

Minimum Wage Legislation and Employment Performance in Zambia: An ARDL Bounds Testing and Granger Causality Analysis, 1997 to 2024

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Abstract

Zambia's statutory minimum wage for domestic workers was revised only six times between 1982 and 2024. During the long intervals between revisions, inflation repeatedly eroded the real value of the wage floor by more than a third before a new adjustment was gazetted. Whether this pattern of infrequent, discrete, and partially eroded wage floors has measurable consequences for aggregate employment remains an open empirical question. Using annual data for 1997–2024, this study examines the short-run effects, long-run association, and direction of predictive causality between the real statutory minimum wage and employment performance in Zambia. The employment-to-population ratio is modelled as a function of the real statutory minimum wage for domestic workers, real gross domestic product growth, and consumer price inflation. Augmented Dickey–Fuller, Phillips Perron, and Kwiatkowski Phillips Schmidt Shin tests indicate a mixture of I(0) and I(1) series, with no variable integrated of order two, thereby supporting use of the autoregressive distributed lag bounds-testing procedure of Pesaran, Shin and Smith (2001). The bounds-test F-statistic is 3.792, which lies within the inconclusive band at the 10 per cent level. The error-correction coefficient is nevertheless negative and highly significant at -0.951 ($t = -3.572$), providing supplementary evidence of rapid adjustment, conditional on the existence of a stable long-run relation. The estimated long-run elasticity of the employment-to-population ratio with respect to the real minimum wage is positive at 0.078 and statistically significant at the 1 per cent level; its sign and significance are preserved under dynamic ordinary least squares, trend-augmented static regression, and static regression with heteroskedasticity- and autocorrelation-consistent standard errors. In the short-run, the real minimum wage has a positive but-statistically insignificant coefficient, while pairwise, multivariate, and Toda–Yamamoto tests do not establish a robust causal ordering in either direction. The aggregate evidence therefore provides no support for a negative employment effect of the statutory wage floor over the period examined. The positive long-run association is consistent with monopsonistic and efficiency-wage mechanisms, but it may also reflect common trends, limited enforcement, and extensive informality. The findings consequently support cautious regularisation of minimum-wage policy while underscoring the greater importance of structural transformation and employment-intensive growth.

1. Introduction

The statutory minimum wage occupies an unusual place in economic policy. Few instruments are so widely used, so readily understood by the public, and yet so persistently disputed among economists. The International Labour Organization defines it as the floor of the wage structure, intended to protect workers at the lower end of the earnings distribution (Carr and Stuart, 2023). Its principal policy rationale is to reduce wage inequality and improve the living standards of low-income workers (Arribas, Cardenas and Rial, 2024). Evidence generally confirms that an enforced wage floor raises earnings near the bottom of the distribution. The unresolved question is whether, and under what labour-market conditions, those gains come at the cost of employment.

The theoretical case for job losses is straightforward. In a competitive labour market in which firms hire until the marginal revenue product of labour equals the wage, a legislated wage above the market-clearing level creates an excess supply of labour, and the burden of the resulting unemployment falls disproportionately on low-skilled and inexperienced workers. Against this stands a body of theory in which the competitive assumption fails. Where employers possess wage-setting power, a moderate wage floor can raise employment as well as wages by moving the labour market towards the competitive outcome (Manning, 2003; Boal and Ransom, 1997). Where wages influence effort, turnover and the quality of applicants, a higher wage can be self-financing through productivity gains (Shapiro and Stiglitz, 1984; Akerlof and Yellen, 1986). The empirical resolution of this dispute is therefore an inherently local matter, contingent on the structure of the labour market in which the wage floor is imposed.

Zambia offers a particularly instructive setting in which to examine this question, for three reasons. The first is the character of the policy instrument itself. The Minimum Wages and Conditions of Employment Act was enacted in 1982 and requires biennial revision, yet the statutory minimum wage for domestic workers was adjusted only six times in the four decades to 2024, in 1997, 2002, 2006, 2012, 2018 and 2022. Adjustments are therefore large, discrete and infrequent rather than incremental. The second is the consequence of that infrequency. Because the intervals between revisions have consistently exceeded the period over which inflation halves the purchasing power of a fixed nominal amount, the real wage floor has followed a repeated saw tooth pattern of sharp upward correction followed by prolonged erosion. Between the revision of 2012 and the revision of 2018 the real value of the domestic worker minimum wage fell by approximately thirty-seven per cent, and between 2018 and 2021 it fell by approximately thirty-five per cent. The policy that firms actually face is therefore far less binding, on average, than the nominal headline suggests. The third reason is the structure of the labour market itself. Trade union density is low, enforcement capacity is limited, and informal employment absorbs the majority of the labour force, which together attenuate the mechanical transmission from legislated wage floors to labour costs across the economy as a whole.

Despite the importance of this policy question, the Zambian evidence remains limited and largely descriptive. Mufungo (2023), studying the retail sector in Lusaka Province, reports that small firms responded to wage increases by reducing hours or postponing recruitment, whereas larger firms absorbed the additional cost through productivity gains or price adjustments. Zimba (2022) finds comparable heterogeneity among small-scale businesses in Lusaka. The Zambia Institute for Policy Analysis and Research (2023), in its review of the 2022 adjustment, focuses chiefly on compliance rather than employment outcomes. None of these studies directly estimates the aggregate time-series relationship, nor do they use a cointegration framework that distinguishes short-run adjustment from a possible long-run equilibrium. The present study addresses that gap.

The general objective is to examine the relationship between minimum-wage legislation and employment performance in Zambia. Three specific objectives guide the analysis: first, to estimate the short-run response of employment to changes in the real statutory wage floor; second, to test for and estimate a long-run relationship between the two variables; and third, to assess predictive precedence in both directions using complementary causality tests.

The study makes four contributions. First, to the authors' knowledge, it is the first application of the autoregressive distributed lag bounds-testing framework to minimum wages and aggregate employment in Zambia. Second, it constructs a deflated real minimum-wage series rather than relying on the nominal statutory figure—a necessary distinction given the scale of inflation-induced erosion. Third, it tests the central long-run coefficient across four estimators and examines causality through three complementary procedures, thereby reducing dependence on any single specification. Finally, it maintains a careful distinction among a statistically insignificant short-run effect, a positive but non-causal long-run association, and formally inconclusive evidence of cointegration. That distinction is essential to a credible interpretation of the policy implications.

The remainder of the paper is organised as follows. Section 2 reviews the theoretical and empirical literature. Section 3 sets out the data, the model and the estimation strategy. Section 4 presents and discusses the results. Section 5 concludes and draws out the policy implications.

2 Literature Review

2.1 Theoretical framework

Three theoretical traditions offer competing predictions about the employment consequences of a statutory wage floor, and the empirical work reported below is best understood as an attempt to discriminate between them in the Zambian context.

The neoclassical model of a competitive labour market provides the benchmark. Firms and workers are price takers, firms hire until the marginal revenue product of labour equals the wage, and the equilibrium wage clears the market. A wage floor set above that equilibrium raises the quantity of labour supplied and reduces the quantity demanded, and the difference is involuntary unemployment concentrated among workers whose productivity lies below the mandated wage. The prediction is unambiguous: employment falls, and it falls most sharply among the young and the low-skilled. This model supplies the expected negative sign against which the estimated coefficients in Section 4 are evaluated.

Efficiency wage theory, originating in the observation of Leibenstein (1957) that higher wages improve nutrition and therefore productivity in poor economies, and developed in the shirking model of Shapiro and Stiglitz (1984) and the fair wage formulations surveyed by Akerlof and Yellen (1986), reverses part of that logic. If effort, turnover and the quality of the applicant pool respond positively to the wage, then a firm paying above the market-clearing wage may face a lower cost per efficiency unit of labour. A legislated increase that moves firms towards the efficiency wage need not reduce employment, and the productivity response may offset the direct cost increase. The theory is particularly relevant in settings where monitoring is costly and turnover is high, which describes much of Zambian domestic and low-skilled service employment.

The monopsony model provides the third and, for present purposes, the most important alternative. Where a small number of employers dominate a local labour market, or where geographic, informational and skill-related frictions restrict worker mobility, employers face an upward-sloping labour supply curve and set wages below the marginal revenue product (Manning, 2003). Employment under monopsony is inefficiently low. A wage floor imposed between the monopsony wage and the competitive wage raises both wages and employment, and only a floor set above the competitive wage generates the neoclassical disemployment result. Boal and Ransom (1997) show that the empirical relevance of this model turns on the elasticity of labour supply facing the individual employer, which is low precisely where labour markets are thin and mobility is constrained. Zambia's low union density, limited enforcement capacity and geographically segmented labour markets suggest that monopsonistic wage-setting is a live possibility rather than a theoretical curiosity.

The three models therefore predict a negative, an ambiguous and a positive employment response respectively. The empirical task is to determine which prediction the Zambian data support, and the framework adopted in Section 3 is deliberately agnostic on the sign of the minimum wage coefficient.

2.2 Empirical evidence

The international literature has moved decisively away from the confident disemployment findings of the mid-twentieth century. Card and Krueger (1994) found no employment loss following the New Jersey minimum wage increase relative to neighbouring Pennsylvania, and the border discontinuity method they pioneered has since been applied widely. Cengiz, Dube, Lindner and Zipperer (2019) use a bunching estimator on United States data and find that minimum wage increases raise wages at the bottom of the distribution with negligible effects on the total number of low-wage jobs. Doucouliagos and Stanley (2019) conduct a meta-regression analysis of eighty two studies and report that the aggregate employment effect is small once publication selection is accounted for, while the wage effect is substantial. Dube (2019) reaches similar conclusions in a review of the international evidence commissioned by the United Kingdom government. Working in the opposite direction, Neumark (2019) surveys the developing country evidence and argues that where the statutory floor is genuinely binding and strongly enforced, both significant positive wage effects and appreciable disemployment effects reappear.

Findings for the members of the Organisation for Economic Cooperation and Development are similarly divided. Kar (2018) applies the augmented mean group estimator to a panel of member countries and reports that minimum wage legislation raises employment in the long run, with a one per cent increase in the minimum wage associated with a 0.17 per cent increase in the employment rate, and notes that although country specific coefficients vary with institutional structure they are predominantly positive. Kim and Lee (2018) estimate the effect on labour demand and supply separately in a panel of twenty-five member countries and find that a higher minimum wage reduces labour demand without affecting labour supply, with a 10 per cent increase reducing employment by approximately 0.7 per cent. Sorkin (2015) draws attention to the horizon over which effects are measured, showing that short-run elasticities can be close to zero while long-run elasticities are substantially larger as firms adjust their capital intensity.

African evidence is sparser but converges on a broadly similar conclusion. Bhorat and Kanbur (2019) show that the ratio of the minimum wage to the mean wage is higher in low-income African countries than in middle income countries, indicating a comparatively aggressive stance, yet report that introducing or raising the minimum wage generally produces a small negative effect or no measurable negative effect on employment. They emphasise that the variation around this average is wide and is driven substantially by differences in compliance, and they document a positive association between the Kaitz index and non-compliance. Obeng (2015) finds no significant negative employment effect of minimum wage increases in Ghana and concludes that economic growth and sectoral composition dominate wage regulation as determinants of employment. Asuamah and Nyarkoh (2023) reach the same conclusion for Ghana while identifying displacement within specific sectors and among low-skilled and young workers. Okey, Adjor and Ketor (2020) apply fixed effects and dynamic panel estimators to a set of African countries and report a negative relationship between minimum wages and employment consistent with the neoclassical view, alongside positive effects of income and education on employment. Ikeje and Islam (2020) document the prevalence of low-wage work in Nigeria, finding that approximately eighty per cent of jobs fall into this category and that women are disproportionately represented in vulnerable low paid employment. Nwaka, Guven and Olubiyi (2016) examine the joint effects on employment and poverty across sub-Saharan Africa and find that the interaction between formal wage regulation and a large informal sector substantially modulates the measured outcome.

The Zambian literature is limited and is dominated by case study designs. Mufungo (2023) applies a mixed-methods approach to the retail sector in Lusaka Province and documents heterogeneous firm responses by size. Zimba (2022) examines small-scale businesses in retail and hospitality in Lusaka District and finds adjustment through price increases, efficiency improvements and a shift towards part-time and temporary contracts. Banda and Musonda (2022) report reduced hiring and shorter hours in urban centres following adjustments. Kaluba (2021) finds improvements in worker earnings and morale where compliance is enforced. Lupiya (2018) reports that minimum wage laws reduced the employment of working-age adults by 0.7 percentage points, with young adults disproportionately affected. The International Labour Organization (2022) documents persistent enforcement and compliance difficulties among small businesses in Zambia.

Two conclusions emerge from this literature. First, much of the recent international and African evidence points to small or negligible aggregate employment effects, although negative effects remain possible where wage floors are strongly binding and enforcement is effective. Second, the Zambian evidence is predominantly sectoral and descriptive. No existing study estimates the aggregate relationship in a time-series framework that separates short-run dynamics from longer-run association and tests predictive precedence. The present study addresses that specific gap, while recognising that aggregate time-series evidence cannot identify heterogeneous effects across workers, firms, or sectors.

3 Data and Methodology

3.1 Data and variable construction

The study employs annual time series data for Zambia covering the period 1997 to 2024, giving 28 observations. The starting year is dictated by the first postliberalisation revision of the statutory minimum wage under Statutory Instrument 119 of 1997, which provides a consistent policy baseline, and the terminal year by data availability at the time of writing.

The dependent variable is the employment-to-population ratio, defined as the proportion of the working-age population that is employed, drawn from World Bank World Development Indicators data based on International Labour Organization modelled estimates. This measure is preferred to the unemployment rate in the Zambian context because open unemployment is a poor indicator of labour market slack where subsistence and informal activity absorb workers who would otherwise be recorded as unemployed. The variable is expressed in natural logarithms so that estimated coefficients on the logged minimum wage carry an elasticity interpretation.

The principal explanatory variable is the real statutory minimum wage for domestic workers. The nominal series is compiled from the relevant Statutory Instruments, namely Statutory Instrument 119 of 1997, Statutory Instrument 2 of 2002, Statutory Instruments 56 and 57 of 2006, the 2012 Instrument, the 2018 Instrument and the 2022 Instrument, with the pre-redemption amounts converted at the 2013 rate of one thousand old kwacha to one new kwacha. The nominal series is then deflated to constant 2023 kwacha using the consumer price index. Domestic workers are used as the reference category because this schedule has been revised consistently across the whole sample period and because domestic work is the segment in which the statutory floor is most likely to bind. The real series is expressed in natural logarithms.

Two control variables are included. The rate of growth of real gross domestic product captures the state of aggregate demand, since employment creation is expected to respond positively to economic expansion. The rate of consumer price inflation captures macroeconomic instability, which may depress labour demand by raising input costs and by increasing uncertainty about future real returns. Both are expressed in percentage terms rather than logarithms because the growth series takes negative values in 1998 and 2020. A binary policy indicator taking the value one from 2018 onwards is retained for use in robustness testing, since the 2018 revision represents the largest single real increase in the sample.

Table 1: Variable definitions, measurement and expected signs

Variable	Symbol	Measurement	Source	Expected sign
Employment to population ratio	lnEPR	Employed persons as a percentage of the working age population, natural logarithm	World Bank WDI, ILO modelled estimates	Dependent
Real minimum wage	lnRMW	Statutory monthly minimum wage for domestic workers deflated to constant 2023 kwacha, natural logarithm	Statutory Instruments; deflated using World Bank CPI	Ambiguous
Economic growth	GDPG	Annual percentage growth of real gross domestic product	World Bank WDI	Positive
Inflation	INF	Annual percentage change in the consumer price index	World Bank WDI	Negative
Policy indicator	D2018	Binary variable equal to one from 2018 onwards, zero otherwise	Statutory Instrument of 2018	Ambiguous

Source: compiled by the authors.

3.2 Model specification

Employment performance is modelled as a function of the real statutory wage floor and of macroeconomic conditions, which gives the general functional relationship

$$\ln EPR_t = f(\ln RMW_t, GDPG_t, INF_t) \dots (1)$$

The empirical model is estimated within the autoregressive distributed lag framework developed by Pesaran, Shin and Smith (2001). The approach accommodates a mixture of I(0) and I(1) regressors, provided that none is I(2), and brings short-run dynamics and lagged level relationships into a single equation. It is also well suited to comparatively small samples. That advantage must not be overstated: with only 28 annual observations, lag selection, statistical power, and finite-sample inference remain important constraints. The general autoregressive distributed lag specification is

$$\ln EPR_t = \beta_0 + \sum \beta_i \ln EPR_{t-i} + \sum \alpha_j \ln RMW_{t-j} + \sum \delta_k GDPG_{t-k} + \sum \zeta_l INF_{t-l} + u_t \quad (2)$$

For estimation and testing this is reparameterised as the conditional error correction model

$$\Delta \ln EPR_t = c + \gamma \Delta \ln EPR_{t-1} + \sum \theta_j \Delta \ln RMW_{t-j} + \sum \phi_k \Delta GDPG_{t-k} + \sum \psi_l \Delta INF_{t-l} + \lambda_1 \ln EPR_{t-1} + \lambda_2 \ln RMW_{t-1} + \lambda_3 GDPG_{t-1} + \lambda_4 INF_{t-1} + \rho t + \varepsilon_t \quad (3)$$

In equation (3), the coefficients on the differenced terms measure short-run responses, while the lagged level terms jointly test for a long-run relationship; t denotes a deterministic linear trend. The long-run elasticity with respect to variable X is recovered as the negative ratio of the coefficient on the lagged level of X to the coefficient on the lagged dependent variable, with delta-method standard errors. When a stable level relationship exists, the coefficient on the lagged dependent variable can be interpreted as the speed at which disequilibrium is corrected. Because the bounds evidence is inconclusive, that interpretation is treated as conditional rather than definitive.

3.3 Estimation and testing strategy

The empirical procedure follows six steps. First, descriptive statistics and a pairwise correlation matrix are computed, with variance inflation factors used to check for multicollinearity. Second, the order of integration of each series is established using the Augmented Dickey–Fuller and Phillips Perron tests, in which non-stationarity is the null hypothesis, together with the Kwiatkowski–Phillips–Schmidt–Shin test, in which stationarity is the null hypothesis. The joint use of tests with opposing nulls provides a more reliable classification than reliance on either alone, which matters in a sample of this size. Lag length for the Augmented Dickey–Fuller regressions is selected by the Akaike information criterion subject to a maximum of three lags.

Third, the lag structure of the conditional error-correction model is selected by the Akaike information criterion over a restricted grid, with a maximum of two lags on the dependent variable and one lag on each differenced regressor; candidate specifications must also satisfy the reported residual diagnostics. Fourth, cointegration is assessed using the F-statistic for the joint significance of the lagged level terms in equation (3), compared with the critical-value bounds of Pesaran, Shin and Smith (2001). An F-statistic above the upper bound rejects the null of no level relationship; a statistic below the lower bound fails to reject it; and a statistic between the bounds is inconclusive. Given the low power of the procedure in short samples, three supplementary diagnostics are reported: the sign, magnitude, and significance of the error-correction coefficient; Johansen trace and maximum-eigenvalue statistics adjusted for small-sample bias; and the stability of the long-run coefficient across alternative estimators. These checks inform interpretation but do not override an inconclusive bounds test.

Fifth, the long-run coefficients are estimated from the preferred conditional error correction model and are then re-estimated by dynamic ordinary least squares with one lead and one lag of the differenced regressors, by static regression augmented with a deterministic trend, and by static regression with heteroskedasticity and autocorrelation consistent standard errors. Agreement across these estimators is treated as evidence that the long-run result is not an artefact of a particular specification. Sixth, the direction of causality is examined using pairwise Granger causality tests on the differenced series, block exogeneity tests within a four-variable vector autoregression, and the Toda–Yamamoto (1995) procedure, which augments a levels vector autoregression by the maximum order of integration and tests restrictions on the first k lags only, thereby remaining valid irrespective of the cointegration status of the system.

Residual diagnostics comprise the Breusch Godfrey Lagrange multiplier test for serial correlation at one and two lags, the Breusch Pagan Godfrey and Harvey tests for heteroskedasticity, the Lagrange multiplier test for autoregressive conditional heteroskedasticity, the Jarque Bera test for normality and the Ramsey regression specification error test for functional form. Parameter stability is assessed using the cumulative sum and cumulative sum of squares of recursive residuals. All estimation is carried out in Python using the statsmodels and arch libraries.

4 Results and Discussion

4.1 Descriptive statistics

Table 2 reports summary statistics for the variables in levels and in logarithms over the 28 observations from 1997 to 2024.

Table 2: Descriptive statistics, 1997 to 2024

Statistic	EPR (%)	RMW (2023 ZMW)	lnEPR	lnRMW	GDPG (%)	INF (%)
Mean	53.829	1236.290	3.985	7.079	5.115	14.492
Median	53.650	1261.405	3.982	7.140	4.925	11.720
Maximum	59.000	1884.810	4.078	7.542	10.300	26.788
Minimum	49.872	585.040	3.910	6.372	-2.790	6.429
Standard deviation	2.537	341.748	0.047	0.301	2.784	6.811
Skewness	0.282	-0.094	0.214	-0.655	-0.760	0.456
Kurtosis	2.068	2.270	1.999	2.809	4.535	1.699
Jarque Bera	1.454	0.808	1.449	1.945	3.741	2.802
Probability	0.483	0.668	0.485	0.378	0.154	0.246
Observations	28	28	28	28	28	28

Source: authors' computation from the compiled dataset.

The employment-to-population ratio averaged 53.83 per cent over the period, ranging from a minimum of 49.87 per cent in 2005 to a maximum of 59.00 per cent in 2024, with a standard deviation of 2.54 percentage points. The comparatively narrow dispersion reflects the structural persistence of the Zambian labour market, in which subsistence and informal activity absorb much of the variation that would otherwise appear as fluctuation in measured employment.

The real minimum wage averaged 1236.29 kwacha per month in constant 2023 prices, but the range is striking. The minimum of 585.04 kwacha, recorded in 2005 after nine years without adjustment following the 1997 Instrument and three years of erosion following the 2002 Instrument, is less than one third of the maximum of 1884.81 kwacha recorded in 2018 immediately after the largest single revision in the sample. A standard deviation of 341.75-kwacha, equivalent to 28 per cent of the mean, confirms that the real wage floor faced by Zambian employers has varied far more than the sequence of six nominal revisions would suggest.

Economic growth averaged 5.11 per cent with a standard deviation of 2.78 percentage points, reaching 10.30 per cent in 2010 during the copper price boom and contracting by 2.79 per cent in 2020 during the pandemic. The negative skewness of minus 0.760 and the excess kurtosis of 4.535 reflect the influence of that single sharp contraction. Inflation averaged 14.49 per cent with a standard deviation of 6.81 percentage points, ranging from 6.43 per cent in 2011 to 26.79 per cent in 1999. The Jarque Bera statistics for all six series are individually insignificant at the 5 per cent level, so the null hypothesis of normality is not rejected for any variable.

4.2 Trend analysis

Figure 1 plots the employment-to-population ratio against the real minimum wage over the sample period, with vertical reference lines marking the six years in which the statutory minimum wage was revised.

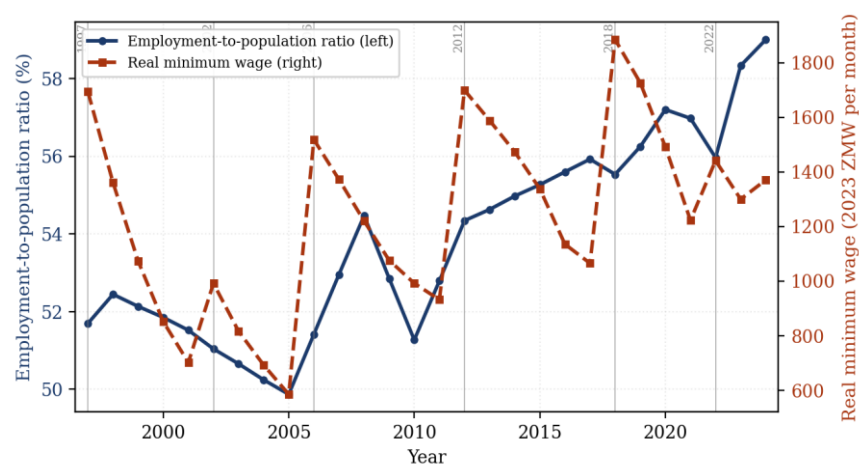


Figure 1: Employment to population ratio and the real minimum wage, 1997 to 2024

Source: authors' construction. Vertical lines mark years in which the statutory minimum wage was revised.

Three features of the series are apparent. The first is the sawtooth path of the real minimum wage. Each revision produces a sharp upward step, after which inflation erodes the real value continuously until the next Instrument is gazetted. Following the 1997 Instrument the real wage floor fell from

1693.75 kwacha to 701.58 kwacha by 2001, a decline of fifty-nine per cent. Following the 2006 Instrument it fell from 1518.40 kwacha to 931.91 kwacha by 2011, a decline of thirty-nine per cent. Following the 2012 Instrument it fell from 1699.37 kwacha to 1065.24 kwacha by 2017, a decline of 37 per cent. Following the 2018 Instrument it fell from 1884.81 kwacha to 1222.81 kwacha by 2021, a decline of 35 per cent. The statutory floor is therefore binding to a materially different degree at different points in the revision cycle, and any analysis using the nominal series alone would misrepresent the policy that employers actually face.

The second feature is the behaviour of the employment-to-population ratio. The series declined gradually from 51.70 per cent in 1997 to 49.87 per cent in 2005, recovered through the commodity boom years to reach 54.47 per cent in 2008, fell back to 51.27 per cent in 2010, and has trended upwards since, reaching 59.00 per cent in 2024. The upward movement from 2018 onwards, from 55.53 per cent to 59.00 per cent, coincides with the two most recent minimum wage revisions but also with a broader structural trend that predates them, a point to which the econometric analysis returns.

The third feature is the absence of any visible contemporaneous disemployment response. In none of the six revision years does the employment-to-population ratio fall sharply in the following period. The 2018 revision, the largest real increase in the sample at approximately seventy-seven per cent, was followed by increases in the employment ratio in each of the next two years. This is a descriptive observation rather than a causal one, since other factors were moving simultaneously, but it is inconsistent with a strong and immediate neoclassical adjustment.

Figure 2 makes the erosion mechanism explicit by plotting the nominal and real minimum wage series together.

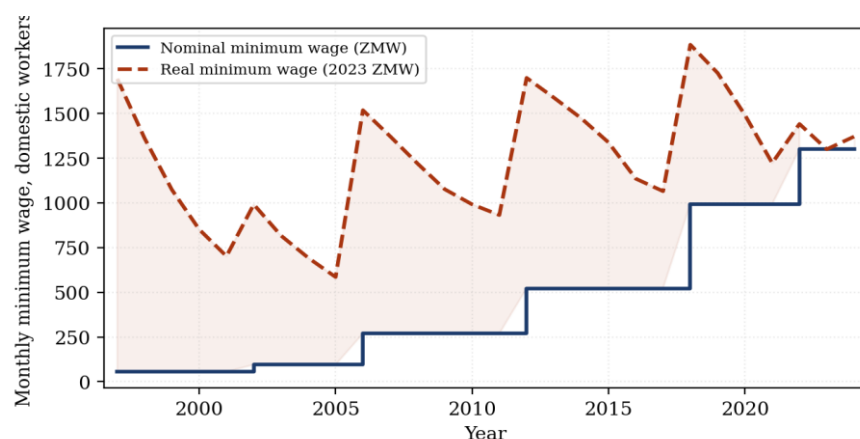


Figure 2: Nominal and real statutory minimum wage for domestic workers, 1997 to 2024

Source: authors' construction from Statutory Instruments, deflated using the consumer price index

The nominal series is a step function that rises monotonically from 55 kwacha per month in 1997 to 1300 kwacha per month in 2022 and 2024. The real series, by contrast, oscillates around a broadly flat mean and stands at 1370.72 kwacha in 2024, only marginally below its 1997 value of 1693.75 kwacha in constant terms and well below its 2018 peak. Over twenty-seven years, and after six nominal revisions that increased the headline figure more than twenty-three fold, the purchasing power of the statutory floor for domestic workers has not risen. This is the central policy fact about Zambian minimum wage legislation, and it substantially conditions the interpretation of the econometric results that follow.

4.3 Correlation and multicollinearity

Table 3 reports the pairwise correlation matrix for the four variables entering the model.

Table 3: Correlation matrix

	lnEPR	lnRMW	GDPG	INF
lnEPR	1.000			
lnRMW	0.604	1.000		
GDPG	-0.406	-0.317	1.000	
INF	-0.441	-0.431	-0.206	1.000

Source: authors' computation.

The employment-to-population ratio is positively correlated with the real minimum wage at 0.604, and negatively correlated with both economic growth at minus 0.406 and inflation at minus 0.441. The positive simple correlation with the real minimum wage is notable but must be interpreted with caution, since both series contain trending components. The negative correlation between employment and growth is at first sight counterintuitive; it reflects the fact that Zambia's highest growth years occurred during the capital-intensive copper boom of the mid 2000s, when the employment ratio was at its lowest, which is itself an illustration of the weak employment intensity of Zambian growth. No pairwise correlation among the regressors exceeds 0.44 in absolute value, and all variance inflation factors are below 1.6, well under the conventional threshold of five. Multicollinearity is therefore not a material concern.

4.4 Unit root tests

Table 4 reports the results of the stationarity tests. For the Augmented Dickey–Fuller and Phillips Perron tests the null hypothesis is the presence of a unit root, so that a small p-value indicates stationarity. For the Kwiatkowski–Phillips–Schmidt–Shin test the null hypothesis is stationarity, so that a statistic exceeding the critical value indicates non-stationarity.

Table 4: Unit root and stationarity test results

Variable	Form	ADF (c)	Prob.	ADF (c,t)	Prob.	PP (c)	Prob.	KPSS (c)
lnEPR	Level	0.114	0.967	-3.674	0.024	0.433	0.983	0.612**
	First difference	-5.136	0.000	-5.277	0.000	-4.231	0.001	0.237
lnRMW	Level	-2.893	0.046	-4.095	0.006	-3.145	0.023	0.353
	First difference	-4.755	0.000	-4.608	0.001	-6.920	0.000	0.179
GDPG	Level	-2.841	0.053	-2.898	0.163	-2.985	0.036	0.170
	First difference	-8.222	0.000	-8.241	0.000	-8.235	0.000	0.122
INF	Level	-2.091	0.248	-2.359	0.402	-1.877	0.343	0.426
	First difference	-5.978	0.000	-4.798	0.000	-6.653	0.000	0.186

Source: authors' computation

The results establish a mixed order of integration. The logged employment-to-population ratio is unambiguously non-stationary in levels: the Augmented Dickey–Fuller statistic of 0.114 and the Phillips Perron statistic of 0.433 fail to reject the unit root null by a wide margin, and the Kwiatkowski–Phillips–Schmidt–Shin statistic of 0.612 rejects the stationarity null at 5 per cent. All three tests agree that the first difference is stationary, so the series is I(1). The apparent rejection in the trend-augmented Augmented Dickey–Fuller specification, with a statistic of minus 3.674 and a p-value of 0.024, indicates that the series is closer to trend stationarity once a deterministic trend is allowed, which is why a trend term is retained in the preferred specification.

The logged real minimum wage presents a genuinely borderline case. The Augmented Dickey–Fuller statistic of minus 2.893 rejects the unit root null at the 5 per cent level, as do the trend-augmented statistic of minus 4.095 and the Phillips Perron statistic of minus 3.145, while the Kwiatkowski–Phillips–Schmidt–Shin statistic of 0.353 fails to reject stationarity. On this evidence the series is best classified as I(0) or as being on the boundary between I(0) and I(1). Economic growth is stationary in levels on the Phillips Perron and marginally on the Augmented Dickey–Fuller criterion, and is therefore treated as I(0). Inflation fails to reject the unit root null in levels under both tests but is clearly stationary in first differences, and is treated as I(1).

This combination—an I(1) dependent variable, one regressor at the I(0) boundary, one I(0) regressor, and one I(1) regressor—is compatible with the bounds-testing framework. No series is classified as I(2), so the procedure is admissible, subject to the finite-sample limitations already noted. Because the Johansen framework ordinarily assumes a common integration order, it is not used as the primary method; its results are reported only as a supplementary sensitivity check.

4.5 Bounds test for cointegration

Table 5 reports the bounds test results for the preferred specification and for the alternative deterministic specification without a trend.

Table 5: ARDL bounds test for a long-run relationship

Specification	F-statistic	10% I(0)	10% I(1)	5% I(0)	5% I(1)
Case V: unrestricted intercept and trend	3.792	3.47	4.45	4.01	5.07
Case III: unrestricted intercept, no trend	0.181	3.17	4.14	3.79	4.85
Error correction term, Case V	-0.951	t = -3.572		p = 0.003	

Source: authors' computation.

In the preferred trend-inclusive specification, the computed F-statistic of 3.792 exceeds the lower bound of 3.47 at the 10 per cent level but remains below the upper bound of 4.45; formally, the result is inconclusive. At the 5 per cent level, the statistic falls below the lower bound of 4.01, so the null of no long-run relationship cannot be rejected. Without a trend, the F-statistic is only 0.181, far below every relevant bound and therefore offering no evidence of cointegration. The implication is clear: taken on its own, the bounds test does not establish a cointegrating relationship between the real minimum wage and employment in Zambia at conventional significance levels.

Three considerations bear on this result. First, statistical power is limited: the preferred equation has twenty-six usable observations, thirteen estimated parameters, and only thirteen residual degrees of freedom. Second, the lagged level coefficient on employment is -0.951 ($t = -3.572$; $p = 0.003$), which is negative and precisely estimated. Banerjee, Dolado and Mestre (1998) show that the t-statistic on this coefficient can provide additional evidence on a level relationship, although the relevant non-standard critical values are more demanding than those of the conventional t-distribution. The estimate is therefore supportive, not conclusive, evidence of error correction.

The third consideration is corroborating evidence. Table 6 reports Johansen trace and maximum-eigenvalue statistics, both unadjusted and adjusted for small-sample bias by the factor proposed by Reinsel and Ahn (1992).

Table 6: Johansen cointegration test (supplementary evidence)

System	Hypothesis	Trace statistic	Adjusted trace	5% critical value	Decision
Bivariate: lnEPR, lnRMW	$r = 0$	17.308	14.646	15.494	Marginal
	$r \leq 1$	0.080	0.067	3.841	Not rejected
Four variable system	$r = 0$	52.819	36.567	47.855	Marginal
	$r \leq 1$	27.050	18.727	29.796	Not rejected

Source: authors' computation.

On the unadjusted statistics the Johansen procedure identifies exactly one cointegrating vector in both the bivariate and the four-variable systems at the 5 per cent level. Once the small-sample correction is applied the trace statistics fall below their critical values in both systems, so this evidence too is marginal rather than decisive.

Taken together, the evidence for the second research objective indicates that evidence of a long-run relationship between minimum-wage legislation and employment is suggestive but not conclusive. The negative adjustment coefficient, the unadjusted Johansen result, and the cross-estimator stability reported below are supportive. Nevertheless, the preferred bounds test is inconclusive at 10 per cent and fails to reject the null at 5 per cent; the small-sample-adjusted Johansen statistics are likewise below their critical values.

4.6 Long-run estimates

Table 7 reports the long-run coefficients derived from the preferred conditional error correction model together with three alternative estimates.

Table 7: Long run coefficient estimates from four estimators

Variable	ARDL long run	DOLS (1,1)	Static + trend	Static HAC	Static + D2018
lnRMW	0.078*** (0.026)	0.123*** (0.038)	0.042*** (0.012)	0.049** (0.024)	0.027 (0.019)
GDPG	0.001 (0.003)	-0.009*** (0.003)	-0.002** (0.001)	-0.007*** (0.002)	-0.004 (0.002)
INF	0.003 (0.002)	-0.004** (0.002)	0.001*** (0.000)	-0.003** (0.001)	-0.002*** (0.001)
D2018					0.055*** (0.017)
Adjusted R squared	0.522	0.811	0.884	0.457	0.669
Observations	26	25	28	28	28

Source: authors' computation.

The central result concerns the sign and precision of the long-run coefficient on the real minimum wage. In the preferred autoregressive distributed lag specification, the estimate is 0.078 with a standard error of 0.026 ($t = 3.051$; $p = 0.009$). Dynamic ordinary least squares yields 0.123, the trend-augmented static regression 0.042, and the static regression with heteroskedasticity- and autocorrelation-consistent standard errors 0.049. Each is positive and statistically significant at conventional levels. Thus, across these estimators, the association lies between approximately 0.04 and 0.12. This cross-estimator consistency strengthens confidence in the sign of the association, but it does not by itself establish cointegration or causal identification.

Interpreted as an elasticity, the preferred estimate associates a 10 per cent increase in the real statutory minimum wage with a 0.78 per cent increase in the employment-to-population ratio in the long run specification. At the sample mean employment ratio of 53.83 per cent, this corresponds to approximately 0.42 percentage points. The magnitude is modest and, importantly, opposite in sign to the standard competitive-market prediction. The result therefore provides no evidence of a negative long-run employment association in the aggregate data; it should not be read as proof that a higher minimum wage causes employment to increase.

The single exception in Table 7 is the final column, in which the introduction of the 2018 policy indicator reduces the minimum wage coefficient to 0.027 and renders it insignificant, while the indicator itself is estimated at 0.055 and is significant at one per cent. The natural reading is that the policy indicator and the real minimum wage are competing to explain the same post-2018 upward movement in employment, and that the indicator, being a simple level shift, absorbs the variation more efficiently. This is a caution against over interpreting the continuous minimum wage coefficient, and Section 4.9 returns to it.

The control variables behave less consistently across specifications. Economic growth carries a negative long-run coefficient in three of the five specifications, which is contrary to the expected sign. This apparent anomaly reflects the composition of Zambian growth over the sample period. The years of most rapid expansion were driven by copper prices and by capital-intensive mining investment, which generate limited direct employment, while the years of slower growth coincided with expansion in agriculture and services, which are more labour-absorbing. The result is best read not as evidence that growth destroys jobs but as evidence that the employment intensity of Zambian growth is low and has been declining, which is itself a finding of some policy significance. Inflation carries a negative and significant coefficient in the static and dynamic ordinary least squares specifications, consistent with the expectation that macroeconomic instability depresses labour demand, though it is insignificant in the preferred specification.

4.7 Short run dynamics and error correction

Table 8 reports the estimated conditional error correction model in full.

Table 8: Conditional error correction model estimates

Regressor	Coefficient	Standard error	t-statistic	Probability
Panel A: short run dynamics				
Constant	3.151	0.957	3.292	0.006
$\Delta \ln \text{EPR}(-1)$	0.475	0.212	2.236	0.044
$\Delta \ln \text{RMW}$	0.026	0.018	1.484	0.162
$\Delta \ln \text{RMW}(-1)$	-0.018	0.016	-1.086	0.297
ΔGDPG	-0.003	0.001	-2.039	0.062
$\Delta \text{GDPG}(-1)$	-0.004	0.002	-1.628	0.128
ΔINF	0.000	0.001	-0.007	0.994
$\Delta \text{INF}(-1)$	-0.001	0.001	-1.589	0.136
Trend	0.005	0.001	3.816	0.002
Panel B: lagged level terms				
$\ln \text{EPR}(-1)$	-0.951	0.266	-3.572	0.003
$\ln \text{RMW}(-1)$	0.074	0.028	2.622	0.021
$\text{GDPG}(-1)$	0.001	0.003	0.245	0.811
$\text{INF}(-1)$	0.003	0.002	1.793	0.096

Source: authors' computation

Panel A speaks directly to the first research objective: the concurrent change in the real minimum wage carries a coefficient of 0.026 with a standard error of 0.018, giving a t-statistic of 1.484 and a p-value of 0.162. The coefficient is positive but statistically insignificant. Its one-period lag is negative at minus 0.018 and equally insignificant, with a p-value of 0.297. Taken together these results indicate that changes in the real statutory minimum wage have no visible short-run effect on aggregate employment in Zambia in either direction. The point estimate on the contemporaneous term rises to 0.036 and approaches significance with a p-value of 0.072 when the 2018 policy indicator is added to the specification, but the qualitative conclusion is unchanged.

The lagged dependent variable in first differences carries a coefficient of 0.475 with a p-value of 0.044, confirming appreciable inertia in the adjustment of employment. Employment growth in one year is positively associated with employment growth in the next, which is consistent with hiring and separation processes that unfold over more than a single year and with the structural persistence noted in the descriptive statistics.

The coefficient on the lagged level of the employment ratio in Panel B is -0.951 with a standard error of 0.266 ($t = -3.572$; $p = 0.003$). It is negative, less than unity in absolute value, and highly significant. Conditional on the existence of a stable level relationship, its magnitude would imply that approximately 95 per cent of a deviation is corrected within one year. For an annual model, this is a rapid adjustment rate and is more appropriately understood as adjustment within roughly one to two years. Because the bounds test is inconclusive, the coefficient is treated as corroborating evidence rather than as independent proof of cointegration.

The deterministic trend carries a coefficient of 0.005 with a p-value of 0.002, implying an underlying upward drift in the employment-to-population ratio of approximately half a percentage point per year that is independent of the minimum wage, of growth and of inflation. This captures demographic and structural forces, including the expansion of the working-age population, rising female labour force participation and the continued growth of informal self-employment, which operate on employment independently of the variables modelled here.

4.8 Diagnostic and stability tests

Table 9 reports the residual diagnostics for the preferred model.

Table 9: Diagnostic test results

Test	Null hypothesis	Statistic	Probability
Breusch Godfrey LM (1 lag)	No serial correlation	$F = 0.557$	0.470
Breusch Godfrey LM (2 lags)	No serial correlation	$F = 0.428$	0.662
Breusch Pagan Godfrey	Homoskedasticity	$F = 1.650$	0.191
Harvey	Homoskedasticity	$\chi^2 = 19.277$	0.082
ARCH LM (1 lag)	No ARCH effects	$LM = 0.237$	0.627
Jarque Bera	Normality of residuals	$JB = 0.746$	0.689
Ramsey RESET	Correct functional form	$F = 0.022$	0.885
Durbin Watson	No first order autocorrelation	$d = 2.173$	Inconclusive region not entered

Source: authors' computation.

None of the reported diagnostic tests rejects its null at the 5 per cent level. The Breusch–Godfrey Lagrange multiplier tests at one and two lags ($p = 0.470$ and $p = 0.662$) provide no evidence of serial correlation, consistent with the Durbin–Watson statistic of 2.173. The Breusch–Pagan–Godfrey and Harvey tests ($p = 0.191$ and $p = 0.082$) do not reject homoskedasticity, while the autoregressive conditional heteroskedasticity test ($p = 0.627$) finds no evidence of volatility clustering. The Jarque–Bera statistic of 0.746 ($p = 0.689$) does not reject residual normality, and the Ramsey specification-error test ($p = 0.885$) provides no evidence of functional-form misspecification. These diagnostics are reassuring, although their power is necessarily limited in a sample of this size.

Importantly, this clean diagnostic profile was not obtained automatically. Specifications estimated without a deterministic trend exhibited significant first-order serial correlation, with Breusch–Godfrey p -values below 0.01, which would have invalidated inference drawn from them. The selection of the preferred specification was therefore made jointly on the Akaike information criterion and on the requirement that all residual diagnostics be satisfied, rather than on the information criterion alone.

Figure 3 presents the cumulative sum and cumulative sum of squares of recursive residuals with their 5 per cent significance bands.

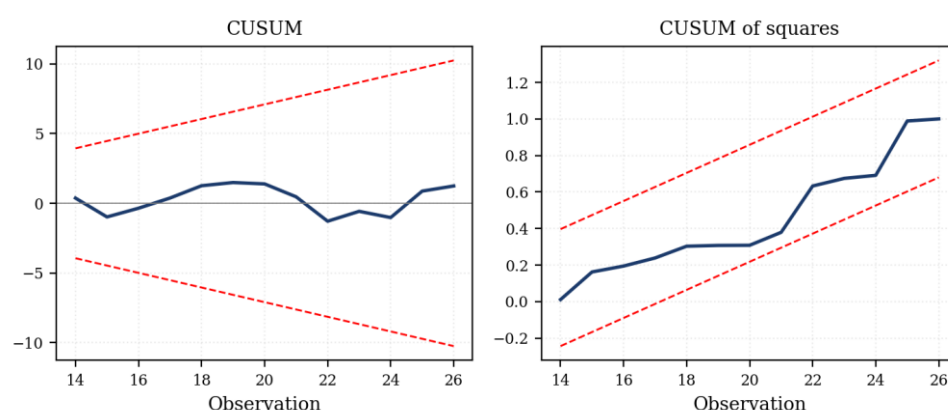


Figure 3: Parameter stability tests

Source: authors' computation. Dashed lines denote 5 per cent critical bounds.

Both cumulative statistics remain within their 5 per cent critical bounds throughout the sample. The recursive-residual diagnostics therefore provide no evidence of parameter instability over 1997–2024. This is reassuring given that the period spans the copper boom, the global financial crisis, the 2018 minimum-wage revision, and the 2020 pandemic contraction. It does not, however, exclude breaks that these low-power tests may fail to detect.

4.9 Granger causality

Table 10 reports the results of the causality tests, which address the third research objective.

Table 10: Granger causality and Toda Yamamoto test results

Test procedure	Lag order	RMW to EPR	Probability	EPR to RMW	Probability
Panel A: pairwise Granger causality, first differences					
Pairwise F test	1	2.606	0.120	0.076	0.785
Pairwise F test	2	4.644	0.022	1.103	0.351
Pairwise F test	3	2.831	0.069	1.076	0.386
Panel B: block exogeneity in the four variable VAR, first differences					
VAR block F test	1	1.311	0.255	0.611	0.437
VAR block F test	2	2.250	0.114	0.157	0.855
Panel C: Toda–Yamamoto augmented VAR in levels					
Modified Wald χ^2	1	1.409	0.235	5.230	0.022
Modified Wald χ^2	2	3.855	0.146	5.313	0.070
Modified Wald χ^2	3	8.347	0.039	3.623	0.305

Source: authors' computation.

The results do not support a robust causal ordering in either direction. In Panel A the pairwise test at two lags produces an F-statistic of 4.644 with a p -value of 0.022, which would reject the null that the real minimum wage does not Granger cause employment at the 5 per cent level. That result, however, is not sustained. At one lag the p -value is 0.120 and at three lags it is 0.069, so the finding is specific to a single lag order. More importantly, when the two control variables are introduced into a four-variable vector autoregression in Panel B, the evidence disappears entirely: p -values of 0.255 at one lag and 0.114 at two lags provide no basis for rejecting the null. The bivariate result therefore appears to reflect omitted variable bias, with the real minimum wage proxying for macroeconomic conditions that independently influence employment.

Panel C reinforces this reading. The Toda–Yamamoto procedure, which is robust to the order of integration and to the cointegration status of the

system, returns p-values of 0.235, 0.146 and 0.039 for the hypothesis that the minimum wage does not cause employment at augmentation orders one, two and three respectively. Only the third of these is significant at 5 per cent, and a result that appears at one lag order out of three, in the specification with the fewest degrees of freedom, cannot be regarded as robust evidence.

Predictive causality from employment to the minimum wage is equally fragile. The Toda–Yamamoto statistic is significant at one lag ($p = 0.022$) and marginal at two lags ($p = 0.070$), whereas the pairwise and vector autoregression tests uniformly fail to reject, with p-values between 0.351 and 0.855. Taken together, the results do not establish systematic policy responsiveness to labour-market conditions. They are consistent with—but do not by themselves prove—the institutional record of six revisions at irregular intervals of four to fifteen years. Minimum-wage policy in Zambia appears episodic rather than systematically countercyclical.

For the third research objective, the evidence therefore indicates that no robust Granger causal relationship can be established between changes in minimum wage levels and employment outcomes in Zambia over the period examined.

4.10 Discussion

The results present a coherent, though necessarily qualified, account. Changes in the real statutory minimum wage have no measurable short-run association with aggregate employment. The estimated long-run association is positive and stable across several estimators, yet the formal evidence needed to describe that relationship as cointegration remains inconclusive. Neither does the analysis establish robust predictive precedence in either direction. The central interpretive task, therefore, is to explain why the data reveal no negative aggregate employment response—not to claim that they prove a positive causal effect.

The most persuasive explanation rests on two features of the Zambian labour market. The first is the limited reach of the statutory floor. Informal employment accounts for the majority of Zambian workers, and enforcement capacity in the formal sector is constrained. A wage floor that binds for a minority of employment relationships, and that is imperfectly enforced even among those, cannot be expected to generate large aggregate effects of either sign. Bhorat and Kanbur (2019) make precisely this point in their finding that noncompliance rises with the Kaitz index across African economies, and the International Labour Organization (2022) documents the compliance difficulties facing small Zambian businesses directly. The second feature is the erosion documented in Figure 2. Because the real value of the wage floor declines by a third or more between revisions, the average degree to which it binds over the cycle is considerably lower than the headline nominal figure suggests. In effect, Zambian minimum wage policy delivers a periodic real increase that is then rescinded by inflation, and a policy that does not sustain a binding constraint cannot generate a sustained employment effect.

The positive long-run coefficient admits three interpretations. First, under monopsony, employers of low-skilled labour may possess wage-setting power because local labour markets are thin and worker mobility is constrained; a wage floor between the monopsony and competitive wages may then raise both wages and employment (Manning, 2003; Boal and Ransom, 1997). Second, efficiency-wage mechanisms may reduce turnover or increase effort sufficiently to offset part of the direct labour-cost increase (Shapiro and Stiglitz, 1984). Third, and most cautiously, the coefficient may partly capture a common trend: both the employment ratio and real minimum wage rose after 2017. Table 7 supports this concern because the post-2018 indicator absorbs the effect and renders the continuous minimum-wage coefficient insignificant. Alongside the non-robust causality results, this sensitivity means that the evidence is stronger against a large negative aggregate effect than it is in favour of a positive causal effect.

These findings are broadly consistent with the recent literature. They align with Obeng (2015) and Asuamah and Nyarkoh (2023) for Ghana, where no significant aggregate disemployment effect was found and where growth and sectoral composition were judged more important determinants of employment than wage regulation. They align with the meta analytic conclusion of Doucouliagos and Stanley (2019) that aggregate employment effects are small once publication selection is accounted for, and with the survey evidence of Dube (2019). They align with the finding of Bhorat and Kanbur (2019) that African minimum wage increases generally produce small or unmeasurable negative effects. They are consistent in direction, though not in magnitude, with the long-run positive elasticity of 0.17 reported by Kar (2018) for member countries of the Organisation for Economic Cooperation and Development.

They differ, however, from two strands of evidence. The first is the panel finding of Okey, Adjor and Ketor (2020) of a negative relationship across selected African countries. The second is the Zambian estimate of Lupiya (2018), that minimum wage laws reduced the employment of working-age adults by 0.7 percentage points. The discrepancy with the latter is instructive, since Lupiya examines general workers in Lusaka using a combination of qualitative and quantitative methods over a shorter period, whereas the present study estimates an aggregate national relationship over 28 years. Displacement within a particular urban sector is fully compatible with a null aggregate effect if displaced workers are absorbed elsewhere, most obviously in informal activity. The two findings should be read as complementary rather than contradictory, and together they suggest that the composition of employment may respond to wage regulation even where its level does not.

The growth coefficient also warrants caution. Its negative or negligible long-run estimate across specifications, combined with a strongly positive deterministic trend, is consistent with weak employment intensity in Zambia's capital-intensive growth episodes. It should not be interpreted as evidence that economic growth causes employment to fall: the coefficient may reflect sectoral composition, common trends, measurement, or omitted variables. Even so, it directs attention towards the capacity of structural transformation—not output growth alone—to generate sustained employment.

5 Conclusion and Policy Implications

This study examined the relationship between minimum-wage legislation and employment performance in Zambia using annual data for 1997–2024. The analysis employed an autoregressive distributed lag bounds-testing framework, supplemented by Johansen tests, dynamic ordinary least squares, pairwise Granger tests, vector autoregression block-exogeneity tests, and the Toda–Yamamoto procedure. Three conclusions emerge, corresponding to the study's three specific objectives.

On the first objective, changes in the real statutory minimum wage have no statistically significant short-run effect on the employment-to-population ratio in Zambia. The estimated contemporaneous coefficient is positive at 0.026 but carries a p-value of 0.162, and its lagged value is negative and equally insignificant. Employment in the short run is governed principally by its own inertia, with a lagged adjustment coefficient of 0.475 significant at 5 per cent.

On the second objective, evidence of a long-run relationship is suggestive but not conclusive. The bounds-test F-statistic of 3.792 lies within the inconclusive band at 10 per cent and below the lower bound at 5 per cent. Johansen statistics that identify one cointegrating vector before small-sample adjustment fall below their critical values after adjustment. Conversely, the coefficient on lagged employment is -0.951 and significant at 1 per cent, while the long-run minimum-wage elasticity is positive and significant across four estimators, ranging from 0.042 to 0.123. The preferred estimate of 0.078 associates a 10 per cent increase in the real minimum wage with an increase of approximately 0.42 percentage points in the employment ratio. This is a robust association within the reported specifications, not a causal estimate.

On the third objective, no robust causal ordering is established. Isolated rejections at particular lag orders do not persist across the pairwise, multivariate, and Toda–Yamamoto procedures. The weak evidence of causality from employment to the minimum wage is consistent with an episodic policy process, but it cannot by itself confirm the institutional mechanism.

Overall, the aggregate time-series evidence does not support the proposition that Zambia's statutory minimum wage reduced employment over 1997–2024. The results are consistent with monopsonistic wage setting and efficiency-wage mechanisms, but they are equally consistent with limited enforcement, extensive informality, and a wage floor whose real value is repeatedly eroded by inflation. The findings therefore do not justify treating job destruction as an established consequence of moderate adjustments. By the same token, they do not justify the stronger claim that higher minimum wages cause employment to rise, nor do they exclude displacement within particular sectors or groups.

Four policy implications follow.

- The revision cycle should be regularised, but cautiously. The statutory requirement of biennial review should be implemented consistently. Figure 2 shows that inflation reduces the real value of the floor by a third or more between revisions. Indexation to consumer prices between formal reviews could limit this erosion, provided that reviews also consider productivity, compliance, sectoral conditions, and distributional effects. The aggregate results do not reveal a detectable employment cost, but the study's sample and design cannot rule out such costs for particular groups or sectors.
- Enforcement and compliance mechanisms warrant greater attention than the level of the floor. A minimum wage that is not enforced protects no one, and the finding that changes in the statutory level produce no detectable aggregate employment response is consistent with a floor that binds for only a minority of employment relationships. Investment in labour inspection capacity, particularly in domestic work and small-scale services, would raise the effective coverage of existing legislation more reliably than further nominal increases.
- Employment policy should look beyond wage regulation. The estimated trend coefficient of 0.005 per year is both large relative to the estimated minimum wage effect and highly significant, while the long-run coefficient on economic growth is negative or negligible. Zambian employment outcomes are being driven by demographic and structural forces and are not responding strongly to output growth. Raising the employment intensity of growth, through support for labour-absorbing sectors such as agriculture, agro-processing, construction and services, is likely to yield larger employment gains than any feasible adjustment to the wage floor.
- Complementary measures should address distributional composition. The contrast between the null aggregate finding of this study and the sectoral displacement reported by Lupiya (2018) and Mufungo (2023) suggests that wage regulation may alter the composition of employment even where it does not alter its level. Targeted support for small firms during adjustment periods, phased implementation for the most affected sectors and active monitoring of youth employment would help to ensure that the distributional gains of a higher wage floor are not offset by displacement within particular segments of the labour market.

5.1 Limitations

Four limitations merit explicit recognition. First, the sample contains only 28 annual observations, which limits the power of the bounds and causality tests and makes the estimates sensitive to lag structure and deterministic terms. Second, the employment-to-population ratio is based on modelled International Labour Organization estimates; smoothing in those estimates may understate genuine year-to-year variation and mechanically increase persistence. Third, the domestic-worker schedule is an imperfect proxy for the economy-wide bite of minimum-wage regulation. A Kaitz index based on relevant earnings, or a statutory schedule weighted by sectoral employment shares and compliance, would measure bindingness more directly. Fourth, aggregate analysis cannot identify heterogeneous effects by sector, firm size, age, gender, formality, or hours worked. The estimates are therefore aggregate and associational and should not be generalised to every worker or establishment.

5.2 Directions for further research

Two priorities emerge for future research. First, establishment- or worker-level analysis by sector, firm size, age, gender, and formality would determine whether the composition, hours, or quality of employment respond even when the aggregate employment level does not.

Second, explicit measurement of policy bindingness using a Kaitz index, coverage and compliance indicators, or sector-weighted statutory schedules would permit stronger tests of whether the weak aggregate response reflects limited enforcement rather than an absence of behavioural adjustment. Higher-frequency data and credible quasi-experimental designs around future revisions would further improve causal identification.

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Author Contribution

M.M contributed to the conception, design of the study, data collection, analysis, interpretation of data and drafting of the manuscript. B.H. contributed to the conception, design of the study, interpretation of data and editing and reviewing of the manuscript. All authors approved the final manuscripts.

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