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Research article

Unemployment invariance hypothesis and labor supply: A test for 31 American countries

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Abstract: This study examined the empirical validity of the unemployment invariance hypothesis (UIH) using a sample of 31 American countries and annual data from 1991 to 2023. While previous literature often focused on single-country analyses, an existing study for Latin America covered only six countries and relied on short time spans, limiting its ability to capture full economic cycles. This paper expanded both the temporal and geographical scope, enabling more accurate cross-country comparisons. The results generally reject the UIH, with significant implications for economic policy in both cases—whether the hypothesis is accepted or not. In countries where UIH is rejected, the discouraged worker effect (DWE) tends to outweigh the added worker effect (AWE). These findings highlight the need for country-specific labor policies, which can be better designed based on the estimates presented.

Keywords: cointegration; added worker effect; discouraged worker effect; unemployment invariance; unemployment rate; labor force participation rate

JEL Codes: E24, C10, J64, J68

1. Introduction

One version of the so-called unemployment invariance hypothesis (UIH), originally proposed by Layard et al. (1991), asserts that the long-term unemployment rate is independent of observed changes in labor supply. Although this hypothesis has been debated by Layard et al. (1991), it has been highly relevant in the macroeconomic and labor economics literature, as it argues that conventional economic policies aimed at modifying the labor force have limited or no effects on long-term unemployment. On the other hand, if this hypothesis does not hold, it becomes necessary to delve into specific labor market phenomena, such as the added worker effect (Woytinsky, 1940) or the discouraged worker effect (Long, 1953, 1958), and to examine whether changes in the unemployment rate and the labor force participation rate are directly or inversely related.

Most empirical research on the UIH yields mixed results regarding its validation. This represents a limitation that has hindered the generalization of findings and restricted the applicability of the conclusions across diverse economic and social contexts. Moreover, the existing literature shows significant gaps, as many of the available studies rely on different datasets and time periods, which complicates cross-country comparisons and the generalization of economic policy measures. Addressing this gap is crucial for understanding how different institutional and socioeconomic contexts may influence the validity of the UIH and, consequently, determine which labor policies are most effective for each country.

The main objective of this study is to estimate whether the UIH holds in a broad and diverse sample of 31 countries across the Americas. To this end, we use a homogeneous database developed by the World Bank within a common time frame—from 1991 to 2023—which allows for a more legitimate comparison between the economies of the countries analyzed. In a first step, we classify the countries into two groups: those in which the UIH holds and those in which it does not—something that already has significant implications for economic policy. In a second step, for the countries where the UIH is not verified, we determine whether the added worker effect (AWE) or the discouraged worker effect (DWE) predominates, in order to provide guidance for policymakers in those countries.

Our contribution to the existing literature occurs along several dimensions. On the one hand, we simultaneously conduct a test of the UIH for 31 countries. To date, few studies have undertaken such a broad cross-country comparison. Moreover, the American continent has not been extensively studied. A notable exception is the analysis by Maridueña-Larrea & Martín-Román (2024a), who examined the hypothesis for Latin America using quarterly data from 2006 to 2019. One key contribution of our study is substantially extending the time horizon (1991–2023). This is important because, when studying long-term cyclical phenomena, it is necessary to rely on a long time series that includes enough years of both economic expansion and periods of recession or slowed growth. Additionally, we significantly expand the geographical coverage—since Maridueña-Larrea & Martín-Román (2024a) only analyzed six countries—by including both developed and developing economies, which represents a novelty in the literature, as no existing studies addressed the UIH using such a large and heterogeneous sample.

To address this objective, the study will employ the cointegration econometric methodology proposed by Johansen (1991), a technique widely used to identify stable long-term relationships in time series analysis. This methodology has been extensively applied to test the UIH as well as to discern the predominance of the AWE or DWE. It will allow us to obtain results on regional heterogeneity in labor market responses to changes in the economic cycle, providing specific guidance for public policies tailored to different national realities.

The findings of this research reveal the significant heterogeneity in labor market responses to changes in the unemployment rate, offering partial support for the UIH in the Americas. Notable differences are observed in both the validation of the hypothesis and the magnitude of the discouraged worker effect (DWE) and the added worker effect (AWE), indicating that labor market dynamics are strongly shaped by country-specific structural factors. Moreover, the results show that long-term effects—especially in the post-pandemic context—can substantially alter the relevance of the UIH, underscoring the need for methodological approaches that incorporate varying time horizons and data frequencies.

The paper is structured into seven sections, including this introduction. Section 2 reviews the existing literature on the topic. Section 3 outlines the theoretical framework underpinning the research objectives. Section 4 describes the methodology used to test the UIH and its related effects. Section 5 details the dataset, the countries selected for the analysis period, and the characteristics of the series used for modeling. Section 6 presents the results obtained, and finally, Section 7 provides the conclusions and policy recommendations derived from the findings.

2. Literature review

One version of the UIH, originally proposed by Layard et al. (1991), argues that, in the long run, fluctuations in the labor force are independent of changes in the unemployment rate, implying the joint absence of both AWE and DWE effects. However, various studies have shown that this hypothesis does not always hold empirically, giving rise to an extensive debate about the specific conditions under which these effects emerge (Karanassou & Snower, 2004; Phelps, 1994; Rowthorn, 1999).

On the other hand, the unemployment rate is one of the most commonly used indicators to assess labor market conditions (Nemore et al., 2021). However, its ability to fully reflect labor market dynamics may be limited when significant changes in the labor force participation rate occur simultaneously, giving rise to the AWE and DWE phenomena[1].

The AWE, initially proposed by Woytinsky (1940), describes the increase in labor force participation when household members enter the labor market in response to the job loss of the primary earner. In contrast, the DWE, identified by Long (1953, 1958), reflects the reduction in the labor force participation rate when unfavorable labor market conditions lead potential workers to stop actively seeking employment during recessionary periods.

Maridueña-Larrea & Martín-Román (2024a) argue that the empirical validation of the UIH is crucial from an economic policy perspective. If the hypothesis is confirmed (i.e., there is no long-run relationship between the two variables), interventions aimed at modifying labor supply would have no effect on long-term unemployment. Conversely, the presence of a long-term relationship between unemployment and labor force participation would prompt an investigation into whether the AWE or DWE predominates, providing valuable insights for the design of effective and well-targeted policies.

[1] For a recent and formal view of these phenomena, see Martín-Román (2022).

The theoretical proposal of the UIH has sparked intense academic debate, driving numerous empirical studies that assess its validity across different economic and geographical contexts. In this regard, the literature reviewed clearly reveals a wide range of findings that highlight the complexity of the phenomenon. Specifically, there is a clear trend toward the widespread rejection of the UIH, implying that in most countries analyzed, there is a significant relationship between the activity rate and the unemployment rate, leading to the presence of either the AWE or the DWE.

In Table 1, we identified 11 studies that have globally tested the UIH for advanced economies since at least 2004[2]. The hypothesis was accepted in only two cases, while it was rejected in the remaining nine. Of these nine cases, six proceeded to analyze complementary effects on labor force participation dynamics, revealing a clear predominance of the DWE (in four out of six studies).

Support for the DWE is found in Österholm (2010) for Sweden, Kakinaka & Miyamoto (2012) for Japan[3], Tansel & Ozdemir (2018) for Canada, and Paternesi Meloni (2024) for OECD countries, employing a panel of time series data. On the other hand, evidence leaning toward the AWE is presented by Emerson (2011) for the United States, as well as by Nemore et al. (2021) for Italy. In the United Kingdom (Karanassou & Snower, 2004) and South Korea (Yilanci & Ozgur, 2024), although the UIH was also rejected, no specification was made regarding the presence of either the DWE or AWE at the aggregate level.

When data are disaggregated by gender in these economies, the literature reveals that the rejection of the UIH presents clearly differentiated patterns by sex. For example, studies conducted in Sweden, Japan, Canada, and various OECD countries show a consistent rejection of the UIH among men, with a marked predominance of the DWE, whereas for women, the results exhibit a more balanced distribution between the AWE and the DWE.

On the other hand, Table 1 also includes empirical evidence from emerging and developing countries, compiling seven studies that have examined the validity of the UIH. Although the number of studies identified for these economies is lower than that for advanced economies (11 studies in total), it is worth noting that the number of individual countries assessed in emerging and developing contexts slightly exceeds that of advanced economies—11 countries versus 10, respectively—when considering only those studies that focus on a single country.

Unlike the clear bias toward rejection of the UIH in developed economies, results in emerging and developing countries are more balanced, with five cases accepting and six rejecting the hypothesis based on aggregated data. It is worth noting that empirical analysis of the UIH in these countries began approximately 12 years later than in advanced economies, with 2016 marking the starting point according to the reviewed literature. Among the studies that supported the UIH are Oţoiu & Ţiţan (2016) in Romania, Tansel et al. (2016) in Turkey, Cheratian et al. (2022) in Iran (only for the female sample), and Khan et al. (2024) in Pakistan. In the Latin American context, Maridueña-Larrea & Martín-Román (2024a) also confirm the validity of the hypothesis for countries such as Brazil and Mexico.

[2] The International Monetary Fund (IMF) classifies economies into two main groups in its World Economic Outlook (WEO) report: (1) advanced economies and (2) emerging and developing economies. For further reference, see https://www.imf.org/en/Publications/WEO/weo-database/2025/april/groups-and-aggregates.

[3] Liu (2014) also rejected UIH in this country, although the author did not specify which effect predominated.

Among the six cases that rejected the UIH in emerging and developing economies, there is a clear tendency toward the presence of the DWE. Individually, Congregado et al. (2021) did so for Poland, as did Maridueña-Larrea & Martín-Román (2024a) for Chile and Uruguay. Using a panel of 52 African countries, Raifu & Adeboje (2022) provided evidence for five regions across the continent. This result is consistent with previously reported evidence for advanced economies, where the DWE also predominates. However, when analyzing data disaggregated by gender, it is notable that among men, the UIH is mostly supported (with Poland, the African countries, and Chile being the exceptions). In contrast, for women, there was no clear consensus regarding the validity of the UIH, as the hypothesis was rejected in six cases but supported in the remaining six.

The reviewed literature offers a comprehensive and nuanced perspective on the UIH. In advanced economies, studies rejecting the hypothesis clearly predominate, revealing labor dynamics that are sensitive to economic fluctuations and a significant presence of specific effects such as the DWE or AWE, depending on the particular context and gender differentiation. In contrast, in emerging and developing economies, the results are more heterogeneous, with no clear trend regarding the overall acceptance or rejection of the UIH. Although at an aggregate level there is some similarity with advanced economies in terms of the predominance of the DWE, a gender-disaggregated analysis reveals distinct patterns, especially toward the acceptance of the UIH among men, but with mixed results for women. This complexity highlights the importance of each country's specific institutional and economic context, underscoring the need for deeper and more differentiated analyses to design effective labor policies, as suggested by Hlasny (2023a, 2023b) and Maridueña-Larrea & Martín-Román (2025).

From a methodological standpoint, the reviewed studies have applied various techniques, mainly conditioned by data availability and the time horizon analyzed, with monthly or quarterly frequencies being the most common. A recurring methodological approach used to capture long-term dynamics and assess the existence of equilibrium relationships between unemployment and labor force participation is the cointegration technique proposed by Johansen (1991). Our research also follows this approach. In this way, our findings are largely comparable with those of the studies reviewed in Table 1.

Special mention should be made of the work by Maridueña-Larrea & Martín-Román (2024a). In that article, the UIH was studied for a small group of Latin American countries, and as such, it can be considered a clear precursor to the present research. However, the current article overcomes two clear limitations of the previously cited study. First, it significantly expands the spatial scope by considering 31 economies—with very different levels of development—instead of the 6 analyzed by Maridueña-Larrea & Martín-Román (2024a). This is particularly important for a study that aims to be positioned within the field of comparative economics. Second, and even more importantly, the temporal scope is substantially extended[4]. This is especially relevant when studying long-term phenomena such as the UIH. Furthermore, when examining cyclical phenomena like the AWE and DWE, it is necessary to have data covering complete economic cycles, and time horizons spanning several decades are highly recommended. Additionally, this study also includes some years following the COVID-19 crisis, thus providing very recent evidence.

[4] In the article by Maridueña-Larrea & Martín-Román (2024a), the country with the longest time span is Peru, with 14 years (from 2006 to 2019), while the countries with the shortest time span are Brazil and Chile, with only 8 years (from 2012 to 2019). In this paper, time series covering 33 years are used.

Table 1. Evidence from the literature on the UIH, AWE, and DWE.

Authors	Country	Period	UIH validity			Econometric approach
			Aggregate data	Male data	Female data	
Advanced economies						
Karanassou & Snower (2004)	United Kingdom	1964–1997	Rejected	n.a.	n.a.	ARDL cointegration
Österholm (2010)	Sweden	1970M1–2007M4	Rejected: DWE	Rejected: DWE	Rejected: DWE	Johansen cointegration VAR
Emerson (2011)	United States	1948M1–2010M2	Rejected: AWE	Rejected: DWE	Rejected: AWE	Johansen cointegration VAR
Kakinaka & Miyamoto (2012)	Japan	1980M1–2010M10	Rejected: DWE	Rejected: DWE	Accepted	Johansen cointegration VAR
Liu (2014)	Japan	1983Q1–2010Q4	Rejected	n.a.	n.a.	Johansen VAR-panel cointegration
Nguyen Van (2016)	Australia	1978M2–2014M12	Accepted	Accepted	Accepted	Johansen cointegration VAR
Tansel & Ozdemir (2018)	Canada	1976M1–2015M12	Rejected: DWE	Rejected: AWE	Rejected: DWE	Johansen cointegration VAR
Altuzarra et al. (2019)	Spain	1978Q2–2016Q4	Accepted	Accepted	Rejected: DWE	Johansen cointegration VAR
Nemore et al. (2021)	Italy	1998M1–2019M7	Rejected: AWE	Rejected: AWE	Rejected: AWE	Johansen VAR-Markov switching
Yilanci & Ozgur (2024)	South Korea	1999M6–2023M1	Rejected	Rejected: DWE	Accepted	Johansen cointegration (Fourier approach)
Paternesi Meloni (2024)	OECD countries	1960–2019	Rejected: DWE	n.a.	n.a.	Panel SVAR
Emerging and developing economies						
Oţoiu & Ţiţan (2016)	Romania	1996Q1–2012Q4	Accepted	Accepted	Accepted	Johansen cointegration VAR and first difference OLS
Tansel et al. (2016)	Turkey	1988Q3–2013Q4	Accepted	Accepted	Accepted	Johansen cointegration VAR
Congregado et al. (2021)	Poland	1995Q1–2016Q4	Rejected: DWE	Rejected: DWE	Rejected: DWE	Johansen VAR with structural breaks
Raifu & Adeboje (2022)	African countries	1991–2018	Rejected: DWE	Rejected: DWE	Rejected: DWE	Panel cointegration and DOLS
Cheratian et al. (2022)	Iran	2005Q2–2019Q1	n.a.	n.a.	Accepted	Time-series and panel cointegration

Continued on next page

Authors	Country	Period	UIH validity			Econometric approach
			Aggregate data	Male data	Female data	
Maridueña-Larrea & Martín-Román (2024a)	Brazil	2012Q1–2019Q4	Accepted	Accepted	Rejected: AWE	Johansen cointegration VAR
	Chile	2012Q2–2019Q4	Rejected: DWE	Rejected: DWE	Rejected: AWE	
	Ecuador	2010Q4–2019Q4	Rejected: AWE	Accepted	Accepted	
	Mexico	2011Q4–2019Q4	Accepted	n.a.	Accepted	
	Peru	2006Q3–2019Q1	Rejected: AWE	Accepted	Rejected: DWE	
	Uruguay	2008Q1–2019Q4	Rejected: DWE	Accepted	Rejected: AWE	
Khan et al. (2024)	Pakistan	1990–2021	Accepted	Accepted	Accepted	Johansen–Gregory–Hansen cointegration

Notes: AWE = added worker effect; DWE = discouraged worker effect; n.a. = not applicable.

3. Theoretical framework

To illustrate the UIH, we follow the so-called basic Layard–Nickell–Jackman (LNJ) model (Layard et al., 1991), as adapted by Karanassou & Snower (2004) and further modified here to explicitly address the research question of this article. Within this theoretical framework, equilibrium in the labor market results from what has been termed the "battle of the markups".

First, the price-setting (PS) function states that the mark-up of prices (P_t) over expected wages (W_t^e) depends negatively on the unemployment rate (u_t), as shown in Equation (1a)—or alternatively Equation (1b)—where α_0 and α_1 are positive coefficients. According to the standard theoretical setting, all variables (except the unemployment rate) are expressed in logarithmic form:

$$\boldsymbol{PS}: P_t - W_t^e = \alpha_0 - \alpha_1 u_t, \tag{1a}$$

$$\boldsymbol{PS}: W_t^e - P_t = \alpha_1 u_t - \alpha_0, \tag{1b}$$

The economic rationale behind the negative relationship between this markup and the unemployment rate is straightforward. It mirrors the reasoning behind the downward-sloping labor demand curve in a neoclassical framework: diminishing returns to labor.

On the other hand, the wage-setting (WS) function reflects, for instance, collective bargaining agreements, and establishes that the markup of (log) wages (W_t) over expected (log) prices (P_t^e) depends negatively on the unemployment rate, as shown in Equation (2). Again, β_0 and β_1 are positive coefficients.

$$\boldsymbol{WS}: W_t - P_t^e = \beta_0 - \beta_1 u_t, \tag{2}$$

The rationale underlying Equation (2) is quite well established today. It appears in textbooks such as Blanchard (2009), as well as in the macroeconomic literature on collective bargaining (e.g., Cabo & Martín-Román, 2019). When the unemployment rate rises, the bargaining power of unions weakens relative to that of employers.

Finally, in the long run, expectations align with actual wages and prices, so that $W_t^e = W_t$ and $P_t^e = P_t$. Solving the system formed by Equations (1b) and (2), the long-run equilibrium unemployment rate (u_t^*) is obtained, as shown in Equation (3):

$$u_t^* = \frac{\alpha_0 + \beta_0}{\alpha_1 + \beta_1}, \tag{3}$$

Now, let us build on the framework developed by Karanassou & Snower (2004), which explicitly incorporates the labor force into the analysis.[5] Accordingly, Equation (1b) becomes Equation (4), and Equation (2) becomes Equation (5):

$$\boldsymbol{PS'}: W_t - P_t = \alpha_1 u_t - \alpha_0 - \alpha_2 L_t, \tag{4}$$

[5] For the sake of simplicity, we do not consider the capital stock here, as it is not relevant to our empirical analysis.

$$\boldsymbol{WS'}: W_t - P_t = \beta_0 - \beta_1 u_t - \beta_2 L_t \quad (5)$$

It is worth noting that Equations (4) and (5) describe how the aforementioned markups are affected by changes in the labor force, given a certain unemployment rate[6]. As before, α_2 and β_2 are positive coefficients.

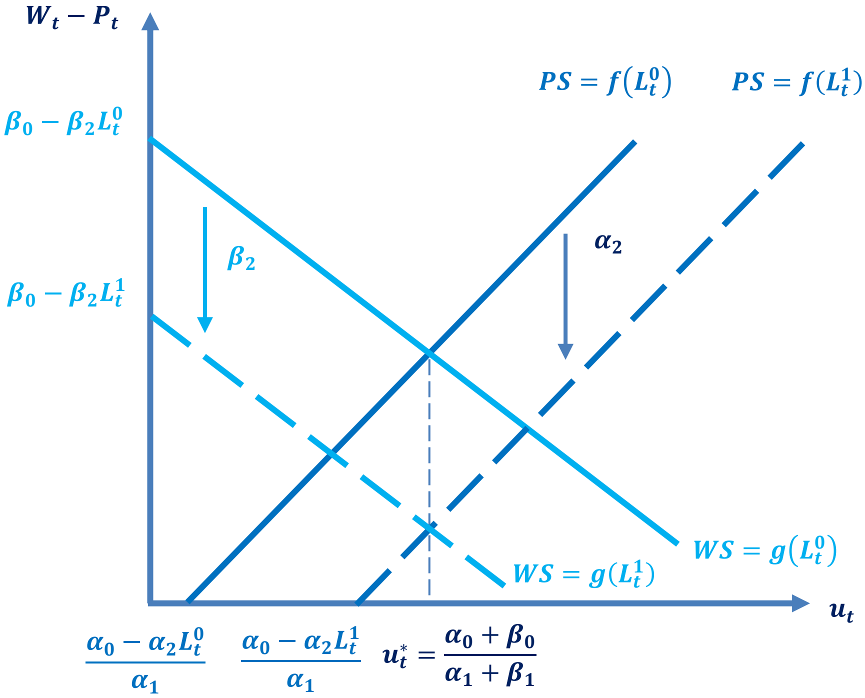

Figure 1. Unemployment invariance hypothesis. Source: Own elaboration.

Figure 1 illustrates the labor market equilibrium associated with this theoretical framework. The vertical axis displays the difference between wages and prices, while the unemployment rate is shown on the horizontal axis. This graphical device provides an alternative to representing equilibrium in the real wage–employment space[7].

Let us consider, for instance, an increase in the labor force (i.e., $L_t^1 > L_t^0$ in Figure 1)[8]. The negative effect on the wage–price differential in Equation (4) leads to a downward shift of the PS curve.

[6] It is important to emphasize that Equations (4) and (5) model the effect of the labor force on the markup of (log) wages over (log) prices, given a certain unemployment rate level. This approach is justified because the unemployment rate can be approximated by the difference between the log of the labor force and the log of employment: $u_t \approx L_t - E_t$. Consequently, the labor force is implicitly included in the definition of the unemployment rate. The graphical implications of this assumption are illustrated in Figure 1.

[7] In any case, it is also widely used. See Blanchard (2009) for a textbook explanation of the correspondence between the two graphical representations.

[8] Assuming a constant working-age population, it is straightforward to translate changes in the labor force into variations in the labor force participation rate.

The economic rationale is that when the labor force expands, given a constant unemployment rate, firms' search costs decrease, which in turn pushes the PS curve downward. On the other hand, the rise in the labor force also shifts the WS curve downward. As established in Equation (5), given a constant unemployment rate, a larger pool of job seekers weakens the relative bargaining power of unions, thus causing the shift.

Solving the system composed of Equations (4) and (5) yields the equilibrium condition for this extended labor market model. Accordingly, the long-run unemployment rate is given by Equation (6):

$$u_t^* = \frac{\alpha_0 + \beta_0}{\alpha_1 + \beta_1} + \frac{\alpha_2 - \beta_2}{\alpha_1 + \beta_1} L_t, \tag{6}$$

From Equation (6), it follows that if $\alpha_2 = \beta_2$, the UIH holds. In other words, when this condition is satisfied, changes in the labor force do not affect the long-run equilibrium unemployment rate. Graphically, the UIH is represented in Figure 1 as an equivalent downward shift of both the PS and WS curves. Whether this condition is met is an empirical question, which the remainder of the paper aims to explore in detail.

4. Methodology

This research employs a time series econometric approach based on the cointegration analysis proposed by Johansen (1991) to assess the existence of long-run equilibrium relationships between the labor force participation rate and the unemployment rate across each of the 31 countries in the Americas. The application of this methodology enables an empirical verification of whether the UIH holds in these countries, aligning with prior studies such as those by Österholm (2010), Nguyen Van (2016), and Maridueña-Larrea & Martín-Román (2024a).

To conduct the analysis, a vector autoregressive model (VARM) is specified, originally introduced by Sims (1980) and widely used in time series analysis (Zivot & Wang, 2003). First, the appropriate lag order is determined using information criteria and a significance test. Once the optimal specification is established, Johansen's trace and maximum eigenvalue cointegration tests are applied to assess both the presence and the number of long-run relationships. If cointegration is confirmed, the model is reformulated as a vector error correction model (VECM), which allows for a joint analysis of both the long-run equilibrium convergence and the short-run adjustment dynamics of the variables (Alizade, 2024; Fanchette et al., 2020).

The VECM can thus be considered a restricted version of the VARM, as it imposes that the endogenous variables maintain long-run equilibrium relationships (Mellander et al., 1992). In addition, the model explicitly captures the short-run adjustments that correct transitory deviations from equilibrium. As such, it is well-suited for the study of nonstationary series that evolve jointly through cointegration, providing a comprehensive framework for interpreting economic dynamics. Specifically, the VECM is specified as follows, in line with Altuzarra et al. (2019)[9]:

[9] An alternative formulation can be found in Lütkepohl & Kratzig (2004) or Hamilton (1994).

$$\Delta y_t = \alpha\beta' y_{t-1} + \sum_{t=1}^{p-1} \Gamma_i \Delta y_{t-i} + \varepsilon_t, \tag{7}$$

where y_t is a (K×1) vector of variables integrated of order one, 1 I(1). The matrices α and β are (K×r) parameter matrices with rank r<K. $\Gamma_1, \ldots, \Gamma_{P-1}$ are (K×K) coefficient matrices, and ε_t is a (K×1) vector of innovations that are independently and identically distributed. If the unemployment rate and the labor force participation rate in each country exhibit a long-run equilibrium relationship—that is, they are cointegrated—the matrix Π will have a rank equal to one. This implies the existence of two (2×1) vectors, α and β, each of rank one, such that $\Pi = \alpha\beta'$ and $\beta' y_t$ is stationary. In this context, α reflects the short-run adjustment effects within the VECM, while β serves as the cointegration vector that captures the long-run equilibrium relationship between the variables. The sign and magnitude of the parameters in β will provide insights into the predominance of the DWE or the AWE across the American region.

The results of this process will be presented in sections 5.2 and 6, and will verify the integration order of the series; also, if all series are found to be I(1) in levels and I(0) in first differences, the stability of the model will be tested, followed by the application of Johansen's cointegration test to determine the cointegration rank r among the variables. If no cointegration relationship is found, the UIH will be accepted. However, if at least one cointegrating vector is identified, then the cointegration vector will be estimated to quantify the magnitude of the AWE or DWE in the 31 economies analyzed.

5. Data

5.1. Series to be used

The variables used in this study were obtained from the World Bank database. They are annual in frequency and cover the period from 1991 to 2023. The variables considered are detailed below:

$$\text{Labor force participation rate (LFP)} \equiv \frac{\text{Labor force}}{\text{Working age population}},$$

$$\text{Unemployment rate (UR)} \equiv \frac{\text{Unemployed population}}{\text{Labor force}},$$

The use of annual data offers two key methodological advantages. First, the time series in this study consists of 33 observations, which are sufficient to estimate cointegration relationships when the model incorporates error correction terms, as demonstrated by Johansen (2002). Second, it avoids the mandatory seasonal adjustment required for monthly or quarterly data—a process that, if performed using inappropriate criteria, may bias estimators, particularly in economies with strong seasonal components. By dispensing with this adjustment, the analysis gains econometric robustness and internal consistency, capturing long-run fluctuations between labor force participation and unemployment more cleanly.

Based on this information, Figures 2 and 3 illustrate the temporal evolution of the unemployment rate and the labor force participation rate for each of the 31 economies in the

Americas over the 1991–2023 period. Additionally, to enable a comparative analysis of the average level of each indicator, Figure 4 presents a georeferenced representation of the annual average of both variables across the entire study period.[10]

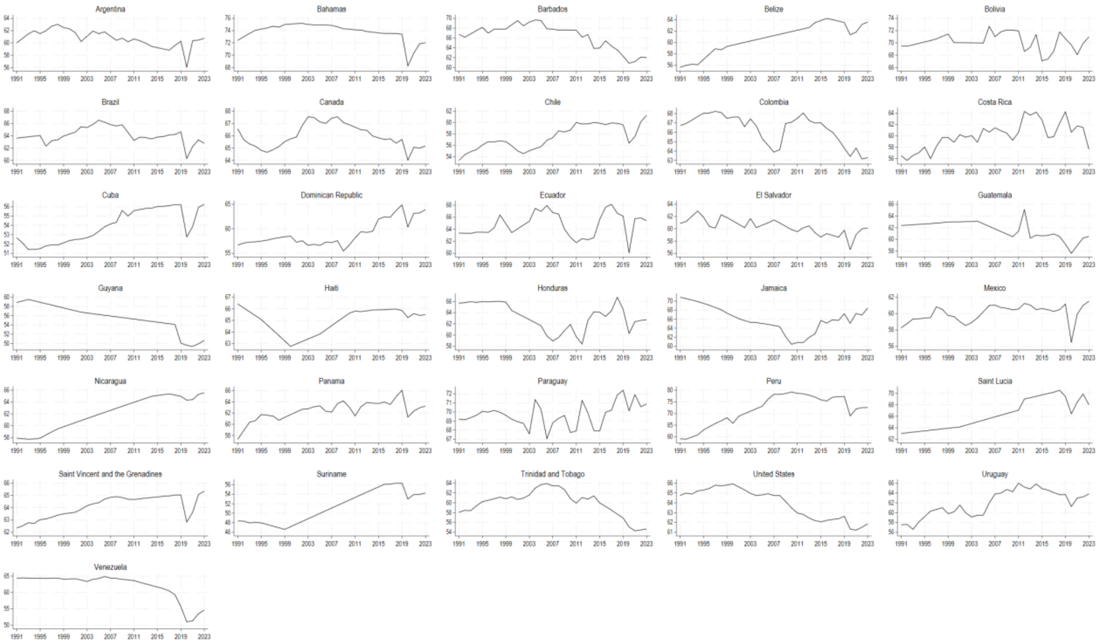

Figure 2. Labor force participation rate: selected countries (1991–2023). Source: World Bank. Own elaboration.

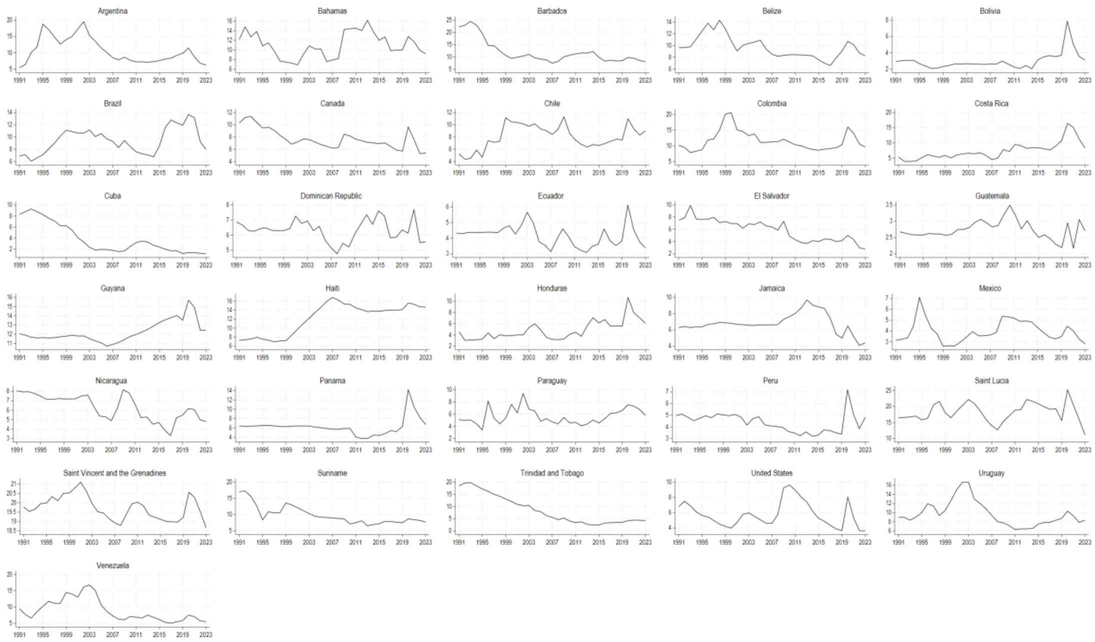

Figure 3. Unemployment rate: selected countries (1991–2023). Source: World Bank. Own elaboration.

[10] Coordinates for georeferencing were taken from the World Bank; for further reference, see https://datacatalog.worldbank.org/search/dataset/0038272/World-Bank-Official-Boundaries.

In the case of the unemployment rate, the lowest figures correspond to Guatemala (approximately 2.7%) and Bolivia (around 3%), while the highest values are observed in Saint Vincent and the Grenadines (close to 19.7%) and Saint Lucia (approximately 18.2%). Between these extremes, countries such as Colombia, Argentina, and Uruguay report rates above 8%, whereas others, such as Ecuador and El Salvador, fall within an intermediate range of 4% to 6%.

As for the labor force participation rate, the lowest average is found in Suriname (around 52%), while the Bahamas presents one of the highest values (close to 74%). Economies such as Peru and Bolivia also exceed 70%, whereas others, including Cuba and Chile, are situated around 54% and 58%, respectively. Most of the remaining countries report participation rates ranging between 60% and 66%.

The comparison of both indicators suggests heterogeneous labor market dynamics across the region. Some countries display relatively low unemployment rates combined with moderate or high participation levels, while others exhibit higher unemployment levels, with participation rates ranging from around 60% to over 70%. Overall, this information illustrates the diversity of labor market conditions across the Americas and provides a descriptive overview of how both variables are distributed across the 31 economies analyzed.

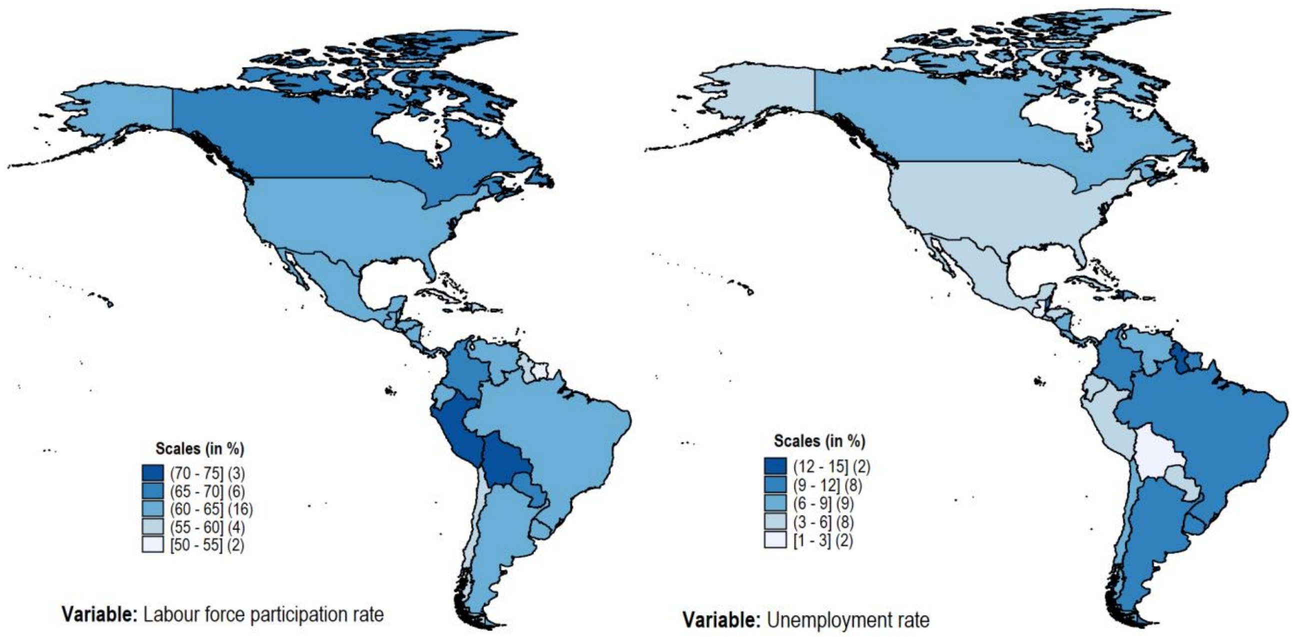

Figure 4. Labor force participation and unemployment rates in the Americas (average for the period 1991–2023). Source: World Bank. Own elaboration.

5.2. Properties of the series

Given that the time series variables used in this study may be affected by issues of spurious regression, it is necessary to examine their order of integration through unit root tests. To strengthen the robustness of the analysis, and considering potential structural shifts in economic series caused by the COVID-19 pandemic, a clearly differentiated two-stage procedure is proposed.

Although labor market rates may initially be assumed to be stationary—given that they are naturally bounded between zero and one—the literature indicates that this condition does not necessarily guarantee stationarity. In particular, two theoretical frameworks provide a solid conceptual foundation for considering that variables such as the unemployment rate—and, by extension, the labor force participation rate—may follow nonstationary processes over certain periods.

The first is the well-known hysteresis hypothesis (HH), put forward by Blanchard & Summers (1987, 1991) and Lindbeck & Snower (1988), which argues that unemployment may exhibit persistence due to rigidities in the labor market, generating unit root processes. This theory was originally anticipated by Phelps (1972) and later supported empirically by several studies in OECD countries (Caporale & Gil-Alana, 2014; Marques et al., 2017) and in Latin America (Ball et al., 2013), including specific cases such as Colombia (Iregui & Otero, 2003), Brazil and Chile (Gomes & da Silva, 2008), and México (Trejo García et al., 2017).

The second framework is the chain reaction theory (CRT), developed by Karanassou & Snower (1997, 1998), which suggests that prolonged adjustment processes in response to economic shocks may explain the apparent nonstationarity observed in certain variables, even when they are theoretically expected to be stationary (Martín-Román et al., 2023). Accordingly, in the first stage of the analysis—corresponding to the pre-pandemic period (1991–2019)—three conventional unit root tests are applied: augmented Dickey–Fuller (ADF), Phillips–Perron (PP), and Kwiatkowski–Phillips–Schmidt–Shin (KPSS). Subsequently, in the second stage, covering the full period (1991–2023), this set of tests is complemented with the structural break unit root test proposed by Zivot & Andrews (1992). The inclusion of this test responds to the methodological need to capture structural changes caused by exceptional events, such as the COVID-19 pandemic in 2020. Failure to account for such shifts may compromise the statistical power of conventional tests in the presence of highly persistent unemployment processes (Cuestas et al., 2011).

This two-stage comparative analysis allows for the evaluation of whether the integration properties identified previously remain valid once structural breaks are considered. It enables not only the detection of unit roots in economic series but also an assessment of the temporal stability of the results. Methodologically, the ADF, PP, and Zivot & Andrews (ZA) tests adopt the null hypothesis of a unit root, while the KPSS test assumes stationarity as the null. Based on the results of each test, it is determined whether the series are stationary in levels or require differencing to achieve stationarity. The detailed results are presented in Appendix A, Tables A.1 and A.2.

Following the application of the ADF, PP, and KPSS tests to the pre-pandemic period (1991–2019), it is observed that in most countries, both the labor force participation rate and the unemployment rate are integrated of order one [I(1)] in levels and stationary in first differences. Although full agreement among the three tests was not always reached, at least one of them supported the presence of a unit root in levels and stationarity in first differences in almost all economies. The main exception is El Salvador, where all three tests indicated that both rates were stationary in levels, which led to its exclusion from the cointegration analysis during the first stage. As a result, 30 of the 31 countries were retained for the cointegration analysis for the period 1991–2019.

In the second stage (1991–2023), which incorporates the years 2020 through 2023—marked by the COVID-19 pandemic—the ZA structural break test is included to control for potential abrupt changes in the series. Under this scheme, 21 of the original 31 countries maintained the I(1) condition in levels, confirming stationarity in first differences. Among the 10 excluded cases, El Salvador again exhibited stationarity in levels, even with the structural break test. Of the remaining nine countries, six (Barbados, Brazil, Costa Rica, Cuba, Nicaragua, and Saint Vincent and the Grenadines) showed a unit root in levels for the unemployment rate series; in Chile and Trinidad and Tobago, the same was true for the labor force participation rate. Jamaica displayed nonstationarity in both series—even in first differences—leading to its exclusion at this stage as well.

Consequently, a mixed cointegration analysis is proposed: one that provides results for the 30 countries validated during the 1991–2019 period, and another for the 21 countries that confirm an integration order of I(1) in levels and I(0) in first differences over the full period 1991–2023. This comparative approach allows for the examination of how results evolve for the same economy before and after the impact of the pandemic, while also avoiding the exclusion of demographically significant countries in the region, such as Brazil and Chile, whose series were suitable for cointegration analysis only during the pre-pandemic stage. In doing so, the study aims to deepen the understanding of the behavior of the unemployment rate and labor force participation rate in the Americas, offering a perspective that captures both structural stability prior to the pandemic and the effects that followed this global shock.

6. Results

6.1. Baseline results

Once the integration order of the series was determined for the two periods of analysis (pre-pandemic, 1991–2019, and the full cycle, 1991–2023), a VAR model was estimated to analyze the relationship between the unemployment rate and the labor force participation rate in each country.[11] A specific dummy variable for the year 2020 was included to capture the exogenous impact associated with the pandemic and to prevent this atypical event from distorting the estimated dynamics of the variables.

After confirming certain stability conditions, the VECM was estimated. This model incorporates the cointegrated variables and allows for the examination of potential long-run equilibrium relationships among them. The cointegration rank (r) was determined using Johansen's cointegration test, whose robustness was enhanced by incorporating the critical values proposed not only by Mackinnon et al. (1999) but also those by Osterwald-Lenum (1992) or both the trace and maximum eigenvalue tests. Additionally, following the recommendation of Lettau & Ludvigson (2001), a constant term was included in the cointegration equation.

This methodological process—from identifying the integration order to selecting the optimal cointegration rank—provided the foundation for the subsequent analyses. The detailed results of the cointegration test are presented in Appendix B, Table B.1. A summary of the main findings regarding the validity of the UIH is presented below.

[11] To rule out potential specification issues, tests for autocorrelation (Lagrange multiplier), heteroskedasticity (White's joint test), and normality (Jarque–Bera statistic) were performed. The results confirmed the absence of serial correlation at the 5% level, as well as homoskedasticity and normally distributed residuals.

Table 2 reports the results of the cointegration test applied to each country to assess the validity of the UIH. Among the 30 countries analyzed, only 21 satisfied the required stationarity conditions in both study periods (1991–2019 and 1991–2023), thus constituting the final sample eligible for this cointegration-based approach. Within this group, nine countries accepted the UIH in both periods, while in three others, the hypothesis was consistently rejected.

When comparing these findings with previous evidence (Table 1), notable discrepancies emerge in cases such as the United States, Ecuador, and Peru. While Emerson (2011) and Maridueña-Larrea & Martín-Román (2024a) rejected the UIH for these economies (using different frequencies and periods), the present study finds support for the hypothesis, suggesting that the extended time span and the use of annual data may capture long-run dynamics and structural changes that are not always observable in shorter or higher-frequency samples. On the other hand, the results are consistent with the literature for countries such as Canada (Tansel & Ozdemir, 2018) and Uruguay (Maridueña-Larrea & Martín-Román, 2024a), confirming the rejection of the UIH and, consequently, the possible presence of AWE or DWE in their labor market behavior.

Table 2. Results of the cointegration analysis (21 countries with I(1) series in both sub-samples: 1991–2019 and 1991–2023).

1991–2023 / 1991–2019	Rejection of the UIH	Acceptance of the UIH
Rejection of the UIH	Belize, Canada, Uruguay	Guatemala, Mexico
Acceptance of the UIH	Argentina, Bolivia, Guayana, Honduras, Panama, Saint Lucia, Suriname	Bahamas, Colombia, Ecuador, United States, Haiti, Paraguay, Peru, Dominican Republic, Venezuela

Note: According to the Johansen test (5% level), acceptance (rejection) of the UIH indicates no evidence (evidence) of cointegration between UR and LFPR.

In the remaining nine cases, results vary depending on the selected time frame. For instance, Guatemala and Mexico support the UIH for the full cycle (1991–2023) but reject it for the 1991–2019 period. The opposite occurs in countries such as Argentina, Bolivia, Guyana, Honduras, Panama, Saint Lucia, and Suriname, where the UIH holds during the 1991–2019 interval but is rejected once the sample is extended through 2023. This pattern suggests that, following the pandemic, some economies experienced labor market distortions associated with this exceptional event.

The COVID-19 pandemic generated unprecedented disruptions in global labor markets, including sharp increases in unemployment, declines in labor force participation, and sectoral reallocation of employment. Even where headline unemployment rates rebounded relatively quickly, labor force participation often lagged behind, with persistent disparities by gender, age, and skill groups. As emphasized in recent work, the pandemic also catalyzed structural changes in labor preferences—most notably the expansion of remote work and the increased salience of work–life balance—while exposing and amplifying pre-existing inequalities (Deng et al., 2023; Guven et al., 2023).

In this context, the mixed post-pandemic dynamics between unemployment and labor force participation can be interpreted through several interrelated mechanisms. First, job-search behavior may have shifted due to mental-health challenges and increased care responsibilities, affecting labor-market attachment and search intensity (Gabriel et al., 2021; Kwak et al., 2025; Xu & Savickas, 2022). Second, government support policies—such as wage subsidies and more generous unemployment benefits—may have stabilized household income and, in some cases, reduced the urgency or intensity of job search (Finamor & Scott, 2021; Porras-Arena et al, 2024; Soares & Berg, 2022; Wanberg et al., 2020). Third, labor supply preferences may have changed with expanded remote-work opportunities and a rebalancing between paid and unpaid work (Bernstein & Martinez, 2021; Krumel et al., 2023; Zhang et al., 2024). Taken together, these channels provide a parsimonious explanation for why unemployment and participation may display heterogeneous trajectories after the pandemic.

From a policymaking perspective, this could be attributed to the difficulty of effectively correcting long-run imbalances in an environment that remains unstable. Conversely, some countries may have implemented structural reforms or active labor market policies aimed at mitigating or reversing the effects of the crisis. As a result, when considering the full cycle, new AWE or DWE dynamics may emerge that were not evident prior to 2020.

Table 3 presents the results of the cointegration analysis for the sample of countries that meet the required stationarity conditions during the 1991–2019 period. Out of a total of nine countries, the UIH is accepted in five and rejected in four. Our findings continue to reveal discrepancies for certain Latin American economies. For instance, when comparing our results with the previous literature (Table 1), we observe that the UIH was previously rejected for Chile using quarterly data, whereas in our analysis, based on annual data, it is accepted. Nonetheless, our results are consistent with prior studies for countries such as Brazil, where the UIH is consistently rejected regardless of whether quarterly or annual data are used.

Table 3. Results of the cointegration analysis for 1991–2019 (9 countries meeting the I(1) requirement only in this sub-sample).

Period	Rejection of the UIH	Acceptance of the UIH
1991–2019	Barbados, Costa Rica, Saint Vincent and the Grenadines, Trinidad and Tobago	Brazil, Chile, Cuba, Jamaica, Nicaragua

Note: According to the Johansen test (5% level), acceptance (rejection) of the UIH indicates no evidence (evidence) of cointegration between UR and LFPR.

Overall, considering 30 countries (excluding El Salvador), the UIH is rejected in 7 countries and accepted in 14. In addition, 9 countries exhibit mixed dynamics (accept–reject): 7 would align with rejection in 1991–2023 (or with acceptance in 1991–2019), and 2 with rejection in 1991–2019 (or with acceptance in 1991–2023). If the mixed cases are added to the corresponding outcome by period, the consolidated totals are as follows: 16 countries reject the UIH and 23 accept it.

Based on the results rejecting the UIH, Table 4 classifies the countries according to the sign of their cointegration vector and the selected reference period.[12] Out of a total of 16 countries, only six exhibit an AWE pattern, while 10 display evidence consistent with the DWE. Therefore, the findings suggest a predominance of the DWE across the Americas, although this pattern appears to be highly dependent on the specific period chosen for the analysis.

[12] For further reference, see Appendix B, Tables B.2 and B.3.

Table 4. Effects derived from the rejection of the UIH (16 countries for periods 1991–2019 and 1991–2023).

Effects	Countries coinciding for 1991–2019 and 1991–2023			Only in 1991–2019
	Both periods	Only in 1991–2019	Only in 1991–2023	
AWE	-	Mexico	Argentina, Guyana, Honduras	Barbados, Costa Rica
DWE	Belize, Canada, Uruguay	Guatemala	Bolivia, Panama, Saint Lucia, Suriname	Saint Vincent and the Grenadines, Trinidad and Tobago

Note: AWE/DWE classification is based on the sign of the long-run relationship between UR and LFPR implied by the estimated cointegrating vector.

Figure 5 illustrates the evolution of the cointegration vectors for the 10 countries that reject the UIH and exhibit a DWE. The left panel shows that the magnitude of these vectors is heterogeneous among countries that rejected the UIH in both periods of analysis. These effects range from mild in Canada to more pronounced in Uruguay and Belize, where, on average, a one-percentage-point increase in the unemployment rate is associated with a decrease in the labor force participation rate of 0.05%, 0.74%, and 1.06%, respectively[13]. The right panel displays the countries that rejected the UIH in at least one of the two periods (see Table 4). The DWE exceeds 3% in terms of its impact on the participation rate per percentage point increase in unemployment in economies such as Guatemala (6.64), Panama (5.55), Bolivia (3.58), and Saint Lucia (3.05). In the remaining countries—Suriname, Trinidad and Tobago, and Saint Vincent and the Grenadines—the effect ranges from 0.25% to 1.75% in absolute value.

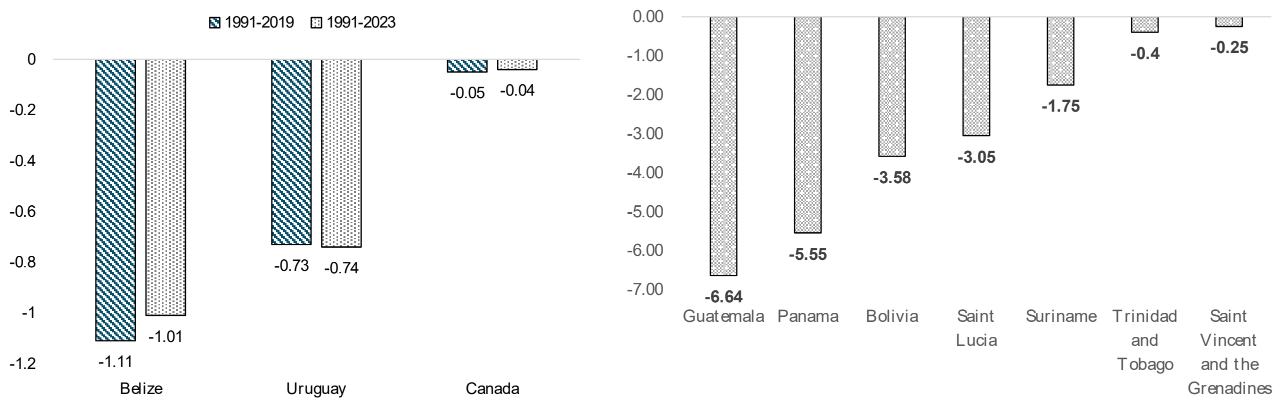

Figure 5. Cointegration vectors: countries rejecting the UIH and exhibiting a DWE.

Figure 6, in turn, presents the cointegration vectors for the countries that exhibited an AWE. Honduras and Barbados show the highest values (16.82 and 9.79, respectively), suggesting a strong labor supply response to increases in unemployment. In contrast, Mexico (1.36), Costa Rica (1.24),

[13] It is worth noting that the cointegration vector remained relatively stable across periods (1991–2019 and 1991–2023).

and Argentina (0.2) display much more moderate coefficients. Thus, the variations in the magnitude of the AWE—similar to those of the DWE—not only reflect high heterogeneity but also the economic and institutional configuration of each country. These findings highlight the importance of designing targeted policies to strengthen labor markets, as noted by Maridueña-Larrea & Martín-Román (2024b).

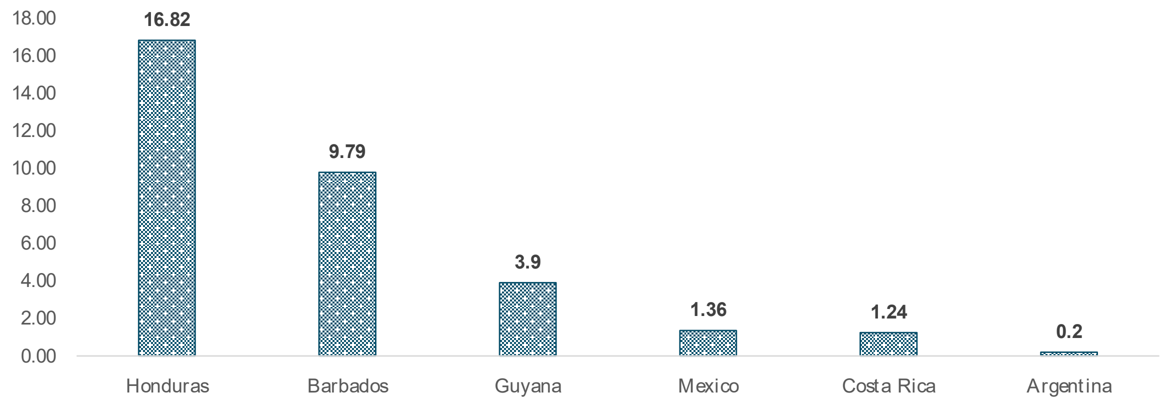

Figure 6. Cointegration vectors: countries rejecting the UIH and exhibiting AWE.

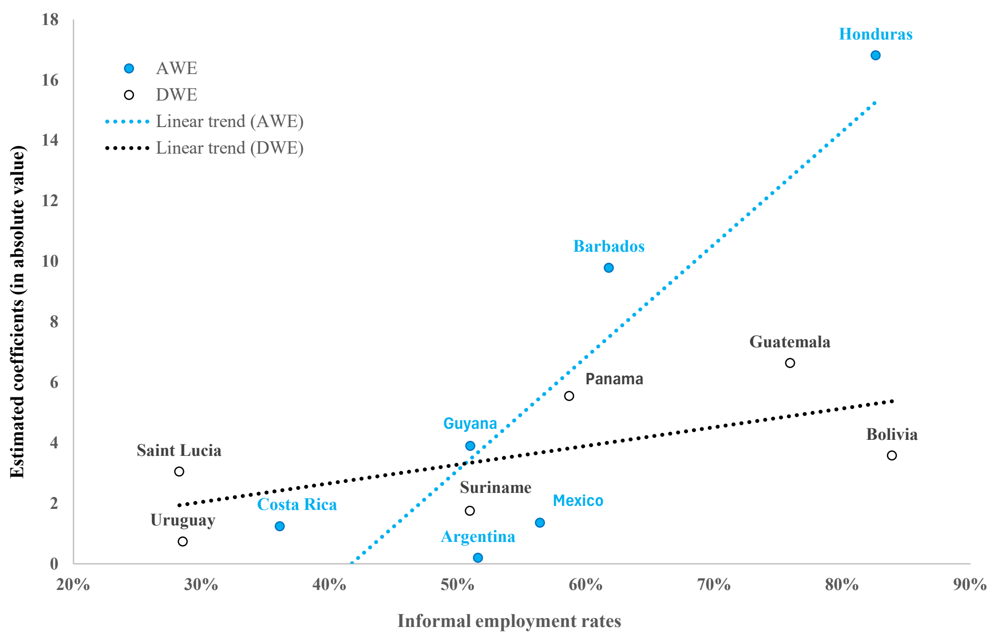

Figure 7. Relationship between informality and AWE-DWE (economies with UIH rejection). Note: Belize, Canada, Saint Vincent and the Grenadines, and Trinidad and Tobago are excluded because informality data were not available in the database used for the analysis. Source: ILOSTAT, Labour Force Statistics (LFS), accessed February 2025.

Overall, the findings reveal substantial heterogeneity in labor-market responses across the region, underscoring the importance of examining multiple time horizons (pre- and post-pandemic) and data frequencies. While most countries support the UIH, in those where it is rejected, distinct adjustment patterns emerge that align with either the DWE or the AWE, with considerable variation in their magnitude. Taken together, these differences reflect not only the structural and institutional features of each economy but also the ways in which economic agents adjust to changes in the unemployment rate.

An additional conjecture to interpret the observed heterogeneity relates to informality, which is particularly prevalent among the emerging and developing economies in our sample (15 out of 16 countries where the UIH is rejected). The literature suggests that, in highly informal settings, the adverse effects of crises may be intensified by deepening contractions in the formal economy, increasing labor vulnerability and complicating fiscal policy responses (Colombo et al., 2019; Razafindrakoto et al., 2024). In the same vein, using informal employment rates compiled from the International Labor Organization, we observe a positive association between informality and the estimated intensity (in absolute value) of both effects for the countries where AWE/DWE mechanisms are detected (see Figure 7).

Although these channels are not formally tested in this study, this evidence provides an additional interpretation and delineates a future research agenda aimed at linking effect predominance and intensity to informality, social protection coverage, and countries' level of development, as well as their demographic characteristics and institutions. In particular, advancing this agenda calls for approaches that jointly model these determinants and examine their interactions, alongside more robust policy evaluations to derive implications for inclusive labor-market strategies (Congregado et al., 2020; Manerkar & Chandran, 2024).

6.2. Robustness analysis

Given the variability observed in the cointegration vectors of the 16 countries analyzed, it is essential to subject these results to additional testing to validate the identified dynamics. Although the annual series used (33 observations for 1991–2023 and 29 for 1991–2019) are sufficient to estimate long-run relationships under the approach by Johansen (2002) and span a complete economic cycle, their moderate length may introduce bias into the results. To strengthen our conclusions, we apply the methodology proposed by Hjalmarsson & Österholm (2010), which involves imposing restrictions on the cointegration relationships to confirm their validity.

According to these authors, Johansen's tests may be biased when the variables possess near-unit roots (rather than exact unit roots), increasing the risk of detecting spurious relationships. To mitigate this issue, two restrictions are imposed on the estimated cointegration vectors: β = (1 0) and β = (0 1). If both restrictions are rejected, the cointegrated relationship is considered valid. Otherwise, it is inferred that the supposed cointegration is driven by the stationarity of a single variable—labor force participation in the first case, or unemployment in the second—without strong evidence of a long-run linkage. The results of this procedure are presented in Appendix B, Table B.4.

The imposition of restrictions on the estimated cointegration relationships was rejected, thereby reaffirming the results in half of the countries where the UIH was not confirmed. However, in the cases of Canada, Guatemala, Guyana, Honduras, Panama, Saint Lucia, Suriname, and Saint Vicent and the Grenadines, the results should be interpreted with caution, as the observed long-run relationship may be driven by a single stationary variable.

7. Conclusions and policy recommendations

The research findings clearly highlight the complexity and heterogeneity of labor market behavior in the relationship between labor force participation and unemployment rates, partially validating the UIH in the Americas. The observed diversity in the acceptance or rejection of the UIH, as well as the marked differences in the magnitude of the DWE and AWE effects, illustrate how labor market dynamics across the region are far from uniform and depend heavily on the specific institutional, economic, and social characteristics of each country.

From an analytical standpoint, the results underscore the methodological importance of adopting extended time horizons and varying data frequencies in economic analysis, as such approaches allow for a better identification and understanding of the structural nature of the observed effects. In particular, the contrast with previous evidence for countries such as the United States, Ecuador, Chile, and Peru suggests that annual data and longer periods provide additional insights capable of detecting effects that often go unnoticed in analyses based on short or high-frequency series. Moreover, the finding that some countries changed their validation or rejection of the UIH once the post-pandemic period was considered highlights the need to recognize and assess the persistent impacts of exogenous shocks on long-term labor trajectories.

Finally, the robustness checks emphasize the need to interpret cointegration relationships cautiously, especially in countries where results may be driven by the stationarity of a single variable. These findings suggest the importance of future studies that delve into additional econometric techniques and expand the scope of available datasets to produce even more robust and country-specific conclusions. We also note that cointegration-based frameworks typically rely on stable long-run relationships and largely linear adjustment dynamics over the sample period. Future research may therefore use complementary time-series approaches to explore, where relevant, whether the results are sensitive to alternative dynamic specifications and potential cyclical asymmetries.

From a policy perspective, the identification of a predominantly DWE in the region signals the urgent need to strengthen mechanisms that promote active job search during periods of economic slowdown. For countries experiencing pronounced impacts—such as Bolivia, Belize, Suriname, and Uruguay—priority should be given to reinforcing job intermediation services, training and reskilling programs, and conditional social protection measures tied to active job-seeking. Such initiatives could help reduce labor market discouragement, reactivate participation, and foster a more sustainable and equitable economic recovery.

In countries where the AWE prevails, it becomes essential to address the potential precariousness resulting from a surge of additional workers entering the labor market. For these economies, it is crucial to implement policies that ensure not only the quantitative expansion of employment but also its quality. In this regard, policies aimed at enhancing labor formalization, contract stability, and adequate working conditions are key to constructively harnessing the added worker effect.

In addition to the policy recommendations already outlined, the findings of this study point to several complementary lines of action that could be highly relevant for policymakers across the region. First, it is essential to develop high-frequency labor market monitoring systems capable of detecting early shifts in labor force participation linked to fluctuations in unemployment. These early-warning mechanisms would enable timely interventions, particularly during economic downturns, when the DWE or the AWE may intensify rapidly.

Furthermore, given the evident differentiation of these effects across gender and other demographic variables, it is advisable to move toward more targeted labor policies. For instance, in contexts where AWE is more pronounced among women, the expansion of public childcare services, equitable parental leave policies, and measures to promote shared caregiving responsibilities can help prevent female labor market entry from occurring under conditions of informality or precarity.

Additionally, the results of this research call for a rethinking of counter-cyclical macroeconomic policies. In countries where the UIH does not hold, stimulus measures that explicitly consider the structural linkages between unemployment and labor force participation may be more effective in fostering both job recovery and labor reactivation. Incorporating the UIH framework into the design of fiscal and monetary policies could enhance the coherence and impact of economic responses to downturns.

Finally, it is crucial to institutionalize regular ex post evaluations of labor policy outcomes. This would allow for a systematic assessment of their effectiveness in reducing labor market discouragement or preventing job quality deterioration. In contexts of high economic volatility, such evaluations become even more vital, as labor dynamics may undergo structural changes that require agile and empirically grounded policy responses.

Use of AI tools declaration

During the preparation of this paper, the authors used [ChatGpt4o] to improve writing. After using this tool/service, the authors reviewed and edited the content as necessary and take full responsibility for the content of the publication.

Author contributions

Conceptualization, Á.M.-R.; methodology, Á.M.-R.; software, Á.M.-L.; validation, Á.M.-R. and Á.M.-L.; formal analysis, Á.M.-R. and Á.M.-L.; investigation, Á.M.-R. and Á.M.-L.; resources, Á.M.-L.; data curation, Á.M.-L.; writing—original draft preparation, Á.M.-L.; writing—review and editing, Á.M.-R.; visualization, Á.M.-R. and Á.M.-L.; supervision, Á.M.-R.; project administration, Á.M.-R. and Á.M.-L.; funding acquisition, Á.M.-R. All authors have read and agreed to the published version of the manuscript.

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Conflict of interest

All authors declare no conflicts of interest in this paper.

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