficients requires the assumption that growth in agriculture is uncorrelated to the contemporaneous and lagged error terms, conditional on past growth in manufacturing, the country and year fixed effects, and any other covariates included in  $Z_{i,t}$ :

$$E[\epsilon_{i,t} \Delta \ln(\text{VA}_A)_{i,s}] = 0, \quad \forall s \leq t \quad (2)$$

### 4 Benchmark Estimates

The assumption in (2) is likely to fail due to omitted variable bias, motivating the use of temperature as an instrument for growth in agriculture. Before describing those estimates, however, it is

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<sup>5</sup>Although the WDI dataset goes beyond 2006, that is the last year for which measures of population-weighted temperature are available. To ensure comparability across specifications, I restrict the sample to 1960-2006 in all models.

useful to establish benchmarks — based on the identifying assumption in (2) — against which to compare them. This section describes two such benchmarks. The first is obtained by estimating equation (1) with annual data and the within estimator. The second is obtained by estimating it with growth in manufacturing over non-overlapping five-year periods, in order to establish a medium-term association that does not use annual variation for coefficient identification.

Columns (1)-(3) in table 3 show the model in (1) estimated with annual data and the within estimator. Each specification includes two lags of growth in manufacturing<sup>6</sup>. Column (1) only includes contemporaneous growth in agriculture which, as we can see, has a positive and statistically significant association with growth in manufacturing. But this contemporaneous effect is small: an increase in agricultural growth by one percentage point is associated with an increase in growth in manufacturing of only 0.10 percentage point. Columns (2) and (3) include up to three lags of growth in agriculture. Only the first two lags are positive and individually significant (deeper lags are also insignificant, but they are not reported).

The dynamic specifications in columns (1)-(3) allow us to compute cumulative effects by assuming a permanent increase in growth in agriculture.<sup>7</sup> As table 3 shows, the estimated long-run multipliers are statistically significant — the multiplier in column (2) implies that a permanent increase in growth in agriculture by one percentage point is associated with an increase of 0.32 percentage point in growth in manufacturing after enough years elapse.

The empirical model allows for both fixed effects and lags of the dependent variable. The within transformation therefore creates a mechanical correlation between the lagged dependent variable and the error term (Nickell, 1981). The resulting bias converges to zero as the number of time periods increase, so it is unlikely to be large given the average of 36 observations per country. Still, to address this potential problem, column (4) shows the specification in column (2) estimated using a system GMM procedure in the spirit of Arellano and Bond (1991), and Arellano and Bover (1995). The procedure applies forward orthogonal deviations to the variables in order to eliminate panel-specific fixed effects, using lags of the untransformed variables as instruments for the transformed variables. The identifying assumption is:

$$E[\epsilon_{i,t+1}^* (\Delta \ln(\text{VA}_M)_{i,s-1}, \Delta \ln(\text{VA}_A)_{i,s})'] = 0, \quad \forall s \leq t \quad (3)$$

where  $\epsilon_{i,t+1}^*$  indicates the forward orthogonal deviation of  $\epsilon_{i,t}$ .<sup>8</sup> If there is no autocorrelation in the error term of second (or higher) order, the first (and deeper) lags of the variables in levels are valid instruments for the transformed lagged dependent variables.<sup>9</sup> The existence of second-order autocorrelation in the error term can be tested using the data, but a causal interpretation of the coefficients on agricultural growth still hinges on the assumption in (2). The coefficients on agricultural growth show a small decline with the system GMM procedure, but growth in manufacturing shows greater persistence, so that the long-run multiplier declines only slightly.

Columns (5)-(7) present the second set of benchmark estimates, which modify the empirical model in two ways. First, I measure growth in manufacturing over non-overlapping five-year

<sup>6</sup>Deeper lags were never statistically significant, and their inclusion did not alter the results.

<sup>7</sup>The long-run multiplier that embodies this assumption of a permanent increase is given by  $\sum_{j=0}^q \beta_j (1 - \sum_{n=1}^p \alpha_n)^{-1}$ , where  $\beta_j$  and  $\alpha_n$  are the coefficients on growth in agriculture and manufacturing, and  $q$  and  $p$  are the respective number of lags, as defined in (1).

<sup>8</sup>The time subscript reflects the practice of storing orthogonal deviations one period late, for consistency with other commonly used transformations, such as first differences (Roodman, 2006). In other words,  $\epsilon_{i,t+1}^* = c_{i,t} (\epsilon_{i,t} - T_{i,t}^{-1} \sum_{j=t+1}^{T_{i,t}} \epsilon_{i,t})$ , where  $T_{i,t}$  indicates the number of available future observations, and  $c_{i,t} = \sqrt{T_{i,t}/T_{i,t} + 1}$  is a scaling factor.

<sup>9</sup>The system GMM approach also estimates equations with the untransformed variables, now using lags of the transformed lagged dependent variables as instruments. These instruments are by construction purged of correlation with the unobserved fixed effects, and it is also assumed that they are uncorrelated with other components of the contemporaneous error term. The use of these additional moment conditions to estimate a 'stacked' system is shown to increase efficiency (Arellano and Bover, 1995).

periods, starting with 1960-1965 (I annualize the result for ease of interpretation). Second, I decompose the right-hand-side variables into two types: a set of ‘flow’ variables measured as annual averages over each growth period, including growth in agriculture; and a set of ‘stock’ variables measured at the beginning of each growth period (see Caselli et al., 1996, for further discussion of this specification).

Since the transformed dataset has few time periods, I estimate all the specifications using the Arellano-Bond-Bover procedure. The specification in column (5) includes only period dummies and one lag of the dependent variable, in addition to growth in agriculture. In turn, the specifications in columns (6) and (7) attempt to attenuate omitted variable bias by including two different sets of flow and stock variables. These variables capture external and domestic macroeconomic conditions, such as the initial level of GDP per capita and the external terms of trade (for a detailed description of the variables in each set, see table B.1 in the Appendix). As we can see, the estimates in columns (6) and (7) imply that an increase by one percentage point in average growth in agriculture is associated with an increase in annual growth in manufacturing of between 0.49 to 0.52 percentage points over a five-year period. These medium-run estimates exceed the corresponding short-run OLS estimates obtained using annual data, as well as the sum of lagged coefficients in columns (2) or (4). I consider these to be the benchmark medium-run estimates obtained without external instrumental variables.

## 5 Instrumental Variable Estimates

The benchmark estimates above address important sources of bias, but in order to obtain causal estimates one needs a source of variation in agricultural output that is uncorrelated with relevant omitted variables. Dell et al. (2012a) provide a candidate instrumental variable for growth in agriculture: annual variation in country-level average temperature.

To construct their measure, Dell et al. (2012a) aggregated monthly local temperature measurements available in the larger Terrestrial Air Temperature and Precipitation Gridded Monthly Time Series dataset (Matsuura and Willmott, 2009). The original measurements were interpolated from a number of weather stations, and then made available on a spatial grid with a resolution of  $0.5^\circ \times 0.5^\circ$  of latitude and longitude (at the equator, each grid node corresponds to approximately  $56\text{km}^2$ ). Dell et al. (2012a) weighted these local measurements by local population, using a survey conducted by the Global Rural-Urban Mapping Project in 1990 (CIESIN et al., 2004). This weighting scheme rests on the assumption that land near populated areas is cultivated more intensively than land in remote areas. Other weighting schemes yielded little change to the authors’ estimates of the reduced-form impact of temperature on a number of economic outcomes (Dell et al., 2012b).

The instrumental variable estimates in this section maintain the following identifying assumption:

$$E[\epsilon_{i,t} \ln(\text{wtem})_{i,s}] = 0, \quad \forall s \leq t \quad (4)$$

where  $\ln(\text{wtem})_{i,s}$  denotes the log of population-weighted temperature. I further discuss this assumption, along with other robustness checks, in section 6.

### 5.1 ‘First Stage’ and Reduced Form Regressions

Columns (1)-(4) in table 4 examine changes in population-weighted average temperature as exogenous shifters of countrywide agricultural supply. They show fixed-effects regressions of growth in agriculture against up to three lags of the logarithm of contemporaneous temperature ( $\ln(\text{wtem})_{i,t}$ ).

Year dummies and two lags of growth in manufacturing are also included, since they are part of the structural model of interest.

All specifications show a negative contemporaneous effect of higher temperature on growth in agriculture. The effect is practically large and statistically significant. An increase in average temperature of one 'within' standard deviation (about 0.025 log points, or 0.5° C) is predicted to cut between 0.82 and 1.17 percentage points of the contemporaneous growth rate in agricultural value added, according to the estimates in columns (1) and (2). This is a large short-run decline, as the unconditional mean of agricultural growth is only 2.6%/yr.

The specifications in columns (2)-(4) add deeper lags of average temperature. As we can see, as temperature returns to its long-run average after a shock, crop yields tend to return to normal. This fact is shown in the positive and significant coefficient of the second lag of temperature — i.e.: holding contemporaneous average temperature constant, higher temperature in the previous year is expected to increase agricultural growth in the current year. Deeper lags of average temperature are statistically insignificant, suggesting that the effect of a one-off temperature shock on agricultural growth is confined to the short run.

These findings confirm the findings in Dell et al. (2012a) regarding the effects of average temperature on agricultural growth in poor countries.<sup>10</sup> They also resonate with the broader literature on the effects of temperature on crop yields, based both on controlled experiments and reduced-form estimates (see, e.g. Adams et al., 1990; Mendelsohn et al., 2000; Parry et al., 2007; Deschenes and Greenstone, 2007; Guiteras, 2009).

Columns (6)-(8) show that higher-than-average temperature hurts growth in manufacturing. In fact, the reduced-form relationship between temperature and growth in manufacturing resembles the relationship between temperature and growth in agriculture, as indicated by the signs of the contemporaneous and lagged coefficients. This finding is reassuring, since reduced-form estimates are free of the bias inherent in instrumental variable estimates (Angrist and Pischke, 2008). More importantly, it suggests that short-run variation in growth in agriculture is driving the reduced-form effect of temperature on manufacturing.

## 5.2 Baseline Results

Table 5 presents different specifications of the empirical model in equation (1) with the log of average temperature as an instrument for growth in agriculture. Column (1) displays the results of estimating a just-identified version of the model — with no lags of growth in agriculture, and only contemporaneous temperature as an instrument. By having as many excluded instruments as endogenous variables, this specification is least likely to suffer from weak-instrument bias (Angrist and Pischke, 2008). It estimates a positive, large, and statistically significant effect of growth in agriculture on contemporaneous growth in manufacturing.

Column (2) adds one lag of temperature to the instrument set.<sup>11</sup> The estimate remains large, but about one and a half decimal point lower than in the just-identified model. An increase in growth in agriculture by one percentage point is now predicted to increase contemporaneous growth in manufacturing by 0.54 percentage points — over five times higher than its counterpart estimated by OLS (compare it with column 1 in table 3).

Column (4) adds two lags of growth in agriculture, with a total of three lags of temperature in the instrument set. The additional lags have positive coefficients, and these are also

<sup>10</sup>Dell et al. (2012a) define poor countries as those whose PPP-adjusted GDP per capita is below the sample median in the first year the country enters their dataset, which includes both developing and developed countries.

<sup>11</sup>I estimate the overidentified models by GMM, although the results change little if two-stage least squares are used instead.

higher than their counterparts in the models estimated by OLS. Conditional on current agricultural growth, however, the effects of past agricultural growth are too imprecisely estimated to be deemed individually significant. This fact indicates that annual variation in temperature, with its strong mean-reverting character, is best suited for identifying short-run effects of agricultural growth. Seen from a different angle, the coefficients identified on the basis of short-run variation in temperature do not reflect longer-term changes to the economic landscape — such as technical change, or trade and macroeconomic policies — that could follow in the wake of persistent changes in growth in agriculture.

As a result, even though the implied long-run multipliers in the specifications with lagged agricultural growth are statistically significant, I adopt the short-run coefficients of the specifications without lags (such as that in column 2), as well as their associated long-run multipliers, as the baseline estimates. The addition of lags of agricultural growth hardly changes these contemporaneous coefficients.

Columns (3) and (5)-(7) address two potential shortcomings of the specifications just described. First, to address concerns with Nickell bias, columns (6) and (7) show the same specifications as those of columns (2) and (4), but now estimated using the Arellano-Bover-Bond procedure. As we can see, the estimated short-run effect of growth in agriculture is about a decimal point lower (but still significant). At the same time, growth in manufacturing exhibits more persistence, so differences in the long-run multipliers are small.

In turn, the specifications in columns (3) and (5) are estimated by the Limited Information Maximum Likelihood (LIML) method, to address concerns with weak instrument bias. These estimates are close to those obtained in columns (2) and (4), providing clear evidence against bias. Simulations show that LIML brings significant improvements in median bias relative to standard standard methods in finite samples with multiple weak instruments. As instrument strength improves, the difference in median estimates between LIML and standard methods declines, indicating a decline in median bias in the latter (Flores-Lagunes, 2007).

Formal tests of weak identification confirm these findings. These testing procedures, originated by Anderson and Rubin (1949), are based on the joint significance of the external instruments in reduced-form regressions like those of table 4.<sup>12</sup> Table 5 reports an Anderson-Rubin Wald statistic that is robust to autocorrelation and heteroskedasticity, as well as a closely related LM statistic proposed by Stock and Wright (2000). As we can see, both statistics lead us to reject the null hypothesis of weak identification in all specifications.

The LIML method also improves inference in the presence of weak instruments by reducing size distortions (Stock and Yogo, 2005). To illustrate this property, table 5 reports the Cragg-Donald Wald statistic, which is based on the joint significance of the excluded instruments in explaining the endogenous regressors. The computed values should be measured against the critical values obtained by Stock and Yogo (2005), which indicate cutoffs for maximum levels of size distortion. These critical values were reported only when available. Both the Cragg-Donald statistic and the critical values, however, assume i.i.d. disturbances.<sup>13</sup> A comparison of columns (3) and (2) reveals lower critical values for the model estimated by LIML, indicating that, given the strength of the instruments, LIML suffers less size distortion than the standard GMM estimator.

In sum, the instrumental variable estimates reveal a large and statistically significant short-run effect of growth in agriculture on growth in manufacturing. If we take the parsimonious LIML and

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<sup>12</sup>If the exclusion restriction is valid, the reduced-form coefficients can be considered a function of both of the effect of growth in agriculture on growth in manufacturing (i.e. the coefficients of the structural equation of interest) and of the effect of temperature on growth in agriculture (i.e. the coefficients on the excluded instruments in the first-stage regression). A weak first-stage relation would thus lead to insignificant coefficients in the reduced-form regression (for more details, see Baum et al., 2007).

<sup>13</sup>I also report the closely related Kleibergen-Paap Wald statistic, which is robust to heteroskedasticity and autocorrelation, but without accompanying critical values.

system GMM estimates in columns (3) and (6) as the baseline, an increase in growth in agriculture by one percentage point is expected to raise contemporaneous growth in manufacturing by between 0.47 and 0.56 percentage points.

The calculated long-run multipliers are also statistically significant and of similar magnitude (0.53 and 0.58). They suggest that if the average country in the sample were to permanently increase the rate of growth in agriculture by 1.8 percentage points (reaching the same rate exhibited by China during the sample period), the predicted long-run increase in growth in manufacturing would range between 0.95 and 1 percentage point. Such a sustained effort to raise growth in agriculture would be a remarkable achievement, as the sample mean of growth in agriculture, as seen in section 3, is only 2.6%/yr. Even though the long-run multipliers should be interpreted with caution — as they were estimated on the basis of short-run variation in the instrument —, they suggest that the payoff of sustained increases in agricultural growth on industrial growth would be substantial.

## 6 Interactions and Robustness Checks

This section tests whether the results are robust to cross-country heterogeneity, non-macroeconomic effects of growth in agriculture, changes in the sample, and the influence of outliers. It also provides further discussion on the validity of the exclusion restriction in (4).

The sample used in this paper includes only developing countries, as opposed to the related studies of Dell et al. (2012a) and Shifa (2014). Yet growth in agriculture may affect the manufacturing sector differently depending on country characteristics such as the share of agriculture in GDP, or the degree of openness to trade. In table 6, I explore potential sources of heterogeneity in two ways.

First, I estimate the baseline specification (including only contemporaneous growth in agriculture) with a full set of year dummies interacted with quartile rank dummies. Each quartile rank dummy indicates whether, at the time it entered the sample, a country belonged to that quartile of the distribution of a given characteristic of interest. The characteristics include PPP-adjusted GDP per capita, the share of agriculture in GDP, the share of imports and exports in GDP, and the share of agricultural exports in GDP (for detailed sources and definitions, see table B.1 in the Appendix).

These year-rank interactions absorb variation across countries belonging to different quartiles; as a result, the coefficients are estimated only on the basis of variation within each quartile. In other words, only similar countries (in terms of initial conditions) provide a yardstick to evaluate the effect of agricultural growth in each country. Columns (1)-(4) show that little is changed by adding year-rank interactions to the baseline LIML specification.<sup>14</sup> The only noticeable change occurs when the characteristic of interest is the share of agricultural exports in GDP, leading to an increase in the point estimate to 0.62.

Second, I add the interaction between growth in agriculture and a dummy indicating whether, at the time it entered the sample, a country was above the median in the distribution of each of the characteristics above. Columns (5)-(8) show the results ('poor' indicates a below-median initial GDP per capita).<sup>15</sup> As we can see, the interaction terms are too imprecisely estimated to be statistically significant. Column (6) suggests a stronger effect of agricultural growth in countries with a higher share of agriculture in GDP. In turn, columns (7) and (8) suggest a weaker

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<sup>14</sup>Due to the limited number of countries, it is not possible to repeat this exercise with the Arellano-Bover-Bond system GMM estimator, as the total number of instruments far exceeds the number of cross-sectional units (Roodman, 2009).

<sup>15</sup>To estimate these specifications, I used the fitted values of a first-stage regression as the instrument for growth in agriculture (for a discussion, see Wooldridge, 2010).

effect in countries that are more open, or more oriented towards primary exports. One could find justification for these differentials in the theoretical literature: as discussed in section 2, demand-side complementarity between agriculture and industry hinges on the response of intersectoral relative prices to increases in agricultural output. In small economies, greater openness to trade may link many domestic relative prices to their counterparts in world markets, thereby narrowing the scope for demand-side complementarity. But the available data warrants no firm inference in this respect.

Table 7 shows the remaining robustness checks, with LIML estimates in the top panel, and Arellano-Bover-Bond estimates in the second panel. Column (1) reproduces the baseline specification for ease of comparison.

I first examine whether the effects so far credited to macroeconomic channels in fact take place through political channels. An influential literature has shown that climate shocks increase the likelihood of civil unrest and regime changes, especially in least developed countries (Miguel et al., 2004; Hidalgo et al., 2010; Brückner and Ciccone, 2011). These forms of conflict emerge because climate shocks hurt crop yields and rural livelihoods, so they are not independent of agricultural growth. But the accompanying unrest may confound the more strictly macroeconomic effects that motivate this paper.

To address this problem, I extend the baseline specification by adding up to two lags of an indicator of civil conflict. This dummy variable indicates the existence of conflict of any type and extent; it is based on Marshall (2013), who codes the severity of episodes of civil violence, civil war, ethnic violence and ethnic war. The coefficients on the conflict dummies are negative and statistically significant, but column (2) shows that the coefficient on agricultural growth remains similar to the baseline levels.

Columns (3) and (4) show specifications estimated with an expanded sample (including countries with at least 20 consecutive observations), and with a reduced sample (including only countries with more than 30 consecutive observations). The estimates increase with the expanded sample, but remain in the the 0.5-0.6 range.

The specifications in columns (5) and (6) examine the sensitivity of the results to influential observations. The specification in column (5) excludes the observations whose residuals (obtained from the baseline specification) are in the top or bottom percentiles. In turn, the estimates in column (6) are obtained with a Winsorized (at 1%) dependent variable. As expected, these procedures reduce the influence of outliers and lead to an overall reduction of about a decimal point in the estimates. In the models estimated by LIML, however, the decline in the short-run coefficients is partly compensated by higher persistence of growth in manufacturing, leading to only modest changes in the long-run multipliers. The largest reduction occurred in the specification with excluded outliers estimated by the Arellano-Bover-Bond procedure. In all cases, however, the estimated impact of growth in agriculture remained practically large and statistically significant.<sup>16</sup>

Finally, a brief comment about the validity of the exclusion restriction in (4) is in order. There is no question that annual variation in a country's average temperature is exogenous to growth in manufacturing or agriculture. But one may claim that variation in temperature has a direct impact on growth in the manufacturing sector. For example, controlled experiments have documented a decline in measures of worker productivity in non-agricultural activities with high temperatures (Seppänen and Vuolle, 2000; Seppänen et al., 2006).

Controlled studies, however, are largely inconsequential to the main identification strategy of this paper. First, this strategy relies only on deviations of a country's annual temperature from its long-term average. As seen in section 3, there is little annual variation in average temperature within countries (the overall within-country standard deviation is only 0.5°C, with imperceptible

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<sup>16</sup>These restrictions on influential observations are implemented only as a robustness check, as no a priori reason exists to believe that the information they convey is less legitimate than that of the remaining observations.

differences across regions). By contrast, controlled studies find that large increases in air temperature are required for perceptible declines in productivity — for example, a 6°C increase in temperature from a neutrality threshold of 25°C is required to cause a 10% decline in quantifiable measures of worker performance (Seppänen et al., 2006). These large swings in annual or seasonal temperatures with respect to a region’s historical averages are implausible. Moreover, by virtue of their controlled design, these studies fail to account for organizational adaptations that, over the course of a year, could offset the effects of unusually high temperatures given a region’s historical record.

To be sure, it is possible that large but localized variation in temperature could have perceptible effects on local non-agricultural output, while having only a modest effect on the countrywide average temperature over a year (see, e.g. Zivin and Neidell, 2010, for a study using local U.S. data collected at daily frequency). But it is difficult to claim that aggregate variation in the dataset is systematically driven by such events. At any rate, given the impossibility of directly testing the exclusion restriction, it is important to be cautious when interpreting the estimates as literal predictions, as opposed to indicators of complementarity between agriculture and industry.

## 7 Potential Mechanisms

Section 2 discussed several channels through which the agricultural sector could relieve macroeconomic constraints on industrial growth, namely saving, demand, foreign exchange, and fiscal constraints. This section provides suggestive evidence on the impact of growth in agriculture — still instrumented by temperature — on proximate measures of these channels. The regressions in this section therefore replace growth in manufacturing in the empirical model in (1) with other outcomes of interest.

Table 8 summarizes the results. It shows that a one-off increase in agricultural growth leads to a decline in the terms of trade between agriculture and the total economy, to higher shares of gross capital formation and gross saving in GDP, to higher growth of real GDP per worker, to higher growth in agricultural exports, and to a decline in the share of food in total merchandise imports. These effects are statistically significant, but confined to the short-run. By contrast, the impact on the share of merchandise trade in GDP is positive but imprecisely estimated, while the impact on the share of trade tax revenues in GDP is near zero, indicating that, as agricultural growth accelerates trade tax revenues increase on a par with GDP (for detailed variable definitions, see table B.1 in the Appendix).

I estimate all specifications by system GMM using the Arellano-Bover-Bond moment conditions, including four lags of the dependent variable in the case of variables defined as ratios, and two lags in the case of variables defined as growth rates.<sup>17</sup> With one exception, I add no other controls, so as not to restrict the channels through which growth in agriculture can affect the outcome variables.<sup>18</sup>

In general, the coefficients suggest that large shocks to growth in agriculture are required to produce noticeable effects on most of these outcome variables. Large short-run shocks, however, are empirically plausible, given the within-country volatility exhibited by agricultural growth. For example, an increase in agricultural growth by one ‘within’ standard deviation (8.5 percentage points) is expected to raise the contemporaneous share of gross capital formation in GDP by 1.82 percentage points, while lowering the share of food in merchandise imports by one percentage point.

<sup>17</sup>I also report a single-tailed test of the hypothesis that the sum of lagged coefficients is less than one, to address concerns with non-stationarity of the variables in ratios.

<sup>18</sup>The exception is the model with the terms of trade between agriculture and the total economy as a dependent variable. That model controls for up to two lags of real GDP growth, so the counterfactual is characterized by a given rate of GDP growth.

The estimates in this section are consistent with the channels of complementarity described earlier in the paper (and further investigated in the formal model of section 8). A reduction in the intersectoral terms of trade, for example, is a mechanism indicating that higher growth in agriculture can serve as an autonomous source of demand for industrial goods (for example, by lowering the cost to workers of meeting their inelastic demand for food, and thus liberating income to be spent on industrial goods). In turn, the expansion of agricultural exports and the substitution of food imports are likely to raise the capacity to import industrial inputs, while the expansion of trade tax revenues may help fund public investment in infrastructure.

At the same time, however, the empirical strategy in this section identified reduced-form effects, which reflect short-run changes in other macroeconomic variables that may be induced by growth in agriculture. It does not identify channels of complementarity free from these induced changes. With this caveat in mind, it is still worth noting that growth in agriculture improves indicators of different constraints that could bind industrial growth.

## 8 A Model of Growth Complementarity

In order to cast light on the empirical findings above, I now present a simple model of growth complementarity between agriculture and industry. It extends the basic framework of ‘gap’ models to examine how agriculture can ease different constraints on industrial capital accumulation.

Gap models start by identifying the main equality constraints that have to be satisfied as accounting identities in equilibrium. The literature’s central contribution is the notion that one of them will be binding on the feasible level of investment. The difference between this binding *ex ante* constraint and its less stringent counterparts is called a ‘gap’. Since all equality constraints have to be satisfied *ex post*, the literature has posited a range of mechanisms through which all gaps are eventually eliminated (see, e.g. Chenery and Bruno, 1962; Bacha, 1990; Taylor, 1994).

I examine the interaction among three macroeconomic equality constraints: the aggregate saving-investment balance; the balance of payments identity; and the government’s capital account. Regarding the saving-investment balance, I also consider two model closures: one in which industrial accumulation is determined by saving, and one in which it is determined by an independent investment function. I refer to the first case as the saving-constrained case, and to the second as the demand-constrained case. For clarity, I consider the elimination of gaps between two equality constraints at a time, with a focus on the contribution of agricultural development to industrial accumulation.

### 8.1 General Framework

Table 9 summarizes the equations of the model. Equations (9.1)-(9.6) describe the general framework shared by all model closures. Equation (9.1) shows that agricultural output — a pure consumption good denoted by  $A$  — is produced with labor and a fixed input (e.g. effective land), under conditions of diminishing marginal returns. Agriculture is traditional in the sense of Lewis (1954), with employment determined as a residual after industrial firms make their hiring decisions. I set the total labor force equal to unity and disregard population growth, so that  $0 < l_a < 1$  is the employment share of agriculture.

Agriculture is traditional also in that labor incomes are linked to the value of the average product. But I allow for class differentiation and the appropriation of rents. Equation (9.2) thus shows that the product wage in agriculture is proportional to the average product of labor ( $\omega_a$ ), with  $0 < \beta < 1$  denoting the share of output appropriated by the rentier class.

Equation (9.3) shows that industrial output ( $M$ ) is produced with labor and capital according to a Leontieff production function, where  $q$  and  $\sigma$  are the output-labor and output-capital coefficients. The stock of utilized capital is the binding constraint on industrial output, determining industrial employment and thus rural employment as a residual. Industrial firms operate in imperfect competition and strive to maintain a desired level of excess capacity in order to respond to unforeseen demand shocks. For simplicity, I only consider growth paths along which utilization is at the desired level, so that  $\sigma$  can be treated as a constant.<sup>19</sup>

Domestic industrial goods can be used either for consumption or for investment. But to model how a shortage of foreign exchange can hinder accumulation, I assume that the private capital stock in industry ( $K_m$ ) is a composite of domestic ( $K_d$ ) and imported ( $K_i$ ) investment goods. They are utilized in fixed proportions, with  $0 < m < 1$  denoting the share of imported goods. The value of the capital stock in domestic currency is given by equation (9.4), where  $p_m$  is the price of the domestic industrial good,  $p_k^*$  is the world price of the imported capital good, and  $x$  is the nominal exchange rate.

To allow agriculture to contribute net foreign exchange earnings, I assume that part of its output, denoted by  $E_a$ , is exported. I examine this contribution of agriculture under two implicit assumptions. The first is that domestic industry faces an inelastic world demand. The second is that endogenous changes in the nominal exchange rate are unlikely to smoothly correct current account imbalances (for evidence, see Chinn and Wei, 2013). Since it is difficult to establish a foothold in new markets, this is the typical short-run scenario in many semi-industrialized economies. As a result, these economies have often adopted contractionary monetary and fiscal policies to correct current account imbalances in the face of external borrowing constraints, giving rise to stop-and-go cycles (Ocampo, 2003). The present model shows that agricultural exports can obviate the need for contractionary policies in this scenario.<sup>20</sup>

To ensure that the model reflects this scenario with maximum simplicity, I disregard exports (or imports) of the domestic industrial good, and the short-run impact of exchange rate movements on the trade balance. In addition, I follow Taylor (1989) and further assume that  $xp_k^* = p_m = p_k$ , where  $p_k^*$  is given and  $p_m$  is a domestically-determined price, since manufacturing output is not traded (see equation 9.5). Doing so allows me to treat the industrial capital stock as if it were homogeneous.

Assuming that foreign capital goods comprise all of the economy's imports, equation (9.11) gives the balance of payments identity. The terms  $\kappa$  and  $e_a$  denote net external borrowing and agricultural exports as a ratio of the capital stock, while  $g^e$  is the rate of accumulation that, for given values of  $\kappa$  and  $e_a$ , is consistent with the balance of payments identity.

The quantity of agricultural exports is determined through a quota auctioned off by the government. This stylized assumption has two advantages. First, it collapses export-promoting policy levers into a single exogenous variable. Second, it links agricultural exports to government revenues, illustrating one of the typical forms of indirectly taxing agriculture in developing countries (see, e.g. Rao, 1989a).

Quota rents are the only source of funds to finance public capital expenditures, which are entirely directed to the domestic investment good. Equation (9.6) thus shows that public investment ( $I_g$ ) is equal to the quantity of agricultural exports times the difference between the economy's external terms of trade ( $p_a^*/p_k^*$ ) and the domestic terms of trade of agriculture ( $p = p_a/p_m$ ), which I assume to be always positive.

<sup>19</sup>In other words, I only consider 'warranted' growth paths (Harrod, 1939). For an analysis of the convergence to these paths in labor-surplus economies, see Nakatani and Skott (2007), and Skott (2010).

<sup>20</sup>To be sure, in the medium and long runs developing countries have a broader menu of choices to reconcile external balance with growth targets. For a discussion of the use of competitive exchange rates under an elastic long-run demand for industrial exports, see Razmi et al. (2012).

In sum, the model's setup is such that an exogenous increase in the domestic supply of agricultural goods affects wage income, as well as the level and composition of domestic demand for industrial goods. In turn, an exogenous increase in agricultural exports raises foreign exchange receipts and government revenues.

## 8.2 Saving and Foreign Exchange Constraints

I begin by considering the classical case of saving-driven investment in industry, in which agricultural growth can raise saving out of profits by lowering industrial costs. Following Lewis (1954), nominal wages in agriculture determine nominal wages in industry, up to an exogenous constant of proportionality (assumed equal to unity). Thus, the product wage in industry ( $w_m/p_m$ ) is a positive function of the average product in agriculture ( $\omega_a$ ) and the domestic terms of trade ( $p$ ), as in equation (9.8).

Social classes have different saving propensities, with no consumption out of profits and rents, and no saving out of wages. Workers in agriculture and industry divide their consumption expenditures between the industrial and the agricultural goods, and they share the same preferences. Equation (9.9) shows the equilibrium condition of the market for agricultural goods, where  $f(\cdot)$  denotes the quantity of the agricultural good demanded by each worker. This equation determines the domestic terms of trade ( $p$ ) as a function of the average product in agriculture ( $\omega_a$ ).

Industrial firms are the only private agents that accumulate capital. Since the government capital account is assumed to be always balanced, equation (9.10) shows that the saving-constrained rate of private capital accumulation ( $g^s$ ) is equal to total domestic and foreign saving as a ratio to the industrial capital stock ( $\kappa$  denotes net borrowing from abroad as a ratio of the industrial capital stock).

Under these conditions, the following proposition can be established:

**Proposition 8.1.** *If the wage bill in the industrial sector exceeds the value of agricultural rents, an increase in labor productivity (and thus output) in agriculture will raise the saving-constrained rate of industrial accumulation if  $\eta < 1 - \epsilon$ .*

where  $\eta > 0$  is the income elasticity of demand for the agricultural good and  $\epsilon > 0$  is the elasticity of substitution between the two goods. Both of these elasticities are assumed constant regardless of the level and composition of consumption.

*Proof.* See the Appendix. □

Engel's law — the proposition according to which the share of agricultural goods in total consumption declines with the level of income — plays a central role in ensuring growth complementarity between the two sectors in this setting. It implies a value of  $\eta$  below one. Since demand for food in developing countries is likely to be price-inelastic (i.e. with a low  $\epsilon$ ), plausible Engel's effects are likely to be strong enough to ensure that the inequality in proposition 8.1 holds in most cases (for a classic statement of a similar proposition, see Jorgenson, 1961). The proposition would thus be valid in countries where voluntary net lending from agricultural rentiers is not too high relative to the size of the industrial sector — a condition likely to be verified in most developing countries (Karshenas, 1995).

The intuition behind the proposition is as follows. An increase in agricultural output — as a result of technical progress, or of a short-run weather shock — raises both agricultural rents and industrial costs in terms of the agricultural good. The  $\eta < 1 - \epsilon$  condition, however, ensures that the corresponding changes in terms of the industrial good go the opposite way: the deterioration in the terms of trade will be large enough as to raise industrial profits and reduce the value of

agricultural rents in terms of the industrial good. In other words, higher wage earnings will be more than proportionately spent on industrial goods, translating into lower unit labor costs for industry and higher saving out of profits. The net effect on total domestic saving will be positive if the industrial wage bill — which weighs on saving out of industrial profits — exceeds agricultural rents.

The additional *ex ante* saving elicited by growth in agriculture, however, may not be absorbed by domestic demand due to balance-of-payments constraints. As mentioned above, I capture this possibility by assuming that a share  $0 < m < 1$  of total investment is comprised of an imported investment good.

To see how the saving and foreign exchange constraints interact, consider panel (a) in Figure 1. It plots the accumulation rate consistent with goods market equilibrium (denoted by  $g^s$ , from equation 9.10), and the accumulation rate consistent with the balance of payments identity (denoted by  $g^e$ , from equation 9.11), both as a function of  $\kappa$ . The resulting schedules are upward-sloping, although  $g^e$  is steeper since  $0 < m < 1$ . At point  $A$ , both relations hold as accounting identities, but industrial growth is constrained by domestic saving and net borrowing from abroad adjusts endogenously. If proposition 8.1 holds, an increase in labor productivity in agriculture will shift the  $g^s$  schedule up. With the real value of agricultural exports given, the increased demand for imported capital goods is accommodated by higher net borrowing from abroad. The economy moves to point  $A'$ , where industrial growth is higher. In other words, if accumulation is constrained by domestic saving, proposition 8.1 ensures that higher agricultural output enhances industrial growth.

But most developing countries face restrictions in international capital markets if the ratio of net borrowing to GDP rises above a perceived threshold of sustainability. Therefore assume that  $\kappa$  cannot exceed  $\kappa^{\max}$ . At point  $B$  in figure 1(a), the economy is constrained by foreign exchange availability. Higher agricultural output would raise *ex ante* saving, but the impossibility of importing additional capital goods would generate a demand shortfall in the market for industrial goods. *Ex post* saving would then be brought into line with demand in Keynesian fashion, with a reduction in the income of those with a positive propensity to save.<sup>21</sup> The economy would return to point  $B$ .

Growth in agriculture can relieve this foreign exchange constraint if a rise in exports accompanies the rise in output. For example, consider a coordinated increase in agricultural output and exports (i.e.  $\omega_a$  and  $e_a$ ) so as to leave the domestic terms of trade ( $p$ ) constant. This exercise would shift the  $g^e$  schedule up, and would have an ambiguous effect on the *ex ante*  $g^s$  schedule (it would raise saving out of rural rents, but reduce saving out of industrial profits). But as long as the foreign exchange constraint remains binding, this coordinated increase would raise equilibrium industrial growth by allowing for higher imports of capital goods. The economy would move from point  $B$  to point  $B'$ .

### 8.3 Demand and Foreign Exchange Constraints

A Keynesian-type demand constraint can emerge if two conditions are satisfied. First, if investment decisions are made independently of saving decisions; and, second, if a mechanism exists to endogenously bring actual saving into line with desired investment. Following Kaldor (1957)

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<sup>21</sup>With capacity utilization at the long-run desired level, the adjustment requires dissaving through a decline in industrial profits. As further discussed in section 8.3 below, a Keynesian macro closure ensures that a demand shortfall in the market for industrial goods reduces the rate of industrial price inflation relative to nominal wage growth, raising  $w_m/p_m$  endogenously and reducing saving out of industrial profits. This closure would in general detach urban earnings from rural earnings, however, so that a demand-constrained Keynesian closure is in general incompatible with an exogenous wage premium between the two sectors. In addition, since balance-of-payments constraints do not directly actuate on individual firms, in practice the reduction in aggregate demand tends to be achieved through contractionary fiscal or monetary policies.

and Robinson (1962), I examine endogenous changes in functional distribution in the industrial sector as one such mechanism. As it turns out, functional distribution cannot respond to demand fluctuations in the market for industrial goods if the intersectoral terms of trade are determined by market equilibrium and the nominal wages in industry are pinned down by the value of rural wages, as in the saving-constrained economy just described.<sup>22</sup> Rather, a demand constraint requires that industrial product wages vary endogenously to clear the market for industrial goods, as long as the participation constraint of industrial workers is satisfied (i.e.  $w_m > w_a$ , as in equation 9.12).

The aggregate saving-investment balance now gives equation (9.14), where  $g^i = h(w_m/p_m)$  is the independent investment function, and  $h' < 0$ . With capacity utilization at the desired level,  $w_m/p_m$  is uniquely and negatively related to the industrial profit rate. In turn, the equilibrium condition for the market for agricultural goods is now given by (9.15).

Equations (9.14) and (9.15) jointly determine the domestic terms of trade and the industrial product wage, which in turn determines the equilibrium rate of accumulation. The following proposition can be thus be established:

**Proposition 8.2.** *Provided the market for industrial goods is stable, an increase in labor productivity in agriculture will raise the demand-constrained rate of industrial accumulation if  $\eta < 1 - \epsilon + \frac{E_a}{\omega_a l_a - E_a}$ .*

*Proof.* See the Appendix. □

The intuition behind this result is as follows. An increase in agricultural output raises saving out of agricultural rents in terms of the agricultural good, but it also causes a decline in the terms of trade of agriculture. If Engel's effects are strong enough, the decline in the terms of trade will more than offset the increase in agricultural rents in terms of the agricultural good. As a result, total *ex ante* saving will decline in terms of the industrial good.<sup>23</sup>

The counterpart of this reduction in *ex ante* saving is excess demand for industrial goods. The reason is that the initial decline in agriculture's terms of trade reduces the cost to industrial workers of meeting their income-inelastic demand for food, freeing up income which is more than proportionately spent on industrial goods. If the equilibrium is stable (i.e. industrial saving reacts more strongly to changes in profitability than industrial investment), the market is cleared by an increase in the price of the industrial good relative to the nominal wage. The ensuing redistribution towards industrial profits will raise desired accumulation and bring about an increase in saving out of industrial profits to finance it. This result — whereby an *ex ante* decline in saving results in an *ex post* increase in saving and accumulation — is akin to the Keynesian paradox of thrift. It also resonates with previous analyses of two-sector models with a demand constraint (see, e.g., Taylor, 1982; Skott and Larudee, 1998).

Panel (b) in figure 1 depicts, in the  $(g, \kappa)$  space, the interaction of the goods market equilibrium schedule in (9.14), denoted by  $g^d$ , and the balance of payments identity in (9.11), denoted again by  $g^e$ . The  $g^d$  schedule is downward-sloping: with an independently given investment function, a rise

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<sup>22</sup>Mathematically, the introduction of an independent investment function requires the introduction of a new endogenous variable. For a classic exposition of the problem in a one-sector model, see Marglin (1984). For restatements in two-sector models, see Dutt (1992) and Rada (2007).

<sup>23</sup>The decline in *ex ante* agricultural saving in terms of the industrial good will also be greater the higher the share of exports in total agricultural output, as they are a component of agricultural incomes that does not contribute to the domestic supply of food. An increase in this domestic supply will thus have a higher proportional impact on the terms of trade if its initial value is low relatively to total agricultural incomes, easing the need for strong Engel's effects. It is easy to see that in a closed economy the condition for an increase in agricultural productivity to raise equilibrium growth is the same in the saving-constrained and demand-constrained regimes (with different interpretations of the underlying economic mechanisms).

in  $\kappa$  has to be the counterpart of lower agricultural exports, which reduce government revenues and its purchases of domestic investment goods for capital formation (see equation 9.6).

At point  $A$ , the economy is demand-constrained. Under the conditions of proposition 8.2, higher productivity in agriculture shifts the  $g^d$  schedule up and raises equilibrium growth (to  $A'$ ), at the cost of higher net borrowing from abroad. By contrast, if the rate of net borrowing is at the maximum level, even if *ex ante* aggregate demand is high enough to sustain growth above point  $B'$  when  $\kappa = \kappa^{max}$ , the economy will be unable to import the required capital goods. This foreign exchange constraint lowers effective investment demand, so that the *ex post*  $g^d$  schedule passes through point  $B$ .<sup>24</sup> Now increases in agricultural output are ineffective if exports don't increase as well.

As before, a policy experiment designed to raise both agricultural output and exports while leaving the terms of trade constant can ease this binding foreign exchange constraint by shifting the  $g^e$  schedule up. Equilibrium growth would thus rise to a point like  $B'$ .

## 8.4 Fiscal and Demand Constraints

As discussed in section 2, technical complementarity may exist between public infrastructure and private investment projects. As a result, private projects which would be profitable given expected demand may not be undertaken in the absence of complementary public investment. The model's set-up, where  $e_a$  is determined through a quota imposed and auctioned by the government, lends itself to analyzing the typical case in which growth in agriculture can fund public investment even when the sector's incomes are not directly taxed (see section 2).

To see how a fiscal constraint interacts with a demand constraint, assume that private accumulation responds only to private profitability as long as a high enough level of public investment is provided. If public investment falls short of this minimum, however, it becomes a binding constraint on private accumulation. The investment function can then be recast in discontinuous form, as in equation (9.16), where  $\alpha > 0$  is a constant of proportionality.

Panel (c) in figure 1 shows the interaction between the demand and the fiscal constraints. The  $g^d$  schedule shows equation (9.14) in the  $(e_a, g)$  space.<sup>25</sup> The curve describing the scaled rate of public capital formation ( $\alpha I_g / K_m$ ), in turn, is concave. On the one hand, higher exports raise quota rents for given domestic terms of trade ( $p$ ). But, on the other hand, they raise the domestic terms of trade, and thus lower the wedge relative to border prices created by the policy.

As drawn, the diagram shows an economy constrained by private demand between points  $B$  and  $C$ . At point  $A$ , however, public capital formation is low enough to become a binding constraint on private investment. The so-constrained private investment causes the  $g^d$  schedule to shift down so that it, too, intercepts  $A$  — an endogenous decline in the profit rate ensures saving-investment balance. Higher agricultural exports can ease the binding fiscal constraint and move the economy from  $A$  to  $A'$ , where industrial growth is higher.

## 8.5 Summary

The model exemplified basic macroeconomic mechanisms through which growth in agriculture can remove binding constraints on the rate of industrial accumulation in developing countries. In the

<sup>24</sup>In practice, the decline in investment demand in the face of a balance-of-payments constraint is brought about by contractionary policies, as the balance-of-payments constraint does not act upon individual firms.

<sup>25</sup>A *ceteris paribus* increase in agricultural exports has ambiguous effects on aggregate demand for industrial goods. On the one hand, it restricts domestic agricultural supply and increases the terms of trade, reducing demand out of industrial wages. On the other hand, it increases government revenues and capital expenditures, which are directed to domestic industrial goods. As drawn, the schedule embodies the assumption that the first effect predominates. The substantive results do not depend on this assumption.

case of saving and demand constraints, Engel's effects in the final demand for agricultural goods — an uncontested stylized fact — play a pivotal role in ensuring growth complementarity between the two sectors, echoing previous analyses (e.g. Jorgenson, 1961; Taylor, 1982). When combined with net export expansion, growth in agriculture can also ease foreign exchange constraints on industrial investment. Finally, if public and private capital are complementary, growth in agriculture can ease fiscal constraints on private accumulation, as it can increase real fiscal revenues either directly or indirectly (see also Rao, 1993).

The model shows that supply shocks to agriculture can sway industrial firms in their output growth and capital accumulation decisions, casting light on key structural mechanisms behind the empirical finding that growth in agriculture elicits growth in manufacturing.

## 9 Conclusion

Development economists have long examined macroeconomic channels through which development in agriculture can support the expansion of high-productivity activities, particularly in manufacturing. The complementarity between the two sectors has also been a centerpiece of historical studies of industrialization. Efforts to identify causal effects using country-level datasets, on the other hand, are comparatively recent. This paper makes a contribution to this literature.

Using variation in population-weighted average temperature as an instrument for growth in agriculture, my baseline estimates show that it has large short-run impacts on manufacturing: an increase in annual growth in agriculture by one percentage point is estimated to raise contemporaneous growth in manufacturing by between 0.47 and 0.56 percentage point. Baseline estimates of the long-run effects of a permanent increase in agricultural growth are also large (between 0.53 and 0.58 percentage points), although short-run variation in the instrument is better suited to capture short-run effects.

The paper also presented a stylized two-sector model to illuminate the channels through which such complementarity might arise. The model focused on agriculture's classic roles as a source of saving, foreign exchange, fiscal revenues, and a home market for industry. It demonstrated that even short-run supply shocks to agriculture can induce industrial firms to revise their output growth and capital accumulation decisions, and that persistent increases in agricultural growth — due to continuing technical change and factor accumulation — can lead to persistent increases in industrial growth. Further empirical estimates showed that growth in agriculture favorably affects indicators of the channels of complementarity highlighted in the model.

My findings have two broad implications. The first is negative in outlook: given that large negative shocks to agricultural output are frequent in the sample, especially among poorer countries, my estimates suggest that their cost in foregone industrial output can be severe, and that they may abruptly interrupt processes of industrial growth. The use of modern inputs in agriculture to reduce its vulnerability to weather shocks is important not only for increasing income security among rural residents, but also for preventing regressive structural change in the economy.

The second implication is that sustained agricultural development should be a key ingredient of policies promoting long-term industrialization. A thorough discussion of the type of agricultural development most conducive to industrial growth, however, is beyond the scope of this paper (for a discussion, see Rao and Caballero, 1990). Suffice it to say that successful cases of industrialization were often accompanied by technical dynamism in agriculture and a broad distribution of its benefits. Enabling conditions have included equitable land tenure systems and, in countries characterized by hidden unemployment, a pattern of technical change focused on increasing output and the utilization of rural labor. In successful cases of industrialization in Asia, the widespread adoption of land-augmenting innovations, such as intensive cropping, fertilizers, and high-yielding varieties, has proven to be conducive to these goals (Smith, 1959; Ishikawa, 1967; Lee, 1979; Kay,

2001). They stand in contrast to innovations designed to primarily economize on labor, which characterized some Latin American experiences and arguably worsened the problem of hidden unemployment (De Janvry, 1978; Sanders and Ruttan, 1978). Achieving such technical dynamism, as well as an adequate output response to demand fluctuations, has also required public investment in research and extension, and in complementary rural infrastructure. It is therefore essential to examine the forms these policies may take in contemporary processes of development.

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# Appendix

## A Proofs of Propositions

### A.1 Proof of Proposition 8.1

Since the export quota is an exogenous policy variable, it follows from (9.9) that

$$\frac{dp}{d\omega_a} = \left( \frac{\eta - 1}{\omega_a \epsilon} \right) p \quad (\text{A.1})$$

where  $\eta > 0$  and  $\epsilon > 0$  are the constant elasticities of income and substitution. From (9.10) and (A.1), it follows that

$$\frac{\partial g^s}{\partial \omega_a} > 0 \Leftrightarrow [\beta - (1 - l_a)] \frac{\eta - 1}{\epsilon} > -[\beta - (1 - l_a)] \quad (\text{A.2})$$

If  $(1 - l_a) > \beta$ , (A.2) will be satisfied if  $\eta < 1 - \epsilon$ . Note that  $(1 - l_a) > \beta$  implies  $(1 - l_a)\omega_a p_a > \beta\omega_a p_a$ , i.e., that the wage bill in the industrial sector exceeds the value of agricultural rents.

### A.2 Proof of Proposition 8.2

From (9.15), it follows that

$$\frac{dp}{d\omega_a} = -pl_a \left[ \frac{1 - \frac{f^A}{\omega_a} \eta}{l_a f^A \epsilon + (1 - l_a) f^M \epsilon + (1 - l_a) f^M \eta} \right] \quad (\text{A.3})$$

where, for notational ease, I define  $f^A = f[(1 - \beta)\omega_a, p]$  and  $f^M = f[w_m/pm^{1/p}, p]$ , that is, the per-worker demand for agricultural goods in the agricultural and industrial sectors. The equation above also relies on the assumption of constant elasticities of income ( $\eta$ ) and substitution ( $\epsilon$ ).

If the Keynesian stability condition holds, i.e.,  $\left| \frac{\sigma}{q} \right| > |h'|$ , the necessary and sufficient condition for an increase in  $\omega_a$  to raise accumulation is for it to reduce the value of agricultural saving in terms of the industrial good, that is:

$$\frac{dg^D}{d\omega_a} > 0 \Leftrightarrow \frac{d \left[ \frac{\beta\omega_a l_a p}{K_m} \right]}{d\omega_a} = p + \omega_a \frac{dp}{d\omega_a} < 0 \quad (\text{A.4})$$

Combining (A.3) and (A.4), along with the market equilibrium condition of  $\omega_a l_a = f^A l_a + f^M (1 - l_a) + E_a$  yields Proposition 8.2.

## B Data Sources and Variable Definitions

Table B.1: Data Sources and Variable Definitions

General	Source/Definition
Sectoral real value added	WDI
Population-weighted average temperature	Dell et al. (2012a)
<b>Table 3</b>	
<i>Set 1 : Stocks</i>	
Human capital stock	PWT 8
PPP-adjusted GDP per capita (log)	PWT 8
<i>Set 1 : Flows</i>	
Real exchange rate undervaluation (log)	Rodrik (2008), calculated with data from PWT 8
External terms of trade (log)	PWT 8
<i>Set 2: Stocks</i>	
Polity II score	Marshall et al. (2013), updated.
<i>Set 2: Flows</i>	
Merchandise trade (% of GDP)	PWT 8
Real exchange rate volatility	PWT 8, defined as standard deviation over each period.
Gross capital formation (% of GDP)	PWT 8
Government consumption (% of GDP)	PWT 8
<b>Table 6</b>	
PPP-adjusted GDP per capita	PWT 8
Share of agriculture in GDP	WDI
Agricultural exports	FAO
Merchandise trade (% of GDP)	PWT 8
<b>Table 7</b>	
Civil conflict	Marshall (2013)
<b>Table 8</b>	
Gross capital formation (% of GDP)	WDI
Gross saving (% of GDP)	WDI
Labor Productivity	PWT 8, defined as real GDP per person engaged
Agricultural exports	FAO
Food imports over merchandise imports	WDI
Merchandise trade (% of GDP)	PWT 8
Trade tax revenues (% of GDP)	Baunsgaard and Keen (2010)

*Note:* WDI: World Development Indicators. PWT 8: Penn World Table version 8.0. FAO: Food and Agriculture Organization.

Figure 1: Pairwise Interaction of Saving, Foreign Exchange, Demand, and Fiscal Constraints

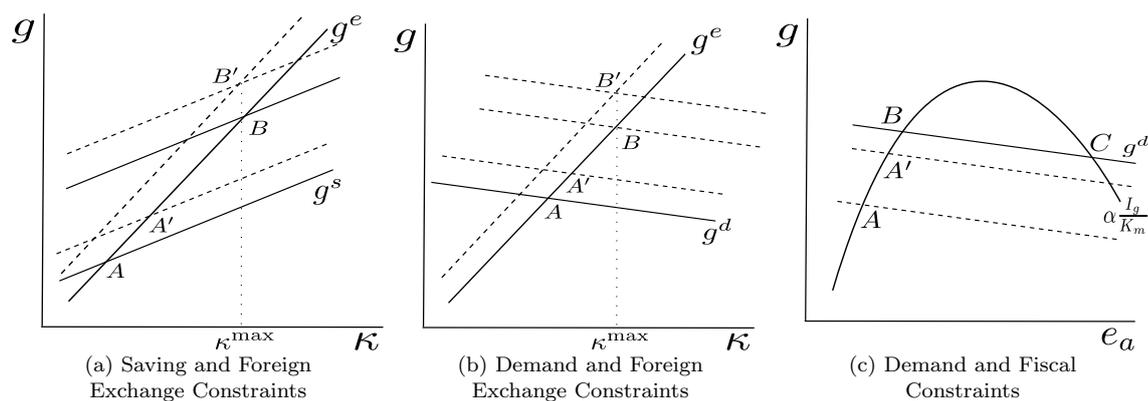


Table 2: Descriptive Statistics of Growth and Temperature

	Mean	Standard Deviation			Countries
		Overall	Within	Between	
<b>Growth in Agriculture (%/yr.)</b>					
All Countries	2.67	8.62	8.54	1.22	62
Asia and Pacific Islands	3.01	4.78	4.74	0.69	13
Latin America and Caribbean	2.57	6.00	5.92	1.02	17
Sub-Saharan Africa	2.27	9.81	9.74	1.19	24
Middle East and North Africa	3.79	14.16	14.08	1.63	7
<b>Growth in Manufacturing (%/yr.)</b>					
All Countries	4.51	8.55	8.17	2.51	62
Asia and Pacific Islands	6.85	7.69	7.23	2.81	13
Latin America and Caribbean	3.29	6.74	6.58	1.53	17
Sub-Saharan Africa	3.68	10.30	10.04	2.28	24
Middle East and North Africa	5.78	7.60	7.49	1.36	7
<b>Weighted Temperature (<math>^{\circ}\text{C}/\text{yr.}</math>)</b>					
All Countries	21.83	4.48	0.49	4.48	62
Asia and Pacific Islands	22.76	4.90	0.34	5.00	13
Latin America and Caribbean	21.19	4.04	0.51	4.13	17
Sub-Saharan Africa	23.22	3.78	0.48	3.83	24
Middle East and North Africa	17.37	3.83	0.63	4.07	7

*Notes:* The decomposition of the overall standard deviation was obtained by using the following transformation:  $\tilde{y}_{i,t} = y_{i,t} - \bar{y}_i + \bar{y}$ . Where  $y$  denotes the variable of interest,  $i$  denotes countries,  $t$  denotes years,  $\bar{y}_i$  denotes the average of  $y$  across time in country  $i$ , and  $\bar{y}$  denotes the overall average of  $y$ . The within-country standard deviation is the standard deviation of  $\tilde{y}_{i,t}$ , while the between-country standard deviation is given by the standard deviation of  $\bar{y}_i$  across all countries. For details about the data sources, see the Appendix.

Table 3: Benchmark estimates without temperature as an instrument.

	(1)	(2)	(3)	(4)	Model with five-year averages		
					(5)	(6)	(7)
	OLS/FE	OLS/FE	OLS/FE	SGMM	SGMM	SGMM	SGMM
$\Delta \ln (\text{Agr. V.A.})$	0.116*** (0.041)	0.143*** (0.044)	0.145*** (0.044)	0.116** (0.047)	0.592*** (0.143)	0.491*** (0.137)	0.527*** (0.161)
$\Delta \ln (\text{Agr. V.A.})_{t-1}$		0.109*** (0.030)	0.114*** (0.030)	0.094*** (0.031)			
$\Delta \ln (\text{Agr. V.A.})_{t-2}$		0.060*** (0.021)	0.071*** (0.024)	0.051** (0.022)			
$\Delta \ln (\text{Agr. V.A.})_{t-3}$			0.031 (0.023)				
Country FE	Y	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y	Y
Controls	N	N	N	N	N	set 1	sets 1+2
$\sum \Delta \ln (\text{Agr. V.A.})$	0.116	0.312	0.361	0.261	0.592	0.491	0.527
p-value	0.007	0	0	0.001	0	0.001	0.002
Lags of $\Delta \ln (\text{Man. V.A.})$	2	2	2	2	1	1	1
$\sum \Delta \ln (\text{Man. V.A.})$	0.054	0.028	0.048	0.082	0.187	0.130	0.084
p-value	0.331	0.615	0.269	0.145	0.032	0.118	0.280
Long-run Multiplier	0.123	0.321	0.380	0.284	0.728	0.565	0.576
p-value	0.008	0	0	0	0	0.001	0.004
Num. of GMM Instruments				52	14	20	33
Lags of GMM Instruments				2	2	2	2
AR(2) test (p-value)				0.331	0.343	0.323	0.168
Hansen test (p-value)				0.182	0.158	0.683	0.242
Countries	62	62	62	62	62	55	55
Avg. Obs. per Country	36.048	35.952	35.194	35.952	6.226	6.309	6.309

\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

*Notes:* The dependent variable is growth in real value added in the manufacturing sector. Standard errors robust to arbitrary forms of correlation within countries are in parentheses. Columns labeled "OLS/FE" indicate the fixed-effects estimator, while columns labeled "SGMM" indicate the Arellano-Bover-Bond system GMM estimator (see description in the text). All specifications include a set of year or period dummies. The specifications in columns (6) and (7) also include the sets of control variables described in the text (for detailed variable definitions for each set, see the data appendix). The SGMM estimates of the specifications in columns (5)-(7) were obtained using two lags of the lagged dependent variable and of all all endogenous flow variables (with the exception of the external terms of trade) as instruments for the transformed equation. The instrument matrix was collapsed to avoid instrument proliferation (Roodman, 2009).

Table 4: ‘First stage’ and reduced form regressions

	‘First Stage’ Regressions				Reduced Form Regressions		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	OLS/FE	OLS/FE	OLS/FE	OLS/FE	OLS/FE	OLS/FE	SGMM
ln(wtem)	-0.338*** (0.094)	-0.469*** (0.144)	-0.468*** (0.141)	-0.526** (0.210)	-0.234** (0.088)	-0.293*** (0.088)	-0.276*** (0.080)
ln(wtem) <sub>t-1</sub>		0.354** (0.165)	0.317* (0.170)	0.403* (0.218)		0.112 (0.074)	0.129* (0.076)
ln(wtem) <sub>t-2</sub>			0.151 (0.106)	0.116 (0.116)		0.085 (0.082)	0.139 (0.091)
ln(wtem) <sub>t-3</sub>			-0.097 (0.104)	-0.090 (0.127)		-0.007 (0.062)	-0.008 (0.059)
Country FE	Y	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y	Y
Region-Year FE	N	N	N	Y	N	N	N
Controls	N	N	N	N	N	N	N
F-Stat of ln(wtem)	13.055	6.313	3.301	1.841	7.006	3.379	3.954
p-value	0.001	0.003	0.016	0.132	0.010	0.015	0.006
$\sum \ln(\text{temp})$	-0.338	-0.115	-0.096	-0.097	-0.234	-0.103	-0.016
p-value	0.001	0.166	0.265	0.412	0.010	0.402	0.332
Lags of $\Delta \ln(\text{Man. V.A.})$	2	2	2	2	2	2	2
$\sum \Delta \ln(\text{Man. V.A.})$	0.017	0.024	0.025	0.036	0.054	0.085	0.118
p-value	0.550	0.423	0.417	0.243	0.337	0.063	0.037
Countries	62	62	62	62	62	62	62
Avg. Obs. per Country	36.065	36.065	36.032	36.032	36.145	36.113	36.113
Long-run Multiplier					-0.247	-0.113	-0.018
p-value					0.011	0.402	0.331
Num. of GMM Instruments							53
Lags of GMM Instruments							2
AR(2) test (p-value)							0.457
Hansen test (p-value)							0.117

\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

*Notes:* The dependent variable in the “first-stage” regressions (columns 1-4) is growth in real value added in agriculture. The dependent variable in the reduced-form regressions (columns 5-7) is growth in real value added in manufacturing.  $\ln(\text{wtem})$  indicates the log of population-weighted average temperature. All specifications include a set of year dummies and two lags of growth in manufacturing. Specification (4) includes a set of region-specific dummies, where the regions are Middle-East and North Africa; Sub-Saharan Africa; Latin America and the Caribbean; Asia and Pacific Islands. Standard errors robust to arbitrary forms of correlation within countries are in parentheses. Columns labeled “OLS/FE” indicate the fixed-effects estimator, while columns labeled “SGMM” indicate the Arellano-Bover-Bond system GMM estimator (see description in the text).

Table 5: Baseline estimates with temperature as instrument for agricultural growth.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	GMM	GMM	LIML	GMM	LIML	SGMM	SGMM
$\Delta \ln (\text{Agr. V.A.})$	0.684**	0.547***	0.559***	0.581***	0.581***	0.476**	0.455***
	(0.280)	(0.185)	(0.190)	(0.182)	(0.185)	(0.185)	(0.157)
$\Delta \ln (\text{Agr. V.A.})_{t-1}$				0.157	0.157		0.201
				(0.193)	(0.200)		(0.191)
$\Delta \ln (\text{Agr. V.A.})_{t-2}$				0.112	0.112		0.046
				(0.145)	(0.147)		(0.166)
Country FE	Y	Y	Y		Y	Y	Y
Year FE	Y	Y	Y		Y	Y	Y
Lags of $\Delta \ln (\text{Man. V.A.})$	2	2	2	2	2	2	2
$\sum \Delta \ln (\text{Man. V.A.})$	0.043	0.045	0.045	0.030	0.030	0.102	0.089
p-value	0.423	0.399	0.394	0.629	0.631	0.050	0.274
Long-run Multiplier	0.714	0.573	0.586	0.876	0.876	0.530	0.771
p-value	0.015	0.003	0.004	0.004	0.007	0.012	0.036
Lags of $\ln(\text{temperature})$	0	1	1	3	3	1	3
Anderson-Rubin Stat	7.127	14.457	14.457	14.060	14.060		
p-value	0.008	0.001	0.001	0.007	0.007		
Stock-Wright Stat	6.354	10.078	10.078	10.150	10.150		
p-value	0.012	0.006	0.006	0.038	0.038		
Kleibergen-Paap Stat.	13.062	6.316	6.316	2.388	2.388		
Cragg-Donald Stat	14.489	14.131	14.131	6.678	6.678		
10% Max Size Crit. Val.	16.380	19.930	8.680	n/a	n/a		
15% Max Size Crit. Val.	8.960	11.590	5.330	n/a	n/a		
Hansen test (p-value)		0.479	0.481	0.998	0.998	0.444	0.253
Num. of GMM Instruments						51	53
Lags of GMM Instruments						2	2
AR(2) test (p-value)						0.636	0.564
Countries	62	62	62	62	62	62	62
Avg. Obs. per Country	36.048	36.048	36.048	35.919	35.919	36.048	35.919

\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

*Notes:* The dependent variable is growth in real value added in the manufacturing sector. Standard errors robust to arbitrary forms of correlation within countries are in parentheses. Columns labeled 'GMM' indicate that the GMM method was used for estimating the models with more instruments than endogenous variables. Columns labeled 'LIML' indicate that the Limited Information Maximum Likelihood method was used instead. Columns labeled 'SGMM' indicate the use of the Arellano-Bover-Bond system GMM method (for a description, see the text). The Stock-Yogo critical values for maximal size distortion were computed for the Cragg-Donald F statistic, which assumes i.i.d. disturbances. They were reported only when available.

Table 6: Robustness Checks and Interaction Effects

	Year-Quartile Rank Dummies				(5)	(6)	(7)	(8)
	(1)	(2)	(3)	(4)				
	gdp per cap.	agr. share in gdp	agr. exp. in gdp	trade in gdp	poor	agr. share in gdp	agr. exp. in gdp	trade in gdp
$\Delta \ln$ (Agr. V.A.)	0.554***	0.513***	0.625***	0.555***	0.543***	0.525***	0.611***	0.678***
	(0.185)	(0.183)	(0.201)	(0.171)	(0.173)	(0.183)	(0.232)	(0.254)
Interaction					0.025	0.291	-0.208	-0.249
					(0.230)	(0.337)	(0.236)	(0.234)
Country FE	Y	Y	Y	Y	Y	Y	Y	Y
Year FE	N	N	N	N	Y	Y	Y	Y
Year-Rank FE	Y	Y	Y	Y	N	N	N	N
Lags of $\Delta \ln$ (Man. V.A.)	2	2	2	2	2	2	2	2
$\sum \Delta \ln$ (Man. V.A.)	0.042	0.054	0.051	0.046	0.045	0.047	0.047	0.041
p-value	0.397	0.193	0.344	0.342	0.396	0.384	0.378	0.445
Long-run Multiplier	0.578	0.542	0.659	0.582	0.569	0.551	0.641	0.707
p-value	0.003	0.005	0.002	0.001	0.002	0.005	0.009	0.007
Lags of $\ln$ (temperature)	1	1	1	1	1	1	1	1
Anderson-Rubin Stat	15.854	15.693	12.791	15.680	15.599	16.183	15.639	14.461
p-value	0	0	0.002	0	0	0	0	0.001
Stock-Wright Stat	10.281	10.136	10.253	11.300	10.509	12.474	10.752	10.113
p-value	0.006	0.006	0.006	0.004	0.005	0.002	0.005	0.006
Kleibergen-Paap Stat.	8.119	6.094	4.268	5.650	3.104	2.515	4.859	4.468
Cragg-Donald Stat	15.124	14.267	10.800	15.088	12.692	9.408	14.892	14.777
10% Max Size Crit. Val.	8.680	8.680	8.680	8.680	7.03	7.03	7.03	7.03
15% Max Size Crit. Val.	5.330	5.330	5.330	5.330	4.58	4.58	4.58	4.58
Countries	62	58	59	61	62	62	62	62
Avg. Obs. per Country	36.048	35.966	36.424	36.082	36.048	36.048	36.048	36.048

\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

*Note:* The dependent variable is growth in real value added in the manufacturing sector. All specifications were estimated by the Limited Information Maximum Likelihood method. Standard errors robust to arbitrary forms of correlation within countries are in parentheses. Columns (1)-(4) include a full set of interaction dummies between year and the rank of a country in the quartiles of each characteristic at the time the country enters the sample. The characteristics of interest are listed in the column headers (for detailed variable definitions, see the text and the data appendix). The variable 'interaction' in columns (5)-(8) denotes the interaction term between growth in agriculture and a dummy indicating whether a country lies above the overall median in the distribution of each characteristic at the time the country enters the sample ('poor' indicates whether a country lies below the overall median of GDP per capita). The Stock-Yogo critical values for maximal size distortion were computed for the Cragg-Donald F statistic, which assumes i.i.d. disturbances.

Table 7: Additional Robustness Checks

	(1) Baseline	(2) Conflict	(3) > 20 obs	(4) > 30 obs	(5) Truncated Outliers	(6) Winsorized Dep. Var.
LIML Estimates						
$\Delta \ln$ (Agr. V.A.)	0.559*** (0.190)	0.577*** (0.203)	0.627*** (0.210)	0.548*** (0.177)	0.470*** (0.178)	0.486*** (0.168)
Long-run Multiplier	0.586	0.609	0.666	0.580	0.542	0.534
p-value	0.004	0.005	0.003	0.003	0.008	0.004
Hansen test (p-value)	0.481	0.396	0.450	0.493	0.462	0.603
Countries	62	61	70	54	62	62
Avg. Obs. per Country	36.048	36	34.214	37.796	34.113	36.048
Arellano-Bover-Bond SGMM Estimates						
$\Delta \ln$ (Agr. V.A.)	0.476** (0.185)	0.492** (0.203)	0.565** (0.243)	0.418** (0.183)	0.286** (0.126)	0.384** (0.145)
Long-run Multiplier	0.530	0.558	0.634	0.481	0.347	0.443
p-value	0.012	0.017	0.022	0.027	0.036	0.012
AR(2) test (p-value)	0.636	0.679	0.584	0.761	0.216	0.878
Hansen test (p-value)	0.444	0.468	0.356	0.220	0.172	0.481
Countries	62	61	70	54	62	62
Avg. Obs. per Country	36.048	36	34.214	37.796	34.177	36.048

\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

*Note:* The dependent variable is growth in real value added in the manufacturing sector. The specifications in the top panel were estimated by the Limited Information Maximum Likelihood method using the within transformation. The specifications in the bottom panel were estimated by system GMM using the Arellano-Bover-Bond moment conditions. Standard errors robust to arbitrary forms of correlation within countries are in parentheses. All specifications include a full set of year dummies and use the contemporaneous plus one lag of temperature as instruments for growth in agriculture. Column (2) includes up to two lags of a dummy indicating the existence of civil conflict (see the text and the Appendix for detailed variable definitions). Columns (3) and (4) include countries with at least 20 and 30 consecutive observations, respectively. Column (5) drops the observations whose residuals (obtained from the baseline specification) are in the top or bottom percentiles. Column (6) uses a Winsorized (at 1%) dependent variable.

Table 8: Impact of Agricultural Growth on Potential Mechanisms

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	agr/total	gross	gross	labor	agr.	food	merch.	trade
	t.o.t.	cap. form.	saving	prod.	exports	imports	trade	tax rev.
	(log)	(% gdp)	(% gdp)	(growth)	(growth)	(% total)	(% gdp)	(% gdp)
$\Delta \ln (\text{Agr. V.A.})$	-0.687**	0.222*	0.441***	0.407*	1.853**	-0.128*	0.474	-0.055
	(0.309)	(0.113)	(0.146)	(0.223)	(0.792)	(0.068)	(0.402)	(0.037)
$\Delta \ln (\text{Agr. V.A.})_{t-1}$	-0.056	-0	-0.036	-0.104	1.642	-0.072	0.048	-0.032
	(0.439)	(0.099)	(0.222)	(0.185)	(1.037)	(0.088)	(0.430)	(0.033)
$\Delta \ln (\text{Agr. V.A.})_{t-2}$	0.256	-0.094	0.141	0.021	0.359	0.010	0.252	-0.061
	(0.290)	(0.120)	(0.161)	(0.138)	(0.827)	(0.071)	(0.379)	(0.059)
Country FE	Y	Y	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y	Y	Y
Controls	Y	N	N	N	N	N	N	N
Lags of Dep. Var	4	4	4	2	2	4	4	4
$\sum$ Lagged Dep. Var.	0.953	0.928	0.950	0.222	-0.353	0.957	0.932	0.978
$H_o: \sum < 1$ (p-value)	0.996	0.886	0.990	1	1	0.806	0.981	0.736
Long-run Multiplier	-10.370	1.768	11.005	0.417	2.849	-4.442	11.393	-6.730
p-val	0.381	0.535	0.061	0.433	0.033	0.465	0.380	0.638
Num. of GMM Instruments	59	59	55	53	52	54	56	41
Lags of GMM Instruments	2	5	1	2	2	2	2	2
AR(2) test (p-value)	0.624	0.477	0.248	0.402	0.632	0.796	0.455	0.584
Hansen test (p-value)	0.137	0.126	0.145	0.563	0.224	0.072	0.849	0.304
Countries	62	62	62	57	59	62	57	54
Avg. Obs. per Country	36.516	37.129	37.452	34.947	38.220	23.806	37.719	23.574

\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

*Note:* The dependent variables are as follows. Column (1): the log of the terms of trade between agriculture and the total economy; column (2): gross capital formation as a share of GDP; column (3): gross saving as a share of GDP; column (4): annual growth rate of real GDP per person engaged; column (5): annual growth rate of real agricultural exports; column (6): food imports as a share of total merchandise imports; column (7): the sum of merchandise exports and imports as a share of GDP; column (8): trade taxes as a share of GDP (see the data Appendix for detailed definitions). All specifications were estimated by system GMM according to the Arellano-Bover-Bond procedure with three lags of temperature as external instruments for agricultural growth. Standard errors robust to arbitrary forms of correlation within countries are in parentheses. Columns (1) includes two lags of GDP growth as controls.

Table 9: Model Equations

General Framework:

$$A = \phi l_a^{1-\delta}, \quad \phi > 0 \text{ and } 0 < \delta < 1 \quad (9.1)$$

$$\frac{w_a}{p_a} = (1 - \beta)\phi l_a^{-\delta} = (1 - \beta)\omega_a, \quad 0 < \beta < 1 \quad (9.2)$$

$$M = \text{Min}(q(1 - l_a), \sigma K_m) \quad (9.3)$$

$$p_k K_m = (1 - m)p_m K_d + m x p_k^* K_i \quad (9.4)$$

$$x p_k^* = p_m = p_k \quad (9.5)$$

$$I_g = \left[ x \frac{p_a^*}{p_m} - p \right] E_a = \left[ \frac{p_a^*}{p_k^*} - p \right] E_a \quad (9.6)$$

$$p = \frac{p_a}{p_m} \quad (9.7)$$

Saving Constraint:

$$\frac{w_m}{p_m} = (1 - \beta)\omega_a p \quad (9.8)$$

$$\omega_a l_a = f[(1 - \beta)\omega_a, p] + E_a, \quad f_1 > 0 \text{ and } f_2 < 0 \quad (9.9)$$

$$g^s = \left[ 1 - (1 - \beta)\frac{\omega_a}{q_m} p \right] \sigma_m + \beta \frac{\omega_a l_a}{K_m} p + \kappa \quad (9.10)$$

Foreign Exchange Constraint:

$$g^e = \frac{1}{m} \left( \frac{p_a^*}{p_k^*} e_a + \kappa \right) \quad (9.11)$$

Demand Constraint:

$$w_m > w_a \quad (9.12)$$

$$g^i = g^i(r_m) = h\left(\frac{w_m}{p_m}\right), \quad h' < 0 \quad (9.13)$$

$$g^i = \left[ 1 - \frac{w_m}{p_m} \frac{1}{q_m} \right] \sigma_m + \beta \frac{\omega_a l_a}{K_m} p + \kappa \quad (9.14)$$

$$\omega_a l_a = f[(1 - \beta)\omega_a, p] l_a + f\left[\frac{w_m}{p_m} \frac{1}{p}, p\right] (1 - l_a) + E_a \quad (9.15)$$

Fiscal Constraint:

$$g^i = \begin{cases} h\left(\frac{w_m}{p_m}\right) & \text{if } h\left(\frac{w_m}{p_m}\right) \leq \alpha \frac{I_g}{K_m} \\ \alpha \frac{I_g}{K_m} & \text{if } h\left(\frac{w_m}{p_m}\right) > \alpha \frac{I_g}{K_m} \end{cases} \quad (9.16)$$