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AGRICULTURAL WAGES AND FOOD PRICES IN EGYPT: A GOVERNORATE-LEVEL ANALYSIS FOR 1976-1993

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ABSTRACT

The trend in real agricultural wages in Egypt is described well by an inverted U-shaped curve with a peak around 1985. But the rise and fall of real wages masks a complex dynamic process by which nominal wages adjust in response to changes in food prices. We use governorate-level panel data for 1976–1993 to explore the nature of this adjustment process. Our results indicate that nominal wages adjust slowly. There is a significant negative initial impact of rising food prices on real wages, though wages do catch up in the long run.

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1. INTRODUCTION

How quickly and how far do wages adjust to changes in food prices? This is an old question, yet, there is relatively limited empirical work that sheds light on this issue for developing countries.[1] Apart from an obvious interest in this question from a labor market perspective, it is also of great relevance to assessing the distributive impact of food price policy. It has often been noted that the distributional effects of changes in food prices depend critically on the assumed model of wage determination (see de Janvry and Subbarao (1984) and Sah and Stiglitz (1987), for instance). Indeed, there is considerable evidence to suggest that agricultural wages are often an important determinant of rural and hence, national poverty.[2] Yet, wage determination models in the literature range the full spectrum from fixed wage models to others with complete wage flexibility, and the choice of an appropriate model remains contentious for the labor markets of most developing countries.[3]

This question is also of particular interest for Egypt. First, it is a question that has remained largely unaddressed for Egyptian labor markets, especially rural labor markets,

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[1] Despite the general dearth of empirical work on this topic, a useful analysis for rural Bangladesh can be found in Ravallion (1987) and Boyce and Ravallion (1991).

[2] For time-series evidence on the importance of agricultural wages as a determinant of rural poverty in India, see Datt and Ravallion (1998).

[3] For a critical review of alternative models of wage determination for rural labor markets in developing countries, see Datt (1996).

despite a long-standing tradition of work on agricultural wages in Egypt.[4] But, more notably, this question has added significance in the context of the current debate on the reform of the Egyptian food subsidy system. An important element of this debate concerns the policy options for the government to reduce its food subsidy budget with minimal adverse welfare consequences for the poor. The welfare effects of food subsidy changes are not limited to just the direct consumption effects. It is also important to consider the induced income effects such as those operating through the wage response in the labor market. There is also the related issue of the extent to which the food subsidy operates like a wage subsidy to the employers. Two additional concerns are (1) will a reduction in subsidy lead to a parallel increase in nominal wages, thus eroding the international competitiveness of Egyptian products? (2) Alternatively, if nominal wages are sticky, is a reduction in subsidy more likely to lead to political unrest? Answers to these questions depend on the nature and speed of the wage adjustment process. It is also important in this context to distinguish between the short- and long-run wage responses.

The identification of an appropriate model of the wage adjustment process is largely an empirical issue. Given that wage adjustment mechanisms are inherently dynamic processes, their successful modeling critically depends upon the availability of long-term data on wages and potential wage determinants, including food prices. Fortunately, such

[4] This goes back to some of the early work by Hansen (1966, 1969) who, for instance, questioned the usefulness of the subsistence wage theory in explaining Egyptian agricultural wages.

data exist for Egypt, and despite some limitations, it is possible to collate these data for such an analysis.

In this paper, we use these data to estimate a dynamic (panel data) model of the determination of agricultural wages at the governorate level. The nominal agricultural wage in a given governorate and time period will thus be estimated as a function of current and past values of a number of variables, including inter alia food and nonfood prices, agricultural and nonagricultural productivity, workers' remittances from abroad, and a measure of labor supply. The model will be used to study the nature of the agricultural wage adjustment process, and to identify, in particular, the short- and the long-run response of nominal agricultural wages to changes in food prices. The paper is organized as follows. In the next section, we review an antecedent in the literature to illustrate some key issues that deserve to be addressed in an analysis of the wage-food price relationship. This discussion is intended to motivate the analytical approach adopted in our study. Section 3 describes our wage data and presents the unconditional trends in real wages by governorate. The specification of our agricultural wage model is discussed in Section 4. Section 5 discusses model estimation issues. Our tests for contemporaneous, short- and long-run homogeneity conditions are presented in Section 6, while Section 7 discusses results from our preferred estimates of the econometric model. In Section 8, we present a simulation on the wage impact of food price changes. Section 9 discusses some caveats and extensions, and some concluding observations are offered in the final section.

2. AN ANTECEDENT IN THE LITERATURE

While there is a sizable literature on trends in agricultural wages for Egypt, most of the literature is descriptive in nature. There has been surprisingly little work on the modeling of agricultural wages in Egypt. One exception is de Janvry and Subbarao (1983) which, though dated, is an obvious point of departure for our study. De Janvry and Subbarao used cross-sectional (inter-governorate) data to estimate agricultural wage functions. Their best estimate of the wage function was the following:

$$\begin{array}{l} w^{m} = 19.28 + \underset{(3.68)}{0.79}\,EMGO + \underset{(1.97)}{3.19}\,WHPR - \underset{(-0.32)}{0.003}\,LABM \\ \qquad n = 15, \qquad R^2 = 0.68 \end{array}, \qquad \textbf{(1)}$$

where w^m is the average male wage rate in the governorate during 1974-78, *EMGO* is a measure of emigration of labor from rural areas, *WHPR* is the price of wheat, and *LABM* is a cropping pattern weighted index of demand for male labor (*t*-ratios in parentheses). At the sample means, their estimates indicate a wheat price elasticity of the nominal male wage of 0.5, implying that about 50 percent of the increase in wheat prices is passed on in higher money wages. There are several reasons to be cautious about interpreting their results.[5]

[5] To be fair to de Janvry and Subbarao, the investigation of the wage-price relationship is not the singular focus of their study. Nevertheless, a discussion of this study is useful for motivating some key features of our approach.

First, de Janvry and Subbarao's is a cross-sectional study, and their results are best interpreted as estimates of the short-run wage response. The failure of nominal wages to catch up with changes in food (wheat) prices does not appear to be a highly probable description of the steady-state equilibrium in the agricultural wage labor market. Wage adjustment processes are typically sluggish in nature, and can thus entail potentially large differences between the short- and long-run responses. Cross-sectional studies, such as de Janvry and Subbarao's, are, by construction, unable to disentangle the short-run from the long-run effects. Yet the ability to isolate these effects can be an important element in understanding the welfare consequences of proposed changes in food price policy, and hence a useful guide to how, if at all, such policy changes should be phased in.

The second issue relates to the fact that de Janvry and Subbarao's estimates are based on only 15 observations. Apart from contributing to the imprecision of the estimated parameters, the limited number of observations also constrained the range of wage determinants the authors could introduce into their analysis, thus potentially exposing their estimates to omitted variable bias. Rural labor markets are typically segmented; the failure to control for relevant regional factors (both observed and unobserved) can vitiate results on the estimated wage response to price changes.[6]

We hope to address some of these concerns in this paper.

[6] We refer to segmentation here in order to stress the need for adequately allowing for local/regional factors in wage determination. The importance of local factors is quite unexceptional in the context of labor markets in most settings, including those in the relatively developed countries. For rural labor markets in developing countries, due to a number of informational, infrastructural and/or institutional constraints, one could expect a lower order of spatial integration of the labor market, and to that extent a greater influence of local factors in wage determination.

3. AGRICULTURAL WAGE DATA AND UNCONDITIONAL TRENDS

Our data on agricultural wages come from the Ministry of Agriculture and Land Reclamation (MALR) and were complied by the Agricultural Economics Research Institute (MALR, Cairo).[7] These data were collated at the governorate level. The data are for 18 governorates in Egypt, and span the period 1976-1993, although the exact period covered varies by governorate. Table 1 shows the governorates included in the study and the period covered for each of them.

Although the original wage data were available on a bi-monthly (twice a month) basis, we aggregated the data up to three observations per year (corresponding to four-month periods) using simple averages. This aggregation was motivated by several considerations. First, for some governorates, there were a number of missing values in the original wage data, and aggregation over a longer period enabled us to plug many of these data gaps. Second, the averaging was also motivated by a desire to attenuate random measurement error in the reported wage data. Finally, the rural food and general Consumer Price Index (CPI) data from the Central Agency for Public Mobilization and Statistics (CAPMAS) are only available once every two months, and data on most other potential determinants of agricultural wages are available only on an annual basis. Thus, the gains from additional temporal disaggregation of the wage data were quite limited.

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[7] This has been the key source of agricultural wage data for Egypt, and has been used in most studies on agricultural wages, including Fitch, Ali, and Mostafa (1980), de Janvry and Subbarao (1983), Assaad and Commander (1994), and Richards (1994).

Table 1 Pattern of real wage growth across governorates

	Governorate	Estimation period	Turning point	Average rate of growth up to 1985 (percent per year)	Average rate of growth after 1985 (percent per year)
	Lower Egypt				
1	Alexandria	1981–1993	1985.5	6.5	–9.9
2	Behera	1976–1993	1985.4	10.3	–8.2
3	Gharbia	1976–1993	1985.1	9.4	–8.4
4	Kafr El-Sheikh	1976–1985		11.3	
5	Dakahlia	1976–1993	1984.1	7.4	–10.5
6	Damietta	1976–1993	1985.2	11.4	–9.9
7	Sharkia	1976–1990	1985.3	13.2	–6.1
8	Ismailia	1976–1990	1984.6	8.0	–5.7
9	Menoufia	1976–1993	1984.2	9.6	–13.2
10	Kalyoubia	1976–1985		9.7	
	Upper Egypt				
11	Giza	1976–1993	1984.7	8.6	–9.3
12	Beni-Suef	1976–1985		7.9	
13	Fayoum	1976–1993	1985.4	7.3	–5.8
14	Menia	1976–1993	1983.5	6.2	–12.2
15	Asyout	1976–1993	1984.4	8.9	–11.3
16	Suhag	1976–1985		8.9	
17	Qena	1976–1985		5.8	
18	Aswan	1976–1985		6.7	

Note: The average growth rates are derived from the estimated parameters of model (2). The time trends (linear or quadratic) for all governorates were highly significant.

Our aggregated wage data thus comprised of three observations per year corresponding to the three "seasons" for the months of December-March, April-July, and August-November, respectively.[8]

The real daily agricultural wage rates for the nine governorates for which we have a complete time series for the full period (1976–1993) along with the rural CPI are graphed in Appendix 2, Figures 2 to 10. The figures suggest only limited regional diversity. In Table 1, we present evidence on unconditional trends in real agricultural wages,[9] allowing for quadratic time trends with seasonal dummy variables, estimated as follows:

$$\tilde{w}_{jt} = \alpha_j + \beta_{1j} t + \beta_{2j} t^2 + \delta_{2j} SEAS2 + \delta_{3j} SEAS3 + \epsilon_{jt} , \qquad \textbf{(2)}$$

where $j = 1, \ldots 18$, and $t = 1976(1), \ldots 1993(2)$; $\tilde{w}_{jt}$ is the natural logarithm of the *real* daily agricultural wage (nominal wage deflated by the General Rural Consumer Price Index) at date t in governorate j; t and t^2 are linear and quadratic time trends; *SEAS*2 and *SEAS*3 are dummy variables taking values of unity for the months of April-July and August-November, respectively, and zero otherwise; and ϵ_{jt} is a governorate-specific disturbance term.

Table 1 shows that the quadratic terms in time trends were not significant for a number of governorates, including Kafr El-Sheikh, Kalyoubia, Suhag, and Aswan. But

[8] Season 1 corresponds fairly closely to the lean season, which generally lasts from December to February.

[9] The term "unconditional" refers to trends without controlling for any wage determinants.

these are also the governorates with usable wage data only up to 1985. Similarly, in the case of two other governorates for which we have wage data only up to 1985, Beni-Suef and Qena, we find that both the linear and quadratic trends are positive, although only the quadratic trends are significant. For all the other 12 governorates for which we have usable wage data beyond 1985, we find that linear time trends are significant and positive, while the quadratic terms are significant and negative. The implied turning points in real wages are all internal to our estimation period. The turning points for the 12 governorates are also shown in Table 1. It is notable that they all lie within the narrow interval between mid-1983 and mid-1985. Most of the real wage turning points are clustered around 1985, including those for governorates with a shorter estimation period of up to 1990 only.

This finding on the turning points in real agricultural wages is consistent with similar observations on wage trends in the literature, albeit mostly made at the national level.[10] However, while the similarity across governorates in the real wages turning points (and in wage trends in general) suggests the operation of some common determining forces, one should be careful in interpreting this as evidence for a high level of spatial integration of the agricultural labor market.

The intra-year variation in wages gives us an opportunity to look into seasonal effects. There is only limited evidence of seasonality in wages. Using a restricted version of model (2) with common effects for the April-July and August-November seasons, we

[10] See Assaad and Commander (1994) and Richards (1994), for instance.

found seasonal effects to be insignificant. However, there is some weak evidence of declining seasonality over time.[11] In a model incorporating both seasonal effects and season-time interactions, we found that at the start of our period, wages were significantly higher (by 3 percent on average) during the April-July and August-November periods (a significant, positive common effect for these seasons); this is consistent with the known pattern of seasonality in the demand for agricultural labor.[12] But we also found the season-time interactions to be negative, though not significant. Upon eliminating the season-time interactions, the seasonal effects became altogether insignificant.

4. SPECIFICATION OF THE MODEL

While we are primarily interested in the relationship between agricultural wages and food prices, to correctly identify that relationship it is important to control for other determinants of wages. Thus, our model of agricultural wages includes variables reflecting conditions on both the labor demand and the supply side. Besides the price variables (described further below), our vector of explanatory variables consists of the following:

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[11] This has been sometimes noted in the literature; see, for instance, Richards (1994).

[12] Season 1 corresponds fairly broadly to the lean season, which generally lasts from December to February (Commander and Hadhoud 1986).

YLD : yield per feddan (value of output of 10 major crops at constant prices per feddan of cropped area);

AREA : total cropped area in the governorate for all crops;

POP : total population of the governorate;

YPUB : the value of public sector industrial output per capita in the governorate normalized by the rural food price index;

YPVT : the value of private sector industrial output per capita in the governorate normalized by the rural food price index;

XR : the (nominal) exchange rate;

REMIT : the value of workers' remittances from abroad in constant LE (normalized by the rural food price index);

SEAS2 : a seasonal dummy variable assuming the value 1 for April-July, and 0 otherwise;

SEAS3 : a seasonal dummy variable assuming the value 1 for August-November, and 0 otherwise.

The rationale for the inclusion of these variables in the agricultural wage model is briefly described.

The yield variable is included to capture the direct and indirect effects of agricultural productivity on labor demand, and hence on wages. However, due to data limitations, the coverage of our yield variable had to be limited to ten crops only (accounting for about 55 percent of national cropped area during 1993).[13] We, therefore, also included the total cropped area (for all crops) in the governorate as an explanatory variable to pick up any additional labor demand effects. We also allow for nonagricultural sources of labor demand. This is done by including measures of industrial output among the explanatory variables. We distinguish between public- and private-sector output, thereby allowing for potentially differential effects of output growth in the two sectors. For the most part, the industrial output variables are measures of economic activity in the formal sector, and hence should be interpreted as measures of the formal (nonagricultural) sector labor demand.[14]

On the supply side, the governorate population is used as a proxy for the rural labor force. Time-series data on the size of the labor force are not available even at the national level. However, even if those data were available, it is arguable that agricultural labor market conditions, including the wage rate, would have influenced the size of the rural labor force. The total population of the governorate, on the other hand, is more likely to be uncorrelated with the model's error term.

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[13] CAPMAS (1995). For further details on the construction of this yield index and the data sources used, see Appendix 1. The underlying area, production, and price data have also been used by Rady, Omran, and Sands (1996), who provide more details on the data.

[14] Due to lack of data, we are unable to include variables representing labor demand originating in the informal sectors, such as services and construction. But, to the extent that economic activity in these sectors co-varies with that in the industrial sector, the public and private industrial output variables would serve as potential proxies.

International migration of labor has been widely argued to be an important influence on labor market outcomes in Egypt.[15] However, no time series-data on migration are available at the governorate or even the national level, but we allow for migration effects by including in our model a measure of workers' annual remittances from abroad. Needless to say, this is only a second-best solution to the limitations of available data. The remittances are expressed in real Egyptian pounds (LE). Thus, the effects of exchange rate changes, especially the depreciation since the late 1980s, are already reflected in the measure of remittances. We nonetheless also include the exchange rate as an additional explanatory variable to allow for labor market effects other than those occurring through workers' remittances.

We also allow for seasonal effects, and a time trend to capture omitted but trended variables, including those associated with secular changes in the macroeconomic environment. The specification of a common time trend across governorates appears justified in the light of our results on the unconditional wage trends discussed above, indicating similar trends across governorates (see Section 3).

It is arguable that the above set of explanatory variables nevertheless omits several potential wage determinants both on the demand and the supply side, for instance, human capital and infrastructural development differentials across governorates. We try to address this problem by exploiting the panel aspect of our data to allow for unobserved

[15] See, for instance, de Janvry and Subbarao (1983), Commander and Hadhoud (1986), Commander (1987), Adams (1991), Fergany (1991), Richards (1994), and Serageldin and Wouters (1996).

governorate-specific determinants of wages. The ability to allow for such cross-sectional effects is important even for data-rich settings, as it is seldom possible to adequately account for a potentially large set of wage determinants using observable data. But it is particularly important for our application, given the current state of available data for Egypt, some of whose limitations have already been discussed above.[16]

Finally, by allowing the current wage to depend on lagged wages, our model also incorporates sluggishness in the wage adjustment process that is typical of the labor market response in most settings.

Incorporating the considerations discussed above, we model agricultural wages as an autoregressive process within a dynamic panel data framework. We begin with a fairly general autoregressive distributive lag (AD) formulation.[17] In particular, we start with an AD(4,4) specification of the model allowing for 4th-order lags in both the dependent and independent variables.[18]

$$w_{jt} = \alpha_0 + \sum_{i=1}^{4} \alpha_i w_{jt-i} + \sum_{i=0}^{4} \beta_i p^{f}_{jt-i} + \sum_{i=0}^{4} \gamma_i^{\prime} x_{jt-i} + \delta_0 t + \delta_1 t^2 + u_{jt}, \qquad \textbf{(3)}$$

[16] For further discussion of the implication of some of these data limitations, also see section 9 below.

[17] See Hendry (1995) on the generality of even a simple dynamic specification such as AD(1,1), which nests a wide variety of empirical dynamic models as special cases.

[18] We initially began with an AD(3,3) formulation, which seemed a natural choice, given that our data set has three observations per year, but residual autocorrelation led us to introduce an additional lag.

where $u_{jt} = \eta_j + \nu_{jt}$ for $j = 1, ..., 18$; $t = 1, ..., 53$; and where w_{jt} is the natural log of nominal daily agricultural wage rate for governorate j and time t, p_{jt}^{f} is the natural log of the rural food price index, x_{jt} is a vector of explanatory variables, and the error term u_{jt} consists of η_j, the unobserved governorate-specific effect, and a random (time and governorate-varying) component v_{jt}. All variables are measured in natural logarithms.

A theoretical motivation for a dynamic specification of the wage function is readily provided. For instance, such a model could be derived from a competitive model of the labor market, and a partial adjustment hypothesis, along the following lines.[19] Let us assume, without loss of generality, that labor demand (L^d) is a function of the wage rate (w) and a set of nonwage factors, Y, determining labor demand, and similarly labor supply (L^s) depends on the wage rate and a set of nonwage determinants, Z. One could thus think of the competitive equilibrium wage (w^*) determined by the market clearance condition ($L^d = L^s$):

$$w_t^* = w^*(Y_t, Z_t)\,. \qquad \textbf{(4)}$$

Now let's assume (realistically) that adjustment to the competitive equilibrium takes time and there is only partial adjustment during the same time period. One could postulate that the speed of wage adjustment depends on the level of excess demand, with wages increasing in response to positive excess demand. Thus,

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[19] The following discussion closely follows the formulation in Ravallion (1987).

$$w_{t+1} - w_t = f(L^d(w_t, Y_t) - L^s(w_t, Z_t)) \quad where\ f' > 0\ . \qquad \textbf{(5)}$$

Or, quite generally,

$$w_{t+1} = g(w_t, Y_t, Z_t)\ . \qquad \textbf{(6)}$$

On introducing additional lags in the wage adjustment process (which, in any case, is largely an empirical issue), and on linearizing equation (6), we obtain a model analogous to model (3) above. A further point should also be noted: equation (6) does not presuppose competitive market clearance, and is also compatible with some noncompetitive models of wage determination. For example, the equilibrium wage in equation (4) could be determined as a bargaining equilibrium, and costs of renegotiation could easily justify a less-than-complete adjustment in the short run, leading to a wage equation such as equation (6).[20]

While our primary concern is to examine food price effects, we also considered including the rural CPI for all commodities or the (derived) rural CPI for nonfood items to explore the possibility of relative price effects on wages.[21] However, this proved to be

[20] For various characterizations of the agricultural labor market in the Egyptian case, see de Janvry and Subbarao (1983), Grabowski and Sivan (1986), Commander (1987), and Richards (1994).

[21] The nonfood component of the rural CPI was derived from the food and general indices using the average rural food consumption share 64.314 percent obtained from the Income and Expenditure Survey 1990-91(calculated from the weighting diagram reported in CAPMAS (1996)).

infeasible as the different components of the CPI were virtually collinear. Denoting as p and p^n the natural logarithms of the rural general and nonfood price indices, respectively, the correlations between these indices were *corr* (p^f, p) = 0.999; *corr* (p^f, p^n) = 0.993; and *corr* (p, p^n) = 0.997. Or, put differently, in expectation, p and (p^f (as also p and p^n) differ by a constant. It is because of this extreme collinearity that we cannot identify a *relative* food price effect on the nominal wage rate. It also implies that beginning with a general wage model (suppressing the lags) such as

$$w_{jt} = constant + \beta_1 p_{jt} + \beta_2 p_t^f + \gamma' x_{jt} + \ldots + u_{jt}, \qquad \textbf{(7)}$$

we can respecify that model as

$$w_{jt} = \alpha_0 + \beta p_{jt}^f + \gamma' x_{jt} + \ldots + u_{jt}, \qquad \textbf{(8)}$$

where $\beta = \beta_2 + \beta_1$, which is analogous to model (3).

Before moving on to estimation issues, it is also useful to note that while all explanatory variables in our wage model are time-varying, not all of them vary by governorate. In particular, the exchange rate, the rural food price index, and the remittance variables are constant across governorates. The exchange rate is, of course, a national variable, but for the other two variables, we are constrained by the available data that do not permit regional disaggregation. (The implications of this for our results are discussed in Section 9.) A detailed discussion of the data sources and the construction of the model variables can be found in Appendix 1. As mentioned before, our final data set consists of an unbalanced panel for 18 governorates. The minimum number of time observations for any governorate is 30 and the maximum is 53.

5. MODEL ESTIMATION

Given that the correlation between w_{it} and η_j also implies a correlation between w_{it-I} and η_j, none of the usual estimators (including ordinary least squares, fixed or random effects estimators) yields consistent estimates for model (3). This is a standard result for panel data models with lagged dependent variables (see Baltagi 1995, for instance). One approach to consistent estimation can be based on the generalized method of moments (GMM) estimator for dynamic panel data models proposed by Arellano and Bond (1991). This estimator involves using a differenced version of the above model:

$$\Delta w_{jt} = \delta_0 + \sum_{i=1}^{4} \alpha_i \Delta w_{jt-i} + \sum_{i=0}^{4} \beta_i \Delta p^f_{jt-i} + \sum_{i=0}^{4} \gamma_i^{/} \Delta x_{jt-i} + \delta_1 (2t - 1) + \Delta v_{jt}, \qquad \textbf{(9)}$$

where the operator Δ denotes first differencing. Notice that while first-differencing eliminates the unobserved governorate-specific effect, it induces a first-order moving average (MA(1)) structure for the transformed error $\Delta\ v_{jt}$. Model (3) can be consistently estimated using a set of moment conditions. If the original errors v_{jt} in the levels model (2) are not serially correlated, valid moment conditions can be based on the lagged values

w_{jt-5} and other exogenous variables.[22] It is important to note that the consistency of the GMM estimator depends heavily on the assumption that there is no second-order serial correlation in the errors of the differenced model, i.e., $E(\Delta v_{jt}\, \Delta v_{jt-2}) = 0$. We will test this condition for our estimates reported below.

We follow a fairly conservative approach in our treatment of the other explanatory variables allowing for potential endogeneity of all variables in the *x*-vector (except, of course, for the seasonal dummies). These variables are treated as predetermined rather than strictly exogenous.[23] Thus, we only include 5th-order lags of the explanatory variables in the instrument set. The instrument set also includes the moment conditions related to the lagged dependent variable.

The estimates of model (3) are reported in Appendix 3 (Table A3.1). We will not discuss these initial estimates in detail here, but it is useful to point out that the consistency requirement of zero second-order autocorrelation is satisfied. Similarly, the Sargan test for overidentifying restrictions is also accepted.[24] However, the large number of estimable parameters associated with the AD(4,4) formulation does induce a loss in the precision of

[22] See Arellano and Bond (1991) and Sevestre and Trognon (1996) for further discussion of this estimator for dynamic panel data models. The GMM estimator has been set up as a GAUSS-based program; further details on the implementation of the estimator are given in Arellano and Bond (1988).

[23] Thus, the variables x_{jt} are not only allowed to be correlated with the unobserved governorate-specific effects η_j, but they are also allowed to be contemporaneously correlated with the random error v_{jt}.

[24] This test is based on the covariance between IV residuals and a set of instruments that need not have been used in the estimation. This covariance should be zero if the model is correctly specified, and the choice of instruments is valid.

the estimates. We thus proceed to test a number of data-consistent restrictions on the model leading to a more parsimonious specification.

In particular, we tested for zero-parameter restrictions on insignificant model parameters. This was done in two steps. We first tested the exclusion of all parameters that had absolute t-ratios of less than 0.5. Sixteen parameters were thus found to be jointly insignificant. On excluding these parameters, we further tested for exclusion restrictions on parameters that had absolute t-ratios of less than unity. Four additional parameters were thus excluded at the second stage. The 20 deleted parameters related to the following variables: lagged nominal wages, Δw_{-2}, Δw_{-4}; lagged food price index Δp^f_{-3}; current and lagged yield per feddan, ΔYLD_{-1}, ΔYLD_{-3}; current and lagged total cropped area, $\Delta AREA$, $\Delta AREA_{-1}$, $\Delta AREA_{-3}$; lagged total population, ΔPOP_{-3}; current and lagged public-sector industrial output per person, $\Delta YPUB$, $\Delta YPUB_{-3}$; lagged private-sector industrial output per person, $\Delta YPVT_{-1}$, $\Delta YPVT_{-2}$, $\Delta YPVT_{-3}$; lagged real exchange rate, ΔXR, ΔXR_{-1}; lagged real remittances, $\Delta REMIT_{-1}$, $\Delta REMIT_{-2}$; and seasonal dummy variables, $\Delta SEAS2$, $\Delta SEAS3$.[25] These restrictions were readily accepted, and on imposing these restrictions, we obtain the estimates reported in Table 2. The bottom of Table 2 also reports the joint (Wald) tests of these restrictions.

[25] The parameters excluded at the second stage were for the following variables: $\Delta AREA_{-3}$, $\Delta YPUB$, $\Delta YPVT_{-2}$, and $\Delta REMIT_{-2}$.

Table 2   Dynamic panel data estimates of the agricultural wage model

Variable	Parameter estimate	t-statistic
Constant	–0.00675	–0.396
Lagged nominal wage		
Δw_{-1}	0.44585	5.060
Δw_{-3}	0.38087	4.442
Current and lagged food price index		
Δp^f	0.16212	1.678
Δp^f_{-1}	0.31587	2.644
Δp^f_{-2}	–0.37751	–3.678
Δp^f_{-4}	0.43975	3.311
Current and lagged yield per feddan		
ΔYLD	0.18440	1.290
ΔYLD_{-2}	–0.19131	–1.199
ΔYLD_{-4}	0.17496	1.162
Current and lagged total cropped area		
$\Delta AREA_{-2}$	1.49920	3.020
$\Delta AREA_{-4}$	–1.00183	–1.979
Current and lagged total population		
ΔPOP	–1.51529	–3.081
ΔPOP_{-1}	1.49461	3.204
ΔPOP_{-2}	–0.52410	–1.740
ΔPOP_{-4}	0.60868	1.922
Current and lagged public-sector industrial output per person		
$\Delta YPUB_{-1}$	0.10753	2.105
$\Delta YPUB_{-2}$	–0.07473	–1.602
$\Delta YPUB_{-4}$	0.11354	2.914
Current and lagged private-sector industrial output per person		
ΔYPVT	–0.03590	–0.935
$\Delta YPVT_{-4}$	0.11312	3.432
Current and lagged real exchange rate		
ΔXR_{-2}	–0.06126	–1.920
ΔXR_{-3}	–0.22401	–4.073
ΔXR_{-4}	0.28331	5.767
Current and lagged real remittances		
ΔREMIT	–0.03769	–1.776
$\Delta REMIT_{-3}$	0.15643	4.359
$\Delta REMIT_{-4}$	–0.08635	–2.719
Time trend		
(2 t – 1)	–0.00033	–2.333
Sargan's test: df = 34	19.694	p = 0.973
Test for first-order serial correlation df = 18	–4.583	p = 0.000
Test for second-order serial correlation df = 18	0.382	p = 0.703
Wald tests for zero parameter restrictions: df = 16	2.202	p = 1.000
df = 4	1.568	p = 0.815

Note: These are parameter estimates for model (4). Number of governorates = 18. Number of observations = 695. The instrument set includes {1; instruments based on GMM conditions $E(w_{jt-5}\,\Delta v_{jt}) = 0$; p^f_{-5}; YLD_{-5}; $AREA_{-5}$; $POPT_{-5}$; $YPUB_{-5}$; $YPVT_{-5}$; XR_{-5}; $REMIT_{-5}$; $SEAS2$; $SEAS3$; t; t^2}. See text for discussion of the tests for zero parameter restrictions. All test statistics are distributed as χ^2 with degrees of freedom as noted.

6. TESTING HOMOGENEITY RESTRICTIONS AND THE PREFERRED ESTIMATES

Alternative hypotheses regarding the speed and completeness of the wage adjustment process can be introduced in the form of three different homogeneity conditions on the nominal wage model. We refer to these as the contemporaneous, the short-run, and the long-run homogeneity conditions.[26]

Contemporaneous homogeneity implies that increases in food prices are *fully* passed on to nominal wages during the same time period. In our model, contemporaneous effects refer to those occurring within the same four-month period. Contemporaneous homogeneity therefore requires

$$\beta_0 = 1, \quad \alpha_1 = \alpha_2 = \alpha_3 = \alpha_4 = \beta_1 = \beta_2 = \beta_3 = \beta_4 = 0 . \qquad \textbf{(10)}$$

Short-run homogeneity implies complete adjustment in the short run, where the *short run* in our model is identified with a 16-month period. A 16-month periodicity is implied by our specification because it uses lags up to the 4^{th}-order, where each lag has a four-month duration. Thus, short-run homogeneity is satisfied if

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[26] Testing for the homogeneity condition as a property of the equilibrium wage is common in the applied labor market literature. The condition is grounded in the notion that in equilibrium the *real* wage is determined by *real* variables. Economic theory suggests this property of the equilibrium wage insofar as labor demand and supply functions are homogenous of degree zero in all prices and nominally-expressed variables such as unearned incomes (i.e., only relative prices matter). Or equivalently, the market *nominal* wage is homogeneous of degree one in all prices and nominally-expressed variables. In our model, this amounts to homogeneity of the nominal wage in current and lagged food prices and lagged wages. The other variables are already expressed in real terms. The exchange rate, though "nominal," is already, by definition, a relative price, i.e. the relative price of the domestic to the foreign currency.

$$\alpha_1 = \alpha_2 = \alpha_3 = \alpha_4 = 0, \quad \Sigma_{i=0}^{4} \beta_i = 1 . \tag{11}$$

Long-run homogeneity implies complete adjustment only in the long run. For our model, this is written as the condition

$$\sum_{i=1}^{4} \alpha_i + \sum_{i=0}^{4} \beta_i = 1 . \tag{12}$$

This is a weaker condition than *short-run* and *contemporaneous* homogeneity conditions. Contemporaneous homogeneity implies short-run homogeneity, which, in turn, implies long-run homogeneity, but *not* vice versa.

To facilitate discussion of further results, it is useful to note that model (2) can also be rewritten somewhat differently to permit a convenient test of the long-run homogeneity restriction:

$$\begin{aligned} \Delta w_{jt} = \alpha_0 &+ (\alpha_1 - 1)\Delta w_{jt-i} + (\alpha_1 + \alpha_2 - 1)\Delta w_{jt-2} \\ &+ (\alpha_1 + \alpha_2 + \alpha_3 - 1)\Delta w_{jt-3} + (\alpha_1 + \alpha_2 + \alpha_3 + \alpha_4 - 1)(w_{jt-4} - p^f_{jt-4}) \\ &+ \sum_{i=0}^{3} \left(\sum_{k=0}^{i} \beta_k \, \Delta p^f_{jt-i} \right) + \left(\sum_{i=1}^{4} \alpha_i + \sum_{i=0}^{4} \beta_i - 1 \right) p^f_{jt-4} \qquad , \\ &+ \sum_{i=0}^{4} \gamma_i' x_{jt-i} + \delta_0 t + \delta_1 t^2 + u_{jt} \end{aligned} \tag{13}$$

where $u_{jt} = \eta_j + v_{jt}$. The long-run homogeneity condition can thus be easily tested as a zero parameter restriction on the coefficient for p^f_{jt-4} , i.e.,

$$\Sigma_{i=1}^{4}\alpha_i + \Sigma_{i=0}^{4}\beta_i = 1 .$$

Notice that the earlier estimates in Table 2 already imply that the parameters α_2, α_4, and β_3 are zero. Incorporating these restriction, model (5) can be rewritten

$$\begin{aligned}\Delta w_{jt} = \alpha_0 &+ (\alpha_1 - 1)(w_{jt-i} - w_{jt-3}) + (\alpha_1 + \alpha_3 - 1)(w_{jt-3} - p^f_{jt-4}) + \beta_0 \Delta p^f_{jt} \\ &+ (\beta_0 + \beta_1)\Delta p^f_{jt-1} + (\beta_0 + \beta_1 + \beta_2)(p^f_{jt-2} - p_{jt-4}) \\ &+ (\alpha_1 + \alpha_3 + \beta_0 + \beta_1 + \beta_2 + \beta_4 - 1)p^f_{jt-4} + \sum_{i=0}^{4}\gamma_i^{\prime} x_{jt-i} + \delta_0\, t + \delta_1\, t^2 + u_{jt}\end{aligned} \quad , \qquad \textbf{(15)}$$

where $u_{jt} = \eta_j + v_{jt}$.

An estimable form of the model is derived (as before) by applying first-differences to equation (6), thus yielding

$$\begin{aligned}\Delta^2 w_{jt} = \delta_0 &+ (\alpha_1 - 1)\Delta(w_{jt-i} - w_{jt-3}) + (\alpha_1 + \alpha_3 - 1)\Delta(w_{jt-3} - p^f_{jt-4}) \\ &+ \beta_0 \Delta^2 p^f_{jt} + (\beta_0 + \beta_1)\Delta^2 p^f_{jt-1} + (\beta_0 + \beta_1 + \beta_2)\Delta(p^f_{jt-2} - p^f_{jt-4}) \\ &+ (\alpha_1 + \alpha_3 + \beta_0 + \beta_1 + \beta_2 + \beta_4 - 1)\Delta p^f_{jt-4} + \sum_{i=0}^{4}\gamma_i^{\prime}\Delta x_{jt-i} \\ &+ \delta_1(2t - 1) + \Delta v_{jt}\end{aligned} \quad , \qquad \textbf{(16)}$$

where the operator Δ denotes first-differencing, while Δ^2 denotes second-differencing. Our estimates of the model in this form are shown in Table 3, after further restricting the model to exclude the insignificant parameter associated with the current food price term $\Delta^2 p^f_{jt}$.

We find that contemporaneous homogeneity is strongly rejected by our data (see the test reported at the bottom of Table 3). Thus, static formulations of the wage-food price relation, such as that used by de Janvry and Subbarao (1983), find no support in our data. The lagged effects are important, and an appropriate formulation of the wage adjustment process therefore warrants a dynamic specification.

Our results also indicate a clear rejection of short-run homogeneity, implying that increases in food prices are *not* fully passed on to nominal wages *in the short run*, i.e., within a 16-month period. In fact, we find that the short-run, wage-food price elasticity is well below unity with a point estimate of 0.27. The rejection of contemporaneous and short-run homogeneity implies that the functioning of the agricultural labor market does not fully insulate the agricultural workers' real wages against food price increases.

A key result of Table 3 relates to the proposed test of long-run homogeneity of the real agricultural wage with respect to changes in food prices (equation [8]). The estimated parameter for Δp^f_{jt-4} is 0.24, and it is statistically insignificant; we are unable to reject long-run homogeneity at better than 15 percent level of significance. Thus, despite their sluggish response, nominal wages do catch up with higher food prices in the long run.

Our preferred estimates of the agricultural wage model are derived by imposing long-run homogeneity and also setting the insignificant parameter on $\Delta(p^f_{jt-2} - p^f_{jt-4})$ to zero. The Wald test for these two restrictions is easily accepted (see Table 3). The last two columns of Table 3 present our preferred estimates. Before we discuss these

Table 3 Dynamic panel data model of agricultural wages: Preferred estimates

Variable	Parameter estimate	t-Statistic	Parameter estimate	t-Statistic
Constant	0.00354	0.219	0.02057	1.798
Lagged nominal wage and food price index				
$\Delta w_{-1} - \Delta w_{-3}$	–0.57303	–6.470	–0.60647	–7.465
$\Delta w_{-3} - \Delta p^f_{-4}$	–0.20732	–2.087	–0.26587	–3.039
$\Delta(\Delta p^f_{-1})$	0.40902	3.814	0.33073	4.030
$\Delta p^f_{-2} - \Delta p^f_{-4}$	0.02047	0.224		
Δp^f_{-4}	0.23624	1.348		
Current and lagged yield per feddan				
ΔYLD	0.13412	0.947	0.04014	0.326
ΔYLD_{-2}	–0.19500	–1.206	–0.16812	–1.097
ΔYLD_{-4}	0.17153	1.124	0.14581	1.012
Current and lagged total cropped area				
$\Delta AREA_{-2}$	1.65917	3.360	1.63241	3.541
$\Delta AREA_{-4}$	–1.10099	–2.160	–1.20159	–2.503
Current and lagged total population				
ΔPOP	–1.67718	–3.432	–1.49717	–3.293
ΔPOP_{-1}	1.50276	3.178	1.17871	2.867
ΔPOP_{-2}	–0.55826	–1.833	–0.61169	–2.118
ΔPOP_{-4}	0.77201	2.528	0.89689	3.444
Current and lagged public-sector industrial output per person				
$\Delta YPUB_{-1}$	0.11208	2.168	0.08757	1.912
$\Delta YPUB_{-2}$	–0.07372	–1.559	–0.05290	–1.219
$\Delta YPUB_{-4}$	0.11767	2.985	0.09006	2.682
Current and lagged private-sector industrial output per person				
$\Delta YPVT$	–0.05220	–1.387	–0.03909	–1.154
$\Delta YPVT_{-4}$	0.11510	3.447	0.10244	3.376
Current and lagged real exchange rate				
ΔXR_{-2}	–0.04913	–1.560	–0.05056	–1.685
ΔXR_{-3}	–0.23362	–4.213	–0.20804	–4.314
ΔXR_{-4}	0.27661	5.574	0.26778	5.747
Current and lagged real remittances				
$\Delta REMIT$	–0.04569	–2.180	–0.04066	–2.162
$\Delta REMIT_{-3}$	0.16434	4.558	0.14425	4.720
$\Delta REMIT_{-4}$	–0.09049	–2.820	–0.08931	–2.928
Time trend				
(2 t – 1)	–0.00038	–2.729	–0.00046	–3.606
Sargan's test: df = 35; 37	22.175	p=0.955	27.521	p=0.872
Test for first-order serial correlation df = 18	–4.380	p=0.000	–4.342	p=0.000
Test for second-order serial correlation df = 18	0.206	p=0.837	–0.051	p=0.959
Test for contemporaneous homogeneity df = 6	341.06	p=0.000		
Test for short-run homogeneity df = 3	68.30	p=0.000		
Wald tests for zero parameter restrictions df = 2	2.84	p=0.242		

Note: These are parameter estimates for model (7). Number of governorates = 18. Number of observations = 695. The instrument set includes {1; instruments based on GMM conditions $E(w_{jt-5}\,\Delta v_{jt}) = 0$; p^f_{-5}; YLD_{-5}; $AREA_{-5}$; $POPT_{-5}$; $YPUB_{-5}$; $YPVT_{-5}$; XR_{-5}; $REMIT_{-5}$; $SEAS2$; $SEAS3$; t; t^2}. All test statistics are distributed as χ^2 with degrees of freedom as noted.

estimates in detail, note that the estimates satisfy the consistency requirement of no second-order serial correlation of the errors v_{jt}. Similarly, the Sargan test for overidentifying restrictions is also satisfied.

7. DISCUSSION OF RESULTS

The following additional observations can be made on the results presented in Table 3.

Current nominal wages in agriculture significantly depend on past nominal wages. Nearly three-quarters of the nominal wage during the past four *seasons* (16 months) is directly passed on to the current wage, of which about 60 percent is passed on during the first four months.[27] The elasticity of the current wage with respect to wages in the past four *seasons* (estimated as $\Sigma_i \alpha_i$) is 0.73; the elasticity with respect to wage in the last *season* is about 0.6.

This sluggishness in the wage response introduces a wedge between the short- and long-run wage effects. The long-run wage elasticities are thus considerably higher than the short-run elasticities. The ratio of the long- to short-run wage elasticities is given by $1/(1-\Sigma_i \alpha_i)$, which is estimated at about 3.76.

This has important implications for an assessment of the impact of changes in food prices. For instance, while the short-run elasticity of the nominal wage to food prices is

[27] Recall from earlier discussion that a season refers to a four-month period.

given by $\Sigma_i \beta_i$, the long-run elasticity is given by $\Sigma_i \beta_i / (1 - \Sigma_i \alpha_i)$. From our acceptance of the long-run homogeneity condition, we already know that the long-run elasticity of the nominal wage to food prices is unity, while the short-run elasticity is much lower, at about 0.27. Thus, over the short run (16-month period), only a little over one-quarter of the food price increases are absorbed in higher nominal wages. While they eventually catch up, real wages decline substantially in the short run.

Agricultural yields turn out to be an insignificant determinant of wages. This may be partly on account of the fact that our yield index had to be based on only ten crops. We do find the measures of total cropped area to be significant, though it is the *growth* in cropped area (rather than the level) that seems to matter; the restriction that the parameters on $\Delta AREA_{-2}$ and $\Delta AREA_{-4}$ add up to zero is statistically acceptable. The short-run elasticity of the agricultural wage with respect to growth in cropped area is around unity.

Population growth has a negative effect on agricultural wages. Again, it is the *growth* in population that has a negative effect on the level of wages. The restrictions that the parameters on ΔPOP and ΔPOP_{-1} as well as those on ΔPOP_{-2} and ΔPOP_{-4} add up to zero are readily satisfied. Thus, it is the growth in population during the last season as well as the last 8–16 months that have a negative impact on wages.

Our results are indicative of the importance of nonagricultural sources of labor demand. Increases in both public and private industrial output (per capita) have a positive impact on agricultural wages. The short-run impact of increases in output originating in the public sector is about twice that of those in the private sector.

We obtained mixed results on the effects of workers' remittances from abroad. While current remittances tended to depress agricultural wages, remittances a year ago had a strong positive effect. Yet, remittances from four seasons ago also had a significant negative effect on wages. The hypothesis that the combined short-run effect of remittances over the 16-month period (and hence the long-run effect) is zero could not be rejected. The exchange rate does have an independent effect on wages. But here again we have mixed results, and the hypothesis of zero long-run effect cannot be rejected.

We believe that the external labor demand effects are better picked up by the trend variables. Our results do indicate significant time trends. In particular, they suggest a quadratic time trend with an inverted U-shape. This is consistent with the findings for unconditional time trends discussed earlier. The turning point suggested by the parameter estimates in Table 3 is around 1983.5, which is also broadly consistent with the unconditional turning points presented earlier in Table 1. These trends are also consistent with the pattern of emigration and external labor demand as conditioned by international oil prices (Richards 1994).

Conditional on all the other determinants of agricultural wages, we found no evidence of seasonal effects. Seasonal effects are, of course, defined in terms of the four-month intervals introduced earlier, which may not adequately represent the agricultural peak and slack seasons across governorates. The observed lack of seasonality in our results should therefore be interpreted with caution.

8. A SIMULATION ON THE IMPACT OF FOOD PRICE CHANGES

What do the above results mean for the dynamics of the wage response to food price changes? We use the results from the estimated model to simulate the impact of a 10 percent increase in the food prices. From earlier discussion, we already know that eventually nominal wages would also increase by 10 percent. But it is also of interest to ascertain the dynamic time path of the nominal wage response, and determine in particular how quickly the eventual goal of full adjustment is realized. The wage-response time path can be discerned from the parameter estimates presented in Table 3. The results are graphed in Figure 1. The wage response is sluggish. By the end of year 1, the nominal wage rises by only about 1 percent, or about one-tenth of the food price increase is made good by end of the first year.[28] But wage adjustment is somewhat more rapid thereafter. By the end of year 2, nearly half of the food price increase is "recovered" in higher nominal wages, over two-thirds by year 3, about 80 percent by year 4, nearly 90 percent by the end of year 5, and about 95 percent by year 7.

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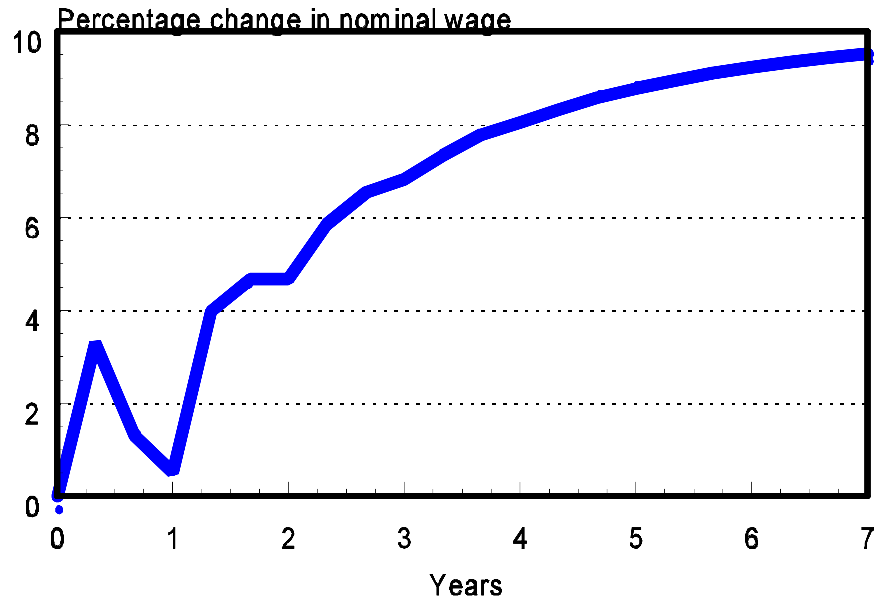
[28] The fluctuation in the wage response during the first year in Figure 1 is an artifact of the discrete four-month periodicity imposed by our data. With more frequently observed data, the stabilizing influence of the long-run effect would have kicked in sooner, producing a smoother wage response.

Figure 1 Simulated impact of a 10 percent increase in food prices on the nominal agricultural wage



The results are thus indicative of reasonably prolonged adverse real wage effects of higher food prices. These adverse short-term welfare effects should be of particular concern to policymakers contemplating reductions in the food subsidy program.

9. CAVEATS AND EXTENSIONS

(i) QUALITY OF WAGE DATA

As discussed in Section 3, our wage data do not extend up to 1993 for all 18 governorates. We opted to truncate the wage data for some governorates for some of the later years on account of very limited variation in nominal wages.[29] More generally in

[29] This data issue is discussed further in Appendix 1.

informal discussions, it has been suggested that the quality of the wage data has deteriorated in more recent years. It is thus useful to consider the general possibility of a time-trended measurement error in the wage data, which could be described as follows:

$$w_{jt}^{*} = w_{jt} + \phi_0 t + \phi_1 t^2 + u_{jt}^{*}, \qquad \textbf{(17)}$$

where w^{*}_{jt} is the actual wage, and $\phi_0 t + \phi_1 t^2 + u_{jt}^{*}$ is the measurement error (ME) with a quadratic time trend. The trend component of ME is readily subsumed under the quadratic trend variables in model (3). But the dynamic specification would still result in biased estimates even if it is uncorrelated with the error process. However, there are two reasons why such ME does not seem to be a serious concern for our results. First, the presence of such ME ought to be reflected in a violation of the Sargan test for instrument validity, but the Sargan test is satisfied by the estimates in Tables 2 and 3. Second, if such ME were indeed characteristic of our data, an implication would be that consistent estimation could no longer be based on the condition $E(w_{jt\ s}\, \Delta v_{jt}) = 0$ for a model that includes (s–1) lags in the dependent variable. Instead, we would need an additional lag in the GMM conditions, using $E(w_{jt\ s-1}\, \Delta v_{jt}) = 0$.[30] However, this condition for consistent estimation under ME, as in equation (16) above, is already satisfied by the estimates reported in Table 3. We also reestimated the unrestricted models in Table 2 and Appendix Table 5 using the condition $E(w_{jt-6}\, \Delta v_{jt}) = 0$, which made little difference to the results.

[30] Thus, for instance, for a model with w_{t-3} as a regressor, valid instruments could be based on w_{t-5}.

A related issue concerns the unresolved ambiguity on the length of the working day, i.e., the number of hours of work the daily wage data refer to. In particular, there is the concern that average hours of work per day have declined over the years (see discussion in Appendix 1). Statistically, this issue is analogous to the case discussed above. A systematic (negative) trend in hours of work that is not reflected in the measured daily wage is equivalent to a ME representation where the wage for a standard-length workday is measured with error which has a systematic time trend. Thus, the arguments in (i) apply to this case too.[31]

(ii) REGIONAL VARIATION IN FOOD PRICES

The rural food CPI data we have used are not spatially disaggregated; the price data are thus invariant across governorate. While we allow for governorate-specific effects, this still matters insofar as there are differential trends in food prices across governorates. Let these heterogeneous trends be represented as

$$p_{jt}^{f} - p_{t}^{f} = \tau_{j} + \tau_{0j}\, t + \tau_{1j}\, t^{2} + \xi_{jt}\,. \qquad \textbf{(19)}$$

To assess the significance of such heterogeneity in spatial price trends, we experimented with the possibility of differential linear and quadratic trends across governorates in the estimable versions of model (3). In particular, we distinguished four different types of

[31] A similar argument also applies to potential measurement error related to the possible decline in the in-kind component of wages.

governorates corresponding to the four different periods of estimation used in this study (see Table 1). Differential trends were allowed for each of the four different types. However, we were unable to reject the null of no differential in the trends (Wald(6) = 4.17; p = 0.65). Thus, spatial heterogeneity in food price trends does not appear to be a serious concern for our results. Similar comments also apply to the remittance variable, which, too, is available only at the national level.

(iii) INDIRECT FOOD PRICE EFFECTS THROUGH HIGHER LABOR DEMAND

The food price effects in our estimated models are conditional on the labor demand variables, notably agricultural yield and area. It could be argued that this underestimates the total food price effects insofar as these effects also operate through higher demand for agricultural labor. Table 4 presents estimates of our model suppressing the agricultural yield and area variables. The results are both qualitatively and quantitatively similar to those in Table 3. The contemporaneous and short-run homogeneity conditions are still rejected, while long-run homogeneity is still statistically acceptable. The short-run elasticity of nominal wages to food prices is 0.2, but it is not significantly different to the earlier estimate of 0.27. There are two potential reasons why the controlling for agricultural yield and cropped area seems to make little difference to the results. First, it is not uncommon to find that the price response of *aggregate* yield and area (as opposed to those for individual crops) is typically low. Second, the government's food subsidy policy drives a wedge between producer and consumer prices. For instance, a fixed retail

Table 4 Dynamic panel data model of agricultural wages: Estimates without yield and cropped area variables

Variable		Parameter estimate	t-statistic	Parameter estimate	t-statistic
Constant		0.00606	0.468	0.00854	0.935
Lagged nominal wage and food price index					
$\Delta w_{-1} - \Delta w_{-3}$		–0.56025	–7.097	–0.58513	–7.834
$\Delta w_{-3} - \Delta p^f_{-4}$		–0.16222	–1.873	–0.19780	–2.517
$\Delta(\Delta p^f_{-1})$		0.27502	3.020	0.28661	3.782
$\Delta p^f_{-2} - \Delta p^f_{-4}$		–0.08259	–1.036		
Δp^f_{-4}		–0.01006	–0.072		
Current and lagged total population					
ΔPOP		–1.28093	–3.092	–1.18108	–2.988
ΔPOP_{-1}		1.34795	3.196	1.22960	3.348
ΔPOP_{-2}		–0.33863	–1.322	–0.33802	–1.360
ΔPOP_{-4}		0.91778	3.663	0.85160	3.883
Current and lagged public-sector industrial output per person					
$\Delta YPUB_{-1}$		0.09670	2.106	0.07554	1.796
$\Delta YPUB_{-2}$		–0.02751	–0.664	–0.01290	–0.335
$\Delta YPUB_{-4}$		0.11394	3.325	0.09627	3.202
Current and lagged private-sector industrial output per person					
$\Delta YPVT$		0.01115	0.383	0.00821	0.295
$\Delta YPVT_{-4}$		0.09468	3.313	0.09521	3.571
Current and lagged real exchange rate					
ΔXR_{-2}		–0.01968	–0.719	–0.01867	–0.701
ΔXR_{-3}		–0.18837	–3.877	–0.19446	–4.407
ΔXR_{-4}		0.24023	5.520	0.24117	5.780
Current and lagged real remittances					
$\Delta REMIT$		–0.02456	–1.375	–0.02898	–1.747
$\Delta REMIT_{-3}$		0.12904	4.238	0.13265	4.900
$\Delta REMIT_{-4}$		–0.08311	–2.977	–0.08506	–3.132
Time trend					
(2 t – 1)		–0.00032	–2.830	–0.00035	–3.384
Sargan's test:	df = 40; 42	44.881	p=0.275	49.113	p=0.210
Test for first-order serial correlation	df = 18	–5.002	p=0.000	–5.090	p=0.000
Test for second-order serial correlation	df = 18	1.338	p=0.181	1.313	p=0.189
Test for contemporaneous homogeneity	df = 6	400.07	p=0.000		
Test for short-run homogeneity	df = 3	99.18	p=0.000		
Wald tests for zero parameter restrictions	df = 2	1.82	p=0.403		

Note: These are parameter estimates for model (7). Number of governorates = 18. Number of observations = 695. The instrument set includes {1; instruments based on GMM conditions $E(w_{jt-5}\,\Delta v_{jt}) = 0$; p^f_{-5}; YLD_{-5}; $AREA_{-5}$; $POPT_{-5}$; $YPUB_{-5}$; $YPVT_{-5}$; XR_{-5}; $REMIT_{-5}$; $SEAS2$; $SEAS3$; t; t^2}. All test statistics are distributed as χ^2 with degrees of freedom as noted.

price of 5 piasters per loaf of bread is maintained while excess demand at that price is met through imported wheat, while the producer price of wheat appears to fluctuate around the world price (Badiane and Kherallah 1998). For the subsidized food products, producer prices are thus relatively insulated from changes in consumer prices.

10. CONCLUSION

Induced, labor market effects of food price policy have been a longstanding concern in development policy debates, and are arguably a matter of substantial relevance to the rationalization of the food subsidy program in Egypt. It is important to consider these induced or second-round effects, as these can sometimes be substantial enough to tip the balance for or against proposed policy changes. At a minimum, they serve to alert policymakers to some less apparent implications of the proposed changes. In this paper, we have attempted to shed light on the existence and importance of second-round wage effects of food price changes in the context of agricultural labor markets in Egypt over the period 1976–1993.

There are several distinguishing features of our study. First, it covers a longer and more recent period than other studies on agricultural wages in Egypt. Second, the study uses governorates as the cross-sectional units of observation, thereby allowing us to take into account observed and unobserved governorate-specific determinants of agricultural wages. This is important because to successfully isolate the response of wages to food prices, one needs to control for other influences on wages. We also exploit the panel aspect of our data to address a number of data limitations related to the potential presence of certain kinds of measurement error in the wage data, and the lack of cross-sectional variation in some of the wage determinants. Finally, the study also allows for past wages to affect current wages (a dynamic wage response) that is typical of the wage adjustment process in most labor markets. This makes for not only a more realistic modeling of the wage adjustment process, but it also allows us to trace out the immediate and longer-term welfare effects of food price policies such as changes in the food subsidy system.

Our analysis points to the volatility of agricultural wages over the period 1976–1993. The typical real agricultural wage in Egypt increased rapidly up to about 1985 and declined rapidly thereafter. The pattern is remarkably similar across the governorates. However, hidden behind these simple time trends is a complex dynamic pattern of wage adjustment to changes in food prices and other determinants.

We find there is considerable sluggishness in the wage response. Nearly three-quarters of the nominal wage during the past 16 months is directly passed on to the current wage, of which about 60 percent is passed on during the first four months. An important implication of this sluggish adjustment is that long-run wage responses are significantly larger than the short-run responses.

Our analysis strongly rejects the notion of instantaneous or quick wage adjustment. In particular, we strongly reject the hypothesis that increases in food prices are *fully* passed on to nominal agricultural wages during the same four-month period (which is the temporal unit of analysis used in our study) or even during the same year. The short-run elasticity over a 16-month period is estimated to be about 0.27. However, nominal wages do fully catch up with higher food prices in the long-run. But the process takes time, up to five years for a 90 percent adjustment (and up to seven years for a 95 percent adjustment). In the interim, there are significant adverse welfare effects associated with real wage declines. Another way to interpret our results is that the food subsidy operates like a wage subsidy to employers only in the long run. Over the short run, it is more like a subsidy to the workers and cuts in that subsidy can be expected to hurt the workers.

APPENDIX 1

NOTES ON THE DATA

WAGES

Data on agricultural wages have been collected by the Ministry of Agriculture for many years. According to the Ministry of Agriculture sources and others familiar with these data, data collection proceeds as follows. The wage data are collected at the village level by the agricultural co-op representatives who ask around the village to establish what the going wage is. The data are reported for various tasks, for both adult male and boy's labor.[32] Data are collected for various crops and operations at the village-level and then reported to a *markaz* (district) office. The *markaz* officer averages the various village level wages and reports them to the governorate office. These data are then reported to the central office in Cairo, although there does not appear to be a standardized form or way of reporting the data. For instance, one governorate statistics office might report a wage for each agricultural operation currently being performed. The statistics officer in Cairo then takes a straight average of those numbers. Another office might report a range of wages for the region, such as reporting that wages currently range from 3 to 5 LE. The

[32] There does not seem to be an official age cutoff for boys. Although data are reported as men's and boys' wages, women and girls also participate in agricultural labor. We were often told that boys and women earn the same wage, which is generally about half of the men's wage. However, the evidence in Commander (1987) suggests that this statement is not necessarily true. Children often make far less than half of what men do and women's wages are generally higher than children's. Commander also finds that children's and women's real wages did not rise as quickly and as much as men's wages. Women's wages in the 1960s were about 60 percent of men's wages, with children's wages being about half. In 1984, these rates had fallen slightly, with women earning about 58 percent of what men did and children earning 41 percent of men's wages.

Cairo office then takes an average of that range. Other governorate offices might only report one number, having done the averaging themselves. In other words, aggregation may occur at various points and in various ways during the collection and reporting process. Commander (1987, 43) believes that the level of aggregation and method of collection may "disguise substantial variation in the wage trend both within and across regions." Fitch, Ali, and Mostafa (1980) are also somewhat critical of the wage data, pointing out that failure to weigh averages correctly may lead to errors. Still, they claim that when compared with another sample of data collected in the 1976–77 Farm Management Survey, the data show similar patterns. According to them, because of the way the data are collected, there may be a small lag in terms of upward adjustments and some seasonal variation may not be reflected in the data. Commander (1987) also suggests that the official wage data understate seasonal wage variations. All this points to potential measurement error in the wage data, some of which is, however, attenuated by our procedure of averaging the wage data over four-month periods.

Although there is some evidence on the variation of wages by agricultural operation (see, for instance, Richards, Martin, and Nagaar 1983), on the primary sample data sheets that we examined, there was little or no variation in wages across agricultural operations within a governorate, which suggests that the use of an average wage across different types of agricultural operations in our analysis is not likely to be a serious problem.

As mentioned in the text, for some governorates there are missing observations in the wage data. This could be due to the absence of any agricultural activity during that period or due to nonreporting. We have dealt with this problem by averaging the wage data over four-month periods, which greatly reduces the number of missing observations. We end up with three wage observations per year, which correspond to the three "seasons" of December-March (season 1), April-July (season 2), and August-November (season 3). Even after the averaging, we were left with a few missing data points. These include the following: Damietta: 1977, season 3; 1985, season 3; 1987, seasons 2, 3; Fayoum: 1978, season 3.[33]

The reliability of the wage data for many governorates appears to have declined in the recent years, especially since the mid-1980s.[34] In examining the nominal wage data, we found that a number of governorates began reporting a constant men's wage of around LE 5 in 1985, with minimal or no variation since then. These included Qena, Kafr El-Sheikh, Kalyoubia, Beni-Suef, Suhag, and Aswan. In addition, Sharkia and Ismailia showed limited variation in nominal wages during 1985–90 and no variation after 1990. Because of these problems with the later data, we decided to limit the period of analysis for these eight governorates: 1976–1985 for the former group of six, and 1976–1990 for the latter two. For Alexandria, our available wage data only start in 1981, and hence the period of analysis is limited to 1981–1993. Table 1 lists the different governorates and the years of data that were included for them.

[33] These observations were excluded from analysis by including a dummy variable for each of the five cases. In later runs of the model, the estimated parameters for the five dummy variables were used to construct estimates of the missing wage data.

[34] See Richards (1994), who also remarks on this problem.

In some governorates, much of the seasonal variation in wages also seemed to decline since the 1980s, but this may be due more to data reporting problems (mentioned above) than to reduced variability in the observed wage.[35] This decline in seasonality appears to be more pronounced in some governorates than in others. Again, the problem is mitigated by the temporal aggregation of wages over the four-month periods.

Another question concerning the data was whether the daily wage reflects the entire payment received by the worker, including any in-kind receipts. We were unable to resolve this issue completely. Based on discussions in the field, we got various answers to the question: "Do workers get in-kind payment?" Some of the Sharkia governorate workers said that the farmer provides the worker with a meal while he or she is working. In Cairo, we were told that no in-kind payments are made.

Fieldworkers in Sharkia also explained that there is a growing tendency towards piece rates, i.e., negotiating a set amount for a given task rather than a daily wage. According to the people we interviewed, the laborers preferred this method of payment. Although this method of payment may be growing in popularity, the Ministry of Agriculture does not explicitly collect these data and therefore they are not included in our analysis. Again, this is of special concern only insofar as these payments adjust differently than wages.

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[35] Commander (1987) offers a somewhat different explanation for the decline in seasonality. Although he does not rule out the conclusion that reporting problems are responsible for some wage differentials, he also suggests that workers' expectations have changed as a result of increased alternative opportunities to agricultural wage labor and that has made wages more sticky.

There is also some unresolved ambiguity concerning the length of the working day. The consensus in Sharkia was that a full working day was about 6 hours. Workers start around 7 AM and finish at 1 PM. Those interviewed stated that 10 to 15 years ago, the workday was longer, with laborers starting around 7 AM, working until 1 PM, then taking a lunch break and returning to the field for another hour or two of work. From these discussions, we gleaned that in Sharkia, hours of work dropped from 8 to 6 hours around 1985. Migration opportunities and the development of nearby industrial centers were given as reasons why agricultural workers refused to work longer hours. It may also be true that in some governorates, the workday is shorter than in others.[36]

One of the best documented sources of reduction in hours of work is Commander's 1987 study, which suggests hours of work actually declined earlier than 1985. Based on his study of four delta villages, he states that while hours of work in the early 1960s averaged 7–8, by 1984, this number had declined to 5.3 hours. He also provides detailed information on how hours of work vary for different agricultural tasks. In general, there seems to be a consensus about a drop in the daily hours of work (with a decrease of about 2 hours often reported), but less so on the period over which that change occurred.

Another problem is that hours may fluctuate seasonally, as Fitch, Ali, and Mostafa (1980) point out. Similarly, Richards, Martin, and Nagaar (1983) state that peak labor demand in agriculture, at least in Sharkia, where they carry out their study, is May/June

[36] Nabil Habashi of the Agricultural Economics Research Institute, MALR, Cairo, suggested that in Menia, the workday is closer to 5 rather than 6 hours.

and October. By contrast, seasonal unemployment peaks during the months of January, February, August, and December. They argue that rural wages are somewhat sticky and therefore wages do not adjust completely to seasonal variations in labor demand.

PRICES

Our price data are based on the Consumer Price Indices (CPI) published by the CAPMAS. These data, which are published on a monthly basis for urban areas, and once every two months for rural regions, have been widely used as deflators in applied work on Egypt. Currently, separate rural CPI estimates are calculated for Upper and Lower Egypt. However, this only began in the late-1980s, when the base year was also changed from 1966/67 to 1986/87. For our study, which covers the period 1976 to 1993, we were thus unable to use the regionally disaggregated CPI data. Besides the general indices for all commodities, the consumer price indices are also available separately for the category food, beverages, and tobacco, as also some nonfood categories. The food price data are generally viewed as accurate, since the prices are obtained from the markets on a regular basis.[37]

[37] Personal communication from Ragui Assaad. Within the nonfood categories, CPI is reported for clothing and footwear, rent, power and fuel, furniture and equipment, medical care, transport and communication, recreation and education and miscellaneous. Figures for rent and transport may be less reliable, since an assumption is made that housing and transportation are primarily publicly provided at subsidized rates, which tends to underestimate inflation in those two items. Richards (1991) and Commander (1987) mention that critics of the CPI argue it underestimates inflation. Commander (1987), for instance, states that inflation during the 1970s was underestimated, suggesting that growth in real wages may be overstated. Most analysts have, nevertheless, still used the CPI data, as these still provide the best available deflators for extended time periods.

For our analysis, we used the food CPI data for rural Egypt, although we found the food and nonfood components to be virtually collinear.

The CPI data are broken down into two time periods. The period January 1974 to June 1988 has a base year of 1966/67. The data for the second series were first published in July 1988 with 1986/87 as the base year. The change in the base to 1986/87 was also associated with an increase in the basket of commodities from 245 to 402, with a large increase in modern technological products, such as televisions. We spliced the indices for the two periods using conversion factors implicit in the published data, (CAPMAS 1996) using overlapping observations for the new and the old base years. We then renormalized the entire CPI series to May 1996 for ease in interpretation. As our temporal unit of observation is a four-month period (see discussion of the wage data above), we have taken a simple average of the two successive rural CPIs (reported every two months) to derive values corresponding to the temporal unit of our analysis.

AGRICULTURAL PRODUCTION, YIELD, AND TOTAL CROPPED AREA

We used detailed governorate-level data on agricultural production, area, and yield for different crops collected by the Ministry of Agriculture.[38] We first constructed a measure of agricultural output at the governorate level using data for the following ten crops: broadbeans, soybeans, corn (Summer), tomatoes (Nile), cotton, potatoes

[38] We are grateful to Mohammed Omran of USAID, Cairo, for providing us with a copy of these data. These data have also been used by Rady, Omran, and Sands (1996), who also provide more details on the data.

(Summer), wheat, potatoes (Nile), rice (Summer), and sesame. We had to limit our production measure to these ten crops because these were the only crops for which a governorate-level time-series was available for our period of estimation. Needless to say, these ten crops are not exhaustive of agricultural production; in 1993, they accounted for 55 percent of total cultivated area.[39] In particular, we are concerned about the omission of *berseem* (clover). *Berseem* is an increasingly important feed crop and by 1993 accounted for 22 percent of total cultivated area. Since data for berseem are unavailable at the governorate level for the early years in our sample, we were unable to include it in our measure of agricultural output.

To mitigate the effects of this omission, we also included the total cropped area in the governorate (including *berseem* area) as an additional explanatory variable representing demand for agricultural labor. Annual governorate-level data on the total cropped area were obtained from CAPMAS.

Our measure of agricultural production was constructed as follows. We first constructed average prices for each of the ten crops using national crop price data for the years 1991/92, 1992/93, and 1993/94. These average crop prices were then used as constant weights for aggregating the quantities of crop output in each governorate for each year. Our measure of agricultural output can thus be interpreted as the governorate- and year-specific value of *real* agricultural output (at average 1991–94 national prices).

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[39] Calculated using 1993 cultivated area figures from the CAPMAS *Statistical Yearbook of 1994* (1995).

We divide this real output by the total cropped area for the ten crops in the governorate for any given year to obtain our measure of agricultural yield.

POPULATION

Ideally, we would have liked to use governorate-level data on the labor force. However, no continuous series are available on labor force statistics. In addition, the existing data are considered somewhat unreliable or noncomparable, as the method of collection has changed over time. Employment data sources include the Labor Force Sample Survey (LFSS) and the Employment, Wages and Hours of Work Survey. The LFSS data are considered a better source, but they still have problems. For one thing, the series is incomplete with a number of missing years. For instance, in 1976 and 1986, no LFSS survey was conducted because the Population Census was taking place. In 1985, too, no survey was conducted. The 1987 and 1989 survey results were never published and hence data for these years remain unavailable. In 1988, a different methodology was used, and therefore the data are not comparable to other years.[40]

As a proxy for labor force, we instead used population estimates from CAPMAS sources. These estimates are reported annually in the LFSS Survey for each governorate, separately for male and female, and rural and urban populations. Printed tabulations of the LFSS data from CAPMAS were available for 1976–1984, 1986, and 1990–1994. As for

[40] Discussions with Ragui Assaad helped clarify these points concerning the state of the employment data for Egypt.

the missing years, we also obtained additional (unpublished) estimates of population for the years 1986–1995 from CAPMAS sources. But these turned out to be slightly different than the population estimates from the LFSS tabulations. For the years 1987–1989, we therefore adjusted the original figures (from the unpublished source) using the ratio of the population figures from the two sources for one overlapping year, 1986. For 1985, however, because no data were available, our population estimates were based on a log-linear interpolation between 1984 and 1986.

INDUSTRIAL OUTPUT

Data on governorate-level industrial output were obtained from CAPMAS. These data are available on an annual basis, and are available separately for industrial output originating in the public and private sectors. These are useful as measures of alternative, nonagricultural sources of labor demand, especially those originating in the formal sectors to which these data are likely to be confined.

We also looked into the possibility of using nonagricultural wage data for our analysis. One possible source for these data is the survey on Employment, Wages and Hours of Work conducted by CAPMAS, which includes formal-sector wage data. However, data from this survey are available only for a limited number of years (1976–1978, 1982, and 1983–1984). Because of these data gaps, we were unable construct a usable measure of nonagricultural wages.

EXCHANGE RATE AND WORKERS' REMITTANCES

Although international migration is widely regarded as an important factor in the operation of rural labor markets in Egypt, data on the number of migrants are difficult to obtain. No continuous series exist at the national level, let alone at the governorate level. While governorate population estimates (described above) may capture out-migration, they do not distinguish between local and international movements. One possibility is to use workers' remittances from abroad. Although no governorate-level data are available on remittances, the *Balance of Payments Statistics Yearbook*, published by the International Monetary Fund, provides national, annual figures on workers' remittances. The remittance series was reported in SDRs for the period 1976–1987 and then in U.S. dollars for the later period (1988–1993). We converted the SDR figures to U.S. dollars using the exchange rate data published in the *Balance of Payments Statistics Yearbook* as well as the *International Financial Statistics Yearbook*.

As noted in Commander and Hadhoud (1986, 171), it is likely that remittance figures are underestimated, as many workers may not be remitting through official channels. This is particularly true for the earlier period, when exchange rates were fixed. In the late 1980s, Egypt carried out a number of devaluations of its currency in an attempt to correct the overvaluation as part of its structural adjustment policies, which have continued into the 1990s (Handoussa 1991). We therefore also allowed for the (real) exchange rate as an additional argument in our specification.

APPENDIX 2
REAL AGRICULTURAL WAGES AND THE RURAL CONSUMER PRICE INDEX

Figure 2 Real wages and rural food price index: Behera

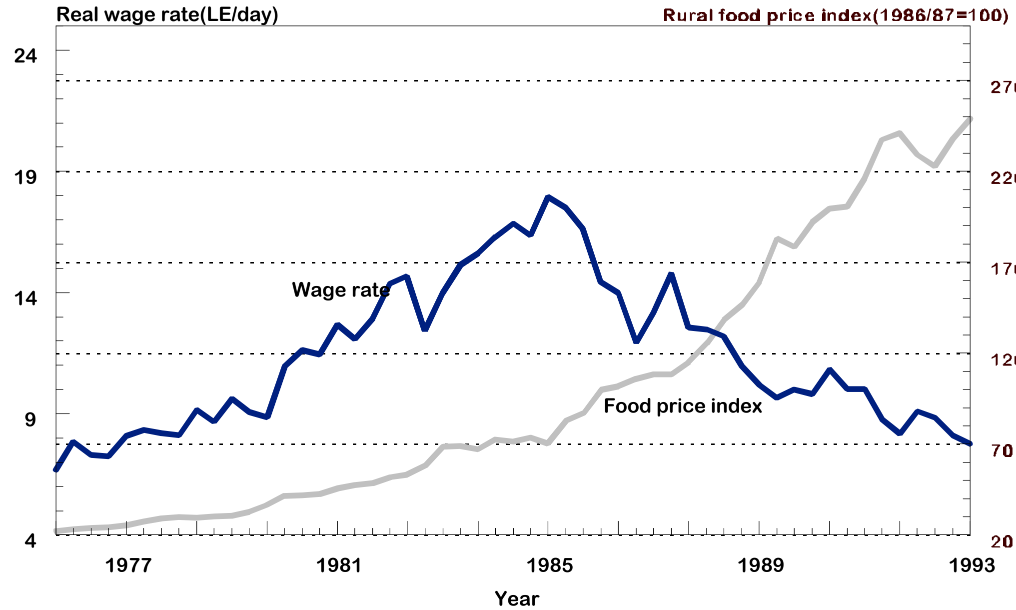


Figure 3 Real wages and rural food price index: Gharbia

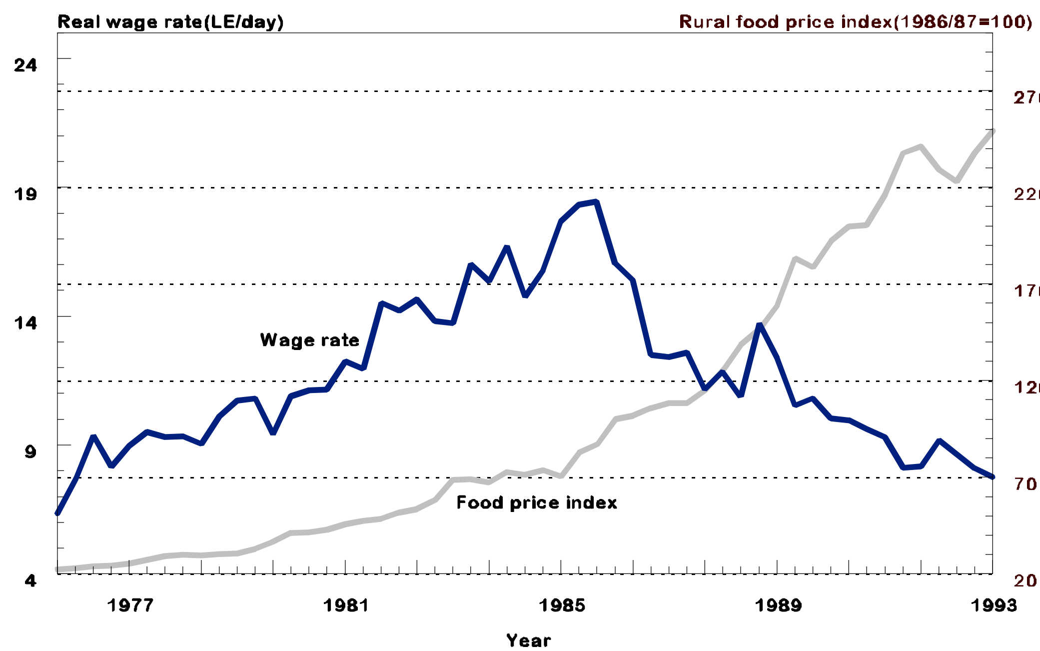


Figure 4 Real wages and rural food price index: Dakahlia

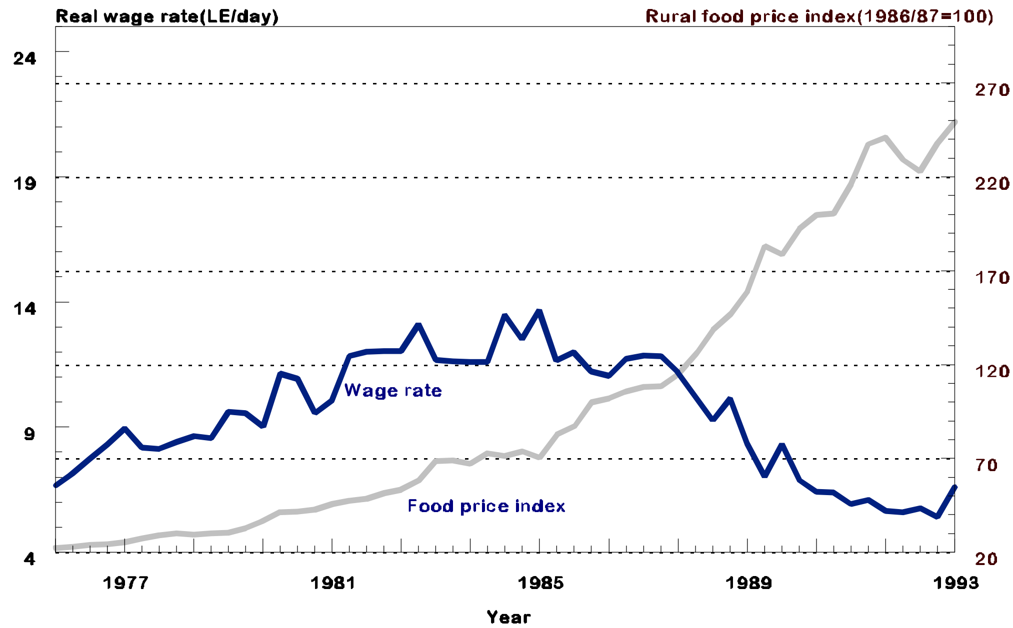


Figure 5 Real wages and rural food price index: Damiett

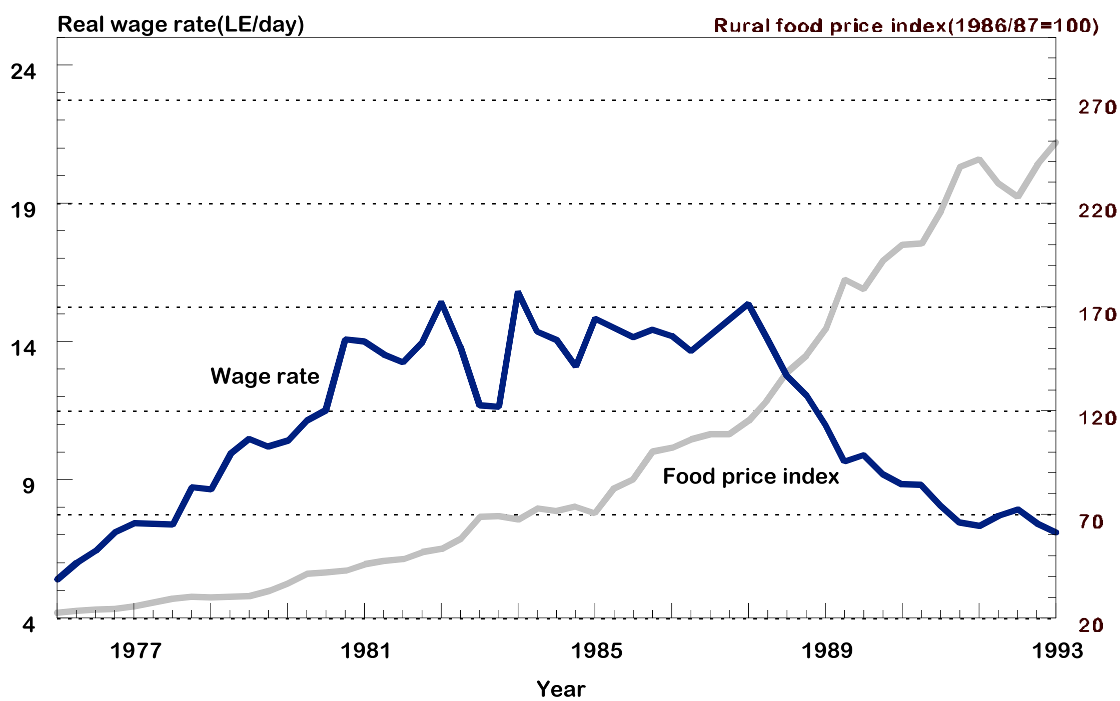


Figure 6 Real wages and rural food price index: Menoufia

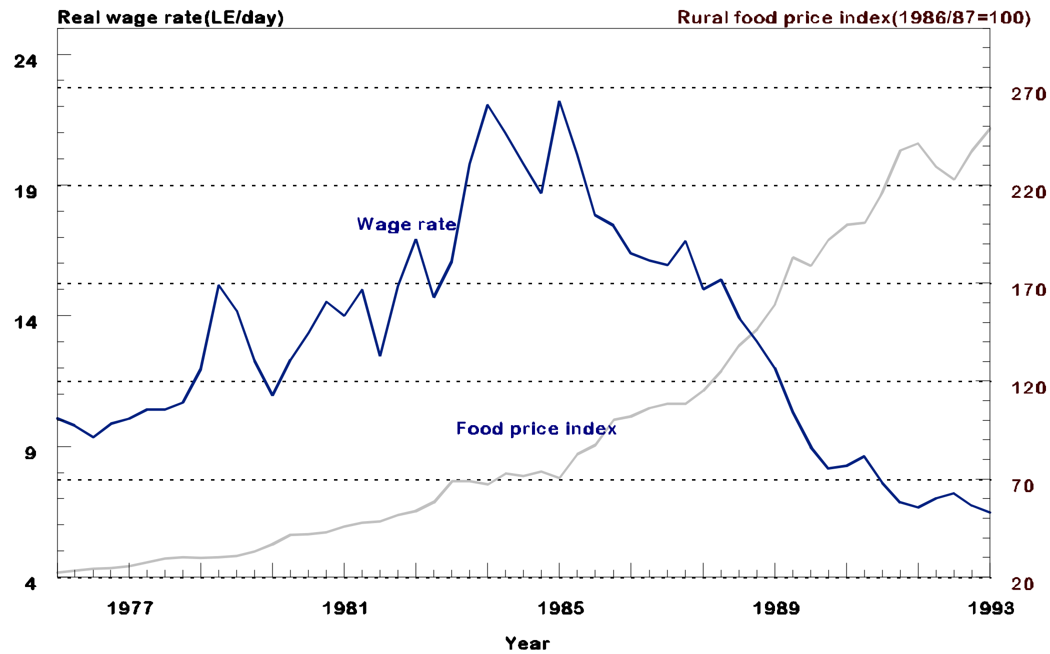


Figure 7 Real wages and rural food price index: Giza

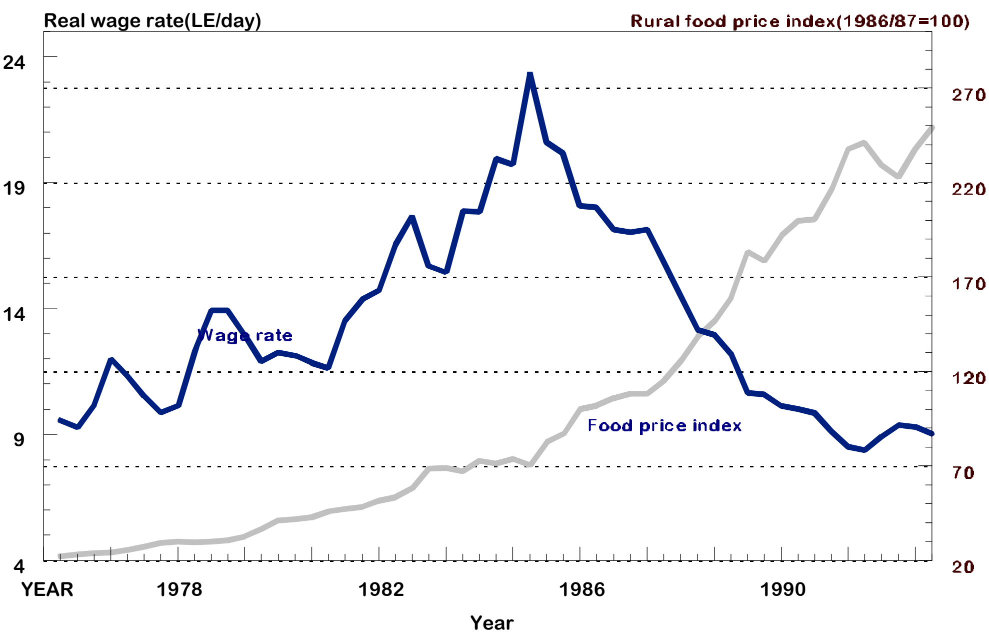


Figure 8 Real wages and rural food price index: Fayoum

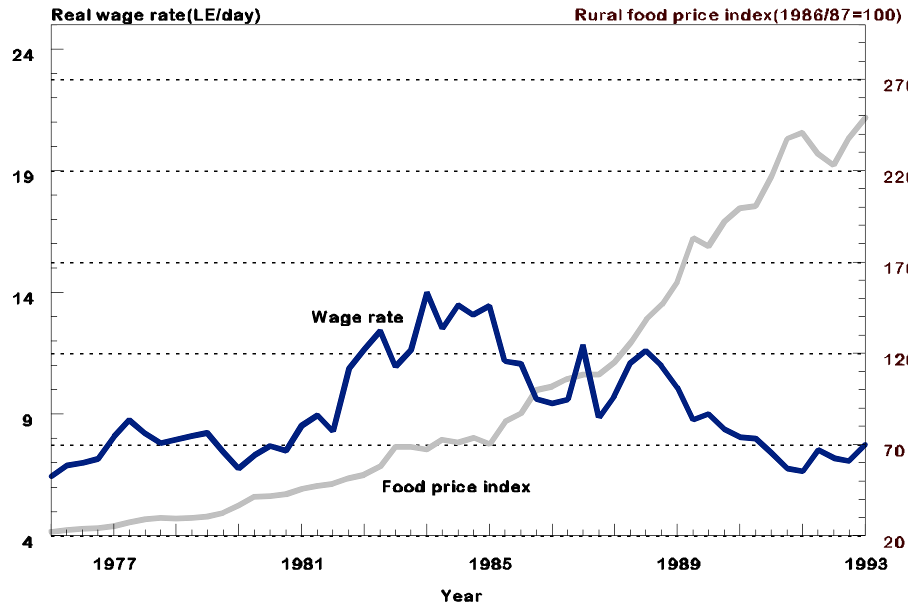


Figure 9 Real wages and rural food price index: Menia

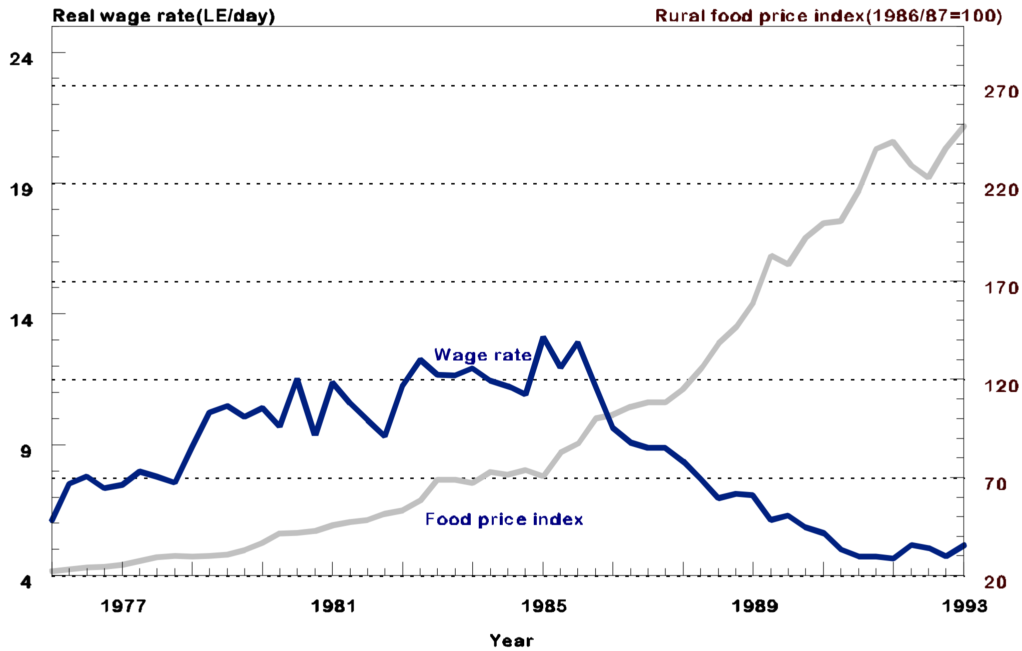
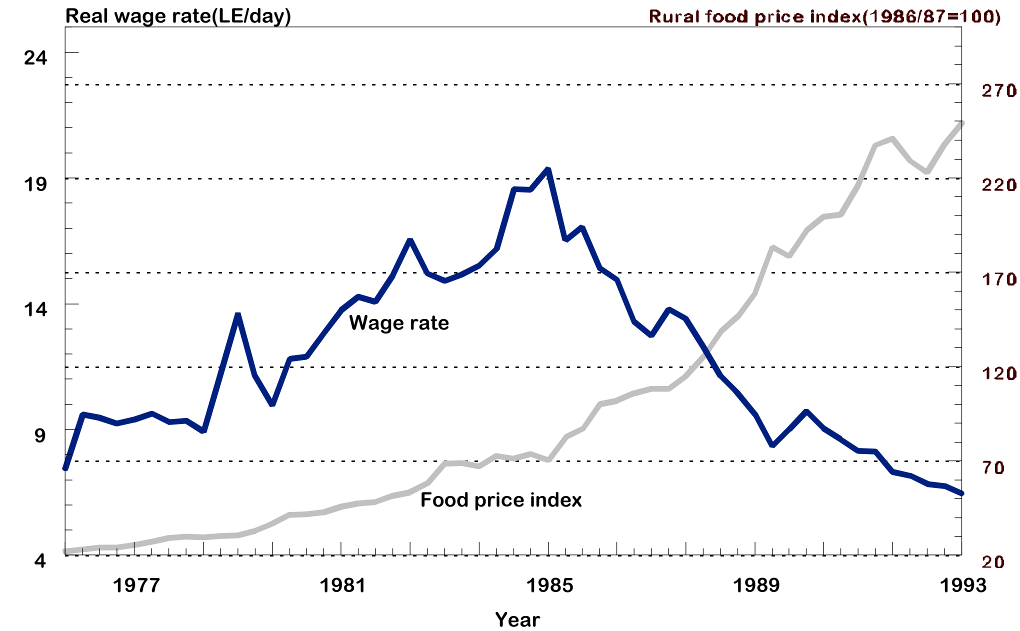


Figure 10 Real wages and rural food price index: Asyout



APPENDIX 3:
INITIAL ESTIMATES OF THE AGRICULTURAL WAGE MODEL

In Table 5, we present the initial dynamic panel data estimates of the agricultural wage model (3). These correspond to the completely unrestricted AD(4,4) specification of the model.

Table 5 Dynamic panel data model of nominal agricultural wages: Initial estimates

Variable	Parameter estimate	t-statistic
Constant	–0.03362	–0.796
Lagged nominal wage		
Δw_{-1}	0.51138	1.821
Δw_{-2}	–0.01319	–0.039
Δw_{-3}	0.37356	0.903
Δw_{-4}	–0.02989	–0.107
Current and lagged food price index		
Δp^f	0.21747	1.121
Δp^f_{-1}	0.36459	1.193
Δp^f_{-2}	–0.31802	–1.372
Δp^f_{-3}	0.07408	0.294
Δp^f_{-4}	0.69621	1.700
Current and lagged yield per feddan		
ΔYLD	0.71590	1.717
ΔYLD_{-1}	–0.08209	–0.178
ΔYLD_{-2}	–0.25360	–0.866
ΔYLD_{-3}	0.11846	0.315
ΔYLD_{-4}	0.21830	0.709
Current and lagged total cropped area		
$\Delta AREA$	–0.38725	–0.359
$\Delta AREA_{-1}$	0.23219	0.198
$\Delta AREA_{-2}$	1.79714	1.309
$\Delta AREA_{-3}$	1.85028	0.993
$\Delta AREA_{-4}$	–2.21824	–1.269
Current and lagged total population		
ΔPOP	–2.22746	–1.696
ΔPOP_{-1}	2.24792	1.737
ΔPOP_{-2}	–0.73830	–1.051
ΔPOP_{-3}	–0.23513	–0.282
ΔPOP_{-4}	0.89118	1.122
		(continued)
Current and lagged public-sector industrial output per person		
$\Delta YPUB$	0.11862	0.811
$\Delta YPUB_{-1}$	0.18096	1.208
$\Delta YPUB_{-2}$	–0.18760	–1.369
$\Delta YPUB_{-3}$	0.01542	0.103
$\Delta YPUB_{-4}$	0.18385	1.490
Current and lagged private-sector industrial output per person		

Variable	Parameter estimate	t-statistic
ΔYPVT	–0.14270	–1.146
$\Delta YPVT_{-1}$	–0.03013	–0.287
$\Delta YPVT_{-2}$	0.06953	0.593
$\Delta YPVT_{-3}$	0.02849	0.245
$\Delta YPVT_{-4}$	0.13131	0.988
Current and lagged real exchange rate		
ΔXR	–0.04294	–0.337
ΔXR_{-1}	–0.04339	–0.425
ΔXR_{-2}	–0.05950	–0.579
ΔXR_{-3}	–0.30226	–2.518
ΔXR_{-4}	0.31511	3.423
Current and lagged real remittances		
ΔREMIT	–0.04308	–0.679
$\Delta REMIT_{-1}$	0.01076	0.180
$\Delta REMIT_{-2}$	0.04716	0.697
$\Delta REMIT_{-3}$	0.18835	2.176
$\Delta REMIT_{-4}$	–0.11271	–1.781
Seasonal dummy variables		
ΔSEAS2	0.00080	0.012
ΔSEAS3	–0.00128	–0.016
Time trend		
(2 t – 1)	–0.00027	–0.854
Sargan's test: df = 14	3.781	p=0.997
Test for first order serial correlation df = 18	–1.111	p=0.267
Test for second order serial correlation df = 18	–0.564	p=0.573

Note: These are unrestricted parameter estimates for model (4). Number of governorates = 18. Number of observations = 695. The instrument set includes {1; instruments based on GMM conditions $E(w_{jt-5}\,\Delta v_{jt}) = 0$; p^f_{-5}; YLD_{-5}; $AREA_{-5}$; $POPT_{-5}$; $YPUB_{-5}$; $YPVT_{-5}$; XR_{-5}; $REMIT_{-5}$; $SEAS2$; $SEAS3$; t; t^2}.

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