The Dynamics of Food Price Convergence in Ethiopia

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Abstract

This paper examines the dynamics of relative price convergence of nine agricultural commodities among regions in Ethiopia using a panel dataset of 18-year monthly prices collected by CSA in two periods (1996-2004 and 2005-2013). Panel unit root tests, fixe-effects, and half-life method were employed to estimate the rate and speed of relative price convergence of commodities. The findings markedly indicate low rate and speed of relative price convergence and considerably persistent relative price shocks unadjusted among regions, suggesting the need to design proactive market policy intervention in improving convergence of commodity prices in Ethiopia.

Keywords: Food price, commodity, price convergence, fixed-effects, Ethiopia.
JEL codes: Q11, Q13.

1. Introduction

The burgeoning literature on commodity price adjustment suggest the importance of spatial, temporal, vertical and intercommunity price integration for the presence of market efficiency. A priori economic theory suggests that spatially efficient markets increase supply and decrease price of goods and flow of goods and services in between food deficit and surplus areas, generating positive net welfare effects. Spatial food insecurity can be

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even be prevented before triggering famine if spatial arbitrage is efficient in distributing produces from surplus areas to deficit areas (Webb *et al.,* 1992; Tschirley, 1995).

Since grain trade liberalization in 1990, a number of short- and long-run policy measures including establishment of commodity exchange, banning of export of food stables like cereals following food price inflation, direct price setting (price ceilings) on some basic food items (oil, wheat flour, sugar), and privatization of many parastatals were implemented to improve domestic commodity market performance in Ethiopia..The most important question in this regard is whether or not these market reform measures and policies have resulted in domestic market performance in Ethiopia. One of the common indicators of the success of domestic market performance is price convergence. Price convergence shows the degree to which prices for goods in spatially differentiated markets, regions or countries have moved together, or converged towards unity (Susanto *et al.* 2008; Bukenya and Labys, 2005; Cecchetti, *et al.* 2002; Parsely and Wei, 1996). Price divergence, as opposed to this, shows the degree to which prices have moved apart as an indicator of poor internal market performance mainly attributable to inadequacy of market information, institutions and infrastructure. It is expected that increases in marketing infrastructure and globalization are broadly suggestive of relative commodity prices in spatially differentiated markets adjusting towards unity over time.

The primary focus of most previous studies in Ethiopia was, however, to measure the spatial, vertical or intercommunity price transmission among markets (se Webb *et al.*, 1992; Dercon,1995; Asfaw and Jayne, 1997, 1998;Gebremeskel *et al.*, 1998; Eleni,2001; Kindie *et al.*,2006; Kindie, 2007; Degye *et al.*,2009). They also have various limitations in terms of market coverage, methodology, and many empirical irregularities. The assumption of spatial price adjustment that commodities are homogenous is not testable and would frequently be unsatisfied. Commodity markets are different in terms of seasonality, geographical location, and other factors, which are difficult to be accounted for by such price adjustment models.

Because of these problems, there is still a wider gap of adequate and relevant empirical evidence on the Ethiopian commodity marketing system.

Multivariate price series models frequently employed for the purpose of spatial, vertical and inter-community price adjustment in Ethiopia have some common limitations and difficulties. The first limitations is that researchers are forced to focus on limited number of markets and commodities due to the fact that these methods are data-intensive, leading to model misspecification emanating from unavoidable omission of many important markets and commodities. Second, the existence of unobserved market-specific effects is ignored, and the assumption that markets are static is admitted. But the market-specific unobserved heterogeneities can contribute more towards the dynamics of price adjustment. The other most important limitation is that they are point estimates, unable to capture the dynamic pattern of price adjustment.

This paper employs rigorous dynamic panel data models which can account for non-time dependence, transportation costs, and unobserved quality differences of commodities. The presence of market or region fixed effects in the estimation also suggests the relative version of the law of one price (LOP), with an advantages over the absolute LOP, which assumes that transaction costs vary proportionately over time. The attempt to use wholesale and retail prices at supply and destination markers in multivariate forecasting models cannot capture these unobserved cross-section-dependent heterogeneities. Accordingly, this study was conducted to generate new and reliable empirical evidence on the dynamics of price convergence of major commodities traded and consumed in Ethiopia. Using an 18-year monthly price series of 9 commodities in 7 regions (including the benchmark region), the dynamic process of food price convergences was investigated over three periods (1996-2004, 2005-2013, and 1994-2013). The remaining part of the paper covers theoretical and empirical framework in Part 2, dataset and analytical methods in Part 3, and presentation and discussion of empirical findings in Part 4. Finally some concluding remarks are presented in Part 5.

2. Theoretical and Empirical Framework

2.1. Theoretical framework

Theoretical and empirical literature proposes many alterative models of univariate and multivariate time series forecasting. If two or more non-stationary time series follow a common longrun path, a test for cointegration would lead to a stationary linear combination of the time series (Engle and Granger, 1987; Johansen, 1988). Co-integration analysis is useful because estimated coefficients from cointegrated regressions will converge at a faster rate than normal, they are super consistent. The two widely employed models of multivariate time series are the vector autoregressive (VAR) and vector error correction (VEC) (Johansen, 1995). The VEC specification restricts the longrun behavior of endogenous price variables to converge to their longrun equilibrium relationships and allow the shortrun dynamics. Johansen proposes two different likelihood ratio tests of the significance of canonical correlations, the trace test and maximum eigenvalue test (Johansen and Juselius, 1990; Johansen, 1995). The trace test tests the null hypothesis of r cointegrating vectors against the alternative hypothesis of k cointegrating vectors. The maximum eigenvalue test, on the other hand, tests the null hypothesis of r cointegrating vectors against the alternative hypothesis of $r+1$ cointegrating vectors. Since the critical values used for the maximum eigenvalue and trace test statistics are based on a pure unit-root assumption, they will no longer be correct when the variables in the system are near-unit-root processes. Thus, the real question is how sensitive Johansen's procedures are to deviations from the pure-unit root assumption.

The very limitation of VAR and VEC model is that they both assume linear price adjustment, which might not be always true. Unlike models with linear price adjustments, the threshold vector error correction (TVEC) and Markov-switching vector error correction (MSVEC) models are appropriate candidates for estimation of short-run and long-run adjustment parameters under the presence of nonlinear price adjustments (e.g. Meyer, 2008; Tadesse *et al.* 2008; Reziti *et al.* 2008; Saghaian, 2008). A threshold introduces nonlinearities into the functional relationship and "specifies the operation modes of the system". The threshold principle is the local approximation over the states, or the introduction of regimes via thresholds. Such regime-dependent parameter stability of some time series is usually referred to as threshold behavior (Tong, 1990). The TVEC model and the MSVEC models are the two regime-dependent econometric models for price transmission analysis. The assumptions regarding the nature of their regime-switching mechanisms are fundamentally different so that each model is suitable for a certain type of nonlinear price transmission (Ihle and Cramon-Taubadel, 2008).

However, all the above multivariate price series models have at least three common limitations and difficulties. First, they are data intensive which force researchers to focus on limited number of markets and commodities rather than including all important markets and commodities in their analyses. Second, they ignore the existence of unobserved market-specific effects. The market-specific unobserved heterogeneities can contribute more towards the dynamics of price adjustment. Finally, they are point estimators, unable to capture the dynamics of price adjustment process.

As evidenced by many scholars including Banerjee (1999), Maddala and Wu (1999), Hadri (2000), Levin *et al.* (2002), Im, *et al.* (2003), Pindyck (2004), and Bai and Ng (2010), panel estimators have methodological power to capture temporal and spatial dynamics of price convergence which cannot be accounted for by simple time series or simple cross sections. Panel estimators are powerful to any other estimators for the fact that they allow to control for individual or market-specific heterogeneity, better to study any dynamics of price adjustments, give time ordering of marketing events, and enable to solve omitted variable problems which are the major sources of divergent implications from previous studies in Ethiopia. The panel unit root test procedures are more relevant than are other methods if the panel is of moderate size (between 10 to 250 markets or regions) and large observation (25 to 250 per market or region) (Levin *et al.,* 2002; Goldberg and Verboven, 2005).

One of the first unit root tests developed for dynamic panel model is that of Levin, Lin, and Chu (LLC) (Levin *et al*., 2002). Their test is based on analysis of a simple autoregressive equation:

$$\Delta y_{i,t} = \sim_i + \text{¢}_i t + \text{ª}_t + {}^{\circ}_i y_{i,t-1} + \mathrm{s}_{i,t}, \quad i = 1,\ldots,N, \; t = 1,\ldots,T \tag{1}$$

This model allows for two–way fixed effects ($\sim$ and ª) and market–specific time trends. The market–specific fixed effects (FE) are important sources of heterogeneity, since the coefficient of the lagged dependent variable is restricted to be homogeneous across all markets of the panel. The test involves the null hypothesis $H_0 : {}^{\circ}_i = 0$ for all against the alternative $H_a : {}^{\circ}_i < 0$ for all i with auxiliary assumptions under the null also being required about the coefficients relating to the deterministic components. The LLC test assumes that the individual units (or regions in this case) are cross-sectional independent[2].

The panel unit root tests by Im, Pesaran, and Shin (IPS) (Im *et al.*, 2003) extends the LLC framework to allow for heterogeneity in the value of ${}^{\circ}_i$ under the alternative hypothesis. Given the same equation, the null and alternative hypotheses are defined as:

$$H_0 : {}^{\circ}_i = 0, \qquad H_a : {}^{\circ}_i < 0, \quad i = 1,\ldots,N_1; \; i = N_1 + 1, \; N_1 + 2,\ldots,N. \tag{2}$$

Under the null hypothesis, all series in the panel are non stationary processes; under the alternative, a fraction of the series in the panel is assumed to be stationary. This is in contrast to the LLC test, which presumes that all series are stationary under the alternative hypothesis. The errors, $\mathrm{s}_{i,t}$,

[2] The LLC test may be viewed as a pooled Dickey–Fuller (FD) or ADF test, potentially with differing lag lengths across the units of the panel. It is applicable to small-large panels (small number of markets or regions and longer time series).

are assumed to be serially auto correlated, with different serial correlation properties and differing variances across units. IPS test proposes the use of a group–mean Lagrange multiplier (LM) statistic to test the null hypothesis. The ADF regressions can be computed for each unit, and a standardized statistic[3] computed as the average of the LM tests for each equation.

2.2. Empirical Evidence in Ethiopia

There have been different studies on commodity market performance in Ethiopia since the commodity market liberalization in 1990. Webb *et al.* (1992) have studied the spatial integration of cereal markets in Ethiopia to detect the spatial efficiency of local markets and their contributions in alleviating food shortage that has occurred because of drought in different geographical locations. Relatively better methods of analyzing market integration were used after Dercon (1995) who used cointegration technique to analyze market integration in Ethiopia and verified that most market prices were cointegrated with Addis Ababa price. After grain trade liberalization in Ethiopia, the major breakthrough in grain market efficiency studies was the one conducted by Grain Market Research Project (GMRP). With this project, Asfaw and Jayne (1997, 1998), Gebremeskel *et al.* (1998), and Eleni (2001) analyzed the response of Ethiopian grain markets to market liberalization policy and market structure, conduct, and performance of many grain markets. They estimated descriptive measures of market integration and identified many constraints in the grain marketing system.

Kindie *et al.* (2006) have analyzed the dynamics of six white wheat markets (Nazreth, Shashemenie, Jimma, Addis Ababa, Dire Dawa, and Mekelle) using vector autoregressive (VAR) model assuming the first three to be surplus and the last three to be deficit markets of wheat. They used the VAR model as a better alternative to address the simultaneous interaction of markets by identifying markets with common factors for policy intervention.

[3] Adjustment factors are used to derive a test statistic that is distributed standard normal under the null hypothesis. In the use of the IPS test a group–mean bar statistic, where the statistics from each ADF test are averaged across the panel are used. Im *et al.*(2003) demonstrate that their test has better finite sample performance than that of LLC.

They selected these markets as they were major supply and consumer markets. The monthly wholesale price levels were tested for causality using VEC mechanism (VECM) in which case Nazreth and Shashemenie were found to be price leaders while there was no exclusive price leadership of wheat markets in the country. This study was relatively more empirical than previous studies conducted in Ethiopia and the weak parameter estimates signify the considerations to be taken in sampling of relevant wheat markets to avoid risk of misspecification of price interdependence. Kindie (2007) analyzed the spatial integration of wheat markets between Ambo and Addis Ababa using autoregressive distributive lag (ARDL) approach to estimate the dynamics of price transmission and recommended that it was cost-effective to initiate price intervention measures in Addis Ababa rather than in local markets.

Degye *et al.* (2009) used a bivariate instrumental variables method to estimate shortrun and longrun dynamic adjustment coefficients between 10 markets of wheat (Addis Ababa, Nazreth, Adaba, Diksis, Ambo, Debre Birhan, Woliso, Shambu, Dessie, and Mekelle). They found that Nazreth was the dominant wheat market dictating price formation in the country which was in line with the results of Kindie *et al*. (2006). Kindie (2009) analyzed the price formation process for three cereal commodities (teff, wheat, and maize) in three markets (Addis Ababa, Nazreth, and Shashemene) using an ECM to identify a commodity which leads price formation in these three markets for cost-effective price stabilization measures. The results show that maize price plays a leadership role in the dynamics of teff and wheat prices at all market pairs except that of Addis Ababa teff market. For these three commodities, Rashid (2011) also analyzed the price leadership role by adding other three markets (Dire Dawa, Jimma, and Mekelle) but employed VAR and VEC models. He found the same result in that maize was the most significant in exacerbating price variability with respect to the persistence of shocks to itself and the two other cereals.

Degye (2011) analyzed the consumer price risk involved in Addis Ababa market for seven commodities (wheat, horse beans, beef, cow milk, egg, banana, and onions) selected from different commodity groups. After classical decomposition of the overall price risk involved into seasonality, predicable trends, and other unpredictable sources of price risk, he found that seasonality was the major source of consumer price risk.

3. Dataset and Empirical Models

3.1. The dataset

This paper utilizes monthly price series of commodities widely traded and consumed in the four major regions (Amhara, Tigray, Oromia and SNNP) and three urban centers (Addis Ababa, Dire Dawa and Harari) in Ethiopia for the period 1996 to 2013. It considers nine commodities in six regions (panels) surveyed for 211 rounds (monthly observations) per commodity. Construction of panel data of these regions and commodities allows estimation of relative price convergence of each commodity by a pooled sample of 1266per commodity.

Major commodities used in the analysis were selected by their weights in the consumption bundle used by Central Statistical Agency (CSA) to compute the consumer price index (CPI) at national and regional level. The weight of each group of commodities is indicated in Table 1. According to the 2013 estimation of the CPI by the CSA, food covers about 57% of the consumption bundle in Ethiopia (CSA, 2013).

Based on the weight of each commodity in the consumption bundle, nine commodities were selected for the analysis. Three cereals (teff, wheat, maize), livestock and livestock products (beef, sheep), fruits (banana, orange), and vegetable and tuber crops (onions, potato) were the major commodities widely traded and consumed in these regions of Ethiopia. To conduct the price and consumption survey in these regions, CSA has used 100 markets and a maximum of 193 commodities (CSA, 2013). The CSA aggregates the market level monthly average prices to regional and national levels in order to compute the national and regional CPI by commodities and commodity groups. But to minimize problems related to aggregation and to secure reliability of the aggregation process in this study, monthly prices were aggregated to regional and national levels by the author.

Table 1: Major food groups and their weights from consumption expenditure

No.	Major group	Weights (2006 base year)
1.	**Food**	0.570
1.1.	Cereal, un-milled	0.112
1.2.	Milk, cheese and egg	0.095
1.3.	Food taken away from home	0.054
1.4.	Pasta, bread etc.	0.040
1.5.	Coffee and tea leaves	0.040
1.6.	Oils and fats	0.036
1.7.	Meat	0.032
1.8.	Pulses	0.030
1.9.	Vegetables and fruits	0.020
1.10.	All other food items	0.111
2.	**Non-Food**	0.430
2.1.	House rent, construction materials, water and fuel and power	0.187
2.2.	Clothing and footwear	0.120
2.3.	Furniture, furnishing, household equipment and operation	0.070
2.4.	All other non-food items	0.053
	Total	1.000

Source: Adapted from CSA (2013).

To investigate the dynamics of relative price convergence of commodities among regions, the 18-year monthly price data was divided into two periods, nine years each (1996-2004 and 2005-2013). The first period is characterized by depressed commodity prices and the second by skyrocketing food prices exhibited since 2004/05 in Ethiopia. The point estimates from each period were compared between each other as well as to the pooled data (1996-2013). The changes in the speed and rate of price convergence were used to report on the rate and speed of relative price convergence of the commodities among regions in Ethiopia.

3.2. Benchmark price

Relative prices of all regions were computed as a ratio of the benchmark price in Addis Ababa. Selection of a benchmark region and price requires adequate information on the supply chain channels, origin and destination markets, and commodities traded and consumed in each region. The benchmark region is assumed to have the supply of and the demand for all commodities under consideration. Addis Ababa is the best candidate benchmark market in Ethiopia since it has trade relationships with all regions either by supplying to or receiving from the regions. This is partly verified by the relative prices computed by selecting Addis Ababa price as a benchmark (Table 2). Moreover, Addis Ababa is the best candidate because it is geographically separated from all regions except Oromia. There is no any other region satisfying the assumption of geographical separation in spatial price convergence and integration.

Table 2: Mean values of monthly relative prices of commodities by regions (1996-2013)

Region	Teff	Wheat	Maize	Beef	Sheep	Banana	Orange	Onions	Potato
Tigray	0.94	1.06	1.07	0.89	0.80	1.24	1.11	1.38	1.75
Amhara	0.78	0.93	0.92	0.81	0.67	1.03	0.82	1.22	1.06
Oromia	0.83	0.89	0.82	0.89	0.66	0.81	0.77	1.26	1.03
SNNP	0.82	0.88	0.75	0.85	0.54	0.50	0.49	1.39	0.94
Harari	1.03	1.18	1.11	1.12	0.99	1.17	1.06	1.23	1.28
Dire Dawa	1.06	1.15	1.11	1.18	0.99	1.07	0.91	1.14	1.21
Mean value	0.93	1.04	0.92	0.92	0.69	0.96	0.85	1.36	1.34
Standard deviation									
Between	0.12	0.13	0.16	0.15	0.16	0.27	0.22	0.10	0.29
Within	0.07	0.12	0.14	0.11	0.14	0.20	0.18	0.27	0.28
Overall	0.13	0.17	0.20	0.18	0.20	0.32	0.27	0.29	0.39

Source: Author's computation.

Regions with relative prices of commodities less than unity are expected to be suppliers of those commodities while those with relative prices greater than unity are destination markets of commodities supplied from Addis

Ababa[4]. It indicates the trade relationships of all regions with Addis Ababa for all commodities under study. The computed relative prices are as expected in the real trade relationships among regions in Ethiopia.

Teff is supplied to Addis Ababa from all regions with the exception of Harari and Dire Dawa (destination regions). Addis Ababa supplies wheat received from other regions to Tigray, Harari and Dire Dawa. Maize is supplied by Addis Ababa to deficit regions (Tigray, Harari and Dire Dawa). The relative price of beef in Harari and Dire Dawa was greater than unity, which was not supported by reversed trade relationships with Addis Ababa since there was no evidence of such trade reversal. This was so because beef in these regions was supplied by other regions not included in this analysis, like Somali region. The case for beef also holds true for live animal trade like sheep. Relative prices of sheep were less than unity for all regions possibly because there were interregional trade relationships which influence price levels directly and indirectly.

3.3. Empirical model of price convergence

Using point estimates of price adjustment, results from previous studies suggest the existence of market integration across markets in Ethiopia (Kindie *et al.*, 2009; Degye *et al*., 2009; Rashid and Negassa, 2011). However, there is no empirical evidence generated so far on both absolute and relative price convergence of commodities among markets and regions in the country. Moreover, many markets and regions and commodities are not yet analyzed due to methodological and data problems discussed earlier. This study differs from previous studies in that it employs the panel unit root tests, fixed-effects, and half-life methods to examine the rate and speed of relative price convergence of major commodities traded and consumed among the major regions and cities in the country. These procedures are

[4] Relative price greater than or less than unity may not necessarily be an indicator of supply or destination regions. Relative prices of regions might be greater than or less than unity without trade relationships of the commodities if there are other interregional trade relationships. Price could also be indirectly related through the reference market in the absence of direct trade relationships.

considered more powerful than the conventional unit root tests and models of simple time series and simple cross-sections. At the minimum, the methodologies improve the power of unit root tests because they provide a larger number of data points and use the variation across regions to improve estimation efficiency (Susanto *et al.*, 2008). The empirical knowledge on the speed and rate of relative price convergence of commodities among regions has a paramount importance in designing and implementing market intervention policies.

To apply the methods of panel unit root tests on monthly prices, Addis Ababa was selected to compute the relative prices of all other regions following the method used by Parsley and Wei, (1996), Susanto *et al.*(2008), Cecchetti *et al.*(2002), and Goldberg and Verboven (2005).

The general dynamic panel model employed is specified as (Levin *et al.*, 2002)

$$\Delta RP_{i,t} = \sim_i + \text{ª}_t + {}_i RP_{i,t-1} + \sum_{h=1}^{h=k} \text{¥}_h \Delta RP_{i,t-h} + \text{£}_{i,t} \quad \textbf{(3)}$$

Where $RP_{i,t}$ is the log of relative prices (the log of the price level of region i relative to the price level of benchmark[5] market j at time t: $RP_{i,t} = \ln\left(\frac{P_{i,t}}{P_{j,t}}\right)$.

In equation (3), $\sim_i$ is a region-specific constant to control for non-time-dependent heterogeneity across regions, Δ is the first difference operator, ª_i is a common time effect in market i, ¥_h is the historical effect of the relative price, k is the lag length, and $\text{£}_{i,t}$ is the white noise error term. The lag structure h can be routinely determined on a variety of information criteria as in a univariate Dickey- Fuller (DF) or Augmented Dickey-Fuller (ADF) test to account for possible serial correlation (Dickey and Pantula, 1987; Dickey and Fuller, 1979).

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[5] The natural benchmark for $RP_{i,t}$ is zero. Identification of benchmark market requires the knowledge on supply and terminal markets for each type of commodity traded.

The interest in equation (3) is the coefficient on the lagged relative price, $_i$, which represents the rate of convergence[6]. Under the null of no price convergence, $_i$ is equal to zero $(H_0 : \ _i = 0 \ \ for \ \ all \ \ i)$, suggesting that a shock to $RP_{i,t}$ is permanent. That is, the LLC test specifies the null hypothesis of H_0 against the alternative hypothesis of $H_a : \ _i = \ \ < 0$, where   is the normalized adjustment coefficient. The major limitation of the LLC test is that $_i$ is the same for all observations. To relax this assumption, Im *et al*. (2003) proposed an extension of the LLC procedure by allowing $_i$ to differ across groups. They tested the null hypothesis $H_0 : \ _i = 0$ against the alternative that $H_a : \ _i < 0$ for at least one market.

To conduct the LLC test, several steps are performed. The first task is to remove the influences of time effects by subtracting the cross-sectional averages from the data:

$$\overline{RP}_{i,t} = RP_{i,t} - \frac{1}{N}\sum_{i=1}^{i=N} RP_{i,t} \qquad \textbf{(4)}$$

Second, for each market, the first difference of relative prices $(\Delta RP_{i,t})$ is regressed on its lagged values $(\Delta RP_{i,t-h})$, a constant $(\sim_i)$, and a trend $({}^{a}{}_i)$ and save the residuals from equation (1) as $\hat{e}_{i,t}$. Third, the lag of relative prices $(RP_{i,t-1})$ is regressed on the same variables in the second step to obtain the residuals $\left(\hat{v}_{i,t-1}\right)$. Fourth, the residuals $\hat{e}_{i,t}$ is regressed on $\left(\hat{v}_{i,t-1}\right)$ without a constant. If we denote the residuals from the final

[6] The closer the rate of convergence to zero, the longer is the estimated half-life of a shock.

regression as $\hat{\pounds}_{i,t}$, the regression standard error obtained from this regression can be defined as

$$\hat{\pm}_{\pounds_i} = \sqrt{\left(T - k - 1\right)^{-1} \sum_{t=h_i+2}^{T} \hat{\pounds}_{i,t}^{2}} \; . \qquad \textbf{(5)}$$

This estimate of the standard error will be used to normalize the residuals from the contemporaneous and lagged auxiliary regressions, $\hat{e}_{i,t}$ and $\hat{v}_{i,t-1}$, for controlling heterogeneity across individual regions. The normalized values of the residuals from the two auxiliary regressions are:

$$\tilde{e}_{i,t} = \frac{\hat{e}_{i,t}}{\hat{\pm}_{\pounds_i}} \text{ and } \tilde{v}_{i,t-1} = \frac{\hat{v}_{i,t-1}}{\hat{\pm}_{\pounds_i}} \qquad \textbf{(6)}$$

Finally, the panel OLS of the normalized residuals will be run to obtain the adjustment estimates (Levin *et al*., 2002; Susanto, *et al.*, 2008)[7]:

$$\tilde{e}_{i,t} = \quad \tilde{v}_{i,t-1} + \tilde{\pounds}_{i,t} \qquad \textbf{(7)}$$

The IPS model is a unit root test for heterogeneous dynamic panels based on the mean-group approach. This approach is similar to the LLC model, in that it allows for heterogeneity across sectional units. Instead of pooling the data, the IPS model uses separate unit root tests for the N cross-section units

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[7] Under the null hypothesis $H_0 : \quad = 0$, the regression statistic $(t \quad)$ has a standard normal limiting distribution for the model that does not include both an individual specific mean and a time trend; however, it diverges to negative infinity for the model having an individual-specific mean, but does not contain a time trend and the model that contains both an individual-specific mean and time trend. To obtain a standard normal distribution, the *t* statistic can be adjusted (denoted as *t-star*: t^{*}) using the ratio of longrun and shortrun standard deviations computed by the procedures of adjusting standard normal distribution for the LLC model (Levin *et al.,* 2002).

(regions). The *t* statistic (denoted as, $\bar{t}$, t-bar) is the average of *t* statistics for all N regions. Let $t_{i,T}$ denote the t statistics for testing unit roots, and let $E(t_{i,T}) = «$ and $V(t_{i,T}) = \pm^{2}$, then

$$\sqrt{N}\left(\frac{\bar{t}_{N,T} - «}{\pm}\right) \sim N(0.1), \text{ where } \bar{t}_{N,T} = \frac{1}{N}\left(\sum_{i=1}^{i=N} t_{i,T}\right) \qquad \textbf{(8)}$$

The problem of the above equation is computing the mean, «, and the variance, $\pm^{2}$. Assuming that the cross sections (regions in this case) are independent, the standardization of the t-bar statistics, using the means and variances of $t_{i,t}$ evaluated under $=0$, denoted as w-t $(\bar{t}_{¤}^{*})$ bar statistic is given by (Im *et al.*, 2003):

$$\bar{t}_{¤}^{*} = \frac{\sqrt{N}\left(\bar{t}_{N,T} - E\left(\bar{t}_{N,T}\right)\right)}{\sqrt{\mathrm{var}\left(\bar{t}_{N,T}\right)}} \qquad \textbf{(9)}$$

where $\bar{t}_{N,T}$ is the average *t* statistic for each region or individual unit, and $E\left(\bar{t}_{N,T}\right)$ and

$ar\left(\bar{t}_{N,T}\right)$, respectively, are its mean and variance, respectively.

4. Empirical Results

4.1 Patterns of price convergence

The purpose of this study is to analyze the rate and speed of relative food price convergence of commodities in Ethiopia. It considers 18-year monthly data of consumer prices of 9 commodities widely consumed and traded among the seven regions and markets in Ethiopia (including the benchmark market). In order to investigate the rate and speed of relative price convergence of these commodities, the study period was classified into three as period 1 (1996-2004), period 2 (2005-2013), and the full period (1996-2013). As a preliminary investigation of the pattern of price convergence to unity, the mean values of relative prices of commodities during the two periods were tested for the presence of significant difference. As reported in Table 3, the tests for five commodities (cereals, banana, and potato) suggest the presence of mean relative price difference between the two periods. The null hypothesis that there is no mean difference between the two periods was rejected at one percent level. But there is not significant relative mean price difference for the other commodities. This preliminary test leads to further investigation of the expectations with more rigorous methods.

Table 3: Two-sample mean-comparison test of relative prices between the two periods

Commodity	Mean value			t-value
	1996-2004	2005-2013	1996-2013	
Teff white	0.88	0.94	0.91	-8.23***
Wheat white	1.05	0.98	1.01	7.86***
Maize	1.02	0.90	0.96	11.52***
Beef	0.95	0.96	0.96	-0.83
Sheep 10-15 kg	0.72	0.71	0.72	0.66
Banana	0.86	1.08	0.97	-13.47***
Orange	0.85	0.87	0.86	-1.10
Onion	1.27	1.27	1.27	-0.40
Potato	1.23	1.19	1.21	1.87**
Number of observations	642	624	1266	

Notes: *** indicates 1% significance level.

Source: Author's computation.

4.2 Panel unit root tests

Before estimating the rate and speed of monthly price convergence of commodities among regions in Ethiopia, the panel dataset was tested for the presence of unit roots (Table 4). The study employs two widely applied methods of unit root tests in balanced and heterogeneous panels, the IM-

Pesaran-Shin (IPS) and the Levin-Lin-Chu (LLC). The use of IPS panel unit root test in panel data accounts for the limitations of LLC method, which could only be employed in a strongly balanced data. The lag structure in the ADF regression was routinely identified by Akaike's Information Criteria (AIC).

The results from the two methods of panel unit root tests were consistent. As shown in table, all point estimates were negative, as expected, and strongly significant at one percent level. Therefore, the null that panels contain unit roots was rejected and the alternative that panels are stationary was accepted in the LLC tests. Similarly, the IPS test rejects the null hypothesis of unit roots regardless of the sample periods. Both tests confirm that the panel dataset of the relative prices were stationary at their first differences suggesting that linear panel estimators could be employed to estimate rate of price convergence of commodities among the regions.

Table 4: LLC and IPS tests of panel unit roots for monthly relative price changes in Ethiopia

Percentage change in relative prices	**LLC test (adjusted t*): Ho: Panels contain unit roots: Ha: Panels are stationary**			**IPS test (W-t-bar): Ho: All panels contain unit roots; Ha: Some panels are stationary**		
	1996-2004	**2005-2013**	**1996-2013**	**1996-2004**	**2005-2013**	**1996-2013**
Teff	-29.3***	-11.6***	-7.4***	-28.8***	-19.7***	-2.0***
Wheat	-21.8***	-22.4***	-12.6***	-23.5***	-25.4***	-25.3***
Maize	-26.0***	-8.6***	-13.4***	-27.6***	-18.9***	-26.4***
Beef	-2.8***	-10.1***	-4.7***	-17.5***	-18.5***	-23.9***
Sheep	-12.8***	-13.1***	-3.6***	-21.6***	-21.3***	-25.6***
Banana	-23.5***	-28.3***	-1.5***	-27.2***	-27.2**	-25.1***
Orange	-16.4***	-25.3***	-17.6***	-22.3***	-25.6***	-27.4***
Onion	-11.4***	-3.8***	-14.2***	-19.5***	-16.0***	-26.6***
Potato	-11.3***	-12.9***	-2.7***	-20.5	-21.7***	-25.4***

Notes: *** indicates 1% significance level.
Source: Author's computation.

4.3 Rate of price convergence

The magnitude or rate of relative price convergence of commodities among the regions was measured by parameter estimates for lagged percentage relative price changes. The general dynamic panel model of relative price convergence (%) is a function of previous-month relative price (log) and the cumulative effects of its own lagged differences. The rate of convergence in this paper is defined as the effect of previous-month relative price (%) on the contemporaneous percentage relative price changes of a commodity over the years (months) and across regions. The model was estimated by the FE method since the inter-temporal price difference in each region has fixed-effects on the percentage relative price changes. Table 5 summarizes the FE estimation of rate of relative price convergence of commodities among regions in Ethiopia. The rates of convergence of relative price changes for the nine commodities are comparatively reported over the three periods.

Table 5: Fixed-effects estimation of rate of price convergence (%) among regions

Change in annual relative prices (%)	1996-2004		2005-2013		1996-2013	
	Convergence rate	**Standard error**	**Convergence rate**	**Standard error**	**Convergence rate**	**Standard error**
Teff	-0.27***	0.03	-0.19***	0.03	-0.13***	0.02
Wheat	-0.37***	0.04	-0.32***	0.04	-0.21***	0.02
Maize	-0.35***	0.03	-0.57***	0.05	-0.26***	0.03
Beef	-0.24***	0.04	-0.26***	0.04	-0.17***	0.02
Sheep	-0.43***	0.05	-0.29***	0.04	-0.34***	0.03
Banana	-0.21***	0.03	-0.32***	0.04	-0.07***	0.01
Orange	-0.43***	0.04	-0.45***	0.03	-0.32***	0.03
Onion	-0.04	0.04	0.02	0.02	-0.04	0.03
Potato	-0.50***	0.05	-0.44***	0.04	-0.36***	0.03

Notes: The optimum lag lengths routinely identified by AIC for each commodity and period are rounded off to their positive integers in order to be used in the estimation of convergence rate in the FE model.

*** and ** indicate significance level at 1% and 5%, respectively.

Source: Author’s commutation

As shown, all the coefficients, with the exception of onion, were negatively and strongly significant and the signs of all parameter estimates were as expected. The rate of relative price convergence among regions in the 18-year period was generally low for all commodities. The maximum and the minimum convergence rate estimated for the whole period was 36 percent for potato and 7 percent for banana, relatively more perishable. A unit percentage shock of relative prices of potato in previous-year had only 0.5 percent change in the current-year relative price convergence with the benchmark market (Addis Ababa). It is generally evidenced that the rate of convergence was very low (7%-36% for the whole period) because greater proportion of the price shocks were unadjusted among regions. Low or nearer-to-zero rates of convergence are indications of permanent shocks which could not be adjusted among regions. Food items covering about 57 percent of the consumption expenditure in Ethiopia have exhibited very low rates of relative price convergence which would deplete the welfare effects expected from spatial and temporal arbitrage (CSA, 2013).

Under the null of no price convergence, the adjustment coefficient for onion was equal to zero, suggesting that the price shock to the relative price was permanent. Due to its perishable nature, there was no significant convergence in the relative prices of onion over the 18-year period and across regions. This evidence is supported by previous studies which verify that marketing agents involved in the production and marketing of onion, particularly onion producers, have used to face exceptionally high price risk in Ethiopia (Degye, 2011).

Some food commodities have undergone increased rate of monthly relative price convergence in the 18-year period. The maximum change in the rate of convergence from the first to the second period was for maize (22%) followed by banana (11%) and potato (5%). This is a good indication of improved domestic market performance favoring the prices of these commodities. However, there are some commodities with significant price divergence as an indication of poor market performance. The relative prices of teff, wheat and sheep were rather diverging at a rate of 8 percent, 5 percent and 14 percent, respectively. The findings point out that the market intervention measures implemented in the country had both negative and positive repercussions. This down turn of price performance of some commodities is possibly explained by the supply of and the demand for the commodities among regions, and the price stabilization measures related to direct price setting (price ceilings) by the government.

4.4 Speed of price convergence

The average speed at which commodity prices move toward parity price is commonly measured by half-life method. Speed of relative price convergence of commodities among regions in this study is measured by half-life of relative price shocks. Half-life of a shock may be defined as a period in which the marginal change in the stationary component of the response becomes half of the initial response. The closer the rate of convergence to zero, the longer is the estimated half-life of a shock. It is computed as $-\ln(2)/\ln(Rho)$ where rho is the correlation of errors (within and between) panels. The estimated half-life of commodities is reported in Table 6.

Speeds of relative price convergence of five commodities under consideration were fairly long in the first period and short in the second period, indicating increased responsiveness of markets in the regions. In the first period, regions tended to slowly and partially respond to relative price shocks of these commodities. In the second period, they required a few days to adjust to half of the relative price shocks with the benchmark market. For instance, the time required to adjust to half of the shocks was reduced by about 38 percent (11.5 days) for potato, 31 percent (9.4 days) for wheat, 30 percent (9.2 days) for teff, and 25 percent (7.6 days) for orange. The findings markedly show that some commodities have experienced considerable improvement in the speed of their price convergence. The computed half-life[8] of relative price shocks was adjusted among regions and over time as

[8] The half-life method estimates only the time required to adjust to half of the initial relative price shock. The remaining half of the shock persists as a permanent shock which might not be adjusted among regions.

rapidly as possible. On the other hand, banana and maize were characterized by price divergence at a rate of 41% percent and one percent, respectively.

Table 6: Estimated half-life of shocks as a measure of speed of relative price convergence

Commodity	1996-2004		2005-2013		1996-2013	
	Half-life	**Days**	**Half-life**	**Days**	**Half-life**	**Days**
Teff	0.69	20.8	0.39	11.6	0.30	9.0
Wheat	0.73	21.9	0.42	12.5	0.32	9.7
Maize	0.62	18.5	0.63	18.8	0.34	10.3
Beef	0.42	12.7	0.40	12.1	0.29	8.7
Sheep	0.49	14.7	0.44	13.3	0.43	13.0
Banana	0.76	22.8	1.17	35.0	0.25	7.4
Orange	0.90	27.1	0.65	19.5	0.50	15.0
Potato	0.72	21.6	0.34	10.1	0.37	11.1

Notes: The speed of relative price convergence can be converted to days by changing the (unit rate or percentage) half-life into 30 days rate (i.e., half-life multiplied by 30 and the product divided by the rate, 1 or 100).
Source: Author's commutation.

The attempt to categorize the 18-year period into two using relatively observable market and policy events has generally resulted in some meaningful implications on the patterns of rate and speed of food convergence among regions and cities in Ethiopia. Market policy measures initiated in Ethiopia before and after 2004 have some positive and negative implications on the performance of commodity markets and prices across regions and cities in Ethiopia.

5. Concluding Remarks

This paper investigates the dynamics of monthly relative price convergence of major commodities consumed and traded in seven regions and cities of Ethiopia. It utilizes monthly consumer prices of nine commodities in these regions and cities. A pane dataset comprising the regions (panels) over 211 months with 1266 data points was constructed for each commodity. Relative price levels and changes were computed based on the benchmark price in Addis Ababa.

Panel unit root tests were conducted by employing two widely used techniques, the LLC and IPS unit root tests. Rate and speed of relative price convergence of commodities among regions were measured by dynamic adjustment parameter estimates and half-life of relative price shocks, respectively. The empirical results indicate that the panel datasets of the log-normalized relative prices were stationary at their first differences. The estimated rates of relative price convergence of commodities among regions were generally low. The speed of relative price convergence of commodities measured by the half-life method was, however, mixed. There was significant improvement in the rate and speed of relative price convergence of some food commodities among regions during the 18-year period.

The findings generally suggest the need to initiate and implement marketing policies designed to optimize the net welfare effects emanating from domestic market performance and interventions. Improving marketing infrastructure, information and institutional arrangement and marketing policies will improve market performance by improving the responsiveness of actors to price shocks. Interventions related to value-adding activities, particularly for perishable commodities, will improve convergence to unitary price among region of Ethiopia. It is also a policy imperative to promote sustained supply-side intervention in order to increase production and productivity of agricultural commodities characterized by low or declining relative price convergence.

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Appendix Table 1: Optimum lags (month) specified by AIC in the ADF regression of panel unit root tests

Percentage change in relative prices	Periods		
	1996-2004	2005-2013	2005-2013
Teff	0.83	2.50	3.50
Wheat	1.33	1.50	2.83
Maize	0.83	2.50	2.83
Beef	3.00	2.50	3.50
Sheep	2.00	2.00	3.67
Banana	1.17	1.17	3.50
Orange	1.67	0.67	2.50
Onion	2.50	4.50	3.00
Potato	2.50	2.33	4.50

Source: Authors' computation.